



Kari Takala

Studies in
Time Series Analysis of
Consumption, Asset Prices
and Forecasting

Bank of Finland Studies
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and Forecasting**

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Abstract

This collection of seven papers deals with three different areas of econometric applications: consumption, asset prices, and forecasting. The papers apply techniques related to the analysis of unit roots and cointegration methods.

The first paper deals with consumption theories and formulates an error-correction forecasting model for consumption. A single cointegration relationship is found between consumption, income and net wealth, which is in line with the permanent income hypothesis. The second paper studies the excess sensitivity of consumption to current disposable income. Estimating the coefficient with time-varying techniques, we notice a decline in the coefficient during the period of financial deregulation toward the end of the 1980s and a rise during the recession. Third paper takes a closer look at how useful consumer barometer variables can be in forecasting variables such as consumption and inflation.

The first paper on asset prices, is based on the theory of cointegration between house and stock prices, which asserts that real after-tax risk-adjusted returns on assets should coincide in the long run. This paper presents a model for house prices that uses stock prices as a leading indicator to improve the forecasting of housing prices. Another paper on asset prices considers cointegration between house prices and inflation, and finds eg that house prices adjust to consumer prices in the long run and that no excess real appreciation, apart from rental income, is derived from house ownership.

The two last papers deal with bankruptcy forecasting and testing for nonlinearities and chaos. It is asserted that bankruptcies can be interpreted as error-correction between supply and demand. Many tests have been developed to study the presence of nonlinearities in economic series. The results of testing unambiguously support that there are strong nonlinearities in economic data, but the evidence for chaos is weak.

Key words: cointegration, asset prices, forecasting, nonlinearity, bankruptcy

Tiivistelmä

Kokoelma sisältää seitsemän erillistä selvitystä kolmelta soveltavan aikasarjaekonometrian erityisalueelta: kulutustutkimuksesta, varallisuushinnoista ja ennustamisesta. Selvityksissä käytetään hyväksi viimeksi kuluneen vuosikymmenen aikana laajasti sovellettuja ekonometrisia yksikköjuuri- ja yhteisintegroituvuustekniikoita.

Kulutusfunktioon liittyvistä selvityksistä ensimmäisessä muodostetaan kulutuksen virhekorjausmalli, jossa virhekorjaustermi on estimoitu kulutuksen, tulojen ja nettovarallisuuden systeemistä. Tulokset ovat sopusoinnussa pysyväistulohypoteesin kanssa. Seuraavassa tutkimuksessa selvitetään kulutuksen liikaherkkyyttä nykytuloihin nähden kiinteiden ja vaihtuvaparametristen mallien avulla. Vaihtuvaparametristen estimointien perusteella likviditeettirajoitteisten kuluttajien osuus väheni jonkin verran rahamarkkinoiden vapautumisen aikana, mutta kasvoi 1990-luvun lamakautena. Kolmannessa selvityksessä taas analysoidaan erilaisten barometrikyselymuuttujien hyödyllisyyttä mm. kulutuksen ja inflaation ennakoinnissa.

Sijoituskohteiden hintoja mallintavista artikkeleista ensimmäinen perustuu hypoteesiin asuntojen hintojen ja pörssikurssien yhteisintegroituvuudesta. Hypoteesin mukaan sijoituskohteiden reaaliset veronjälkeiset riskittömät tuotot eivät pitkällä aikavälillä poikkea toisistaan. Täsmennetyssä virhekorjausmallissa asuntojen hintojen kehitystä ennakoivana indikaattorina käytetään pörssikursseja. Tuoreemmassa sijoituskohteiden hintoihin liittyvässä mallinnuksessa selvitetään asuntojen hintojen ja kuluttajahintojen yhteisintegroituvuutta. Tulokset viittaavat siihen, että asuntojen hintojen muutokset sopeutuvat pitkällä aikavälillä yleiseen inflaatiovauhtiin ja ettei asunnon omistamisesta ole odotettavissa ylimääräistä reaalista arvonnousua.

Kahdessa viimeisessä tutkimuksessa selvitetään konkurssien ennustamista sekä epälineaarisuuden ja kaaoksen testausta taloudellisilla aikasarjoilla. Konkursseja voidaan analyysin mukaan pitää kysynnän ja tarjonnan tasapainovirheen indikaattorina. Viime vuosikymmenen aikana on esitetty useita stokastisten aikasarjojen epälineaarisuutta mittaavia testisuureita. Estimointien perusteella näyttää siltä, että taloudelliset muuttujat ja niiden väliset riippuvuudet sisältävät usein epälineaarisia riippuvuuksia, mutta todistusaineistoa deterministisestä kaaoksesta taloudellisissa aikasarjoissa on vähänlaisesti.

Asiasanat: yhteisintegroituvuus, sijoituskohteiden hinnat, ennustaminen, epälineaarisuus, konkurssit

Foreword

I would like to thank all the colleagues I have had the pleasure to work with on these and other papers. I'm also grateful to the collaborators on the papers for permitting me to use the joint papers in this collection, as well as the publishers – the Bank of Finland, Finnish Economic Papers, Physica-Verlag and Government Institute for Economic Research – for allowing reprinting of the papers.

I would also thank the advisers and referees as well as the official examiners for the study – Pasi Holm and Erkki Koskela – for their guidance. Needless to say that they are not responsible for the remaining deficiencies. In particular I must thank Professors Matti Virén and Erkki Koskela for their advise over the years. Without them, this study most likely would never have been completed. Lastly I'm also indebted to Glenn Harma and Malcolm Waters for correcting the language and to Päivi Lindqvist for finalising the layout of the papers.

Helsinki, December 2001

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1 Introduction

This introduction briefly describes the state of time-series econometrics as it forms the methodological basis for this collection of empirical papers. First, recent trends in time-series econometrics as regards forecasting and cointegration are described at a general level. Then in the summary section these are examined more closely in the context of the papers.

Although economics is often regarded as a highly theoretical science, in essence economics is an empirical discipline. It is no exaggeration to say that in the final analysis there is only empirical evidence in economics. In principle, economics does not even constitute a science based on strict laws.¹ Economic systems are essentially invented and fictitious rather than discovered and hence economic theories are more like tools than representations of the actual economy.² Economic laws are more or less generalisations of behavioural regularities that are true only in statistical or probabilistic terms.

Economic theories themselves can be derived by generalising from observed phenomena or via algebraic derivation and logical inference based on modelling. But in any case they need to be verified (or falsified) by empirical testing. Probably the hardest job in empirical econometrics is to falsify all the irrelevant hypotheses, since there are usually lots of possible explanations for observed phenomena. Economic observations are almost without exception blurred and obscured and seemingly contain much measurement error and other kinds of noise. Therefore economic relationships are essentially stochastic, which necessitates the use of statistical theories, models and inference to verify economic theories. Statistical theories and tools such as estimation methods are always needed when economic theories are verified in practice.

However, this is not to say that economic theories are useless. On the contrary, theories are essential in guiding our thinking, in giving incite into what are the important questions, in clarifying the structure of problems faced by real economies, and in warning us of the

¹ There are only a few relationships referred to as 'laws' in economics. These are typically like Okun's law, which merely points out that there is negative correlation between economic growth and the unemployment rate and that growth must exceed a certain rate in order to have a lowering effect on the unemployment rate. Many laws in classical physics are deterministic rather than statistical. In economics only relations similar to the quantity theory of money $MV = PT$ come close to being such laws.

² See Judge, Griffith, Hill, Lutkepohl and Lee (1985, p. 2).

uncertainties in conducting economic policy. Economic theory is concerned with explaining relationships among economic variables, whereas statistical models are concerned with drawing conclusions from limited data. Solid economic theory is therefore always one of the most important ingredients of empirical econometric work. As a social science, economics has remained largely non-experimental and has been based widely on statistical testing of economic hypothesis. Theories may be created in many ways, but empirical evidence must be replicable if it is to be useful in an ever-changing economic environment.

It is likely that in most cases there is no single theory that will entirely explain a set of observed economic phenomena or variations therein. Thus we need several theories in order to explain eg consumption behaviour. Consumers can be classified into a few groups of economic agents, each group behaving according to different rules of conduct. When we are building large models or even single behavioural equations, it is difficult to take all of these partly conflicting theories into account. Therefore we must adopt the most important features in our empirical models and perhaps to allow for some serial autocorrelation in the error terms because of the missing theories.

2 Statistical models and economic analysis

The most important application of economic theories is to improve economic policy. For instance, Gourieroux and Monfort (1997) define the goal of empirical economic analysis as that of highlighting the economic mechanism and decision-making. Forecasting could be added to these. Many organisations prepare their policies by making various calculations and set their control instruments on the basis of predictions as to macroeconomic activity.

In formulating and testing interrelations between economic variables, system methods are needed. The link between economic theory and its adequacy with respect to reality is a statistical model. By definition, a statistical model is a parameterised family of probability measures (Johansen, 1995). More broadly described, a statistical model gives the set of assumptions on which the empirical sample data is generated. The difficult task in testing economic hypothesis is that they are always conditional on specific assumptions about the statistical model. Thus, in testing economic theories, we must be sure that the underlying statistical method is adequate for the

circumstances. Otherwise we cannot be sure whether we are testing the theory or just the statistical assumptions behind the applied method.

Varying modelling assumptions has led into a wide range of specific econometric models, and usually there is no single reliable strategy to choose between different statistical models. For instance, linear regression model diagnostics as such are often insufficient as a guide to choosing between models. Nor does simple maximisation of the coefficient of determination or use of model selection criteria produce an unambiguous solution. Usually we end up with few plausible specifications from which we must make a choice.

Statistical models are usually developed in order to describe fluctuations in data and to express economic theory in terms of the parameters of a statistical model. After the researcher has proved by testing that the model is data-congruent, an economic hypothesis can be formulated as parametric restrictions that are then subject to confirmation or rejection based on test statistics. The purpose of a statistical model is to provide a framework for analysing and testing different economic theories against observed data. In many cases economic theories contain abstract concepts, eg permanent income or natural rate of unemployment, that are not directly observable and hence must be replaced by empirical proxies. Therefore we cannot be certain whether a theory is good. Even testing is insufficient to make this determination, due to mismeasured concepts. For instance, Kotlikoff (1989) asserts that in economics the operational variables are very different from the theoretical concepts they intend to measure. From the statistical viewpoint, it is rather straightforward to determine whether a specific economic model is a valid description of the sample data. Unfortunately, reality may also change over time, so that a model based on a fixed regime or fixed parameters can be a bad choice, especially for forecasting. For example, financial deregulation in the late 1980s changed the role of the real interest rate in determining consumption. The elasticity of interest rate became negative after 1986, when lending rate regulation was abolished. Since then, interest rate elasticity has been affected by the 1990s recession and tightening of liquidity constraints at that time.

3 Economic forecasting

One of the most important reasons for econometric modelling is forecasting. Economic theory is important in guiding empirical specifications, eg in selecting exogenous variables, functional form, order of integration, nature of relationships etc. However, the gap between theory and empirical implementation is often a wide one, albeit recently developed econometric testing methods have been highly useful in model specification and in choosing between various models.

Economic forecasts are generally necessary to successful policy-making. During the past twenty years time series analysis has greatly changed the practice of empirical econometrics, due to both improvements in the statistical tools and the appearance of powerful econometrics software for PCs, which has facilitated the testing of economic hypotheses and provided easy access to a variety of new modelling techniques. Theory and testing based on linear models and relaxation of the basic regression assumptions has moved economic analysis toward more realistic assumptions about dependencies between variables, relationship forms, distributions and other aspects of statistical models.³ The forecasting of nonlinear processes and non-stationary series has been developed to account for special data features associated with obvious nonlinearities, eg in production functions (see eg Granger and Teräsvirta, 1993).

Univariate time-series models have been widely used in quick forecasting tasks, but when more emphasis is placed on causes and effects of different explanatory factors, one must resort to model-based forecasting.

Many useful modelling tools have been developed to help with specification of the forecasting model. For instance, for selecting variables to include in a forecasting model, the Granger causality test is very helpful in that it can efficiently eliminate spurious regressors from dynamic models. Thus this test is crucial for finding the relevant exogenous variables for forecasting models and checking whether feedback relations are present. Only by knowing the endogeneity/exogeneity of model variables can defects such as simultaneity bias be avoided.

³ Widely used econometric software packages nowadays provide a wide range of possibilities for testing the assumptions behind an OLS regression: weak-exogeneity, residual non-autocorrelation, homoscedasticity, normality, model stability and non-multicollinearity.

Suppose we could agree on what variables to include in a model. It is now possible to study the model's stability, ie variation over time of parameters, using recursive, rolling or time-varying estimates. Changing parameters are important to recognise in the process of producing a forecast; otherwise significant bias may appear in the forecasts. In many cases this type of analysis has helped us to determine the relevant theories for explaining relationships between economic variables. Encompassing tests have been included in packages that help us choose between rival theory specifications of models explaining a particular dependent variable.⁴ Ex post forecast comparison and calculation of accuracy measures are often also needed in choosing between forecasting models.

4 Recent trends in applied econometrics

For over a decade now econometrics has been going through a revolution concerning analysis of unit roots and cointegration. Variables that have time-varying means and variances are non-stationary and may have unit roots. And non-stationary variables may have common trends and may form stationary linear combinations (usually based on equilibrating long-run relationships). A cointegrating relationship based on economic theory can be extremely useful in economic forecasting, since it would imply rules of convergence toward a specified long-run equilibrium. For instance, cointegration between consumption and disposable income implies a stationary saving rate (Engle and Granger, 1987). In earlier economic studies, spurious regressions emerged from misleading inference based on analysis of unit root variables. Now, identification of cointegrating relations helps us find the proper economic mechanisms.

Traditional statistical concepts concerning stochastic processes, eg non-stationarity and exogeneity, have been highly scrutinised and extended and integrated into the basics of time series analysis. The essence of time-series analysis is the measurement of relationships between groups (vectors) of economic variables. These new cointegrating VAR methodologies have also helped to model both the

⁴ A model is said to encompass a rival formulation if it can explain the results produced by the rival formulation. The rival model contains no information that could be used to improve the preferred model. A good model should not only explain the data but should also explain why rival models fail to account for the same data. The model that encompasses the other should also explain the latter's forecast errors.

long-term equilibrium and short-term dynamic relationships. Another major feature of time series models is, of course, the dynamic adjustment that characterises the variables in the models. Economic adjustment is rarely frictionless.

Thus it is possible that the major recent advances in time series analysis have been achieved by introducing ideas for solving the treatment of long-term relationships (cointegration), economic causality and specification of econometric models. The concept of cointegration was originally formulated by Granger (1981) to describe a well-known statistical property in an economic context. However, models entailing error-correction-type equilibrating processes had been used in economics since the early 1960s. Cointegrating relationships have since formed the backbone of many long-term restrictions also in large macroeconomic models that require long-term anchoring for long-term simulations. Besides the ongoing progress in respect of cointegrating relationships, further gains have been made in unit root limit theory. Econometric research of this type may also radically affect the way in which non-stationary variables are forecasted (Clements and Hendry, 1999). Many areas of utmost importance to macroeconomics have benefited from these advances in statistical methods.

Common trends and equilibrium error-correction terms based on cointegration have been used in many key behavioural equations such as consumption and investment functions, foreign demand and price equations, and other equations based on optimisation. Macroeconomic price series are also often connected with the path of the general price level, which involves some sort of equilibrating mechanism of returns. During the 1980s we saw many advances in the analysis of linear systems and especially for VAR models. Testing for common trends, analysis of cointegrating rank and specification tests for conditioning variables have given empirical model construction a more profound statistical basis.

Roughly speaking, the problem in economic modelling is to divide the variation in an observed dependent variable into systematic conditional dependence and irregular residual components. The problem is that there is an almost infinite number of ways to do this. In time series analysis it is assumed that the data consist of a systematic pattern (usually a set of identifiable components) and random noise, which usually makes the pattern difficult to identify. Most time series analysis techniques involve some form of filtering out the noise in order to make the model and the essential relationships clearer.

More recently, econometric research has focused on questions not directly related to linear regression models but rather concerning nonlinear models, nonparametric estimation and non-stationary processes with structural breaks. Special statistical assumptions are also needed to analyse dependent variables with limited range, eg discrete-choice variables and time series panels from different countries. In order to allow for different assumptions concerning statistical convergence criteria or statistical distribution of residuals, more attention has been paid to robust estimation procedures that are not based on the normal distribution.

Although many economic relationships are felt to be actually nonlinear, empirical testing is often aimed at merely approximating these relations using linear models, in which the individual contributions of the separate exogenous variables are simply summed together. Recently many tests have been developed to study whether stochastic series include non-linearity. This non-linearity arises mostly from budget constraints, implicit limits, buffer stocks and institutional regimes. In practice, the most serious problem is that there are too many possibilities for modelling nonlinear relationships and we lack effective strategies for choosing between the different functions. Economic theory in itself very rarely suggests any specific statistical family for an economic functional relationship, and this causes difficulties right at the start of the specification.

In practice, the data generating process for an observed economic time series is always unknown. Because in the end economic evidence has to be empirically tested, economic theories must be revealed by sequential and successive data modelling and testing. While economic theory is a good guide for searching for economic invariance, no theory can be approved without empirical testing, since economic relationships are most likely stochastic. However, it is clear that most economic debates and contradictory arguments can be solved only by means of empirical estimation and testing.

5 Summary of the papers

So far we have broadly discussed the role of economic theory and recent advances in time-series econometrics. Next we will consider in greater detail how these are connected within the topics discussed in this collection of papers. In this section we focus on how developments in econometrics methods are reflected in the papers. The articles below are classified under three major topics, although

they share some common dimensions beyond the broad field of research. For instance, the two papers on consumption and asset prices both use cointegration techniques and end up proposing specific forecasting models. Even though the last group of papers includes some that emphasise forecasting models, these papers also deal with wider econometric issues such as nonlinearities. The data used in most of the papers consist of publicly available time series, produced mainly by Statistics Finland.

Many of the papers also deal with certain aspects of forecasting. Takala (1995), Barot and Takala (1998) and Takala and Pere (1991) all specify explicit forecasting models. Possible cointegration properties between variables are also taken into account to deal with long-run adjustment to the equilibrium level, which always affects the long-run forecasts.

6 Consumption determination

Theories of consumer optimisation with rational expectations and emerging cointegration techniques have been developed in detail and applied to consumption analysis.⁵ The literature on consumption is vast, but one of the most coherent treatments of consumption analysis can be found in Deaton (1992).⁶

It is fair to say that today the life-cycle permanent income hypothesis (LC-PIH) is the premier basis for analysis of the consumption function. In the basic LC-PIH model wealth is simply accumulated saving, which is available for consumption. The permanent income hypothesis asserts that consumption depends on current labour income, expected present value of earnings and cumulated net wealth.

The papers by Takala (1995, 2000) consider the analysis of consumption and specification of the consumption function. In many respects consumption analysis has been one of the key areas in applied econometrics, both theoretically and methodologically. The Takala (1995) paper tests for cointegration between consumption, income and net wealth and finally proposes an error-correction model for explanatory and forecasting purposes. In the literature, net wealth has

⁵ Engle and Granger (1987) used consumption and income as an example of a cointegration relationship in their pioneering article.

⁶ Another broad survey worth mentioning is Muellbauer and Lattimore (1995).

seldom been included directly in the cointegration relation.⁷ The estimation is operationalised in a multivariate cointegration framework. The emphasis is also on the role of wealth in the consumption function. One key issue is how many cointegration relations there exist between consumption, income and net wealth. In accord with permanent income hypothesis, the tests using Finnish data favour the existence of just one cointegration relationship, which greatly simplifies the construction of an error-correction model for consumption.

Net wealth (at market value) is included in the system to allow for capital gains in real estate wealth, which means that wealth is not simply accumulated financial savings. Wealth was also disaggregated into financial assets, real estate wealth and debt, in order to confirm the existence of one cointegration relationship and to test the liquidity of the different assets, in order to determine the best wealth measure for consumption function. The tests suggest that a broad concept of net wealth, including all resources available for consumption, would be the preferred specification. It was also found that the average propensities to consume are different for financial wealth and real estate wealth.

The results from the cointegration system show that neither disposable income nor net wealth can be regarded as weakly exogenous in the system nor can either be excluded from the system of endogenous variables. Thus there does not exist a separate, statistically meaningful consumption function. Despite this, effort was made to explain which variables in the past had the greatest effect on consumption, and an error-correction consumption function was estimated. In building empirical models, one of the key elements is their stability in forecasting. In addition to the integrated core of the consumption model, ie income, net wealth, few weakly exogenous stationary variables; the expected real interest rate, inflation, unemployment rate and relative prices of consumption subgroups were found to be significant. The paper shows that the Finnish consumption function is more stable and accurate when real net wealth is included in the cointegrating relation as part of permanent income.

After Robert Hall's (1978) contribution using the Euler equation approach, the benchmark for the life-cycle/permanent-income hypothesis has been the idea that consumption should follow a random

⁷ A related paper that uses two cointegration relations is Hendry and von Ungem-Sternberg (1981) and its updated version, Hendry (1994). The wealth effect is taken into account in Patterson's (1991) study.

walk.⁸ Therefore consumption changes should be unpredictable white noise and orthogonal to any lagged publicly available information, including lagged changes in income. However, in many studies consumption growth has been observed to be sensitive to changes in income, which has generally been interpreted as indicating significant credit constraints.

The Takala (2000) paper tests in more detail the excess sensitivity of consumption to changes in disposable income. The Euler approach allows direct estimation of the parameters of the utility function. The problem lies in determining the proper functional form of the utility function. Several alternatives appear in the literature, but iso-elastic functions have gained the most attention recently. Of course there are some drawbacks also in the Euler equation approach, eg it does not hold under liquidity constraints. The specific formulations of the consumption function tested here are modifications of the Campbell and Mankiw (1991) model, with total consumption divided into the consumption of life-cycle consumers and that of liquidity-constrained consumers, who are able to spend only current disposable income in each period.

It is thought that most consumers are able to plan their consumption over a fairly long horizon, which means that consumption is determined by permanent income, which indicates persistent ability to spend. The problem is to determine whether consumption tracks current income too closely to support the PIH and, more generally, which is the best theory of consumption. Since we do not have actual macroeconomic data on consumption over the life cycle of different generations or of liquidity-constrained consumers, we must make assumptions on the characteristics of these groups and estimate a mixed function that identifies also the liquidity-constrained consumers, ie those who are forced to spend their entire current income.

This paper performs excess-sensitivity testing of consumption with Finnish data. It also investigates the probable causes of this sensitivity using different instruments and conditioning variables in time-varying sensitivity estimations. Most of the estimates indicate that about a half of the consumers were liquidity constrained during 1980–1998, which is among the highest in international comparisons. However, in

⁸ The Euler equation approach to consumption is not actually a parametric consumption model. It merely sets out the first order conditions that an optimal consumption path should fulfil. As these are only first order conditions, they are necessary but not sufficient for an optimal consumption plan. On the basis of these conditions on current and past consumption, one can test different models against each other.

estimating the excess sensitivity coefficient using time-varying techniques, we see a decline in the coefficient during the period of financial deregulation in the late 1980s and an increase during the recession. Finally this parameter has declined by the end of 1990s significantly. This is an important finding since the use of constant parameter models may induce biased forecasts.

Fixed parameter estimates of disposable income are not highly sensitive to assumptions on normality, estimation method or instrument variables. However, the excess-sensitivity parameter is sensitive to assumptions about which variables are allowed to have time-varying effects. The results are relatively unstable unless we agree on the proper conditioning variables and the model specification by conditioning estimates.

Djerf and Takala (1997) takes a closer look at how useful consumer barometer studies can be in forecasting economic variables such as consumption, consumer prices, household borrowing and saving. Since the rational expectations revolution, much effort has been expended in studying the information content of subjective expectations of households and consumers and their predictive power.

Data on business and consumer confidence (barometer data) are often ordinal or categorical, rather than continuous. The qualitative nature of barometer data implies some loss of information, but often the data are easier to collect and less costly. There are other problems with such data, eg thresholds (Zimmermann, 1997). The purpose of survey data is to complement our knowledge, via greater timeliness or sometimes by measuring economic variables that cannot be directly observed. Since the expectations revolution in economics in the 1970s, a great deal of international effort has gone into the development of barometer variables so as to improve forecasting. In Finland, however, the significance of these variables has been recognised only fairly recently.⁹ Barometer studies are an even more recent tool in consumer economics than eg are confidence surveys concerning manufacturing output, investment plans etc.

Comparisons between consumer sentiment variables and real economic activity indicators sometimes show a very close correspondence, eg between household expectations and realised changes.¹⁰ In some cases the expectations horizon is shorter than the

⁹ Carroll et al (1994) study the usefulness of consumer sentiment in respect of the United States and Berg and Bergström (1996) of Sweden.

¹⁰ Because barometer series are usually stationary, they do not involve cointegration. However, it has been observed that barometer variables often correlate strongly with annual differences in actual economic variables.

planning horizon, but even then expectations contain some true predictive power. Using Granger causality tests it was shown that household expectations can be truly forward-looking and not simply extrapolations from the past. An interesting case is private non-durable consumption, where predictive power is limited according to Hall's random walk hypothesis of consumption. Consumer sentiment variables are usually more volatile than actual macroeconomic changes and can also produce false signals.¹¹

Based on Finnish consumer barometer data, it seems that sentiment variables can be helpful particularly as regards small scale forecasting models and pure time series models. For policy uses, the fact that a consumer confidence index predicts consumption gives some scope for fiscal policy, since consumption can be affected through government influence on household expectations. For monetary policy targeting, household inflation expectations seem to be useful in assessing the real interest rate.

7 Asset prices and cointegration

Theory and statistical models are often much more closely aligned in financial econometrics than elsewhere in economics. For example, the hypothesis that asset prices should follow a random walk or martingale model can be very explicitly tested by statistical methods. Financial economics has more often focused on the returns of different assets rather than the prices themselves. Although the theoretical price of an asset is the expected accumulated net capital income, valuation changes divert in the short run the price from its equilibrium path.¹² This is natural in the sense that capital markets should equate returns whereas prices can vary widely, depending on market demand-supply imbalances. Uncertainty plays a major role in the pricing of financial assets. The predictability of asset returns is related to their conditional distributions and in particular to the manner in which conditional distributions evolve over time.

The two papers in this collection that deal with asset prices begin by sketching the theory of cointegration between two key assets –

¹¹ For example, consumer confidence suggested a slowdown in growth in early 1994, which did not occur, but consumers did not foresee the temporary growth slowdown of the Finnish economy in late 1995.

¹² Campbell, Lo and MacKinley (1997) provide a good review of the methods and testing of financial markets.

housing and shares – based on the result from general capital market equilibrium that real after-tax risk-adjusted returns on different assets should coincide in the long run.¹³ The total yield for any asset is composed of capital income and valuation. The latter refers to the appreciation (valuation), which depends on supply-demand balance, and usually dominates changes in prices of shares and housing. If an asset is not traded, ie held forever, capital income is the sole yield component. This gives us a clear reference point for the long-run return. In practice this means eg that the price of a dwelling is the discounted present value of the expected rents and the price of a share of stock is the present value of the future dividends.

Economic behaviour is based on optimal predictions, which are essentially statistical expectations with respect to defined information sets. Although economic agents usually do not use statistical models to form their expectations, the process can be imitated by estimating simple time series models. Assessing the predictability of asset returns has been a major interest of many economists and investors. So far the basic theories emphasising the efficient markets hypothesis have been widely accepted. Long-run returns and therefore hypotheses on how prices evolve over the long run are important for detecting contradictions vs underlying theories on random walks for prices. The hypothesis of long-run cointegration between major household assets prices has important implications for household portfolio management. In making return-risk choices, households are also making key investment decisions concerning lifetime wealth and hence lifetime consumption and utility. From the forecasting standpoint, martingale and random walk models imply unpredictability of price changes and define the weak-rationality of these pricing models in terms of market expectations.

The two papers use the assumption of efficient markets as a reference basis. The efficient markets hypothesis is of course an abstraction, like so many other useful concepts in economics, eg perfect competition, representative agent, equilibrium level. The more efficient the market, the more random and unpredictable the price changes. The asset price models used in these papers are based on present value models, in which prices are represented as discounted present values of future capital income. For shares, capital income is dividends and, for housing, capital income is rent. Thus persistent changes in capital income have larger effects on asset prices than do

¹³ Closely related papers on theoretical foundations and treatment are Campbell and Shiller (1987, 1988).

temporary changes in dividends. Price bubbles can also lead to price volatility without introducing predictability of returns. To the extent that price movements are explained by changes in returns they cannot be regarded as bubbles.¹⁴

The first paper, Takala and Pere (1991), formulates and tests the efficient markets pricing hypothesis and uses a two-step Engle-Granger error-correction model based on this restriction. Campbell et al (1997) note that one of the oldest questions in financial econometrics is whether asset prices are predictable. This paper introduces a model for housing prices in which share prices are a leading indicator that improves the forecasting of housing prices. Housing and share price indices are found to be unit root processes, but not pure random walks, which means that they could be cointegrated. Based on single-equation cointegration tests, these price indices were found to be cointegrated in the period 1970–1990. Thus an error-correction model was specified for forecasting changes in housing prices. Since this paper was written, much has happened in the market environment, particularly as regards the Finnish stock market. The Helsinki stock exchange has opened up largely because of advances in information technology. Nokia's domination of the HEX stock exchange and foreign ownership has not loosened the cointegration with the housing markets (especially in Greater Helsinki), probably contrary to expectations.

The more recent paper, Barot and Takala (1998), considers the long-run cointegration between housing prices and inflation, using the multivariate Johansen approach. The null hypothesis is that in the long-run housing prices follow and adjust to overall consumer prices and no excess real gain, apart from rental income, is received from housing ownership. The approach here is analogous to that of Villani (1982), with minor modifications. We define the price of housing to consist of three present value components: housing services, housing ownership tax advantages, and mortgage finance. An expression is needed for capital-user cost of housing ownership. We must also calculate the real return on housing ownership, net of financing and tax costs. To simplify the presentation, we ignore specific housing financing arrangements available in Finland and Sweden, and instead assume that the lending rate equals the opportunity cost of capital.

¹⁴ Present value models can also be formulated to have time-varying returns. In principle, high prices should predict large dividends or rents and low prices low returns. Even if the variability of expected returns is small, their persistence can lead to large changes in asset prices.

We postulate a long-run relationship between consumer prices and housing prices. The estimated system models are based on cointegration analysis of housing prices and consumer prices, and so we need to provide an explanation for the equilibrium mechanism between these variables. In the user-cost calculations, consumer goods and housing are alternative expenditure items, and due to cointegration their relative price is assumed to be stable. The consumption capital asset pricing model (CCAPM) implies that marginal utility of consumption is equal to the real return on housing (eg Breeden 1979). Briefly, cointegration follows from the fact that marginal utilities from different types of consumption tend to be equal, which leads to a stationary marginal rate of substitution. Because in the long run marginal utilities equal real after-tax returns, housing prices appreciate at the rate of inflation and there are no excess real returns. Equal expected returns will also lead to cointegrated price indices.

Thus the paper gives the important insight that real housing prices are a stationary variable in the long run. In addition, housing prices and consumer prices have the same growth rates in the long run, which can be used to estimate equilibrating housing prices and enable us to compare actual and equilibrium housing prices in each period. In principle, this asserts that houses are risk-free assets, like consumption in general; otherwise, a risk premium would have to be introduced for (risky) returns. Equilibrium is defined as a state in which there is no inherent tendency to change. However, equilibrium could be classified further into local and global equilibria, depending on whether shocks can be persistent or not.

Given the emphasis on price stability in monetary policy, the concern caused by rapid increases in housing prices, reflecting eg income prospects and credit growth, is quite understandable. It is believed that such rises could provide an early indication of inflationary pressure. Housing prices typically move in long cycles, due to long-lasting excess demand or supply, and hence are partly predictable. Volatility in housing prices is caused by the fact that the supply of houses does not react rapidly to changes in housing demand. Changes in real income growth and housing lending often affect short-run movements in housing prices. However, the growth rate of real income should affect only rental income in the long run.

Because housing prices tend in the long run to rise at the rate of overall inflation, housing is a good inflation hedge. However, in the long run, the real return on housing is equal to the explicit or (for owner-occupied housing) implicit rental income from housing ownership. The estimation results also show that changes in the

general price level are transmitted into housing prices fairly quickly but that inflation is surprisingly insensitive to housing prices, even though housing expenditure gets a weight of almost a quarter in the CPI.

8 Forecasting applications

Traditionally time-series analysis and classic time-series decompositions have focused on forecasting the time-paths of variables. In principle the technique has been to extrapolate predictable components of univariate models (eg trend and seasonality) into the future. In a multivariate context, dynamic economic models can be used for forecasting variables and testing dynamic economic hypotheses, and one can investigate the causal relationships in these models. When forecasts are produced with OLS regression models, we need to know or extrapolate future values of the exogenous variables in order to forecast the time paths of the endogenous variables. In forecasting complex causality relationships among exogenous variables can give a rise to numerous forecasted paths for endogenous variables.¹⁵

In the time series context, stability conditions are related to stationarity conditions. Since economic models aim to uncover dynamic paths of time-series, those components that can be extrapolated are summed up to make a forecast. The trend is the integrated 'permanent' part of the series, which changes the mean of the series. The seasonal components exhibit regular annual variation, and the irregular component is the stationary noise, which exhibits a tendency to revert to zero.

With structural models, forecasts are based on lagged and contemporaneous relationships, and on conditional expectations of these predetermined variables. Economic forecasting is one of the main reasons for building economic models, though simple univariate models can also be specified for forecasting purposes. In addition to dynamic economic relations, these models use the statistical properties of the time-series processes, which often leads to fairly similar forecasting performance. A requirement of a successful forecasting

¹⁵ An implicit assumption in using regression methods is that the structure of the model remains unchanged. If the model is unstable poor forecasting performance is to be expected.

model is that the explanatory variables be weakly exogenous and that the dependent variables not Granger-cause the independent variables.

The study of bankruptcies by Takala and Virén (1996) sets out the ingredients of an equilibrium model for bankruptcies and uses cointegrating system models to explain and forecast bankruptcies. Various macroeconomic variables are found to explain the observed number of bankruptcies, and it is also found that the number of bankruptcies is also an important indicator of current economic activity. In addition, we consider and test bankruptcies as a means of eliminating excess capacity from the commodity market by treating bankruptcies as an equilibrium error between total demand and supply. Even when the economy is at the peak of a boom, some bankruptcies occur because of changes in tastes, misplaced investment or erroneous corporate strategy. However, during a boom the number of bankruptcies is on average lower than during recessions. If total demand and supply are cointegrated, bankruptcies could be regarded as approximating the cointegration error between these variables, and bankruptcies could be used as a Granger causal predictor of changes in demand.

The empirical part of the paper attempts to explain Finnish bankruptcies. The impact of macroeconomic variables on the number of bankruptcies has been studied earlier in eg Altman (1983), Hudson and Cuthbertson (1993) and Laitinen (1990). Bankruptcies have been found to depend on economic activity, interest rates and overall indebtedness ratios. This paper shows that bankruptcies are strongly related to the business cycle and even more strongly to indebtedness, real interest rates and asset prices. The importance of these financial variables increased when the financial markets were liberalised in the 1980s. Although there is much seasonal and cyclical variation in bankruptcies, its long-run ratio to the number of firms is relatively stable and represents a kind of 'equilibrium bankruptcy ratio'. When the number of bankruptcies is high, it is likely that correcting forces for reducing bankruptcies will be strong. The number of bankruptcies is not likely to be explosive. Even, more likely the ratio of bankruptcies to number of firms will at least be stationary.

One might interpret the cointegration relation between production, real bank lending and bankruptcies as being a reflection of the fact that outside forces or fundamentals drive the stochastic common trends in these variables. There are also other stationary shocks eg in interest rates, change in terms of trade, inflation etc that divert the equilibrium paths of these variables from what is expected. From the

viewpoint of efficient forecasting, these stationary variables could be added to an error-correction specification to improve predictions.¹⁶

The last paper, Takala and Virén (1997), takes a closer look at nonlinearities in economic time series and tests for the chaos and stochastic properties of real and financial economic time series. Chaos analysis deals with the analysis of deterministic time series, ie series that evolve according to given rules, without any stochastic error term. An important property of chaotic behaviour is that the variable depends heavily on the initial state of the process. Small differences in the initial state can lead to large differences in the outcome, even after just a few iterative rounds of estimation. Economic systems are dynamic, but it is not likely that they are deterministic (and hence sensitive to initial conditions) in any meaningful sense. This says that economic systems are flexible to changing environment rather than deterministic.

Estimation of a specific nonlinear model could be based on information on the functional form. The flexibility of many nonlinear models can lead to overfitting within the sample period. In such cases, a priori analysis of the nature of the phenomenon is important. Selection of model type is much more difficult among nonlinear models. Nonlinearities may arise eg when there is large idle production capacity that could be put into operation, and a slump could arise quickly as a consequence of losses in foreign trade. It could be argued that government or central banks might react in a nonlinear manner to certain economic indicators such as currency reserves, current account or inflation. Thus it is likely that policy instruments are used in a discrete way that makes estimates of ordinary policy reaction regression equations also biased. Nonlinearities could be associated with particular features of time series, eg a financial crisis, which induce asymmetric adjustments.

One way to reveal nonlinearities in a data set is to compare residual variances of post-sample forecasts to those of linear forecasts. Granger and Teräsvirta (1993) emphasise the importance of first testing for linearity vs nonlinearity before trying to build a nonlinear model. In a traditional regression model, this could be done via

¹⁶ In different economic applications, the speed of adjustment to long-run equilibrium may differ considerably. If adjustment is slow – due eg to regulation or adjustment and transaction costs – the cointegration relation may be blurred by stochastic shocks.

Ramsey's Reset tests. The stability of parameters could give important hints of nonlinearity in OLS estimations.¹⁷

Economic time series seem to behave randomly and are difficult to predict. Moreover, there is little evidence of deterministic behaviour in economic time series other than identities. Testing for the presence of deterministic generation of time series such as returns on shares could reveal predictable temporal structures. On the whole, there is no strong evidence in favour of the presence of deterministic laws or dimensional chaotic generators in economics. One characteristic of a nonlinear model is sensitivity to the initial values of the process, which reflects to some extent the permanency of shocks that affect the process.

The aim of the paper was to determine whether indeed there are signs of nonlinearities in these series. Thus, we carried out a set of tests like those in Lee, White and Granger (1993). At this stage, most of these tests were applied to univariate models, although a multivariate application would obviously be more interesting. In analysing the series, attention was also paid to the distinction between nominal and real series. This can be motivated by the fact that nonlinearities are presumably quite different with nominal and real variables.¹⁸ Thus, it is interesting in this respect to compare a typical real series, eg industrial production, to a nominal series such as share prices.

This paper contains a set of tests for nonlinearities in economic time series.¹⁹ Some of these tests correspond to standard diagnostic tests for revealing nonlinearities, but there are also few more recent developments in modelling nonlinearities. The latter test procedures make use of chaos theory (long-memory) models and some asymmetric adjustment models. The test results unambiguously support the existence of strong nonlinearities in the data, but the evidence for chaos is weak. Nonlinearities are detected not only in a univariate setting but also in some preliminary investigations dealing with a multivariate case. Also some differences between short and long-term behaviour can be detected.

¹⁷ Testing nonlinearity includes testing temporal dependence of the residuals of the estimated model. In a sense these tests resemble portmanteau tests as a form of lack of fit tests. Under broad conditions, the first order asymptotic distribution is the same as for the true residuals and converges at speed $T^{1/2}$.

¹⁸ See also Mullineux and Peng (1993).

¹⁹ Empirical tests were carried out with very long time-series (almost 900 observations) covering Finnish monthly data for 10 macroeconomic time series for the period 1920–1996.

One common explanation for nonlinearities is chaotic behaviour. The behaviour of financial variables in particular has been analysed from this viewpoint (see eg DeGrauwe et al, 1993; Greedy and Martin, 1994; Peters, 1993 and Vaga, 1994). These analyses focused on testing for the existence of chaos. Theoretical analyses have been presented mainly as examples of various possibilities in which chaos could arise. Economic time series include much noise, and stochastic features that cannot be modelled by any simple deterministic rule.

Most monetary series – eg relative prices, changes in price level and money aggregates – show some form of nonlinear behaviour. Prices are often more volatile than the real series, since they serve as a market clearing device. Monetary phenomena are sometimes based on valuations that could be adjusted without any relevant cost. In the market clearing process, it is often (but not always) easier to change the price than the quantity. Although prices could easily move in either direction, crises in the market generate excessively large negative (or positive) changes.²⁰ Nominal price rigidities would also have similar effects. Thus it may be no surprise that the real exchange rate, share prices and inflation seem to adjust asymmetrically to shocks.

However, in many cases real economic variables also vary in a nonlinear way. Obvious evidence of nonlinear adjustment can be seen eg in the apparent persistent tendency toward cyclical behaviour of most important production variables.²¹ Whether the nonlinearities in these real series arise from the series' generating process itself or from random shocks is largely an empirical question. So far no agreement has emerged as to whether real or monetary phenomena are responsible for business cycles. We hope that our estimates concerning the nonlinearity of these series will shed some light on this issue as well.

This collection of papers entails economic applications using time series techniques to analyse economic relationships. Regarding the use of time series analysis, asset market data is often used as a forerunner to new techniques. The field of economic applications of advanced time series techniques has recently included integrated analysis of unit roots and stationary short-run adjustment. Most of the papers in this

²⁰ This affects the volatility of these series. Another major observation as to the origin of 'price shocks' relates to their unstable variance over time. It has been verified that in many cases price changes – eg in the stock market – cluster significantly. Forecasting price changes is therefore a harder task for economic agents than is forecasting smoother real variables.

²¹ See Pfann and Palm (1993) for details.

collection deal with these topics and are connected closely with forecasting. And once one is able to explain an economic variable, it becomes also easier to forecast.

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The consumption function revisited: An error-correction model for Finnish consumption

Kari Takala

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Abstract

The permanent income hypothesis asserts that consumption depends on current labour income, expected present value of earnings and cumulated net wealth. Based on this idea, a three variable cointegration system, including consumption, income and net wealth, is tested and approved. Net wealth is included into the system in market values to allow capital gains in real estate wealth, which means that wealth is not simply accumulated savings. Wealth is also disaggregated into financial assets, real estate wealth and debt, in order to confirm the existence of one cointegration relationship and to test the proper wealth concept for the consumption function. The tests suggest that a broad concept of net wealth is preferred for the consumption function. It is also found that the average propensities to consume are different for financial wealth and real estate wealth.

The results from the cointegration system show that neither disposable income nor net wealth can be regarded as weakly exogenous in the system nor can either one be excluded from the system of endogenous variables. So, in fact, there does not exist a separate statistically meaningful consumption function. Despite this, effort was made to explain which variables in the past had the greatest effect on consumption, and so-called error-correction consumption function was estimated. In addition to the $I(1)$ core of the consumption model including income, net wealth, few weakly exogenous stationary variables; the real interest rate, inflation, unemployment rate and relative prices of consumption subgroups were found to be significant.

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1 Introduction

The relationship between consumption and income has long been one of the most important issues in macroeconomic model building and forecasting. Although much of this attention is due to the large share of private consumption in GDP, the relationship has been the subject also of profound theoretical consideration since Hall's (1978) Euler equation formulation using rational expectations. It is quite clear that the simplest formulations of the permanent income hypothesis (PIH) with exogenous income, constant interest rates and additive utility functions, cannot explain all the stylized facts about consumption. Yet the usual starting point, with consumption smoothing as a consequence of optimization over some time horizon, has not been subjected to much scrutiny. However, it is clear that all the stringent restrictions that have been placed on the consumption function cannot be adequate or harmless. But when it comes to the relaxing assumptions to offer a better theory, opinions differ. Some writers think that **consumers cannot strictly speaking form their expectations rationally** or that subjective expectations are poorly specified. This has led to the conclusion that aggregate expectations may be rational, but there are aggregation problems in the representative agent consumer theory. In this respect individuals may have myopic expectations because of uncertainty as to the future earnings (Flavin, 1981, Hall & Mishkin 1982, Muellbauer & Murphy, 1993). Another relaxation of restrictions has emerged through the error-correction model, where rationality is maintained only in the long-run, and agents are allowed to make mistakes in the short run (Hendry 1993, p. 179).

Few others think that **the representative consumer** approach has reached its limits in describing the variety of phenomena that determine consumption and savings (eg. Attanasio & Browning 1993). Savings is motivated by numerous factors that cannot be presented within a single framework. Consumers can therefore be classified into different groups with respect to age, socio-economical status and prosperity, which affect consumption eg. through liquidity constraints (Campbell & Mankiw 1990, Sefton & Veld, 1994).

Quite recently many of the stylized facts about consumption and the role of the life-cycle model in explaining them have been challenged in the light of international time series and cross-section evidence (Carroll & Summers 1989). They emphasize the role of the perfect capital market assumption as a major obstacle to household consumption smoothing over lifetime. **Capital markets are not efficient enough** to discount the value of future earnings, social security etc. (see also King 1986). For example, borrowing against pensions is usually limited, partly because of the uncertainty of the time of death. Constraints on capital markets have also been proposed to be responsible for the excess sensitivity of consumption to current income. However, in many ways this explanation too is not sufficient. For instance consumption exhibits excess smoothness with respect to permanent income, whereas consumption is too sensitive to current disposable income. In fact, this tells that the optimization horizon may be much shorter than remaining lifetime. In the PIH consumption is proportional to physical non-human wealth and expected value of earnings. As consumers are trying to ensure themselves against variation in earnings (eg. due to unemployment), they could retain stable (liquid) asset income ratios.

This idea has been formulated in the buffer-stock view to savings (Carroll 1992).

This paper starts by reviewing shortly the basic LC-PIH theory and empirical problems concerning the determinants of private consumption and the problem of predicting it (chapter 2). This paper concentrates on the role of wealth in the consumption function. The starting point is a three variable cointegration relationship between consumption, real disposable income and broadly-defined net wealth, based on the LC-PIH formulation of the consumption function. Recent applications of this framework include Brodin & Nymoen (1992) for Norway, Berg & Bergström (1993) for Sweden and Patterson (1994), Hendry (1994) for United Kingdom.

Empirically, the choice between consumption and saving may also have been subject to change, since in Finland savings have only recently carried a strongly positive real return. During a long period of liquidity constraints and restricted capital movements, the major share of household saving was closely tied to the requirements for obtaining housing loans and was under-compensated due to the regulation of deposit rates. During the past 5–7 years the situation has changed radically. The wealth effects associated with financial liberalization and thereafter the consumption function have also raised questions about their role in the large forecasting errors of the late 1980s (chapter 3). Financial deregulation produced first large potential capital gains to owners of housing wealth. The house price bubble bursted in early 1990s and since household indebtedness had almost doubled, the collapse in consumption emerged.

However, the main purpose of this paper is to provide a better specification of the consumption function for forecasting and conditioning for dynamic simulations. As an introduction we review the unit root properties of consumption, income and wealth (chapter 4). This is essential for the specification of the error-correction consumption function. Forecasting consumption requires also strong exogeneity of the explanatory variables, which entails Granger non-causality and weak exogeneity. Therefore we look at, whether a consumption function exists in Finland. This task includes testing **what is the precise form of the cointegration system of consumption** and what variables could be used as weakly exogenous variables (chapter 5). This paper emphasizes the role of capital market restrictions by disaggregating net wealth into financial wealth, real estate wealth and debt. The liquidity of these assets is found to be different in financing consumption. Although the emphasis is on the empirical model, the model should be based on a theoretically adequate description of the stylized facts of consumption in Finland, so that it can produce reliable long-run forecasts.

Attention is also paid to the problem of finding structural breaks and outliers during the period of **financial market deregulation** in the 1980s and to the effects of the easing of liquidity constraints (chapter 6). For the present period, the credit crunch might have to be taken into account, as it might have reversed or offset the effects of deregulation with respect to liquidity constraints.

2 The life-cycle/permanent income hypothesis

According to the life-cycle/permanent income hypothesis (LC-PIH) consumption depends on current labour income, net wealth and the present value of expected earnings. If the capital market is perfect, consumers can smooth their consumption by lending or borrowing at the same interest rate even with respect to their expected labour income. In practice it is quite clear that capital markets are not perfect in two respects. First there is an interest margin between lending and borrowing rates and secondly, more or less binding liquidity constraints weaken the ability to borrow against uncertain future earnings. Information and transaction costs reflect an important part of this uncertainty. The close relationship between aggregate disposable income and consumption indicating liquidity constraints has been seen one of the most important empirical arguments against the LC-PIH (Deaton 1992).

Here, we present only the canonical model of the permanent income hypothesis (PIH). In this model the following restrictions are made

- i) the (real) interest rate is constant
- ii) tastes do not change intertemporally (constant time preference)
- iii) there is no transitory consumption
- iv) the utility function is additively separable over time and between consumption, leisure and other goods
- v) households live infinitely long

With these assumptions we can write

$$c_t = (r/1+r)[A_t + H_t] = (r/1+r)[A_t + \sum_{i=0}^{\infty} (1+r)^{-i} E_t y_{t+i}],$$

where

- r = real interest rate
- A_t = nonhuman wealth at the end of the period t
- H_t = human wealth
- y_t = labour income at time t

Denoting $R = 1/1+r$, we can write the present value of the future earnings at period t as

$$H_t = \sum_{i=0}^{\infty} r^i E_t y_{t+i}$$

and capital income becomes $[r/(1+r)] A_t = rR A_t$. If we try to assess, what assumptions are not realistic or harmless, we may conclude that from the viewpoint of short-run consumption function, the constant real interest rate and weak exogeneity of disposable income could be suspected. The existence of noise in consumption, constant time preference or length of the economic planning horizon may not introduce that large biases.

Since the future path of income is uncertain, consumption plans will be revised as new information about future labour income becomes available, ie.

$$\begin{aligned}\Delta c_t &= \frac{r}{(1+r)} \left[\Delta E_t y_t + \sum_{i=1}^{\infty} (1+r)^{-i} \Delta E_t y_{t+i} \right] \\ &= \frac{r}{(1+r)} \sum_{i=0}^{\infty} (1+r)^{-i} \Delta E_t y_{t+i}\end{aligned}$$

Therefore the revisions in expected earnings are reflected in the changes of consumption. If real interest rate is constant and therefore no capital gains exist, nonhuman wealth affects only the level of consumption (Flavin, 1993). We can easily see that nonhuman wealth evolves by recursion as

$$A_{t+1} = (1+r)[A_t + y_t - c_t].$$

If we substitute the basic consumption equation into this equation, we get the equation for a change in nonhuman wealth

$$\Delta A_{t+1} = (1+r) [y_t - (1-R)H_t].$$

If real interest rate is constant, nonhuman wealth depends only on income and human wealth. If we allow unexpected capital gains in PIH, we relax the assumption of constant real interest rate. Capital gains mean also that wealth cannot be regarded as accumulated savings. With respect to the consumption function collapse in late 1980s, the exclusion of the capital gains has been surely crucial. If unexpected capital gains are defined as the present value of revision in the expected capital income, it can be shown that unanticipated capital gains will affect consumption similarly as revisions in labour income (Flavin 1981).

However, PIH is not that explicit about the definition of nonhuman wealth. The best guess would be to use some broad Hicksian definition of 'permanent' income. In PIH the flow of capital return should include capital income from financial assets and capital services received from real estate wealth. We may argue that REPIH is very restrictive in excluding capital gains, since a major part of variance in housing wealth is related to valuation changes, that correlate with interest rates and income expectations. Capital gains due to real estate wealth prices contain important forward-looking aspects. Net wealth can be accumulated through financial saving, amortization of loans or by capital gains. With finite lifetime the principal capital could be consumed also. Even if in the long run real interest rate is constant, interest rate variation affects the consumption-saving choice in the short run through intertemporal price effect.

In the simple national income and product account (NIPA) identity $s_t = y_t - c_t$ savings could be viewed as a change in net wealth, ie. $s_t = \Delta w_t$, only if capital gains are excluded.

In general the change in net nonhuman wealth can be written as the following

$$\begin{aligned}\Delta A_t &= A_t - A_{t-1} = P_t Q_t - P_{t-1} Q_{t-1} = \\ &P_t Q_t - P_t Q_{t-1} - P_{t-1} Q_{t-1} + P_t Q_{t-1} = \\ &P_t \Delta Q_t + \Delta P_t Q_{t-1} \approx \\ &s_t + \Delta G_t\end{aligned}$$

where s_t is net savings and ΔG capital gains. The first term in the equation reflects the accumulation of wealth through net savings. The second term describes the change in the value of assets due to price changes. NIPA concept of property income in principle includes the imputed rental income from owner-occupied housing but not capital gains on illiquid assets.

If there are no capital gains, the change in net wealth is equal to net savings. It may be assumed that in nominal terms financial wealth and debt are not subject to depreciation; in real terms depreciation due to expected inflation is taken into account in interest rate premiums, as the nominal interest rate consists of the real rate of interest and the expected inflation rate. Real estate wealth is subject to depletion and technical depreciation even when it is not a consequence of asset consumption. The total return on an asset comprises of capital income and capital gain due to changes in relative prices.

The life-cycle framework retains the important long-run cointegration between consumption and lifetime disposable income. It is useful to separate the long-term effects from the short-run variation. For example it can be said that the real interest rates affects consumption in the short run but that the effect will diminish in the long run if the real interest rate is weakly stationary. Therefore the **interest rate** is needed as a stationary explanatory factor in the consumption function, but it is **not essential for the I(1) core of the long-run cointegration relationship**. The real interest rate affects the intertemporal distribution of consumption, but it is not a resource that can be consumed. Long-run forecasts are therefore unaffected by stationary interest rates, whereas short-run predictions of course gain from adding them. Various factors – in fact too many to list – affect consumption in the short-run, some of these are changes in social security, tax reform, capital gains, deregulation of financial markets and changes in inflation expectations (see Berg and Bergström 1993).

In the short run disposable income could be used either for consumption or saved as net wealth. Accumulated savings consist of real estate investment and net financial (financial wealth – debt) assets, which are also available for financing consumption. Therefore, it is not surprising that wealth affects consumption, since it carries information about past savings and the ability to borrow and gain capital income. However, it is fair to say that the most difficult part to model in the life-cycle framework is the present value of human capital. Income expectations are not observable and there exists no true market that evaluates the present value of earnings.¹ Uncertainty about the value of human wealth could easily be a multiple of other potential sources of uncertainty.

¹ One may however speculate that eg. a slave market could to some extent measure the present value of human capital, although there are certain moral hazard problems involved.

So far it was assumed that in the PIH income is **exogenously determined**. If income and nonhuman wealth are not exogenous, the consumption function does not exist as such. It is very likely that there exists an endogenous feedback between real estate wealth and income expectations, since eg. housing prices correlate with expected income, being as they represent present values of housing services. In the end these are empirical matters. Muellbauer & Murphy (1993) argue that if there is anything to be gained from the rational expectations hypothesis, there should be a gain in modelling income as well as consumption. There are many stylized empirical facts that contradict the strict versions of LC-PIH consumption models. It was already refuted by Friedman (1957) himself, who argued that consumer optimization takes place over a period of three years or slightly longer rather than over a lifetime. Although nobody has truly contradicted the core idea that consumers try to smooth consumption, empirical tests have not been completely successful.² Most of the formulations lack empirically meaningful and testable implications for situations where liquidity constraints exist for large groups of consumers. One important exception is Campbell and Mankiw (1990). Uncertainty of labour income varies over time, and savings is largely based on the behaviour of a limited number of wealthy consumer-investors. These stylized facts are not satisfactorily incorporated into life-cycle models. In some studies uncertainty in earnings has been taken into account by including unemployment rate or the change in unemployment rate into the consumption function (eg. Carroll 1992). The topic in this paper is to show how inclusion of net wealth into consumption function will improve the results. In particular we present that disaggregating net wealth will reveal few important features in the PIH framework.

Asset price changes could lead to capital gains and the amount of financial saving is also more sensitive to changes in real interest rates among the "true savers". Therefore the **distribution of an asset portfolio** will matter for the **choice between saving and consumption**. These effects will surely be present in the short run, but may vanish in long run considerations. Anyway these propositions tell that changes in saving behaviour cannot be understood without taking into account changes in the wealth portfolio. Since wealth is owned in unequal shares among the wealthy and ordinary savers, there is a need for different models for ordinary liquidity-constrained consumers and rich savers.

The riskiness of earnings is related to macroeconomic income risks eg. through changes in unemployment. People often save during their pension years because of probable health expenses. However, this savings motive is related rather to attempts to smooth consumption than to prepare for income risk.³ Skinner (1988, p. 248) emphasizes the effect of earnings uncertainty for the precautionary motive of savings. Skinner also remarks that the closer earnings are to a random walk the more important is precautionary savings.

Empirical studies from as far back as the 1950s have found significant differences in the savings rates of different occupation groups. In Finland the savings and accumulation of wealth has differed profoundly between different age and socio-economical groups (eg. Vilmunen & Viren 1991). There is direct

² For Finnish empirical evidence, see Takala (1995).

³ According to Skinner (1988) precautionary savings could account about 56 percent of aggregate life-cycle savings.

evidence that the riskiness of earnings affects the willingness to save. This emphasizes the fact that aggregate savings is also related to overall changes in income risk.⁴ The discussion above shows that there exists various stylized facts in actual behaviour, which are difficult to incorporate into PIH. Even though basic PIH captures the important long-run features, it requires some tuning concerning the short-run behaviour. As a summary it could be said that different assets have somewhat different role with respect to consumption. Real estate wealth is held mainly for service flow. Only if income expectations are favourable they could be used as collateral for consumer credits. Liquid assets are held for transaction services and as buffer-stock for consumption smoothing. Long-term financial wealth like time deposits and bonds are invested to produce capital income. Debt is used to acquire housing wealth or durables.

3 The data and recent history

The data used is largely based on the seasonally adjusted series constructed for the Bank of Finland BOF4 model. In addition, we use more recent measures of net wealth in our calculations. From the theoretical point of view, using seasonally adjusted data is unfortunate, since causal and dynamic relations could be seriously affected.⁵

Other limitations are present as well. In Finnish national accounting, consumer durables are included in total consumption based on purchases. This means that the depletion of these investment commodities is exaggerated and the services flow from durables is underestimated. Therefore also the share of durables in total consumption is underestimated. We expect that durables may be more sensitive to changes in user costs and therefore real interest rates than non-durable consumption. This is another reason why durables were separated from the consumption of non-durables and the final **consumption function was estimated with non-durables**.⁶ It could be argued that it is not proper to

⁴ Against plausible theoretical considerations Skinner (1988) finds that precautionary savings is less important among those occupational groups that face the largest variance in earnings, eg. among the self-employed and salesmen. With Finnish cross-section questionnaire data it was found that surprisingly those who did not expect unemployment increased saving most due to growth in unemployment. Most households indicated an increased motive for saving due to income uncertainty, but only those that did not expect unemployment were able to save (Takala 1995).

⁵ In addition, in testing a consumption function hypothesis, the data should be in per capita form to avoid unnecessary complications due to demographic effects caused by eg. a growing population. However, since we are interested in providing a useful macroeconomic forecasting equation, the data was not deflated for population.

⁶ Patterson (1985) emphasizes that a proper treatment of durables requires changes to definition of income. If durables are added into consumption, their depreciation have to be taken into account as well. An alternative way to handle durables is to define income net of depreciation (and value losses) of the stock of durables. Consumer behaviour may be based on some Hicksian type of 'permanent income', which keeps the net wealth constant, but durables purchases are paid from saving left over from non-durable consumption out of disposable income or financial portfolio reallocation.

exclude durables only, since income has to be adjusted as well eg. by reducing the expenditure on durables from income. The LC-PIH applies to the consumption of non-durables and the service flow from durables. Unfortunately no official measure of the service flow from durables exists. It may be worthwhile to try also how a moving average of durable purchases added to non-durables will alter the results.

Although the time series for total private consumption and for consumption excluding durables do not differ greatly, this is merely because of the relatively small volume of durables than to a similarity of behaviour (figure 1). The consumption of services and non-durable goods are however closely related (figure 2). Another likely reason for the distinct behaviour of durables is that liquidity constraints reduce durables spending disproportionately when income declines (Carroll & Summers 1989).

The very first variable that has been used to explain consumption is, of course, **disposable income**, which represents readily spendable funds. Other resources that could be used to finance consumption are gross wealth and debt. However, one might expect that real estate wealth and debt have lower spendability than eg. deposits, bonds or stocks. With slightly myopic and impatient borrowing-constrained consumers it would be expected that wealth and debt are not regarded as homogenous with respect to consumption possibilities.

The close relation between consumption and real disposable income is apparent both in levels and differences (figures 3–4). In fact, this observation has been used in various applications as an example of a cointegration relationship. Cointegration between consumption and real disposable income would imply that **the saving rate is stationary** and closely related to the error-correction term between these level variables (Engle and Granger 1987, Campbell 1987).

In early consumption function specifications, it was already noted that net wealth could be used as additional variable. Instead of just using real disposable income, one could construct a proxy for permanent income by using disposable income and real net wealth. This approach has been used eg. by Brodin and Nymoén (1989, 1992), although it did not prove to be totally successful in Norway.

The only problem with this type of approach is that there are no proper official statistics for sectoral or household wealth. Sectoral debt classifications also have deficiencies. In this study we compare two separate net wealth concepts and their appropriatedness for the task at hand. A more proper and broader concept of household net wealth on a quarterly basis is available only from 1979 onwards. This data is constructed from various sources of information and could be regarded as the best available disaggregated market valued asset portfolio data.⁷ However, a relatively good narrow approximation

⁷ The constructed financial wealth includes cash, bonds and different types of deposits (tax-free, taxable, withhold-taxed deposits, time and currency deposits). Only relatively minor assets like life insurance and pension saving funds are excluded at the moment. Real estate wealth includes agricultural wealth (fields, real estates and equipment and cattle), forest estates, durables (cars, boats etc.), entrepreneurial wealth, summer cottages, housing wealth and stocks. Only stocks in firms outside the stock market are excluded. The debt measures, disaggregated into housing loans, consumer credit and other debt, were taken from the Statistics Finland accounts. The involvement of disaggregated net wealth assets is shown in the Appendix.

could be constructed based on data from the Bank of Finland (BOF4) model including cash, deposits, housing wealth and bank lending and covering 1960 onwards (figure 5).

Already at this point it is necessary to discuss the particular difficulties concerning the sample period, especially the late 1980s'. If we compare savings as measured in the national accounts (disposable income – consumption expenditure) and the change in household financial wealth, we see a strong negative correlation from about 1985 onwards. Financial wealth (cash, deposits, bonds) does not include any asset that is subjected to capital gains except through inflation, since stocks were included in real wealth. Expected inflation is included as a premium in nominal interest rates. Even though the relationship between savings and changes in financial wealth would be expected to be positive in normal conditions, the availability of foreign lending accelerated the growth of indebtedness. In fact household indebtedness almost doubled between 1985–1991 (see Brunila & Takala 1993). Savings declined as households turned to investment in housing (house purchases) and loan amortizations rather than financial deposits have increased (figure 6).

The peak in financial asset accumulation at the end of 1988 is related to foreign lending used in real wealth and private business and corporate purchases, which returned back to the banks as deposits. Business sales were introduced by the expected new tax on capital gains as from the beginning of 1989. In a closed economy with capital movements, such a change in financial assets would not have been possible, since rising interest rates would have balanced such an investment and borrowing boom. Foreign lending was fuelled also by the banks' competition for market share and the restructuring of bank ownership.

The opening up of the financing sector to foreign borrowing can also be seen from the difference between the national accounts savings rate and the Hicksian wide savings rate measure, which measures the change in potential consumption resources. However, the main consequence for savers has been the positive return on financial saving after 1990. Financial deregulation has also affected household spending and increased debt service costs, at least temporarily. Households' gross debt servicing peaked at almost 30 percent of disposable income. When the overspending had ended, the saving rate recovered quickly. Increasing real returns on financial assets followed because of rising international interest rates and the onset of the withholding tax on capital income in 1991. This, together with and uncertainty about future earnings due to high unemployment, can be regarded as the main reasons for the increased saving rate.

4 Univariate properties of consumption and income

Prior to cointegration analysis, the relevant series are typically pretested for order of integration. A rough rule for a feasible regression is that the integration orders of both the left and right sides should be the same. There is no way in which a nonstationary variable can be explained successfully with nothing but stationary variables. The same applies in most cases in the opposite direction, unless some of the right-hand side variables are cointegrated. For example if

one cannot reject the hypothesis that logs of consumption and income are cointegrated, it is not possible to reject the hypothesis that the saving rate is stationary.

The random walk behaviour of non-durable consumption is quite apparent. This could be confirmed already from the autocorrelation function of the residuals of differences of the basic variables. Serious autocorrelation is present in real disposable income and particularly in real net wealth, but the first **differences of log consumption are relatively free of autocorrelation** (table 1). Table 1 presents unit root tests for the basic logarithmic series. Although the evidence for unit root is not always unambiguous, the presence of unit root for the levels of series is not rejected in most cases.

The permanent income hypothesis is based on the idea that consumption depends only on permanent income. Since permanent income is unobservable, it has been proxied by different functions of current and lagged earnings that measure expected labour income plus some function of wealth indicating expected capital income. In many applications a proxy for wealth has been used as part of permanent income, since it is always possible to exhaust the capital itself in order to finance consumption. This was also the main argument in the early formulations of the life-cycle hypothesis. Several empirical studies have subsequently taken into account the role of imperfect capital markets and emphasized the distinction between the spendability of different types of wealth. Cash and other forms of financial wealth are readily available for consumption, but there are limits in the spendability of real estate wealth and debt at least in the short run.

From the viewpoint of cointegration, it is reasonable to assume that the log of consumption and log of real disposable income are integrated of order one. If these variables are also cointegrated, this means that the savings rate is stationary.⁸ However, if we consider real savings to be integrated of order one, it could happen that the cumulative of it, namely real net wealth, could be even I(2) (Hendry 1993 p. 211). So far our unit root tests have shown no indication of real net wealth to be anything more than I(1).⁹ According to performed test saving rate was weakly stationary ie. integrated of order zero. Inflation and real interest rate should be weakly exogenous variables in the consumption equation. The evidence on their stationarity is however not that convincing for this relatively short period. In a steady-state solution consumption and income would have the same long-run growth rates, but there is no obvious restriction to the growth of real net wealth in logs. If the saving rate is stationary white noise, the cumulant of savings, ie. net wealth, would be only I(1). Therefore it seems that in the long run we may not be able to rule out the possibility that log real net wealth could not be I(1). To account for this possibility, we have tried different functions of wealth in the consumption function. Disaggregated net wealth, real return of net wealth and the ratio of net wealth to disposable income were compared as potential proxies for the wealth concept.

⁸ From the decomposition of disposable income, we have $Y = C + S \Leftrightarrow C/Y + S/Y = Y/Y \Leftrightarrow S/Y = 1 - C/Y$, which implies the following; $\log(S/Y) = \log(1) - \log(C/Y) = -\log(C/Y)$.

⁹ Consumption functions are formulated in real terms. By deflating consumption expenditure and other variables with prices we assume zero-degree homogeneity in prices.

Figure 1.

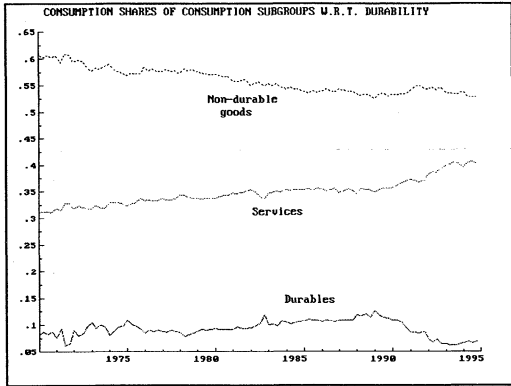


Figure 2.

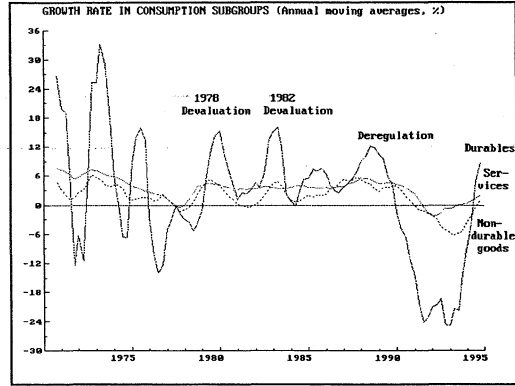


Figure 3.

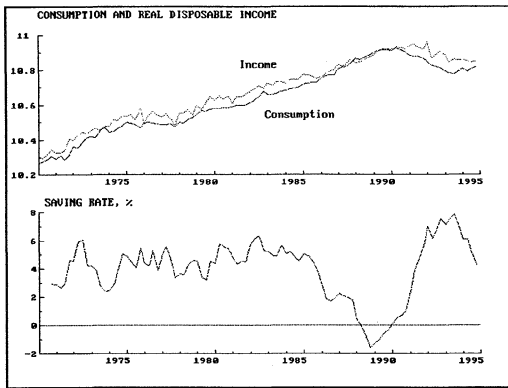


Figure 4.

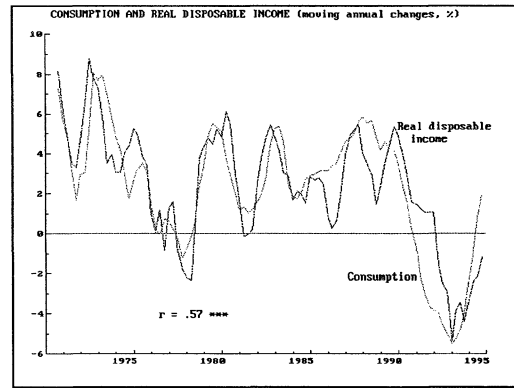


Figure 5.

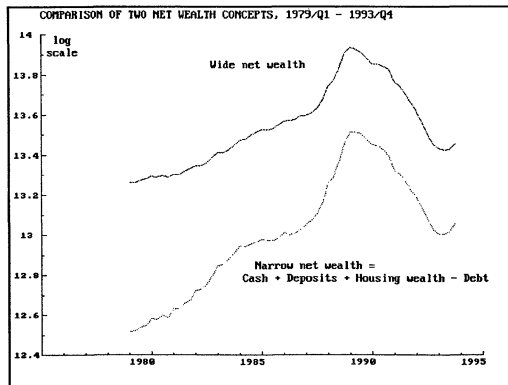


Figure 6.

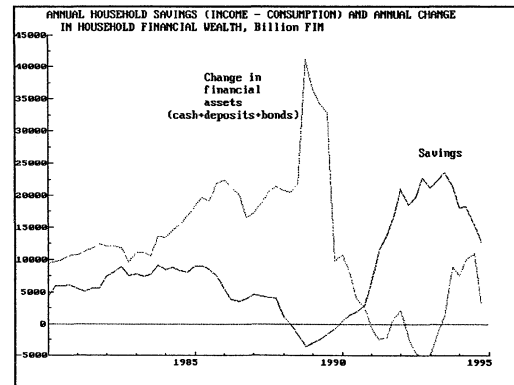


Table 1. **Testing for the order of integration, Augmented Dickey-Fuller unit root and time trend tests**
 Non-durable consumption, real disposable income
 and real net wealth, 1970/Q1–1993/Q4

Autocorrelations from differences, lags 1–6,

Lag	1	2	3	4	5	6
LC:	.101	.065	.246 *	.130	.206	.100
LNONCD:	.191	.054	.216 *	.296 **	.203	.099
LRYD:	-.443 **	.065	.070	.137	-.259 **	.132
LRNW:	.626 **	.548 **	.452 **	.482 **	.227 *	.186

Critical value at 5 % significance level: $\pm 2/\text{SQR}(84) = \pm 0.213$
 1 % significance level: $\pm 2.3/\text{SQR}(84) = \pm 0.252$

DICKEY-FULLER AND AUGMENTED DICKEY-FULLER TESTS WITH LAGS 1 AND 4
 (McKinnon's 95 % critical values)

Variable	DF-statistic		ADF(1)-statistic		ADF(4)-statistic	
	Without trend	With trend	Without trend	With trend	Without trend	With trend
LNONCD	-2.422	.069	-2.298	-.508	-1.840	-1.931
LRYD	-1.709	-3.164	-2.173	-1.547	-2.057	-1.594
LRNW	-1.678	.335	-1.316	-1.699	-1.748	-4.212
DLNONCD	-7.933	-8.407	-5.894	-6.326	-2.144	-2.540
DLRYD	-15.404	-15.741	-9.504	-9.999	-3.764	-4.336
DLRNW	-4.470	-4.495	-3.039	-3.032	-2.953	-3.017
SAVRATE	-7.617	-7.590	-4.686	-4.650	-2.991	-2.949
D4LCPI	-1.071	-2.608	-1.644	-2.805	-1.473	-2.445
RRLBN	-0.848	-2.738	-1.217	-2.833	-1.264	-2.729
95 % Crit. values	-2.892	-3.457	-2.892	-3.458	-2.893	-3.459

Variables:

- LC = Log of consumption
 LNONCD = Log of non-durable consumption
 LRYD = Log of real disposable income
 LRNW = Log of real net wealth
- DLNONCD = Log difference of non-durable private consumption
 DLRYD = Log difference of real disposable income
 DLRNW = Log difference of real net wealth
 SAVRATE = Saving rate
 D4LCPI = Annual log difference of consumer price index, %
 RRLBN = Real (deflated by CPI) lending rate for new loans, %

5 Testing for cointegration hypothesis

We use Johansen's (1988) VAR cointegration procedure in testing linear relationships between consumption, income and net wealth, since it has several advantages over the original two-step Engle-Granger (1987) method. First, VAR modelling allows us to look for the number of cointegration vectors, whereas the Engle-Granger procedure for testing cointegration applies only to the case of one cointegration vector. Secondly, in a VAR system it is easier to separate long-run cointegration relations from short-run dynamic responses. Statistically, the VAR method combined with ML estimation is also more efficient. In the Engle-Granger procedure, inference about the cointegration vector depends upon nuisance parameters and is sensitive to finite-sample bias. Therefore two-step procedures lack power in comparison to the modified VAR approach. Analysing the full VAR system of equations eliminates the possible single equation bias that would likely disturb the two-stage results. Therefore also the Engle-Granger two-stage regression method and multivariate VAR method will produce different results. Johansen's reduced rank regression framework is also more suitable for system tests concerning the full conditioned model, such as the weak exogeneity and exclusion tests.

The main advantage of the ECM representation is the explicit separation of the long-run relationship between the modelled variables and the short-run variation around equilibrium paths. The economically most meaningful part is included in the long-run equilibrium relations, which usually reflect some sort of optimizing behaviour. The optimization is based on the long-run relationship, but also the short-run dynamics must be modelled, if a proper description of the process is to be obtained.

Another advantage of the Johansen method is related to the rapid convergence of the OLS-parameter estimates, which enables more reliable forecasts based on cointegrating vectors. With economic time series, sample sizes are often limited and superconsistent parameter estimates are most welcome. Despite the small sample bias present in parameter estimates, long-run asymptotic relationships dominate the sources of bias. The VAR procedure also takes into account the short-run dynamics of the endogenous variables while estimating the cointegration vector. This is the main reason for more reliable estimation results (Muscatelli & Hurn 1992). The VAR procedure allows testing of various other hypothesis concerning the cointegration vector β , since the coefficients include the long-run equilibrium conditions.¹⁰

If cointegration exist among the system variables, the consumption function would need to be modelled in an error-correction form. Interpretation of the cointegration relation between consumption and real permanent income seems to be that outside forces or fundamentals drive the stochastic common trends in both variables. Cointegration is usually thought to be a long-run relationship between levels, but it can also emerge eg. at seasonal frequencies. There are also other short-term shocks from real interest rates, taxation and the stock market, which

¹⁰ The Johansen procedure was performed to test cointegration with the VAR -system estimation available on Microfit 3.0 and a RATS subprogram CATS written by Johansen and Juselius (1991) and improved by Hansen (1993). Regression models were spesified and estimated with PcGive 8.0 by Doornik and Hendry (1994).

divert consumption and income from their planned time paths. From the viewpoint of efficient forecasting, these stationary variables could be added to the ECM specification to improve predictions.

The starting point in cointegration analysis is the static long-run equation. The first question is whether we should include net wealth in the cointegration specification. Estimation problems emerge already at this stage. We should not forget that the parameters of the static equation may not be stable. The next question concerns the number of cointegration relationships.

The simplest case is the consumption-income relationship introduced in Engle and Granger (1987) as an example of cointegration dependence

$$\Delta c_t = \mu + \beta_1 \Delta y_t + \beta_2 z_t - \gamma(c-y)_{t-1} + \varepsilon_t, \quad \gamma \geq 0$$

where c is consumption, y is income and z are stationary weakly exogenous variables. Hendry and von Ungern-Sternberg (hereafter HUS, 1981) already used the difference between income and liquid wealth as another error-correction term affecting consumption

$$\Delta c_t = \mu + \beta_1 \Delta y_t + \beta_2 z_t - \gamma_1(c-y)_{t-1} - \gamma_2(w-y)_{t-1} + \varepsilon_t, \quad \gamma_1 \geq 0, \gamma_2 \leq 0.$$

where c is consumption, y is income, w is (liquid) wealth. In addition the specification may include weakly exogenous stationary variables (z) like real interest rate, inflation etc. to improve the short-run dynamics of the model.

In HUS consumers try to maintain long-run proportionality between both consumption and income, and in addition between wealth and income. It should be noted that ECM terms have different signs, since accumulation of saving will be reflected in the wealth-income ratio. Even though HUS and Hendry (1994) uses liquid assets instead of total net wealth, it could be argued that the stock-flow integral correction mechanism (ICM) due to disequilibria in the consumption-income ratio is more understandable if a broader concept of accumulated purchasing power is used (Patterson, 1991). Of course, we may assume that liquid assets are used more directly in smoothing consumption and that real estate wealth or time deposits regarded as investment. The HUS approach pays special attention to the dimensionality problem of net wealth being a stock variable and income and consumption being flow variables.

Savings should at least correspond closely to changes in financial wealth, but as we have seen this does not seem to be the case with empirical data. Savings and changes in net wealth have been also loosely related. Real wealth measured in market values varies with demand, which affects consumption through prices and is therefore related to net saving only in the long run.

Consumption can be financed through income, financial wealth or borrowing. The role of real estate wealth is a bit complicated, since owning eg. a house means that one gets either imputed income or can get rental income. If real estate wealth is used for non-durable consumption it must first be transformed into financial form by being sold. Limiting wealth to liquid assets is connected to spendability and accumulation of savings. Consumers usually want to consume housing in some proportion to their income, in fact many empirical studies show that housing consumption has unit elasticity. Consumers may therefore want to preserve some constant relationship between housing wealth and income. This

raises another important viewpoint on real estate wealth. It was noted already in Muellbauer and Murphy (1990, p. 364) and King (1990) that asset prices can proxy income expectations. As the market price for a house reflects the present value of housing services, it carries information about future labour income. This argument emphasizes the role of real estate wealth in the consumption function from the point of view of forward-looking consumers. Therefore the role of net wealth and real estate wealth is bit obscure in the context of consumption function. The inclusion of real estate wealth can be motivated by their role in producing services and their ability to predict future earning. Another variable that can be used in consumption functions is indebtedness as it predicts changes in income expectations as well (see Takala, 1995).

In HUS and Hendry (1994) net liquid assets were found to be significant, therefore we must ask, does the fungibility, liquidity and the form of capital yield affect the consumption behaviour in the short run and even in the long-run.

In a sense net wealth is merely cumulated savings, which could be consumed almost alike disposable income. Therefore, it is not only labour (or disposable) income that forms a proper measure for permanent income. Net wealth could be used as a buffer stock to smooth variations in disposable income. Therefore, a straightforward rival to this specification is the following formulation were only one cointegration vector and thereafter one equilibrium error emerges.

$$\Delta c_t = \mu + \beta_1 \Delta y_t + \beta_2 \Delta w_t + \beta_3 z_t - \gamma (c - y - w)_{t-1} + \varepsilon_t, \gamma \geq 0.$$

This ECM specification can be derived directly from a first-order autoregressive-distributed lag model (see Banerjee et. al. 1993, p. 48–49). The inclusion of net wealth stock into the equation as such could be interpreted as a flow variable, if capital return is proportional to the wealth stock. Choosing between these specifications is an empirical matter. The error-correction term is used in the equation to keep a record of the divergence from the long-run proportionality between consumption, income and wealth. The latter alternative, is theoretically consistent with the REPIH, but HUS is also consistent with this specification if $c_t = \beta_1 y_t + \beta_2 w_t$ satisfies homogeneity restriction $\beta_1 + \beta_2 = 1$, which implies existence of one cointegrating vector (Brodin & Nymoén, 1992). Even if consumption and income are flow variables and REPIH presents consumption smoothing in terms of yield for human wealth (earnings) and nonhuman wealth (capital income and imputed services) with infinite horizon, in practice the real capital stock could be consumed as well.

If adding net wealth to the consumption function leads to better parameter constancy, we may expect a more stable cointegration system to be found among consumption, income and net wealth. The insight in PIH is to give a theoretical basis for the long-run solution, which does not generally hold to liquid assets only. Since binding liquidity constraints were removed only recently in Finland, we can also test whether a proxy for liquid financial assets or net wealth performs best during the pre-liberalization period.¹¹

¹¹ Brodin and Nymoén (1989, 1992) found that net wealth, including housing wealth at market value, is an essential part of a household's lifetime income and life-cycle budget constraint. However, according to cointegration tests homogeneity restriction of consumption proportional to income and wealth was rejected. Unfortunately their wealth measure was somewhat inefficient eg. by excluding illiquid financial assets and few real estate assets.

6 Empirical results

6.1 Consumption and the liquidity of different assets

Traditional consumption maximizing life-cycle theory is often regarded as being indifferent to the composition of the net wealth. Liquidity or other aspects of financing do not enter explicitly into the consumption-saving decision. This is natural only if capital markets are assumed to be perfect, and real riskless after-tax returns for different assets are equal.

Berg and Bergström (1993) emphasize that the elasticities of different types of assets may not be the same in the consumption function and analyse the cointegration relationship with disaggregated net wealth. It has been proposed that since the spendability of different assets varies, liquidity constraints on borrowing and transaction costs in real estate markets will increase the elasticity of financial wealth in consumption.¹²

In practice we are faced with transaction and information costs, and as a consequence it is cheaper to finance consumption by using liquid financial assets than fixed-size (indivisible) real estate wealth. Borrowing using real estate wealth as collateral is also somewhat restricted. These considerations also reflect the basic assumption of Friedman's permanent income hypothesis. Therefore the liquidity composition of the asset portfolio will affect optimal consumption. The proper measure of consumption used in the consumption function will also depend on the structure of the asset markets. The optimal individual portfolio – regarding net wealth as accumulated savings – will depend on the interest margin between borrowing and lending, the collateral ratio of real estate holdings etc. As a consequence of financial deregulation, it has been argued that illiquid assets may have become more spendable during the 1980s (eg. Muellbauer & Murphy 1993).

The effect of **different asset structures** on the consumption-wealth ratio is studied in Pissarides (1978). Pissarides formulated a special form of transaction costs, where illiquid assets cannot be sold for full value until held for a certain number of periods. On the other hand, no transaction costs are assumed in the case of consumer goods. This formulation has a few advantages. First, the timing of income payments will affect the timing of consumption. This may explain to some extent the higher correlation observed between current income and consumption as compared to what is predicted by the permanent income hypothesis. This also means that restrictions in the asset market impose restrictions on consumption too (Pissarides 1978, p. 292).¹³ Secondly this formulation is in accordance with the buffer stock theory of savings as it explains why the short-run marginal propensity to consume (MPC) is smaller than MPC out of permanent income (Carroll 1992).

¹² In fact Berg and Bergström (1993) present evidence that net financial wealth and debt were not significant predictors of consumption in Sweden prior to mid-1980, probably because of financial regulation.

¹³ With few assumptions, Pissarides formulates a model in which consumption is a linear-homogenous function of lifetime wealth. Liquidity differences in assets affect consumption through future discount factors, which depend on transaction costs, maturity of illiquid assets and, of course, rates of return.

In fact the effect of the portfolio implies a hump-shaped consumption-wealth ratio against age. This form is familiar from several studies of income and savings profiles.¹⁴ Pissarides also argues that the **liquidity aspect of assets** is not an exogenous effect with respect to consumption, but rather an **endogenous property of net wealth**. The liquidity of a portfolio is chosen together with the consumption-savings decision. In practice the liquidity structure of the portfolio should tell something about the subjective time preference of the investor-consumer. The testable implication of this theory is that the composition of the portfolio could be used as a predictor in the consumption function.

Table 2 presents several static consumption functions which test the choice of the wealth concept in the equation and decompose the asset portfolio and thereby improve the forecasting performance. Testing with two proxies for net wealth shows that including wealth increases explanatory power and improves the diagnostics of the equation. Decomposition of net wealth indicates that the **influence of growth of financial wealth on consumption is also greater than that of real wealth**. The marginal propensity to consume by borrowing is also significantly different from zero, although we must not forget that financial wealth and debt are highly multicollinear. Regressions therefore show that the composition of the asset portfolio could certainly be used as an additional regressor in the consumption function. Models 6 and 7 indicate that the ratio of financial wealth to net wealth (FW/NW) or to real wealth (FW/RW) could also be used as regressors.

These static model estimates propose that including real net wealth improves the explanatory power, stability and residual diagnostics of the consumption equation. Disaggregation of net wealth into financial assets, real wealth and debt also give us some idea about the relative importance of these assets in financing consumption. Since the equation is estimated in logs, the parameters are also the elasticities.

Like Berg and Bergström (1993), we note that the elasticities of different assets do vary significantly. The average propensity to consume out of financial wealth is much higher than that out of real wealth. The stability of the static equation parameters was also tested. It is also useful to look at the rolling regression estimates of the static equation. According to figure 7 there have been significant changes in the parameters. The most important observation is the increased sensitivity of consumption to real income during 1986–1987, and the gradual decline after 1989. This may reflect the increasing importance of income expectations and the easing of liquidity constraints for expected labour income. The importance of net wealth in determining consumption has increased from the beginning of 1980s, but the effect does not appear to be strong because of financial liberalization. This is surprising, since the effect of financial deregulation should reduce the relative importance of current income for consumption and increase the significance of net wealth (see Bayoumi 1992).

Recursive estimation of the long-run static equation showed also that the significance of the composition of the asset portfolio has increased during the 1980s. This observation coincides closely with additional UK evidence by King (1990), who emphasises the role of wealth as a means to finance the observed

¹⁴ Finnish panel data also reveals a similar shape for the consumption profile in cross-sections (Sullström and Riihelä 1993).

over-spending. In Finland there is also evidence about over-spending, which could be found eg. from consumption-income ratios (figure 8). This indicates that monetary deregulation did affect purchases of durables. The decline in the household saving rate can be found in the deterioration of the current account balance as well.

From the static equations in table 2, we get the cointegrating regression Durbin-Watson (CRDW) test for the presence of cointegration. These tests already propose that cointegration could be found. Table 3 presents a cointegration analysis of a five-variable VAR system, where net wealth is disaggregated into financial assets, real estate wealth and debt. Disaggregated quarterly data on wealth was available from only 1979 onwards. In the multivariate framework, the exclusion of system variables can be tested as well. Exogeneity of the regressor with respect to a particular parametrization assumes Granger non-causality in the feedback relation.¹⁵ Among these variables only one cointegration relationship could be found. Recursive trace and maximum eigenvalue tests are plotted in figures 9–10. The variables in the system are not stationary and **none of them could be excluded from the system** without breaking it. If we assume that there exists two cointegrating vectors, the common zero degree homogeneity for both vectors ie. unit elasticity of consumption was rejected.

In addition the parameter homogeneity tests performed on the β -vector suggest that although the cointegrating relationship is strongest between consumption and real income, wealth variables should not be forgotten. According to LR-tests a broad definition of net wealth should be used in the analysis, and if this is not available net financial wealth (financial assets – debt) should be used. Comparing different wealth concepts in three variable settings, however, we could not reject the null hypothesis of one cointegration relationship in any of the three variable VAR systems. Adding durable consumption to the system does not change the result of single cointegration relationship either. Replacing real net wealth with either real financial wealth or real wealth is irrelevant, perhaps because of a common trend in all the wealth variables. According to **the eigenvector homogeneity test, financial wealth, net financial wealth and net wealth all satisfy the condition**. These results are however somewhat sensitive to the period chosen and the lag length of the VAR system.

With assumption of one rank (one c_i -vector), the tests on β -vector confirmed that the long-run elasticity of consumption with income and wealth is unity. This result holds even if we replace consumption with non-durable consumption.¹⁶ The stability of the β -vector is almost amazing (figure 11). The break down during

¹⁵ Prediction of a particular variable assumes only weak exogeneity, but for behavioural equations strong exogeneity is needed (Engle, Hendry & Richard 1983). However, Granger causality is a property of the data generating process, whereas exogeneity is a property imposed for the model specification and parametrization. Regressors in a model are said to be super-exogenous, if the parameters in interest are weakly exogenous and invariant to changes in the marginal densities of the weakly exogenous variables. Super-exogeneity is needed to ensure that ECM representation is a reduced form of a forward-looking intertemporal optimization.

¹⁶ The zero-degree homogeneity of the static equation was found only for the wider measure of net wealth. Replacing non-durable consumption with total consumption including durable purchases the homogeneity is rejected for 1979/Q1–1993/Q4. The homogeneity property between total consumption, disposable income and net wealth may not be accidental, since the system corresponds the aggregate budget constraint of the household sector.

1988–1989 matches with the peak in financial wealth and debt due to foreign borrowing. Even if we found only one cointegration relationship from the system including disaggregated wealth components, we may think eg. that financial asset and bank lending (household debt) could be cointegrated in the long-run. In addition it may be that consumers may try to keep on relatively constant ratio between housing debt and real estate wealth as a target. Therefore we tested also a couple of other parameter restrictions on β .

It turned out that ratios of different assets to income were clearly non-stationary. According to portfolio theory, it is also quite unlikely that separate asset-income ratios would be stationary, since the relative return on assets varies in time. Changes in rates of return would affect the portfolio allocation and we may rather expect that the whole net wealth is kept constant with respect to income.

However, according to ADF tests there was no indication for net wealth-income ratio to be stationary. This is quite obvious if we simply plot the ratio (figure 12).

In addition to homogeneity tests and exclusion tests for individual asset, the exclusion of net wealth could be tested in the system context. The restriction test for exclusion of net wealth ie. $\beta_3 + \beta_4 + \beta_5 = 0$ was rejected with probability level 0.003. In table 2 we tested and rejected the equality of financial wealth and real estate wealth coefficient. The same restriction $\beta_3 = \beta_4$ posed on the system confirmed the same result with probability level 0.002. However, as well as before in table 2 the tested long-run proportionality of real estate wealth and real debt $\beta_4 + \beta_5 = 0$ was not rejected. This may reflect two phenomenon. Firstly debt is mostly used for financing purchases of real estate wealth like housing. In addition real wealth is used as collateral for bank loans. As we have seen not only inclusion of wealth but also disaggregation of wealth seems to increase our insight about consumption changes. These observations correspond to some extent those in Patterson (1984). Even with disaggregated wealth, we got a very close long-run unit 'permanent' income elasticity.

Table 2. **Consumption function with different wealth variables,**
Dependent variable: Logarithmic consumption

	Long period: 1970/Q1–93/Q4		Short period: 1979/Q1–1993/Q4		Model 5	Model 6	Model 7
	Model 1	Model 2	Model 3	Model 4			
LRYD	.996 (400.83)	.886 (50.72)	.706 (32.88)	.65 (28.24)	.67 (31.38)	1.05 (322.51)	1.00 (318.97)
LRNW		.092 (6.31)	.231 (13.53)			.004 (2.66)	.004 (2.83)
LRFW				.28 (6.12)	.27 (6.27)		
LRRW				.15 (7.93)	.14 (7.16)		
-LRDEBT				-.14 (-5.88)			
-LRDEBT(-1)					-.12 (-6.07)		
FW/NW						-0.37 (-2.11)	
FW/RW						-.43 (-2.51)	
R ²	.999	.999	.999	.999	.999	.999	.999
CRDW	1.10	1.30	1.41	1.88	1.87	0.76	0.79
P-values for residual diagnostics:							
AR(5)	.000	.000	.119	.922	.904	.000	.000
ARCH(4)	.020	.168	.328	.986	.955	.516	.866
HETERO.	.001	.000	.628	.219	.070	.019	.018

ADDITIONAL WALD TESTS FOR DISAGGREGATED WEALTH COEFFICIENTS

Exclusion	Model 4		Model 5	
	F(1,56)	P-val.	F(1,56)	P-val.
LRFW	37.45	.000	39.35	.000
LRRW	62.90	.000	51.19	.000
LRDEBT	34.53	.000	36.84	.000
All together	208.16	.000	223.09	.000

Proportionality between wealth effects

	Model 4		Model 5	
	F(1,56)	P-val.	F(1,56)	P-val.
LRFW = LRRW	5.64	.021 *	5.46	.023 *
LRRW = LRDEBT	0.07	.798	0.18	.673
LRFW + LRDEBT = LRRW	0.01	.951	0.06	.807

Table 3.

Testing the number of cointegration vectors

(VAR model with 2 lags, a trend, no seasonal terms)

Estimation period: 1979/Q1–1993/Q4, 58 observations

System variables: (LC, LRYD, LRFW, LRRW and LRDEBT)

ie consumption, income, financial and real wealth and debt

Hypothesis		Eigen- value	Maximum Eigenvalue test		Matrix Trace test	
Null	Alt.		Test Stat.	95% Cr.v.	Test stat.	95% Cr.v.
$r = 0$	$r = 1$.5321	41.014	33.461	79.584	68.524
$r \leq 1$	$r = 2$.3073	19.822	27.067	38.569	47.210
$r \leq 2$	$r = 3$.2247	13.739	20.967	18.747	29.680
$r \leq 3$	$r = 4$.0785	4.414	14.069	5.008	15.410
$r \leq 4$	$r = 5$.0109	.593	3.762	.593	3.762

 β eigenvectors and α adjustment coefficients (chosen $r = 1$)

Variable	β -coeff.	Normalized β	α -coeff.	Normalized α
LC	13.61	-1.00	-.021	.296
LRYD	-9.07	.67	.046	-.629
LRFW	-5.20	.38	.036	-.491
LRRW	-1.81	.13	.042	-.596
LRDEBT	2.73	-.20	-.006	.086

Estimated Long Run Matrix (π)

	LC	LRYD	LRFW	LRRW	LRDEBT
LC	-.2961	.1973	.1130	.0393	-.0594
LRYD	.6294	-.4195	-.2402	-.0835	.1263
LRFW	.4913	-.3274	-.1875	-.0652	.0986
LRRW	.5764	-.3842	-.2200	-.0765	.1157
LRDEBT	-.0859	.0573	.0328	.0114	-.0173

LR-TESTS FOR STATIONARITY, EXCLUSION AND WEAK EXOGENEITY

Critical values presented at 95 % significance level

Variable	STATIONARITY (DF = $p - r = 4$)		EXCLUSION (DF = $r = 1$)		WEAK-EXOGENEITY (DF = $r = 1$)	
	$\chi^2(4)$	CR.Value	$\chi^2(1)$	CR.Value	$\chi^2(1)$	CR.Value
LC	34.67	9.49	20.15	3.84	3.30	3.84
LRYD	33.10	9.49	17.92	3.84	3.71	3.84
LRFW	35.05	9.49	13.70	3.84	2.26	3.84
LRRW	34.71	9.49	13.85	3.84	4.36	3.84
LRDEBT	34.68	9.49	8.00	3.84	0.04	3.84

TESTING HOMOGENEITY RESTRICTIONS:

Changes in consumption proportional to income changes

 $H_0: \beta_1 + \beta_2 = 0$, i.e. $\beta_1 = -\beta_2$ LR-test statistic for $H_0: \chi^2(1) = 3.832$, ($P = 0.050$)

Changes in consumption proportional to income and financial wealth changes

 $H_0: \beta_1 + \beta_2 + \beta_3 = 0$, i.e. $\beta_1 = -(\beta_2 + \beta_3)$ LR-test statistic for $H_0: \chi^2(1) = 0.096$, ($P = 0.757$)

Changes in consumption proportional to income and net financial wealth changes

 $H_0: \beta_1 + \beta_2 + \beta_3 + \beta_5 = 0$, i.e. $\beta_1 = -(\beta_2 + \beta_3 + \beta_5)$ LR-test statistic for $H_0: \chi^2(1) = 1.640$, ($P = 0.200$)

Changes in consumption proportional to income and net wealth changes

 $H_0: \beta_1 + \beta_2 + \beta_3 + \beta_4 + \beta_5 = 0$, i.e. $\beta_1 = -(\beta_2 + \beta_3 + \beta_4 + \beta_5)$ LR-test statistic for $H_0: \chi^2(1) = 0.039$, ($P = 0.843$)

Figure 7.

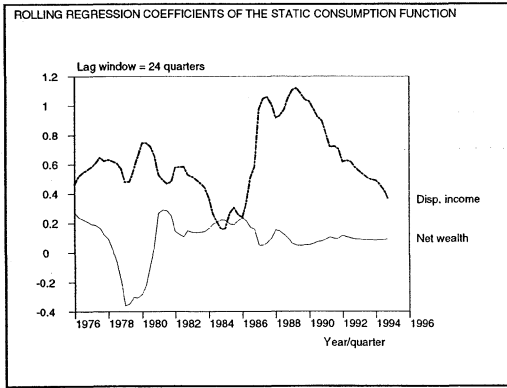


Figure 8.

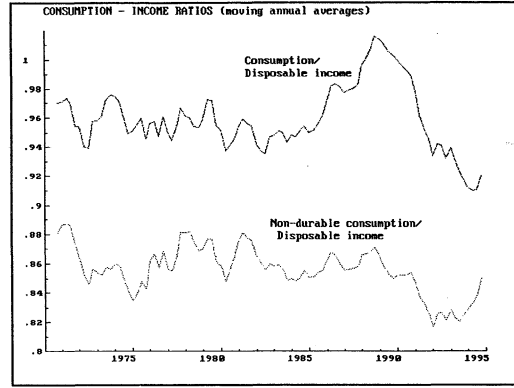


Figure 9.

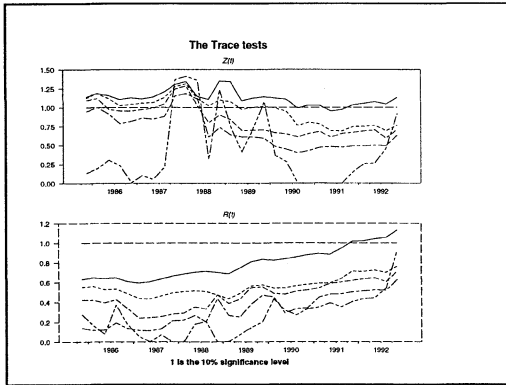


Figure 10.

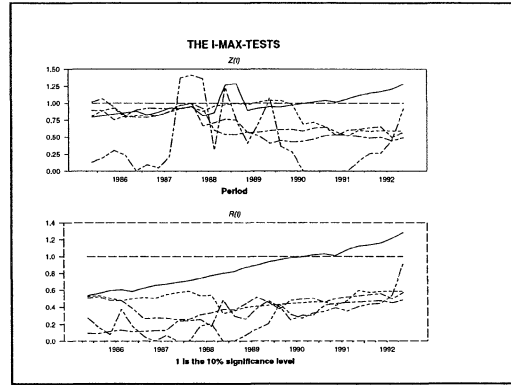


Figure 11.

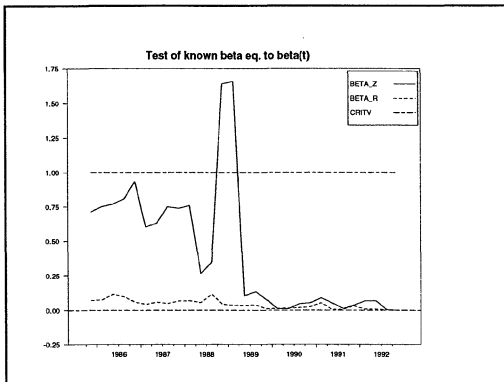
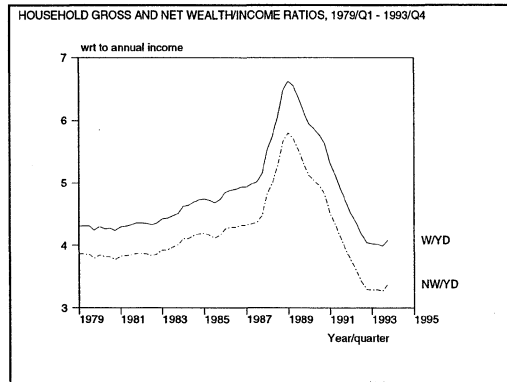


Figure 12.



6.2 Cointegration system estimations

A closer look at cointegration was started with a two-variable (total consumption and real disposable income) model. According to trace and maximal eigenvalue tests, no conclusion about cointegration could be drawn at the 5 % significance level. Although the normalized β coefficients were relatively close to each other, proportionality between parameters is rejected at 2.2 percent probability level. Table 4 presents the results from a more successful estimation between **three integrated variables** (now non-durable consumption, real income and the narrow proxy for real net wealth). In Finland durable consumption is measured on the basis of purchases. The accounting procedure exaggerates the depletion of durables, by which consumption is overestimated and the stock of durables is underestimated. We expect that the consumption of durables is more sensitive to interest rates and income expectations than non-durable consumption. The consequence is that durable consumption is much more volatile and has reacted strongly eg. to devaluations and financial deregulation.

Tests and plots show quite unambiguously that there is only one stationary cointegrating vector between the variables. However, it is always possible that this relation is a part of an even larger system. For example we could include the real market interest rate in the system, but still just one cointegration vector appears. The real interest rate should be stationary and therefore not included in the core of the cointegration system for consumption. The appearance of the interest rate can be motivated by its role as a measure of the opportunity cost of net wealth. Including the real interest rate as an additional explanatory variable in the system makes net wealth less significant, as might be expected according to the present value formulae of real estate wealth.

In order to test for exclusion, LR-tests were performed to see whether any of the endogenous system variables could be eliminated from the long-run relations. The exclusion of a certain variable from the cointegration vector was again tested by setting the particular β vector relating to the variable in the system zero. Table 4 indicates also that aggregated real net wealth cannot be excluded from a proper specification of the consumption function. The proportionality test between β 's shows that we cannot reject proportionality between consumption, income and wealth. Therefore it is quite clear that the assumption of a cointegration relation between consumption, real disposable income and real net wealth holds. The parameter restriction of a zero homogeneous system cannot be rejected at the 5 percent significance level.

We have emphasized the necessity of including net wealth in the consumption equation. Therefore we must discuss the role of wealth in this context. It was mentioned already earlier that net wealth is time-aggregated widely defined Hicksian savings. According to the REPIH, saving (and borrowing) is used to smooth consumption with respect to changes in expected labour income. The buffer-stock view of savings gives another interpretation of wealth in the model by considering wealth as a means to prepare against the uncertainty of varying income.

There are also other reasons for having wealth in the consumption function. We have seen eg. that the average propensity to consume out of different types of wealth differs between financial wealth and real estate wealth.

Table 4.

Testing the number of cointegration vectors

(VAR model with 2 lags, a trend, no seasonal terms)

Estimation period: 1970/Q1–1993/Q4, 94 observations

System variables: (LNONCD, LRYD, LRNW, i.e. Non-durable consumption, real income and 'narrow' real net wealth)

H(rank)	Eigenvalue	Maximum Eigenvalue test $-T \cdot \text{Ln}(1-\lambda_i)$		Matrix Trace test $-T \cdot \sum \text{Ln}(1-\lambda_i)$	
		Test Stat.	Crit. value $\lambda_{\max}(.95)$	Test stat.	Crit. value $\lambda_{\text{trace}}(.95)$
$r \leq 2$.026	2.463	3.76	2.463	3.76
$r \leq 1$.074	6.990	14.06	9.453	15.41
$r = 0$.235	24.140	20.97	33.593	29.68

STANDARDIZED β' EIGENVECTOR (chosen $r = 1$)

Variable	β -coeff.	Normalized β
LNONCD	6.67	-1.00
LRYD	-5.99	0.90
LRNW	-0.40	0.06

STANDARDIZED α Coefficients (Error-correction loadings)

Variable	α -coeff.	Normalized α
LNONCD	.0003	-.002
LRYD	.107	-.714
LRNW	.042	-.281

Estimated Long Run Matrix (π)

	LNONCD	LRYD	LRNW
LNONCD	.002	-.002	-.000
LRYD	.715	-.642	-.043
LRNW	.281	-.253	-.017

LR-TESTS FOR STATIONARITY, EXCLUSION AND WEAK EXOGENEITY

Critical values presented at 95 % significance level

Variable	STATIONARITY (DF = $p - r$)		EXCLUSION (DF = $r = 1$)		WEAK-EXOGENEITY (DF = $r = 1$)	
	$\chi^2(2)$	CR.Value	$\chi^2(1)$	CR.Value	$\chi^2(1)$	CR.Value
LNONCD	20.65	5.99	11.41	3.84	1.81	3.84
LRYD	19.63	5.99	11.12	3.84	13.73	3.84
LRNW	21.87	5.99	3.92	3.84	4.69	3.84

TESTING STRUCTURAL PARAMETER RESTRICTIONS:

Proportionality tests:

Changes in non-durable consumption proportional to income and net wealth

 $H_0: \beta_1 + \beta_2 + \beta_3 = 0$, i.e. $\beta_1 = -(\beta_2 + \beta_3)$ LR-test statistic for $H_0: \chi^2(1) = 0.887$, (P = 0.346), $\chi^2_{.95}(1) = 3.84$

Changes in non-durable consumption proportional to changes in real income

LR-test statistic for $H_0: \chi^2(1) = 1.479$, (P = 0.224), $\chi^2_{.95}(1) = 3.84$

6.3 Recursive VAR estimations

Recursive analysis of long-run parameters relaxes the assumption of parameter constancy. Following Hansen and Johansen (1992) it is possible to fix the short-run parameters and consider changes in the long-run parameters recursively. Parameter constancy can then be tested by means of recursive estimation. Since our estimation period might include large structural changes due to financial liberalization and deep recession, this is crucial for the existence of a stable consumption function.

Figures 13 and 14 present the recursive estimates of the trace and maximum eigenvalue test statistics. The tests confirm once again that only one cointegration relation could be found between variables, since the test values are above the scaled critical values for only one eigenvalue for most of the time. The test statistic graphs are generally upward sloping toward the end of the 1980s. However, the stability of the test statistics is not perfect. The assumption of one cointegration rank is not seriously under dispute. According to the Z-representation the cointegration relationship for the largest eigenvalue is somewhat stronger, but in the R-representation both test statistics are more stable. The Z-representation presents the stability of the rank in a recursion where the sample size increases from t to T , while the R-representation answers the question of cointegration rank constancy when the short-run parameter estimates are fixed over the process of estimating the long-run parameters recursively.¹⁷

Figures 15–16, which show the one-step prediction error test and prediction error tests for the three series (in order consumption, income and wealth), which identify a few outliers in the series. The most extreme outliers could be found in the first quarter of 1988. This observation agrees with other studies dealing with the boom in bank lending, consumer debt and housing loans (Brunila and Takala 1993). Housing loans were allowed to be tied to long-term market interest rates from the beginning of 1988. In the residual series we also see that the beginning of 1992/Q2 has significant outliers in income and wealth equations perhaps due to postponed tax refunds.

The constancy of the β parameter can be seen from figure 17, which presents evidence on the question of whether the $\beta(T)$ at the end of the sample period could be considered to be spanned by $\beta(t)$.¹⁸ Figure shows that the β -vector based on level autoregressions (R-representation) is always below the 5 % critical value and therefore constant. Z-representation based on difference autoregressions does violate β -constancy for same time during 1982–85, but after that constancy is preserved.

Hansen and Johansen (1992) argue that structural changes in the loadings (α) and cointegration vectors (β) will be reflected in the time paths of the estimated recursive eigenvalues. The plotted eigenvalues are shown in figure 18. In the largest eigenvalue, signs of structural breaks are not obvious and we could accept the hypothesis of a constant cointegration relation.

¹⁷ Hansen and Johansen (1992) emphasize that from the point of view of the stability of the cointegration relation, the stability in the tests of the R-representation is more relevant. The Z-representation is based on an estimation of VAR model, where all the parameters are updated recursively, whereas in R-representation the short-run parameters are fixed.

¹⁸ The selection of $\beta(t)$ is arbitrary, but if there is no particular period of interest, it is useful to select $\beta(T)$ since it is the estimate with the smallest sample variation (Hansen and Johansen 1992).

Figure 13.

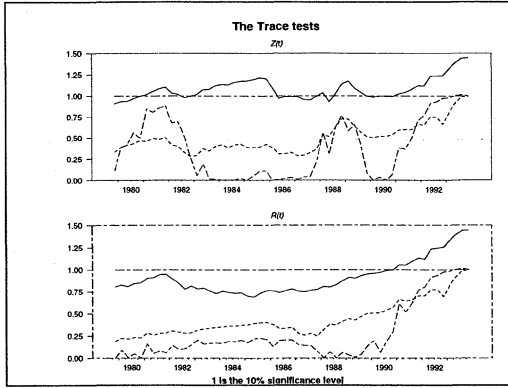


Figure 14.

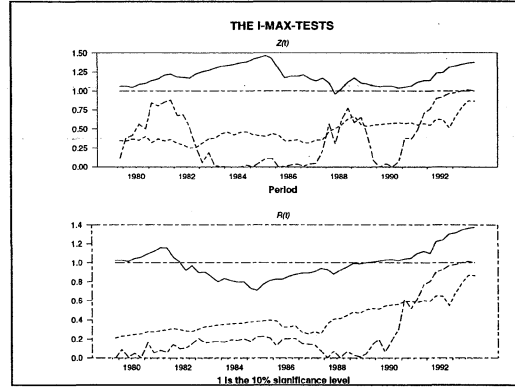


Figure 15.

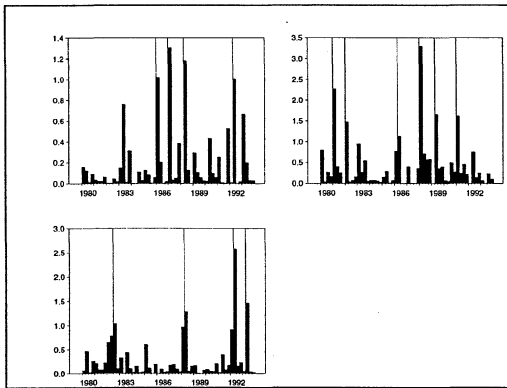


Figure 16.

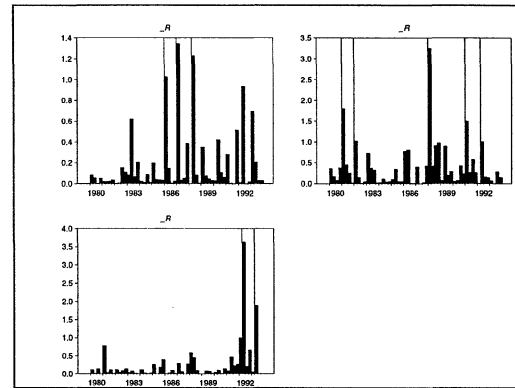


Figure 17.

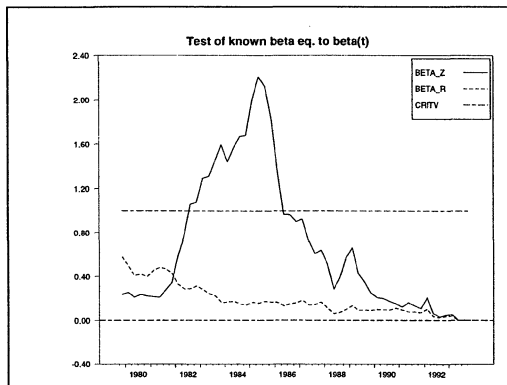
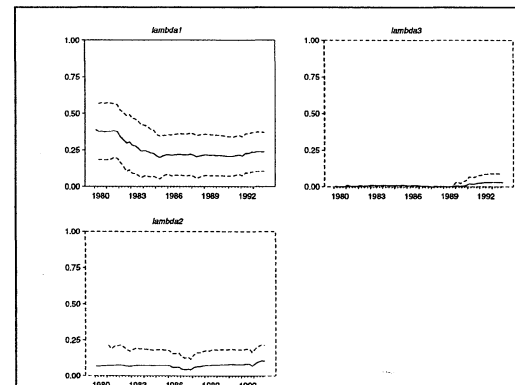


Figure 18.



6.4 Specification of the consumption function

In table 5 we finally present a specification of an error-correction consumption function for differences of non-durable consumption. The variables included in the core of the cointegration vector and the error correction term are all highly significant. Because of the endogeneity of disposable income and net wealth, the model was estimated using instrumental estimation with lagged values of income as the instruments.

Figure 19 compares the savings rate and the estimated error-correction term (ECT) of the cointegration relation. Since non-durable consumption is almost proportional to the proxy for permanent income and therefore highly correlated with real disposable income, the correlation between ECT and the savings rate is not a surprise.¹⁹ This gives us motivation to interpret savings as a proxy for equilibrium error in the adjustment of consumption. Campbell and Mankiw (1991, p. 729) argue, however, that ECT (lagged savings) cannot be seen to represent any kind of disequilibrium. This is mostly a matter of semantics, since as income dominates the consumption determination it causes the error-correction term to correlate strongly with the saving rate, as mentioned. In practice we do not observe the error-correction term directly, but we surely have some idea about savings behaviour. Stabilizing feedback effects coming through ECT may also reflect changes in assets that affect consumption, ie. when consumption and income are not equal, there is saving, which affects cumulative savings. It must be remembered that a pure difference equation without ECT has no equilibrium solution, but it could still be consistent with a steady state solution.

According to Granger causality tests, the saving rate anticipates strongly and negatively changes in expected income with lags 1–3 quarters. This conclusion agrees closely with Campbell's (1987) results, which follow directly from the cointegration relationship between consumption and income. One implication of the cointegration relationship is that the equilibrium-error term Granger causes at least one of the cointegrated variables. Savings Granger causes both consumption and especially income growth. In this sense, savings has an anticipatory role a predicting a decline in income, which is in accordance with the precautionary motive for savings. The dependent variable and model fit are compared in figure 20.

The specification also includes a few stationary variables which are supposed to explain the short-run adjustment in consumption. At first inflation was included, and it turned out to be highly significant. In the regressions, inflation was also separated into expected and unexpected inflation with an AR(5) model, but both components turned to be significant and realized inflation was left in the equation.

The relationship between consumption and inflation also demands a closer look. Inflation may affect consumption in several ways. One explanation tells us that when prices rise rapidly consumers cannot distinguish between changes in relative prices and changes in the overall price level. In order to safeguard themselves from inflation and falling purchasing power, consumers accelerate consumption (especially of durables). Another way to maintain purchasing power is

¹⁹ The ECT was calculated from the unrestricted static long-run equation, which did not impose the homogeneity restriction. The reason for this that homogeneity did not quite hold with the narrow net wealth for period 1970–93.

to demand a higher inflation premium for saving. Therefore as expected inflation increases, so do nominal interest rates. This also affects gross interest income. As households are net lenders to other sectors, their interest income will rise because of inflation.

HUS (1981) argued that large increases in nominal interest receipts are balanced by capital losses in financial assets, but whereas gross interest income is included in disposable income, capital losses are not. Therefore the national accounting statistic do not fully reflect the economically perceived real income.²⁰ However, it is clear that increasing unexpected inflation could cause major losses to owners of non-indexed financial assets like deposits and on the other hand, capital gains to debtors. In Finland this interpretation does not apply to the latest fall in the saving rate in the late 1980s as inflation has rapidly declined. Rather it is more likely to be due to increased spending in durables and housing. If inflation could be responsible for the decline in the saving rate during financial liberalization, there should be a significant negative correlation between changes in financial assets and inflation. Such a phenomenon could not be found in the late 1980s. Muellbauer & Murphy (1989) argue that consumers who are not liquidity constrained would be affected by changes in real interest rates, which reduces the willingness to borrow. On the other hand, households could be affected by nominal interest rates as well, since the nominal burden from debt will increase the debt service payments and therefore reduce consumption.

In Finland debt service costs have not been very sensitive to interest rates, because prior to 1988 housing loans were tied to the central bank base rate, which has been changed only by political decision. What has affected debt service costs is the increasing indebtedness. Since most consumers must be forward looking in their consumption, we proxied the intertemporal price of consumption with an auxiliary regression for the expected real interest rate. In some specifications a distinctive nominal interest rate effect was found. However, the presented specification in table 5 rejected the equality of real lending rate and inflation coefficients.

As an empirical observation it seemed that the dynamics of income, wealth and stationary variables can be specified in various alternative ways. However, it maybe more helpful to specify the model with 'forcing' exogenous variables than with lagged endogenous, even though they may reflect the slowness in adjustment. The results from the regression equation can be presented in a nice way by presenting the endogenous variable as a decomposition of additive contribution components (regression coefficient times the explanatory variable). From figure 21 it can be seen eg. that the explanatory power of net wealth and real interest rate has substantially increased during 1980s'. The importance of unemployment has emerged merely during the recession of 1990s'. On the other hand the effect of inflation has decreased and the significance of expected real interest rate increased.

²⁰ Hendry and von Ungern-Sternberg conclude that if the income elasticity of consumption is unity in the long run, the fall in the consumption-income ratio during the 1970s must be related to incorrect measurement of income due to inflation effects.

In the model the relative price between durables and non-durables has an overemphasized power, since it has taken the role of constant as well.²¹ It was also found necessary to use few dummy variables to take into account effects of outliers that otherwise would affect parameter estimates. Separate dummies were used for couple of outliers for 1975/Q4–1976/Q1, financial deregulation 1987/Q1–1987/Q4 and the collapse of the Soviet trade during 1991/Q1–1992/Q1.

The consumption of services is included in the consumption of non-durables. Relative price variable is based on price adjustment, which indicates the direction in which consumption will evolve. The purely competitive market story of consumption balancing would indicate that the price of consumption is taken as given and equilibrium in the market is attained by adjusting the volume of consumption. However, it seems clear that to some extent prices adjust as well, and so disaggregating prices should be useful in prediction. Price formation in consumption cannot be just as immediate as supposed. According to Granger causality tests, the prices of durables and non-durable goods will predict the prices of services. This could be caused eg. by nominal wage rigidity and rigid price setting of public services. Lastly the unemployment rate was found to be significant predictor for consumption. Adding unemployment rate into the equation can be interpreted to describe the effect of income uncertainty on consumption (Carroll, 1992).

²¹ Constant was left out of the equation since it did not prove to be significant. Even though the estimate of the relative prices coefficient is 5.77, the contribution is around 4.3. The mean of non-durable consumption was 4.2 %.

Figure 19.

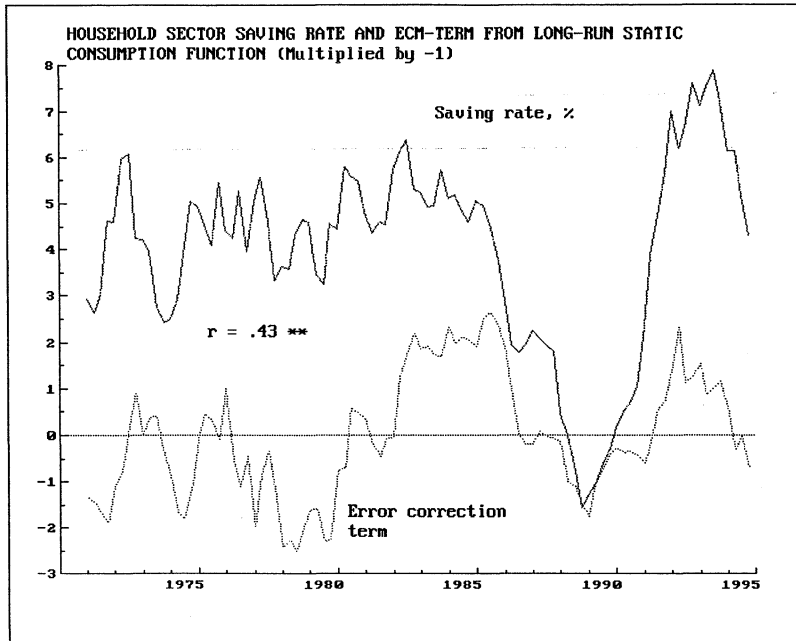


Figure 20.

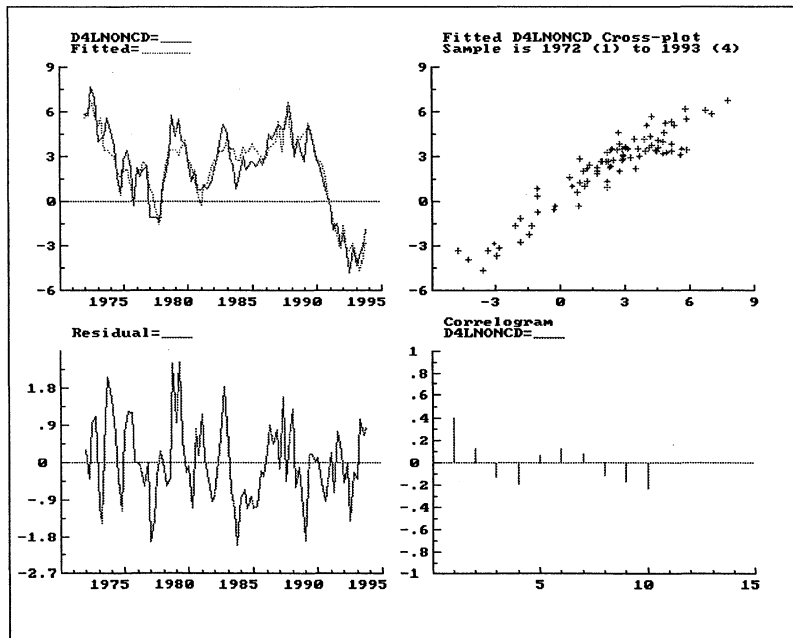


Table 5. **The model for non-durable consumption,**
 Dependent variable: D4LNONCD
 OLS- and IV-estimations, 1972/Q1–1993/Q4

Variable	OLS-estimates			IV-estimates		
	Coefficient	Std.Error	t-value	Coefficient	Std.Error	t-value
D4LRYD	0.277	0.0493	5.617	0.625	0.1902	3.290
D4LRNW	0.070	0.0140	4.981	0.061	0.0186	3.244
ECT-Lag4	-0.310	0.0740	-4.184	-0.613	0.1826	-3.357
D4LCPI	-0.345	0.0576	-5.990	-0.197	0.1062	-1.855
ERRLBN4	-0.171	0.0594	-2.885	-0.061	0.0947	-0.654
PCDrPCND	5.771	0.7543	7.650	2.581	1.9047	1.355
UR	-0.283	0.0549	-5.168	-0.070	0.1305	-0.539
FINDEREG	1.3986	0.6282	2.226	0.630	0.8987	0.701
SOVTRADE	-1.5678	0.5702	-2.749	-2.153	0.7920	-2.719
i1975p4	-2.5842	1.0716	-2.412	-3.995	1.5560	-2.567
i1976p1	2.2305	1.0901	2.046	4.490	1.8194	2.468

Testing parameter restriction: $\beta(D4LCPI) = \beta(ERRLBN)$

OLS: $F(1,77) = 15.960, p = .0000 **$

IVE: $F(1,77) = 5.168, p = .0258 *$

Model performance and residual diagnostics	OLS-estimation		IV-estimation	
	Test	P-value	Test	P-value
$R^2 =$	0.93		0.90	
DW =	1.21		1.50	
$\sigma =$	1.01		1.29	
AR;	$F(5,72) = 4.76$.001**	$\chi^2(2) = 4.45$.486
ARCH;	$F(4,69) = 0.29$.881	0.45	.772
NORMALITY;	$\chi^2(2) = 1.38$.503	4.41	.110
HETEROSCED.;	$F(18,58) = 1.31$.215	0.96	.507
RESET;	$F(1,76) = 0.72$.399		

Auxiliary regression:

The estimated model of real lending rate expectations (ERRLBN4)
 Modelling RRLBN by OLS, 1970/Q1–1993/Q4

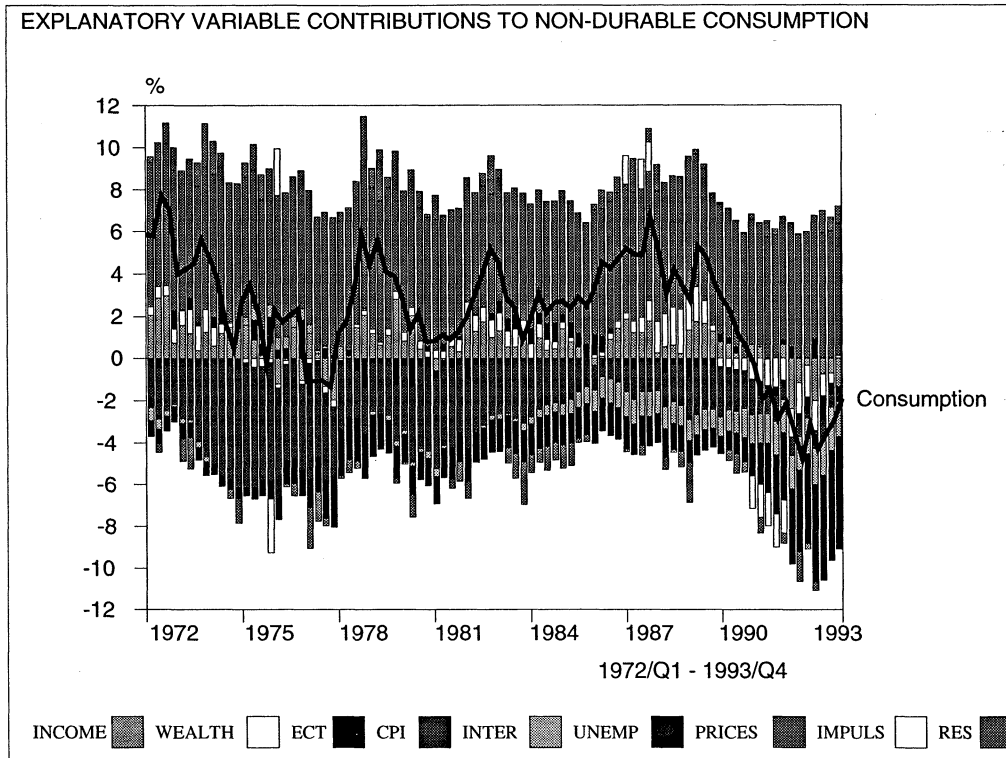
Variable	Coefficient	Std.Error	t-value	t-prob	PartR ²
Constant	-4.030	1.2059	-3.342	0.0011	0.0820
RRLBN_4	0.68187	0.060340	11.300	0.0000	0.5053
RLBN_4	0.51501	0.12992	3.964	0.0001	0.1117

$R^2 = 0.630$ $\sigma = 2.638$ DW = 0434

Variables:

- D4LNONCD = Annual log difference of non-durable private consumption, %
 D4LRYD = Annual log difference of real disposable income, %
 D4LRNW = Annual log difference of real net wealth, %
 ECT-Lag4 = Error correction term from static long-run equation between non-durable consumption, real income and net wealth
 D4LCPI = Log annual difference of consumer price index, %
 PCDrPCS = Relative price between durable prices and services
 ERRLBN4 = Expected real lending rate for new loans, %
 RLBN = Nominal lending rate for new loans, %
 RRLBN = Real (deflated by CPI) lending rate for new loans, %
 UR = Unemployment rate, %
 FINDEREG, SOVTRADE, i1975p1, i1976p4 = Impulse dummy variables

Figure 21.



7 Conclusions

This paper has presented an updated error-correction specification of a consumption function for non-durable consumption in Finland. Based on the permanent income hypothesis, consumption should depend on current labour income, net wealth and the present value of human capital (income expectations). In accordance with this, a cointegration relationship including non-durable consumption, disposable income and net wealth was formulated and tested. Therefore a linear combination of disposable income and net wealth represents a proxy of permanent income, which is used by consumers as an implicit budget constraint in the optimization of consumption.

Estimation results showed that a more stable consumption function could be attained if net wealth is included in the cointegration relation. In fact the results from the disaggregation of net wealth showed quite clearly that there exists only **one cointegration relation** between consumption, income and net wealth. On the other hand, there was no strong evidence of any stable wealth-income relationship.

Data problems were confronted in the construction of net wealth due to the financial deregulation of the late 1980s. Structural changes in the parameters of the cointegration relation were also found. Although our results confirm a distinct wealth effect during deregulation, they also indicate that the more important effect may be due to favourable income expectations. During 1987–90 the marginal propensity to consume out of real disposable income rose significantly above one. Overspending was due to an increased demand for durables and housing purchases, which lead to significant indebtedness.

A recursive static long-run equation and VAR analysis indicated that parameter changes have appeared during the late 1980s, especially in real income. Same conclusion can be drawn from the rolling long-run static equation estimation of the consumption function. Combining this evidence with the effect of financial deregulation hints of changes in income expectations and uncertainty. It seems clear that the role of income uncertainty has to be studied more carefully in the context of the consumption function. Consumption depends on income expectations and it is not possible to model consumption without modelling the income process. Net wealth can be seen to proxy for income expectations through real estate asset prices.

A single equation equilibrium correcting 'consumption function' was estimated in differences by using real disposable income and real net wealth (lagged two quarters) plus a lagged error correction term as the basic variables. Additional stationary variables were used to account for the short-run adjustment; inflation, the expected real interest rate, unemployment and relative price between durables and non-durables. The cointegration analysis also showed that the disaggregation of wealth was needed in order to understand the turbulent period of increasing indebtedness.

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Appendix

Figure 1. **Household sector financial wealth**

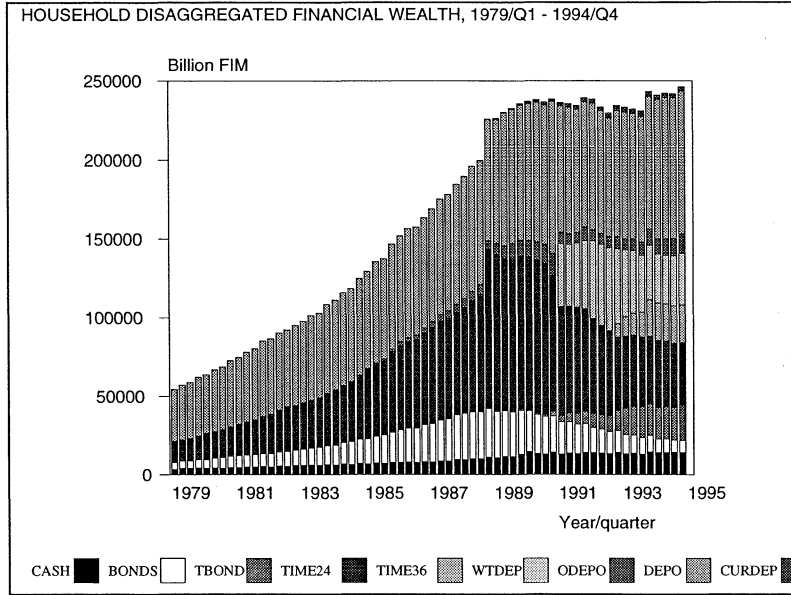


Figure 2. **Household debt**

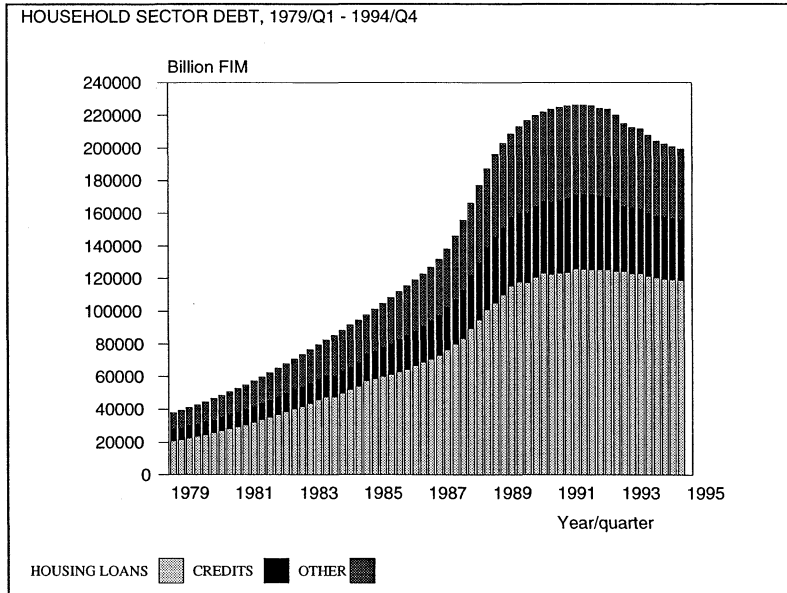
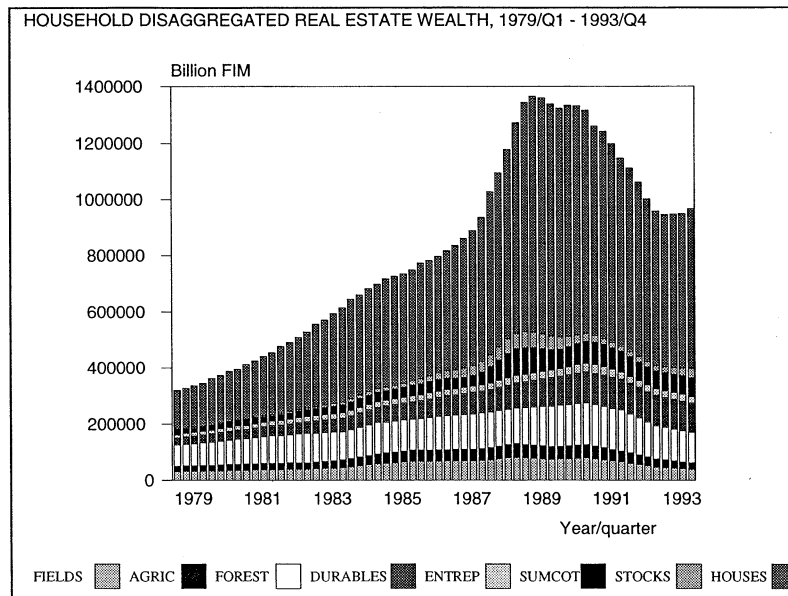


Figure 3. Household sector real estate wealth



Financial wealth

- Cash = Cash holdings
- Bonds = Non-taxed government bonds and other bonds
- Tbond = Taxed bonds (issued from 1989 onwards)
- Time24 = Taxed 24 month time deposits
- Time36 = Taxed 26 month time deposits
- Wtdep = Withholding taxed deposits
- Odepo = Other taxed deposits
- Depo = Ordinary non-taxed deposits
- Curdep = Currency deposits

Household debt

- Housing
 - loans = Household sector housing loans
 - Credits = Consumer credits
 - Other = Other loans including entrepreneurial loans

Real estate wealth

- Fields = Agricultural wealth in form of fields
- Agric = Other agricultural wealth like barns, equipment etc.
- Forest = Forest real estate wealth
- Durables = Household durables, autos, boats etc.
- Entrep = Personal entrepreneurial wealth
- Sumcot = Summer cottages
- Stocks = Stocks in firms Helsinki stock exchange
- Houses = Housing wealth (houses, flats etc.)

Excess-sensitivity testing of consumption with Finnish data

Kari Takala

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Abstract

This paper performs Campbell-Mankiw-type Euler equation tests of excess-sensitivity of consumption to current income based on Finnish data. Estimations are performed using both fixed and time-varying parameter methods. Because the probable causes of excess sensitivity are unknown, parameter stability is tested with different instruments in IV estimation. Most of the fixed parameter estimates indicate that about a half of the consumers may have been liquidity-constrained during 1980–1998. The estimates are not highly sensitive to assumptions about normality, estimation method (IV, GMM or robust estimation) or instruments employed.

Estimating the excess-sensitivity parameter as a time-varying coefficient, we observe first a decline – due to the financial deregulation of the late 1980s – followed by a rise during the recession and finally a declining trend. The time-varying real interest rate tends to smooth out the excess-sensitivity parameter and absorb the observed rise in excess sensitivity during the recession. The estimations suggest that liquidity constraints related to excess sensitivity have probably eased over the course of the estimation period.

Key words: permanent income hypothesis, excess sensitivity, time-varying estimation

* I would like to thank Erkki Koskela, Pasi Holm and Matti Virén for comments, Markku Rahiala for statistical advice and Glenn Harma for checking the language. The usual disclaimer applies.

1 Introduction

Since Robert Hall's (1978) seminal work on the Euler equation approach to optimal consumer behaviour, empirical studies on the consumption function have tried to explain sensitivity of consumption to current disposable income. The main implication of Hall's framework, in which rational forward-looking consumers maximise expected lifetime utility, is that changes in consumption should be unpredictable, ie that consumption itself should follow a random walk and that current consumption should be the best prediction of next-period consumption.

Although numerous studies have rejected the strict random walk property of consumption, it has been recognised that the path of consumption is rather close to a pure random walk. There is an obvious response to the rejection of orthogonality based on publicly available data such as lagged income: liquidity constraints, eg credit rationing, may prevent consumers from optimising and smoothing consumption with respect to lifetime wealth. Most other empirical studies reject orthogonality with respect to lagged income, which has been regarded as evidence in favour of liquidity constraints (eg Flavin 1981, Campbell and Mankiw 1989).

Recent papers (eg Bacchetta and Gerlach 1997) emphasise the importance of credit variables (both availability and cost of credit) in determining current consumption. They also acknowledge that the PIH allows monetary policy to affect consumption only via changes in permanent income. They find that credit market variables predict significantly consumption changes and hence might be useful in controlling overall demand. Between internal financing (eg deposits) and external financing (borrowing), a premium must be taken into account, which can be approximated by the margin between lending and deposit interest rates.¹ However, it should be noted that changes in the price of external financing affect consumers' consumption-saving choices.² It is assumed that this balance sheet effect may be more stringent during recessions, due to informal frictions. Of course this is another hypothesis that should be tested with empirical time series.

¹ In representative agent consumer theory, capital markets are usually assumed to be perfect and therefore consumers are freely able to lend or borrow at the same nominal interest rate, which makes the budget constraint linear and the solution more straightforward without serious loss of generality.

² De Bondt (1999) finds some evidence that in Germany the external finance premium can be used to predict consumption.

Liquidity constraints on households may promote a higher saving rate since households are not able to borrow as much as they would wish. Capital market imperfections may promote growth in the economy since they raise the rate of return via more efficient allocation of capital through investment. For instance, down payments for housing loans may affect the variability of housing prices and subject households to forced saving. Therefore changing the down payment ratio – as often happens during a liberalisation period – does affect the saving ratio as well as economic growth. However, this result applies to closed economies only. In small open economies the rise in the saving rate due to liquidity constraints is likely to be smaller or even ambiguous. With perfect capital markets and unrestricted capital mobility, investment and growth should not affect the saving rate. Actually, it seems that there is always some advantage in information due to country-specific aspects of the financial market and that capital flows are subject to some inertia.³

The research task in this paper is to establish an estimate of the proportion of liquidity-constrained consumers in Finland. A large proportion of liquidity-constrained consumers would confirm excess sensitivity. However, liquidity constraints may be only one of several reasons for excess sensitivity. Excess sensitivity may also arise from non-life cycle behaviour, time-varying interest rates, or income uncertainty, as suggested recently by Madsen and McAleer (2000).

This paper uses the same testing methodology as Bacchetta and Gerlach (1997) and Taylor and Sarno (1998), ie time-varying estimates of coefficients for disposable income are calculated using a simple Campbell and Mankiw (1989, 1991) framework. Time-varying estimates are calculated by means of state space modelling and Kalman filter estimation. Estimates of the proportion of liquidity-constrained consumers may depend on model specification, so it is important to determine what factors affect estimation of the proportion of liquidity-constrained consumers and whether the proportion is varying over time. Thus we want to determine the magnitude of uncertainty involved in estimating the importance of liquidity-constrained consumers for consumption expenditure. In addition, we

³ Jappelli and Pagano (1994) note that in Scandinavian countries consumer credit exceeds 10% of national income whereas in many western European countries credit amounts to well below 3%. These differences are not entirely explained by liquidity constraints; preference for borrowing, demographic features and tax incentives are also important. They also assert that the main reasons for cross-country differences in liquidity constraints include regulation, information on borrowers' default risk and cost of enforcing loan contracts.

are interested in knowing what economic conditions may affect the sensitivity of consumption to disposable income. Of course any evidence of excess sensitivity would comprise an important strand of empirical evidence against the permanent income hypothesis, which is still – despite its deficiencies – the major modern theory of the consumption function.

The rest of the paper is organised as follows. We start by deriving the Euler equations in section 2. The proposed Campbell-Mankiw-type testable equations presented in the literature are also reviewed in this section. Since we are interested in the importance of liquidity constraints for consumption and in whether the proportion of liquidity-constrained consumers has changed over time, we estimate both fixed- and time-varying-coefficient models. In single-explanatory-variable models, the constant and the regression coefficient are obviously correlated. Section 3 reviews also the time-varying properties of the excess-sensitivity parameter with and without time-varying constant and real interest rate. Attention is paid also to the relevance of credit variables recently emphasised in several empirical papers. Finally in section 3.4 we look at the time-varying estimates and consider some plausible explanations for the observed variation in the sensitivity parameter. Concluding remarks are given in section 4.

2 Euler equation approach to consumption

2.1 Consumer choice and growth of consumption

Consumers' intertemporal optimisation determines the conditions under which consumers smooth their consumption over the life cycle. The life-cycle permanent income hypothesis (LC-PIH) consumption model asserts that consumers finance their spending on the basis of a more permanent conception of lifetime (or shorter planning horizon) wealth/income than is afforded by current disposable income. Although permanent and current incomes are in practice correlated for many reasons, it has been suggested that the sensitivity of consumption to income reveals the proportion of liquidity-constrained consumers (King, 1986, p. 72 and Campbell and Mankiw, 1991).

The standard representative consumer utility maximisation problem with finite time horizon, T , can be expressed as maximisation of the objective function (see eg Taylor and Sarno, 1998 or de Bondt, 1999)

$$V_0 = E_0 \sum_t (1 + \rho)^{-t} U(C_t), \quad t = 0, \dots, T \quad (1)$$

subject to the budget constraint represented by the asset evolution equation

$$A_{t+1} = (1 + r_t)(A_t + Y_t - C_t), \quad (2)$$

where A_t is real net assets at the start of period t , which can be regarded as inherited wealth at time $t = 0$ in the intergenerational framework. Labour income is denoted with Y_t . Consumption is usually defined as total non-durable goods and services excluding user costs (Deaton, 1992, p. 10).

The subjective time preference, ρ , which is used to deflate future utilities, is assumed to be non-negative and constant, even though it is likely to vary over time (and hence over age). Since it is unobserved, it is almost always assumed to be constant and equal to the mean of the real interest rate. Capital income is received on wealth and current savings, while the real return can be time varying, r_t , because of a varying nominal interest rate or inflation. Therefore r_t is also stochastic and may follow a random walk process, which means that consumers assume that changes in the real interest rate are always permanent so that the real rate does not return to an 'equilibrium level'. Changes in the real interest rate represent efficient discounting of new information on resources available for consumption. Consumers are assumed to be able to borrow or lend freely at this rate, subject to lifetime budget constraints. The first-order conditions for maximisation of V_0 subject to the budget constraint form the Euler-equation

$$U'(C_t) = (1 + \rho)^{-1} E_t [(1 + r_t) U'(C_{t+1})] \quad (3)$$

Absent further assumptions, this condition has no closed form solution. In order to proceed, we must first assume an explicit form for the utility function. The interest rate will determine how steeply consumption rises with age. The form of the utility function will also affect optimal consumption growth, which will always depends on the real interest rate and subjective time preference (ie the exchange rate between utilities in different periods). Consumption growth will be the steeper, the higher the interest rate and the lower the time rate of preference. These effects follow via the endogenous asset affect, according to which a higher yield is obtained on savings by shifting

consumption to later periods. Impatience in utility simply places more weight on current vs future consumption.⁴

With convex marginal utility, there is always a precautionary motive for saving. And with iso-elastic preferences, there is an infinite penalty for zero consumption; hence consumption is always non-negative. Cochrane (1991, p. 760) notes that, although Euler equations are partial equilibrium propositions, they hold even without any further assumptions about the utility function and thus they can be studied without using fully specified macroeconomic models. Therefore the Euler conditions can be used to study preferences, but when we want to draw further conclusions about consumption, we need to specify properties of utility functions, risk-aversion, the effect of credit on consumption etc. Thus Euler-equation-type optimisation conditions can at best provide a good theoretical starting point; they cannot be regarded to be the final reference eg for successful forecasting models.

2.2 Euler equation models

Hall's (1978) original theory of consumer choice under uncertainty with rational expectations led to widespread testing of its major implication, that of the unforecastability of consumption growth. Random walk properties of consumption changes are conditional on the information set, ie information on past income or any other publicly available information. Even if the real interest rate and subjective time preference are not exactly equal, the RW model could provide a close approximation to the consumption path. It should be remembered that this strong implication about unforecastability of non-durables consumption is based on very restrictive assumptions concerning the entire consumption optimisation framework.⁵

⁴ In the Euler equation, future-period variables are based on expectations, which are subject to change. Thus consumers will also revise their consumption plans as they get older and gain information about divergences eg between expected and actual incomes. However, there is no point in committing to future consumption any sooner than necessary.

⁵ In practice the time series properties of consumption are affected by differences between the consumer's actual planning horizon and data frequency. Using this line of argument we can also approach the question of appropriate consumption planning horizon by asking which consumption time aggregation will maximise the random walk properties of consumption and which definitions of consumption most precisely meet the RW properties.

The list of restrictions includes (see Muellbauer and Lattimore, 1995, p. 237 or Taylor and Sarno, 1998, p. 224)

- a) no credit constraints
- b) non-stochastic additively separable quadratic utility function
- c) the same constant subjective discount rate for each consumer ($\rho_{it} = \rho$, for all i, t)
- d) constant real interest rate ($r_t = r$), which is equal to the subjective discount rate
- e) rational expectations of non-capital income.
- f) no kinks or other non-linearities in the budget constraint
- g) no consumption habits or adjustment costs
- h) coincidence of consumer planning and decision horizons and data frequency.

Since the publication of Hall's (1978) paper, the many theoretical and empirical studies have relaxed the assumptions and tested their relevance. Of course relaxing these restrictions will also affect the income-sensitivity of consumption. Here we concentrate on estimating different variants of models and applying alternative assumptions as to the interest rate in order to determine how sensitive the estimates are to different conditioning instrument variables. Especially with the link to portfolio choice, it is natural to think that r_t is not constant over time. Therefore it is believed that financial conditions have an effect on excess sensitivity. In any case we must make further assumptions in order to proceed to testable empirical models. Further problems arise because different theoretical models can imply similar types of testable empirical models, which are often, unfortunately, observationally equivalent and so cannot be distinguished from each other.

The relationship between consumption and the interest rate can be analysed in greater detail. We can consider the constant-elasticity-of-substitution (CES) utility function $U(C_t) = C_t^{1-\gamma} / (1-\gamma)$; ($\gamma > 0$). The Euler condition (3) can be then written as

$$E_t \left[(C_{t+1} / C_t)^{-\gamma} r_t \right] = 1 / (1 + \rho).$$

If the joint vector of consumption growth, $\Delta c_t = \log(C_{t+1}/C_t)$, and log of the real interest rate (r_t) are assumed to have stationary log-normal distributions, their distributions can be characterised by the mean μ_t

and variance σ^2 .⁶ Assuming log-normality of these variables, the Euler condition can be further manipulated to obtain (see Hahm, 1998, p. 296 and Hansen and Singleton, 1983, p. 253)

$$E_t \Delta c_{t+1} = -\rho / \gamma + \gamma / 2\sigma^2 + 1/\gamma E_t r_{t+1} \quad (4)$$

where σ^2 is the conditional variance of $(\Delta c_{t+1} - 1/\gamma r_{t+1})$. It should also be noted that with a CES function the coefficient for the real interest rate should be positive. Another commonly used parameterisation of the utility function is the iso-elastic utility function

$$U(C_t) = \delta(\delta - 1)^{-1} C_t^{(\delta-1)/\delta} \quad (5)$$

where $\delta (= 1/\gamma)$ is the intertemporal elasticity of substitution, which depends inversely on the curvature of the utility function. This elasticity of intertemporal substitution indicates the degree to which consumers are willing to defer current consumption in order to obtain a higher expected real return. The curvature of the utility function is important since it defines the consumer's attitude to risk. In the case of an iso-elastic utility function, the log of consumption and the real interest rate are assumed to be jointly normally distributed and the Euler equation (3) is reduced to

$$E_t \Delta c_{t+1} = \delta(E_t r_t - \rho) + \text{Var}_t(\Delta c_{t+1} - \delta r_t) / 2\delta \quad (6)$$

where $\text{Var}_t(\Delta c_{t+1} - \delta r_t)$ denotes the conditional variance. Even if consumption and the interest rate are not normally distributed, this formulation approximates the true Euler equation.

The conditional variance term is in practice absorbed in the intercept term, but we can either test the ARCH properties of the residual series or the growth of consumption directly (see Taylor and Sarno, 1998, p. 223). De Bondt (1999, p. 4) notes also that the 'constant' in this model is not strictly constant since it includes the conditional variance. Therefore we have additional grounds for assuming that the constant is also time varying. If this conditional term is found to be sufficiently stable, based eg on testing for the ARCH effect, it can be assumed constant, in which case equation (6) can be simplified as

⁶ See Hansen and Singleton (1983, p. 253).

$$\Delta c_{t+1} = \mu + \delta r_t + \varepsilon_{t+1} \quad (7)$$

where ε_{t+1} is not correlated with any lagged variables but may be correlated with r_t . Thus we have arrived at a random walk model with constant drift. The constancy of the drift can also be evaluated via time-varying parameter estimation. In the testing section of the paper we examine how relaxing this assumption affects the results. A similar estimable model is derived from equation (4), now with $\alpha = -\rho/\gamma + \gamma/2\sigma^2$, $\beta = 1/\gamma$ (Hahm, 1998, p. 297).

In a simple Euler-equation-type specification, the coefficient of the real interest rate can also be a time-varying parameter. This parameter would probably be a function of demographic factors such as age, education, family composition and employment status, which change over a person's lifetime. The coefficient of the real interest rate is the elasticity of intertemporal substitution, which measures the extent to which consumers are willing to trade consumption for expected real return. With an additive time-separable utility function, the intertemporal elasticity of substitution is inversely related to relative risk aversion.⁷ Model (7) implies that an increase in the real interest rate will induce an increase in future consumption (assuming $\delta > 0$) by the amount future consumption is increased by higher interest rates and how much of future consumption is transferred to the present. The magnitude of the elasticity of intertemporal substitution determines this ratio.

Following Hansen and Singleton (1983) and assuming consumption and the real interest rate to be jointly conditionally lognormal and homoscedastic, the basic Euler equation (3) simplifies to

$$E_t \Delta c_{t+1} = \mu + \delta E_t r_{t+1} \quad (8)$$

where μ is constant and lower-case variable names indicate logs. Now simply follow Campbell and Mankiw (1989, 1991) in assuming that consumers can be divided into two groups, viz those whose behaviour is described by the permanent income hypothesis (equation 7 above) and those who may be liquidity-constrained 'rule of thumb' consumers who spend their entire current income and are described by

⁷ Hahm (1998) argues that estimates of intertemporal elasticity of substitution could be around 0.3 for US data. Numerous studies have asserted that this elasticity could be close to zero, but this would imply (implausible) infinite relative risk aversion.

$$E_t \Delta c_{t+1} = (1 - \beta_t) \mu + \beta_t \Delta y_{t+1} + (1 - \beta_t) \delta r_{t+1} \quad (9)$$

This equation includes an additional term for the disposable income of the liquidity-constrained consumers, which gives rise to a time-varying drift and an interest rate effect for the life-cycle consumers. Arranging terms, we express the model as

$$E_t \Delta c_{t+1} = \beta_t \Delta y_{t+1} + (1 - \beta_t) [\mu + \delta E_t r_{t+1}]$$

which can be written in the empirically testable form

$$\Delta c_{t+1} = (1 - \beta_t) \mu + \beta_t \Delta y_{t+1} + (1 - \beta_t) \delta E_t r_{t+1} + (1 - \beta_t) \varepsilon_{t+1} \quad (10)$$

where the parameter β_t gives the proportion of income that the liquidity-constrained consumers receive directly in period t . Equation (10) reveals already that there is a negative relationship between the constant and the disposable income parameter. After having estimated this model and obtaining estimates for β_t , we can get estimates of the drift for life-cycle consumers, $(1 - \beta_t) \mu$, and the intertemporal elasticity of substitution, δ . If δ is close to zero, as suggested, consumption growth evolves as a weighted sum of the change in disposable income (for liquidity-constrained households) and a random walk with drift representing the consumption behaviour of life-cycle consumers. Moreover, the coefficient of the real interest rate (or residual) cannot be tested without further assumptions as to intertemporal elasticity of substitution, δ . If the expected real interest rate is constant or $\delta=0$, the equation collapses to a single explanatory variable model, with negatively correlated constant and regression coefficient.

In the original Campbell and Mankiw (1989) paper, β_t was estimated as a fixed constant. The null hypothesis, confirming the PIH, would be simply $\beta_t = 0$. This model cannot be tested by OLS, since positive correlation between ε_{t+1} and Δy_{t+1} (which is possible) would render estimates of β upward biased and inconsistent. Thus IV estimation would be required. Originally Campbell and Mankiw (1989) used lagged instruments to estimate the response of consumption, whereas Hall (1978) used also contemporaneous instruments. Lagged instruments can be justified if agents, when forming expectations, do not have information on macro-variables such as income. Since β_t can be time varying, the proportion of liquidity-constrained consumers is allowed to change over time, depending on credit conditions, interest rate, business cycle conditions etc. Taylor and Sarno (1998) and Girardin, Sarno and Taylor (2000)

construct an index of financial liberalisation, k_t , which is expected to approximate β_t .

Gigardin, Sarno and Taylor (2000, p. 363) propose and estimate a time-varying model

$$\Delta c_{t+1} = (1 - \beta_t)\mu + \beta_t \Delta y_{t+1} + (1 - \beta_t)\delta r_{t+1} + (1 - \beta_t)\epsilon_{t+1} \quad (11)$$

with a logistic transformation

$$\beta_t = \phi_0 + [2 \exp(-\phi_1 k_t) / (1 + \exp(-\phi_1 k_t))]$$

applied to β_t , which forces this parameter to vary in the closed interval $[0, 1]$. This is reasonable since the income share of the liquidity-constrained consumers cannot exceed unity. The parameter for financial deregulation, k_t , was approximated by the ratio of household total credit to disposable income, ie the indebtedness ratio. They estimated this model with nonlinear IV-estimation with twice-lagged instruments.

Bacchetta and Gerlach (1997) and Taylor and Sarno (1998) examine the time-varying features of the excess-sensitivity parameter, β , using a linear specification and assuming β_t to be strictly between zero and one.⁸ The explanatory variables in the testing equation included both random and fixed variables such as real disposable income, real housing loans, real consumer credit and the interest rate wedge between lending and borrowing rates. They suggest that credit growth and the interest rate wedge can be used to measure the tightness of credit conditions. Their consumption measure included non-durable goods and services, and the instruments were lagged by at least two periods, due to reporting lags that prevent consumers from using current information at time $t-1$ when they are forming their expectations. Otherwise the time averaging error of a random walk in consumption implies that the error term will be an MA(1) process.⁹

A recent paper by Ludvigson (1999) suggests that credit growth should also be included among the explanatory factors for life-cycle savers' consumption, which is smoothed via debt (and amortisations) even over temporary changes in income. The expected growth of

⁸ It should be noted that these papers also assume the constant to be fixed, whereas here we also consider the case where the constant is allowed to be time varying.

⁹ First-order serial correlation in the error term may lead to inconsistent estimates, if only one-period lagged instruments are used. In studies like Campbell and Mankiw (1989) shocks in tastes were assumed to give rise to an additional MA(1) process in the error term, which affects the period of time for which error terms should be uncorrelated.

credit is used as an additional explanatory variable in the model. Credit growth is expected to correlate negatively with income growth for life-cycle households. Liquidity-constrained consumers' consumption is unaffected since they simply use their entire disposable incomes for consumption. The reason for this is that the borrowing limit varies with current income, since many banks tie their lending to current earnings and creditors do not have any better information on permanent income. In principle there are two channels by which credit supply can affect consumption: resources available for consumption (ie disposable income and available credit, which varies with the direct credit supply) and current credit shocks that affect future credit constraints of liquidity-constrained consumers.

$$\begin{aligned} \Delta c_{t+1} = & (1-\beta_t)\mu + \beta_t\Delta y_{t+1} + (1-\beta_t)\delta E_t r_{t+1} + \gamma(1-\beta_t)E_t \Delta d_{t+1} \\ & + (1-\beta_t)\varepsilon_{t+1} \end{aligned} \quad (12)$$

where d_t refers to net debt used to finance non-durables consumption.¹⁰ Ludvigson (1999, p. 444) also discusses the financial deregulation-type effect, which discretely relaxes binding liquidity constraints and causes a quick jump in debt to a new and higher level and hence an increase in consumption. In principle this should make consumption less sensitive to expected changes in current income, since the consumption increase is due to higher debt. Thus it may be useful to study the time-varying paths of both disposable income and credit supply to see which of these explanatory variables increases eg during a period of financial deregulation. The null hypothesis is that a higher credit ceiling will lead to a reduction in the sensitivity of consumption growth to expected income and increased use of credit. This change will be related to the disposable income of liquidity-constrained consumers. It should be noted that in practice the relaxing of liquidity constraints affects primarily the marginal propensity to consume durables, whereas consumption of non-durable goods and services is less affected.

A related approach has been used by De Bondt (1999), who suggested that besides life-cycle savers and liquidity-constrained consumers, one could consider a separate group of consumers who have access to external financing in addition to current income. This group can be considered to be quite similar to liquidity-constrained consumers. It is proposed that the external finance (credit supply) term

¹⁰ These equations must be estimated by GMM because of the expectations.

could include a financial accelerator term associated with the business cycle. Thus the estimated model can be written as

$$\begin{aligned} \Delta c_{t+1} = & (1 - \beta_{1t} - \beta_{2t})\mu + (\beta_{1t} + \beta_{2t})\Delta y_{t+1} + (1 - \beta_{1t} - \beta_{2t})\delta E_t r_{t+1} \\ & + \alpha_1 \Delta EFP_t + \alpha_2 \Delta EFP_t^* BC_t + (1 - \beta_{1t} - \beta_{2t})\epsilon_{t+1} \end{aligned} \quad (13)$$

where EFP is the external finance premium (interest rate margin) and BC a business cycle indicator. The last term approximates the financial accelerator effect. Alternative versions of this basic model can be estimated applying on different parameter restrictions based on whether liquidity-constrained consumers spend only their disposable income and whether the EFP effect varies over the business cycle (nonzero α_2).¹¹

In this approach de Bondt (1999) concentrates on explaining the β -estimates. The variable that describes the effectivity of liquidity constraints, ie the external finance premium (EFP), indicates variations in the relative price of credit. The Campbell and Mankiw study focused more on restrictions based on availability of credit, whereas the time-varying- β approach tends to give more weight to the price of credit in financing consumption. De Bondt (1999) uses the interest rate margin between deposits and bank lending as a proxy for the wedge between external and internal cost of financing consumption. This effect can be separated from the pure volume liquidity constraint, but it can also be analysed with it simultaneously, since the two cannot be entirely separated, being essentially mere reflections of the disequilibrium observed in any market. In most markets, such adjustment usually involves changes in both quantities and prices, rather than in just one of these.

It should be noted that the above testable equations have been written with time-varying constants, whereas the original studies by Campbell and Mankiw (1991), Bacchetta and Gerlach (1997), de Bondt (1999) and Ludvigson (1999) all treat the constant as fixed. It is possible to free the constant to be time varying and to view it as including all the other time-varying effects. Obviously we must allow the real interest rate to have a time-varying coefficient. In addition we may also allow other conditioning variables, eg credit, to have time-varying parameters so as to enable testing of the reactions of the disposable income parameter. In the empirical part of the paper we

¹¹ De Bondt clearly notes an interrelationship between the excess-sensitivity parameter and the constant, but, because he performs only fixed parameter estimations, he treats the constant as fixed.

estimate versions of the proposed model based on different assumptions as to the time-varying parameters.

3 Empirical results from Euler equation tests

3.1 Excess-sensitivity testing

Excess-sensitivity testing for consumption tries to clear up the puzzling sensitivity of current consumption changes to lagged information on disposable income. Current consumption should already incorporate all the relevant information in disposable income and other factors affecting permanent income as consumption should be proportional to permanent income. Permanent income is unknown as such in the short-run, but can be approximated as a linear combination of disposable income and net wealth. In the short run, stationary variables such as the real interest rate will affect the permanent income estimate.

Because of information limitations empirical testing of excess sensitivity should be based on income information lagged by two quarters. On the aggregate level, information lagged by one quarter is not readily available when current consumption decisions are made. While this is not the case for nominal interest rates, it does apply to the real interest rate due to the one-quarter lag in publishing the inflation rate.¹² Two-quarter information lags are also conceivable in respect of certain other real economic and financial variables.

In direct testing of Euler equations one need not make any particular ad hoc assumptions about the error term, because it is directly the forecast error in consumption. It is necessary only to assume orthogonality of the error term (forecast error) and explanatory variable distribution (past information set). Forecast errors need fulfil only the standard weak stationary properties. Special attention could be paid to the relationship between consumption, credit and wealth. The special role of wealth is understandable since net wealth clearly can be used to finance or smooth consumption.¹³

¹² Nominal interest rates are available daily, and these in fact measure the cost of next-period consumption in terms of an alternative investment such as a deposit.

¹³ In particular it would be interesting to see whether conditioning on different types of wealth, ie different types of financial assets, debt or real estate wealth, leads to different impacts on excess-sensitivity tests.

3.2 Results of Euler equation estimations

Tables 1 and 2 present the results of the fixed-parameter excess-sensitivity regressions. In addition to the IV estimates, we present GMM estimates in order to compare consistent estimators without imposing normality or other full density assumptions about the distribution of the residual. Another reason for using GMM is that we can include variables that entail expectations (in the absence of actual expectations data) as regressors (Hamilton, 1994, p. 422–424).

In these models the instruments in the IV and GMM estimations have been lagged in most cases by 2 or 3, due to reporting lags, MA(1) errors in consumption measurement of durables, and errors in the measurement of consumption.¹⁴ Modelling of expected variables should in principle be done using instrumental methods. However, for the purpose of comparison, we also report OLS estimates. As regards the market interest rate, we also used one lag, since such information is surely available to consumers before consumption decisions are made. Table 1 reveals that the real interest rate is not a very efficient instrument for consumption whereas disposable income clearly seems to be.

The diagnostic tests for the IV estimations clearly indicate that the estimated models do not produce data-congruent representations of consumption determination. The residuals are seriously serially correlated in all the equations and autoregressive heteroscedastic in most of them. The null hypothesis of no autocorrelation was rejected with very high probability (p-values) in every estimated equation. Thus the estimated models cannot be regarded as sufficient dynamic representations of non-durables consumption. Autocorrelation can be reduced by switching from annual to quarterly differences, but the model would still not pass all the relevant diagnostic tests. The specification of the consumption model presented in tables 1 and 2 seems to be overly simplified.

In all the models the excess-sensitivity parameter is significantly greater than zero and hence the simple PIH implication of the null hypothesis, $\beta = 0$, is rejected. Estimates of the excess-sensitivity

¹⁴ Based on experiments with different lags, it seems that the results concerning excess-sensitivity parameter are not particularly sensitive to lag length unless one uses the contemporaneous changes of the explanatory variables. There was also serious residual correlation when one-period lagged instruments were used.

coefficient are clustered quite tightly around 0.5.¹⁵ However, it should be noted that in a single explanatory variable regression the excess-sensitivity parameter and the constant are not estimated independently and that they are negatively correlated. Therefore a large constant is associated with a small sensitivity parameter and vice versa.¹⁶ In the estimated models with only lags of real interest rate as instrument, the correlation between the constant and β -coefficient is very high, while in other models it is around -0.5 . In table 1 the constant terms are concentrated around 1.1%, and thus assuming that β is 0.5 implies that the drift of life-cycle consumers is 2.2%. According to the GMM estimations in table 2, the annual drift could be even higher, ie around 3%. But from table 2 we also notice that in general the GMM estimates for β are lower in all cases other than those in which only lagged values of disposable income were used as instruments.

¹⁵ Our estimates for disposable income are significantly higher than those presented in Taylor and Sarno (1998) for the United Kingdom, which were about 0.2–0.25, or in Bacchetta and Gerlach (1997) for four countries (US, UK, France and Canada), which were in the range .244–.349, or even in Agell and Berg (1996) for Sweden, which were in the range .27–.35. Gigardin, Sarno and Taylor (2000) estimated the proportion of liquidity-constrained consumers to be on average 0.36 for France in 1970–1993, but according their time-varying estimations this ratio has converged to zero at the end of the estimation period. However, a rather high ratio of about one-half was originally found by Campbell and Mankiw (1989) for US data, and Cushing (1992) and Ludvigson (1999) found ratios of around 30–40%. De Bondt (1999) also reports a high proportion of liquidity-constrained consumers for the United Kingdom and Belgium. Brunila's (1997) GMM income parameter estimates using Finnish data from 1960–1993 were .36 and .25, depending on the specification.

¹⁶ The excess-sensitivity parameter in this regression is negatively correlated with the constant, because in a one explanatory variable regression model $\text{Cov}(\alpha, \beta) = -\bar{Y} \text{Var}(\beta)$, where \bar{Y} is the mean of disposable income.

Table 1. **Excess-sensitivity regressions,
1980Q4 – 1998Q4**
Instrumental method estimation

$$d4lnoncd = \alpha + \beta * d4lryd$$

	Constant β -param.		Instruments (lags)		R ²	DW	Residual diagnostics, p-values		Instrument equation 1. stage		Parameter correlation		Wald test for $\beta=0$	
							AR(5)	ARCH(4)	R ²	Corr(α, β)	p-values			
OLS	1.168 (5.01)	0.497 (8.52)			0.51	0.45	0.0000	0.0000			-0.4794	.0000		
IV	0.971 (2.19)	0.602 (2.88)	rs(1), rs(2)		0.07	0.49	0.0000	0.0071	0.08		-0.8872	.0287		
IV	0.985 2.694	0.594 3.689	rrs(2), rrs(3)		0.11	0.49	0.0000	0.006	0.14		-0.8282	.0071		
IV	1.237 (5.15)	0.46 (6.60)	d4lryd(2,3)		0.35	0.31	0.0000	0.0790	0.70		-0.5490	.0000		
IV	1.173 (4.92)	0.494 (7.22)	d4lryd(2,3), d4lrhoulo(2,3)		0.49	0.45	0.0000	0.0000	0.73		-0.5384	.0000		
IV	1.071 (4.52)	0.549 (8.29)	d4lryd(2,3), d4lrhoulo(2,3) spread(2,3)		0.67	0.47	0.0000	0.0000	0.78		-0.5229	.0000		
IV	1.116 (4.74)	0.525 (8.07)	d4lryd(2,3), d4lrhoulo(2,3) spread(0-3), d4lrconcre(2,3)		0.68	0.46	0.0000	0.0000	0.80		-0.5245	.0000		
IV	1.31 (5.45)	0.422 (6.19)	d4lryd(2,3), ect(4), ur(2)		0.75	0.41	0.0000	0.0000	0.75		-0.5316	.0000		
IV	1.165 (5.04)	0.499 (8.21)	d4lryd(2,3), d4lnoncd(1) ect(4), rrs(2,3) d4lrhoulo(2,3), margin(2,3)		0.92	0.45	0.0000	0.0000	0.89		-0.5034	.0000		

*) t-values are presented in parenthesis below the constant and β -parameter estimates. The dependent variable is total consumption excl. durables purchases, ie consumption of non-durable goods, semi-durables and services. For IV estimation, the diagnostic F-test was performed as an autoregressive test (AR(5) test) for residual serial correlation with 5 lags. The ARCH(4) tests are calculated for residual autoregressive heteroscedasticity with 4 lags. The null hypothesis is no ARCH, which would be rejected if the test statistic is too high. This test is done by regressing squared residuals on a constant and lagged squared residuals. Parameter correlations are calculated on the basis of parameter covariance divided by the cross-product of parameter standard deviations, ie $Corr(\alpha, \beta) = Cov(\alpha, \beta) / \sqrt{Var(\alpha) * Var(\beta)}$.

Variables:

d4lnoncd = Annual change in non-durable consumption, %
d4lryd = Annual change in real disposable income, %
rs = Short-term market interest rate (3 month Helibor rate), %
rrs = Real short-term market interest rate, %
d4lrhoulo = Annual change in real housing loans, %
d4lrconcre = Annual change in real consumer credit, %
spread = Interest rate spread between long and short-term rates, %
margin = Interest rate margin between lending and deposit rates, %
ect = Error-correction term between income and consumption in cointegration equation
ur = Unemployment rate, %

Table 2. **Excess-sensitivity regressions,
1980Q4 – 1998Q4**
Generalised moment method estimation

$$d4lnoncd = \alpha + \beta * d4lryd$$

Estimation Method	Constant	β -parameter	Instruments (lags)	R ²	DW	Wald tests for restriction	Parameter correlation
						$\beta=0$ p-values	Corr(α, β)
OLS	1.168 (5.01)	0.497 (8.52)		0.51	0.450	0.0000	-0.4794
GMM	1.192 (2.32)	0.504 (2.73)	rs(1), rs(2)	0.50	0.453	0.0064	-0.7714
GMM	1.549 (2.78)	0.405 (2.01)	rrs(2), rrs(3)	0.47	0.394	0.0445	-0.8737
GMM	1.374 (3.14)	0.42 (4.77)	d4lryd(2,3)	0.49	0.407	0.0000	-0.6876
GMM	1.579 (3.88)	0.457 (4.86)	d4lryd(2,3), d4lrhoulo(2,3)	0.48	0.413	0.0000	-0.1741
GMM	1.51 (4.41)	0.464 (6.36)	d4lryd(2,3), d4lrhoulo(2,3) spread(2,3)	0.48	0.421	0.0000	-0.6251
GMM	1.539 (5.88)	0.417 (6.17)	d4lryd(2,3), d4lrhoulo(2,3) spread(0-3), d4lrconcre(2,3)	0.48	0.400	0.0000	-0.5766
GMM	1.510 (3.60)	0.492 (5.32)	d4lryd(2,3), ect(4), ur(2)	0.49	0.431	0.0000	-0.7351
GMM	1.856 (6.92)	0.459 (7.58)	d4lryd(2,3), d4lnoncd(1) ect(4), rrs(2,3) d4lrhoulo(2,3), margin(2,3)	0.44	0.380	0.0000	-0.6719

Variables:

d4lnoncd = Annual change in non-durables consumption, %
d4lryd = Annual change in real disposable income, %
rs = Short-term market interest rate (3 month Helibor rate), %
rrs = Real short-term market interest rate, %
d4lrhoulo = Annual change in real housing loans, %
d4lrconcre = Annual change in real consumer credit, %
spread = Interest rate spread between long and short-term rates, %
margin = Interest rate margin between lending and deposit rates, %
ect = Error-correction term between income and consumption in cointegration equation
ur = Unemployment rate, %

However, these estimates are average estimates, whereas we are more interested in whether the excess-sensitivity parameter changes over time and whether it depends on economic conditions such as credit availability and cost. We can start examining the variability of the excess-sensitivity parameter by performing recursive or rolling estimations of the model. We performed the tests for non-durables consumption. Figure 1 presents the rolling regression estimates. The average excess-sensitivity parameters are 0.60 for total consumption and 0.50 non-durables consumption for the period 1980Q4–1998Q4. Overall, the parameter estimates are quite similar for the two consumption measures. The excess-sensitivity parameter estimate is however smaller for non-durables than for total consumption. This result seems reasonable since it is natural that liquidity constraints would be more stringent for purchases of durable goods included in total consumption. The excess-sensitivity parameter also clearly correlates with the saving rate (figure 2).

The negative correlation between the constant and the β -parameter was very clear for the recession period of the 1990s, for which the excess-sensitivity parameter reaches about 0.9 in the total consumption equation and 0.75 in the non-durables consumption model. The time-path of the excess-sensitivity parameter remains roughly the same if the interest rate variables are included as additional exogenous variables. There is a slight decline in the β -parameter in the 1980s, period of overheating, and a more significant rise during the recession years of 1991–1994.¹⁷

Table 3 gives test results for the basic excess-sensitivity model augmented by the real interest rate. The β -estimates are generally only a slightly smaller than the GMM estimates in table 2. Results for β -estimates are similar except that the real interest rate gets a positive coefficient that is not significant and the efficiency of the instruments is higher. The correlation between real interest rate coefficient and constant is highly negative, but the β -term and interest rate coefficient are correlated only in half of the models. Therefore, based on the β and δ parameter estimates, there is not much evidence of compensation between real interest rate effects and the income term.

It is common to use OLS regression in testing for excess sensitivity, but it should be remembered that OLS produces a BLUE

¹⁷ However, Attanasio and Weber (1993) show that estimates of the elasticity of intertemporal substitution are very different when estimated with aggregate data and average cohort data. The aggregate data estimates are substantially lower. They also assert that excess-sensitivity tests can be useful in finding the sources of aggregation bias.

only if the errors are independently and identically distributed. If the errors are not normally distributed, other unbiased estimators may be preferred. OLS gives heavy weight to large outliers and if outliers are not taken into account explicitly they will cause bias in the excess-sensitivity parameters. One way to reduce the weight of outliers is to use the L1 norm for errors and so minimise absolute deviations.¹⁸ Therefore it is worthwhile to check whether the estimates of excess sensitivity change when this type of non-normality is assumed.

One way to do this is to use a robust statistical technique that relaxes the normality assumption and is less outlier-sensitive than is the squared-error weighting of the OLS. The main advantage is that with robust methods less weight is given to outliers, but other types of distributional problems related to normality may arise. However, since the results from these performed median regressions were close to the OLS estimates, ie generally between .46 and .5, it seems likely that the assumption of normality in the error terms is not highly misleading.

In connection with excess sensitivity, it would be of interest to know how severe the liquidity constraints are. In practice there cannot be a single measure for liquidity constraints, but Taylor and Sarno (1998) propose a unified index derived from a combination of different indicators. Considering financial regulation and its effects on liquidity constraints, King (1986) already separated credit rationing from the wedge between borrowing and lending rates. In some empirical applications, the unemployment rate has been used, since for most households the primary income risk is associated with unemployment. The margin between borrowing and lending interest rates has also been used as a direct proxy for liquidity constraints. In practice there may also be structural-type constraints, eg on individuals who have defaulted on payments or have atypical jobs.

The estimated fixed parameter of the model by Gigardin et al (2000), with logistic transformation of the excess-sensitivity parameter, turns out to be somewhat different than that of the equation (11) model for France by Gigardin, Sarno and Taylor. The mean of the excess-sensitivity parameter was 0.44 and was clearly significantly different from zero. The drift for consumption was 3.37, but the parameter ϕ_1 was not significantly different from zero, which suggests a constant β . Another feature worth mentioning was that while the estimated elasticity of substitution was not significant for France, it

¹⁸ Least absolute value models are conditional to the 0.5 quantile and hence are also called median regressions.

turned out (contrary to expectations) to be negative and significant for Finnish data.

Comparison between the estimated models based on specifications (12) and (13) showed that all the explanatory variables were significant, but in the Bondt-type-model real interest rate on new household loans had a positive sign. In the Bondt-type specification (13), the interaction term between external finance premium and business cycle (here GDP) was highly significant, while the change in real consumer credit also proved to be highly significant for the Ludvigson-type specification (12). Otherwise the diagnostics of the Bondt-type model were better, but both models suffered from residual autocorrelation. Both models failed to encompass the other, but the test statistics were somewhat stronger for the Ludvigson-type specification (see Appendix 1).

Table 3. **GMM estimates of interest-rate-augmented Euler equation, 1980Q4 – 1998Q4**
Generalised Method of Moments

$$d4lnoncd = \alpha + \beta * d4lryd + \delta rrs$$

	Constant	β -parameter	δ	Instruments (lags used)	R ²	DW	$\beta=0$	Corr(α,β)	Corr(α,δ)	Corr(β,δ)
OLS	1.81 (3.98)	0.484 (8.32)	-0.111 (-1.62)		0.52	0.45	.0000	-0.35	-0.87	0.14
GMM	1.887 (1.87)	0.512 (2.81)	-0.134 (-0.81)	rs(1), rs(2)	0.51	0.46	0.005	-0.28	-0.87	-0.11
GMM	2.269 (2.43)	0.384 (1.81)	-0.160 (-1.11)	rrs(2), rrs(3)	0.49	0.39	0.0696	-0.62	-0.79	0.13
GMM	12.438 (0.67)	0.375 (1.13)	-1.993 (-0.61)	d4lryd(2,3)	-4.62	0.15	0.258	-0.66	-1.00	0.64
GMM	1.599 (0.99)	0.462 (4.92)	-0.007 (-.03)	d4lryd(2,3), d4lrhoulo(2,3)	0.47	0.42	.0000	-0.22	-0.97	0.03
GMM	1.134 (1.90)	0.411 (4.80)	0.102 (.91)	d4lryd(2,3), d4lrhoulo(2,3) spread(2,3)	0.40	0.37	.0000	0.04	-0.81	-0.49
GMM	1.459 (3.18)	0.319 (5.17)	0.116 (1.20)	d4lryd(2,3), d4lrhoulo(2,3) spread(0-3) d4lrconcre(2,3)	0.29	0.30	.0000	0.44	-0.87	-0.66
GMM	0.446 (0.46)	0.375 (3.61)	0.237 (1.67)	d4lryd(2,3), ect(4), ur(2)	0.26	0.31	.0000	-0.53	-0.85	0.11
GMM	0.837 (2.71)	0.298 (5.19)	0.211 (3.43)	d4lryd(2,3), d4lnoncd(1) ect(4), rrs(2,3) d4lrhoulo(2,3), margin(2,3)	0.21	0.28	.0000	-0.01	-0.72	-0.55

Variables:

d4lnoncd = Annual change in non-durable consumption, %
d4lryd = Annual change in real disposable income, %
rs = Short-term market interest rate (3-month Helibor), %
rrs = Real short-term market interest rate, %
d4lrhoulo = Annual change in real housing loans, %
d4lrconcre = Annual change in real consumer credit, %
spread = Interest rate spread between long and short-term rates, %
margin = Interest rate margin between lending and deposit rates, %
ect = Error-correction term between income and consumption in cointegration equation
ur = Unemployment rate, %

Figure 1.

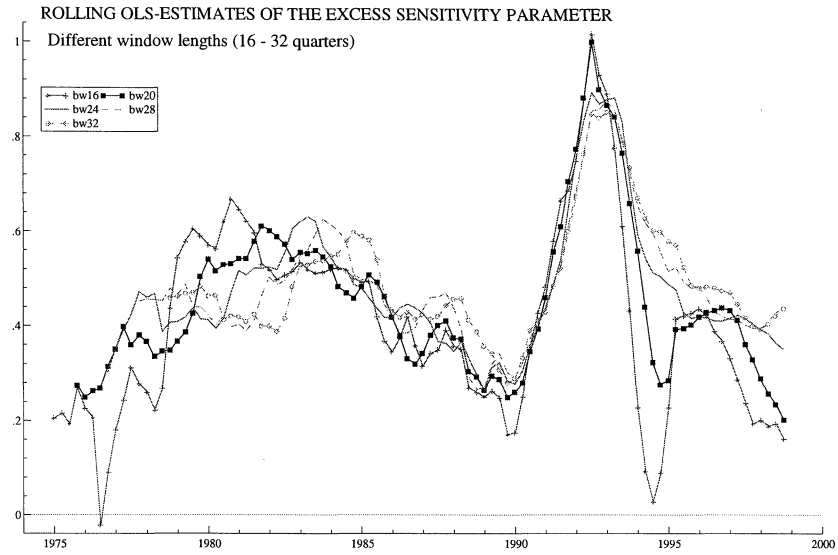
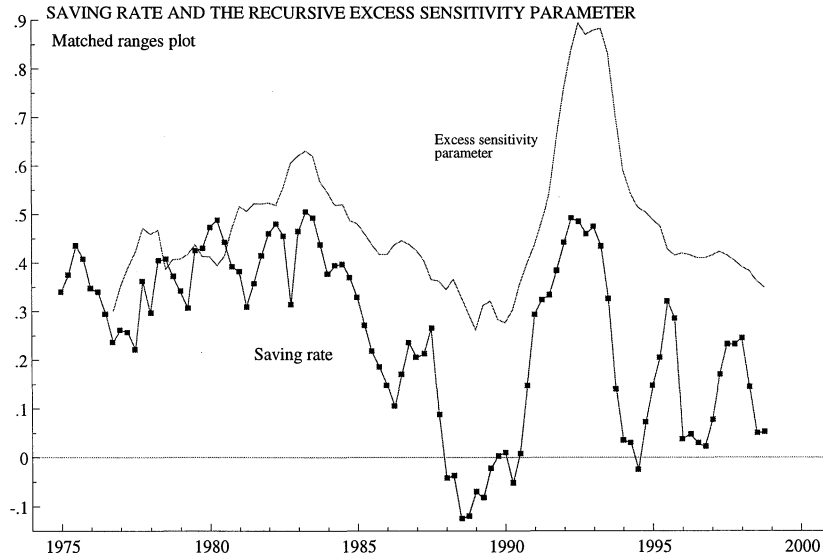


Figure 2.



3.3 Time varying estimation of the consumption function

It is quite plausible that the excess-sensitivity parameter would vary eg with the degree of financial market regulation, income uncertainty and household indebtedness. It is therefore unlikely that financial market deregulation would have had no effect on debt financing of consumption. Therefore a state space representation and Kalman filter estimation can be used to assess the unobserved and most likely time-varying excess-sensitivity parameter. In fixed parameter estimation, one should at least check on the stability of the parameter. An obvious preference for this choice rests on the presumption that liquidity constraints are likely to be related to cyclical movements in real economic activity that induce variability in savings and subsequently affect bank lending.

The estimation and testing of the time-varying parameters is performed by assuming that the β -coefficient follows a random walk, as there is no self-evident theory to justify a different assumption. Random walk updates for parameters could be a natural choice when there are many separate shocks affecting income (expectations) that are not independent of the noise in observations of consumption. So the main reason for selecting a time-varying parameter is that the state contains noisy measurement. The random walk would imply that shocks to the random coefficients persist forever. This option enables one to study the robustness of the excess-sensitivity parameter estimates. The state summarises information on past and present vis-à-vis the consumption generating process. Additional exogenous regressors with fixed coefficients can be taken into account, as can variables for which one would like to have random coefficients (state variables). Such time-varying parameter methods can be used to study the stability of the excess-sensitivity model.¹⁹

We also tried to explicate more precisely the effect of wealth on consumption changes. A good starting point is simply to consider the different sources of consumption financing. The primary source of financing non-durables consumption is current disposable net income,

¹⁹ The state space formulation assumes that error-terms of the measurement and parameter transition equation are uncorrelated with each other in all time periods. This may not necessarily hold for income parameter revisions and excess sensitivity, eg error-terms may have some common ground.

and a secondary source is previously accumulated financial savings.²⁰ Borrowing from the financial markets is occasionally used with housing wealth as collateral, and when labour income has finally declined more permanently real estate wealth must be realised. External sources of financing non-durables consumption include bank lending, which is the market solution for handling this sort of uncertainty. Housing purchases can be financed by bank loans or mortgages from building companies, but they usually require some initial savings and an amortisation plan that may restrict or even crowd out other forms of consumption. Durables purchases can be financed also by other forms of credit eg consumer credit or car finance loans. However, there is no absolute rule or usual pattern indicating that internal sources of finance (eg saving deposits) are always used before resorting to external sources of finance (debt), even though the 'shadow' financing cost of the former is usually lower.

3.4 Empirical results for time-varying estimations

The results for the time-varying estimations were, in contrast to those for the fixed parameter estimations, sensitive for the assumptions as to which variables were allowed to change over time. Comparison of results for non-durables vs total consumption shows that including durables in the model leads on average to higher sensitivity parameters, which again accords with the presumption that durables purchases are much more sensitive to financing opportunities and households' liquidity. Durables consumption has also much larger variance over business cycles than non-durables consumption. However, the hump in the parameter path in the early-1990s recession is more pronounced for non-durables consumption.

The standard procedure has been to allow only disposable income to vary along time (eg Bacchetta and Gerlach 1997). As we have seen in number of specifications real interest may also have time-varying effect on consumption. In addition, if we think that in a simple excess-sensitivity equation various other variables affecting consumption are in fact included into the constant, we may also allow the constant to be

²⁰ Mankiw (2000) argues that possibly close to 30% of US households have zero or negative wealth and so many households lack the financial wealth for smoothing consumption.

time varying as well. If constant is not fixed but is allowed to follow random walk, we get sometimes very different results.

Figure 3 compares the path of the time-varying sensitivity parameter without additional explanatory variables to models with fixed and time-varying constant. The income parameter in the fixed constant model is greater than 0.6 at the start of the 1980s but declines slowly during the decade. Then it increases in the 1990s recession hump, after which it steadily declines up to the end of the period, except for the very end of 1998. The minimum proportion of liquidity-constrained consumers is about 0.3. When the constant is allowed to vary over time, it generally smoothes the time path of the sensitivity parameter, though the short-term variation increases. The recession hump vanishes, and the overall picture tells us that liquidity constraints associated with disposable income have been steadily declining. Over the time period 1980–1998 the excess-sensitivity parameter declines from 0.5 to around 0.1.

In addition when real lending rate is modelled as time-varying parameter, the whole picture of the time path of excess-sensitivity changes again, reaching sometimes even negative values. The same happened with real short-term market interest rate. With fixed constant real interest rate and disposable income parameters are negatively correlated, but there is a distinctive declining trend in the excess-sensitivity parameter up to the very end of 1998. In Finland the indebtedness ratio was rapidly increasing after financial deregulation started in the end of 1980s, but it also started to decrease already during the recession (figure 4). Indebtedness ratio began to increase after 1997 due to heavy lending in housing loans. It is very likely that size of excess-sensitivity parameter is related to indebtedness ratio.²¹

Nevertheless we observed that the time-varying estimates were highly sensitive to which additional variables were used in the estimation (eg the wealth and credit variables) and to which variables were assumed to have time-varying effects. Assuming the constant (consumption drift) to be fixed has been commonplace in the literature, but if we are unsure about the conditioning variables, we can certainly check the sensitivity of the results by a using time-varying constant as well. Consumption drift may also vary over time, due eg to changes in income growth and productivity growth.

As a final exercise, we compare the time-varying model estimations with an error-correction model for consumption and

²¹ However, if we allow also the constant to be time varying, we achieve steadily declining excess-sensitivity parameter from .5 to about 0.1, but the time-varying real market interest rate coefficient turns out to be negative over the whole estimation period.

explore the excess-sensitivity parameter within this more ‘fully fledged’ consumption model. The idea in presenting this type of equation is to take into account all the relevant conditioning variables and to compare these results with the Campbell-Mankiw-type augmented Euler equation.²² In addition to cointegrating core variables, this model uses changes in the unemployment rate to reflect income uncertainty and housing prices to account for income expectations. Another important aspect is that, based on the diagnostics, this ECM model is an adequate representation of consumption, whereas the estimated fixed parameter models were not.

Appendix 2 shows the specified ECM model, where the average excess-sensitivity parameter is only 0.15. The corresponding time-varying excess-sensitivity parameter from this model can be seen from figure 5. The time path also shows that the proportion of consumers that depend directly on disposable income has decreased steadily, to about 0.26 at the end of 1990s. As a rough conclusion, we can also agree that there may exist significant liquidity constraints or at least rule-of-thumb behaviour among consumers. However, we notice that in the error-correction model there is no hump in the path of the excess-sensitivity parameter during the recession.

To summarise, it would be good to have an economic explanation that accords with the estimation results. Financial deregulation at the end of 1980s clearly increased the availability of credit in Finland, but the excess-sensitivity parameter does not react very much. Household credit increased by a third of GDP within a few years. The estimation using basic models showed that the proportion of liquidity-constrained consumers increased during recession, but as regards the results for the varying constant we cannot be sure whether this was due to increased sensitivity to disposable income or to something else such as the real interest rate.

To be specific, we can also look for discrete changes in financial legislation, indicating major turning points in credit restrictions and availability (see also Sefton and Veld, 1999). In Finland the late-1980s financial deregulation greatly increased consumption and especially housing purchases. The sequence of financial liberalisation led to the abolishment of both the central bank’s collateral requirements for credit and the ceilings on lending rates for housing loans at the end of 1986. Housing loans were allowed to be tied to 3- and 5-year market

²² A similar comparison was made by Gigardin, Sarno and Taylor (2000). They compared the time-varying version of the Campbell and Mankiw model with the ECM specification and found significant time variation in excess sensitivity, but concluded also that the time-varying model could not be rejected against ECM model.

interest rates from the start of 1988. The abolishment of credit restrictions led to a surge in banks' housing loans, and since demand for housing was rationed. Households were eager to use this opportunity to borrow against housing capital and labour income.

Most of the theoretical literature emphasises the fact that financial deregulation has a direct and distinct effect on the ability of previously liquidity-constrained consumers to borrow against their future labour income. This should reduce the correlation between consumption and current disposable income. In the short run, consumption of durables and housing would increase, which would show up in a lower saving rate. Papers presenting this type of explanation for the lending and consumption boom include Muellbauer and Murphy (1990), Lehmussaari (1990), Koskela, Loikkanen and Virén (1992) and Bayoumi (1993). Based on our estimations, there is a definite decline in the excess sensitivity only after the recession.

However, based on Swedish data, Agell and Berg (1996) offer a different interpretation, which relies on upward revision in expected income, which leads to increased borrowing. They ask the important question of whether excess sensitivity has changed during the 1980s financial deregulation. In addition, they show that increased consumption growth is due not only to the wealth effect in consumption but also to the boom in housing prices. They argue that income growth rates were exceptionally low when borrowing increased. However, it can be argued that an increase in housing prices may also reflect favourable income expectations in the current housing market situation (Muellbauer and Murphy, 1990). This explanation emphasises the increase in the demand for loans while financial deregulation gives more weight to the supply of loans. The reason for growth in borrowing could be linked to income expectations, regime shift in interest rates, or exchange rate policy.

In the Finnish data we saw a rise in the average propensity to consume between total consumption and income and that this effect was limited almost entirely to durables consumption. Liquidity-constrained households may use durables purchases as a tool for smoothing consumption.²³ In fact it can be seen that the size of the APC between non-durables and income was even higher at the start of the 1990s, due to the deep recession and rapid increase in unemployment, which surely should have increased the number of liquidity-constrained households. Agell and Berg (1996) also provide this type of evidence for Sweden by citing the coincidence of the

²³ See Flavin (1999).

consumption boom with the housing boom, but they attribute this to the wealth effect.²⁴ Their main point is that this explanation does not quite fit for the Swedish case of a higher propensity to consume out of income. In fact they present recursive MPC estimates for disposable income that indicate that there is no significant drop in consumption sensitivity to income during the financial deregulation in Sweden in the late 1980s.

Deterministic shifts over the forecast horizon are the main cause of forecasting failures in respect of non-stationary series.²⁵ Model specification and estimation uncertainty are not equally important causes of failure. Thus we must correct for structural breaks or make other types of corrections for non-constancy of parameters. The reason for using time-varying parameters is based on fact that when the assumption of constant parameters fails the model fit is not very helpful in guaranteeing accurate ex ante forecasts.

4 Conclusions

In this paper we have reviewed the techniques for estimating the Campbell and Mankiw model and tested for excess sensitivity of consumption. Econometric modifications that were expected to fine tune the fixed parameter estimates did not result in much of a change in the average estimates of the overall proportion of liquidity-constrained agents. IV estimation has been proposed for testing excess sensitivity, as any lagged instrument is orthogonal to the error term in a simple consumption growth regression relating consumption to disposable income. GMM estimates were also calculated since some of the model variables are based on expectations. However, these estimates were also quite similar.

The estimates of excess sensitivity in Finland are clearly higher (around 0.5) than in a few major countries and even in comparison to similar countries like Sweden. A similar finding of Agell and Berg (1996) was that financial deregulation did not at first seem to affect the income sensitivity of consumption of non-durables, but the in following recession this sensitivity did increase significantly and

²⁴ Agell and Berg (1996) also argue that wage earners had good grounds to revise their expectations about future earnings, and this could be seen to be a major reason for consumption growth.

²⁵ Clements and Hendry (1999, p. 194) note that consumption forecast failure in the United Kingdom in the late 1980s, using the famous DHSY error-correction model, was mainly due to financial deregulation, which caused the consumption bubble.

restrict the purchases of durables. This observation accords with the decline in income expectations, and also to growing income uncertainty as unemployment increased rapidly.

After estimating time-varying excess-sensitivity estimates, we would conclude that the sensitivity of consumption to current disposable income clearly decreased in the late 1990s. Time-varying excess-sensitivity estimations also produce varying results concerning both the proportion of liquidity-constrained consumers and the overall path of the parameter. Time-varying excess-sensitivity coefficient tends to decline rather steadily over the estimation period, from 0.5 to significantly lower levels. Most likely this is a consequence of financial deregulation, though recursive and rolling 'fixed' parameter estimations indicated a very definite hump during the recession.

Finally comparing these results with an error-correction model – allowing only the excess-sensitivity parameter to be time varying – we get the result that the proportion of liquidity-constrained consumers declined rather smoothly from around 35% to 25% by the end of 1990s. All in all these observations have important implications for forecasting models with constant parameters. One should be caution about assuming that the proportion of liquidity-constrained consumers has remained stable and that the effect of disposable income would be not change. Clearly, the elasticity of consumption with respect to current disposable income needs further assessment in consumption forecasting models.

Figure 3.

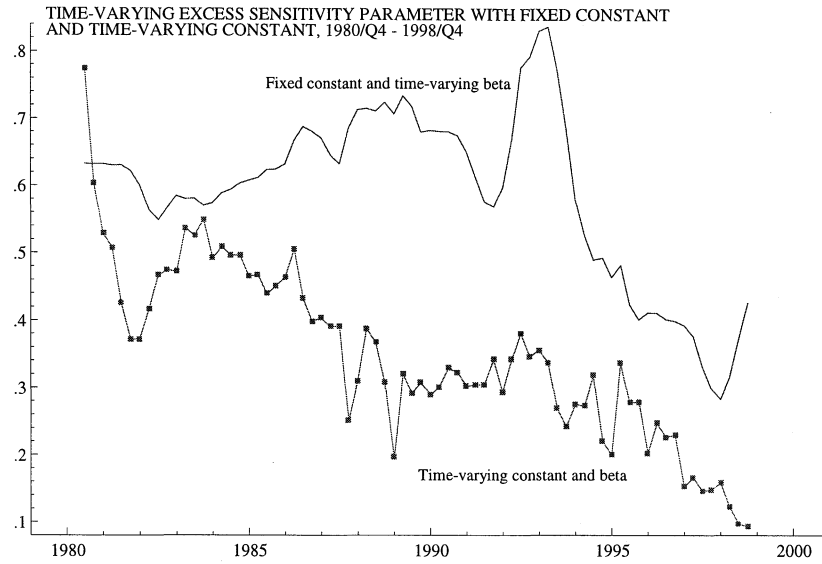
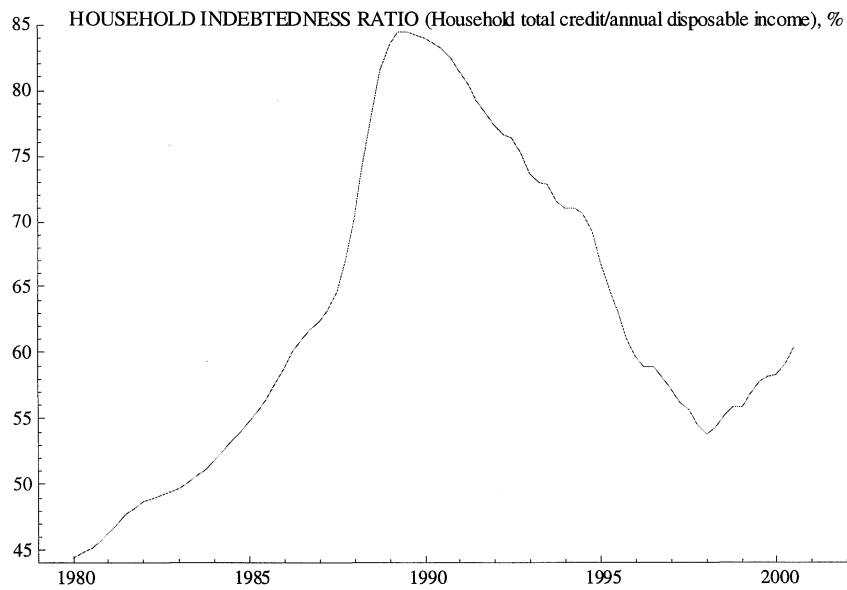
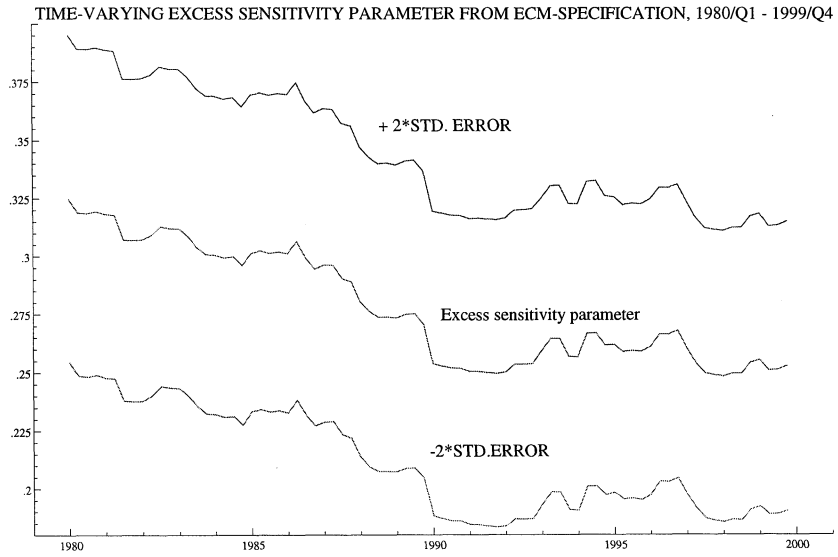


Figure 4.



Source: Statistics Finland

Figure 5.



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Appendix 1.

Comparison of de Bondt and Ludvigson specifications

Table 4. **Comparison and encompassing tests between de Bondt and Ludvigson specifications**
Estimation period: 1980Q4 – 1998Q4

de Bondt model (M1)			Ludvigson model (M2)		
Variable	Coefficient	t-prob	Variable	Coefficient	t-prob
Constant	5.75	0.0004	Constant	1.77	0.0000
d4lryd	0.28	0.0000	d4lryd	0.33	0.0000
margin_1	-1.27	0.0018	rrlbn_1	-0.13	0.0180
rrlbn_1	0.15	0.0352	d4lrconscre	0.10	0.0000
bcefp_1	0.07	0.0000			
R ²	0.742		R ²	0.675	
DW	0.985		DW	0.62	
Diagnostics			Diagnostics		
AR 1- 5 F(5, 63)		0.0000**	AR 1- 5 F(5, 64)		0.0000**
ARCH 4 F(4, 60)		0.2420	ARCH 4 F(4, 61)		0.0078**
Normality $\chi^2(2)$		0.2430	Normality $\chi^2(2)$		0.8569
Xi ² F(8, 59)		0.9032	X ² F(6, 62)		0.0054**
Xi*Xj F(14, 53)		0.1928	Xi*Xj F(9, 59)		0.0165*
RESET F(1, 67)		0.0121*	RESET F(1, 68)		0.6184

$\sigma[M1] = 1.26521$ $\sigma[M2] = 1.41122$ $\sigma[\text{Joint}] = 1.12978$

Model 1 v Model 2	Form	Test	Form	Model 2 v Model 1
-7.08132	N(0,1)	Cox	N(0,1)	-9.10204
5.94499	N(0,1)	Ericsson IV	N(0,1)	6.98645
14.5764	Chi ² (1)	Sargan	Chi ² (2)	26.0594
18.2806	F(1,67)	Joint Model	F(2,67)	20.3302
[0.0001]				[0.0000]

Variables:

d4lnoncd = Annual change in non-durable consumption, %
d4lryd = Annual change in real disposable income, %
rrlbn = Real lending rate for new household loans, %
d4lrconscre = Annual change in real consumer credit, %
margin = Interest rate margin between lending and deposit rates, %
bcefp = interaction term between business cycle indicator and interest rate margin

Appendix 2.

ECM consumption function for quarterly change in real consumption

Data: 1980Q1 – 1999Q4, OLS estimation with new ESA95-data

Dependent variable: Log-difference of real private consumption

Variable	Coefficient	Std.Error	t-value
Constant	0.00519	0.00106	4.872
DLPYR	0.15158	0.06064	2.499
ECT_1	-0.16859	0.05793	-2.910
DURX	-0.00616	0.00213	-2.886
DLRIHX	0.14716	0.03433	4.286
D90Q3Q1	-0.01025	0.00518	-1.978
D82Q4	0.02504	0.00832	3.009
D83Q1	-0.02512	0.00854	-2.938

$R^2 = 0.5777$, $F(7,72) = 14.069$ [0.0000], $\sigma = 0.00821972$, $DW = 2.10$

Residual diagnostics:

Test	Statistic	p-value	Test description
AR(1–5)	$F(5,67) = 2.2106$	[0.0633]	Autocorrelation test
ARCH(4)	$F(4,64) = 0.2909$	[0.8828]	Autoregressive heteroscedasticity
Normality	$\chi^2(2) = 0.5452$	[0.7614]	Normality test
X_i^2	$F(11,60) = 0.7399$	[0.6962]	Heteroscedasticity
$X_i * X_j$	$F(19,52) = 0.8015$	[0.6948]	Linearity
RESET	$F(1,71) = 1.8815$	[0.1745]	Functional mis-specification

Variables:

- DLPYR = Log-difference of the household real disposable income
- ECT_1 = Error-correction term from a homogenous long-run relation between consumption, disposable income and household financial wealth
- DURX = Difference in unemployment rate, %
- DLRIHX = Log-difference in real house prices
- D90Q3Q1 = Impulse dummy = 1 for 1990Q3 – 1991Q1 (recession)
- D82Q4 = Impulse dummy = 1 for 1982Q4 (devaluation)
- D83Q1 = Impulse dummy = 1 for 1983Q1 (devaluation)

Macroeconomy and consumer sentiment: Performance of the Finnish consumer barometer after ten years

Kari Djerf • Kari Takala

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Abstract

The Finnish Consumer Barometer was introduced in autumn 1987. Data were first collected twice a year and from August 1991 until September 1995 quarterly. After Finland joined the European Union in 1995, the survey was adopted as one member of the Harmonised Consumer Survey of the European Communities. Since October 1995, data have been collected monthly.

Performance of the Consumer Barometer has already been evaluated by means of descriptive studies (see Djerf, 1990). As the survey matures, it becomes feasible to make a more thorough study on the usefulness of the survey. Here we are, for example, interested in investigating how consumers were able to predict the long-lasting recession of our economy.

The consumer confidence index and the five questions used for calculating it are compared to various components of Finnish macroeconomic time series. Additionally, we analyse the coincidence of other common measures (unemployment expectations, inflation expectations, etc.) as well as other, less frequently studied indicators such as the willingness to save and borrow with their possible counterparts in the real economy. It is important to evaluate whether consumer assessments about different economic questions are useful in predicting various macroeconomic variables ie we look for the additional information contained in barometer answers. We end by considering the usefulness of various indicators for specific macroeconomic behavioural equations.

Macroeconomy and Consumer Sentiment: Performance of the Finnish Consumer Baro- meter after Ten Years*

Kari Djerf / Kari Takala

1 Introduction

Macroeconomic forecasters have observed the usefulness of barometer surveys for several decades. In Finland, the last 10 years have seen major breakthroughs for barometer indices. Barometer studies are used in predicting corporate activity (both large manufacturing and medium sized firms), investment and consumer expenditure. Without doubt, forward-looking economic agents know something about their future economic behaviour. It is another matter though, whether these expectations, intentions and plans are necessary in forecasting forthcoming economic activity, and whether the consumer sentiment is a cause of the observed activity.

In this paper, we study the relationship between economic activity and consumer assessment about the most important macroeconomic variables. In particular, we are interested in the predictive power included into consumer responses concerning major macroeconomic variables. The similarity in growth rates and coincidence of turning points with the actual changes are also important to record. One major mistake during the deep recession of the 1990s was the negligence of the wealth effect in macroeconomic consumption functions. At least partly these forecast errors would have been avoided by giving more attention to consumer assessment about the growth of the economy. The same applies to Sweden (Berg and Bergström 1996).

We try to assess the usefulness of consumer sentiment variables as leading indicators of economic activity. Related questions are how much lead indicators provide and how reliable they are in giving correct signals about forthcoming activity. In particular we are interested whether consumer sentiment indicators give forecasters early warning of economic slowdown or overheating.

In practical forecasting work, barometer surveys can also be useful for macroeconomic forecasters in giving a fresh update about current economic activity as official statistics lag at least a quarter or so from the current situation. In this sense, the usefulness of sentiment variables is not limited by the role of having leading indicator status; even a coincidence could be enough.

It is possible in principle that consumers have additional information about their forthcoming behaviour that is not available to econometric forecasters. Thus, we may approach this question by classifying the information sets of agents by weak rationality,

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semi-strong and strong rationality. Weakly rational expectations (sentiments) are based on past and present of the variable in interest. Semi-strongly rational expectations can also use any publicly available information related to the variable in interest like official statistics. Strongly rational expectations use any relevant information affecting the variable, including subjective motives that may affect the development of the variable.¹ As a research strategy, we first test the weak rationality of the sentiment expectations and proceed these stronger forms of rationality by estimating few model-based information sets using sentiment variables as additional variables in behavioural equations. The problem in stronger forms of rationality is naturally the ambiguity of the information set. Therefore, an ultimate answer could be out of the scope of any study. However, some sort of qualitative answer is clearly possible.

2 Background of the Finnish Consumer Barometer

The Finnish consumer barometer was introduced in late 1987. In most central European countries – France, Belgium, Denmark, Italy, Holland, UK and Sweden – consumer surveys were started along with the EU survey already in the early 1970s. In the beginning, the consumer survey was done twice a year, but soon it was increased to a quarterly basis. In October 1995, the consumer barometer was then switched to a monthly basis. Over time there have been changes in some of the questions.² Some definitional improvements have also been made, e.g. the assessment about forthcoming inflation is now published as a percentage and not as characteristic values, which are somewhat more difficult to assess.

In most questions, the classification basis is still ordinal in scale, and the average of the characteristic values is used in measuring the overall sentiment. As the typical range of the characteristic value is not known, it is only possible to learn the usefulness of the answers by measuring them for some time. The consumer barometer is done by interviews to reduce the time lag. The time lag from the start of interviews is currently 2–3 weeks. The Finnish survey is based on standard EU definition of overall consumer confidence indicator (CCI), which is made up as an average of five questions (Figure 1):

1. The financial situation of households now compared to 12 months ago
2. The expectations of the financial situation of households over the next 12 months
3. The general economic situation now compared to 12 months ago
4. The expectations of the general (Finnish) economic situation during the next 12 months
5. The advantageousness of purchasing capital (durable) goods at present

¹ In fact, one approach would be to analyse whether the changes in the distribution of the sentiment answers as characteristic values carry information about the uncertainty of economic activity. Currently, the calculation of the characteristic values as the difference between those who expect growth and those who expect decline does not react to the share of those who answer the situation to be unchanged. However, it may well be that the share of those regarding the situation to be unchanged may reflect the uncertainty of the net change. Therefore this ‘second order’ information may be useful sometimes in constructing statistical confidence bounds to responses.

² For example the question of the regularity of saving has been replaced by questions that concern the ability to save, which is more properly related to consumption function theory (see Takala 1995b).

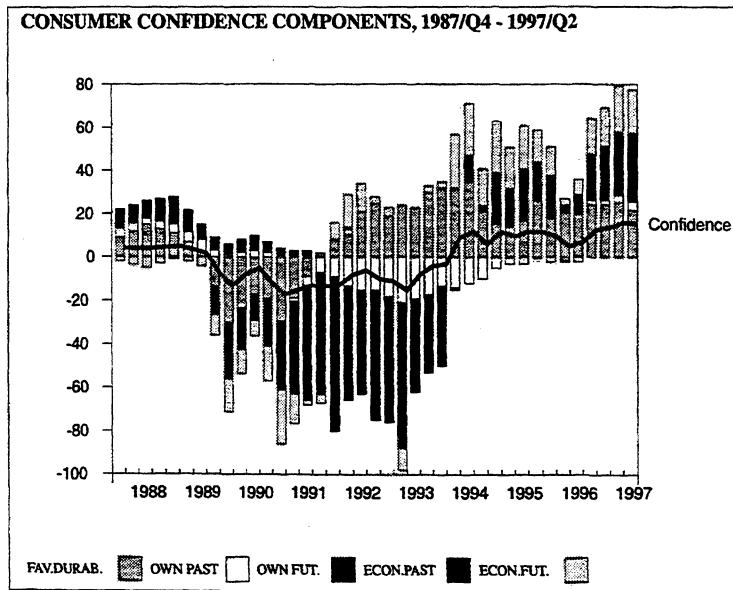


Figure 1 CCI and its Components

In questions 1–4 consumers answer whether the situation is a lot/a little better or worse, about the same or 'do not know'. In the characteristic value these responses are weighted by ± 1 , ± 0.5 or zero respectively. Question 5 contains only response options favourable, unfavourable and neither/nor in addition to the 'do not know' answer.³ Since there are two pairs of questions that are backward and forward-looking and one last question whose timing balance is close to present, it can be said roughly that CCI is balanced to current time. CCI did not indicate significant seasonal variation, but there is slight tendency for consumers to be more optimistic in spring than in the second half of the year. Seasonality was not found either in forward-looking expectations about one's own economic situation or, the general economic situation. In this respect the series are quite different from the Swedish data which showed strong seasonality in both these forward-looking indices (Berg and Bergström 1996). It must be remembered, however, that the Finnish series are much shorter.

The component which correlates the most with the CCI is the past year of the economic situation ($r = .91$). The correlation between components is strongest between the past and future of one's own economic situation. This symmetry may not be surprising since there is high autocorrelation and therefore inertia in household income and consumption. In fact, households are trying keep their consumption as constant as possible according to major consumption function theories. Especially in Finland, income transfers rose rapidly during the last recession, and kept the purchasing power of households

³ The weights 1 and 0.5 seem somewhat arbitrary, and it should be analysed whether a more optimal weighting could be achieved, e.g. by weighting with normal distribution based on central-limit theorem.

much more constant than overall economic growth. The correlation with CCI is stronger to the general economic situation than to a household's own situation. Surprisingly, the perceived favourableness of time for large purchases is slightly negatively, correlated with a household's assessment of its own economic situation.

There is a small bias downwards in the CCI compared to GDP growth, but the major difference lies in the volatility of characteristic values. The standard deviation of CCI is twice that of GDP. Consumer assessments of past economic performance is also much more volatile than that of forthcoming economic growth, which is also typical of optimal forecasts.

If we look more closely at the CCI components, we see that consumers' assessment of their own economic situation seems to be generally much more stable than the assessment of the economy as a whole (Figure 2). One explanation for this could be the fact that the social security transfers received by households rose enormously during the recession. The same phenomenon can also be seen in the share of static (no change) assessment about households' own economic situation (Figure 3). It can also be seen that consumer expectations about the future are in general brighter than the assessment about the past. Partly this may be due to overall economic growth. There are also other signs of an optimism bias. During recession the past year of the economy was assessed to be a lot worse by 40–50 % of households, but this did not affect at all the assessment about forthcoming year of the economy. According to normal inertia of the economic growth, however, some signs of continuing slowdown would have been expected.

It can be seen that the actual variation in consumer assessment takes place in the answers 'a little better/worse', their contributions to the characteristic values are usually larger than those answers indicating 'a lot better/worse'. Consumers are inherently more uncertain about the future than the past as indicated by the 'Do not know'-answers. Of course the amount of information is also asymmetric for these two types of questions, since the past is already gone while the future has not yet happened. The uncertainty reflected in the 'Do not know' answers is systematically higher (9–20%) in the question concerning the disposition to buy durables, while for the other questions their share has been always below 9 percent. The answers concerning the favourable time to buy capital goods increased the level of the confidence index in 1990s (Figure 4).

It is also interesting to examine which CCI components have contributed most to the overall consumer confidence in different periods. The business cycle has clearly affected CCI contributions. As mentioned above, consumer assessments of the past year's economy were most strongly correlated with the CCI index. During the severe recession of May 1990 – November 1992, this correlation declined significantly and expectations about the forthcoming year of the economy contributed more clearly to the CCI. During the growth period of August 1994 to June 1997 the advantage of buying durables lost its significance in relation to other CCI determinants.

We may assume that the share of consumers that responds 'the situation has stayed the same (or is expected to be the same)' reflects the certainty or the probability of the actual outcome (Figure 5). The ordinary consumer confidence indicator measures especially the change in assessment and expectations. It is therefore possible to weight the characteristic value with the uncertainty ($1 - \text{share of static responses}$) concerning the balance of those answers indicating change. So greater uncertainty about the change will be dampened by weighting the characteristic value (Figure 6). Greater certainty about the change will give more weight to the confidence indicator. However, the weighting did not change dramatically the correlation of the CCI with the actual GDP growth

(Figure 7). One technical problem in weighting is related to the limiting maxima and minima of the weighted CCI, although the ranges are now closer to actual percent changes of GDP.

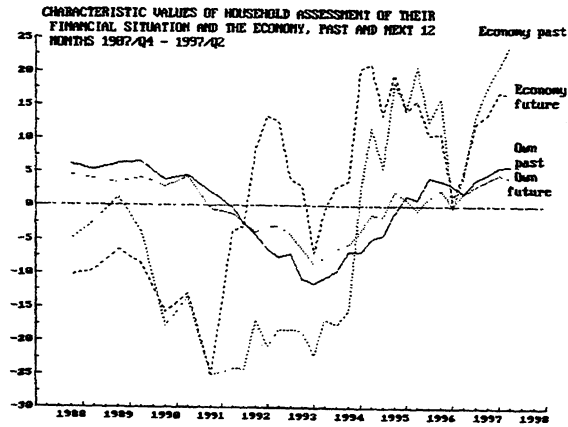


Figure 2

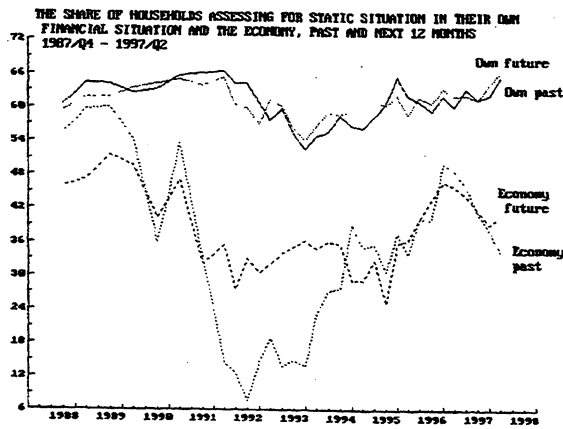


Figure 3

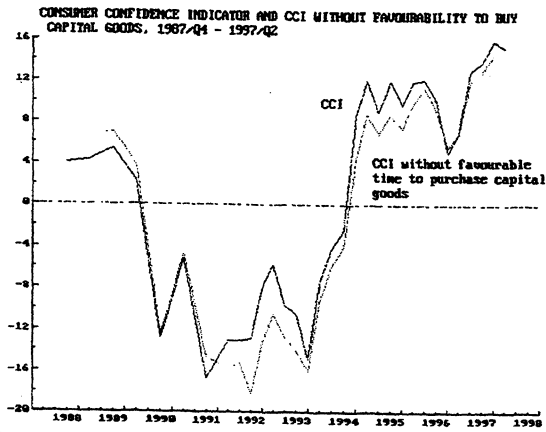


Figure 4

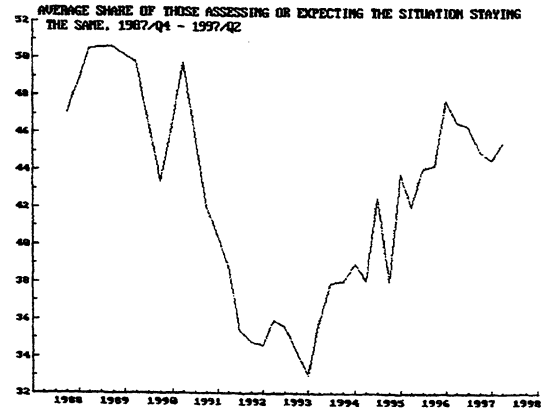


Figure 5

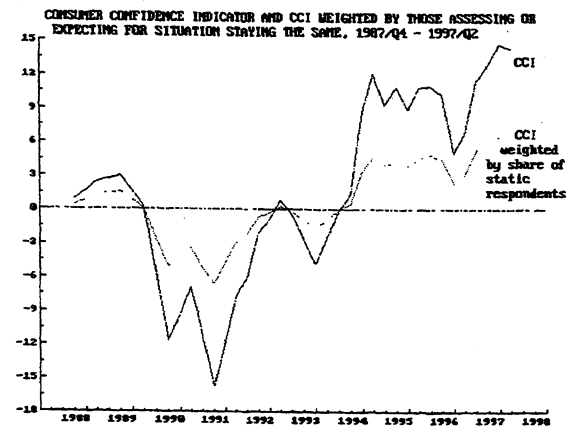


Figure 6

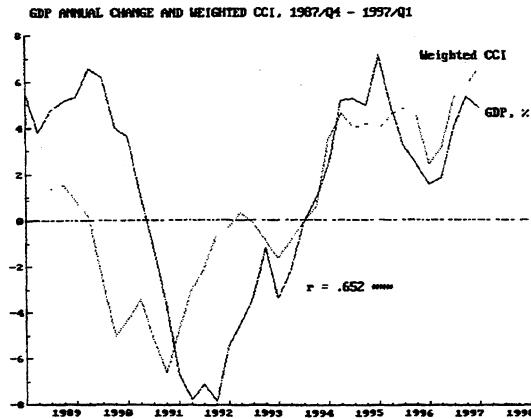


Figure 7

3 Consumer Sentiment Variables and the Macroeconomy

The Finnish consumer barometer study contains a fairly broad amount of questions about main macroeconomic and household-related variables. However, it is not always clear what variable correlates with what barometer question. In this section we try to identify some of the closest relations observed.

The consumer confidence indicator (CCI) itself contains information on both the economy as a whole and the economic situation of the individual household in particular. The developments in these indicators may not always coincide, but aggregating these to the economy level should tell us the overall picture. The key issue is, of course the representativeness of the survey. The sample size in the survey has varied. At present it is about 1,700 individuals from which we normally obtain more than 1,500 responses. The response rate has been about 85 per cent. Each respondent represents about 2,600 individuals. When households are considered each response represents about 1,500 other households. We have not found indications of bias due to nonresponse or measurement errors so far. Nevertheless, the reliability of some questions may not be as good as is needed. That is especially true when the respondents are asked questions of rather infrequent phenomena, for example plans to acquire a house.

One way to assess the quality and usefulness of the questionnaire would be to compare the household assessments about the past to that actually recorded by statistical officials. In Figure 8 we see the GDP for four-quarter growth and consumer assessments on the past 12 months in same range plotting. Clearly, it can be seen that consumer assessments correlate very closely to the actual changes and that turning points are almost identical. Private consumption also contributed a great deal to the 1991–93 recession. In Figure 9, GDP growth is compared to household ex ante predictions one year ahead of the economy with a four-quarter lag to match the time periods. The correlations are quite convincing. In fact they are better than many other macroeconomic forecast models for the early 1990s (Figure 10). However, it can be seen that consumer expectations about the general economic situation gave two major misleading signals and overall

CCI could be better in indicating economic activity (Figure 11). First, consumers expected a slowdown in growth during 1993, which did not happen. Second, households did not foresee the temporary slowdown in growth in late 1995 and the beginning of 1996.

The same pattern holds true for inflation and inflation expectations, too (Figures 12–13). Consumers know very well the past rate of inflation, and overall they seem to predict acceleration or deceleration of forthcoming inflation very well. Households made only one major mistake in inflation expectations in late 1991. Households expected a clear acceleration in inflation during 1992 due to the devaluation of the Finnish markka in November 1991 by 12%. Against the widely discussed devaluation-cycle theory and our past experience, inflation did not accelerate, however, as the price competitiveness of the export firms did not improve enough and the domestic recession did not abate as quickly as expected. Devaluations have previously increased the markka import prices, but this time this inflation channel was not as strong as earlier because of a larger tendency for pricing to the market of importers. However, the household expectations were revised almost optimally in the next barometer round in spring 1992. The floating of the Finnish markka started in September 1992 which further reduced the external value of the markka until late spring of 1993.

Household assessments about their own economic and financial situation seemed to have an in-built optimism during the last recession. The past 12 months have always been more gloomy than that expected ahead in the barometer history (Figure 14). The same applies to the comparison of the lagged expectations and assessments about the recent past (Figure 15). Even though the recession was deeper than most macro-economic forecasters expected, it seems somewhat odd that consumers are as cautious and conservative in predictions as other forecasters. Historically, consumers have shown a systematic optimism bias in assessments of the future, which must be corrected. Household expectations of their own financial situation was regressed on the assessment of the past 12 months' change; the bias was 4.8 and only the current evaluation of the own past was significant.

One would expect CCI to have a close relationship with consumption, but it looks as if CCI has an even closer relationship with durable consumption, or more correctly, with durables purchases as measured in National Accounts (Figures 16–17). In this respect results are similar to those found for Sweden (Ågren and Jonsson 1991). Explaining durable purchases can be achieved by using the question concerning the propitious time to buy durables (Figure 18). Durable purchases have been found to depend on the unemployment rate in many studies, since increasing unemployment means rising income uncertainty. Households seem to foresee the forthcoming changes in unemployment amazingly well (Figure 19). Consumer assessment of other aspects of their behaviour is, in general, also useful. For example, household borrowing plans seem to be efficient predictors for actual borrowing (Figure 20), and consumer assessments of savings correlate with interest rate evolution (Figure 21). Table 1 presents the closest relationships found between consumer sentiment, forward indicators and macro-economic variables. While in most cases it is self-evident what the question measures, a few interesting relationships emerge. Even though the maximum correlation lag did not appear to be exactly four in most expectations with one year horizon, the correlation was in most cases almost as high because of serial correlation.

Table 1 Consumer Sentiment Variables and the Corresponding Macroeconomic Variables Based on cross-correlation functions

Sentiment variable	Macroeconomic variable	Max correlation	Max correlation found on lag
CCI	GDP, %	.761	1
CCI	C, %	.750	1
CCI	CD, %	.847	1
Household own financial future	C, %	.852	2
Future of the economy	GDP, %	.481	4
Inflation expectation	CPI, %	.895	3
Favourable time for large purchases	CD, %	.801	6
Fav. time for saving	Saving rate, %	.582	3
Fav. time for borrowing	Saving rate, %	.650	4
Fav. time for saving	Lending rate, %	.865	0
Plans for raising a loan	Borrowing, %	.939	0
Unemployment expectation	Change in unemployment, %	.734	2

Abbreviations:

- GDP, % = Four-quarter GDP growth, %
- CD, % = Four-quarter change in durables purchases, %
- C, % = Four-quarter change in private consumption, %
- CPI, % = Four-quarter change in consumer price index, %
- CCI = Consumer confidence indicator

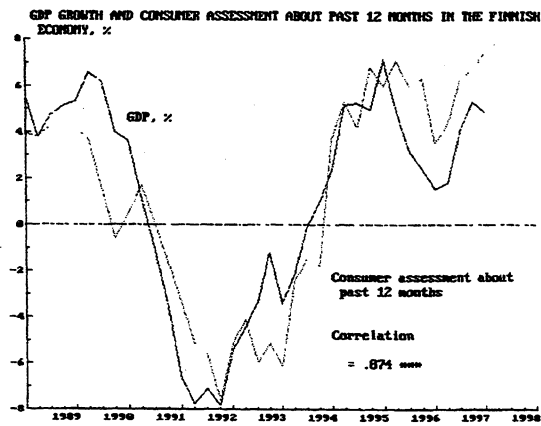


Figure 8

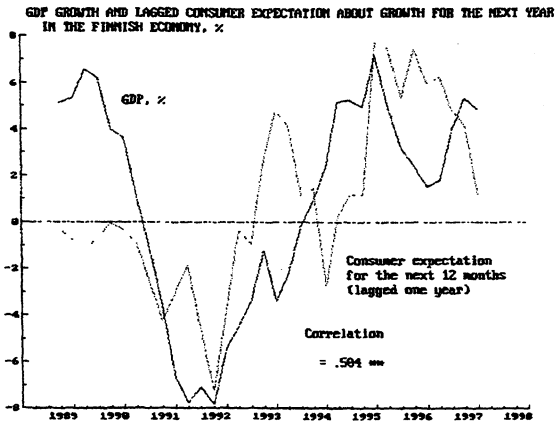


Figure 9

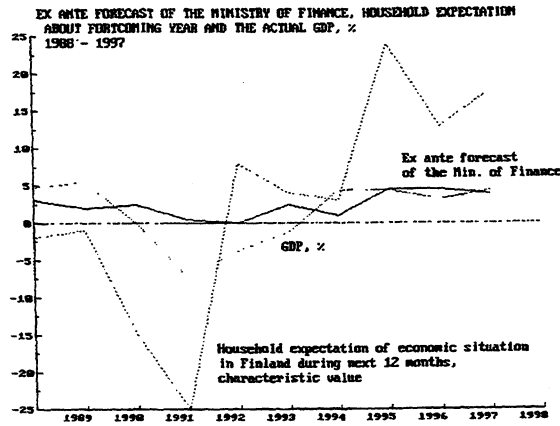


Figure 10

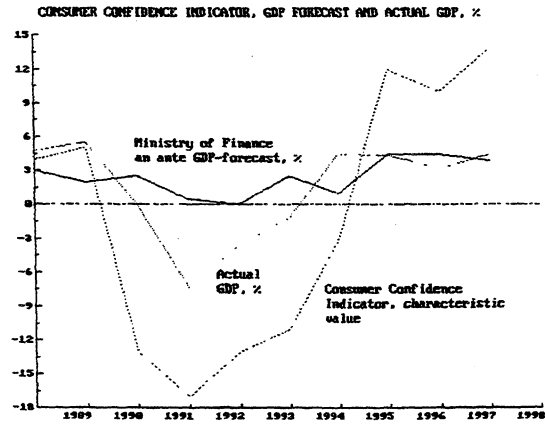


Figure 11



Figure 12

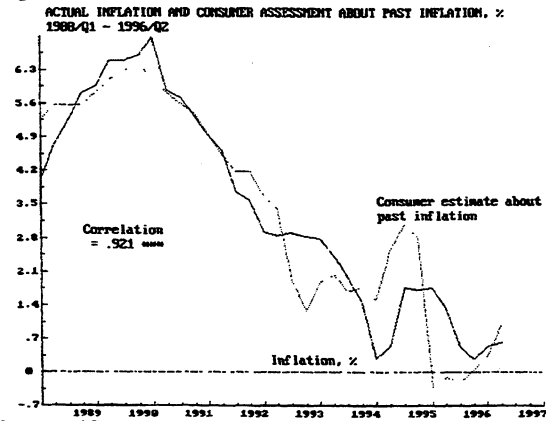


Figure 13

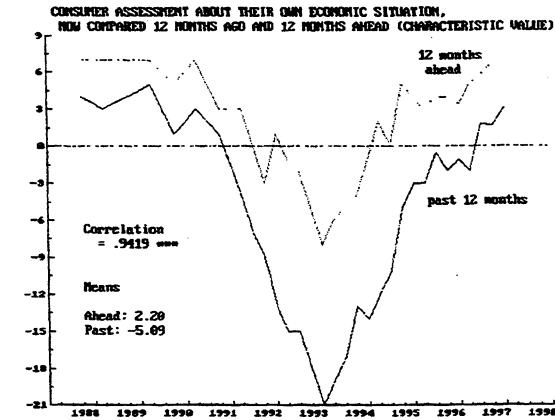


Figure 14

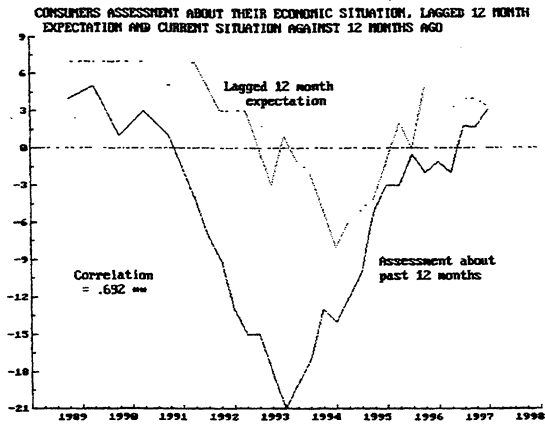


Figure 15

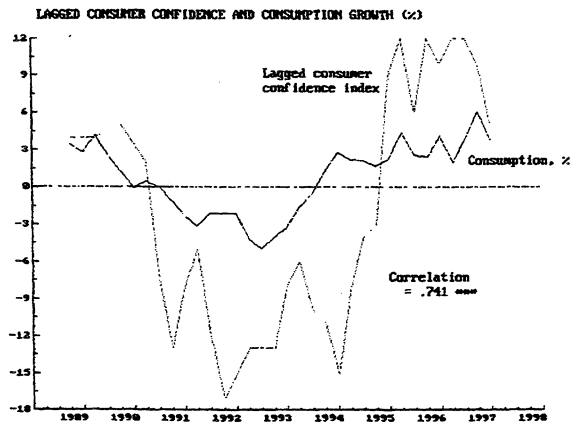


Figure 16

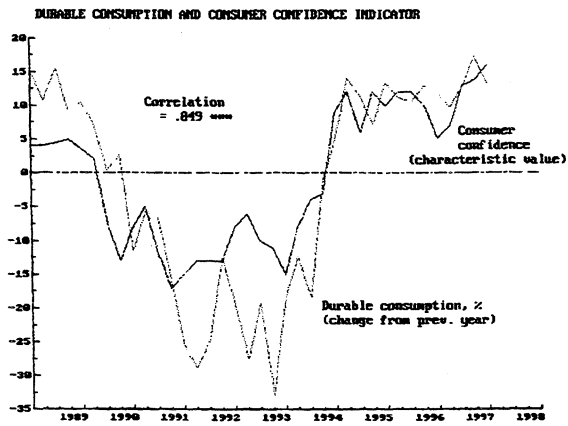


Figure 17

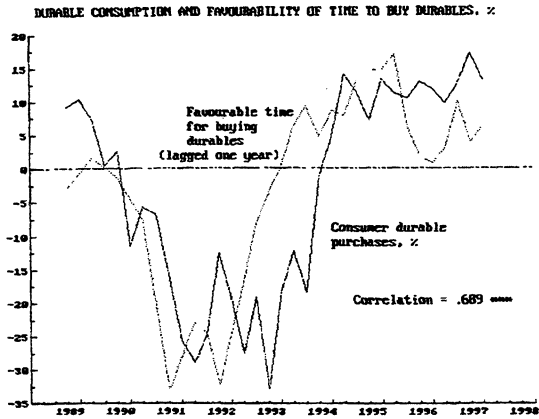


Figure 18

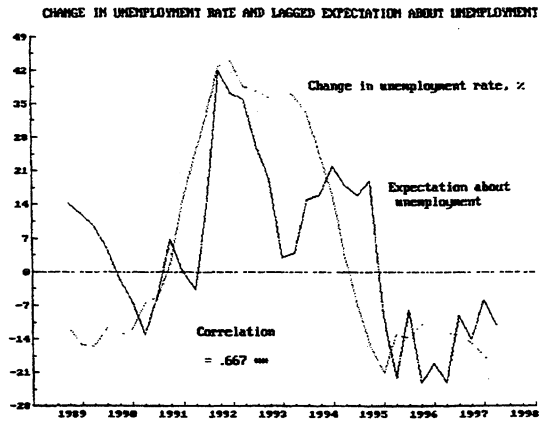


Figure 19

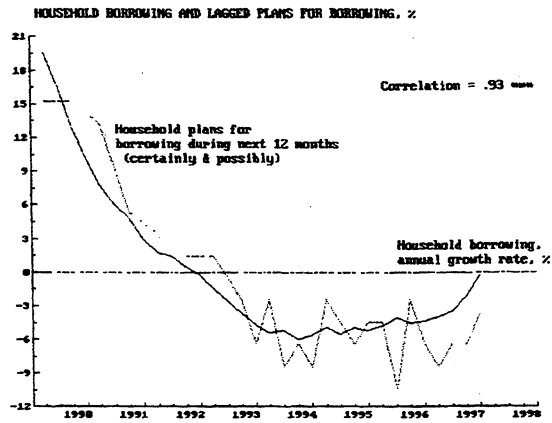


Figure 20

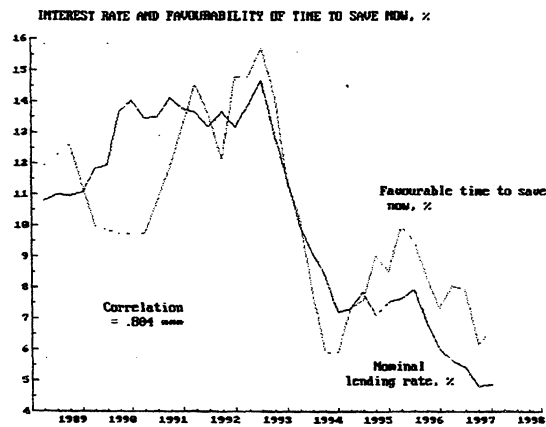


Figure 21

4 Macroeconomic Activity and Confidence Indicators

In this section we ask the basic question whether sentiment variables contain leading or coinciding information about actual macroeconomic activity variables. In addition, the timing of the answers with respect to that actual macroeconomic variation is utterly important as the purpose is to find out the usefulness of sentiment variables for forecasting. The reliability of such information is also important, since false signals are problematic.

The usefulness of the sentiment indicators must be based on views about their role in the forecasting procedure. Dieden and Kennedy (1997) classify two basic views about the status of confidence indicators that summarize either observable or unobservable economic conditions.

It is either the case that sentiment indicators reveal economic conditions not observable from other macroeconomic variables, or they actually measure non-economic psychological factors, which nevertheless affect the economy. In the sense that consumer sentiment variables truly reflect expectations of macroeconomic variables and therefore carry information about forthcoming behaviour, it must be the case that confidence indicators contain also an unobservable component.

In general, it would be surprising if consumer intentions did not reflect future behaviour. Nobody else knows, for instance when a household plans to take a housing loan, considers buying a durable like car, or will decide to save regularly. At the time of the survey, no statistics are available for some of the indicators. As all expectations are formed on the basis of past and current information, consumer sentiment variables are functions of the past. If household expectation formation exhausts the past optimally and sufficiently, there should not be any bias in ex post expectations.

Macroeconomic variables as such cannot include any forward-looking effects, for example decided policy changes, whereas in econometric models, these effects have to be included explicitly through exogenous variables. Expectations of consumers include such policy changes or even household reactions to these changes, and therefore, the

information set of consumers is by definition larger than in backward looking econometric models.

Carroll, Fuhrer and Wilcox (1994) ask whether consumer sentiment indicators include predictive power over consumption, and whether the sentiment indicator contains additional predictive power about consumption not included into any other relevant explanatory variable. They also discuss whether confidence variables simply reflect economic situations or actually cause economic volatility. Do we necessarily need confidence indicators as additional variables, for example in explaining consumption? In the case of consumption, this question has special interest since it accords with the famous hypothesis of Robert Hall (1978), whereby the lagged consumption level should reflect the best forecast for current consumption. Halls theory is based on a utility maximising consumer with rational expectations. This theory also implies that observed consumption apart from durables should follow a random course as only current innovations will disturb consumption from the chosen equilibrium level manifested in the lagged consumption.

If consumer sentiment variables are regarded only as expectations to actual macroeconomic variables, it may likely be that sentiment and corresponding macroeconomic variable are also cointegrated. Rational expectations should also coincide in the long run with the observed development of the variable. The only practical problem is that confidence indicator is presented in differences, i.e. they are stationary. As cointegration may prevail only between integrated variables, consumer sentiment variables should be first aggregated to levels.⁴ We also need theoretical background concerning the relationship to test these properties within a bivariate VAR-system framework. As consumer sentiment and corresponding macroeconomic variables should be cointegrated, the short-run discrepancy (error-correction term) should predict either one or both of the integrated variables. In this case most likely it will predict the forthcoming changes in sentiment.

In the market, only a few 'price variables' are forward-looking, like prices of dwellings which indicate the discounted present value of housing services and therefore income expectations. Long-term interest rates also fundamentally reflect the intertemporal price of consumption, which is compounded from rather constant real interest rate and expected inflation. Otherwise, there is hardly any information about future beliefs and actions of economic agents.

4.1 *Granger Causality Tests*

Granger causality tests can be applied to analyse the predictive causality between two stationary variables. As dependent variable we use four-quarter changes of macroeconomic variables and as explanatory variables various consumer sentiment variables. The purpose is to find out whether consumer sentiment variables include any additional information on the past of the dependent variable. Granger causality tests are performed by two linear regressions, one containing the lagged values of the consumer sentiment variable in the equation and one without. The Granger causality test is then based to the F-test for the lagged consumer sentiment variables as a group. If this F-test is significant, we can conclude that the consumer sentiment variable reduces the forecast error of

⁴ Performed ADF-unit root tests did not prove very convincingly that the sentiment variables would be stationary. The characteristic values calculated are clearly bounded, since they are calculated from the population shares of respondents, and should therefore be stationary in the longer run.

the macroeconomic variable and it is therefore useful at least over the autoregressive univariate model.

In addition, we perform Granger causality tests on the opposite (feedback) direction to see whether consumer sentiment assessment is truly forward-looking and not simply an autoregression of the past of the macroeconomic variable. This test checks the efficiency of the expectation formation. For few variables also the past of the macroeconomic variable is checked (such as GDP, CPI and own financial situation) in the barometer questionnaire. In these cases it is possible to analyse whether consumers have any bias or inefficiency in formulating their expectations. In this paper, we do not investigate what variables likely affect mostly the confidence indicators.

In principle there are caveats with respect to a third factor relating to consumer sentiment expectations and the macroeconomic variable in question (Hsiao 1982). Although in this particular setting, this kind of situation is methodologically unclear, the scope of Granger causality test is limited and should not be regarded as a proof for a strict structural relationship between variables.

The results from the Granger causality tests are presented in Table 2 pairwise, first testing the Granger causality from consumer sentiment indicator to macroeconomic variable and then checking the feedback. The tests are performed for four-quarter changes with different lag lengths covering 1–4 quarters.

Statistically the overall consumer confidence indicator (CCI) seems to predict the overall economic growth very well. This predictive power, i.e. the p-value of CCI, is almost 5 per cent even one year ahead. On the other hand, there is clearly no feedback at all from the GDP growth to CCI. One may think that this relationship is due to the large share of consumption from the overall demand and spending, but according to these tests this is not the full explanation.

Households seem as an aggregate and well-aware of the general Finnish economic situation. This can be seen from the consumer assessment of the Finnish economy with respect to actual GDP change in 1996. It was already noted that of all CCI components, it correlates the most with the past general economic performance. In this case, only the first quarter test is significant on 5 per cent level as the question is also posed in the form economic situation now compared to that 12 months ago.

CCI predicts also household consumption and especially durable consumption. This is not very surprising since it has been emphasized that purchasing durables especially with debt financing is preceded by increased confidence of one's own economic situation. In the case of durables there are some signs of feedback from consumer durables purchases to confidence. Consumers assessments about the favourable time to purchase durables is also an efficient predictor for consumer durables purchases, which is no surprise either.

Consumers can predict the changes in inflation relatively efficiently, as consumer assessment about forthcoming inflation is clearly a very powerful predictor for inflation for about 3 quarters ahead (see also Kuismanen and Spolander 1995). In this case the feedback is also clean, so consumers do not simply extrapolate inflation from past experience.⁵

The consumer barometer includes several questions about saving behaviour; the motives, saving ability, saving patterns, saving targets etc. Graphical comparison already

⁵ Quite recently, this question has been revised into a question presenting the inflation expectations directly in annual percents, which makes it easier to interpret.

shows that the question concerning the favourableness of time to save correlates closely with the interest rate development. It is interesting that, in this case, the interest rate seems to Granger-cause the favourableness of time to save, not the other way around as with most other variables. This observation may reflect the formation of the interest rate in the money market that affects, with a lag, new bank lending rates. However, it could also be the case that if favourableness of saving would indicate changes in the subjective time preference (discount factor between current and future consumption utility), therefore it should Granger-cause interest rate as well. It should be remembered though, that the interest rate is fundamentally a forward-looking variable.

The question concerning the unemployment trend also caused some problems in interpretation. It was found out that actual unemployment rate seems to Granger-cause unemployment rate changes. However, taking the difference from the unemployment rate, i.e. considering the change in unemployment rate, seems to confirm that consumers can predict these changes efficiently.

In general, it appears that consumer sentiment variables are, if not absolutely necessary in predicting the macroeconomic, at least very useful short-term indicators for several macroeconomic variables. The information sets in the bivariate Granger causality tests are, however, limited to the past of these variables, which correspond with the weak form of rationality. In principle, the information set used by households is unlimited, except that the future cannot cause the present or past. The expansion of the information set could be done by considering the role of confidence indicators as additional variables in specified behavioural equations. Therefore, we will consider specified consumption and saving models.

4.2 Spectral Analysis on the CCI and GDP Relationship

Spectral analysis could also be helpful in checking the timing or phase difference between the series. Cross-spectrum analysis can be used in studying the strength of the relationship between confidence indicator and economic activity at different frequencies. In the figures shown here, plotting against frequencies have been transformed into periods (quarters) to ease the comparison. From Figure 22 we see that the phase-shift, i.e. the timing difference between the series is around 3–4 quarters as expected, since the timing balance of CCI is about zero, but GDP growth is measured as four-quarter change. The squared coherence, which is a spectral analysis analogue to correlation, is maximized in the short-run by around 3 quarters (Figure 23). It is possible that there is some information lag in the household measurement of the economy, therefore the foreseen horizon is shorter than four quarters. This is somewhat longer than was seen in the cross-correlation function analysis. The gain measures a unit impulse in the independent variable until a new equilibrium is found. The gain could then be interpreted as the regression coefficient of the particular frequency of the independent variable CCI on the corresponding frequency component of GDP. The estimated gain also has its short-term peak around 3 quarters (Figure 24).⁶

⁶ Spectral analysis between confidence and durable purchases showed the phase-shift to be almost exactly four quarters, as it should be. The same observation can be made between inflation expectations and actual inflation.

4.3 Results from Behavioural Models

In this section we examine whether various consumer sentiment variables add predictive power to the major behavioural model, e.g. household consumption function, saving and borrowing models.

We start with the trite observation that consumer confidence correlates with economic activity. Consumer optimism and pessimism may, therefore simply vary with economic growth as purchasing power varies with labour income. As the average propensity to consume out of disposable income is over 95% in Finland, it appears that confidence varies with consumption, too. One reason that this close relationship arises is the group of rule-of-thumb consumers. Namely Campbell and Mankiw (1991) distinguished two types of consumers. Rule-of-thumb consumers are liquidity constrained consumers who consume all the income they receive. In Finland, the share of liquidity constrained consumers could be about 30–50 percent of consumer total expenditure (Takala 1995b). Those consumers not limited by current income in their spending behave more or less like rational life-cycle optimizers that can smooth their consumption for longer periods. Life-cycle consumers are not supposed to be as sensitive to variation in growth and are therefore more immune to sentiment volatility. In fact, according to Hall (1978), if every consumer was a rational life-cycle optimizer, the changes of non-durable consumption would be an unforecastable white noise.

Carroll, Fuhrer and Wilcox (1994) note that the Campbell-Mankiw model is particularly useful in this context, since consumer sentiment should affect consumption only indirectly through income. The testable hypothesis, therefore, is whether consumer sentiment has additional explanatory power for consumption, even when lagged income is included as a separate regressor.

Table 3 presents a Hall hypothesis test for Finnish quarterly data with and without consumer confidence indicator as regressor for consumption and consumption by durability. From the corrected R^2 measures it can be clearly seen that simple Hall hypothesis is strongly rejected as the CCI is a very powerful predictor for changes in consumption. Taken as a separate regressor sentiment is a much stronger explanatory factor than disposable income. Another approach more closely related to macroeconomic forecasting would be simply to test in the existing specifications for the significance of sentiment variables as additional regressors. This should tell us whether we need to pay attention to consumer sentiment variables as carriers of additional information not included in any other variable we have used in the specification.

Table 4 represents a more full-fledged version of an error-correction model with and without the consumer confidence indicator. For a longer estimation period without the CCI variables, the EC model works fairly well despite some serial correlation in the residuals, which arises from using annual differences. The error-correction term is significant and the other explanatory variables have the expected signs (for details see Takala, 1995a). Using the 4 quarter lag (or any other lag) on CCI as an additional regressor does not however improve the ex post performance of the model. The implication seems to be that CCI contains only information already included in other macroeconomic variables. This casts some doubt on the hypothesis that sentiment variables could be structural causes for observed behaviour.

We found that household expectations of their own economic situation did not correlate at all with the actual annual change of the consumption they are supposed to predict.

On the other hand, this expectation correlated strongly with the change in consumption at the time expectations were formed. This observation corresponds to that found in Berg and Bergström (1996). So in practice this expectation can be used to evaluate the current annual change in consumption, which we also lack because of lags in statistics production. Consumer views about the forthcoming year could be a coincident indicator that discounts beliefs about future permanent income and could therefore be useful in forecasting. We also tried to use consumer expectations about the next 12 months (PCI) instead of CCI, but this did not prove to be as useful (Figures 25–26).

In Table 5, we represent a saving rate model comparing for the same estimation period the significance of consumer sentiment variable concerning favourable time for saving with one lag.

Here it turned out that the sentiment variable did not include any additional information on the saving rate that was not already in the other variables (see also Takala 1995b).

Although the CCI did not prove to be very efficient in predicting the forthcoming changes in consumption, this may be an exception since according to Hall's life-cycle theory consumption is a near random walk process and therefore consumption growth should be unforecastable. Therefore consumer confidence should not contain any useful information for forecasting future consumption growth. In Table 6 we show evidence that prediction of consumer borrowing can be aided by the sentiment variable concerning the favourableness of time to raise a loan. Household assessment about how favourable the time is to raise a loan correlates strongly with the real lending rate and negatively with the favourable time to save assessment. Figure 27 shows clearly that the household assessment for borrowing was in deep trouble during the early 1990s recession. Only quite recently has it recovered, which has been accompanied by a change in borrowing patterns. The barometer questionnaire also includes questions on actual intentions for raising a loan: surely, possibly, maybe not, or no. These variables were not useful in our forecasting model. It might be the case that raising a home loan is such a rare occasion in Finland that there is sizeable measurement error to disturb the indicative power of this question.

One important variable for monetary policy targeting is inflation. In Table 7 we present a monthly inflation model using consumer barometer inflation expectations for the period 1995/m10 onwards, since the question concerning inflation expectations were revised while the consumer barometer was changed to a monthly basis (Figure 28). The other explanatory factors in this specification are money supply (M2), import prices for consumer goods, interest rate and tax tariff index. Due to joining the EU at the beginning of 1995, food prices collapsed and during the adjustment period of 1.5 years a dummy variable was used. In addition a impulse dummy for January 1994 was used. The history so far shows that barometer inflation expectations clearly have potential prediction power over actual forthcoming changes in consumer prices (Figure 29). Using inflation expectations even for this short period as an additional explanatory factor seems to have some value depending on the forecast horizon (see also Kinnunen 1996).

An adaptive element in household inflation expectations also seems to exist. This means that expectations are revised according to actual inflation and latest inflation forecast errors. The symmetry in evaluation of the past inflation and anticipated inflation also casts doubt on the view that expectations are myopic.

Table 2 Granger Causality Tests for Consumer Sentiment Indicator and Macroeconomic Series, 1987/Q4–1996/Q4

$$\Delta y_t = \sum_{p=1}^p \beta_p \Delta y_{t-p} + \sum_{p=1}^p \gamma_p \Delta S_{t-p} + \varepsilon_t$$

$$H_0 = \gamma_1 = \dots = \gamma_p = 0.$$

F-test probability value for different lag lengths

Cause	⇒	Consequence	1 Quarter	2 Quarters	3 Quarters	4 Quarters
GDP, consumer confidence (CCI) and consumer assessment about past year in the economy						
GDP		GDP	.0073**	.0515	.0286*	.0595
GDP		CCI	.4086	.8755	.3972	.6374
Cons. View		GDP	.0038 **	.0745	.0970	.2184
GDP		Cons. View	.2736	.6413	.6620	.9315
Consumption (non-durables, durables), confidence and favourableness of durables purchases						
CCI		C	.0134*	.0779	.1167	.0591
C		CCI	.8186	.5956	.3002	.0951
CCI		CNCD	.0432*	.1958	.3181	.0317*
CNCD		CCI	.8545	.9990	.5700	.1291
CCI		CD	.0002**	.0022**	.0057**	.0339
CD		CCI	.4104	.0705	.0486*	.0744
Fav. Durab.		CD	.0089**	.0040**	.0101*	.0633
CD		Fav. Durab.	.0187*	.1865	.2786	.3605
Inflation expectations and actual inflation						
Inflation exp.		CPI	.0067**	.0212*	.0230*	.0574
CPI		Inflation exp.	.2723	.5710	.5616	.3147
Interest rates (nominal and real), favourable time for saving and ability to save						
Saving fav.		RLBN	.6830	.1297	.3058	.1576
RLBN		Saving fav.	.0486*	.0200*	.0321*	.0344*
Saving abil.		ReRLBN	.0509	.5357	.7758	.7701
ReRLBN		Saving abil.	.0661	.0476*	.1374	.2001
Unemployment expectations and actual unemployment rate						
UR expect.		UR	.0000**	.2086	.4503	.2764
UR		UR expect.	.0921	.0083**	.0148*	.0329*
UR expect.		D4UR	.0000**	.0133*	.0490*	.0825
D4UR		UR expect.	.9785	.1328	.3044	.5178

Table 3 Testing the Hall Hypothesis with Lagged Consumer Sentiment about Future Consumption

Reduced-form evidence on significance of consumer sentiment and its incremental explanatory power

$$\Delta \log(C_t) = \alpha_0 + \sum_{i=1}^4 \beta_i S_{t-i} + \sum_{i=1}^4 \gamma_i Z_{t-i} + \varepsilon_t$$

	F-test for sentiment variable	R ² with sentiment only	R ² with lagged real disp. income only	R ² with sentiment and income	Incremental expl. power of sentiment
Total consumption	.0001**	.552	.010	.624	.614
Durable goods	.0024**	.424	.151	.522	.371
Non-durable goods	.0835	.255	.092	.372	.280
Services	.0686	.268	.120	.429	.309

Table 4 Consumption Function with Additional Sentiment Variable
Dependent Variable: Non-durable Consumption, %

Estimation period: 1976/Q3-1996/Q4			1988/Q4-1996/Q4			
Variable	Coefficient	t-value	t-prob	Coefficient	t-value	t-prob
Constant	4.162	6.217	0.0000	6.893	3.870	0.0007
d4IReYD	0.227	5.627	0.0000	0.189	2.549	0.0176
d4IReYD_1	0.060	1.634	0.1065	-0.025	-0.418	0.6795
d4IReW	0.060	4.174	0.0001	0.037	1.905	0.0688
ECM_4	-0.252	-5.005	0.0000	-0.159	-1.327	0.1971
RRLBN_2	-0.178	-2.940	0.0044	-0.703	-2.926	0.0074
d4UR	-0.357	-3.893	0.0002	-0.190	-1.231	0.2304
d4IPCP	-0.325	-4.459	0.0000	-0.140	-1.019	0.3183
CCI_4	0.001	0.039	0.9693			
Model performance			Model performance			
R ² = 0.8521			R ² = 0.9331			
F(7,74) = 60.921 [0.0000]			F(8,24) = 41.846 [0.0000]			
s = 0.9578			s = 0.8017			
DW = 1.33			DW = 2.46			
RSS = 67.881			RSS = 15.424			
Diagnostics		Test p-value	Diagnostics		Test p-value	
AR 1- 1F(1, 73) =		8.9093 [0.0039] **	AR 1- 1F(1,23) =		2.0647 [0.1642]	
ARCH 1 F(1, 72) =		0.1284 [0.7211]	ARCH 1 F(1,22) =		2.8498 [0.1055]	
Normality Chi ² (2) =		0.5909 [0.7442]	Normality Chi ² (2) =		0.6445 [0.7245]	
Xi ² F(14, 59) =		0.6291 [0.8300]	Xi ² F(16,7) =		0.9891 [0.5403]	
Xi*Xj F(35, 38) =		0.6538 [0.8963]				
RESET F(1, 73) =		1.0557 [0.3076]	RESET F(1,23) =		0.6234 [0.4378]	

Variables:

d4IReYD = Four-quarter change in real disposable income, %
d4IReW = Four-quarter change in real household net wealth, %
ECM_4 = Lagged (4 quarters) error-correction term from long-run static model, %
RRLBN = Real bank lending rate for new loans, %
d4IPCP = Four-quarter change in private consumption deflator, %
d4UR = Four-quarter change in unemployment rate, %
CCI_4 = Lagged (4 quarters) consumer confidence indicator

Table 5 Household Sector Saving Model with and without Additional Sentiment Variable
(Favourable time to save now, percent)
Dependent variable: Saving rate (%), 1988/Q3–1996/Q4

Variable	Without sentiment variable			With sentiment variable		
	Coeff.	t-value	t-prob	Coeff.	t-value	t-prob
d4IReYD	0.4199	5.492	0.0000	0.3928	4.953	0.0000
RePHM	-0.1153	-3.887	0.0006	-0.1112	-3.745	0.0009
UR	0.2054	3.754	0.0008	0.2424	3.865	0.0006
d4UR	0.5127	2.217	0.0349	0.3925	1.563	0.1298
RRLBN	0.5266	2.902	0.0071	0.4579	2.418	0.0226
d4IPCP	1.1837	2.835	0.0084	1.1658	2.810	0.0091
FAVOURABLE TIME FOR SAVING (ONE LAG)				0.0466	1.181	0.2478
Model performance			Model performance			
R ² = 0.907			R ² = 0.912			
s = 1.385			s = 1.375			
DW = 1.84			DW = 1.91			
RSS = 53.67			RSS = 51.03			
Diagnostics	Test	p-value	Diagnostics	Test	p-value	
AR 1- 3F(3,25)	= 2.082	[0.1281]	AR 1- 3F(3,24)	= 3.3216	[0.0367] *	
ARCH 3 F(3,22)	= 0.156	[0.9245]	ARCH 3 F(3,21)	= 0.2400	[0.8674]	
Normality c ² (2)	= 3.185	[0.2034]	Normality c ² (2)	= 3.3117	[0.1909]	
Xi ² F(12,15)	= 1.981	[0.1057]	Xi ² F(14,12)	= 1.368	[0.2964]	
RESET F(1,27)	= 1.689	[0.2047]	RESET F(1,26)	= 1.3963	[0.2480]	

Variables:

d4IReYD = Four-quarter change in real disposable income, %
RePHM = Real house price index
RRLBN = Real bank lending rate for new loans, %
d4IPCP = Four-quarter change in private consumption deflator, %
UR = Unemployment rate, %
d4UR = Four-quarter change in unemployment rate, %

Table 6 Household borrowing model, 1989/Q1–1997/Q1
Dependent variable: Four-quarter change in household debt, %

Variable	Coefficient	t-value	t-prob
Constant	17.592	10.980	0.0000
d4IPHM	0.116	2.653	0.0132
UR_2	-0.764	-8.678	0.0000
d4IPCP_1	0.461	1.541	0.1349
ATTPKUK_4-	1.275	-6.089	0.0000
KBMI14_3	0.055	2.804	0.0092

Model performance	Diagnostics
R ² = 0.9778	AR 1- 3F(3, 24) = 1.2134 [0.3263]
F(5,27) = 238.24 [0.0000]	ARCH 3 F(3, 21) = 1.1155 [0.3652]
s = 1.25424	Normality Chi ² (2) = 1.4636 [0.4811]
DW = 1.26	Xi ² F(10, 16) = 1.6533 [0.1784]
RSS = 42.4742	Xi*Xj F(20, 6) = 2.4563 [0.1344]
	RESET F(1, 26) = 41.653 [0.0000] **

Variables:

d4IPHM	= Four-quarter change in house prices, %
UR	= Unemployment rate, %
d4IPCP	= Four-quarter change in consumer price deflator, %
ATTPKUK	= After-capital tax deduction interest rate on household loans, %
KBMI14	= Favourable time to raise a loan at present, %

Table 7 Inflation model, 1991/M7–1997/M5
Dependent variable: Inflation (d12ICPI), %

Variable	Coefficient	t-value	t-prob
Constant	-0.93932	-5.575	0.0000
d12ICPI_1	0.53403	10.417	0.0000
INFLEXP_3	0.25562	2.299	0.0249
INFLEXP_5	0.17468	1.498	0.1392
EHEL3M_6	0.03570	2.674	0.0096
d12IM2	0.11063	6.212	0.0000
d12IMP90_6	0.03774	2.971	0.0042
d12IPT90	0.21127	9.494	0.0000
Eudummy	-0.24073	-2.659	0.0100
i1994p1	-0.80442	-3.950	0.0002

Model performance	Diagnostics
R ² = 0.9760	AR 1- 5F(5, 56) = 0.6810 [0.6396]
F(9,61) = 275.09	ARCH 5 F(5, 51) = 0.3896 [0.8537]
s = 0.18775	Normality c ² (2) = 1.6127 [0.4465]
DW = 1.77	Xi ² F(16, 44) = 0.4635 [0.9519]
RSS = 2.1503	Xi*Xj F(44, 16) = 0.7302 [0.7986]
	RESET F(1, 60) = 0.0628 [0.8029]

Variables:

- d12ICPI = Inflation measured as CPI annual change from previous year, %
- d12IM2 = Annual change in money aggregate (M2), %
- EHEL3M = Market interest rate (Helibor 3 months), %
- d12IMP90 = Annual change in import prices for consumer goods, %
- d12IPT90 = Annual change in tax tariff index, %
- INFLEXP = Consumer inflation expectations for next 12 months, % (0 for 1990/m1 - 1995/m9)
- Eudummy = Dummy variable for food prices(1 for 1995/M1-96/M5, 0 otherwise)
- i1994p1 = Impulse dummy for 1994/January.

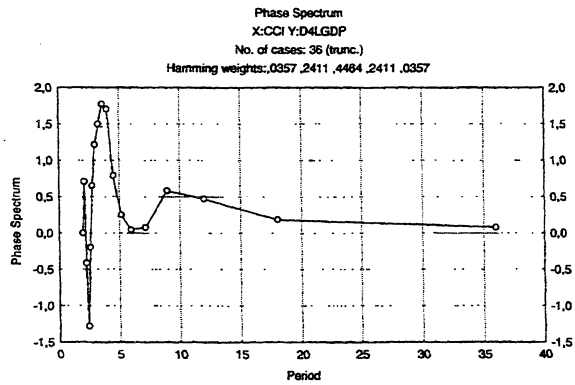


Figure 22 Phase Spectrum between Confidence Indicator and GDP Plotted against Time in Quarters

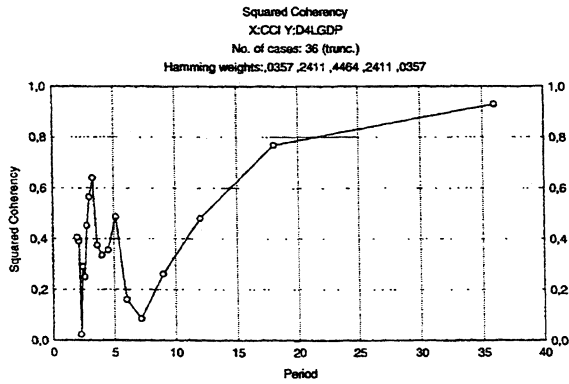


Figure 23 Squared Coherence between Confidence and GDP, %

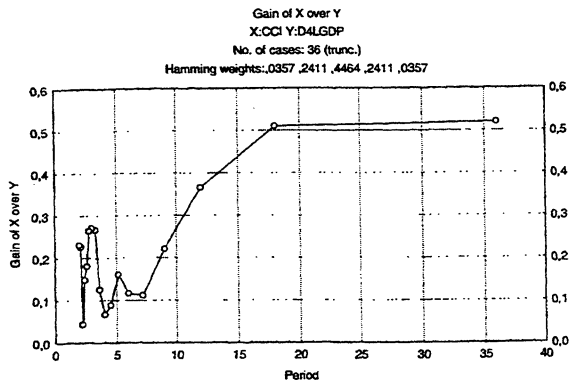


Figure 24 Gain of Confidence Indicator on GDP Growth

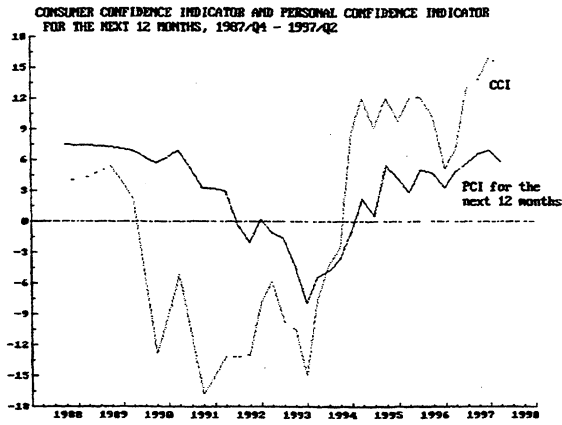


Figure 25

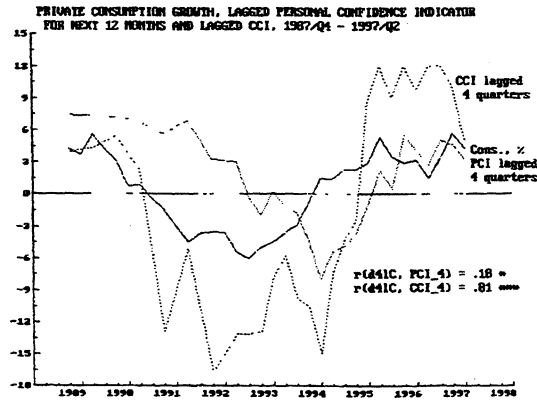


Figure 26

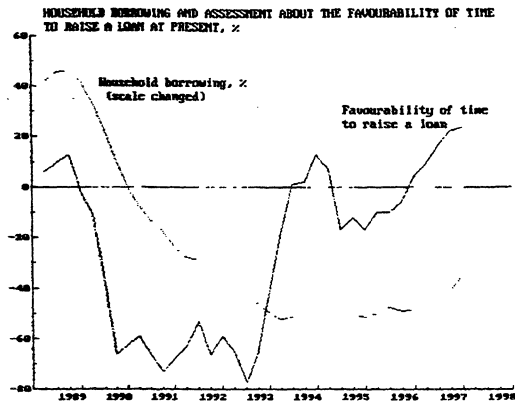


Figure 27

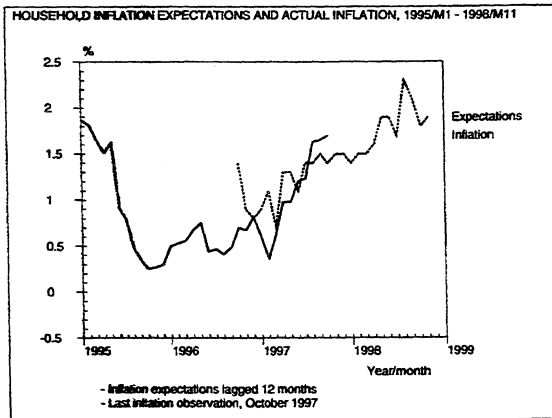


Figure 28

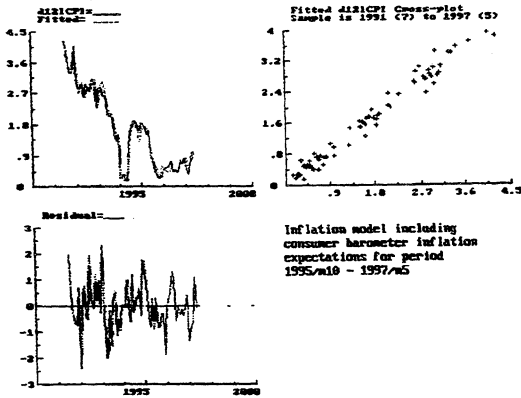


Figure 29 Inflation Model Performance, %

5 Conclusions

Our comparison of consumer sentiment variables and real economic activity have shown some cases of very close correspondence between household expectations and actually realized changes. In some cases, however, the expectation horizon does not reach as far as actually planned, but nevertheless predictive power does exist. The performed tests showed that for most variables there is a one-sided predictive Granger causality from the sentiment variable to the macroeconomic variable. This confirms that household expectations are truly forward-looking and not simply pure extrapolations from the past.

One interesting but probably exceptional case is private non-durable consumption where predictive power is limited, as assumed also according to Hall's random walk hypothesis of consumption. In contrast, for durables, predictive power is more obvious, as recorded also in several other studies.

However, consumer sentiment variables are usually more volatile than actual macroeconomic changes and can therefore produce false signals. For example, consumer confidence predicted a slowdown in growth in early 1994, which did not happen. In addition, consumers did not see the temporary growth slowdown of the Finnish economy in late 1995. The status of sentiment variables varies from expectations to intentions and plans to almost announced commitments to act in a certain way.

A more stringent test for the usefulness of consumer sentiment indicators would be to apply sentiment variables as additional variables in behavioural models. In equations such as the basic consumption function and the household borrowing equation, consumer expectations are potential explanatory factors. It cannot be ruled out altogether that sentiment variables may also have an independent influence on activity. In specifications used in short-term forecasting, expectations to raise a loan seem to be potential explanatory factor. In general it seems that sentiment variables – though not often likely structural causes for macroeconomic variables – can be very helpful in particular small-scale forecasting models and pure time-series models. For policy uses, the fact that CCI predicts consumption gives some scope for fiscal policy as consumption can be affected through government influence on household expectations. In monetary policy, household inflation expectations seem to be useful in assessing real interest rates and guiding inflation targeting.

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Testing the cointegration of house and stock prices in Finland

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TESTING THE COINTEGRATION OF HOUSE AND STOCK PRICES IN FINLAND*

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This paper looks at the efficiency of asset markets and pricing rules with regard to house and stock markets in Finland. House and stock prices are found to have unit roots, which is a necessary condition for efficient markets. However, asset prices are not pure random walks, but instead unit root processes. The unit root part is supposed to reflect the changes in fundamentals, and the error component reflects short run deviations from the market equilibrium.

Evidence on cointegration between asset prices is found. Based on the hypothesis of cointegration, an error correction model is estimated. Deviations from the long run market equilibrium can be used to improve predictions of stock and house prices, e.g. house price predictions can be improved on average by using lagged changes in the stock index and the equilibrium error as a useful indicator of disequilibrium in the cointegrated markets. The presence of such an error correction term in itself shows that asset markets are not fully efficient.

Granger causality between these two aggregate asset prices runs from the volatile stock market to the housing market rather than the opposite way. During the sample period the Finnish money market went through a gradual liberalization process that revealed an imbalance in the asset market caused partly by rationing in the rental housing market. In addition, tax incentives for owner-occupied housing and relaxed liquidity constraints significantly increased the demand for houses during the period 1987–89. Real bank lending proved to be a significant predictor for asset prices, especially for real house prices. Error correction models with additional variables due to liquidity constraints, taxation and demand for assets were used to explain the house prices.

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1. Introduction

The economic theory of capital asset pricing relies heavily on the principles of present value calculations and the hypothesis of efficient capital markets. Present value models tell us that the price of an asset is a function of the expected future yields discounted to the current date. This should apply to all assets such as stocks, land, houses and durables, since they are alternative investment objects. However, the valuations of forward-looking consumer/investors are affected by subjective factors as well, like tastes, asymmetric information, differences in time preference and degree of risk aversion.

At each moment the market demand — supply condition produces equilibrium price and quantity, which is based only on marginal amounts traded of each asset. With durable assets like houses, the prices realized on the market may not be perfect indicators of the value of the whole stock. Housing investment is also volatile and prices of new houses vary with the prices of the old stock. This may make the adjustment process slow, although prices are supposed to carry all the relevant information regarding the asset. On the other hand stock market crashes and speculative bubbles have cast serious doubts on the efficiency of financial markets in rationally reflecting the fundamentals. Therefore it might at first seem unlikely, that strong dependence would be found between markets of separate assets.

In the *short run*, asset prices vary with changes in operating cost, depreciation, alternative asset yield, taxation rules and risk premiums. However, even in the case of durable assets, changes in interest rates are not directly reflected in asset prices, as would be anticipated from the present value formula.

In the *long run*, it seems plausible that the determination of asset prices has to depend on the replacement cost of an asset, which may sometimes be hard to define for exhaustible resources — land, ground water, forest. In an efficient economy these factors of production are used to the point where the value of the marginal product equals the marginal cost of the input. Therefore, it is not surprising that the efficient market hypothesis has implications for asset prices as well. In this paper several implications of the joint hypothesis of

the asset pricing theory and efficient markets are tested. The importance of these implications has recently been rediscovered, since they can be tested with the newly discovered and rapidly developing econometric methodology of unit roots and cointegration.

In the long run, asset prices should be cointegrated, and real after-tax risk adjusted returns should coincide. There is no reason to expect the prices of two different individual stocks to be cointegrated, but since we are dealing here with broad asset markets, this may be the case. This is the question we address. Unfortunately many factors can disturb the expected cointegration, like changing risk premiums, changes in expected inflation and relative prices, tax rules and tax advantages, regulation of supply, portfolio changes induced by changes in subjective time preference, irrational expectations about future returns etc. Several proxies of these variables are used in explaining the divergence of asset prices from their equilibrium values.

Quite recently *statistical models* have been developed to describe parsimoniously the choice and demand for assets. This line of work follows the mean-variance approach to asset demand, according to which the demand for an asset is fully described by two moments of the return distribution, namely the expected return and variance. This approach is taken in the strict version of capital asset pricing model (CAPM), which has thus far been the dominant model.

For example, the efficient market hypothesis applied to present value models with rational expectations implies that asset returns and hence asset prices, should follow a random walk, possibly with drift. Empirically this means that changes in asset prices cannot be predicted with current information apart from the drift. On the other hand, under efficient markets asset prices should be cointegrated, which means that their price paths should not diverge in the long run (Bossaerts, 1988). The pure random walk hypothesis and asset price cointegration are contradictory, since both cannot prevail at the same time (Granger, 1986). If cointegration prevails, this means that the equilibrium error can be used in predicting the corresponding target changes, which cannot hold under the pure random walk property.

We start our analysis with a description of

the data and recent phenomena in the development of asset prices and their determinants, in particular, the effects of the liberalization of the money market which occurred in the late 1980's. Then we proceed in reviewing the implications of the CAPM in tests of prices. In the econometric section of the paper, we review the unit root and cointegration tests proposed in the literature and apply them to our data. After presenting some evidence on cointegration of asset prices, we move to the question of possible causes of short run distortions in the asset market. We try to explain the short run equilibrium error between housing prices and the general stock index with general economic fundamentals and more specific determinants of asset prices. Finally, we construct and test ECM model for house price changes, based on cointegration of asset prices with some additional explanatory factors.

2. Description of the asset market and data

Two basic asset series: housing prices and the general stock index for Finland were analyzed with quarterly data of 1970/1 – 1990/2.

The choice of assets was motivated by their availability and overall importance as investment targets. The house price index for the whole country was used, although the houses in the Helsinki area could be more close substitute for stocks (see also Suoniemi, 1990). The stock index used is the Unitas index of the biggest commercial bank (SYP) in Finland. We have generally considered the stock market to be more efficient than the housing market because of its more rapid reactions to new information, greater volatility and smaller transaction costs. To find out whether this is true, we first review some major aspects of asset prices and their determinants from 1970.

From the nominal prices it is hard to see a close relation between the price levels, only a clear upward trend (figure 1). The real prices of houses and the stock index show already evidence of a long run linear relation, especially if we match the ranges i.e. making a linear transformation for house prices (see figures 2 and 3). The real prices were calculated by deflating nominal prices with the consumer price index (CPI). The base year in figures is 1972. During the period 1972–90 the real price curves intersected about four times. The dramatic increase since 1986 in stock index and, a little later, housing prices

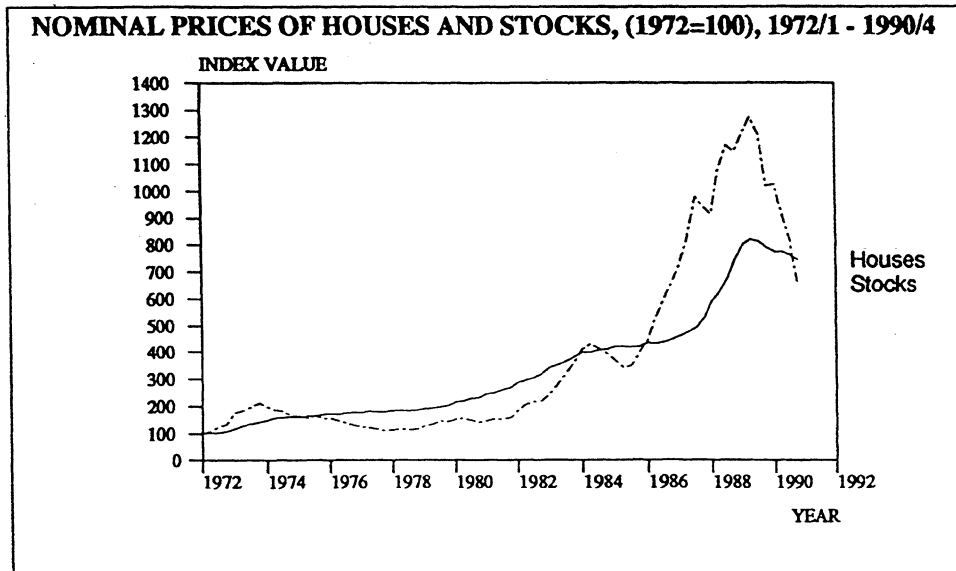


Figure 1.

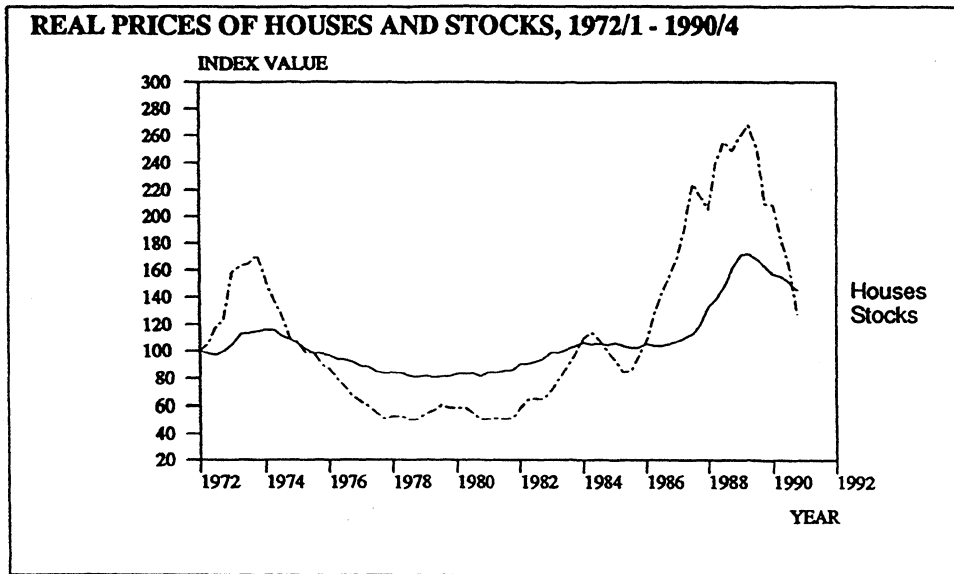


Figure 2.

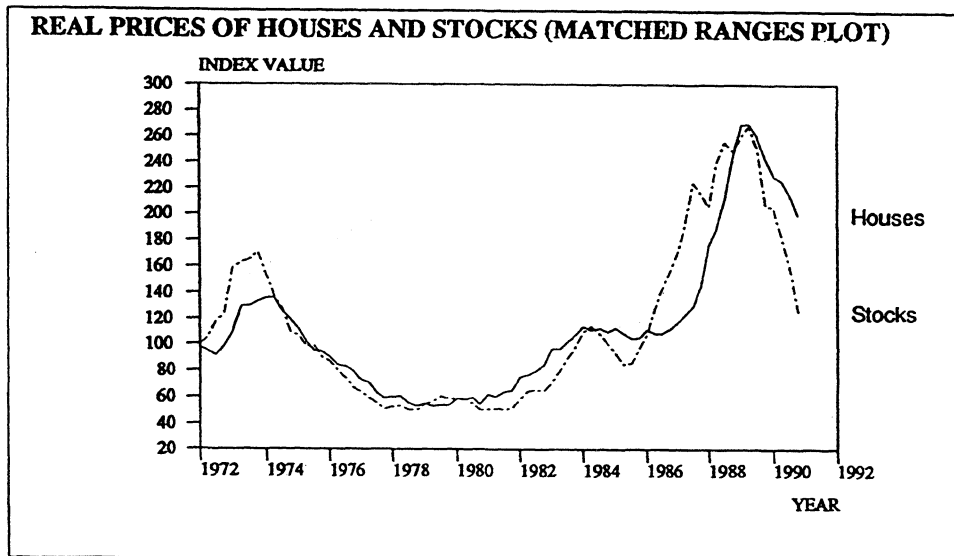


Figure 3.

was mainly a consequence of capital market liberalization and increased bank lending thereafter. The October 1987 stock market crash was not very deep in Finland, and the level of general stock prices recovered to the

pre-crash level already in spring 1988. From spring 1989 real price indexes have fallen as a consequence for a similar overheating like in Norway, West-Germany and Sweden.

The relation between real bank lending and

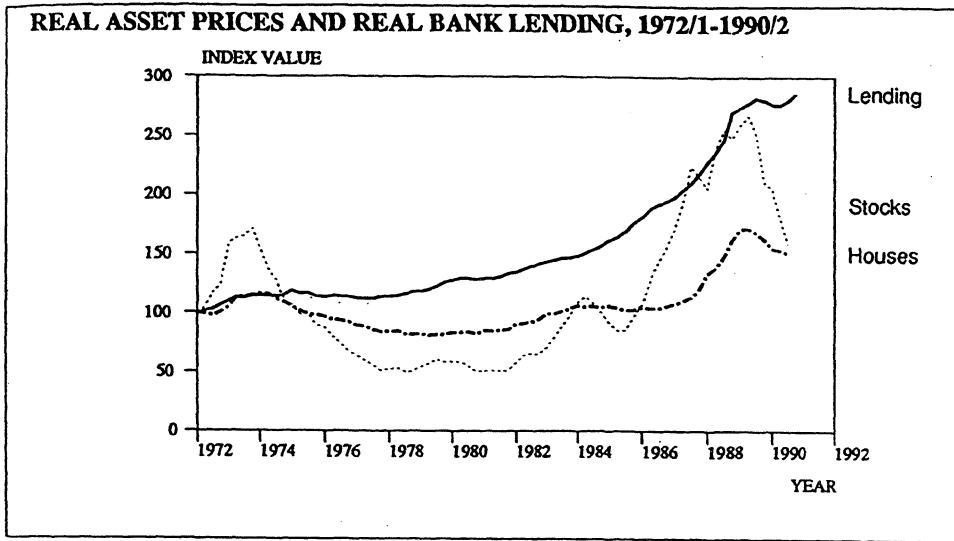


Figure 4.

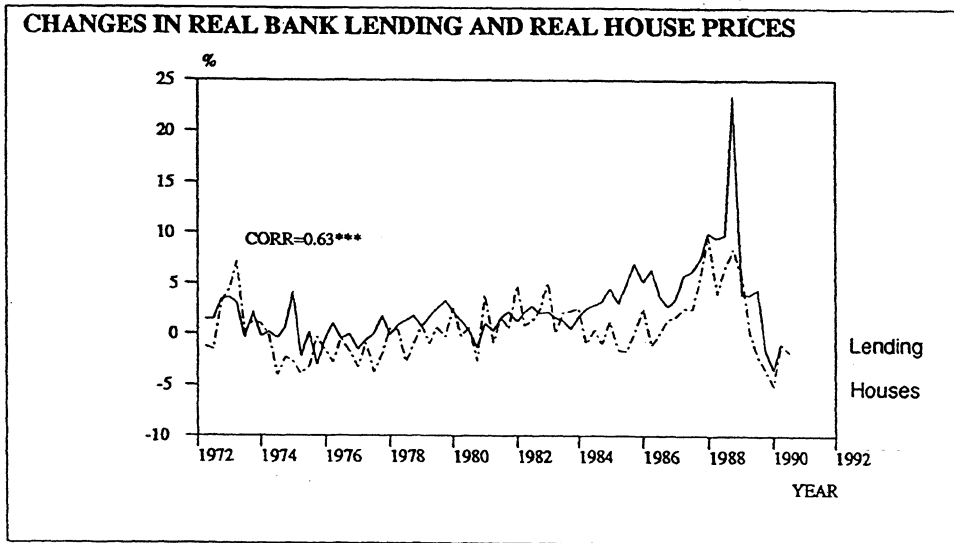


Figure 5.

real house prices, deflated with the consumer price index, can be seen from figure 4. The differences of real bank lending and asset prices can be seen from figures 5 and 6. The instantaneous correlation between real bank

lending and real house price difference was even as high as 0.63. The liberalization of the money market revealed the imbalance in the rationed rental housing market. The very drastic symptom of this was the large demand in-

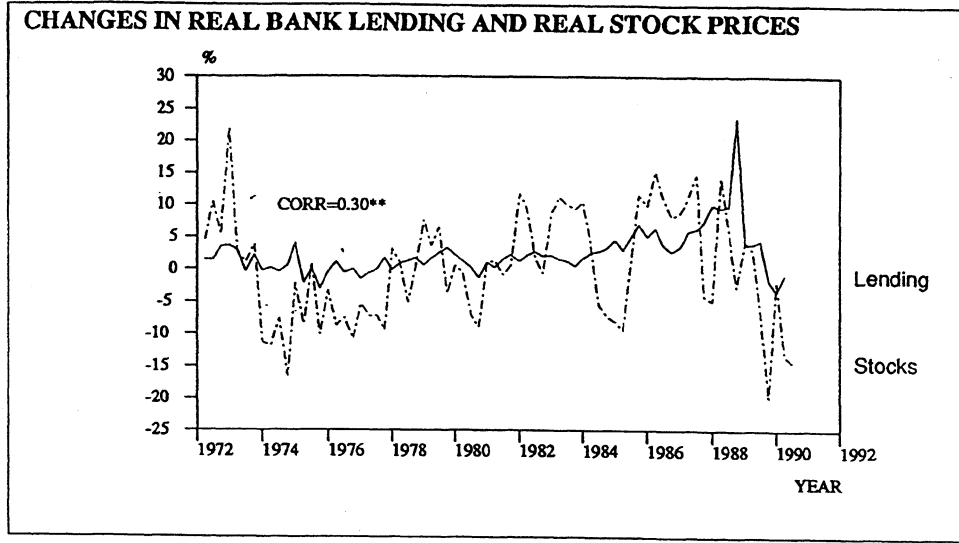


Figure 6.

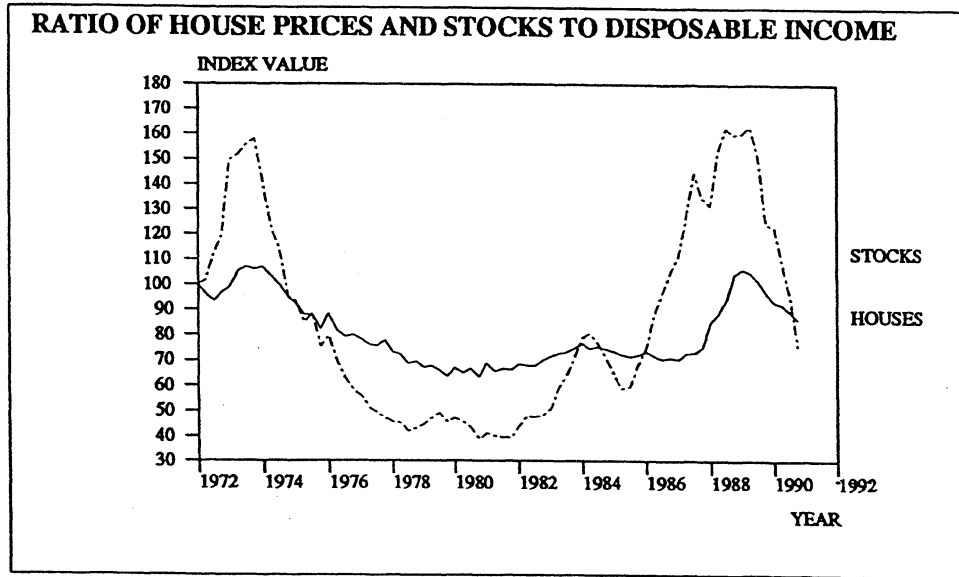


Figure 7.

crease for owner-occupied housing. The increase in demand and bank lending went directly into asset prices, since supply was almost inelastic. The Finnish rental housing market has been rationed since 1960's and tax

incentives for owner-occupied housing have risen the share of owner-occupied housing since then from 56 % up to over 70 %.

The ratios between nominal house prices and nominal stock index to household dis-

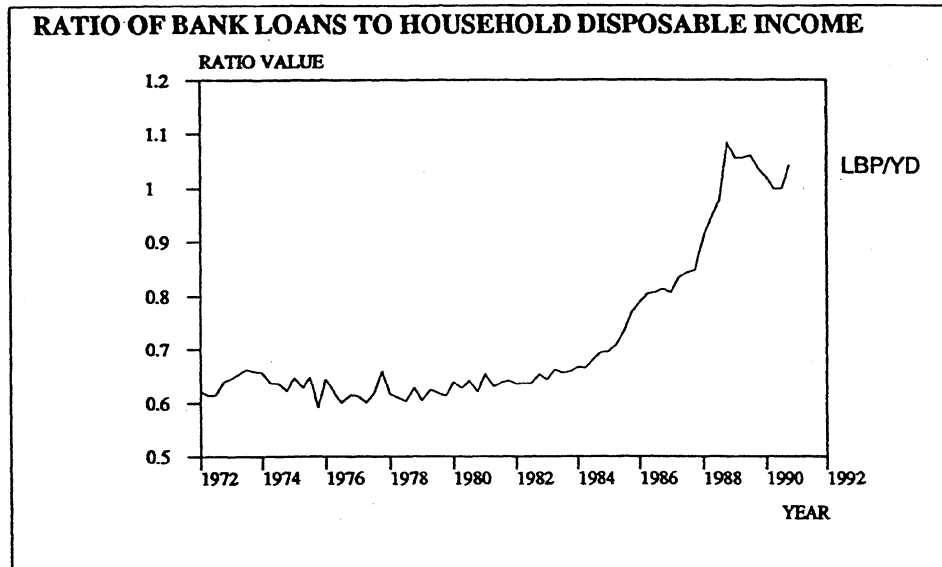


Figure 8.

posable income can be seen from figure 7. In the beginning of 1970's house prices went up following a demand boom. After the 1974 oil crisis, house prices decreased steadily in real terms with respect to income till the end of 1970's. Real construction costs have been quite stable and construction costs and consumer prices increased at about the same growth rates up to 1987.¹

The liberalization of capital markets and foreign capital inflow through firms and banks (households are not yet able to raise foreign loans directly) is the primary candidate to explain the rapid stock and house price rises since 1986. The share of housing loans in total bank lending has however fallen slightly, as consumption loans have risen even faster. Because the value of the housing stock

has risen rapidly, the ratio of housing loans to the value of housing stock has increased only slightly. Nevertheless, it is clear that the debt/asset ratio has been increasing. Figure 8 presents the ratio of outstanding bank loans to household disposable income. The planning horizons of households may have lengthened due to money market liberalization, since liquidity constraints have eased and borrowing against future earnings is easier.

3. Implications of capital asset pricing model

The aim of asset pricing theories is to explain why different assets have different risk premiums. Therefore, the purpose of the capital asset pricing model is to show how expected returns of portfolios are also functions of the riskiness of assets (Rothschild, 1985). The riskiness of an asset is usually defined as the variance of the distribution of the return. The investor's choice of a portfolio is based on risk adjusted expected returns. The core of portfolio theories is that the risk of an asset depends on the return of other assets as well. In this respect CAPM relies on the concept of

¹ It has been suggested by the construction firms that house prices increase, because construction cost rise. This is not quite true, since change in the prices of newly built houses follow changes in the prices of the existing stock, which depends on the overall demand — supply situation. Therefore, the causality runs more likely from housing prices to construction costs (wages, materials) and land prices. Tobin's q approach seems to work quite well for the Finnish housing market (see Takala and Tuomala, 1990).

efficient markets. In fact, in an efficient market, the return on an asset includes a premium for risk that cannot be removed through diversification. More importantly the risk premium is proportionate to the covariance between asset and market returns, not to the variance of the asset itself.

Capital asset pricing theory has many applications in economic modelling. Present value models are often used in assessing the profitability of investment projects, but they can be applied as well to the valuing the price of assets — houses, land, firms and durables. The standard version of the CAPM assumes a number of requirements, like returns should be normally distributed, borrowing and lending possible at risk-free rate and there exist no transaction costs or taxes, which restricts the scope of the theory.

One important implication of the CAPM is that the observed risk premia should be fully explained by the covariance between asset returns and market return. However, in practice we have acknowledged evidence on incompleteness and inefficiency in capital markets and there is substantial evidence on the volatility clustering of asset price changes. There is no direct reason for risk or risk premiums to be constant over time. Still, risk is the main determinant of asset price changes. The varying risk premia over time is due to unanticipated interest rate movements. This risk can be measured by the conditional variance of an asset return. Another implication of the CAPM under constant absolute risk aversion is that the observed mean return and variance of return should be positively correlated across assets (Engle, Lilien and Robins, 1987 p. 394). To take into account the variation in risk, we included the conditional variance of the stock index model into the error correction models.²

² As the CAPM is based on the idea that asset portfolios are held as functions of expected means and variances of the rates of return, shifts in asset demand imply changes in both of these determinants (Engle, 1982, p. 989). The standard CAPM should be a good approximation to asset pricing when the marginal utility of consumption is highly correlated with the return on the stock market, or more generally with the portfolio of tradable assets. Since human capital is not fully tradable, relaxing the liquidity constraints should make assets more sensitive to interest rates and return. The efficiency of the as-

We do not give any full representation of the CAPM model here. We merely list some of the possible ways to test implications of the theory. For a more extent exposition see Elton & Gruber (1987) or Rothschild (1985). In the empirical part, we measure the return of an asset by changes in the asset price as an implication of the present value model (see Campbell and Shiller, 1988).³

Testable implications:

1) The CAPM is based on the idea that all investors choose mean-variance efficient portfolios. Expected return on a risky asset is therefore proportional to the non-diversifiable risk in an efficient market. Risk is measured by the covariance of asset return to market portfolio return. Market portfolio is defined to be the linear combination of all assets, that satisfy the efficient mean-variance condition. Therefore CAPM implies that the asset risk premium is determined by the diversifiable risk, which is measured by the covariance with the market returns and which is called beta, in reference to the regression slope (Bollerslev et. al., 1988 p. 116).

The expected return (change in the asset price) of an asset i is

$$(3.1) \quad E(r_i) = r_f + \beta_i [E(r_m) - r_f],$$

where $\beta_i = \text{Cov}(r_i, r_m) / \text{Var}(r_m)$ and r_f = return on risk-free asset r_m = return on market port-

set market or the rationality of agent behaviour can be tested by calculating the correlation between the market portfolio return and the marginal utility of consumption. The consumption CAPM implies that the returns on market portfolio and marginal utility of consumption should be perfectly negatively correlated (Blanchard and Fischer, 1988 p. 509).

³ The present value model expresses the price of an asset as a deterministic function of expected future yield. With finite time yield can be separated into capital income and appreciation. Here, we assumed that asset is held forever and change in asset price includes all the changes in the discounted future capital income (dividends or imputed rent on housing). All the relevant information about yield is contained in the expectation of an asset price, since the spread between the true asset price and its expectation is stationary. If the yield follows a random walk, so would the asset price, when discount rate is constant. Campbell and Shiller (1988) consider the conditions, when yield and asset price are cointegrated.

folio. In other words, we have the usual expression for efficient portfolios

$$\begin{aligned} \text{Expected return} &= \text{risk free rate} \\ &+ \text{risk premium.} \end{aligned}$$

Risk premium is proportional to beta, which reflects the sensitivity of an individual asset to movements in the market portfolio return. Thus the CAPM predicts that investors in security i will be rewarded only for market related risk.

2) Engle et al. (1987) suggest that the degree of uncertainty in asset returns varies over time and any increase in the expected return can be identified from the risk premium. Investment decisions are always based on expectation, conditional on some undefined information set. We may restrict the information set to current and past values of returns, but it is unlikely that asset price processes will be stationary over time periods long enough to estimate expected returns with any accuracy. The constancy of risk premium and the variability in β can be tested with iterative least squares estimation in asset equation above.

If in the CAPM, the regression coefficient β changes over time, we are faced with the problem of non-constant conditional variance of the error term. This can be analyzed by means of ARCH models. One period rate of return should be ex ante unforecastable in mean, i.e. asset price levels are martingales, but if an ARCH-M model can be identified, economic gain could be achieved through changes in diversification (Engle, 1988).

Portfolio theory emphasizes the interdependence between the risk of particular asset and the risk of other available assets. The riskiness of an asset depends ex post on how much the actual return differs from the expected return. For a riskier asset the required rate of return is higher because of the higher risk premium. But over the efficient market risk, adjusted returns should be equal. Therefore different observed nominal returns simply reflect different expectations of risk.

Since subjective expectations cannot be directly observed from market data, we may assume that rational economic agents do not make forecasting errors on average in assessing this risk. Hence at the market level, expected risk is equal to observed risk. But we cannot assume that the variances of expected

and actual returns coincide. We may assume that the variance of the expected return is smaller than the actual variance, since predictions on return are usually based on a smaller information set. Predictions are made for the systematic part of the actual variation and since there is always some positive error variance, predictions are conservative. In addition to information processing and analyzing, the econometrician's information set is smaller than that of the market participants.

We sum up the hypotheses to be tested as follows;

a. Random walk property and unit roots

H_0 : Asset returns follow a random walk, and asset prices a random walk with drift⁴

Asset returns should follow random walk process under informational efficiency; therefore,

$$(3.2) \quad r_t = r_{t-1} + e_t, \text{ where } e_t \sim \text{NID}(0, \sigma^2).$$

Asset prices should be random walk with drift (μ), which can be tested with

$$H_0: r_t = \mu + r_{t-1} + e_t \text{ against}$$

$$H_1: r_t = \mu + bt + \phi r_{t-1} + e_t,$$

where b is the slope of the deterministic time trend and $\phi < 1$. The appreciation of asset prices were used as a proxy for the returns, although dividends is a form of return for stocks in the short run as well.

b. Cointegration and Granger causality

H_0 : CAPM implies the cointegration of asset prices, where all risky assets are held in the same proportion, in addition risk adjusted returns should coincide (in the long-run) with the return of a riskless asset (Rothschild, 1985).

Ordinary asset pricing models like the CAPM imply complete separation, which

⁴ Exceptions are also possible. For instance, the Lucas (1978) asset pricing model implies that stock prices do not have to have the martingale property, except under risk neutrality (Michener, 1982 p. 166).

makes asset prices perfectly collinear. However this result can usually be rejected with empirical data. If we relax the assumption about perfect separability and replace it by the weaker separating equilibrium, we may preserve approximate collinearity (Bossaerts, 1988). This leads to cointegrated asset prices, which allows for weakly dependent pricing errors. The rise of price of an asset will lead to an increase in the portfolio share of this asset, but the ratio $p_{it}^* s_i / p_{jt}^* s_j = w_{ij}$ should be time-invariant, where s is supply of asset and i, j denote the assets. Unfortunately supply for an asset is rarely known in practice. For housing the supply may be approximated with current stock plus housing investment, but for stocks new share issues are very flexible and it may be hard to define supply.

H_0 : cointegration implies that the stationary equilibrium error term should Granger-cause at least one of the cointegrated variables (Engle & Granger, 1987, Campbell and Shiller, 1988 p. 507).

For finding Granger's predictive causality between variables, non-cointegration should be checked first. If cointegration exists, the equilibrium error should be added as an additional regressor to the causality tests of stationary variables.

4. Econometric methodology

We use the increasingly popular cointegration methodology. In this section we review very briefly the concepts of integration and cointegration and the related tests that we use. See for example Diebold and Nerlove (1988) and Pere (1990) for more elaborate surveys.⁵

⁵ A time series x_t is said to be integrated of order d , if it has a stationary and invertible ARMA representation after differencing d -times. If this is so, we denote $x_t \sim I(d)$. Many economic time series are supposed to be $I(1)$ or some times $I(2)$ series, see for example Nelson and Plosser (1982) and Hall (1986).

The elements x_{it} of a N -variate time series x_t are cointegrated of order d, b if (i) each $x_i \sim I(d)$, $i = 1, \dots, N$, and (ii) there exists a vector $\beta (\neq 0)$ such that $\beta' x_t \sim I(d-b)$. We denote this by $x_t \sim CI(d, b)$. That is time series are cointegrated, if they are integrated of the same order and some linear combination of them is integrated of a lower order than the original series.

The importance of cointegration is due to the empirical finding that many time series resemble $I(1)$ time series and the fact that economic theory often suggests that the levels of the time series should be closely connected. Thus we see that many economic time series should be $CI(1, 1)$. In addition, according to the so called Granger representation theorem [Engle and Granger (1987)] vector x_t has an error correction representation, if the variables are cointegrated. So we have a handy way to model the series, if the data tell us that the series are cointegrated.

We use these concepts to test if stock and house prices are cointegrated. First we explore the order of integration of these series. We use the Dickey-Fuller -tests (Dickey and Fuller (1979)) for this purpose. Their idea is very simple. Let us assume that x_t is an $AR(1)$ process

$$(4.1) \quad x_t = \phi x_{t-1} + u_t,$$

where $u_t \sim NID(0, \sigma^2)$. We can reparametrize (adding $-x_{t-1}$ on both sides) this as

$$(4.2) \quad Dx_t = (\phi - 1)x_t + u_t.$$

This equation could be expanded by adding seasonal terms and linear time trend

$$(4.3) \quad Dx_t = \mu + bt + \sum_{s=1}^3 \Phi_s Q_s + (\phi - 1)x_t + u_t.$$

Our null hypothesis is that x_t is $I(1)$ or $\phi = 1$. If the null is correct, then in the regression (4.2) the estimate of $r = (\phi - 1)$ and its t -value should be near zero. If the t -value differs significantly from zero, we reject the null hypothesis and conclude that $r < 0$ or $\phi < 1$ or equivalently that x_t is stationary. Otherwise we accept the null hypothesis. This is the Dickey-Fuller test (DF).⁶

If it turns out that our data has to be modeled with an $AR(p)$ autoregression, where $p > 1$, to whiten the residuals u_t :

$$(4.4) \quad x_t = \sum_{i=1}^p \phi_i x_{t-i} + u_t.$$

⁶ In our regression here the appropriate distribution theory is nonstandard because of the nonstationarity of x_t . However, we can use the t -value as a test statistic, because the textbook of Fuller (1976) includes its critical values as simulated by Dickey.

This again can be reparametrized to

$$(4.5) \quad Dx_t = r_1 x_{t-1} + \sum_{i=2}^p r_i Dx_{t-i+1} + u_t,$$

where $r_1 = (\sum_{i=1}^p \phi_i - 1)$ and $r_j = -\sum_{i=j}^p \phi_i$, $j = 2, \dots, p$. We can still test the hypothesis null $x_t \sim I(1)$ with the help of the Dickey-Fuller test. The null is easily seen to be equivalent to $r_1 = 0$. The augmented Dickey-Fuller test (ADF) statistic is the t -value of r_1 . The test is otherwise exactly the same as the ordinary Dickey-Fuller test. The critical values in Fuller (1976) apply to this statistic also, if the sample is big enough. If we accept the null hypothesis, we go on to test for cointegration.

We use these same tests to explore the cointegration of our data. After checking that house prices (y_t) and stock prices (x_t) are $I(1)$ we estimate the so called cointegrating regression

$$(4.6) \quad y_t = ax_t + u_t.$$

Now if y_t and x_t are to be cointegrated u_t has to be stationary. We can test this by applying the DF and ADF test to u_t , now called DFR or ADFR respectively. According to the null y_t and x_t are not cointegrated or u_t is integrated. The alternative is that y_t and x_t are cointegrated or u_t is stationary. The test goes exactly as the ordinary DF and ADF tests except that we cannot use the tables in Fuller (1976), because the residuals \hat{u}_t are estimated. Due to the properties of ordinary least squares, the estimated residuals look more stationary than the original residuals and this changes the critical values of the DF and ADF tests. Engle and Granger (1987) have simulated some critical values. We use them as approximate critical values for our time series, which has become a common practice in empirical studies of cointegration. Small sample critical values for ADF test can be found in Blangiewicz & Charemza (1990).

We report also the Durbin-Watson statistic (CRDW) of the regression (4.6). It is a quick check of the first order of autocorrelation of u_t , because $CRDW \approx 2(1 - \hat{r})$, where \hat{r} is the first order sample autocorrelation of u_t . This gives us some indication of the probability of cointegration of y_t and x_t . DW is usually only slightly above zero if y_t and x_t are not cointegrated. If u_t were an AR(1)

process, it could be used as a test for cointegration (see Engle and Granger (1987)). Usually, as in our data, this is not the case, so we don't refer to it as a proper test statistic. In addition the Johansen (1988, 1989) tests for cointegration vectors, available in PC-GIVE, were calculated.

The implications for Granger causality in cointegrated data could be done as follows. If y_t is the housing price index and x_t stock index, then

$$(4.7) \quad z_t = [a_1 \ a_2] [x_t \ y_t]' = a_1 x_t + a_2 y_t$$

The proposition of one-sided Granger causality from equilibrium error term z_t to Dx_t can be seen from the regression

$$(4.8) \quad Dy_t = \mu + \sum_{j=1}^p b_j Dy_{t-j} + \sum_{j=1}^p w_j Dx_{t-j} + \sum_{j=1}^p q_j z_{t-j} + e_t,$$

where the null hypothesis for Granger non-causality is $H_0: q_j = 0$, for all j . The Granger causality implication is a consequence of the expectation theory, namely that the price of an asset is a weighted sum of expectations of the future capital income and appreciation.

We want to emphasize that the results of these various tests are tentative. The tests are known to have low power against very autocorrelated but still stationary alternatives (Dickey and Fuller (1979) and Engle and Granger (1987)). Further, the test size may be distorted by changes in the process generating the data (Hendry ja Neale (1990)) and MA-terms in the series (Schwert (1989)). And last but not least there is the question of the quality of the house and stock indexes. With these caveats in mind let us proceed to the empirical analysis.

5. Empirical results

We begin with a description of the unit root and cointegration characteristics of the series under consideration. Then several Granger causality results are given, after which an estimated ECM is presented, along with its implications regarding equilibrium error.

5.1 Results from unit root tests

The sizes of the AR(1) coefficients for the price series immediately suggested the presence of unit roots in logarithmic levels (table 1). The DF and ADF tests confirmed this conclusion. Despite this, the residuals of these autoregressions were clearly autocorrelated. Therefore prices are not pure random walk processes or martingales. In the augmented Dickey-Fuller test five lags of the dependent variable were

added to the equation to eliminate autocorrelation in the residual.

Table 1 presents different versions of unit root tests with constant linear time trend and seasonal dummy variables for house and stock prices. Both DF and ADF tests showed no evidence against unit roots in the price level series. For log differences of the asset prices unit roots are rejected. Surprisingly both asset price series included significant seasonal variation. For the stock index the first

Table 1. Unit root tests for asset prices;
Dickey-Fuller and Augmented Dickey-Fuller tests
Sample period: 1970/1–1990/2, 82 observations

Variable	RLPHM	RLUNI	DLRPHM	DLRUNI
Dickey-Fuller regressions:				
$x_t = \mu + \phi x_{t-1} + u_t$				
Reg. coeff. (ϕ)	1.02	0.99	0.54	0.54
DF -test	0.83	-0.32	-4.80***	-4.70***
Order of integration	I(1)	I(1)	I(1)	I(1)
$x_t = \mu + bt + \sum_{s=1}^3 \Phi_s Q_{st} + \phi x_{t-1} + u_t$				
Reg. coeff. (ϕ)	1.00	0.99	0.51	0.54
b (t -test)	1.99*	1.07	1.17	0.29
Q_s (t -tests for seasonals)	—	2.17 Q_1	-2.90 Q_2	3.67 Q_1
DF -test	-0.26	-0.73	-4.95***	-4.67***
Order of integration	I(1)	I(1)	I(0)	I(0)
Variable	RLPHM	RLUNI	DLRPHM	DLRUNI
Augmented Dickey-Fuller regressions:				
$x_t = \mu + \phi x_{t-1} + \sum_{j=1}^5 g_j D x_{t-j} + u_t$				
Reg. coeff. (ϕ)	0.98	0.98	0.55	0.53
ADF -test	-0.92	-1.28	-3.40**	-3.19**
Order of integration	I(1)	I(1)	I(0)	I(0)
$x_t = \mu + \phi x_{t-1} + bt + \sum_{s=1}^3 \Phi_s Q_{st} + \sum_{j=1}^5 g_j D x_{t-j} + u_t$				
Reg. coeff. (ϕ)	0.98	0.97	0.51	0.55
b (t -test)	1.21	0.74	1.34	0.30
Q_s (t -tests for seasonals)	—	3.03 Q_1	-2.02 Q_2	3.01 Q_1
ADF -test	-1.01	-1.45	-3.33**	-3.02**
Order of integration	I(1)	I(1)	I(0)	I(0)

Abbreviations in prefixes of variables are L = logarithmic and D = difference, R for real (deflated with the consumer price index). PHM refers to the index of housing prices in the whole Finland and UNI to the UNITAS stock index of the Helsinki stock exchange.

quarter of the year has the highest value; in house prices, change slows down or becomes negative in the second quarter. Except for the real house price level in the DF test, there was no indication of deterministic time trend.

The logarithmic transformation dampened the exponential rise in asset prices during the financial market liberalization period. Since our estimation period covered only 20 years, we cannot be sure whether this simply indicates an exceptional period. The long upward swing of stock and house prices in the 1980's after the oil crises was in any case an excep-

tional period. Some part of it dues to increase in money supply throughout the western world. It seems however, more reasonable that asset prices follow a stochastic rather than a deterministic time trend.

5.2 Cointegration tests

Results of the cointegration tests (CRDW, DFR and ADFR) are shown in table 2. For a quick check the CRDW test between the price of houses and stock price index levels gave no indication of cointegration. Although

Table 2. Cointegration tests for asset prices; Cointegration Durbin-Watson, Dickey-Fuller and Augmented Dickey-Fuller tests, Sample period: 1970/1 - 1990/2

Dependent variable:	RPHM	RUNI	LRPHM	LRUNI
CRDW	0.23	0.25	0.18	0.18
Dickey-Fuller Regression				
$Du_t = -ru_{t-1} + e_t$				
Reg. coeff. (-r)	-0.08	-0.10	-0.06	-0.08
DFR	-1.37	-1.83	-1.18	-1.58
Augmented Dickey-Fuller Regression:				
$Du_t = -ru_{t-1} + \sum_{j=1}^4 b_j Du_{t-j} + e_t$				
Reg. coeff. (-r)	-0.30	-0.26	-0.25	-0.22
ADFR	-4.29***	-4.22***	-3.94***	-3.80***

The critical values for 100 observations are presented in Engle & Granger (1987) and for smaller samples in Blangiewicz & Charemza (1990, p. 306).

The Johansen (1988, 1989) tests of the rank of cointegration
Tested variables: LRPHM, LRUNI

H (rank)	Eigenvalue μ_i	Maximal eigenvalue		Trace test $-T \sum \ln(1 - \mu_i)$	Crit. value $\mu_{trace(.95)}$
		Test $-T \ln(1 - \mu_i)$	Crit. value $\mu_{max(.95)}$		
$r \leq 1$	0.035	2.72	8.08	2.72	8.08
$r = 0$	0.218	18.98	14.60	21.69	17.84

the DW statistic in cointegration regression is smaller than R^2 , this is not a case of spurious regression, but merely a result of the fact that regression and correlation measures the co-movement of variables. Cointegrating regres-

sions without R^2 close to unity should be viewed with caution. Our R^2 on regressing logs of real house prices on corresponding stock index was 0.84.

However, taking residual autocorrelation into account inference, the ADFR test with 4 lags, results in strong rejection of non-cointegration between both real asset prices and their logs. This rejection, i.e. the acceptance of cointegration, was significant with 1 % level in this sample. Also the Johansen (1988, 1989) type of cointegration tests accept cointegration between asset prices with 5 % significance level (table 2).⁷

It seems highly probable that cointegration would be a useful working hypothesis. The relation between asset prices can be interpreted as an equilibrium process, that takes about 1–3 quarters to adjust to a smaller shock in the asset market. Since the equilibrium mechanism seems to be working rather slowly, the error correction term is nearly non-stationary.

Table 3. Granger causality tests for asset prices.

Dependent Dy_t	Explanatory Dx_t	F -test for restriction ¹	
		without null lag	with null lag
Changes in real asset prices			
DLRPHM	DLRUNI	2.11*	1.84
DLRUNI	DLRPHM	1.78	1.56
Real asset prices and real lending			
DLRPHM	DLRLBP	2.72*	5.51**
DLRLBP	DLRPHM	1.76	4.52**
DLRUNI	DLRLBP	1.59	2.16
DLRLBP	DLRUNI	1.91	2.43*
Real asset prices and error correction term			
DLRPHM	ECM	3.45**	12.21**
DLRUNI	ECM	1.77	10.30**
House prices and interest rates			
DLRPHM	RLB	1.31	2.03
DLRPHM	RB	0.39	1.40
DLRPHM	DRLB	1.38	1.78
DLRPHM	DRB	1.90	2.73*
Stock prices and interest rates			
DLRUNI	RLB	1.15	1.57
DLRUNI	RB	2.47*	2.07
DLRUNI	DRLB	1.89	2.05
DLRUNI	DRB	2.44*	2.09
House prices and construction costs			
DLRPHM	DLRPIH	0.35	0.47
DLRPIH	DLRPHM	2.60*	2.34*

¹ Granger causality tests were performed as bivariate autoregressive-distributed lag models (max lag used was 5) as

$$Dy_t = \alpha_0 + \sum_{i=1}^5 \alpha_i Dy_{t-i} + \sum_{j=0,1}^5 \beta_j Dx_{t-j} + h z_{t-1} + e_t$$

with quarterly data. The results for null hypothesis (Granger non-causality from x_t to y_t) $H_0: h = \beta_0 = \beta_1 = \dots = \beta_5 = 0$ can be seen above (see Hendry, 1989 p. 36). F -statistic had degrees of freedom $F(5,65)$ without null lag and $F(6,64)$ with null lag, when cointegration term was not present. With cointegrated variables (y_t, x_t), the error correction term (z_{t-1}) also has to be included in the regressions (see Granger, 1988).

5.3 Granger causality tests

Granger causality (GC) tests were performed for stationary dependent and explanatory variables in a single equation context, since strong exogeneity of the regressors with respect to the parameters of interest is based on the assumption of Granger non-causality from past and present values of the dependent variable. With cointegrated variables this predictive causality can also run through the error correction term, which could be invalidated unless the error correction term is taken into account. The results of the GC tests are presented in table 3.

From visual inspection of the asset prices and Granger causality tests it seems probable that causality runs from the stock index to the house prices rather than vice versa (figure 2).

⁷ Unit root and cointegration tests were also performed with annual data covering 1965–1989 (25 observations). The only house price data (namely bank financed houses in the Helsinki area) available for this period was found in a study of C. Bengs (1989). However, the cointegration tests were quite similar for these series. The CRDW test result was just below the critical rejecting non-cointegration level ($DW = 0.32$ with LPHM regressed on LUNI), but the DFR and ADFR tests indicated rejection of non-cointegration.

This conclusion is further evidenced by the fact that the error correction term Granger causes house prices even at the 1 % significance level. This interpretation accords with the usual assumption of the stock market reacting rapidly in changes in economic fundamentals. House prices seem to be more rigid.

Our hypothesis on the strong impact of real bank lending on real house prices was supported. The instantaneous correlation between real bank lending and real house prices was 0.63 and with lending lagged one period 0.51. Bank lending Granger causes real house prices at a very high significance level. Further, the causality seems to be one-sided, namely increases in lending increase prices, but high prices do not predict high lending. The effect of real bank lending is not so strong in the stock market, although the instantaneous correlation was still significant (see figures 5–6). We must emphasize that real bank lending is merely an useful variable in predicting house prices, not a true causal factor itself.

No Granger causality was found between real logarithmic disposable income and real asset prices. Interest rates and house prices were not significantly correlated. Government bonds and stock shares could be substitutes, since the former predicts opposite changes in latter. One interesting relation was found be-

tween housing prices and construction costs. It appears that increasing house prices predict increases in costs rather, than that rising construction costs being the main reason for rises in housing prices.

5.4 Error correction model for house prices

In table 4 an ECM is estimated for the change in housing prices. There was significant autocorrelation in the dependent variable that was handled with a lag in the dependent variable. The sign of the lagged error correction term was negative as expected from ECM-theory. Table 4 also gives the diagnostics summarizing the adequacy of the formulation. Clearly this basic version of the model is not adequate enough. Problems emerge with respect to autocorrelation of the residuals, heteroscedasticity and misspecification of the functional form.

In table 5 a more elaborated version is presented. In addition to the basic variables, several constructed variables were used to describe the ongoing changes in the asset market due to liberalization of financial market. These additional variables were always lagged at least one quarter (see appendix for constructed variables). A liquidity constraint proxy was constructed on the basis of the difference between the market interest rate (RS) and the bank lending rate (RLB) multi-

Table 4. Error correction model for house prices
Quarterly data, 1970/3 – 1990/2
Modelling DLRPHM by OLS

Variable	Coefficient	STD error	t-value	Partial r ²
DLRPHM 1	.370	.110	3.367	.132
DLRUNI 1	.071	.035	2.013	.051
ECM 1	-.074	.035	-2.090	.055
CONSTANT	.001	.004	.327	.001
Q2	-.018	.006	-3.023	.109
DLRLBP 1	.397	.192	2.064	.054
Model performance		Diagnostics for the residuals		
R ²	0.53	DW		2.24
R ² (diff. + seasonal)	0.37	AR (5); F (5,69)		2.10
F (5,74)	16.93	ARCH; F (4,66)		0.10
		Normality X ² (2)		2.59
		Heterosced. F (9,64)		0.55
		Functional form F (19,54)		3.70**
		Reset F (2,73)		4.77*

plied by the ratio of bank loans to disposable income (LBP/YD). The log value of consumption was used as a proxy for permanent income, since it is a result of optimizing behaviour in light of all relevant information on the budget constraint available to consumer/investors. Innovations in the AR(1) model of disposable income differences were used to account for surprises in income. In addition a couple of tax advantage variables (absolute and relative) were constructed in order to capture the benefit from the tax deductibility of housing loan interest payments and tax exemption of imputed capital income from owner-occupied housing. These variables were based on an estimate of the mean marginal tax rate calculated in the BOF4 -model (Tarkka et al., 1989).

The autoregressive conditional error variance of the stock index (UNI-ARCH) was used as a proxy for the varying risk premium. Risk premium is paid to the asset as a compensation for the variance in the return. There is evidence of volatility clustering in the Finnish stock market, which can be used in the house price equation as well. This term got a

significant (although in many variants unstable) *t*-value. Under the CAPM an increase in conditional variance of the stock index could indicate increasing uncertainty that has a positive effect on housing prices and therefore a increasing yield of housing. Since housing and stock prices are clearly strongly positively correlated, we infer that increasing volatility in the stock market reduces the asset prices.

5.5 Liberalization of the financial market

During our sample period two phases of overheating in housing prices occurred. Both of these phases, 1972–73 and 1987–88, were due to changes in house financing. Finland has gone through the same sort of gradual money market liberalization as the other Scandinavian countries. Interest rate constraints on bank lending to the public were abolished in steps from 1983 to August 1986. In spring 1987 free money market interest rates were allowed to be used as reference rates for non-housing loans. In addition, firms have been free to engage in foreign capital transactions.

Table 5. Error correction model for house prices
Quarterly data, 1970/3–1990/2
Modelling DLPHM by OLS

Variable	Coefficient	STD error	t-value	Partial r ²
DLRPHM 1	.253	.101	2.493	.0826
DLRUNI 1	.077	.033	2.276	.0699
ECM 1	-.069	.033	-2.102	.0602
CONSTANT	.0006	.005	.111	.0002
Q2	-.0135	.005	-2.351	.0742
D88/1	.0610	.020	3.023	.1170
DLRLBP 1	.4724	.181	2.601	.0893
LIQCON 2	-.0006	.000	-1.976	.0536
DLRCV 2	-.3918	.152	-2.562	.0869
UNI-ARC1	.0090	.004	2.251	.0685
Model performance		Diagnostics for the residuals		
R ²	0.66	DW		2.19
R ² (diff. + seasonal)	0.54	AR(5); F(5, 69)		1.95
F(9, 69)	14.76	ARCH; F(1, 67)		0.69
		Normality X ² (2)		0.78
		Heterosced. F(16, 52)		0.44
		Functional form F(26, 45)		0.92
		Reset F(2, 67)		3.43*

Diagnostic tests performed are those readily available in PC-GIVE, (see Hendry, 1989).

The resulting inflow of foreign exchange was channeled into housing market through the commercial banks. In addition, interest rates were rather low in 1987 in anticipation of a depression.

There were other reasons for the rise in households' demand for credit, such as increases in income resulting from better terms of trade. The increase in real bank lending was anyhow a symptom of serious disequilibrium in the housing market, caused partly by the rationing of rental housing and tax incentives for owner-occupied housing, such as the deductibility of interest payments on housing loans.

To account for a possible regime shift after the liberalization, which seems to be the main trigger for the rapid rise in asset prices after autumn 1986, we introduced a step dummy after 1986/3, money market liberalization, which however was not significant. This may not be surprising, since adaptation to new kinds of financial markets is not a once-for-all regime shift, rather an ongoing process that will take several years to accomplish. The explosion of asset prices in the first quarter of 1988, was accounted for by a single observation dummy D88/1. For housing prices there was a smaller outlier in 1986/2. The stock index had an outlier in 1985/3 and 1988/2. The recursive OLS estimations showed also signs of structural change in early 1980 s.

The major impact of real lending on real house prices is accomplished already within three quarters, but the variables are not cointegrated. The possible endogeneity of real bank lending as part of the financial allocation problem remains a bit of a mystery, since we did not get any robust results for the significance of interest rates.

6. Conclusions

We conclude that both asset price series have a unit root and are therefore integrated processes $I(1)$. There was also a feeling that nominal house prices might even be $I(2)$, but this was interpreted to reflect merely the exponential house price increases after the money market liberalization. The explosion in asset prices took place from autumn 1987 up until spring 1989.

The unit root property in the asset prices follows from the fact that weakly efficient

markets reflect the stochastic trend in economic fundamentals. However, asset prices are not purely random walk with white noise errors, but rather unit root processes with stationary ARMA errors. The stochastic error part of these asset price processes is predictable from other variables, such as the evolution of other asset series and pricing errors seen in the error correction term. Therefore, this stochastic error contains the short run discrepancy from the market equilibrium.

In addition, some evidence was presented that the asset prices are cointegrated, which means that there exists a stationary linear combination of them, i.e. $ECT = (y_t - \theta x_t) \sim I(0)$. In fact, strong evidence for the cointegration hypothesis was found with the ADER tests and Johansen tests for real asset prices.

Based on cointegration theory, an ECM model was constructed for asset price changes and several additional explanatory variables were tested for their effects on the slow adjustment in asset prices. Nominal asset prices also reflect to some extent the important arbitrage relation between the real risk adjusted, tax-free returns of assets, which motivates the cointegration hypothesis. However, problems in modelling nominal house prices due to the persistence of their nonstationarity remain, even after first differencing. As mentioned, these problems had a great deal to do with the money market liberalization and explosion of asset prices that followed.

Surprisingly, there was also a clear seasonal variation in house prices, especially in the first and second quarters. It seems natural to suppose that housing prices adjust in the short run whereas quantities adjust only in the long run. Under cointegration, the stochastic trend (unit root component) should carry the information about the change in long run market equilibrium. Cointegration does not have to hold, for example between two single stock price series, but it should hold between broad asset markets, if they are efficient enough. Therefore, cointegration also enables one to predict asset prices with the error correction mechanism.

A few additional variables were used in the error correction models of real house prices, among which real bank lending and stock market volatility, lagged permanent income and liquidity constraints seemed to have some value in predicting house prices.

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Appendix

Cointegration data: quarterly data 1970/1 – 1990/2 from the Bank of Finland model (BOF4) data

Variables

PHM	= Housing prices, whole country 85 = 100
PIH	= Construction price index, 80 = 100
LBP	= Bank lending to public, FIM million
R8870	= Effective dividend of the stock index
YD	= Household disposable income, FIM million
MTAX	= Marginal tax rate estimate for households
UNI	= Unitas price index for stocks, 75 = 100
RS	= Market rate of interest, %
RLB	= Bank lending rate to public, %
RB	= Government tax-free bond yield

Constructed variables

LIQCON	= $(RS - RBL) * (LBP / YD)$; describes liquidity constraints
MRT-LBY	= $(LBP / YD) * MTAX * RLB$; relative tax advantage
MR-LBP	= $MTAX * RBL * LBP$; aggregate tax advantage
ECM	= Error correction term
D88/1, D86/2, D85/3, D88/2	= impulse dummies for single outliers
MM-LIB	= Money market liberalization step -dummy, value = 1 after 86/3
CV	= Total private consumption, FIM million (proxy for permanent income)
YD-INNO	= Income innovation term from RLS -estimation of disposable income or AR(1) plus constant and trend
UNI-ARCH	= Predictor of ARCH -variance from AR(5) -stock index difference equation

If not otherwise mentioned prefix L defers to logarithm, E for expectation, D for difference and R for real (deflated with appropriate price index). As in PC-GIVE notation 1 in the end of the variable refers to one period lag.

House prices and inflation: A cointegration analysis for Finland and Sweden

Bharat Barot • Kari Takala

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House Prices and Inflation: A Cointegration Analysis for Finland and Sweden

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Abstract

Given the emphasis on price stability in monetary policy, the concern caused by recent rapid increases in housing prices are understandable. It is suspected that such rises may provide early indication of mounting inflationary pressure. The purpose of this paper is to formulate and estimate an error-correction system model for housing prices and inflation for forecasting purposes. By using the estimated cointegrating vector, we also get an estimate of the equilibrium level for house prices that might be helpful in analysing the current situation in the housing market and the stance for monetary policy.

Housing prices typically exhibit large cycles, and they are thus predictable to some extent. Volatility is caused by the fact that the supply of houses does not react perfectly to changes in housing demand. However, housing prices and inflation tend to have similar growth rates over the long run. In other words, houses provide a good inflation shelter, but in the long run, the real return to is equal to the explicit or implicit rental income derived from the owning of houses. The estimation results also show that the changes in the general price level are transmitted into house prices rather quickly, but inflation is surprisingly insensitive to housing prices. The equilibrium relationship between housing prices and consumer prices is also affected in the short run by variables such as interest rates, wages and the unemployment rate.

Keywords: House prices, inflation, cointegration

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1 Introduction

Growing interest has been focused on the question of how much rising house prices actually signal increasing inflation pressures in Finland and Sweden. In many countries, house prices tend to follow large swings due to changes in housing demand and the fixed supply of housing. This makes house prices more or less predictable (see Englund and Ioannides, 1997). While the business cycle is usually coincident in these countries, the Finnish house price cycle has recently been somewhat ahead the Sweden. The central banks in both countries are keenly interested in the issue of whether large changes in house prices lead to changes in inflation expectations that could accelerate inflation. In this paper, we examine the house and consumer price relationship and address the general questions: What is the essential relationship between house prices and inflation, and to what extent do house prices adapt to inflation in the long run? We also test whether house and consumer prices are cointegrated and whether house prices are useful in predicting overall inflation.

Our enquiry immediately leads us to ask another question. Is the correlation of house prices and consumer prices spurious or do house prices actually cause and predict inflation? Long-run empirical evidence for many countries seems to confirm that housing expenditure shows near unit elasticity with respect to disposable income (Kennedy and Andersen, 1994). Several related equilibrium ratios such as the ratios between housing expenditure and disposable income and housing debt to housing wealth have been used in house price modelling (see Hendry, 1984). Apparently, house prices follow the cost-of-living index, but not the earnings index. In other words, households experiencing rising real earnings tend to increase their quantity and quality of housing. Housing investment costs tend to follow the general price index rather than long-run earnings.

Housing and non-housing consumption are alternative spending items. Consumer utility maximization should lead to a solution where marginal utilities equate and their ratio corresponds to the relative price between them. If cointegration exists, then the only problem is how to estimate the (dis-)equilibrium at individual points in time. To do this, we ask how the adjustment in such prices evolves over time.

Volatility in house prices has resulted in vast changes in household wealth in Scandinavian countries, specifically in Finland and Norway. Through the wealth effect in consumption house prices indicate changes in expectations and affect the business cycle. The effect of capital gains on consumption and saving depends on (a) the extent to which capital gains are persistent, and (b) how fast the economic agents realize the capital gains by selling and borrowing using a house as a collateral. During the last decade a large part of variance in house prices was attributed to the liberalization of the financial market, which led to the credit boom and thereafter increased demand for housing. Under the assumption of inelastic supply of houses, the only consequence of a credit boom would be a price bubble. The house price collapse was especially strong in Finland due to a coincident recession and collapse in exports. The collapse in the Swedish house prices was partly due to the tax reform in 1991, which drastically cut down the after-tax return on housing investments. Generally speaking, however, a

constellation of forces conspired to cause the significant drops in housing prices that occurred in both countries.

The importance of the housing as a service, a form of consumption and an aspect of wealth cannot be ignored. In Sweden, housing wealth accounts for about 60–70% of household total wealth (Barot, 1995), while in Finland housing wealth constitutes around 55% of household aggregate wealth. As housing constitutes the largest part of household wealth in both countries, changes in house prices probably have effects that extend beyond the housing market. The owner-occupation rate is very high (around 70%) in Finland, while in Sweden it is less than half. This latter fact is also reflected in the correlation between house prices and indebtedness as household debt is mostly housing debt. Thus, increases of household debt reflect increases in house prices through increased housing demand as the housing supply is rather inelastic in the short run. In Finland, house prices are also more volatile than in Sweden.

There are various reasons for the close connection between house prices and inflation. Indirectly, rapidly rising house prices signal inflation pressures through excess demand for housing. House prices may also be among the most important indicators for signalling the future development of income expectations because of their role as present discounted value of housing services. This view is sometimes described as the affordability aspect. House prices may also be useful in predicting general inflation as well. Their macroeconomic importance is clear: a house is the largest investment good most people ever purchase or sell, roughly one-fifth of household expenditure relates to housing. Another feature deriving from the present value formula is the role of interest rates in determination of house prices. Indeed, changes in lending rates generally drive house price development.

Some difficulty in comparing the relationship between house and consumer prices arises because prices are treated differently in Finland and Sweden in calculation of the private consumption deflator (PCP) and consumer price index (CPI). If inflation is measured according to national income accounts, the private consumption deflator does not depend directly on house prices, since owner-occupied housing cost are measured through imputed rents. Therefore, housing prices do not enter into imputed rents apart from the effect that owner-occupied houses and rented houses are substitutes, which induces correlation between house prices and actual rents. In few countries B among them Finland and Sweden B the CPI applies a different approach to owner-occupied housing cost, where costs are measured based on capital user costs. In these calculations, market prices of houses affect inflation directly through depletion of the housing capital. Moreover, the interest rates of housing loans enter the calculations. As a result of the use of the CPI in wage negotiations it is quite conceivable that house prices have contributed directly to inflation. Capital costs of owner-occupied housing may also have significant effect on inflation expectations of households even though they do not directly affect the consumption deflator. In any case, for our purposes of modelling house and consumer prices, we want to eliminate any 'endogenous effects' from the CPI, so we use the NSA consumer deflator.

We embark on our country comparisons by utilizing the extended present value formula as a basis for our forecasting system for housing prices and inflation. Since our goal is to specify a forecasting system, we use variables such as interest rate, wage sum, unemployment rate and bank lending as fundamentals underlying housing prices and inflation. We want to structure the explanatory

variables in as similar a manner for both countries, but Finnish and Swedish housing markets are clearly separate markets. Thus, we do not take into account interrelations between the two. Some interactions are quite close; for example, major events in the stock market could trigger spillover effects into the housing market as well. Therefore we study also consumer, house and stock prices as a larger system. Our framework is a multivariate cointegrating vector autoregression approach, where long-term cointegration restrictions can be taken into account.

2 Determining house prices

House price determination usually follows one of two theoretical approaches, ie, houses are either treated as consumption goods or as investment goods. How inflation is measured depends on the approach chosen. In the following, we review the approaches to clarify and to justify why we postulate the cointegration between house and consumer prices and how our empirical estimations may be deficient.

2.1 House price as discounted present value

Before modelling housing prices, we need a theory of house value. A house can be seen as a capital (or investment) good which provides a flow of real housing services over time,¹ but there are also many further distinguishing features such as location and connections to communal services that are not related to the building as such. House purchases are partly based on such service features, where no clear distinction with the environment can be made. Indeed, neighbourhood is one of the most important determinants of housing prices. On the other hand, such location constraints further contribute to the inelasticity of housing supply. Other factors could be lags in planning and zoning for land use by communal administrators. Property taxes are effectively indirect payments for communal services, so they should be taken into account in calculating net property income.

Thus, the present value of a house may be defined as a subjective discounted sum of forthcoming housing services

$$PV = \sum_{t=1}^N \left[\frac{R_t}{(1+r)^t} \right] \rightarrow R/r, \text{ when } N \rightarrow \infty \text{ \& } R_t = R \quad (1)$$

where PV denotes the present value, R_t is the net rent (gross rent - expenses) at time t , r is the subjective time preference (or in market equilibrium the interest rate) and N the life span of house in years. When the housing market is competitive, $PV = P_h = R/r$, where R is the market rent. Here we simply denote

¹ In National Income and Production Accounts (NIPA) and SNA accounting housing is a investment good, not a durable consumer good like cars and boats, which are accounted into consumption by purchases. Therefore housing services are also accounted as a net rent of the invested capital.

house price (P_h) as equal to housing value, since for single dwelling the square price for each housing unit can be thought to be the same. Of course, the size of dwellings varies and the price per unit of floorspace usually declines as a function of size. However, the size of the dwelling does not vary over time, so house price remains proportional to its housing value. It can be seen from the present value formula that nominal interest rates are negatively correlated to house prices through heavier deflation of housing services. Expected increases in interest rates will reduce housing demand. The equilibrium level of housing demand will be affected by the real interest rate, particularly during periods of steady inflation. In a steady inflation regime, it is common that expected and actual inflation match and no effects related to unexpected inflation appear. The value of a house should have a limiting value of ratio between rent and interest rate when time approaches infinity. In other words, the rent should correspond to the alternative return for the housing investment, ie $R = r P_h$, which may look more familiar. In most cases, the market price or the rent is observable, so an unobservable return rate can be estimated based on them

Thus, while the value of a house can be interpreted as present discounted value of housing services, the present value theory reflects a basically long-run equilibrium approach that smooths over short-run changes in house prices. It should be remembered that all assets have two components of return, namely capital income and appreciation (price changes) ie in case of housing (Varian, 1987):

Housing yield = Rental capital income (explicit or implicit rent) + appreciation,

which is analogous to stock market yield composed of dividends and price appreciation. With assets such as houses, land and stocks, price changes tend to dominate the short-run development of the total yield. For pure financial assets (deposits, debt), nominal appreciation is by definition zero and compensation for inflation is incorporated into the capital income (interest payment) component. With stocks, for example, both the price change and dividend are directly measurable. For owner-occupied houses, however, the rent return is neither directly observable nor realized in money terms; total yield has to be calculated. Further, unlike profits, dividends are usually smoothed and compensated, so stock price changes and dividends correlate negatively very strongly. Thus, we do not analyse the total return of housing, but instead concentrate on the relationship between house prices and inflation. In fact, one of the major arguments put forward in this paper is to show that the expectation of house price appreciation is zero in the long run.² This means that the expected capital gains on housing are not only stationary, but also a zero mean variable. Therefore, the only permanent return component left in the long run for owner-occupied houses is imputed rent B keeping a dwelling empty can be profitable only if there is short-run price appreciation.

We adopt an approach analogous to Villani (1982) with slight modifications to make it more suitable for our purposes. We define the house price consisting of

² In fact the same applies for stock prices also, since eg in Finland the mean of real stock prices (deflated by cost-of-living index) does not differ significantly from zero for period 1922-97.

three present value components namely; housing services, home ownership tax advantages and mortgage finance. Basically, we need to form a capital user cost expression for home ownership and calculate real return for home ownership net from financing expenses and taxation. To simplify the presentation, we do not take into account any specific financing arrangements available in Finland and Sweden, rather we assume that the lending rate equals the opportunity cost of capital. Several assumptions need to be made before we are able to present the equilibrium house price as the supply price for housing services.

We postulate a long-run relationship between consumer and house prices. Since our models are based on cointegration analysis of house and consumer prices, we have to provide explanation for this equilibrium mechanism between these variables. In the user cost calculations consumer goods and housing are alternative expenditure items and by cointegration their relative price is assumed to be stable. Consumption capital asset pricing model (CCAPM) implies marginal utility of consumption to be equal to real return for housing (eg Breeden 1979, Kazemi, 1988). Briefly, cointegration follows from the fact that marginal utilities from different kinds of consumption tend to equate, which leads to a stationary marginal rate of substitution. As marginal utilities equal real after-tax return in the long run, indirectly appreciation of houses equals inflation, and no excess real return emerges. Equal expected returns will also lead to cointegrated price indices.

2.2 The user cost for housing

In order to comprehend the relationship between house and consumer prices, we have to consider the choice between consumption on housing and on other goods, and their price formation. The cointegration between these two prices is not self-evident, since in general there is no obvious economic forces that will converge prices of two separate goods.³ Markets are usually defined for single homogenous goods. As an analogue to the stock market, there is no guarantee that stock prices of two separate firms will develop in the same way. Firms can also be bankrupted and their shares may turn out to be worthless. As stock market indices usually account the prices of limited number of firms, there might be bias due to the bankrupted firms. However, when we aggregate and diversify our holdings between firms, their development will soon converge to the behaviour of the overall stock market index. When we are comparing two wide markets like housing market and stock market in general, there may be some strong long-run relationship between them.

As a result of positive demand shocks in the housing market, eg due to demographic factors, tax changes, the house prices will increase since the supply of houses is fixed in the short run. Housing production will take time and significant lags appear also because of land restrictions, planning and financing time lags, shortages in building material etc. The consequence is that housing market adjusts sluggishly to changes in demand. The only policy recommendation

³ However generally, price indices for goods grouped by use do not converge in the long run; preferences may change, tax treatment may be indexed, etc.

to smooth house prices changes seems to be affecting the variability of the demand.

The calculation of the cost of housing services is not totally straightforward, since the costs and benefits of home ownership do not always appear in money form. Since house is a durable asset, we have to use present value calculations to approximate what consumers truly expect. A house will produce a long lasting stream of housing services with adjustment costs, therefore housing consumption is a dynamic problem. First we have to list and measure all the income and expenses due to owner-occupied housing. For owner-occupants costs are largely unobservable and the imputed income have to be computed. Especially we have to decide how to treat possible capital gains and how to import the whole tax code into calculations.

Broadly speaking the alternative use for housing services supplied by a dwelling is simply consumption to other goods. In practice, renting a dwelling is a direct alternative to buying a dwelling. In technical terms owner-occupied housing services are supplied by a durable housing stock, while non-durable consumption is a flow (Van Order and Villani, 1982). These two different dimensions of the problem makes it suitable to analyse it in an optimum control analysis framework. Buying a large durable item such as a house forces the consumer to finance part of the purchase through lending. It is possible to separate two different interest rates applied to different types of capital. Formulating an optimal control problem does not take into account indivisibilities and treats housing capital as a continuous variable. It is possible to increase housing capital by making housing investment, but housing capital is also subject to technical consumption that depreciates the real housing capital at an almost constant rate. Households can also make financial savings through excess income over consumption and save in the form of amortizations. In this model, financial savings deposited to banks produces interest return, which is lower than the borrowing rate. In analogue to depreciation faced by housing capital, financial saving is subject to inflation, which depreciates the purchasing power. The real return for savings is approximated therefore as nominal interest subtracted by inflation. In addition nominal interest is subject to capital income taxation. In this sense government taxes also purely nominal return caused by inflation.

It can be derived that the user cost of housing capital equals the marginal rates of substitution (MRS) between housing capital and consumption

$$\frac{u_h}{u_c} = [(1 - t_c)r - \pi + \delta]p_h / p, \quad (2)$$

where u_h and u_c are marginal utilities with respect to housing and consumption, p_h and p house and consumer prices respectively, t_c is currently the capital rate (due to tax rules in Finland and Sweden), r is the interest rate (lending and borrowing rates are equal), π is expected house price inflation (appreciation) and δ is real depreciation, which can be assumed to include property taxes and transaction cost (see Brown et al 1997, p. 534).

The marginal rate of substitution reflects the amount of money that has to be given to the consumer to compensate for one lost unit of housing (eg one unit of floorspace). This equation is based on comparison of alternative costs. In the capital market equilibrium, the real rental price of housing must equal the real user cost and clear the market for housing services and housing capital markets. The

expression presents a measure of housing cost with respect to prices of other consumer goods. Housing cost is an increasing function of the after-tax nominal interest rate, while inflation decreases the opportunity cost of housing. If housing stock is fixed in the short run, an increase in the rate of inflation will increase the real price of houses as long as the nominal interest rate rises proportionally and keeps the real interest rate constant. In a similar way, an increase in the capital tax rate for other assets (capital gains on houses are tax exempt) will raise real house prices as the value of the tax advantage increases.⁴ Thus, an important implication concerning the regime for permanently low inflation is that it will lead to lower relative house prices (Holly and Jones, 1997 p. 554).

Naturally, depreciation is also an expense to the homeowner. As said, this formula indicates that when we aggregate different goods across the market, marginal utilities tend to equate as consumers are maximize their utilities (see eg Varian, 1984). Over time the price vector between house and consumer prices forms a stable temporal equilibrium. Assuming asset market equilibrium, the real rental price of housing (R_t/p) equals the real user cost. The rental price R_t clears the housing services market, which gives us

$$\frac{R_t}{p} = \frac{u_h}{u_c} = [(1 - t_c)r - \pi + \delta] \frac{p_h}{p}, \quad (3)$$

For owner-occupied houses true rental price R_t is unobservable, but can be approximated as a function of some income measure (permanent, disposable income or wages), added with possible demographic variables and existing stock of houses. Therefore, the rental price is determined by variables that affect the demand and supply of housing (Breedon and Joyce 1993, Brown et al 1997). The nominal house prices can be written as function of consumer prices, income, demographic variables (DEMO), housing stock (HW) and the user cost expressed in the specification below

$$\ln P_h = f \{ \ln(P), \ln(Y), \ln(\text{DEMO}), \ln(\text{HW}), \ln[(1-t_c)r - \pi + \delta] \}. \quad (4)$$

This specification can directly be estimated either by linear regression or, eg a time-varying coefficient model. Cointegration between house and consumer prices can be modelled using error-correction models (ECM). In empirical models, the user cost variable can be conveniently decomposed into nominal user cost and a capital gains variable as these components are likely to behave very differently.

⁴ In principle capital gains are taxed in Finland and Sweden. In Finland, owner-occupied dwellings are tax exempted after two years holding. Capital gains on non-owner-occupied houses are taxed effectively with lower rate than other capital income, since houses are held on average almost 10 years, which postpones the tax payment (Varian, 1987 p. 204). Due to special tax rules owner-occupied houses are effectively tax exempted in capital income and capital gains taxation.

2.3 Treatment of housing costs in consumer price index

The treatment of housing costs for owner-occupied housing is based on capital user cost formulas in few countries like Finland, Sweden and the UK. The problem arises with the measurement of the unobserved imputed rent for homeowners. For rented flats and houses, gross rents are directly observable and no ambiguity could arise. In national accounts, measurement of home ownership costs is based on corresponding rental expenses, which are likewise also not observed, but could be more reliably evaluated than capital costs.

Applying the user cost formula above, we can rewrite the nominal housing costs for owner-occupiers as

$$\begin{aligned}
 C &= M + r_A(1-t)((1-\delta)H-L) + r_L(1-t)L - \Delta P_h \\
 &= M + r_A(1-t)((1-\delta)H-L) + r_L L - tr_L L - \Delta P_h \\
 &= M + (1-t)r_A(1-\delta)H + (1-t)(r_L-r_A)L - \Delta P_h,
 \end{aligned} \tag{5}$$

where C = housing costs, M = maintenance costs, r_A = alternative return, r_L = lending rate for housing loans, L = housing loans, t = marginal tax rate, H = gross housing wealth, δ = depletion of housing capital and ΔP_h = appreciation of the housing capital.

Housing costs are a sum of maintenance costs, alternative costs accruing from ownership of the housing capital, borrowed capital interest payments and possible appreciation/depreciation of the housing capital. Housing capital is also subject to constant depletion that erodes the housing wealth. All the components except return of the ownership, ie imputed housing income, are observable. In the SNA statistics, only actual money transactions are measured.

Buying large durables such as houses usually involves using external financing and thereby making it necessary to divide capital into own and borrowed capital. In most cases, the lending rate exceeds the return from the alternative investment. If perfect capital markets are assumed (r_L-r_A) disappears as lending rate is assumed to be equal to return of the alternative use. In this case, the equation reduces to

$$\begin{aligned}
 C &= M + (1-t)r_A(1-\delta)H - \Delta P_h \\
 &= M + (1-t)\delta H + (1-t)r_A H - \Delta P_h,
 \end{aligned} \tag{6}$$

This assumption would eliminate the need to include separate interest payment costs into the formula. Maintenance cost can be defined as the cost that keeps the capital value of the housing wealth intact. In this sense the items M and $(1-t)\delta H$ are similar. Assumption of perfect capital markets is however unrealistic. Calculations performed in the National Statistical Offices are based of other kinds of simplifying assumptions. We can first write the previous equation into a form

$$\begin{aligned}
 C &= M + (1-t)r_A(1-\delta)H + (1-t)(r_L-r_A)L - \Delta P_h \\
 &= M + (1-t)\delta H + (1-t)r_A H + (1-t)(r_L-r_A)L - \Delta P_h.
 \end{aligned} \tag{7}$$

In Finland, the third component, which measures the imputed return of the housing capital, and the appreciation term measuring the changes in the value of housing capital are assumed to eliminate each other, ie $(1-t)r_A H - \Delta P_h = 0$ (Lehtinen and Koskenkylä, 1988).

This assumption simplifies the measurement of owner-occupied housing costs significantly since it avoids estimating the unobserved imputed housing income. It also allows us to get rid off the very volatile appreciation component in housing capital. The house price development is not washed out entirely, however, since market house prices are used in valuing the depletion of the housing capital. These simplifications are by no means harmless, at least in the short run. If we think that the total return of the gross housing is compounded of two components, namely return and appreciation, it is more realistic to assume that in the long run there is no real appreciation, which means that appreciation of houses equates to inflation. In this sense, housing capital affords a perfect shelter for inflation, but not much else.

In real terms, only the imputed return is left as a return for housing wealth. This accords better to what we have in mind when we are buying houses and expect return in form of housing services. This imputed rent makes the basis of the housing wealth as an asset, not the expected rise in value due to price changes that may be negative also. Anyway, the formula used in Finnish CPI calculations for owner-occupied housing costs reduces into

$$C = M + \delta H + r_L L, \quad (8)$$

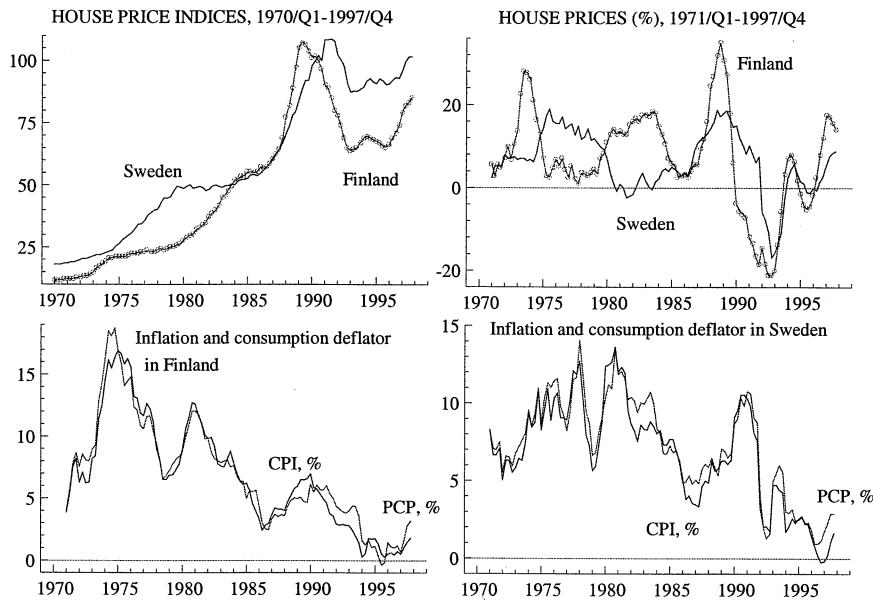
which leaves us only maintenance cost, depletion and interest payments of housing loans as housing costs. In Sweden cost of owner-occupied houses include similarly interest payments, depreciation, insurance, maintenance and preparation costs plus real estate taxes. Up to 1984 value of net housing capital was based on purchasing prices, which were updated by maintenance costs reduced with borrowed capital. The return of own capital has been based on average deposit rate and for borrowed capital the average bank rate accordingly. From 1984 onwards in Sweden depreciation has been calculated from market values of houses with 1.4 percent depreciation rate.

The problem with our forthcoming estimations about the relationship between house prices and inflation makes CPI a slightly endogenous to house prices. Therefore, we prefer to use private consumption deflator (PCP) as a measure of inflation in cointegrating relation. Figure 1 shows the difference in CPI and PCP caused by treatment of houses as either consumption or investment goods. The PCP calculation treats owner-occupied housing as consumption goods and calculates imputed rents for owner-occupied houses. In these calculations, house prices do not directly affect the housing costs.

For the Swedish data we also use a consumption deflator that does not include CPI-type calculations for home ownership. In Sweden, this problem is not as serious as in Finland, since the share of owner-occupied housing is much smaller than in Finland. In Sweden, however, the vast subsidies devoted to housing may affect the efficiency of housing markets. In Finland, subsidies to housing are also considerable, but they are mainly directed at rental housing production. The volatility in housing prices is something that arises from a rigid supply of houses

that does not react to rapid changes in housing demand. If we want to reduce house price volatility, we have to smooth directly housing demand.

Figure 1. **House prices and inflation in Finland and Sweden**



3 Basic data description

In both countries, house prices were affected by the lending boom due to financial liberalization of the late 1980s (Figure 1). The figure also presents house price effects on difference between CPI and PCP. In Finland, the collapse in house prices was most severe during the recession and matches the actual decline in economic activity and household disposable income. During 1990–93, GDP fell altogether 14% in Finland. In Sweden, house prices fell by 20% in a short period (about 18 months), while in Finland nominal house price decline was almost 40% and extended to 1996. In Finland, the volatility in house prices measured from peaks and troughs has also been somewhat larger, which may relate to the higher share of owner-occupied housing and higher variance in economic activity. The possibility to move to rental housing tends to smooth the variation in house prices, since they are substitutes to owner-occupied housing. In both countries, the latest upward swing started in 1996. It is very likely that current swing may persist for at least 3–4 years as in both countries house prices were significantly below their long-run equilibrium levels. Households are still rather cautious in using debt finance in housing purchases and investment.

Shortly after financial deregulation, the Swedish economy began to slide into recession. First escalating interest rates due to a rising budget deficit, then rising unemployment signalling greater uncertainty about the future brought a radical

decline in housing demand. After devaluation in Finland 1991, Sweden followed a similar route in 1992 with a decline in domestic consumption and investment demand. By 1993, house prices had dropped by 30% across the country. Englund et al (1995) estimate that about one-half of a per cent of this fall in house prices was due to the effects of the tax reform 1991, while 8% was caused by the fall in real GDP. The 1991 tax reform implied that the previous prevalent subsidies were removed both for the households and the suppliers i.e. the builders in the building industry. In addition the debt deflation process were prevailing in the Swedish economy up to late 1993. The boom in consumption was originally financed by equity withdrawals and households borrowed against rising housing wealth and used the proceeds to finance consumption, therefore a collapse even in nominal prices was possible. The basic story behind house prices is rather similar both in Finland and Sweden.

The relationship between house prices and consumer prices may look at first hand rather weak in Finland and Sweden, but instantaneous correlations are still high in both countries. In Finland, correlation between annual change in nominal house prices and inflation reaches 0.58; in Sweden it goes as high as 0.35 (Figure 2). There are many reasons for house price collapse in early 1990s both in Finland and Sweden including smaller tax deductions, higher real interest rates and tighter bank lending policy. This led to a slowdown in bank lending in both countries, as it was correspondingly one of the reasons for house price overheating as well. Monetary expansion is one of the main background factors for asset price inflation. At the same time companies found it more reasonable to raise funding from the capital markets, households increased their bank lending. While the bank lending constraints were eased it allowed households to use the tax benefits related to borrowing. Financial liberalization together with tax distortions triggered the boom in house loans.

By treating houses as assets we can also compare house and stock prices for evidence of cointegration as well (Takala and Pere, 1991). In Finland, the relationship between house and stock prices has been closer than in Sweden (Figure 3). Stock market prices in Helsinki and Stockholm have diverted from domestic asset markets mainly because of increase in telecommunication share prices and their globalized ownership. Therefore, cointegration, for example, between stock market prices and house prices based on capital market efficiency is rather difficult to demonstrate at the moment.

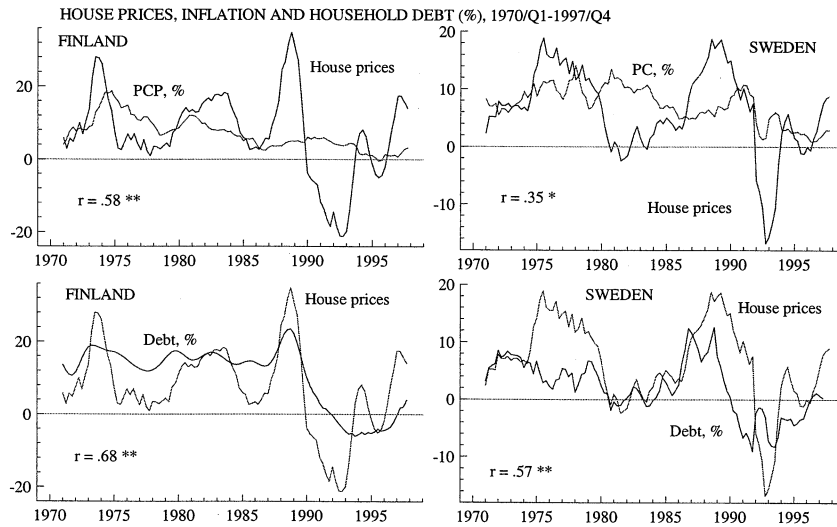
The affordability aspect to housing ownership can be clarified by looking at the house wealth valuation with respect to disposable income (Figure 4). The affordability aspect has frequently been related also to house price ratio to disposable income. In the long-run housing wealth should be proportional or at least stationary to disposable income as housing services consumption is expected to be near unit elastic with respect to total household expenditure. Our time period is however unique in this respect. In Sweden the indebtedness ratio has not yet normalized very much after the deregulation of the credit market in 1980s, while in Finland the indebtedness ratio has declined remarkably. The increased indebtedness in late 1980s has affected also the ratio between market house prices and disposable income. It should be noted that there is statistically significant declining trend in house prices-disposable income ratio in both countries, which is

related to increasing standard of living as earnings index has increased faster than cost of living index.⁵

One reason behind the decrease in the popularity of debt financing after 1990s recession has been the increase in after tax real interest rates. In both countries real after tax rates have become strongly positive during the 1990s (Figure 5). In Finland the deductions of housing interest payments were restricted severely by the introduction of new capital income tax in 1993 that limited the interest deduction rate first to 25%, while the benefit rate had been on average close to 50 earlier when income tax rates were applied. In Sweden similar changes in the tax code had taken place already from the beginning of 1988 and 1989. The tax wedge between real market interest rate and effective after-tax of housing borrowing contributed therefore significantly on the asset price boom and crash. By theory, the size of the tax wedge becomes larger with higher inflation and with higher income tax rate.

In the short run, house prices are largely demand determined. House price development in both countries was also affected significantly during early 1990s by a slowdown in real disposable income. With increased income expectations in the last couple years, house prices have started to rise again.

Figure 2. **House prices and private consumption deflator, %**



⁵ In the UK however, real house prices and real disposable income have followed similar trends. UK changes in income and house prices have also correlated strongly (Brown et al, 1997).

Figure 3. **House, consumer and stock price indices**

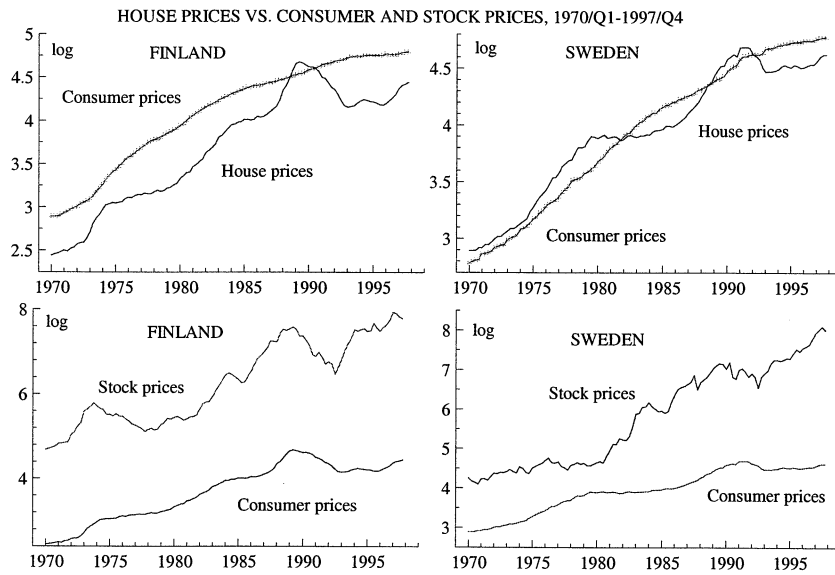


Figure 4. **Housing wealth, house price and indebtedness ratios**

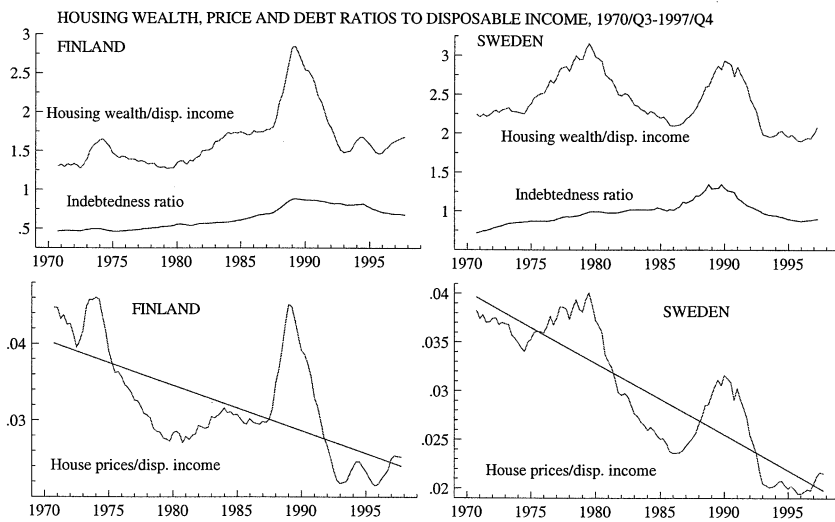
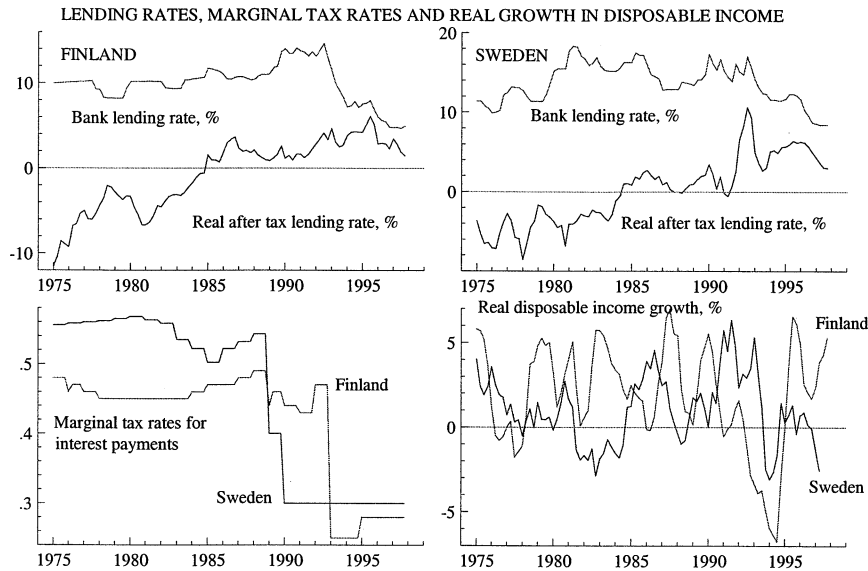


Figure 5. Lending rates, tax rates and income growth



4 Empirical evidence on house price relationship

4.1 General evidence

Table 1 provides and compares the basic statistics for Finnish and Swedish house prices and inflation from the beginning of 1970 up to the second quarter of 1997. The means and standard deviations are surprisingly close to each other for each country, even though our estimation period is not particularly long. Even though means of house price changes are rather close to inflation in both countries, the individual price changes are even closer for each country over the period. For comparison table 1 reports also price changes in housing investment and stock prices.

From the theoretical point of view, it is not surprising that old house prices and housing investment prices converge to each other as housing stock is naturally renewed by housing investment. Although new houses are better outfitted, and there is a quality difference between them, old and new houses are close substitutes, and this keeps their price development tied together (Figure 6). Indeed, old house prices and housing investment prices are also cointegrated. Also, while stock market price changes have been much more volatile than house prices during the sample period, the mean of stock prices is rather close to that of house prices.

Figure 6. **House prices and housing investment prices and shock profiles**

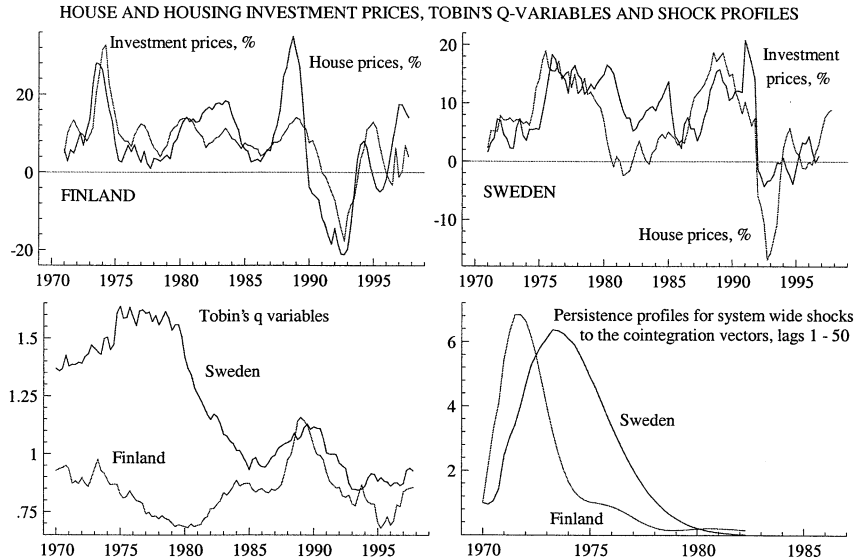


Table 1 (in Appendix 1) reveals that for Sweden the normality of house prices is rejected. This is due to outliers in the data during the period 1990B91. In Finland, inflation includes surprises during early 1970s oil crises. More significant from our viewpoint, however, is the observation that excess return from houses seem not to be different from zero in the long run.

Rising house prices will make the house purchase more difficult for average families. This affordability aspect will limit demand for owner-occupied housing and eventually slows the increase in house prices. The restriction caused by affordability limit becomes even more severe for very small dwellings due to indivisibility and for very large houses because of total value. The consequence of this affordability limitation is stickiness in the house prices. House prices do seem to have a unit root, but there is predictable inertia in these prices that cannot be overlooked. Intuitively this inertia could arise from the production lags in the housing supply.

A variety of unit root tests showed unanimously that the null hypothesis of a unit root cannot be rejected for the indices of house and consumer prices. The same tests applied to log differences of these indices show that house and consumer prices are most likely integrated of order one. Possible cointegration between house and consumer prices prerequisites that these series are integrated.

Additional clarifying information about the nature of the data generating processes can be get through analysing these series by structural time series models (see Harvey 1989). The decomposition of the house prices into trend, seasonal and irregular components reveals that there is no significant drift in house prices and a pure random walk in levels is a more proper description of the process for both countries. There are weak signs of seasonal variation in house prices, but seasonal terms are not statistically significant at the 5% confidence level. It is as hard to try to identify any significant irregular components from the

univariate house price processes, which points out that the shocks to house prices are more likely permanent. These results seem to convince us that housing markets are rather efficient and that house price adjustments do not indicate large deviations from market equilibrium as such. House prices are determined by other fundamental factors like income, interest rates, but house prices work efficiently in equating the demand and supply (see also Kosonen, 1997).

In general, the same observations apply to quarterly consumer prices. They are also close to pure random walk processes that do not include any drift, irregular or seasonal components. It must be emphasized, however, that monthly consumer price indices usually exhibit a clear seasonal pattern, due eg to tax changes at the beginning of the year and seasonality of some food commodities.

The system estimations were started by looking at bivariate house and consumer price systems. The first decision in cointegration testing is the selection of the VAR order used in system rank identification. Frequently model selection criteria would suggest different choices of order. For Swedish house and consumer price system, Akaike's information criterion (AIC), Hannan-Quinn criterion (HQC) and Schwarz's Bayesian criterion (SBC) proposed unanimously that five lags should be optimal. For Finnish price systems VAR orders 2 and 5 were both likely candidates. The diagnostics of the residuals, however, suggested that five lags should be selected to confirm the white noise properties. As the adjustment of house prices with respect to consumer prices seems to be slow due to slow reaction of housing supply selection of a high VAR order seems also more appropriate.⁶ There is slight risk choosing the order too high for over-parametrization, but we also know that the adjustment of house prices is rather sluggish.⁷

Another important task before getting into actual cointegration testing is to find out what kind of deterministic components are included in the house and consumer price system. Testing the existence of deterministic components such as intercepts and trends in the data or cointegration relations would have important implications for the asymptotic distributions of the rank statistics. Certainly, the choice of the cointegration rank is one of the most important single decisions to be made in system modelling, since all the inference on hypothesis is conditional on the chosen rank. In economic applications, a trend is rarely needed in the cointegration space and more likely test results of that kind would rather indicate from other important missing variables. A constant could be included into system to allow a drift in nonstationary series and a overall mean in stationary variables. Including trend would be harder to explain, since we do not expect to find any linear or quadratic trends among price indices or in particular within the cointegrating vectors. Our series also had the same base year, although the means for Finnish house and consumer prices are quite different. For Swedish series, the means were almost identical.

⁶ It is obvious in practice that the lag length chosen affects the significance of the cointegration rank tests. When the VAR order is selected too long, the number of identified cointegration relations tends to get smaller. If the rank chosen is too small, the tested hypothesis is rejected too often (Hansen and Juselius, 1995 p. 8).

⁷ In most applications with quarterly data, a second order VAR is sufficient.

It is thus quite unrealistic to assume that house or consumer prices include any deterministic trends, even though consumer prices look rather linear in logs. However, it remains an empirical matter to include a constant term in the cointegration relations as the differenced series have almost equal means. The selection of the rank and deterministic component should be made jointly (see Hansen and Juselius, 1995 p. 68). For Finnish and Swedish bivariate price systems, in particular, the rank of one and model without deterministic component were chosen, although for Finnish data separate system deletion tests for constant and trend this seemed to cause doubts.

Finally the Finnish model was estimated without any deterministic components. Whereas inclusion of a constant did not change the results of the Swedish model much, including a constant into the unrestricted model affected the Finnish results considerably. This is probably due to the near linear trend in the log of consumer prices. Without a constant, cointegration is found; with unrestricted constant, cointegration tests fail. For the Swedish price system, cointegration is found easily with or without an unrestricted constant.

4.2 House and consumer prices system tests

In order to get a closer look to the relationship between house and consumer prices, we start testing by scrutinizing the simple bivariate system. We test cointegration by using the efficient Johansen maximum likelihood approach, which enables us also to identify several cointegration relations simultaneously. Table 2 presents the ordinary cointegration tests and basic tests for stationarity, exclusion and weak exogeneity for both countries. The long-run exclusion test is based on restriction $H_0: \beta_i = 0$, where i indicates the endogenous variable. The test for weak exogeneity is similarly done by imposing $H_0: \alpha_i = 0$ on the adjustment parameter. The common maximum eigenvalue and trace cointegration tests, which are used for testing the number of cointegration relations, are the following

$$\begin{aligned} \lambda_{max}(r, r+1) &= -T \ln(1 - \lambda_{r+1}) \\ \lambda_{trace}(r) &= -T \ln \sum_{i=r+1}^n (1 - \lambda_i), \end{aligned} \tag{9}$$

where λ 's are the estimated eigenvalues of the characteristic equations, T is the number of usable observations and r is the rank of the long-run matrix (π). The number of the cointegration vectors (r) is always smaller or equal to the number of the endogenous variables (n).

According to cointegration tests, there is likely one stationary linear combination among prices in both countries and the homogeneity tests more or less confirms that house price and consumer prices do not drift apart in the long run. For Swedish data, the homogeneity restriction ($\beta_1 + \beta_2 = 0$) is rejected at 4.1% significance level, while in the Finnish data restriction remains in power. In a bivariate system, two cointegrating relations would mean that original variables are not trended, ie stationary, which surely is not the case. Price indices are clearly non-stationary and neither of the variables could be excluded from the

cointegration relation. For both countries, weak exogeneity is strongly rejected for house prices, while consumer prices are not rejected for weak exogeneity. However, it must be emphasized that results concerning weak exogeneity would easily change depending on the particular system tested or even sensitive to chosen lags in the VAR system.⁸

The theory is not in contradiction with the idea that house prices are endogenous and consumer prices could be exogeneous for both countries. After all, it is more likely that house prices will adapt to consumer prices in the long run than the opposite. Consumer prices reflect the overall expenditure prices, but house prices only the share of housing expenditures. In many countries it has been observed that housing expenditure tends to have unit income elasticity in the long run. In the short run, deviations are bound to exist. In Finland, for example, the share of housing expenditure from total expenditure rose due to recession. In Sweden the share of housing expenditure has been relatively stable around 21–24 percent.

In general, it seems possible to present a system model for house and consumer prices based on cointegration between these series. Cointegration will limit the long-run development of the price series, especially house prices, but we can add an error-correction term into the model as well. This error-correction term should also Granger cause either one of the cointegrating variables. In the short run, there can be a bundle of variables affecting house and consumer prices separately.

Granger causality tests were applied for stationary quarterly differences to find out in which direction the predictive causation runs (Table 3). For Finnish data, Granger causality tests between house and consumer prices revealed little of interest, but in the Swedish data, we found that consumer prices seem to affect house prices strongly in the short run. In both countries, house prices effectively predict changes in the unemployment rate as house prices predict income expectations and income uncertainty. There exists a more direct link between house prices and wages as house prices predict changes in wages. House purchases are based on forthcoming wages as debt financing forms an important part of the transaction. Granger causality tests were also performed for trivariate system blocks. In both countries, house prices were Granger-caused by consumer and stock prices, but there were feedback effects as well in Sweden. Based on these empirical findings, we could formulate a bivariate system model for house and consumer prices. Instead, we will continue to see whether a more general enlarged system might be achieved by adding stock prices into the system.

⁸ These results were obtained by using five lags in the VAR; with only two lags, the Finnish results were just the opposite.

4.3 Estimation of the error-correction model

In order to evaluate more broadly the equilibrium house prices, we turn to analysing larger price systems. With three or more variables, there may be more than one cointegration vector, and the presence of multiple cointegrating vectors can be tested only in the Johansen framework.⁹ Another motive for doing this was to be sure that house and consumer price relationship is not a part of a larger asset price cointegration system. This framework also allows us to analyse a wider scale of hypothesis and the speed of adjustment of the parameters.

In Tables 4 and 5 we present cointegration results from three variable endogenous system of house and stock prices and consumer prices. The null hypothesis is again non-cointegration, which is rejected also in this larger system. In this system we used bank lending rate for new loans as stationary I(0) exogenous variable for the price system.¹⁰ Adding stationary variables into system would otherwise increase the cointegration rank, although robustness of the system can be tested by adding stationary variables into system. The weak exogenous tests performed showed that this intuitively appealing assumption could be made for both countries. The exclusion tests showed, however, that the lending rate cannot be eliminated from the determination of the long-run equilibrium (which does not entirely accord to our beliefs). Economic theory suggests it would be more likely that the stationary interest rate does not belong to the core of the system of trended price indices.

In the trivariate price system the VAR length selected as five for both countries, though for Sweden even this did not seem to be quite sufficient as can be observed from the system diagnostics. The rank of the matrix is equal to the number of its characteristic roots that differ from zero. Quite clearly according to eigenvalues and cointegration tests only one cointegration vector is found, therefore only one linear 'equilibrium' combination between variables is stationary.

Then the rank of the long-run matrix (π) is one and hypothesis of the matrix of cointegrating parameters (β) and their adjustment parameters (α) can be estimated upon this assumption $\pi = \alpha\beta'$, where α and β are both ($n \times r$) matrices. The rank of the long-run matrix also equals the number of cointegrating vectors. If this relationship is homogeneous, the interpretation would be that income can be either consumed or invested into housing assets or into the stock market.

As we should expect according to present value formula, the lending rate could be seen as one determinant of price in both countries. For both countries a four variable system exclusion tests implied that no variable could be eliminated from determining the β -vector. Nearly the same result was duplicated when lending rate was given a status of weakly exogenous variable included into the cointegration space. For Sweden, though, stock prices were now supposed to be excluded. The deletion of the stationary lending rate from the system was strongly

⁹ For instance, adding house investment prices into system would break the weak cointegration between house and consumer prices, since the relationship between house and housing investment prices would dominate the cointegration relation.

¹⁰ Here we have used the Microfit 4.0 package, which allows stationary exogenous variables within the cointegration VAR system (Pesaran and Pesaran, 1997).

rejected for Finnish data (with p-value of .001 in the LR-test) and for the Swedish data as well (with p-value of .002).

Adding stock prices into the system did not alter the finding about only one cointegration relationship, although as alternative asset prices house and stock prices could be cointegrated separately as well. According to estimation results, the relationship between house prices and consumer prices clearly dominates the cointegrating vector found. Since we are particularly interested in the long-run equilibrium of house prices, we must look at the imposed restrictions on the cointegrating β -vector, which have the form: $H_0: \beta = H\phi$.

After testing the zero homogeneity of the β -vector, we found that it somewhat surprisingly rejected Finnish data at a probability level below 1%. Imposing restriction $\beta_1 + \beta_2 = 0$ was also strongly rejecting in this system. For the Swedish data, zero homogeneity was not rejected below 5.6% probability. Results are seldom as clear as one would hope, especially in system estimations, so we simply have to make choices about which route is the most promising to follow in empirical work. The rejection of the homogeneity is however more related to inclusion of the lending rate, rather than inclusion of the stock prices.

Having analysed these trivariate price system tests it appeared that the stock prices could be more properly used as weakly exogenous variables. Stock prices did seem to have a rather minimal effect on the long-run solutions as well. Partly this may reflect the fact that housing and consumer good markets are much broader markets and more important for the whole economy. For example in Finland, the number of individual investors dropped to a half during the recession in early 1990's. In addition in both countries the liberalization of foreign capital movements in late 1980's may have also weakened the inter-relationship between housing and stock markets. Subsequently the internationalization of the stock market during last five years due to telecommunication corporates have likely affected the results.

It is also possible to test, whether adjustment coefficients of system variables adjust significantly to their long-run relationships. The separate zero-restrictions were imposed to α -adjustment coefficients. For Finland zero restriction was strongly rejected for consumer prices, but not for house or stock prices. For Swedish data, zero-restrictions on α -adjustment coefficients were rejected for house and consumer prices, but not for stock prices.

The estimated shock profiles for cointegration vectors tell us that for Finnish data a common shock goes much faster through than for Swedish data (Figure 6). It can be seen, however, that the size of the reaction is initially somewhat stronger in Finland. There may be no obvious reason for this, but the larger share of owner-occupied housing in Finland might be partly responsible.

Before we tackle the problem of estimating an error-correction model for house and consumer prices, it is interesting to look at the estimates of equilibrium house prices. In Figures 7 and 8, we first present equilibrium house price estimates from the bivariate price system with lending rates as exogenous variables. The homogeneity restrictions were imposed for this estimation. In Sweden, the result from the estimates in last quarter of 1997 indicates that house prices were under their long-run equilibrium level. In Finland, it seems that the recent house price increases have already raised prices above the long-run equilibrium level.

In the final error-correction system, we included for simplicity only house and consumer prices as endogenous variables. After-tax lending rate, stock prices,

unemployment rate and wages were used as exogenous variables. Tables 6 and 7 present finally the estimated EC-models for house and consumer prices. Figures 9 and 11 show the actual series and their fitted values plus standardized residuals.¹¹ The residual analysis tells us that for Finnish series there is a period in mid 1980's where the model fit is not quite good in house prices. In Sweden, the bivariate model does have some difficulties to follow the collapse in house price in the beginning of 1990s, but otherwise model fits seem appropriate. The model diagnostics for both countries tell us however, that there may be some important dynamic variables missing from our specification, since there is serial autocorrelation in the residuals. Partly this autocorrelation arises from our specification of using annual differences of the endogenous variables to have a more proper scale for coefficients. The normality tests also alarms for Sweden, which indicates that there are few outliers in the model.¹²

The error-correction terms are highly significant and have the right sign for both countries in the house price equations. The overall size of the coefficient is however small suggesting that adjustment to equilibrium will take a long time once the system has been shocked. In Figures 10 and 12, the impulse response functions are plotted to show how asymmetric the shocks coming from house prices and inflation are. It can be seen that a unit shock in inflation has a very strong and long lasting effect on house prices in both countries, while a unitary shock in house price has smaller effect on inflation. Thus, the preponderance of evidence seems to suggest that house prices are more likely to adjust to inflation than the other way around.

Apart from the endogenous variables, exogenous variables have mostly expected signs in the system equations for house prices and inflation. Increasing after-tax lending rate reduces house prices in both countries. Change in unemployment rate had a 'wrong' positive sign in the Finnish house prices equation, since increasing labour income uncertainty should have a negative effect on housing purchases and therefore to house prices. In the Swedish house price equation, unemployment rate change had a negative sign, but the coefficient was not statistically significant. The timing of increasing and decreasing unemployment rate as well the changes in the rates are also pretty similar in both countries. This arises of course from the close correlation with business cycles and relatively large foreign trade. In both countries wages were used in the inflation equation as a demand pressure variable and it turned out to be valuable. In the Finnish inflation equation bank lending was used in the house price equation as additional explanatory variable, which turned out significant as expected.

¹¹ The ex post forecast tests performed for last three years showed accurate predicting ability for both countries.

¹² While it is possible to take these outliers into account, we have not done so here.

Figure 7. **Equilibrium house prices in Finland**

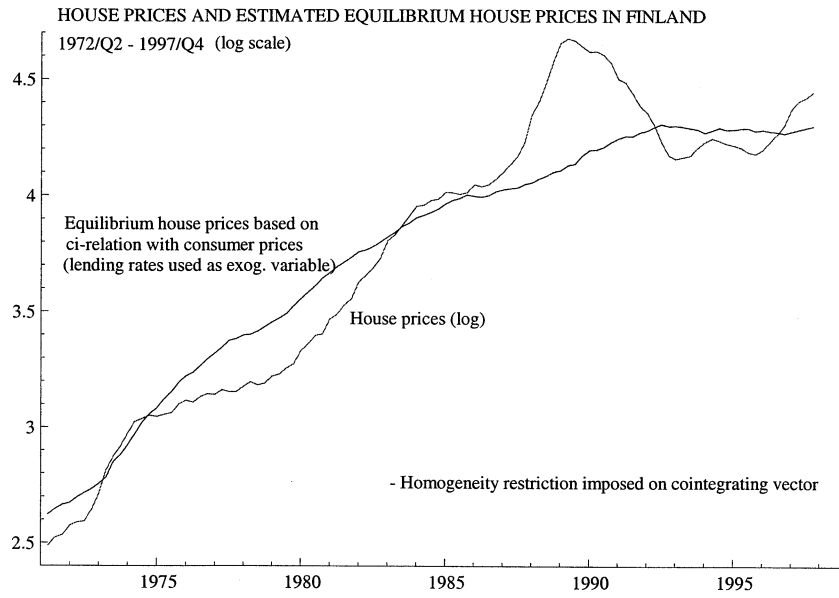


Figure 8. **Equilibrium house prices in Sweden**

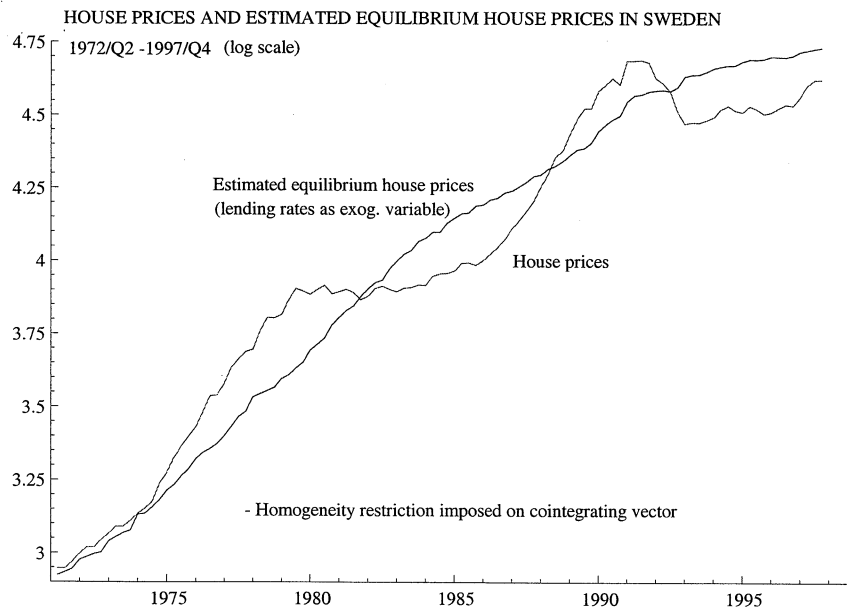


Figure 9. **Finnish house price and inflation system estimation**

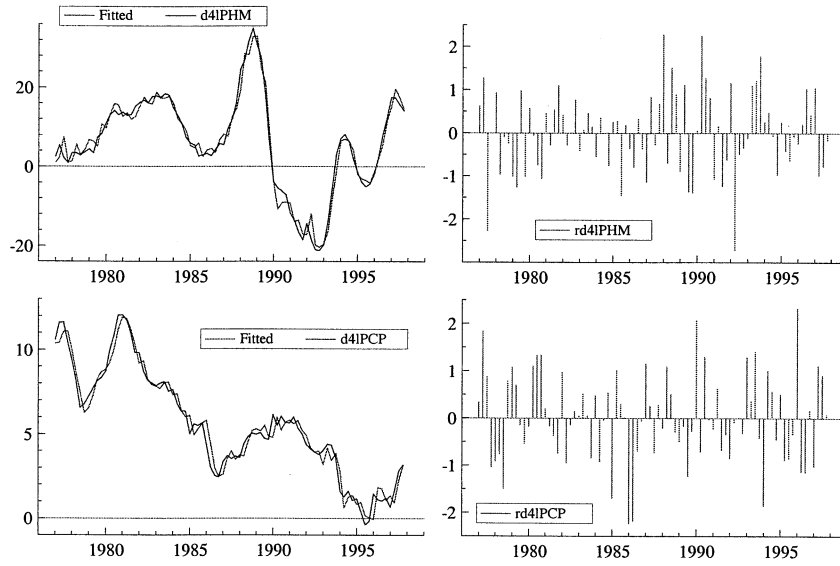


Figure 10. **Impulse responses in Finnish house price and inflation system**

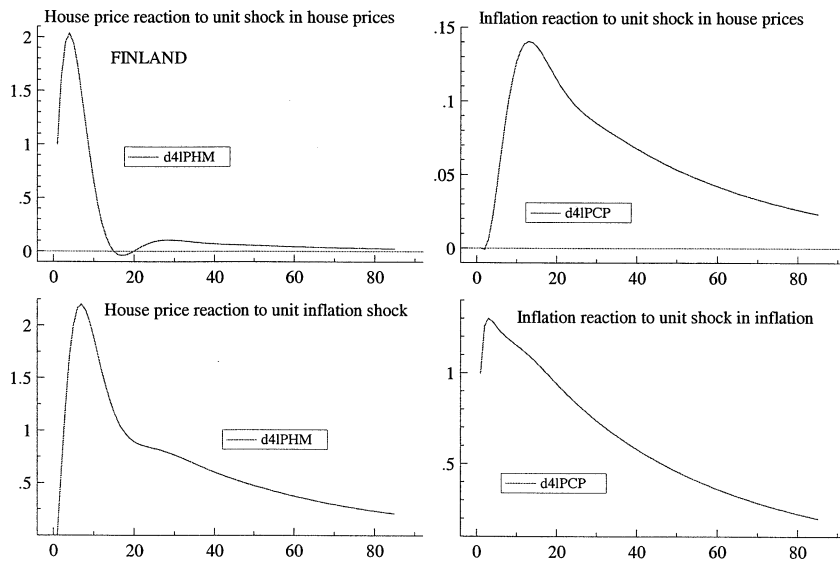


Figure 11. Swedish house prices and inflation system estimation

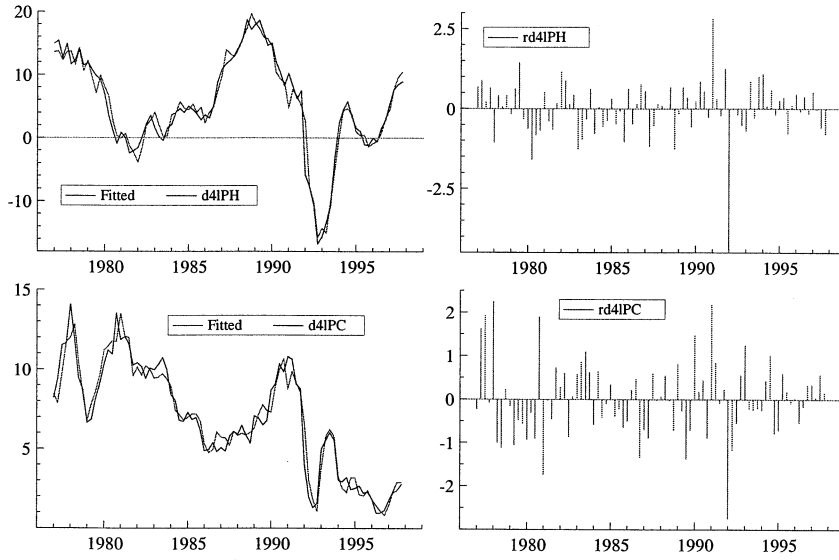
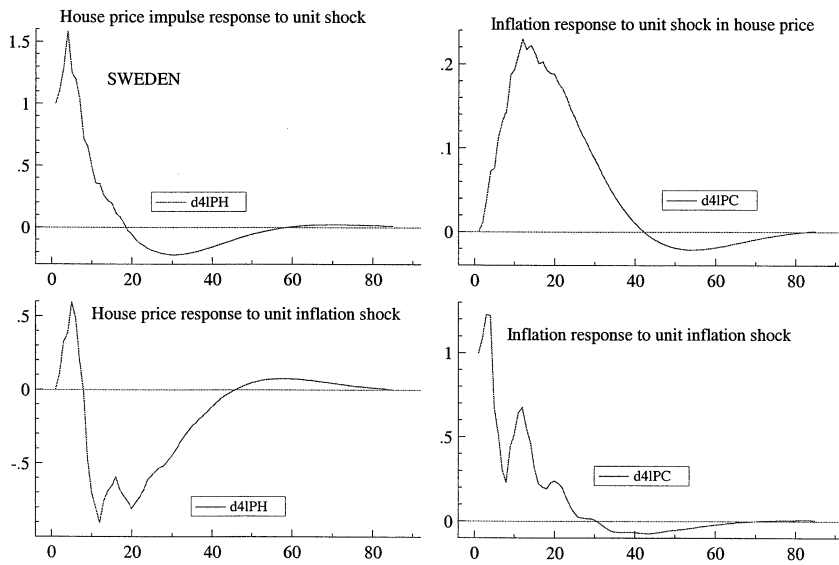


Figure 12. Swedish impulse responses



5 Conclusions

The aim of this paper has been to model house prices and inflation in such a way that the long-run property of cointegration between house and consumer prices is maintained. Cointegration between house prices and consumer goods prices may seem puzzling, since in countries such as the UK, real housing prices show a clear trend that matches well with the trend in real disposable income, while we argue here that real house prices are stationary. Our international comparison shows, however, that the UK may be likely an exception rather than a rule in this respect (see Appendix 2). It is important, of course, to distinguish whether house prices follow consumer price index or earnings index in the long run as the trends in these series likely differ. Housing expenditure has near unit income elasticity, but in Finland and Sweden house price ratio to disposable income has had a significant negative trend at least during last 30 years.

The system modelling between house and consumer prices showed that the general price level is transmitted into house prices rather quickly, but inflation is surprisingly insensitive to changes in house prices. This matches with the fact that house prices are much more volatile than non-durable consumer goods prices. Adding stock prices into the system does not improve the system considerably. We find that while house prices are clearly not neutral with respect to inflation, inflation is not affected much by shocks in house prices. It was found that in addition to the arbitrage relation between consumer and house prices there are many fundamental determinants of house prices like expected wages, unemployment rate, after-tax interest rates, which divert in the short run the actual house price development from their equilibrium level.

These results contribute a little in the way of operational advice for market participants. House prices exhibit large swings in the market, but in the long run, no excess return remains over inflation rate. Eventually and hopefully increasing information about the market equilibrium will reduce also the volatility of prices in the housing market. For monetary policy, asset price inflation does not express a major threat, but volatility in asset prices could be stabilized by trying to smooth housing demand.

Periods of high expected inflation can profoundly affect people's willingness to get into debt. Unexpected inflation can affect debt burden significantly, so a house buyer must also be careful with respect of the equilibrium in the lending market. In principle, trading takes place only at equilibrium prices, but in the long-run situations vary. During the past decade there has been a significant reduction in the subsidies for home ownership both in Finland and Sweden. The effective tax rate for mortgage interest payments deductions were limited in both countries due to capital income taxation reform. In addition, other housing subsidies have been restricted, which have lowered the present value of these tax reliefs. EMU may also work as a stabilizer for asset prices as it stabilizes inflation expectations and interest rates. Unfortunately, stabilizing interest rates may smooth housing investment, but it does not affect the demand for houses, which remains a problem.

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Appendix 1

Table 1. **House, housing investment, stock prices and inflation in Finland and Sweden, %**

Period: 1971/Q1 - 1997/Q2 ¹³

	Mean	Median	Std.dev.	Skewness	Kurtosis	Norm. p-level
House prices						
Finland	7.07	6.60	11.59	-0.26	0.31	.365
Sweden	6.23	6.35	7.45	-0.61	0.58	.037*
Inflation						
Finland	7.07	6.36	4.35	0.64	0.05	.005**
Sweden	7.37	7.23	3.16	-0.31	-0.68	.051
Real house price						
Finland	0.00	0.39	11.13	0.04	0.34	.426
Sweden	-1.14	-0.76	7.00	-0.51	0.06	.061
Housing investment prices						
Finland	7.63	8.18	7.73	-0.30	2.52	.000***
Sweden	7.93	7.47	5.93	-0.10	-0.09	.115
Stock prices						
Finland	11.63	11.75	26.93	0.05	-0.67	.407
Sweden	13.88	11.48	22.29	0.34	0.22	.308

Table 2. **Cointegration tests and basic bivariate ci-system tests for house and consumer prices**

Cointegration tests for Finland with no intercept or trend, in Sweden with no trend component and 5 lags in VAR for 1971/Q2 - 1997/Q2

Hypothesis	Eigenvalues	Max Eigenvalue tests			Trace tests				
Null	Alt.	FIN	SWE	FIN	SWE	95%	FIN	SWE	95%
r = 0	r=1	.1234	.1548	13.82*	17.65**	11.03	13.89*	18.92**	11.03
r ≤ 1	r=2	.0006	.0119	0.65	1.26	4.16	0.65	1.26	4.16

Homogeneity restriction on the bivariate price system $\beta: \beta_1 + \beta_2 = 0$

LR-test: $\chi^2(1)$	Statistic	p-value	
LR-test: $\chi^2(1) = 2.88$	[0.090]		FINLAND
LR-test: $\chi^2(1) = 4.19$	* [0.041]		SWEDEN

LR-tests for stationarity, exclusion and weak exogeneity

	EXCLUSION		STATIONARITY		WEAK EXOGENEITY		95%	99%
	$\chi^2(p-r=1)$	SWE	$\chi^2(r=1)$	SWE	$\chi^2(r=1)$	SWE		
House price	12.55**	8.21**	13.07**	5.21*	13.76**	9.60**	.84	6.65
Cons. price	13.07**	5.21*	12.55**	8.21**	0.19	0.90	3.84	6.65

¹³ It seems that in most major countries real house prices (deflated by consumer prices) have been relatively stationary during the past thirty years, although the ADF(1) test rejects a unit root strictly only for Germany, Belgium, Ireland and Norway (see also Kennedy and Andersen, 1994 p. 32-35). Real house prices were significantly different from zero only for Japan and Finland during 1970-96 (see Appendix 2). In the UK, however, there is undoubtedly a trend in real house prices which matches rather well with the trend in real disposable income (see Brown et al (1997, p. 530). For instance, mortgage rationing, which was abolished in early eighties, might be responsible for non-stationarity (G. Meen, 1990). In addition in the UK imputed rent and the capital gains on owner-occupied houses is not taxed, therefore increases in the tax advantage may also have affected the positive trend in house prices (Holly and Jones, 1997 p. 553). It is also possible that in other above-mentioned countries some sort of rationing, market failure or financial liberalization was responsible for the non-stationarity of real house prices.

Table 3. **Granger causality tests for house prices and inflation, quarterly changes (%)**

F-test probability levels (p) with different lags
Quarterly data, 1970/Q2 – 1996/Q4

		Quarterly differences		
Independent variable	Dependent variable	1-lag	3-lags	6-lags
Finnish data				
PH	PC	.132	.090	.274
PC	PH	.714	.945	.968
PH	UR	.000**	.001**	.004**
UR	PH	.175	.253	.121
PH	YW	.000**	.016*	.056
YW	PH	.304	.497	.473
Swedish data				
PH	PC	.724	.025*	.507
PC	PH	.008**	.041*	.545
PH	UR	.001**	.005**	.000**
UR	PH	.418	.666	.616
PH	YW	.092	.000**	.177
YW	PH	.680	.259	.540

Granger tests for trivariate blocks (with 5 lags)

Dependent	Conditional	χ^2 -test p-value	
		Finland	Sweden
PH	PC, STOCK	.011 *	.008 **
PC	PH, STOCK	.054	.034 *
STOCK	PH, PC	.137	.006 **

Variables:

PC	= Private consumption deflator, %
PH	= House price change, %
UR	= Unemployment rate change, %
YW	= Wages change, %
STOCK	= Stock prices, %

Table 4. Cointegration with no intercepts or trends in the VAR(5) for Finland, 1971Q2–1997Q2

Endogenous: LPHM, LPCP, LHEX (FINLAND)
Weakly exogenous I(0) variables: RLBN

MULTIVARIATE STATISTICS

TRACE CORRELATION = 0.538

TEST FOR AUTOCORRELATION

L-B(26), $\chi^2(195)$ = 176.093, p-val = 0.83
LM(1), $\chi^2(9)$ = 11.522, p-val = 0.24
LM(4), $\chi^2(9)$ = 4.590, p-val = 0.87

TEST FOR NORMALITY

$\chi^2(6)$ = 6.591, p-val = 0.36

TEST FOR EXCLUSION: LR TEST $\chi^2(r)$

r	DGF	$\chi^2(5)$	LPHM	LPCP	LHEX	RLBN
1	1	3.84	9.96	5.92	5.69	6.60
2	2	5.99	12.33	7.76	8.07	7.20

TEST FOR STATIONARITY: LR TEST $\chi^2(p-r)$

r	DGF	$\chi^2(5)$	LPHM	LPCP	LHEX
1	3	7.81	14.05	14.01	14.04
2	2	5.99	4.20	4.33	4.44

TEST FOR WEAK-EXOGENEITY: LR TEST $\chi^2(r)$

r	DGF	$\chi^2(5)$	LPHM	LPCP	LHEX
1	1	3.84	3.37	8.15	1.65
2	2	5.99	3.41	9.99	4.13

List of eigenvalues in descending order:

.1765 .0812 .0129

Null	Alternative	Max. eigenvalue	stat.	Trace statistic	95% CV
r = 0	r = 1	20.3934*	17.68	30.6410*	24.05
r ≤ 1	r = 2	8.8875	11.03	10.2477	12.36
r ≤ 2	r = 3	1.3602	4.16	1.3602	4.16

Estimated Cointegrated Vectors (free and restricted) in Johansen Estimation (Standardized in brackets)

	β -vector	Rest. $\beta_1+\beta_2+\beta_3=0$
LPHM	.82171 (-1.0000)	-.50998 (-1.0000)
LPCP	-.24956 (.3037)	.53481 (1.0487)
LHEX	-.24136 (.2937)	-.024836 (-.0487)

LR Test $\chi^2(1) = 11.4885[.001]$

Estimated Long Run Matrix in Johansen Estimation

	LPHM	LPCP	LHEX
LPHM	-.0104	.0032	.0031
LPCP	-.0229	.0070	.0067
LHEX	.0340	-.0103	-.0099

	LPHM	LPCP	LHEX	t-values
DLPHM	-0.046	0.021	0.010	2.12
DLPCP	-0.022	0.010	0.005	3.65
DLHEX	-0.126	0.059	0.028	1.63

New variables:

LHEX = Log of Helsinki stock exchange price index
RLBN = Bank lending rate for new loans to public, %

Table 5. **Cointegration with no intercepts or trends in the VAR(5) for Sweden, 1971Q2–1997Q2**

Endogenous: LPH, LPC, LSWE (SWEDEN)
Weakly exogenous I(0) variables: RUT

MULTIVARIATE STATISTICS

TRACE CORRELATION = 0.471

TEST FOR AUTOCORRELATION

L-B(26), $\chi^2(195)$ = 234.938, p-val = 0.03
LM(1), $\chi^2(9)$ = 8.368, p-val = 0.50
LM(4), $\chi^2(9)$ = 29.757, p-val = 0.00

TEST FOR NORMALITY

$\chi^2(6)$ = 43.612, p-val = 0.00

TEST FOR EXCLUSION: LR TEST CHISQ(r)

r	DGF	$\chi^2(5)$	LPH	LPC	LSWE	RUT
1	1	3.84	9.50	10.82	0.00	5.76
2	2	5.99	13.84	14.17	0.06	5.77

TEST FOR STATIONARITY: LR TEST CHISQ(p-r)

r	DGF	$\chi^2(5)$	LPH	LPC	LSWE
1	3	7.81	13.85	15.66	17.38
2	2	5.99	2.05	3.96	7.42

TEST FOR WEAK-EXOGENEITY: LR TEST CHISQ(r)

r	DGF	$\chi^2(5)$	LPH	LPC	LSWE
1	1	3.84	4.74	4.70	0.03
2	2	5.99	10.13	10.75	0.53

List of eigenvalues in descending order:

.25316 .065185 .1671E-3

Null	Alternative	Max. eigenvalue stat.		Trace statistic	
		Statistic	95% CV	Statistic	95% CV
r = 0	r = 1	30.0661*	17.68	37.0261*	24.05
r ≤ 1	r = 2	6.9428	11.03	6.9600	12.36
r ≤ 2	r = 3	.017212	4.16	.017212	4.16

Estimated Cointegrated Vectors (free and restricted) in Johansen Estimation (Standardized in brackets)

	β -vector	Rest. $\beta_1+\beta_2+\beta_3=0$
LPH	1.0801 (-1.000)	-.9923 (-1.000)
LPC	-1.3258 (1.2275)	.9516 (.9589)
LSWE	.1222 (-.1132)	.04071 (.0410)

LR-test $\chi^2(1) = 3.6447[.056]$

Estimated Long Run Matrix in Johansen Estimation

	LPH	LPC	LSWE	t-values
LPH	-.0688	.0845	-.0078	
LPC	.0212	-.0259	.0024	
LSWE	-.2280	.2799	-.0258	
DLPH	-0.051	0.053	-0.000	-2.924
DLPC	0.028	-0.029	0.000	2.967
DLSWE	0.018	-0.019	0.000	0.198

New variables:

LSWE = Log of Stockholm stock exchange price index
RUT = Bank lending rate, %

Table 6. House prices and inflation error-correction system for Finland

The present sample is: 1977/Q1 - 1997/Q4, estimation by FIML

Equation 1 for d41PHM

Variable	Coefficient	Std.Error	t-value	t-prob	HCSE
d41PHM_1	1.3733	0.09882	13.896	0.0000	0.10422
d41PHM_2	-0.5319	0.10145	-5.243	0.0000	0.10338
d41PCP_2	-0.3391	0.13529	-2.507	0.0143	0.13981
CI _{fin} _4	-0.0444	0.02123	-2.093	0.0397	0.02008
atRLBN	-0.4193	0.45136	-0.929	0.3558	0.37451
d4UR_1	0.2071	0.23802	0.870	0.3868	0.22524
d41LBH	0.1939	0.05758	3.367	0.0012	0.06082
Constant	2.6365	2.7283	0.966	0.3369	---
$\sigma = 2.3605$					

Equation 2 for d41PCP

Variable	Coefficient	Std.Error	t-value	t-prob	HCSE
d41PHM_4	-0.02632	0.01244	-2.114	0.0378	0.01197
d41PCP_1	1.03690	0.07222	14.357	0.0000	0.07406
d41PCP_3	-0.14225	0.06348	-2.241	0.0280	0.06271
d41YW_2	0.03490	0.02530	1.379	0.1718	0.02814
atRLBN_2	-0.20566	0.11125	-1.849	0.0684	0.10002
d41LBH	-0.23526	0.12984	-1.812	0.0740	0.11452
d41LBH_1	0.28712	0.25126	1.143	0.2567	0.24561
d41LBH_2	-0.02517	0.14760	-0.171	0.8651	0.15558
Constant	1.3481	0.69185	1.949	0.0550	---
$\sigma = 0.6372$					

loglik = -25.860752 log|\Omega| = 0.615732 |\Omega| = 1.85101 T = 84
LR test of over-identifying restrictions: $\chi^2(27) = 48.4967 [0.0068]$ **

correlation of residuals

	d41PHM	d41PCP
d41PHM	1.0000	
d41PCP	-0.020769	1.0000

Diagnostics summary:

	Test statistic	p-value
d41PHM :Portmanteau 9 lags	= 16.745	
d41PCP :Portmanteau 9 lags	= 23.628	
d41PHM :AR 1- 5 F(5, 57)	= 12.690	[0.0000] **
d41PCP :AR 1- 5 F(5, 57)	= 6.201	[0.0001] **
d41PHM :Normality $\chi^2(2)$	= 1.348	[0.5096]
d41PCP :Normality $\chi^2(2)$	= 0.662	[0.7179]
d41PHM :ARCH 4 F(4, 54)	= 0.230	[0.9199]
d41PCP :ARCH 4 F(4, 54)	= 0.423	[0.7907]
d41PHM : χ^2 F(42, 19)	= 0.918	[0.6047]
d41PCP : χ^2 F(42, 19)	= 0.306	[0.9993]
Vector portmanteau 9 lags	= 54.299	
Vector AR 1-5 F(20,130)	= 2.519	[0.0010] **
Vector normality $\chi^2(4)$	= 2.062	[0.7244]
Vector χ^2 F(126,93)	= 0.968	[0.5703]

Finnish model variables:

PHM = House price index
PCP = Private Consumption Deflator
atRLBN = After tax bank lending rate for new loans, %
UR = Unemployment rate, Statistics Finland, %
LBH = Banks loans to households, Million FIM
YW = Wage sum, Million FIM
CI_{fin} = Error-correction term

Table 7. House prices and inflation error-correction system for Sweden

The present sample is: 1977/Q1 - 1997/Q4, estimation by FIML

Equation 1 for d41PH					
Variable	Coefficient	Std.Error	t-value	t-prob	HCSE
d41PH_1	0.89175	0.037252	23.938	0.0000	0.03899
d41PC_1	0.14736	0.11912	1.237	0.2198	0.11318
CIsw_e_4	-0.08791	0.02920	-4.034	0.0001	0.01758
atRUT_1	-0.55529	0.14698	-3.778	0.0003	0.21652
Urswe_4	-0.05600	0.12023	-0.466	0.6426	0.11125
Constant	5.6272	2.0854	2.698	0.0085	---
$\sigma = 2.1596$					

Equation 2 for d41PC					
Variable	Coefficient	Std.Error	t-value	t-prob	HCSE
d41PH_5	-0.06308	0.02700	-2.336	0.0221	0.02486
d41PC_1	0.67041	0.07097	9.446	0.0000	0.07637
CIsw_e_4	0.04867	0.01363	3.571	0.0006	0.01138
d41Wages	0.04147	0.03642	1.139	0.2583	0.03790
d41Wages_1	0.10653	0.03865	2.756	0.0073	0.04527
URswe_1	-0.13394	0.07162	-1.870	0.0652	0.06450
Constant	1.4768	0.91183	1.620	0.1094	---
$\sigma = 1.0541$					

loglik = -59.47386 $\log|\Omega| = 1.41604$ $|\Omega| = 4.12079$ T = 84
 LR test of over-identifying restrictions: $\chi^2(35) = 93.4419$ [0.0000] **

correlation of residuals		
	d41PH	d41PC
d41PH	1.0000	
d41PC	0.27874	1.0000

Diagnostics summary:		Test statistic	p-value
d41PH	:Portmanteau 9 lags	= 27.903	
d41PC	:Portmanteau 9 lags	= 40.014	
d41PH	:AR 1- 5 F(5,55)	= 11.000	[0.0000] **
d41PC	:AR 1- 5 F(5,55)	= 7.411	[0.0000] **
d41PH	:Normality $\chi^2(2)$	= 48.689	[0.0000] **
d41PC	:Normality $\chi^2(2)$	= 15.853	[0.0004] **
d41PH	:ARCH 4 F(4, 52)	= 0.946	[0.4448]
d41PC	:ARCH 4 F(4, 52)	= 3.887	[0.0078] **
d41PH	:Xi2 F(46, 13)	= 0.814	[0.7088]
d41PC	:Xi2 F(46, 13)	= 0.500	[0.9571]
Vector portmanteau	9 lags	= 77.06	
Vector AR 1-5	F(20,134)	= 3.527	[0.0000] **
Vector normality	$\chi^2(4)$	= 47.679	[0.0000] **
Vector Xi2	F(138, 87)	= 1.881	[0.0008] **

Swedish model variables:

PH = House price index
 PC = Private consumption deflator
 atRUT = After-tax lending rate, %
 URswe = Unemployment rate, %
 Wages = Wage sum, SKR
 CIsw_e = Error-correction term

Appendix 2

Figure 13. Real house prices in 12 countries, 1970-1996

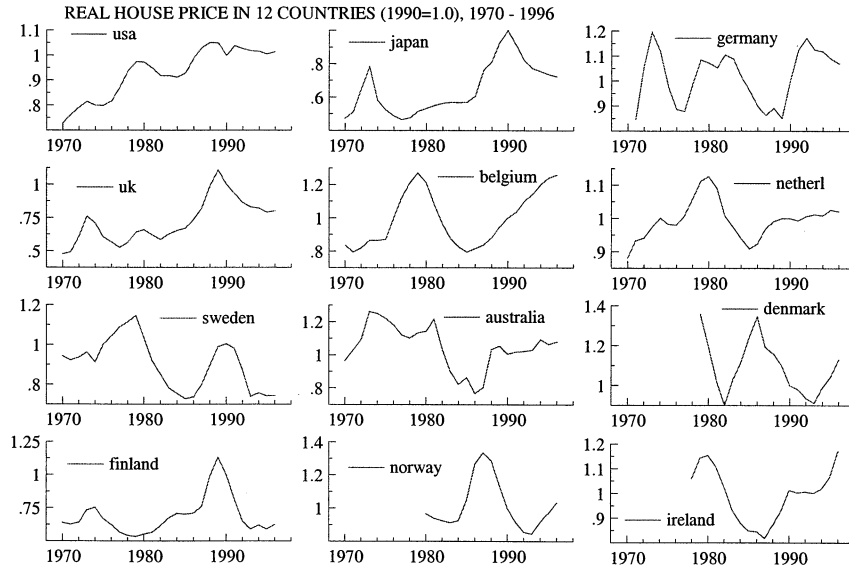


Table 8. Stationarity of real house prices (deflated by CPI) in 14 countries, 1970-1996

Country	Period	Mean (α)	Std. dev.	t-value α # 1	Norm.test p-value	Unit root test		Cointegration test	
						ρ - coeff. test	ADF(1) test	Trace test	β -homo- geneity p-value
USA	70-96	.931	.097	-0.71	0.025*	.899	-1.76	41.76**	.0009**
JAPAN	70-96	.654	.153	-2.26*	0.056	.817	-1.90	18.28**	.394
GERMANY	71-96	1.020	.104	0.20	0.167	.536	-4.07**	20.72**	.894
UK	70-96	.721	.160	-1.74	0.265	.809	-2.36	29.46**	.0003**
BELGIUM	70-96	.995	.161	-0.03	0.018*	.824	-3.64*	38.67**	.0000**
NETHERLANDS	70-96	.994	.056	-0.11	0.351	.746	-2.97	33.62**	.0000**
SWEDEN	70-96	.904	.125	-0.77	0.435	.787	-2.44	47.51**	.0000**
AUSTRALIA	70-96	1.048	.129	0.37	0.529	.732	-2.13	34.95**	.0000**
FINLAND	70-96	.690	.143	-2.17*	0.001**	.702	-2.90	8.35	.8871
IRELAND	78-96	.995	.106	-0.05	0.734	.856	-3.26*	22.70**	.0002**
NORWAY	80-96	1.017	.146	0.12	0.013*	.649	-4.12**	9.63**	.0000**
DENMARK	79-96	1.091	.134	0.68	0.498	.552	-2.20	21.02**	.0001**
FRANCE	85-96	.915	.069	-1.23	0.523	.669	-4.26*	34.02**	.0000**
CANADA	83-96	.927	.123	-0.59	0.602	.593	-2.41	18.50**	.0003**

The presence of a unit root in real house prices is rejected by the ADF tests for Germany, Belgium, Ireland, Norway and France. The mean on real house prices diverts from unity (here arbitrarily 1990=1.0) only for Japan and for Finland during 1970-1996. Critical values for the ADF(1) tests are -2.985 for 5% significance level and -3.72 for 1% significance level for 1970-96.

The trace test between house and consumer prices has been adjusted with the number of degrees of freedom (T-nm) instead of T available in PcFiml, the critical values 95% significance level (see Doornik and Hendry, 1995). In all countries except Finland, cointegration prevails; homogeneity is not rejected for only a few countries (Japan, Germany and Finland). As can be seen from Figure 13, house price cycles can take a long time and adjustment to various shocks is slow.

Bankruptcies, indebtedness and the credit crunch

Kari Takala • Matti Virén

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Bankruptcies, Indebtedness and the Credit Crunch¹

Kari Takala² and Matti Virén³

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Abstract. This paper deals with Finnish bankruptcies. It shows that bankruptcies are strongly related to the business cycle and that they are perhaps even more strongly related to indebtedness, real interest rates and asset prices. The importance of these financial factors probably increased when the financial markets were liberalized in the early 1980s. Although there is a lot of seasonal and cyclical variation in bankruptcies the long run level (especially when adjusted to the number of firms) is almost constant representing some sort of "a natural rate of bankruptcies". What makes bankruptcies so important is the fact that they directly affect production, employment and credit expansion. The credit crunch effect in particular is scrutinized in the paper.

Keywords: bankruptcy, financial distress, credit crunch

1 Introduction

One of the prime purposes of a bankruptcy is to settle accounts with creditors and to establish a market value for the company as a whole. A major task in bankruptcy settlement is to prevent a firm from running into further debt, which is the main concern of the creditors. In most bankruptcies the amount of debt is greater than the value of the assets, which leads to financing costs to both creditors and debtors. Another difficulty crops up here, since a company is usually worth more as an operational unit than as the sum of its separate parts. It has been emphasized that a firm is a functional entity, and

by far the greatest part of its value is imbedded in the cooperation between its employees.

In general, it can be argued that bankruptcies result mainly from bad management, unnecessarily risky or unlucky investment projects or, as in recent times, unexpected rapidly diminishing demand. In Finland, bankruptcies have emerged as a macroeconomic problem recently as a consequence of an unforeseen rapid decline in GDP and in total domestic demand. Several factors, like the collapse in trade with the former Soviet Union, deterioration of the terms of trade, increased foreign indebtedness because of devaluations and rising real interest rates, can be seen as primary causes of recession. However, the first hints from the growth in bankruptcies can be traced back already to the start of financial liberalization in 1983. There seems to be some evidence that the easing of bank lending and weaker ties to clients, and hence less control feedback from firms to banks, is responsible for the large losses seen during the recent recession.

The overheating of the Finnish economy occurred after 1986, as interest rate regulation was abolished and the obstacles to capital movements were gradually removed. During the period of overheating, which lasted from late 1986 up to the spring of 1989, a very large number of new firms were started up. Most of the financing came in the form of bank loans. Thus, bank lending increased at a real rate of 20–30 % during this period, which obviously induced a huge increase in the prices of all financial and real assets. At the same time, indebtedness increased, of course, and when income and asset prices started to fall, indebtedness became a very serious problem.

Indebtedness, in the face of an exceptionally deep recession, was an obvious cause of the wave of bankruptcies and the credit crunch which are studied in this paper. The Finnish case is not, of course, exceptional, although the magnitude of the crisis makes it an interesting case for empirical analysis (for the sake of comparison, see e.g. Gunther, et al. 1995 and Shrieves and Dahl 1995 for the U.S. case).

In this paper we try to develop a macroeconomic model of bankruptcies. For this purpose we first look at certain stylized facts regarding Finland. We make use of data which cover a relatively long period, 1920–1994. The data are monthly, although we use mainly annual frequencies in the empirical analysis. The modelling is based on a cointegration analysis which deals with bankruptcies and certain important macroeconomic variables, both financial and non-financial. In addition to bankruptcies we model bank lending, or strictly speaking credit expansion, and total output. The purpose of this type of modelling is to see the extent to which financial variables, along with cyclical macroeconomic variables, affect bankruptcies and what kind of feedback effect exists between bankruptcies and these variables. Specifically, the credit crunch hypothesis is subjected to testing.

2 Some background on bankruptcies

2.1 Historical background

Under the law, bankruptcy petitions and proceedings are registered by the courts and related data is gathered by Statistics Finland. The number of monthly bankruptcy proceedings¹ have been available since 1922 (see Figure 2). The previous large boom in bankruptcies can be linked with the Great Depression in the early 1930s. The current high level of bankruptcies is clearly unprecedented. Even if the total number firms in existence is taken into account, the level is very high – something one could probably not forecast a decade ago (for U.S. evidence, see e.g. Meehan 1993).

Bankruptcy proceedings can also be analysed using time series components like trend, seasonal component and irregular variation. The original monthly series look quite volatile even in logs. Analysis of the structural time series model shows that the level of bankruptcies has a large variance, but the trend is fairly stable. There is clearly also seasonal variation in bankruptcies, but the pattern of seasonal variation has changed significantly over the decades. Currently, the seasonal peaks are in January and September – November, while the lowest level of proceedings are in the summer and in December. The model estimations also show that the irregular variance of bankruptcies has been a major component of the total variation of bankruptcy proceedings (see Takala and Virén 1994). It could be argued that the process of generating bankruptcies during the Second World War was quite different at least from the period of financial regulation, which lasted from after the war up to 1983.

It is useful to compare the number of bankruptcies to the total number of firms. The number of firms itself has increased faster than population (see Table 1). The structure of production could also affect the number and share of bankruptcies. Industrial companies have been relatively big in Finland, but the increased amount of small service companies may also have raised the number of bankruptcies relative to firms. This may reflect the change in the structure of production, as the number of service firms has increased with rising GDP.

¹ A conceptual distinction could be made between two measures of bankruptcies. We speak of bankruptcy petitions (applications) registered by the courts and bankruptcy proceedings accepted by the courts for further action. In practice there could be several bankruptcy applications made by several creditors regarding the same firm, whereas proceedings register only one case for each firm. It is also possible that the debtor himself could apply for bankruptcy. The bankruptcy resettlement procedure available from 1993 must be applied and approved by the debtor himself.

Table 1. Bankruptcies and number of firms in Finland

Year	Number of bankruptcies	Number of firms	Population in thousands	Bankruptcies/population, %	Bankruptcies/firms, %
1922	725	6763	3228	.022	1.97
1930	1945	10410	3463	.056	2.66
1940	265	15068	3696	.007	0.33
1950	406	24030	4030	.010	0.69
1960	829	32011	4446	.019	1.06
1970	1361	45352	4598	.030	1.24
1980	1057	56134	4788	.018	1.16
1985	2122	109806	4911	.043	2.57
1990	3588	133321	4998	.072	2.12
1991	6155	125121	5029	.122	4.18
1992	7348	125700	5055	.145	5.02
1993	6769	117295	5080	.133	4.93
1994	5502	118000*	5099	.108	3.77*

Starred values are forecasts. Bankruptcies/firms (i.e., the business failure rate) is computed in terms of corporate bankruptcies (i.e., individual bankruptcies are excluded).

The number of bankruptcies varies with the phase of the business cycle. When GDP grows rapidly the ratio of bankruptcies to number of firms is small. This is a result of both a smaller number of bankruptcies and a larger number of new firms start-ups. In a recession the number of bankruptcies will increase while the number of firms increases slowly or even decreases.

The number of bankruptcies depends on various factors. Firms that go into bankruptcy are mainly small companies with heavy debt with respect to cash flow or net profits. In these small companies the personal losses of entrepreneurs are also the largest. Even if we leave out those bankruptcies that have taken place without further demands on debt capital, the equity of bankrupt firms is on average only half of their total debt. In addition to the size of the firm, the industry, the phase in the business cycle and the capital structure predict the probability of bankruptcy.

In addition to these macroeconomic indicators, a few microeconomic indicators have proved to be useful in predicting bankruptcies. The number of payment failures precedes bankruptcies at the firm level as well as in the aggregate. Unfortunately, the series on payment failures covered only the period 1987 to 1994 and cannot be used in the present context.

During the 1980s the distribution of new bankruptcies in different industries was relatively stable. Most of the bankruptcies occurred in

commerce (28 %) and manufacturing (23 %), followed by construction and services, each with about a 16 per cent share (see Figure 3). The devaluation of the markka in November 1991 and the float starting in September 1992 shifted bankruptcies from the export (open) to the closed sector. One worrying feature of the recent bankruptcy boom is the fact that the share of bankruptcies applied for by debtors itself has been increasing. Whereas normal bankruptcy applications are used as means of collecting debt, this is not the case when a debtor itself applies bankruptcy.

Later we cite evidence that bankruptcies could be an indicator of an equilibrium process with supply being equated to diminishing demand. If the slowdown in demand is fast enough, there is no time to cut production and other firms try to keep up their cash flows as well. In this case firms with excess debt will get into difficulties and later on will reach the final dead end. This theory is based on the fact that total demand and supply will be cointegrated in the long-run. Despite the fact that demand and supply are integrated of order one, the bankruptcies/companies ratio will be stationary as one linear combination between these variables. Bankruptcies nevertheless have a positive mean and finite variance.

Bankruptcies are obviously related to employment and unemployment. Bankruptcies directly create unemployment. The causal relationship, however, is more complicated because unemployment can cause bankruptcies via decreased demand. In this study we cannot thoroughly analyze the bankruptcy-unemployment relationship because the historical unemployment data is somewhat deficient. Suffice it to mention that for a short sample period (1960–1993) we found that the causation goes unambiguously from bankruptcies to unemployment, not vice versa.

Money market liberalization seems to have affected the bankruptcy generation mechanism in Finland and other Nordic countries. This can be seen directly from the plot of bankruptcies. The number of bankruptcies started to rise even during 1984 (although the economy grew rather fast, at the rate of 3–5 %, until 1989). The regulation of bank lending kept the bankruptcy figures low up to the mid-1980s. After this regulation was loosened, the tight control of banks ended suddenly. For firms, financing through the stock market also became more attractive. However, debt-equity ratios began to rise slowly already in 1985. In Finland an important turning point in financing was achieved when firms involved in foreign trade started to intermediate foreign loans through their accounts. Banks demanded similar operating room and started to rapidly expand their currency loan portfolios. When the regulation of lending interest rates was abolished in autumn 1986, the supply of bank loans increased rapidly.

The increase in real bank lending rose up to as high as 30 per cent p.a. in 1986–1989. Therefore, an increase in bankruptcy was to be expected sooner

or later. What was unknown at the time was that the economic slowdown would be as steep as it turned out to be. Firms' indebtedness has had the effect of a rising real interest rates very sharply for firms operating in the closed sector of the economy. These problems were not relieved with the devaluation of the markka in November 1991. Firms with foreign debt suffered from the devaluation, and those firms which operated in the domestic sector, i.e. which had only domestic returns, faced the biggest problems. They had large capital costs, wages were sticky (in fact, wage costs even increased because of the unemployment compensation system) and prices could not be increased because of the overall excess supply in the domestic markets. Bankruptcies created further bankruptcies because some bankrupt firms continued their activities under the bankruptcy authority. In many cases, these firms created market disturbances because they demanded much lower prices – they had no cost worries!

The increasing risk of bankruptcy in the late 1980s and early 1990s is to a large extent a consequence of the rapid growth of financing and thereof of the number of firms. Over a half of the bankrupt firms have been operating under five years. Firms 2-3 years in age have had the highest risk of ending up in bankruptcy.

2.2 The credit crunch and bankruptcies

The role of bank lending is crucial in the generation of bankruptcies, since bank debt is the major source of financing to small and medium sized Finnish firms. In every recession bank lending and credit availability decelerate. Therefore, it is useful to look at whether credit availability is now more restrictive than in similar declining phases of previous business cycles. Historically, credit crunches have started from a decline in bank deposits. The current credit crunch, however, is more or less linked to the decline in the asset prices and in collateral values and, of course, to the fall in income and the resulting failure in debt servicing. This has caused debt losses and therefore, through shrinking bank equity, forced banks to cut their lending.

In Finland, as in other Nordic countries, the state has guaranteed the BIS capital ratio requirement of 8 per cent. In this sense the credit crunch is a consequence of a capital crunch. A major reason for the debt losses has been the drop in real estate prices as well as all other asset prices. The capital crunch has especially increased bankruptcies among small and medium size firms in the closed sector. This is natural since they have relied heavily on bank credit for financing. The attendant loss in bank capital is the main difference between the current business cycle and previous ones since the Second World War.

The background for the Finnish case is such that it could come directly from a textbook. The financial markets were liberalized in a situation where the economy was experiencing one of the strongest booms since the second World War. Interest rate control was abolished, first from lending rates and, after about one year, from deposit rates. Capital controls were also abolished. Demand for credit was exceptional high because of a backlog of unsatisfied excess demand, high income growth expectations and relatively low real interest rates.

Before liberalization, there had been a very long period of excess demand and credit rationing. During that period banks had very close relationships with their customers. New customers were carefully scrutinized before they got bank loans and many of them did not get loans at all. In the case of a household, a very long history as a customer and a large downpayment were required.

With liberalization the importance of customer relationships diminished (at least temporarily). Obviously this weakened banks ability to monitor the quality of their customers. Perhaps more important, however, was the fact that banks started to compete for market shares. Banks have relatively few instruments which they can use in competition. In the Finnish case, lending was used as an instruments of competition. Thus, more advantageous lending terms were offered to new customers.² It comes as no surprise that those banks that competed hardest for new clients got a disproportionately large share of the bad clients with high credit risk and, in some cases, even criminal intentions (this is something one might expect on the basis of the principles of adverse selection moral hazard). In particular, savings banks adopted this kind of very aggressive growth strategy which later on led to complete disaster.

Savings banks (and to some extent cooperative banks) tried to expand their lending to the corporate sector. Because of scale considerations and customer/ownership relationships, they had to concentrate on small firms operating in the domestic market. This sectoral concentration created considerable credit risk, which was unfortunately actualized during the recent recession. As for the commercial banks, also they competed heavily for market shares. Much of their resources was used in "ownership races" in which the banks tried to gain or secure ownership in the largest firms.

Still another problem was caused by developments in the real estate market. In Finland, as in most OECD countries (see O'Brien and Browne 1992), all banks increased their lending in the real estate market relative to

² There is some evidence suggesting that banks were very poor in pricing the risks of their loan contracts. The risk premia were in some cases even negative! See e.g. Murto (1993).

the industrial and commercial sectors. It would be an exaggeration to argue that the whole housing boom originated from excessive credit expansion. In Finland the housing boom in the late 1980s was perhaps the most striking in the whole OECD area (see e.g. Loikkanen et al. 1992). When the market collapsed after 1989, banks faced a huge risk exposure. There was huge overcapacity in the construction industry and a huge stock of new unsold houses. House prices fell in real terms more than 50 per cent in the early 1990s (land prices also fell considerably although the drop in stock prices was still much more dramatic).

As a consequence Finnish banks became fragile and financially vulnerable to lower credit quality, declining profitability and deflation of collateral values. Much of this change could be seen as cyclical, but the heart of the problem seems to be structural. It has been estimated that it will take at least until the end of the decade for the increased indebtedness to melt away. The Finnish experience repeats the similar history of a bad slump in the real estate business in the US, Norway and UK (for the US case, see e.g. Syron 1991).

In Finland, the depression lasted until 1993 and an upturn started in 1994. Bankruptcies have not yet, however, decreased to the pre-recession level (nor has unemployment). Perhaps the most important change which has taken place is the decrease in indebtedness (see Figure 2). Bank lending has considerably decreased since 1990 in both nominal and real terms. It can be argued that this is caused by both demand and supply effects. The demand for credit has decreased both because of reduced investment activity and the need to restore a more healthy financial structure. Increased uncertainty may also have contributed to this course of development.

On the supply side, banks have experienced unprecedented credit losses and all banks have been in serious difficulties regarding bankruptcy or merger (or, more probably, government takeover). There are several signs that banks' behaviour has changed towards the pre-liberalization period rules. This, in turn, shows up in more stringent lending conditions, choice of customers, collateral requirements etc. Thus, non-price rationing is again used to some extent. Obviously, it is very difficult to say how much of the decrease in bank lending is caused by demand and supply considerations. In the subsequent empirical analysis we try identify both effects, but quite naturally we are more interested in the supply effects. That is because it is almost self-evident that there are demand effects. Whether there are important supply, or credit crunch, effects is already a more controversial question.

3 Empirical analysis

3.1 Outline of the analysis

In this chapter we model the behaviour of bankruptcies, credit expansion and output. The emphasis is on the analysis of bankruptcies. Thus, we try to find out to what extent bankruptcies depend on the main financial variables: indebtedness, real interest rates and stock prices. In addition, we scrutinize the importance of certain other macroeconomic variables which should matter, especially in a small open economy framework: the terms of trade, the real exchange rate, labour costs and, of course, aggregate demand.

When modelling credit expansion we pay particular attention to the reverse relationship between bankruptcies and bank credit. Thus, we try to determine whether credit expansion – when it is controlled by various determinants of the supply of and demand for bank loans – is indeed sensitive to bankruptcy risk. If a negative relationship can be asserted between bankruptcies and credit expansion, we may conclude that the credit crunch hypothesis is not completely at odds with the (Finnish) data.

Finally, we consider the link between total output and bankruptcies. The question is then whether bankruptcies help in predicting output developments. This question is analyzed with the help of a relatively simple reduced form output growth equation, which also includes stock prices together with some more conventional determinants of output.

In modelling these variables an obvious starting point would be the analysis of co-integration (see Engle and Granger 1987 and Johansen 1991). We make use of this analysis although – at least at this stage – we cannot fully utilize the co-integration framework in building the empirical model for all of these variables. In some earlier analyses (see Takala and Virén 1994 for details) it turned out that bankruptcies, output and credit are co-integrated with one (and no more or no less) co-integration vector.

It is obvious, however, that the cointegration relationship is more complicated, at least in a setting in which we focus on the long-run development of an economy. Complications came especially from certain measurement problems. It is very difficult to get reliable measures of the number of firms and so to get a precise idea of the true importance of bankruptcies. Other problems concern the measurement of debt and financial assets. We have relatively good data on banks' Finnish markka loans to the public but the data on foreign loans is very deficient. Unfortunately, the latter have constituted a significant portion of firms' financing in certain periods of time. We suspect that this "missing credit" problem is also the reason why it

is so difficult to establish a reliable cointegration relationship for the determination of bank loans.

Although we still intend to build a complete dynamic model for the key variables in our study, at this point we adopt a more modest approach by specifying some simple single-equation models for the above-mentioned three variables. The dynamic specifications are also quite "old-fashioned" in the sense that we apply the conventional partial adjustment approach rather than the co-integration cum error-correction model strategy. In the case of bankruptcies, however, we use both approaches in building the estimating models.

As a first step in the empirical analysis, we scrutinize the time series (unit root) properties of the data series. Most of our data are monthly although some key variables are available only on an annual basis. Hence the analysis is carried out with both frequencies. The results from these analyses are reported in Table 2.

It is not difficult to see that the data for output, financial assets and liabilities, as well as for bankruptcies, are characterized by unit roots, while interest rates, terms of trade and the real exchange rate are roughly stationary $I(0)$ variables. This distinction between the variables should obviously be kept in mind when building the estimating models - at least to avoid nonsense regression models.

As far as bankruptcies are concerned there are two quite different alternatives. Either bankruptcies are stationary or some equilibrium error between bankruptcies and, say, indebtedness and demand is stationary. The first alternative is a not a bad approximation, in particular when the number of bankruptcies is adjusted to the number of firms. Then some sort of "a natural rate of bankruptcies" emerges. Unfortunately, the second alternative does also get some support from the data. In fact, the quality of the data is not sufficiently good to allow for discriminating between these two alternative views. Thus, in the subsequent empirical analysis, both alternatives are developed.

Table 2. Unit root tests for the time series

	Annual data	Monthly data
Bankruptcies (b)	-1.087	-1.307
Bank lending (l)	0.038	0.308
Gross Domestic Product (y)	-0.912	..
Industrial production (ip)	-0.243	-0.507
Terms of trade (tt)	-1.904	-3.226
Real exchange rate (fx)	-2.218	-3.038
Real interest rate (rm)	-3.126	-4.639
Consumer prices	-0.500	0.371
Money supply (m1)	-0.558	-0.559
Money supply (m2)	-1.389	-1.720
Stock prices (sx-dp)	-2.847	-2.165
Real wages (w)	0.262	-0.590
Government expenditure/GDP	-2.828	-4.860
Stock exchange transactions	-0.232	-1.559
Business failure rate (b-f)	-3.079	..
Bankrupt firms' debt (db-y)	-1.629	..
Cointegration vector ($\hat{u}1$)	-3.380	..
Cointegration vector ($\hat{u}2$)	-4.099	..
Cointegration vector ($\hat{u}3$)	-4.139	..
Critical values, 5 %	-2.902	-2.865
Critical values, 1 %	-3.524	-3.440

Results are derived for the Augmented Dickey-Fuller test. The model includes a constant term and one (with monthly data four) lagged difference terms. The estimation period is 1925-1994 (1922M5-1994M12).

3.2 The analysis of bankruptcies

The model which we use for bankruptcies in this study is quite similar to earlier bankruptcy equations (see e.g. Altman 1983 and Laitinen 1990). This comes as no surprise because if we start from a standard firm's profit condition we end up with a model which depends on aggregate demand and certain cost variables. To derive the behavioural equation for bankruptcies we may use the following expression for a firm's net wealth (in real terms) as a point of departure:

$$AN_t = (1 + r_t)AN_{t-1} + \pi_t + \tau_t, \quad (1)$$

where AN denotes the firm's net wealth. π stands for profits which are determined by $pq - C(q)$ where p denotes the output price, q output and $C(q)$ production costs. Finally, τ denotes (net) capital gains.

Clearly, AN_t can be negative (and the firm may face bankruptcy) if π and/or $\tau < 0$. More precisely, a negative value of AN_t may actualize if the previous period's debts are large, output prices low, output demand low, production costs high, and capital gains negative. The effect of interest rates r on AN_t is basically ambiguous, but assuming that they have a negative effect on profits, a negative wealth effect also arises.

In a small open economy setting, one may measure p with the real exchange rate fx (and/or with the terms of trade tt). Output demand may be proxied by the Gross Domestic Product y and capital gains by stock prices sx .³ Firms' net wealth creates some measurement problems but the indebtedness ratio (debts/GDP) $1 - y$ may serve for this purpose.

We could then postulate the following relationship between bankruptcies b and possible explanatory variables:

$$b = b(1-y, y, r, fx \text{ \& \ } tt, sx) \quad (2)$$

(+)
(-)
(+)
(-)
(-)

One essential ingredient should still added to this model. That is the persistence bankruptcies. Bankruptcies today cause bankruptcies tomorrow for various reasons. First, other firms suffer credit losses. Second, some bankrupt firms continue operations with much lower operating costs creating an unhealthy competitive environment. Finally, bankruptcies change the operating procedures of other firms and banks for instance in terms of trade credit, collateral and so causing additional liquidity problems. This all implies that bankruptcies (or the business failure rate) depend on the previous periods' bankruptcies.

Here, we face the difficult problem of choosing the reference variable for the number of bankruptcies. It is not at all clear whether we should relate the number of bankruptcies to the number of firms (i.e. to consider the business failure rate) or to some other scale variable. The choice is even more difficult because the number-of-firms variable is quite deficient (the

³ In an open economy setting, negative capital gains may also arise because exchange rate movements. I.e., depreciation of the domestic currency may increase the amount of foreign debt expressed in domestic currency. That is by the way exactly what happened in Finland in 1991-1992. Thus, the effect (real) exchange rate on AN_t is in principal ambiguous.

definition of a "firm" has changed considerably over time). Moreover, cointegration analysis does not give a clear-cut answer to the question of whether the business failure rate is stationary or not.

For these reasons we use some alternative definitions for the dependent bankruptcy variable. The estimating equation is derived from the firm's net wealth expression (1) using in the first place a simple partial adjustment mechanism as a point of departure. The individual variables are introduced into the model so that they are (at least approximately) stationary. Thus, the equation takes the following form:

$$(b-z) = \alpha_0 + \alpha_1(b-z)_{-1} + \alpha_2(1-y) + \alpha_3\Delta y + \alpha_4rm + \alpha_5\Delta sx + \alpha_6fx + \alpha_7gs + \varepsilon, \quad (3)$$

where b denotes bankruptcies and z possible reference variables, $z=n$ indicates population and $z=f$ the number of firms. ε is the error term. The other right-hand side variables have the following definition: $(1-y)$ is the indebtedness rate, Δy is the rate of change in GDP, rm is the real interest rate (government bond yield in real terms), Δsx is the rate of change in stock prices (deflated by consumption prices), fx the real exchange rate index and gs the central-government-expenditure share of GDP. The latter variable is introduced to take the Second World War into account. During the war years, the value of gs was close to 0.5 while in normal years the value has been around 0.1.⁴

The model is estimated with annual Finnish data covering the period 1923–1994. The corresponding OLS and Instrumental Variable estimates are presented in Table 3. In addition to aggregate figures, the table also indicates some estimates for sectoral equation although the data in this respect is quite deficient. In addition, a similar specification is estimated using an error-correction model. To obtain the error-correction term we estimated some alternative long-run (co-integration) equations (see Table 4). The following set of variables was used in these co-integration equations: equation (1): $\{f, y, l, gs\}$, equation (2): $\{n, y, l, gs\}$ and equation (3): $\{n, y, l, w, gs\}$.

The results from the partial adjustment specification and from the error-correction model are qualitatively almost identical. The only difference concerns the long-run properties of these models, which by definition are

⁴ One question which naturally arises here concerns the size distribution of bankrupt firms. Does the increased number of bankruptcies necessarily imply that a disproportionately large number of small firms go bankrupt. This is a difficult question and we cannot answer it because we have not enough data. We have however data on the debt of bankrupt firms (see Figure 3). The time series of real debt db and the number of bankruptcies behave quite similarly, except for the war years. This close correspondence can be interpreted as evidence of the relative invariance of the size distribution of bankrupt firms.

different. Thus, in the case of partial adjustment specification all right-hand side variables also have a long-run effect on bankruptcies while the error correction model says that in the long run the number of bankruptcies is determined by the number of firms (or the scale of the economy), output, debt and labour costs. These variables could also be interpreted as the indebtedness ratio and the functional distribution of income. In the error-correction models the coefficient of the lagged error-correction term (co-integrating vector) is clearly significant, which suggests that the specification is indeed warranted (see Kremers et al 1992). The estimated co-integrating coefficients (see Table 4) also suggest that the specification makes sense. The coefficients of error-correction terms range from -0.25 to -0.48 . Thus, one could argue that a disequilibrium in terms of bankruptcies takes more than two years (but probably no more than four years) to vanish.

Clearly, increasing indebtedness increases bankruptcies. This is well in accordance with a priori theorizing and it is well in accordance with the corresponding Figure 2. In the same way, the overall economic situation, measured by GDP, affects business failures. The effect is not very strong but it appears to be quite systematic in terms of different estimating specifications and estimators. The relationship between OLS and IV estimators indicates that there is indeed some simultaneity between b and Δy . Thus, a fall in output tends to increase bankruptcies, but an increase in bankruptcies tends also to decrease output. It is interesting to note that besides GDP, the real exchange rate index also enters the equation. This variable tells that foreign export markets are very important to Finnish firms. They are always important because the domestic markets are so small. In the case of a recession, this importance may become even more crucial, and from this point of view the level of competitiveness is an essential variable.

The real interest rate effect is also positive. The corresponding coefficient is relatively large and very significant. The economic interpretation is rather straightforward: higher real interest rates make debt costs much higher and if this is not compensated by an increased cash flow, firms face financial problems. Higher real interest rates also reflect tighter money markets, and under such conditions firms may not be able to obtain additional liquidity from the banking sector.⁵

⁵ We made an experiment to control the liquidity effect by introducing the rate of change in narrow money, $\Delta M1$, into the estimating equation. As one could expect, the coefficient of this variable turned out to be negative (increased liquidity decreases bankruptcies) although the coefficient could not be estimated very precisely (the t-ratio remained at the level of one). Also the terms of trade variable tt was used as an additional regressor. Its coefficient could not, however, be estimated very precisely and, therefore, it was chopped from the final specifications.

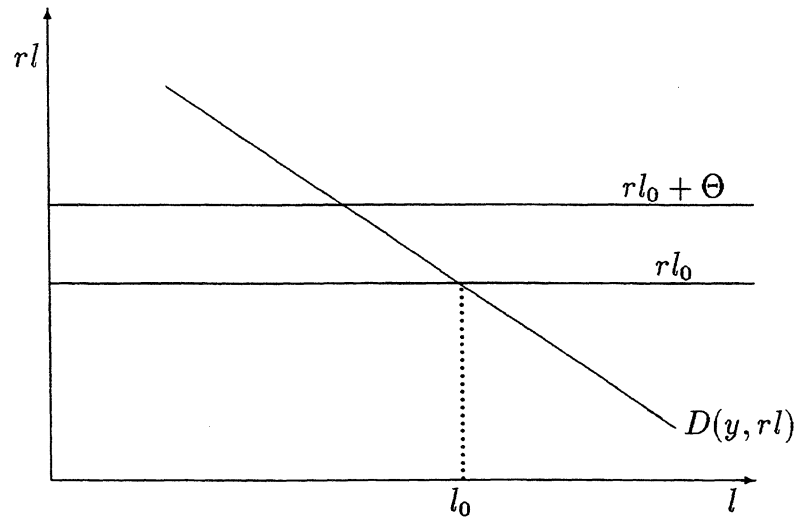
The role real interest rates can also be explained by referring to the role of inflation. Altman (1983) has proposed that increasing inflation reduces competition between firms and shelters inefficiency. It has been said that with high inflation it is easier to raise prices and profits, which lowers the efficiency of the market in a sense by keeping bad products in the market too long.

It is also worth mentioning that the rate of change in stock prices is negatively related to bankruptcies. There is an obvious causal explanation for this finding: an increase in stock prices (as well as in other wealth prices) increases both the value of the firm and the corresponding collateral values, making it easier to handle the liquidity situation.

3.3 Modelling credit expansion

Credit expansion obviously depends on both the supply and demand determinants. Unfortunately, it is not easy to derive a meaningful reduced form equation for the amount of credit. This is generally true but especially so far Finland. The domestic bank loan market was regulated for a period of sixty years (from the mid-1930s to the mid-1980s). The basic form of regulation concerned the banks' average lending rate. Because of this regulation, the supply of and demand for loans were generally not equal. It is generally assumed that the bank loan market was characterized by excess demand. Although we cannot demonstrate that this assumption is true, we should keep it in mind in the subsequent derivation of the credit expansion estimating equation. Thus, the bank loan market is assumed to function in way which is illustrated in Figure 1.

Figure 1. The bank loan market



The demand for bank loans (in real terms) is assumed to depend positively on the scale variable (here, GDP) and negatively on the real interest (lending) rate (rl). Thus, $l^d = D(y, rl)$. The interest rate is exogenously set at some level rl_0 . Banks would not, however, expand their lending to l_0 because that would lower their profits. Instead, they would lend less: the more the regulated interest rate deviates from the equilibrium rate the larger the rationing effect. Rationing could be thought as an exercise in which banks set a rationing premium, say Θ , on the interest rate. In practice, this premium shows up in different non-price rationing terms, as in the downpayment ratio, required length of customer relationship and the collateral requirements. The premium is not constant but depends on the determinants of the supply of bank loans. We may assume that supply depends positively on the interest rate (or, in fact, on the interest rate margin), the stock of deposits and the expected credit losses which, in turn, may be measured by bankruptcies. If the supply of loans is also written in real terms, we may end up with a specification where real loan supply depends on in addition to interest rate(s) and bankruptcies, the real amount of deposits and (negatively) the rate of inflation. The inflation

effect comes via the eroding effect that it has on the real values of both bank deposits and loans.⁶

The rationing premium, Θ would thus depend on the exogenous variables in the following way:

$$\Theta = \Theta(b_{-1}, rl - rd, m_{-1}, \Delta p), \quad (4)$$

(+) (-) (-) (+)

where $(rl - rd)$ denotes the interest margin (for banks) and m_2 the (real) money supply. The latter variable is introduced here as a proxy for bank deposits. The bankruptcy variable appears here with a time lag. Obviously, the existence of a time lag is more an empirical question and therefore we experiment with both speculations (a model with b_t or with b_{t-1} ; see Table 4). As for the interest rate margin we have some data problems and hence we cannot directly apply this variable. In fact, we have only two interest rate series available: the government bond yield, which represents the market rate (rm), and the central bank's discount rate (rd). Because the lending and deposit rates have been tied to this discount rate the difference between rm and rd might reflect an opportunity cost for banks. The higher $(rm - rd)$ the higher banks' financing expenses and the less advantageous is bank lending relative to money market operations. This, in turn, would show up in higher Θ and in lower credit expansion.⁷ In the empirical specification we also replace rl by either rm or rd . Here, rm is used mainly because we want to use the same variable in the bankruptcy and GDP equations.

Thus, we might derive the following linear estimating equation for credit expansion (rate of change in the real amount of bank credit):

$$\Delta l = \alpha_0 + \alpha_1 \Delta l_{-1} + \alpha_2 \Delta y + \alpha_3 \Delta b_{-1} + \alpha_4 \Delta rm + \alpha_5 \Delta \Delta p + \alpha_6 \Delta m_2 + u, \quad (5)$$

⁶ If the loan supply equation is written in terms of nominal loan supply L , which depends on the current period's nominal variables, deflation by the price level may leave the real loan supply to depend on the price level. If, however, supply also depends on the lagged values of exogenous variables, say on lagged deposits, DEP_{-1} , which are here proxied by $M2_{-1}$, then supply in real terms may also depend on the rate of change in prices.

⁷ Here we ignore that fact that Θ may not be a continuous linear function with respect to the exogenous variables. Obviously, if Θ is not linear, the whole bank loan (or credit expansion) equation is not linear. If the excess demand regime changes to an excess supply regime or vice versa, we should probably try to apply genuine disequilibrium models. See, e.g., Quandt (1988). Unfortunately, the performance of such models has not been very good. All in all, there seems to be no satisfactory way of modelling credit markets which have experienced both credit rationing and deregulation (see, e.g. Basu (1994) for more detailed arguments on this problem). In fact, the existence of equilibrium credit rationing may also lead to a similar conclusion although for different reasons.

where Δ denotes the first backwards differencing operator. $m2$ denotes the log real money supply in terms of $M2$ which is used here as a proxy for bank deposits.

Estimation results for this equation are presented in Table 5. The equation performs quite well: the parameters even seem to be stable, which is somewhat surprising given the institutional and demand/supply regime shifts which have taken place in the Finnish financial markets. All the individual variables behave well according to theory. Only the bankruptcy variable is somewhat of an exception in a sense that the lagged level, but not the difference, enters the estimating equation. This might result from asymmetries in the adjustment of credit supply: extending credit and reducing credit might not behave in same way and at least the bankruptcy relationship might be different. The important thing, however, is that the coefficient of the bankruptcy variable α_3 is systematically negative and marginally significant. Thus, there is some evidence of a credit crunch. Notice also that the real interest rate variable is systematically significant (presumably merely reflecting demand behaviour): during depression periods real interest rates tend to increase and, together with increased bankruptcies, they may indeed have adverse credit supply effects.

One additional variable, i.e. the terms trade, turned out to be quite an important ingredient in the credit expansion equation. This variable can be seen as a sort of leading indicator of the state of economy and, particularly, of firms' income expectations. It is no surprise that this variable has a strong positive effect on credit expansion.

3.4 Bankruptcies and output

Finally, we also experimented with the modelling of total output (GDP). The purpose of this experiment was to see whether output growth is affected by bankruptcies (i.e. to see whether causality runs only from output growth to bankruptcies).

One can see that output growth is also almost a random walk, even unrelated with the level of per capita output (see, e.g., Table 2). Given this background it is somewhat surprising that bankruptcies can still help in predicting output growth. The same is not true in terms of other financial and non-financial variables. For instance, a univariate regression relationship between output growth and real interest rates turns out to be the following:

$$\Delta y = .052 - .362gs + .022rm - .075rm(-1) + \hat{u}_1 \quad R^2 = .076, DW = 1.484 \quad (6)$$

(5.58) (2.13) (0.33) (1.14)

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By contrast, the corresponding model for b (or, in fact, $b-f$) turns out to be the following:

$$\Delta y = .072 - .431gs - .041(b-f) + .029(b-f)(-1) + \hat{u}_2 \quad R^2 = .349, \text{ DW} = 1.962 \quad (7)$$

(7.36) (3.09) (4.97) (3.40)

These regression relationships suggest that bankruptcies represent an essential ingredient in the transmission mechanism by which different financial and non-financial shocks affect the economy. The shocks may not show up in direct output effects (as is the case with empirical analyses using with Finnish data) but these effects may well come through bankruptcies. Thus, several VAR model studies which have shown that financial variables are rather unimportant in terms of output determination may have given misleading results just because of the omission of this.

4 Conclusions

Bankruptcies have become an important variable in many countries. The development in Finland has been particularly conspicuous. Bankruptcies have been responsible for very large unemployment and output losses. More importantly, bankruptcies have caused enormous credit losses to banks, which in turn have profoundly affected the capital market and which also have placed a heavy burden on government and taxpayers.

This paper has analyzed the macroeconomic determinants of bankruptcies as well as the consequences of business failures for the financial markets. It is no surprise that bankruptcies behave cyclically. Increased demand and competitiveness reduce bankruptcies and vice versa. In the same way, one might expect that bankruptcies depend (negatively) on real interest rates and (positively) on increases in asset prices. A related factor, which we emphasize in this paper, is indebtedness. It can be argued that indebtedness itself constitutes an equilibrium error-correction term. Excessive indebtedness easily causes a wave of bankruptcies when an economy is hit by a recession with a fall in output (and asset prices) and an increase in real interest rates.

Finnish data provide strong evidence for this argument. This is true for both the stylized facts and the results of empirical analyses. Our analyses also show that bankruptcies affect the growth rate for bank loans. Thus, cyclical fluctuations may increase because bankruptcies lead to a credit squeeze, decreased liquidity, higher real rates, lower asset prices and, finally, to additional bankruptcies (as pointed out e.g. in Stiglitz 1992). Although our results are only preliminary they strongly suggest that the role of bankruptcies

deserves much more attention in future analysis of the relationships between financial markets and the macroeconomy.

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Table 3. Estimates for the bankruptcy equation

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
const.	2.966 (5.74)	985 (2.88)	-1.071 (2.37)	-1.126 (2.44)	-1.140 (0.23)	-297 (0.47)	-1.271 (1.31)	-1.808 (1.76)	.131 (0.31)	.338 (3.14)	.527 (0.88)	.929 (2.26)
b(-1)	.796 (17.88)	.797 (17.84)	.741 (14.52)	.741 (14.50)	.779 (12.39)	.781 (12.36)	.544 (7.01)	.544 (6.88)	.753 (9.31)	.913 (21.28)	.718 (11.74)	.817 (15.60)
1-y	.410 (3.22)	.430 (3.32)	.491 (3.77)	.506 (3.82)	.501 (2.28)	.540 (2.41)	1.058 (3.22)	1.199 (3.50)	.867 (1.80)	2.821 (1.82)	.622 (2.62)	.272 (3.03)
Δy	-1.196 (1.72)	-723 (0.82)	-1.143 (1.72)	-828 (1.00)	-5.417 (4.67)	-4.540 (3.14)	-2.810 (1.54)	-.055 (0.02)	-.282 (1.64)	-.868 (2.70)	-2.007 (1.60)	-.784 (0.93)
rm	.856 (2.29)	.819 (2.16)	1.004 (2.74)	.977 (2.64)	.739 (1.18)	.674 (1.06)	3.057 (3.03)	2.837 (2.74)	.921 (1.03)	.495 (0.20)	2.547 (3.72)	.630 (1.40)
Δsx	-.166 (1.56)	-1.675 (1.57)	-1.150 (1.46)	-1.510 (1.47)	-.333 (1.79)	-.335 (1.80)	-.231 (0.84)	-.233 (0.83)	-.246 (0.98)	-.117 (1.56)	-.253 (1.34)	-1.448 (1.15)
fx	-.850 (5.11)	-.867 (5.16)	-.814 (5.21)	-.826 (5.24)	-1.005 (3.63)	-1.039 (3.71)	-1.259 (3.10)	-1.377 (3.30)	-.261 (0.66)	-3.222 (2.77)	-.919 (3.30)	-.562 (2.91)
gs	-.057 (5.30)	-.056 (5.11)	-.059 (5.63)	-.058 (5.54)	-.037 (2.10)	-.033 (1.95)	-.175 (5.40)	-.168 (5.07)	-1.088 (0.47)	-6.698 (1.03)	-.066 (3.64)	-.044 (3.50)
R2	0.963	0.963	0.960	0.960	0.908	0.907	0.909	0.905	0.793	0.940	0.934	0.920
SEE	0.199	0.199	0.192	0.192	0.333	0.334	0.523	0.533	0.487	0.134	0.360	0.240
DW	2.063	2.016	2.049	2.016	1.556	1.484	1.780	1.700	1.190	1.241	2.096	1.962
LM1	0.392	..	0.358	..	2.206	..	1.055	..	4.393	3.148	0.513	0.069
JB	4.477	4.618	2.328	2.308	2.414	4.057	0.424	0.455	629	472	253	0.269
Chow	4.074	..	3.345	..	1.657	..	3.645	..	9.033	5.483	1.910	3.799
dep.var. estimator	b ols	b iv	b-η ols	b-η iv	b-f ols	b-f iv	db-y ols	db-y iv	BA ols	BI ols	bc ols	br ols

dep. var. denotes the definition of the dependent variable. b denotes bankruptcies, l banks' total credit supply, y real GDP, m real yield on government bonds, sx the Unitas stock index deflated by consumption prices, fx the real exchange rate, gs the share of central government expenditure of GDP, w the real wage rate and M1 the nominal money stock. (b-η) indicates that the number of bankruptcies is divided by population (η). (b-f) that it is divided by the number of firms (f) and (bd-y) that the dependent variable is the debt of bankrupt firms in relation to GDP. In the case of (b-η) specification, also all other relevant variables are divided by population. ba, bi, bc and br denote sectoral bankruptcy variables for agriculture, industry, commerce and other branches, respectively. Due to zero observations, the first two sectoral equations are expressed in levels (not in logs). All variables, except rm and fx are expressed in logs. (-1) indicates that the variable is lagged by one year. (½), in turn, indicates a half-year lag. Numbers in parentheses are unadjusted t-ratios (heteroscedasticity/autocorrelation adjusted t-ratios are so close to these unadjusted ratios that they are not reported). LM is the Godfrey autocorrelation test statistic in the presence of a lagged dependent variable, JB is the Jarque-Bera test for residual normality and Chow is the Chow stability test statistic for the period 1945. Under the null hypothesis, the distribution of LM1 is standard normal, the distribution of JB is chi square with two degrees of freedom while the distribution of Chow is (approx.) F(0.54). OLS denotes the ordinary least squares estimator and iv the instrumental variable estimator. With this estimator, the list of instruments includes lagged Δy, the discount rate (rd), the terms of trade tt, fx, the growth rate of industrial production (ip) and money supply (m1).

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Table 4. Error-correction model estimates for bankruptcies

$$(1) \quad b = 1.556 + .546f - 1.564y + 1.698l - .131gs + \hat{u}1$$

(2.03) (1.84) (4.77) (14.76) (7.78)

R2 = 0.862, SEE = 0.370, DW = 0.576, ADF1 = 3.31.

$$\Delta b = -.027 - 2.115\Delta y + 2.451\Delta l - .047\Delta gs - .245\hat{u}1(-1)$$

(0.86) (3.29) (7.96) (3.48) (3.85)

R2 = 0.553, SEE = 0.191, DW = 1.628, JB = 2.048, Chow = 0.900.

$$(2) \quad b = -89.643 + .143\eta - 4.098y + 2.327l - .119gs + \hat{u}2$$

(8.43) (8.52) (10.97) (20.84) (9.52)

R2 = 0.930, SEE = 0.264, DW = 0.864, ADF1 = 4.10.

$$\Delta b = -.069 + .172\Delta\eta - 2.645\Delta y + 2.372\Delta l - .042\Delta gs - .348\hat{u}2(-1)$$

(1.12) (2.86) (4.15) (7.32) (3.10) (4.12)

R2 = 0.621, SEE = 0.177, DW = 1.642, JB = 1.615, Chow = 0.676.

$$(3) \quad b = -127.167 + .175\eta - 4.809y + 1.411l - .089gs + 3.103w + \hat{u}3$$

(10.78) (11.10) (13.72) (6.86) (7.94) (5.02)

R2 = 0.949, SEE = 0.229, DW = 1.267, ADF1 = 4.23.

$$\Delta b = -.080 + .203\Delta\eta - 3.219\Delta y + 2.050\Delta l - .037gs + 1.775\Delta w - .479\hat{u}3(-1)$$

(1.87) (3.63) (5.61) (7.74) (2.80) (3.32) (5.05)

R2 = 0.677, SEE = 0.164, DW = 1.695, JB = 0.407, Chow = 0.734.

$$\Delta b = -.067 + .190\Delta\eta - 3.068\Delta y + 1.941\Delta l - .039\Delta gs + 1.446\Delta w + .487\Delta rm$$

(1.57) (3.37) (5.31) (7.17) (3.01) (2.56) (1.59)

$$-.072\Delta sx - .427\hat{u}3(-1)$$

(0.83) (3.98)

R2 = 0.693, SEE = 0.163, DW = 1.708, JB = 0.714, Chow = 0.497.

The first equation is the cointegration equation and the latter equation(s) the respective error corrections model(s). ADF1 denotes the Augmented Dickey-Fuller test statistic for unit root (the 5 per cent critical value is 2.90). Otherwise, notation is the same as in Table 1.

Table 5. Estimates for the credit expansion equation

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
const.	-.015 (0.24)	-.097 (2.08)	-.173 (3.87)	-.094 (2.01)	-.145 (3.66)	-.147 (3.68)	-.149 (3.66)	-.145 (3.59)
$\Delta l(-1)$.176 (1.94)	.404 (4.53)	.406 (4.52)	.377 (4.08)	.380 (4.28)	.364 (3.94)	.403 (4.49)	.398 (4.45)
Δy	.399 (2.46)	.395 (2.76)	.403 (2.81)	.259 (1.41)	.373 (2.62)	.296 (1.61)	.425 (3.00)	.329 (2.11)
$b(-1)$	-.036 (3.92)	-.011 (1.82)	-.011 (1.87)	-.012 (1.98)	-1.011 (2.21)	-1.165 (2.29)	-.518 (1.57)	-.952 (1.78)
rm	.420 (2.70)	-	-	-	-	-	-	-
Δrm	-	-.192 (2.24)	-.192 (2.22)	-.205 (2.30)	-.188 (2.17)	-.192 (2.20)	-.195 (2.19)	-.189 (2.15)
Δp	-.813 (6.05)	-	-	-	-	-	-	-
$\Delta \Delta p$	-	-.698 (7.96)	-.699 (7.98)	-.702 (7.94)	-.692 (7.97)	-.694 (7.97)	-.697 (7.89)	-.719 (8.16)
tt	.274 (7.06)	.136 (3.65)	.128 (3.55)	.144 (3.78)	.130 (3.69)	.135 (3.74)	.123 (3.45)	.129 (3.57)
$\Delta m2(-\frac{1}{2})$.444 (3.23)	.474 (4.72)	.478 (4.74)	.507 (4.84)	.465 (4.76)	.483 (4.76)	.477 (4.64)	.457 (4.62)
rm-rd	-	-	-	-	-	-	-	-
R2	.758	.812	.812	.809	0.816	0.815	0.809	0.811
SEE	.046	.040	.040	.041	0.040	0.040	0.040	0.040
DW	1.406	1.821	1.844	1.757	1.838	1.803	1.825	1.840
LM1	2.753	0.421	0.289	..	0.424	..	0.322	0.290
JB	0.215	0.089	0.056	0.224	0.330	0.647	0.033	0.055
Chow	0.418	1.848	2.000	..	1.794	..	2.076	1.915
dep.var.	Δl	Δl	$\Delta(l-\eta)$	Δl	Δl	Δl	Δl	Δl
b.var	b	b	$b-\eta$	b	$b-f$	$b-f$	$db-y$	b
estimator	ols	ols	ols	iv	ols	iv	ols	iv

b.var denotes the definition of the bankruptcy variable. In column (8), it is not lagged (as it is in other equations). Notation is the same as in Table 1.

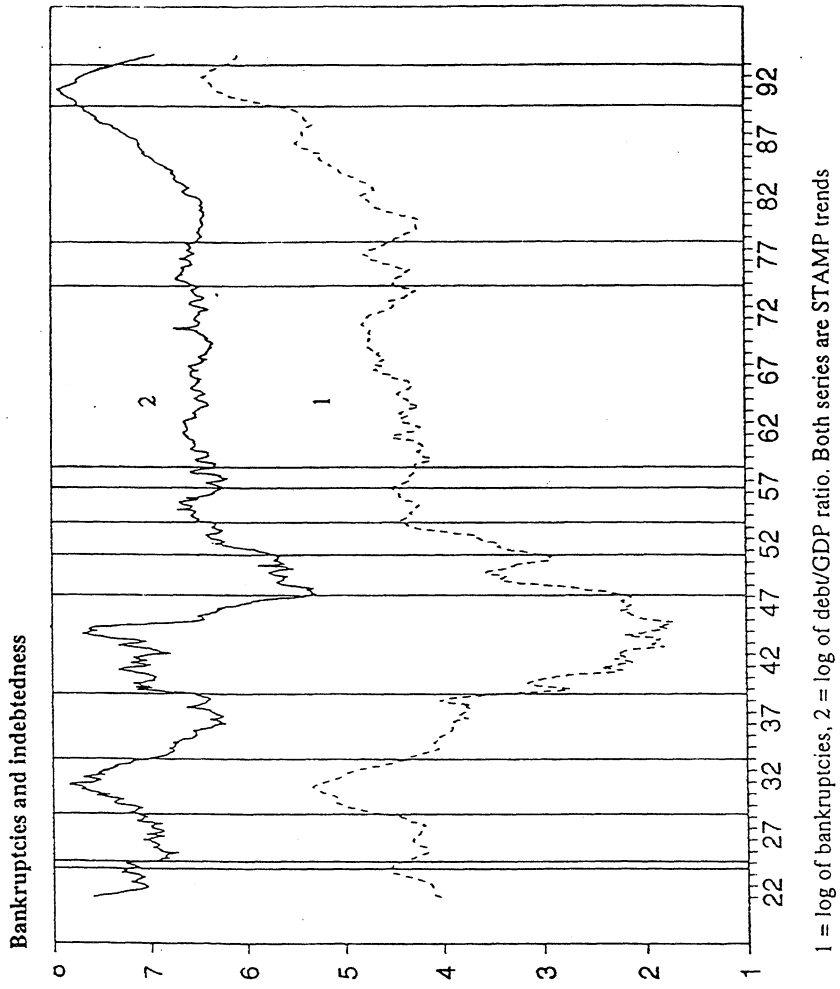
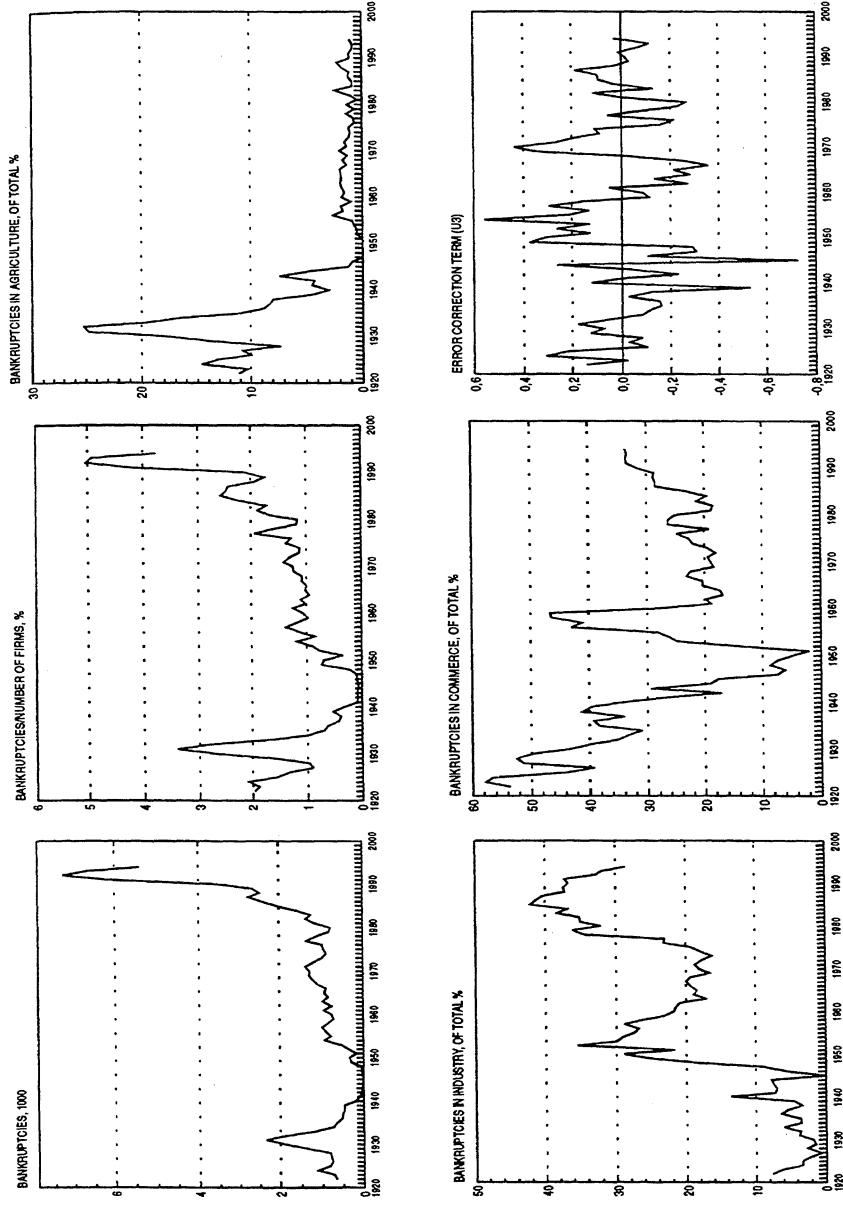


Figure 2.

Figure 3



Testing nonlinear dynamics, long-memory and chaotic behavior with financial and nonfinancial data

Kari Takala • Matti Virén

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Abstract

This paper contains a set of tests for nonlinearities in economic time series. The tests comprise both standard diagnostic tests for revealing nonlinearities and some new developments in modelling nonlinearities. The latter test procedures make use of models in chaos theory, so-called long-memory models and some asymmetric adjustment models. Empirical tests are carried out with Finnish monthly data for twelve macroeconomic time series covering the period 1920–1996. Test results support unambiguously the notion that there are strong nonlinearities in the data. The evidence for chaos, however, is weak or nonexistent. The evidence on long memory (in terms of so-called rescaled range and fractional differencing) is somewhat stronger although not very compelling. Nonlinearities are detected not only in a univariate setting but also in some preliminary investigations dealing with a multivariate case. Certain differences seem to exist between nominal and real variables in nonlinear behaviour.

1 Introduction

Even though economic relationships are thought to be fundamentally nonlinear, most modelling practices start with linear tests and modelling. The obvious reason for this has been the difficulty of choosing from among numerous nonlinear alternatives. Economic theory rarely helps the researcher with anything other than perhaps giving the assumed sign between the two variables. Given the amount of tests and statistical theory based on linear spaces, it has been almost too easy to restrict attention to linear models. However, the poor performance of these models in forecasting e.g. business cycles suggests that maybe things are not so simple.

Apart from some almost self-evident nonlinear functions like the production function or the utility function, nonlinearities have rarely been treated satisfactorily in economics. Although the state of art in nonlinear economics has started to receive more attention, the main problem is that we do not have any clear-cut procedure for approaching these nonlinearities. Up till now, we have had no better advice than just to begin with linear testing and to try to limit the nature of nonlinearities to some well specified class of models.

This paper examines several long Finnish macroeconomic time series. The purpose of the examination is to find out whether there are, in fact, any signs of nonlinearities in these series. We carry out a set of tests analogously to Lee, White and Granger (1993). Most of these tests are applied to univariate models although a multivariate application would obviously be more interesting. When scrutinizing the series we pay special attention to the distinction between nominal and real series. This can be motivated by the fact that nonlinearities are presumably quite different depending on whether nominal or real variables are involved. (For an extensive survey of the literature, see Mullineux and Peng (1993).) Thus, it is of some interest to compare a typical real series, say industrial production, and a nominal series, say stock prices, in this respect.

Most monetary series - like relative prices, changes in price level and money aggregates - display some form of nonlinear behaviour. Prices are often more volatile than the real series, since they play a market-clearing role. Monetary phenomena are based upon valuations that can be adjusted without any significant cost. In the market-clearing situation it is often - but not necessarily always - easier to change the price rather than the quantity. Although prices can easily move in both directions, crises in the market produce excessively large negative (positive) changes. Nominal price rigidities would also have similar effects. Therefore it comes as no surprise that the real exchange rate, stock prices or inflation seem to adjust asymmetrically to shocks.

This affects the volatility of these series. Another major observation concerning the origin of "price shocks" relates to their unstable variance in time. It has been shown that in many cases price changes - e.g. in the stock market - cluster significantly. Forecasting price changes is therefore a harder task for economic agents than forecasting smoother real variables.

Nowadays, a general response to situations of changing volatility (heteroskedasticity) is to use an ARCH model specification. It may well be, however, that the ARCH model is not the proper framework. It is possible that prices possess the so-called long memory property, thus containing permanent components. In particular, the long-memory property shows up in high and

persistent serial correlation over long lags between absolute values of the (linearly filtered) series. It also shows up in so-called rescaled range analysis, which provides estimates of the persistence of time series. Obviously, this kind of long-memory phenomenon is at variance with a linear structure and therefore it may be useful to consider it here as well.

However, in many cases real economic variables also vary in a nonlinear way. Evidence of nonlinear adjustment is provided by e.g. the apparent and persistent tendency for there to be cycles in most important production variables (see, e.g., Pfann and Palm (1993) for details). Whether these nonlinearities in real series arise from the generating process of a series itself or from random shocks is largely an empirical question. So far, no agreement has emerged as to whether real or monetary phenomena are responsible for business cycles. We hope that our estimates of the nonlinearity of these series might shed some light on this issue as well.

One general class of explanations for nonlinearities is chaotic behaviour. Quite recently, there have been numerous theoretical and empirical applications of "chaos theory". In particular, the behaviour of financial variables has been analyzed from this point of view (see, e.g., the books by DeGrauwe et al (1993), Greedy and Martin (1994), Peters (1993) and Vaga (1994)). The analyses have concentrated on testing the existence of chaos; theoretical analyses have mainly been presented as examples of various cases where (deterministic) chaos might arise. Here, we leave the theoretical developments aside and concentrate solely on empirical testing. It is not easy to derive a theoretical model which would be readily applicable to all macroeconomic series which are at our disposal.

Although the analysis mainly deals with univariate models, some preliminary work is done to identify nonlinear relationships between variables. In this context, we do not follow any specific hypothesis concerning the relationships between variables. Rather, we simply make use of a cross-correlation analysis with respect to different moments of our variables. Thus, the analyses represent some sort of first step towards a generalized Granger test for nonlinear relationships. This analysis gives us a general idea of the magnitude and nature of these relationships. An obvious next step is to go back to theory and think about how the findings coincide with different theoretical approaches.

The structure of the paper is very straightforward. First, we look at the data in section 2, then we briefly present the test statistics and illustrate their properties with some simulated data in section 3, and in section 4 we go through the test results for univariate models. The results deal with various diagnostic tests procedures and with a set of analyses of the correlation dimension, rescaled range, time irreversibility, nonlinear adjustment, parameter stability and long memory. In section 5, we scrutinize the results from a cross-correlation analysis between different moments of these series and, finally, in section 6 we present some concluding remarks.

2 The data

The data are monthly Finnish data covering the period 1920M1–1996M9. After data transformations the period 1922M5–1996M9 is covered. Thus, there are 893 observations in each series. The following twelve series are analyzed:

- Industrial production (ip)
- Bankruptcies (bank)
- Terms of trade (tt)
- Real exchange rate index (fx)
- Yield on long-term government bonds (r)
- Consumer price index (cpi)
- Wholesale price index (wpi)
- Banks' total credit supply (credit)
- Narrow money (M1)
- Broad money (M2)
- UNITAS (Helsinki) stock exchange index (sx)
- Turnover in stock exchange (st).

The first four series are real and the subsequent eight nominal. The data are presented in Figure 1. For presentational convenience, most of the series are shown in logs. To get some idea of the timing of changes in these variables the recession periods are marked by shaded areas.

Otherwise, the details of the data are presented in Virén (1992), Autio (1997) and Poutavaara (1996). We merely point out that the ip, bank, credit, M1 and M2 series are seasonally adjusted. This is simply for data reasons – only seasonally adjusted data were available for the prewar period 1920–1938. As for World War II (1939–1945), the data are treated in the same way as for the peace years.

The overall quality of the time series is rather good. Only the interest rate series are somewhat deficient, as can also be seen from Figure 1. The interest rate series suffers from the fact that banks' borrowing and lending rates were administratively fixed from the mid-1930s to the early 1980s and, therefore, bond yields were not genuinely market-based but were, too, indirectly rationed.

3 The test statistics

Data transformations

It is preferable to start testing nonlinearities by estimating the linear model and analyzing the respective residuals. Although economic relationships are more likely to be nonlinear, there is a danger of unnecessary complication if the difference in relation to a linear model is small.

The need for a nonlinear model also depends on the purpose of the model. For short-run forecasting, linear models may suffice, but for long-run forecasts or explanation of apparent nonlinear features more appropriate modelling is needed. Since testing linearity is widely covered in Granger and Teräsvirta (1993), we

discuss only a few basic considerations here. The linearity tests can be divided into two groups, depending on whether a specific nonlinear alternative exists or not. Since our data do not refer to any specific nonlinear formulation, we concentrate on testing against a general nonlinear alternative.

As was mentioned above, we analyze only univariate models. Some kind of basic specification is provided by a linear AR(4), which turned out to be a reasonably good approximation for all time series. In specifying the order of the autoregressive models, we used model selection criteria (SC, HQ, AIC). In order to study the dynamic dependencies between variables, we thought that in the first place it would be best to filter the original series with the linear autoregressive model of the same order. Thus, the residuals are not severely (linearly) autocorrelated. A few exceptions do exist, however, for higher order autocorrelation (for the lag 12, for instance). Anyway, we prefer the parsimonious AR(4) model to more sophisticated specifications.¹ In fact, we also used first log differences for all relevant variables instead of AR(4) residuals. The residuals were qualitatively very similar, suggesting that the AR(4) transformation is not that crucial. For space reasons, the results with the first difference data are not reported here.

Log transformations were applied to most of the series. Thus, only the terms of trade, the real exchange rate and the interest rate series were left untransformed. To assess the validity of this transformation we made use of the Box-Cox transformation. The results of this procedure generally supported the above-mentioned choice. Only in the case of consumer prices and the terms of trade could one not be sure whether or not to make the log transformation.

Dealing with nonlinearities is often easier after the linear dependencies in a time series have already been taken care of. Therefore nonlinear adjustment can be found in a series properly filtered with an autoregressive (linear) model. However, empirical problems do emerge at this point. It often happens, especially in multivariate analysis, that filtering is almost too effective, since all the significant relationships between variables are removed. Therefore unduly long autoregressive lag models that also affect the asymmetry in the series should be avoided.

Standard diagnostic tests

Given the autoregressive model, we compute the following sets of tests: First a set which consists of some basic statistics on the residuals of this linear AR(4) model (see Table 1). These statistics include the coefficients of skewness and kurtosis in addition to the median. We intend to use these data to detect possible asymmetries. The second set of tests consists of traditional specification tests for functional misspecification/nonlinearity. The tests (reported in Table 2) consist of Engle's (1982) ARCH test in terms of lagged squared residuals, Ramsey's (1969) RESET test in terms of higher-order powers of the forecast value of x_t , White's (1980) heteroskedasticity/functional form misspecification test in terms of all squares and

¹ We are well aware that the remaining higher-order autocorrelation might invalidate the subsequent test statistics which are related to the measure of the correlation dimension (see Ramsey, 1990, for details).

cross-products of the original regressors, the Jarque and Bera (1980) test for normality of residuals and Tsay's (1986) nonlinearity test in terms of squared and cross-products of lagged values x_t .² Finally, the Hsieh (1991) third-order moment coefficients are computed. They should detect models which are nonlinear in mean and hybrid models which are nonlinear in both mean and variance but not models which are nonlinear in variance only.

The BDS test for chaotic process

In addition to these "traditional" test statistics we also computed the BDS (Brock, Dechert and Scheinkman) test statistic (see Table 4) and Ramsey's (1990) irreversibility $G_{1,2}$ test. The BDS test is designed to evaluate hidden patterns of systematic forecastable nonstationarity in time series. The test was originally constructed to have high power against deterministic chaos, but it was discovered that it could be used to test other forms of nonlinearities as well (see, e.g., Brock, Hsieh and LeBaron (1991) Frank and Stengos (1988a) and Medio (1992) for details).

The BDS test can also be used as a test for adequacy of a specified forecasting model. This can be accomplished by calculating the BDS test for the standardized forecast errors. The BDS test is then used as a specification test. If no forecastable structure exists among forecast errors, the BDS test should not exceed the critical value. The BDS test has been found to be useful as a general test for detecting forecastable volatility. The key concept here is the correlation dimension, which can be applied in detecting the topological properties of series. For a purely random variable, the correlation dimension increases monotonically with the dimension of the space and the correlation dimension remains small even when the topological dimension of the space (embedding dimension) increases (Brock, Hsieh and LeBaron (1991)).

For a single series x_t for which $x_{t,m}$ is the set of m adjacent values of this time series $x_{t,m} = \{x_t, x_{t+1}, \dots, x_{t+m-1}\}$, also called m -histories of x , the m -correlation integral $C_m(\epsilon)$ is defined as

$$C_m(\epsilon) = \lim_{T \rightarrow \infty} T^{-2} [\text{number of ordered pairs } (x_{t,m}, x_{s,m}), t \neq s, 0 < s, t < N, \\ \text{such that } \|x_{t,m} - x_{s,m}\| < \epsilon],$$

where $T = N - m + 1$ and N is the length of the series. $\|x\|$ denotes the maximum norm (see, e.g. Eckmann and Ruelle (1985)). Now, defining the correlation dimension $d(m)$ as

² For the properties of these test statistics see e.g. Petrucci (1990) and Lee, White and Granger (1993).

$$d(m) = \lim_{\epsilon \rightarrow 0} \frac{\partial \log C_m(\epsilon)}{\partial \log \epsilon}$$

The correlation dimension is based on the fact that, for small ϵ , $C_m(\epsilon) \sim \epsilon^d$. In the case of truly chaotic series, the correlation dimension is independent of m while if the series are random i.i.e. processes $C_m(\epsilon) = C_1(\epsilon)^m$ and hence the (regression) slope of $\log(C)$ on $\log(\epsilon)$ increases monotonically with m .

The purpose of the correlation measure is to describe the complexity of the true series and measure the nonlinear dimension (degrees of freedom) of the process. Tests of chaos concentrate on low-dimensional deterministic chaos processes, since there is no efficient way to tell the difference between high-dimensional chaos and randomness.

Although the correlation dimension can be calculated and interpreted rather easily, there are some major problems with the estimation of this measure, mainly due to fact that economic data are relatively noisy and there are too few observations available (see Ramsey (1990) and Ramsey, Rothman and Sayers (1991) for more details). It can be shown that when the dimension of the data set is based on this Grossberger-Procaccia measure, the estimate of it is necessarily biased because of the following small sample problem: With a finite data set the value of ϵ cannot be too small because otherwise $C_m(\epsilon)$ will be zero and thus $d(m)$ is not defined. By contrast, with large values of ϵ , $C_m(\epsilon)$ saturates at unity so that the regression of $\log(C_m)$ on $\log(\epsilon)$ is simply zero. Thus, the smaller the number of observations, the larger ϵ has to be, and the more biased the estimate of the dimension will be.

Although theory concerns the properties of $C_m(\epsilon)$ as $\epsilon \rightarrow 0$, the reality is that the range of ϵ used in estimating $d(m)$ is far from zero and inevitably increases away from zero as the embedding dimension is increased. Smaller values of ϵ require substantial increases in sample size in order to determine a linear relationship between $\log(C_m(\epsilon))$ and $\log(\epsilon)$. In fact, the relationship is linear only for a narrow range of values for ϵ . Thus, one should be very careful in evaluating single point estimates of $d(m)$. By scrutinizing the entire path of $d(m)$ with respect to ϵ one may obtain a more reliable estimate of the true dimension. Alternatively, one may use the test procedure suggested by Brock, Hsieh and LeBaron (1991) in calculating the following BDS test statistic:

$$BDS(m, \epsilon) = \sqrt{T}(\hat{C}_m(\epsilon) - [\hat{C}_1(\epsilon)]^m) / \sigma(m, \epsilon),$$

where $\sigma(m, \epsilon)$ is an estimate of the standard deviation. The BDS tests whether $C_m(\epsilon)$ is significantly greater than $C_1(\epsilon)^m$, and when this is the case nonlinearity is present. Under the null hypothesis of x_t following i.i.d., and for fixed m and ϵ , $C_{m,T}(\epsilon) \rightarrow C(\epsilon)^m$, as $T \rightarrow \infty$, and $BDS(m, \epsilon)$ has the standard normal distribution. (Notice, however, that $C_m(\epsilon) = C(\epsilon)^m$ does not imply i.i.d.) The power of the test will depend critically on the choice of ϵ .

The BDS test statistic is complicated since it depends on the embedding dimension (m) and the chosen distance (ϵ) related to the standard deviation of the data. The selection of m is important in small samples, especially when m is large, since increasing m means that the number of nonoverlapping sequences will

become smaller. And when the sample is less than 500 the asymptotic distribution may be different from the sampling distribution of the BDS statistic. The selection of ε is even more crucial and failure to detect non-normality in calculating the BDS with small ε is a consequence of too few observations. Brock, Hsieh and LeBaron (1991, p. 52) suggest that for 500 or more observations, the embedding dimension m should be smaller or equal to 5, whereas ε should be 0.5–2 times the standard deviation of the data. In the empirical application, some alternative values of the dimension parameter m and the distance parameter ε are used.

The problem with the BDS test is, however, that it does not have a simple interpretation. Nonlinearity based on the BDS test could be a result of chaos or a nonlinear stochastic process. However, the BDS test was originally designed to test whether the data-generating process of a series is deterministic (chaotic) or not (Granger & Teräsvirta (1993), p. 63). Since the BDS test is based on the null hypothesis that the observations (here AR(4) residuals) are i.i.d., rejection merely reveals that this is not the case. The specific form of nonlinearity is therefore an open question.

As for the practical implementation of the test, it is done here by using the residuals of the AR(4) model as inputs. The use of the autoregressive filter is based on the invariance property of chaotic equations shown by Brock (1986). Brock showed that if one carries out a linear transformation of chaotic data, then both the original and the transformed data should have the same correlation dimension and the same Lyapunov exponents.

In order to get some idea of the implications of deterministic chaos we illustrate the case by comparing a truly deterministic chaos series with a random $N(0,1)$ series. A logistic map model which takes the form $x_t = 4*x_{t-1}(1-x_{t-1})$ is used to generate the chaotic series. Both series contain 2000 observations; the initial value of the logistic map series is 0.3.³ The figure on the following pages illustrates the time paths of these two series (only the first 200 observations are graphed), the respective autocorrelations for 60 lags, two dimensional plots in terms of the current and lagged value of the variable, the correlation dimension estimates with an embedding dimension 2–5 and the BDS test statistics with the embedding dimension 2 over the ε values 0.5–3.0.

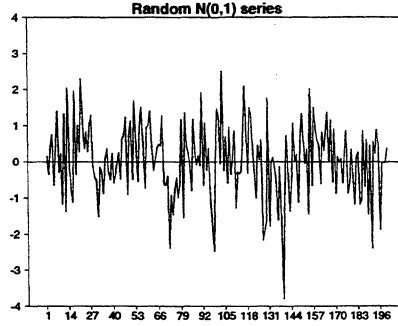
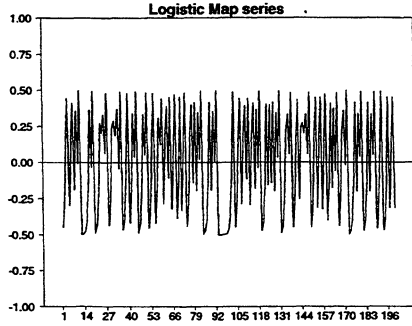
It may be worth mentioning that all tests of chaos depend on the sampling procedure (time aggregation). Thus, if the high frequency chaotic data is measured only infrequently (for instance, daily observations are recorded only monthly), the data appear to be just random data. This property is illustrated by Figure 11 in the end of the paper (see also Table 4 for the BDS statistics in this case).

³ The first value of the series is .300. The series are very sensitive with respect to this initial value. If the initial value is changed to .30001, the new series diverges from the original series after 14 observations and never converges. In addition to the logistic map specification we used the Henon map (which is equally often used a benchmark example). Here, the Henon map takes the following parametrization: $x_{t+1} = 1 + y_t + 1.4*x_t^2$ and $y_{t+1} = .3x_t$ with $x(0) = .1$ and $y(0) = .1$. Needless to say, these series also depend very much on the initial values. Both the logistic map and the Henon map are obviously very simple illustrations of chaotic behaviour and they should not be considered as representative models.

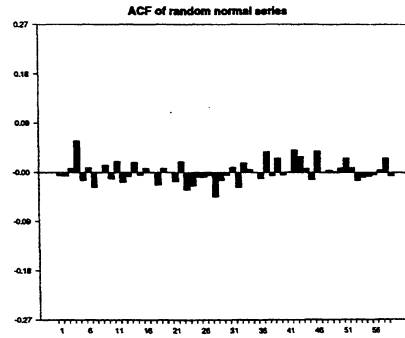
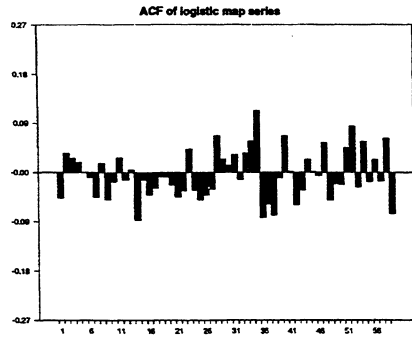
Figure 1. Comparison of logistic map and random series

Logistic map series (adjusted with mean)

Random $N(0,1)$ series



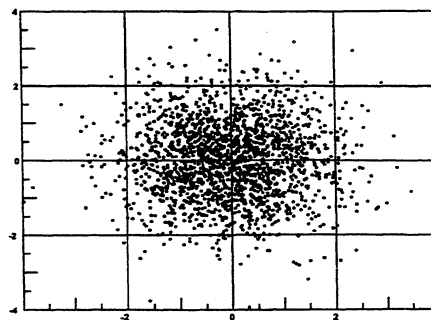
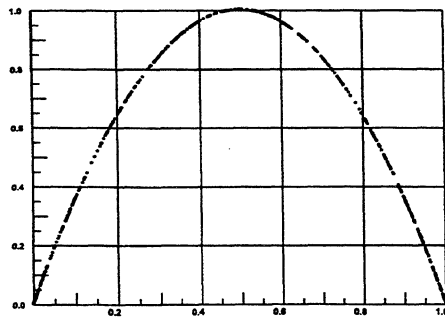
Autocorrelations



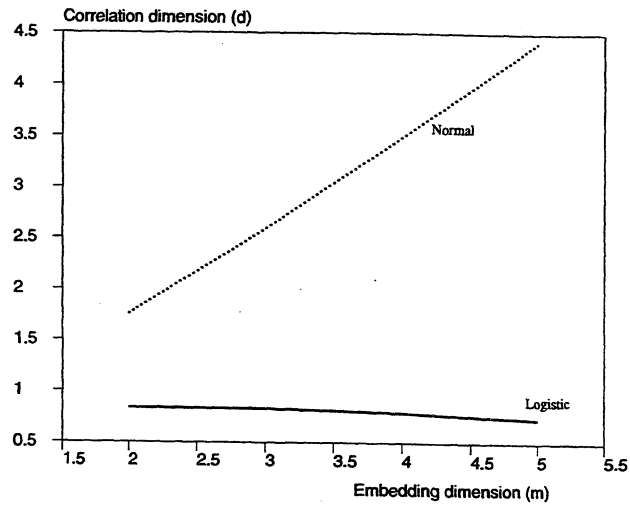
Two-dimensional plots

Logistic map series

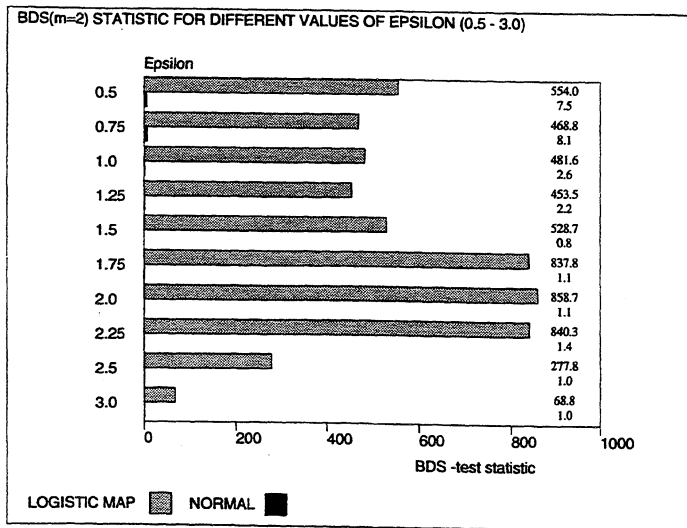
Random $N(0,1)$ series



Correlation dimensions of logistic map and random normal processes

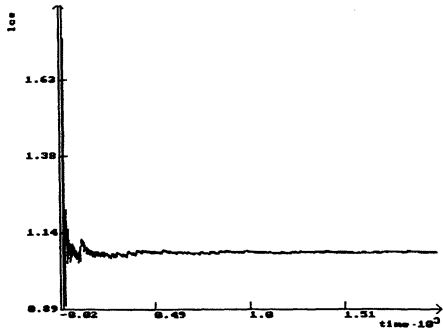


BDS(2) statistics for $\epsilon = 0.5-3.0$

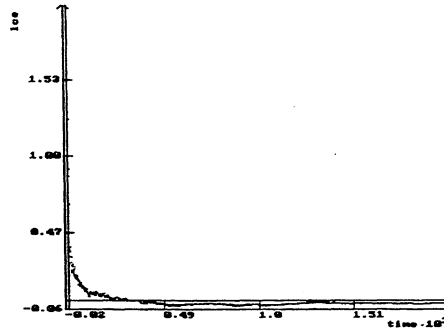


Lyapunov exponents

Logistic map, $L_1 = 1.07$

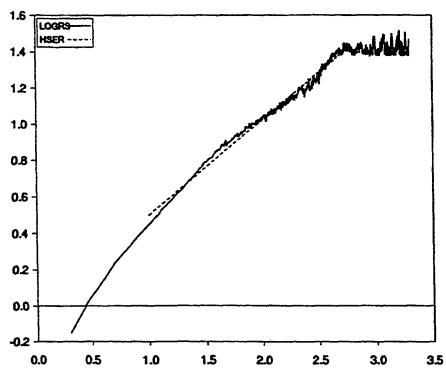


Random $N(0,1)$, $L_1 = -0.02$

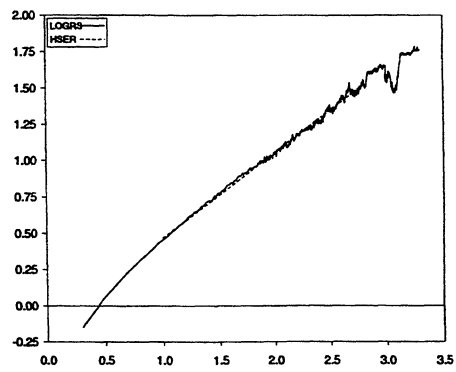


Hurst exponents

Logistic map, $H = .426 (.520/.024)$



Random $N(0,1)$, $H = .588$



The purpose of Figure 1 is to show that the time series and the autocorrelations are quite similar. In fact, one might at first glance consider the logistic map series to be random walk series. The dimension plots show, however, that there is a fundamental difference between these two series. The random $N(0,1)$ series is spread quite evenly over the plane while the logistic map series does not fill enough space at a sufficiently high embedding dimension, which is a generic property of chaotic processes. The clustering of two-dimensional plots also shows up in the dimension estimates (and in the BDS test statistics). The estimate for the logistic map series is about one irrespective of the embedding dimension (it can be shown that the correlation dimension for the logistic map is 1.00 ± 0.02 , see, e.g., Hsieh (1991)). Finally, the BDS test statistics clearly discriminate these two series. Thus, the statistic for random normal series typically fails to exceed the critical value while the test statistic for the logistic map exceeds the critical value by many hundreds.

Lyapunov exponents

The Lyapunov exponents measure the average stability properties of the system on the attractor. Frequently the presence of at least one positive Lyapunov exponent is taken to be the definition of chaos. For a fixed point attractor, the Lyapunov exponents are the absolute numbers of the eigenvalues of the Jacobian matrix evaluated at the fixed point. Thus, the Lyapunov exponents can be considered as generalizations of eigenvalues (see, e.g. Medio (1992) and Frank and Stengos (1988a) for further details).

To define the Lyapunov exponents consider the following N th order dynamic system:

$$\frac{dx}{dt} = F(x),$$

where x is a vector with N components. Consider a trajectory $x^*(t)$ that satisfies this equation and an arbitrarily small positive initial displacement from the start of $x^*(t)$ denoted by $D(t)$. Now, it can be shown under fairly general conditions that, for given $D(0)$, the following limit exists:

$$L_i = \lim_{t \rightarrow \infty} (t^{-1} \ln |D_i(t)|), \quad t = 1, 2, \dots, N.$$

Notice that the Lyapunov exponents are not local properties as one might think. Thus, the values of L_i are independent of the choice of $D(0)$. In fact, one may interpret the exponent(s) to measure the average rate of separation over the entire strange attractor.

A positive Lyapunov exponent measures how rapidly nearby points diverge from each other. A negative Lyapunov exponent, in turn, measures how long it takes for a system to reestablish itself after it has been perturbed. Basically, this is the reason why the Lyapunov exponents offer a way to classify attractors.

The problem is that it is not easy to estimate Lyapunov exponents from experimental data. Wolf et al (1985) have developed a FORTRAN program which estimates the largest exponent L_1 from these kinds of data but it has been shown (see, e.g., Brock (1986) and Brock and Sayers (1988)) that the estimates are very sensitive with respect to the nuisance parameters used in the context of the program. Thus, for instance, large positive Lyapunov estimates may be obtained for pure noise data. Our own experience points in the same direction. Therefore we are reluctant to use the Wolf et al (1985) estimates to characterize our real data.

Quite recently, McCafferty et al (1991) and Dechert and Genacay (1993) have proposed an alternative algorithm using the so-called multilayer feedforward networks, which appear to have superior properties with respect to the Wolf et al (1985) algorithm. This will allow us to rescrutinize the values of the Lyapunov exponents in a more affirmative way. Although we do not go through the analysis of Lyapunov exponents with the real data, we may refer to Figure 1 in the text where the largest Lyapunov exponent is presented for random normal and logistic map time series.⁴ In the case of logistic map series, the exponent is large and positive while with random noise series the exponent converges to a (small) negative value.

The ARFIMA model estimates

Long-term memory often shows up in the form of nonperiodic cycles. This has lead to development of stochastic models that exhibit dependence even over very long spans, such as the fractionally-integrated time series models. The models have autocorrelation functions which decay at much slower rates than those of weakly dependent (mixing) processes. This example, the data generating process of X_t could be the following

$$(1-L)^d X_t = e_t$$

where L is the lag operator and e the white noise term. d can be noninteger which gives a "fractionally differenced" (or order d) time series. Now, if X_t is stationary and invertible for $d \in (-\frac{1}{2}, \frac{1}{2})$ and exhibits a unique kind of dependence that is positive or negative depending on whether d is positive or negative. If d is positive autocorrelation decay very slowly, indeed so slowly that their sum diverges to infinity. If $d < 0$, they sum collapses to zero, instead (see e.g. Lo (1991) for demonstration of the effects of fractional differencing on the autocorrelation function).

Estimating an ARFIMA model is not a easy task. Computational problems are not of second order importance. In addition, one has to find out the proper

⁴ Lyapunov exponents have been estimated in several empirical studies; see, e.g., Frank and Stengos (1988c), Frank et al (1988) and Peters (1993). The results have been somewhat mixed, partly depending on the algorithm (thus, for instance, Frank and Stengos (1988) do not find support for the existence of chaos while Peters's results point in the opposite direction). There is, however, a lot of ambiguity concerning the results because of convergence problems and computational sensitivity.

specification for the estimating model. Here this problem boils down in determining the lag structure for the AR part of the model.

The Hurst exponent (rescaled range analysis)

The Hurst exponent is a new measure which can classify time series in terms of persistence (or "antipersistence"), stability of the data-generating mechanism and the importance of outlier-type observations. Thus, it can distinguish between a random series and a non-random series, even if the random series is non-Gaussian. The Hurst exponent was first applied to natural systems (first, in analyzing water reservoir control within the Nile River Dam project in early 1900) but recently there have been numerous applications to financial data (see, e.g., DeGrauwe, et al (1993) and Peters (1993)).

Computing the Hurst exponent (H) and the related V test statistic requires the following steps:

$$Q_n \equiv \frac{1}{\hat{\delta}_n(q)} \left[\text{Max}_{1 \leq k \leq n} \sum_{j=1}^k (X_j - \bar{X}_n) - \text{Min}_{1 \leq k \leq n} \sum_{j=1}^k (X_j - \bar{X}_n) \right],$$

where

$$\hat{\delta}_n^2(q) \equiv \frac{1}{n} \sum_{j=1}^n (X_j - \bar{X}_n)^2 + \frac{2}{n} \sum_{j=1}^q w_j(q) \left\{ \sum_{i=j+1}^n (X_i - \bar{X}_n)(X_{i-j} - \bar{X}_n) \right\}$$

where \bar{X} is the sample mean $(1/n) \sum_j X_j$. $\hat{\delta}_n^2(q)$ is simply the square root of consistent estimator of the partial sum's variance. If there is no short-run dependence, the variance is simply the variance of the individual terms X_t . In the presence of short-run dependence we have to modify the statistic (following Lo (1991)) and include also the autocovariance terms.

Under the i.i.d null hypothesis, as n increases without bound, the rescaled range Q_n converges in distribution to a well-defined random variable V when property normalized so that

$$\frac{1}{\sqrt{n}} Q_n \Rightarrow V$$

The fractiles of the (non-modified) V statistic are reported in e.g. Lo (1991). As for the Hurst exponent, it can be estimated from the following model

$$Q_n = (\alpha \cdot n)^H$$

where α is the scaling constant.

In detecting long-run dependence, the rescaled range analysis is probably not the most efficient way of doing that. By contrast, estimating the fractional differencing parameter directly (e.g. in the context of an ARFIMA model) would be a better way. Still, the R/S analysis could be useful as a complementary tool in assessing more general features of long-run dependence (for further details, see., e.g., DeGrauwe et al (1993) and Peters (1993)).

According to the statistical mechanics, H should equal 0.5 if the series is a random walk. In other words, the range of cumulative deviations should increase with the square root of time. For many (most?) time series from a natural system, the value of H has turned out to be much higher than 0.5. In surprisingly many cases the value of 0.73 is obtained (see DeGrauwe et al (1993)).

When H is different from 0.5, the observations are no longer independent in the sense that they carry a memory of all preceding events. This memory can be characterized as "long-term memory". Theoretically, it lasts forever. Thus, the current data reflect everything which has happened in the past. Notice that this is something which cannot be taken into account in standard econometrics, where time invariance is assumed.⁵

Now, consider the case where $H < 1/2$ and $H > 1/2$. In the former case, the system is antipersistent or "mean reverting". Thus if the system has been up in the previous period, it is more likely to be down in the next period. By contrast, when $0.5 < H < 1$, the system is persistent or "trend-enforcing". If the series has been down in the last period, then the chances are that it will continue to be down in the next period.

A R/S plot for random $N(0,1)$ and a logistic map series is presented in Figure 1 in the text. Note that the estimated slope (i.e., the Hurst exponent) is 0.59, which is quite close to the theoretical value of 0.5. For finite series, the expected value of H , $E(H)$, is in fact somewhat larger than 0.5. Thus, the value of 0.59 may well fall inside the confidence interval of $E(H)$ ($\text{Var}(H) = 1/n$; see Peters (1994)). The estimated slope of the logistic map series is instead 0.43 (as for the Henon map, an estimate of .38 is obtained; for the logistic map series see Figure 1, see also Peters (1994)), which says (in statistical terms) that this series has no population mean and that the distribution of variance is undefined. Clearly, there is nothing we can forecast with these series.

The Ramsey irreversibility test

The irreversibility test, which has been derived by Ramsey and Rothman (1988) and Rothman (1993), deals with the concept of time reversibility.⁶ Time

⁵ The values of the Hurst exponent can be related to a correlation (C) measured in the following way: $C = 2^{(2H-1)} - 1$. Thus, when $H = 1/2$, $C = 0$, and we are dealing with a random series. Its probability density function may be the normal curve but it does not have to be. By contrast, if H is different from $1/2$, the distribution is not normal.

⁶ A stationary time series $\{x_t\}$ is time reversible if for any positive integer n , and for every $t_1, t_2, \dots, t_n \in z$, where z is the set of integers, the vectors $(x_{t_1}, x_{t_2}, \dots, x_{t_n})$ and $(x_{-t_1}, x_{-t_2}, \dots, x_{-t_n})$ have the same joint probability distributions. A stationary time series which is not time reversible is said to be irreversible. Notice that, by definition, a non-stationary series is time irreversible. See e.g. Tong (1983) for further details.

irreversibility is a concept which is useful in analyzing possible asymmetries (nonlinearities) in economic time series, for instance, in output series. According to the conventional Mitchell–Keynes business cycle hypothesis, cyclical upturns are longer, but less steep, than downturns (see also the "plucking model" of Friedman (1993)) If one traces out the behaviour of cycles in reverse time it can be seen that the symmetric cycle is time reversible and that the asymmetric cycle is time irreversible.

Ramsey and Rothman (1988) propose that the presence of time irreversibility should be checked by estimating a symmetric bicovariance function in terms of x_t . The test statistic which is obtained from this bicovariance function is of the following type:

$$G_{i,j}^k = T^{-1} \sum_{t=1}^T \left[(x_{t1})^i (x_{t-k})^j - (x_t)^i (x_{t-k})^j \right] \quad k=1,2,\dots,K.$$

If the time series is time reversible, $G_{i,j}^k=0$ for all k . As for the choice of exponents, i and j , we assume here that $i=2$ and $j=1$ (here we just follow Ramsey (1990)). In addition, we experiment with the pair $i=3$ and $j=1$. The maximum lag length K is set at 120. To ensure stationarity, we also use here AR(4) residuals instead of the original time series. The significance of the G statistic is tested by computing the confidence limits according to the following formula for the variance of $G_{1,2}^k$:

$$\text{Var}[G_{1,2}^k] = \left(\frac{2}{(T-k)} \right) [\mu_4 \mu_2 - \mu_2^3],$$

where $\mu_2 = E[x_t^2]$ and $\mu_4 = E[x_t^4]$. Assuming that the data are independent and identically distributed $N(0, \sigma^2)$, the right hand side of the above formula can be simplified to $\left(\frac{4}{(T-1)} \right) [\mu_2^3]$. This is clearly a crude approximation because the normality assumption does not hold, nor are the variables uncorrelated. However, it is not at all clear how the variance terms should be computed when x_t is not i.i.d. but follows e.g. some general ARMA(p,q) model (see Ramsey and Rothman (1988) for various experiments). The test statistics and the respective confidence limits are displayed in Figure 8.

A nonlinear adjustment equation

Instead of just computing test statistics for nonlinearity, it would be tempting to estimate a general nonlinear time series model and compare its properties with a linear model. Unfortunately, such a general nonlinear model does not exist nor is there any agreement on a reasonable approximation which could be used to capture the possible nonlinear elements of the data. Still, the situation is not completely hopeless. There are some interesting candidates for a nonlinear

specification. The first which deserves to be mentioned is the threshold model specification introduced by Tong (see e.g. Tong (1983)). Another specification which is clearly worth mentioning is the nonlinear employment (output) equation introduced by Pfann (1992). This (estimating) equation takes the following form:

$$x_t = a_0 + a_1 t + a_2 x_{t-1} + a_3 x_{t-2} + a_4 (x_{t-1} x_{t-2}) + a_5 (x_{t-1}^3 x_{t-2}) + a_6 (x_{t-1} - x_{t-2})^3 + \mu_t,$$

where μ is the random term. According to Pfann (1992) and Pfann and Palm (1993), the parameter of the nonlinear terms can be unambiguously signed in the case of employment equations. Thus, a_4 should be positive (if hiring costs are higher than firing costs, or in general, if the cycle spends more time rising to a peak than time falling to a trough). Moreover, parameter a_5 is expected to be negative if the asymmetry (skewness) of magnitude (i.e. the magnitude of troughs exceeds the magnitude of peaks) is negative and parameter a_6 is also negative if the asymmetry (skewness) of duration is negative (i.e., it takes longer for a series to rise from a trough to a peak than to fall from a peak to a trough).

Although this model may make more sense with (productive) input and output series, we also apply it to all ten (here, in fact, thirteen) Finnish series partly to see whether the real and nominal series can be discriminated on the basis of this equation. The results are reported in Table 9. This table also includes a comparison of this model with a linear alternative.⁷

4 Test results with univariate models

4.1 Results from diagnostic tests

The message of the empirical analyses is quite clear and systematic: the data do not give much support to linear models. Thus, all the test statistics reported in Tables 2 and 3 indicate that at least a linear AR(4) model is in trouble.⁸ According to Table 2, the residuals from the AR(4) model suffer from heteroskedasticity and non-normality. The ARCH(7) statistic is significant for all variables (perhaps excluding the interest rate). Thus, even with real series like industrial output an autoregressive conditional heteroskedasticity effect can be discerned. This is something new. Nobody is surely surprised to find an ARCH effect in stock prices but here a similar result applies to other variables as well.

⁷ Here, we merely replicate the experiments by Pfann (1992). Thus, we take the same detrending procedure (see the second term on the right hand side) and the same lag structure. Obviously, extending the lag length beyond 2 would enormously complicate the model.

⁸ In addition to the test statistics reported in Table 2, we also computed the Keenan (1985) and McLeod-Li (1983) test statistics. Both of these turned out to be highly significant. Thus the marginal significance levels were in all cases well below 5 per cent. The test statistics were also computed for the post-Second World War period. Results were quite similar to those reported in Table 2. Thus the war itself cannot explain why the results lend support to nonlinearities.

Non-normality is clearly a severe problem. It is quite obvious that normality is violated because of outlier observations. Clearly, some observations can be classified as outliers and it might well be that these observations contribute to the rejection of linearity. This can be seen from Figures 2 and 3 which contain the time series and frequency distributions for the AR(4) residuals. In accordance with Table 1, the main problem seems to be excess kurtosis, not so much excess skewness. Although the normality assumption is rejected, the graphs suggest that the distributional problems are not, after all, so severe as the Jarque–Bera normality test statistic suggests.

Unfortunately, there is no obvious remedy for non-normality and outlier observations. One alternative is, of course, to use robust estimators and examine whether the results (e.g., the properties of residuals) change importantly as a result of the change in estimators. In fact, we did do this but it turned out that the results with the least absolute deviations estimator were qualitatively very similar to the OLS results. Another possibility is to reconsider the relevant sampling distributions of the nonlinearity tests statistics in the light of observed behaviour of OLS residuals. Here, we have not yet worked out this alternative.

After these considerations, some comments on the RESET and TSAY nonlinearity test statistics merit note. Both tests do suggest that the (linear) functional form is misspecified for most of the variables. The results are, however, very systematic. Thus, for instance, industrial production and bankruptcies, on the one hand, and narrow money and credit supply, on the other hand, behave in a different way in these tests. Moreover, the test results do not allow us to draw a clear line between real and nominal variables. As far as Hsieh's (1991) third-order moment coefficients are concerned, one can see that with some variables the coefficients are very high. Some of the highest coefficients are, in fact, quite similar to those of the logistic map series! High coefficient values are obtained for the real exchange rate, consumer and wholesale prices, money and – somewhat surprisingly – stock prices. By contrast, the values for industrial production, bankruptcies and the terms of trade are somewhat lower although all of them are not "clean". Thus, nonlinearities do exist and nonlinearities are not only a problem for real variables. Since the third-order moment coefficients are not intended to test models which are nonlinear in variance, one may conclude that the high coefficient values for the nominal series not (only) reflect some ARCH effects but also other sorts of nonlinearities (say GARCH-in-Mean effects or long-memory behaviour).

4.2 Results from analyses of the correlation dimension

Next, we turn to results from the analysis of the correlation dimension. These results are presented as follows: First, the two-dimensional plots of the AR(4) residuals are presented in Figure 5, then the correlation dimension estimates are presented in Figure 6 (Figure 6 consists of two plots showing the correlation integral and the derivative of $C(\epsilon)$ in terms of ϵ ; the respective numerical values are reported in Table 3) and, finally, the BDS test statistics are reported in Table 4.

Unfortunately, the results from these exercises are somewhat different. First, the dimension plots are not consistent with the existence of low-dimensional chaotic behaviour (notice, however, that we just look at things very informally in

two dimensions). Although there are some differences between variables, none of the variables behaves in a chaotic manner. Stock prices may best correspond to a random variable (observations are evenly distributed over the x_t, x_{t-1} plane) while some clustering takes place in consumption and wholesale prices.

As one might expect on the basis of the dimension plots, the estimates of the correlation dimension (the embedding dimension running from 2 to 5) lend very little support to a model of chaotic behaviour. The estimate of $d(m)$ increases almost linearly with the embedding dimension m . Only wholesale prices give an opposite result. The estimate of $d(m)$ remains in the neighbourhood of one even if the embedding dimension is increased to 5. Figure 2 may explain why this result emerges. The behaviour of prices in the 1920s and 1930s was completely different from the rest of the sample period (i.e. the price level was practically stationary during the pre-war period while after the outbreak of the Second World War the rate of inflation turned out to be stationary). If the 1920s and 1930s are dropped from the sample, the correlation dimension estimates behave well in accordance with the other variables.⁹

Somewhat contrary to these results, the BDS statistics turn out to be very high, suggesting that the data-generating mechanism is not linear. The null hypothesis that the series are random i.i.d variates is rejected in all cases with standard significance levels. The same result emerges when ARCH residuals are used instead of OLS residuals. A completely different result emerges, however, when the series are shuffled, i.e. the observations are arranged in a random order. Then the null hypothesis of independent observations is typically not rejected, which suggests that the distributional assumptions are not very critical in terms of the outcome of the BDS statistics. By contrast, the time-series structure is the important aspect which produces the very high values of the BDS statistics.

But how should we interpret this conflicting evidence? Should more stress be given to the correlation dimension estimates or the BDS test statistics. The answer is not easy. Perhaps the best way to summarize this evidence is to conclude that there are definitely some signs of nonlinearity but not necessarily of deterministic chaos.

4.3 Results from the tests for the long memory property

As pointed above, the analyses make use of the ARFIMA model estimates and the rescaled range test statistics. In addition, we carry out more informal tests by scrutinizing the autocorrelation structure of AR(4) residuals in terms of different transformations. This research menu may reflect the fact all of these analytical tools are used in assessing the presence of long-run dependence.

In time series, a long-term memory property is said to be present if absolute values of a stationary variable r_t have significant autocorrelations for long lags, i.e. $\rho(|r_{t-k}|, |r_t|) \neq 0$, when k is large. This property was first noted for speculative price series by Taylor (1986) and thereafter also called the Taylor effect (see

⁹ For the period 1939M9–1996M9 the following set of dimension estimates was obtained: $m = 2$: 1.901 (1.02); $m = 3$: 2.709 (1.30); $m = 4$: 3.617 (1.94) and $m = 5$: 4.226 (1.01). These values are clearly in accordance with the other values in Table 3 and hardly consistent with the existence of deterministic chaos.

Granger and Ding (1993)). In practice, this property implies that the simple random walk model does not hold for stock prices, even if the price changes are serially uncorrelated. This phenomenon also shows up in the rescaled range analysis with the Hurst exponents. High ($H > 0.5$) values of the Hurst exponent imply strong persistence – the fact that the observations (residuals) are independent but they have a memory. Thus, the data are not generated with a random walk but, instead, with a biased random walk or, in other words, with fractional brownian motion.

For instance, if we consider stock price changes, it seems intuitively appealing to observe that they are uncorrelated, but this does not explain anything about the heteroskedasticity found in them. Statistically, stock prices could be martingales with non-constant innovation variance (see e.g. Spanos (1986)). However, from the economic point of view, the problem is to find out whether residual variance from the linear model follows conditional heteroskedasticity (ARCH), a generalized version of it (GARCH), asymmetric power ARCH (A-PARCH as defined in Ding, Granger and Engle (1993)) or some other form of heteroskedasticity appropriate for the particular time series. However, univariate models could be helpful in identification and prediction of the type of heteroskedasticity, but probably insufficient for understanding these processes.¹⁰

Heteroskedasticity in residuals already shows that stronger forms of rational expectations rationality, which imply efficient use of all information, does not hold for higher moments of the process. In fact, expectation errors are not white noise, but rather innovation processes with non-constant variance. The long-memory phenomenon also puts emphasis on the long-term cyclical swings often encountered in economic time series. These cyclical swings could relate to business cycles or even Kutznets and Kontrajev cycles or a tendency to generate serious financial crises like those witnessed in the 1930s and 1980s. However, as Granger and Ding (1993) emphasize, caution should be observed in interpretation, since it is not the series themselves but their absolute values that have the long-memory property.

If the efficient market hypothesis were to hold strictly, the random walk property would imply that r_t is an i.i.d process. In addition, any transformation of r_t , like $|r_t|$ or r_t^2 should also be an i.i.d process (Ding, Granger, Engle (1993), p. 87). The sample autocorrelations of the i.i.d process will have finite variance $1/\sqrt{T}$ and larger correlations for $|r_t|$ will indicate the long-memory property. Ding, Granger and Engle (1993) show that, if $|r_t|^d$ is taken as a yardstick in measuring the strength of autocorrelation for long lags, the long-memory property is strongest around $d = 1$.

In the same way as Ding, Granger and Engle (1993), we found that all variables in our data set showed clear evidence of long memory, and thus the sample autocorrelations for absolute values of residuals were greater than the autocorrelations of squared residuals. This resemblance could indicate that economic time series have characteristics of models not fully described or understood so far.

¹⁰ Granger and Teräsvirta (1993, pp. 51–53) note that a series may have short memory in mean, and long memory in variance, but that the opposite is not so likely, i.e. long memory in mean with short memory in variance. Short memory in mean is often found in stationary series, whereas long memory is present in integrated "level" series.

Series which had $|r_t|$ well above r_t^2 were industrial production, the terms of trade, the real exchange rate and the interest rate. Slightly different were series such as bankruptcies, wholesale prices, money supply (M2) and stock prices, which mostly shared the same characteristics. This could be due to rare, but large discrete changes in these series, such as e.g. the effects of devaluations and strikes. The results from these long-memory tests performed for AR(4)-residuals of our time series are presented in Table 5 below. Graphs of sample autocorrelation functions for the absolute values of the AR(4) residuals are shown in Figure 7.

Among other things, the results indicate that linear filtering with an AR(4) model is not sufficient to remove dependence on the distant past in these series, even though model selection criteria would suggest at most times that the fourth-order autoregressive polynomial should be long enough. Despite the fact that these series have dominant long-run features like unit roots and trends, parsimonious linear models seem unable to account for these. Observations therefore point to the conclusion that trends in economic time series are more likely to be stochastic than deterministic. Hence we come up against nonlinearities again.

The main message is, however, there is significant long-run dependence in all of the real and monetary series. In addition, there seems to be no clear difference between real and monetary variables as regards how fast autocorrelations would die out for long lags. Whether that dependence can be accounted for specific long memory models is analyzed next.

Now, turn to estimation results with ARFIMA models. Estimation is carried out using both original time series and fitting an ARFIMA(4,d,0) model to these series and using the AR(4) residuals and fitting an ARFIMA(1,d,0) model to these series. The ARFIMA/OX program by Doornik is used estimation. Both the Maximum Likelihood (ML) and the Nonlinear Least Squares (NLS) estimators are used here.¹¹

The estimates are presented in Tables 6-8 below. Table 6 contains the NLS estimates for the ARFIMA(4,d,0) model for unfiltered time series (a log transformation is taken for all series except for *tt*, *fx* and *r*). The ARFIMA(1,d,0) estimates are reported in Table 7 (for the AR(4) residuals) for both ML and NLS estimators. Finally, we report in the difference parameter estimates which have been obtained by using the Geweke-Porter-Hudak (1983) approach. The estimated fractional difference parameters correspond here again to the AR(4) filtered series.

All in all, the results lend some support to the long-memory property but the overall evidence is quite moot. In the case of unfiltered data, there are some examples (i.e., *tt* and *r*) in which the fractional difference parameter is significant. With filtered data, the evidence is much weaker (for quite obvious reasons). Turning to the Geweke-Porter-Hudak differencing parameter estimates in Table 8, one may notice that they give a bit more evidence of fractional differencing. Results with money and credit series (in addition to consumption prices) suggest that the differencing parameter may indeed be fractional and thus the series may be long-run dependent. What makes this observation important is the fact the subsequent results with rescaled range analysis point to the same direction.

¹¹There were several computational problems with the Maximum Likelihood estimation, i.e. computational failures and great sensitivity with respect to initial values. The nonlinear least squares alternative performed much better in this respect.

The results for the rescaled range analysis are reported in Table 9 and in Figure 8. Table 8 contains the V statistics for all time series in addition to the estimates of the slope parameters, i.e. the Hurst exponents H . The time series graphs for Q are presented in Figure 8 to illustrate stability of pattern of (possible) short and long dependence.

Looking at Table 8 shows that the Hurst exponent is generally above 0.5. One has, however, to take into account the fact that in finite samples the expected value of H is well above 0.5 (see Peters (1994) for simulated values of $E(H)$). Thus, if one scrutinizes the values of the test statistic V , it comes out that they are generally not statistically significant. In other words, there is only weak evidence of long memory. The series which show long-memory properties are in fact quite the same which showed similar properties in the case of Geweke-Porter-Hudak estimator. Thus, money and credit series together with prices behave in a manner which is consistent with long memory. For all other series, the evidence is less compelling – and for "real" series there is no evidence of long memory.¹²

An interesting question is how long is the long-memory phenomenon. Is there a certain time span – say one year – during which observations are not independent. One could, for instance, argue that for various reasons (see, e.g., Peters (1993)) the stock market is not efficient in the short run but efficient in the long run. In other words, we have some cycles which just reflect these inefficiencies (or, more generally market imperfections). This might show up in a change of the R/S slope. In fact, this kind of reasoning seems to apply to the Finnish stock price series. There is quite a clear change in the slope after 300 data points corresponding to a cycle of about 12 years. In the short run, stock prices seem to follow the random walk model (H equals to .48) while in the long run stock prices can be characterized as independent or even antipersistent. Thus, an estimate of .18 is obtained for the data points exceeding 300 (see Figure 7). Clearly, this finding is consistent with the results obtained by Peters (1993) in terms of behavioral changes but equally clearly the finding is at variance with the long-memory property of stock prices.

4.4 Results from the time irreversibility analysis

A similar result emerges with Ramsey's (1990) irreversibility test statistics reported in Figure 8.1. Although, the confidence limits are only indicative, some signs of nonlinearities can be discerned with all series. Somewhat surprisingly, stock prices do not seem to be the most striking example of this sort of nonlinearity. Thus, for instance, the test results for industrial production tell more about nonlinearities than the results for the stock index (see Figure 8.2). Bankruptcies and banks' total credit supply seem to be more obvious candidates. Perhaps this is something which is in accordance with the observed nature of

¹²The results changed only marginally when the original rescaled range statistic was replaced by the modified rescaled range statistic. As with Lo (1991), the values of V did generally decrease although the changes was on average quite small. More important change took place when the $AR(4)$ residuals were replaced by first (log) differences. The evidence on long memory was in this case even weaker than with the $AR(4)$ residuals.

indebtedness and the relationship between indebtedness, credit supply and bankruptcies (see, for instance, Stiglitz and Weiss (1981) and Bernanke (1983)).

Recently, Rothman (1993) has shown that the Ramsey–Rothman irreversibility test is relatively powerful against the threshold model. Thus, our findings could also be interpreted from this point of view. In other words, there are nonlinearities but not of deterministic chaos type but rather resulting from nonlinear model structure or parameter instability. In the subsequent sections, we consider these alternatives.

4.5 Estimates of adjustment equations

Can anything else be said about the nature of nonlinearities? Tables 2 and 6 suggest that this is the case.¹³ Table 1 indicates that the real series and the nominal series behave in a very different way. The nominal series do not show any signs of negative skewness. Moreover, the nonlinear adjustment equations (reported in Table 6) behave very badly, for instance, in terms of stationarity.¹⁴ It is particularly interesting to compare the behaviour of industrial production and stock prices. Industrial output is characterized by clear negative skewness (in magnitude) while there is no apparent skewness in stock prices. With industrial production, positive residuals are much smaller and obviously more numerous than negative residuals. Intuitively, this makes sense since capacity constraints limit increasing production while a decrease in orders or bankruptcies may lower production more rapidly. With stock prices, there is no difference between positive and negative residuals. Thus, adjustment of stock prices does not contain significant asymmetries. See Figure 9 for details; notice that positive and (absolute values of) negative AR(4) residuals are presented here in an ascending order.

4.6 Results from stability analysis

The adjustment properties could, of course, be scrutinized in a straightforward way by looking at the parameter stability over depressions and booms. Table 10 contains some indicators of parameter stability for the univariate AR(4) which is

¹³ Here, we have introduced three additional real variables: the real interest rate and the (inverses of) money and credit velocities.

¹⁴ With consumer and wholesale prices, there seems to be positive skewness indicating that prices tend to increase faster than they tend to decrease, which obviously makes sense. The behaviour of the long-term interest rate may only reflect this same fact. The real exchange rate, in turn, is characterized by gradual deterioration of competitiveness and once-for-all devaluations of the currency. Money and credit seem to behave in the same way as stock prices in terms of skewness although the estimations results are somewhat different. With bankruptcies, the results represent some sort of puzzle. Industrial output and bankruptcies do not seem to be just mirror images – quite the contrary. Thus, there are some (although not very significant) signs of negative skewness indicating that peaks in bankruptcies are smaller than the corresponding troughs. This clearly indicates that bankruptcies are perhaps more related to financial and institutional variables than just to demand and output.

used as some sort of point of departure in this study. Thus, we have computed the average lag length for depression (the shaded areas in Figure 1) and non-depression periods, the Chow stability test statistic in terms of the sample split and an F-test statistic for the significance of multiplicative (x_{t-1} *depression dummy) terms. It turns out that the stability property is at variance with the data. Moreover, there is some, although not very strong, evidence of asymmetric adjustment in the sense that the average lag length is shorter in depressions than in "normal years".

The stability measures are to some extent consistent with the evidence from the nonlinear adjustment model but some clear inconsistencies also arise. For instance, somewhat conflicting results are obtained for bankruptcies and stock prices. It should be noticed, however, that the classification of observations is based on output behaviour and the cyclical behaviour of other variables, such as stock prices, do not coincide with output movements and, therefore, the results cannot be identical.

Thus, if anything can be learned from this exercise, it is the fact that nonlinearities do seem to exist with the long Finnish series but there are clear differences between nominal and real variables. Thus, it is perhaps futile to analyze all sorts of nonlinearities using a single model as a frame of reference.

5 Testing dependencies between residual moments

The purpose of first applying an autoregressive model to the series is to remove the potential trend component from them. The deterministic or stochastic long-term trend could be removed in other ways as well, e.g. by differencing or modelling with structural time series models and then eliminating the trend component. We proceed by calculating dependency measures of different transformations of these AR(4) residuals.¹⁵ Different moments of residual series and absolute values of residuals are considered as transformations. Therefore we calculate dependence tests from cross-autocorrelations between these univariate residuals as a first step in searching for dynamic relationships.

As can be seen, this procedure looks like an extension of the Granger causality test. However, we start by calculating Portmanteau test statistics without conditioning on past observations of the transformed residuals of the series itself. Portmanteau tests give us potential evidence about the direction and strength of the dynamic dependencies between variables. If the relationship is one-sided, it greatly simplifies the identification of the sources of shocks in these series.

To test whether residuals of the autoregressive model satisfy the properties of independent white noise, this can be seen by calculating the Portmanteau (Q) statistic. This test is designed to detect departures from randomness among the k first auto- or crosscorrelations. The test has the following form

$$Q = T(T+2) \sum_{k=1}^M (T-k)^{-1} r_k^2$$

¹⁵ We also computed the same measures with respect to the ARCH-model residuals. The results turned out to be so close to the results with squared OLS residuals that we do not report them.

where r_k^2 are the squared correlations of the residuals.

This modification of the basic Box-Pierce statistic was first presented in Ljung and Box (1978). The test statistic is asymptotically $\chi^2(M)$ distributed when the original residuals are independent. There is no clear solution in choosing M , but in our case too small values could result in failure to detect dependencies between important higher order lags. As might be expected, increasing M will, on the other hand, lead to lower power of the test (Harvey (1981), p. 211).

The Portmanteau statistic could also be applied to the higher moments or absolute values of stationary series as a general test against non-randomness. McLeod and Li (1983) have shown that squared residuals have the same standard asymptotic variance ($1/T$) as the original series if the residuals are random. In the following tests we assumed the lag order of 24 (2 years) to be large enough to pick up long term dependencies between different moments of residuals. In our application economic theory has rather little to say about the lags between shocks leading to variation in other variables.

Table 11 presents a summary of the estimated Q test statistics. Only the number of significant cases is reported here. The test statistics have been computed both for leads and lags to get some idea of causality. A more detailed report of the results from cross-correlation analysis is available upon request from the authors.

With reference to the table we point out that in general the number of significant values is very high. Almost two-thirds of the coefficients are significant at the 5 per cent level of significance. Particularly in the case of absolute values of the AR(4) residuals, the dependencies are very strong. In accordance with the results from univariate long-memory tests, the results in Table 11 suggest that the long-memory phenomenon also applies to co-movements of different variables - and not only within real and nominal variables but between all macroeconomic variables.

As for the role of different variables, one may note that the bankruptcy variable is very important in terms of the correlation structure. In fact, the number of significant correlations for bankruptcies is bigger than with all other variables. By contrast, the money supply series M1 and M2 and the terms of trade tt are only moderately correlated with other variables.

The test results do not tell very much about causation. In general, the cross-correlation coefficients are of the same magnitude with respect to leads and lags. Therefore, it is very hard to draw any far-reaching conclusions on this matter.

Calculating the contemporaneous correlations between variables does not have any dynamic causal interpretation as it only indicates instantaneous linear co-movement (positive or negative) within a month. As can be seen from Table 11, about one-third of the off-diagonal correlations are significant at the 5 per cent level. The interpretation of (significant) correlations is in most cases rather straightforward. Thus, for instance, consumer prices correlate in an expected way

with, wholesale prices, monetary variables like credit, money aggregate, stock prices and the real exchange rate but not with other real variables.¹⁶

Altogether, the correlations between higher moments of the AR(4) residuals – in the same way as between the absolute values – are so strikingly high that further analysis in a multivariate nonlinear set-up is clearly required. The first step is simply to find out why volatility changes are so much related. In addition, one has to think about a possible explanation to the observed strong co-skewness between variables. Finally, one has also to take into account the fact that the long-memory property also seems to apply to the co-movements of different series – both nominal and real. It seems at least that a (multivariate) ARCH model is not a sufficient or a proper specification to account for these features of the data.

6 Concluding remarks

The empirical analyses presented in this paper have given strong and unambiguous support to the existence of nonlinearities in Finnish historical time series. The univariate case is very clear but it seems that nonlinearities may be even stronger and more important in the multivariate set-up. Obviously, this calls for further research in this area.

It is surely not surprising that the exact nature of nonlinearities cannot be identified. We are inclined to conclude that deterministic chaos is not the probable explanation. It is to be noted that Brock and Potter (1993) arrive at a similar conclusion when they review some recent evidence from macroeconomic and financial data. Another explanation which is often mentioned in this context concerns ARCH and GARCH effects. It is typically found that, after these effects are accounted for, the evidence for nonlinearity and chaos is weakened (see, e.g., Hsieh (1991)). In this study, we found the ARCH effect to be of minor importance. Thus, the explanations for nonlinearities must be sought elsewhere. Nonlinearities may, for instance, reflect neglected nonstationarities but we would prefer to argue in favour of the specific (asymmetric) properties of the short-run (cyclical) adjustment process. There could well be various institutional arrangements and constraints, informational deficiencies, capacity constraints and so on which prevent immediate and symmetric adjustment and which, in turn, explain the empirical findings. Finally, various stability tests clearly indicate that the behaviour of macroeconomic variables is quite different in recession and expansion periods.

It seems highly possible that nonlinearities may change some widely accepted assumptions or results. Thus, for instance, the neutrality of money may not be so good a approximation as it seems in the context of linear models. It may also be that the conventional symmetric adjustment mechanisms represent a very poor framework for dynamic specification. Also, the short and long-run properties of

¹⁶ On the other hand, it is interesting to note that wholesale prices do correlate with both real and monetary variables. Industrial production correlates only with wholesale prices and bankruptcies, but in both cases the sign of the correlation seems to be the opposite than expected. It is also hard to interpret why interest rates correlate positively with stock prices. According to present value formulae, the relation should be just opposite.

different time series and the way in which the corresponding markets function need to be carefully rethought in the light of, for instance, the long-memory results obtained in this study. Finally, it may be that the importance of certain variables (and unimportance of the other variables) in the propagation mechanism of nominal and real shocks in the economy will change a lot if nonlinearities are taken into account. The Finnish data suggest that, for instance, bankruptcies are such a neglected variable.

Table 1. Descriptive statistics for the residuals of a linear AR(4) model

	skewness	kurtosis	median	med(-)	med(+)	stand.dev.
ip	-.66	5.52	.222	-2.058	2.443	.056
bank	-.59	4.70	.747	-12.968	14.141	.306
tt	.66	26.37	.037	-21.609	23.849	2.263
fx	4.18	69.66	-.142	-.725	.540	3.010
r	.45	16.87	.002	-.026	.032	.274
cpi	3.34	30.36	-.126	-.460	.324	.013
wpi	3.04	21.42	-.115	-.506	.325	.012
credit	.04	8.38	.016	-.428	.450	.011
M1	.22	12.55	-.183	-1.612	1.372	.029
M2	-3.08	51.47	-.035	-.525	.472	.011
sx	-.13	5.09	-.096	-2.699	2.337	.050
st	-.085	2.73	-.455	-20.621	21.848	.366

Skewness and kurtosis denote the coefficients of skewness and kurtosis, respectively. Median denotes the sample median, med(-) and med(+) denote the 25 and 75 per cent (quartile) values. In the case of log transformation, the values of the median, med(-) and med(+) have been multiplied by 100. ip denotes (log) industrial production, bank (log) bankruptcies, tt the terms of trade, fx the real exchange rate index, r yield on long-term government bonds, cpi the (log) consumer price index, wpi the (log) wholesale price index, credit the (log) banks' total credit supply, M1 (M2) the (log) narrow (broad) money, sx the (log) Unitas stock price index and st the turnover in stock exchange. The sample period is 1922M5-1996M9. (1) Not significant at the 5 per cent level.

Table 2. Diagnostic test statistics for a linear AR(4) model

	ARCH	RESET2	RESET3	Func. form	WHITE	J-B	TSAY
ip	20.65	0.65	1.72	14.61	11.72	306.88	10.94
bank	18.34	4.51	7.50	15.55	25.02	255.34	27.22
tt	10.69	7.60	8.76	3.74	4.64	2632.10	31.61
fx	8.24	4.18	8.84	2.89	4.70	1121.00	86.74
r	2.91	0.73	0.84	2.26	1.09	1636.60	23.55
cpi	41.43	3.39	3.72	23.05	13.64	473.89	106.39
wpi	8.06	4.36	5.22	8.60	9.44	563.05	29.67
credit	22.68	16.25	8.12	7.14	14.08	746.08	46.96
M1	20.57	5.17	5.96	24.64	9.13	1204.4	137.35
M2	2.94	17.79	10.73	17.34	9.22	1661.2	122.35
sx	44.23	0.12	0.37	41.73	6.88	380.38	48.10
st	17.02	0.46	6.45	18.15	9.26	154.74	52.55

ARCH denotes Engle's ARCH test statistic (with 7 lags), RESET2 test statistic adds the second power of the fitted value as an additional regressor RESET3 includes both the second and third powers of y . Func. form is the F-test of the second power of the explanatory variables and their cross-terms included in the regression. White denotes White' heteroskedasticity/functional form test statistic, J-B the Jarque-Bera test statistic for residual normality and TSAY Tsay's nonlinearity test statistic for 4 lags. 1 % and 5 % denote the critical values of the respective test statistics.

Table 2. continued

	r(1,1)	r(1,2)	r(1,3)	r(1,4)	r(2,2)	r(2,3)	r(2,4)	r(3,3)	r(3,4)	r(4,4)
ip	-.142	.112	-.011	-.119	.114	-.006	.031	-.144	.019	.086
bank	.194	.005	-.101	.021	.115	-.115	-.049	.206	-.019	.192
tt	-.237	-.002	-.123	-.041	.106	.032	.018	-.262	-.083	-.027
fx	-.494	-.370	-.404	.152	-.560	.345	-.634	-.547	.193	-.351
r	-.237	-.049	.013	.038	-.039	-.056	-.046	-.142	-.121	-.418
cpi	.619	.393	-.498	-.598	-.042	-.019	.353	.007	.796	.796
wpi	-.353	.113	-.118	.001	.137	.052	.302	-.378	.124	.044
credit	.124	-.147	-.212	.055	.112	-.113	.148	.069	.198	.009
M1	-.495	-.089	.313	.134	-.837	-.040	.266	-.638	.035	-.297
sx	.298	.188	-.038	.015	-.429	-.133	.148	.058	.115	-.031
random N(0,1)	-.040	-.015	-.011	-.016	-.005	.020	-.050	-.055	-.039	-.015
logistic map	.669	.536	.556	.558	.848	.544	.561	.833	.669	.536

r_{ij} 's are Hsieh's (1991) third-order moment coefficients $[\sum x_t x_{t-i} x_{t-j} / T] / [\sum x_t^2 / T]^{1.5}$.

Table 3. Estimates of correlation dimension with AR(4) residuals

	Embedding dimension				
	1	2	3	4	5
ip	1.22 (0.10)	1.91 (0.11)	2.73 (0.16)	3.53 (0.17)	4.48 (0.31)
bank	1.37 (0.55)	2.07 (0.53)	2.84 (0.46)	3.78 (0.68)	4.82 (1.05)
tt	1.26 (0.88)	1.99 (1.03)	2.74 (1.14)	3.35 (1.00)	3.94 (0.97)
fx	1.06 (0.30)	1.71 (0.50)	2.45 (1.00)	3.32 (2.37)	4.22 (4.99)
r	0.55 (5.51)	0.57 (4.12)	0.59 (2.62)	0.62 (2.37)	0.63 (2.29)
cpi	1.01 (0.16)	1.77 (0.28)	2.52 (0.40)	2.23 (0.40)	3.80 (0.68)
wpi	0.81 (10.08)	1.28 (28.56)	1.37 (82.97)	1.40 (144.18)	1.42 (188.76)
credit	1.00 (0.18)	1.75 (0.33)	2.59 (0.54)	3.50 (1.03)	4.39 (1.55)
M1	1.14 (0.14)	1.87 (0.15)	2.70 (0.15)	3.65 (0.27)	4.69 (0.56)
M2	1.22 (0.10)	1.97 (0.10)	2.83 (0.11)	3.75 (0.19)	4.81 (0.48)
sx	1.17 (0.17)	1.92 (0.18)	2.79 (0.22)	3.61 (0.12)	4.25 (0.24)
st	1.12 (0.21)	1.86 (0.23)	2.72 (0.25)	3.62 (0.30)	4.56 (0.41)
random N(0,1)	1.05 (0.14)	1.85 (0.11)	2.80 (0.16)	3.78 (0.26)	4.61 (0.09)
Henon map	0.97 (0.01)	1.27 (0.03)	1.29 (0.02)	1.30 (0.05)	1.32 (0.05)
Logistic map	0.83 (0.14)	0.90 (0.04)	0.96 (0.01)	0.98 (0.02)	1.00 (0.08)

Numbers inside parentheses are chi-square test statistics for the goodness of fit.

Table 4. **BDS test statistics for the residuals of a linear AR(4) model**

	m=2 ε=0.5	m=3 ε=0.5	m=4 ε=0.5	m=5 ε=0.5	m=2 ε=1.0	m=5 ε=1.0
Original AR(4) residuals						
ip	14.08	20.75	27.47	37.04	12.26	22.62
bank	8.76	13.08	11.68	-6.37	9.95	6.66
tt	11.90	14.84	18.77	24.71	9.52	14.75
fx	13.42	16.29	19.18	22.53	11.48	12.67
r	14.42	17.84	20.74	23.45	10.99	13.83
cpi	11.09	15.31	18.49	22.44	10.45	15.09
wpi	7.33	10.07	12.49	16.02	8.85	12.28
credit	10.95	14.24	17.65	22.62	11.21	17.66
M1	3.28	4.67	5.71	6.45	4.91	7.78
M2	6.70	10.41	13.59	16.83	7.35	14.10
sx	8.01	9.40	10.92	12.25	8.37	13.05
st	2.00	2.66	3.69	4.54	1.82	4.75
random N(0,1)	0.6	0.2	-0.3	-0.4	0.7	0.1
Henon map	165.9	280.8	428.9	717.7	76.0	91.3
logistic map	669.1	881.1	1152.1	1570.0	282.0	250.3
logistic map(4)	2.0	2.0	1.8	-8.1	-1.3	2.0
logistic map(30)	0.7	-2.1	-1.8	-2.2	-1.6	-0.9
ARCH(4) residuals of an AR(4) model						
ip	10.43	14.35	16.16	18.06	5.06	10.40
bank	11.91	14.53	16.63	18.28	9.98	12.39
tt	3.19	4.30	5.57	5.77	1.96	5.34
fx	6.03	8.94	9.08	8.99	1.60	4.40
r	6.21	7.76	8.20	8.40	3.37	5.36
cpi	14.03	14.84	14.63	14.61	9.05	9.50
wpi	10.33	10.63	10.36	9.77	11.08	9.97
credit	13.46	15.25	15.78	16.26	11.31	12.73
M1	6.61	9.21	10.47	10.78	2.80	7.35
M2	6.10	8.44	9.04	9.12	7.55	8.38
sx	9.44	12.98	15.28	17.22	8.03	12.57
st	4.18	5.50	6.56	7.70	4.51	5.87
Shuffled AR(4) residuals						
ip	-2.2	-1.4	-1.0	0.4	-2.5	-1.3
bank	-0.8	-0.2	-0.3	0.4	-0.9	1.1
tt	1.6	2.1	2.0	1.9	1.9	1.7
fx	0.4	1.0	0.7	0.5	1.4	1.5
r	1.9	1.6	1.3	1.2	1.7	1.0
cpi	2.7	2.6	2.3	0.2	0.7	1.2
wpi	-1.0	-1.6	-1.5	-1.2	-1.3	-1.6
credit	0.4	-0.1	-0.6	-0.4	-0.6	-0.8
M1	-0.6	-0.4	-0.3	-0.7	-1.3	-0.2
M2	-1.2	-0.7	-0.7	-0.8	-0.9	-0.3
sx	1.0	1.1	0.6	0.7	1.8	1.6
st	-0.4	0.1	0.1	0.2	-0.9	-0.1
Henon map	0.6	0.3	0.2	0.3	0.1	-0.0
logistic map	1.0	1.3	2.2	0.2	-0.2	1.2

The test statistic is $BDS = T^{1/2}[C_m(\epsilon) - C_1(\epsilon)^m]/\sigma_m(\epsilon)$, where $T = N - m + 1$ and N is the number of observations, $C_m(\epsilon) =$ the correlation integral $= T^{-2}[\text{number of ordered pairs } (i,j) \text{ such that } \|x_{i,m} - x_{j,m}\| < \epsilon]$ where $x_{i,m}$ is the m -history of the time series x and $\sigma_m(\epsilon)$ is the respective standard deviation. Under the null that the series is independently and identically distributed, the BDS has a limiting standard normal distribution. Here, $\epsilon = 0.5$ corresponds to $\epsilon = 0.5 \cdot \{\text{the standard deviation of the residual series}\}$. $\epsilon = 1.0$ is defined in the same way. The shuffled series are obtained by sampling randomly with replacement from the data until a shuffled series of the same length as the original is obtained. The sample period is 1922M5-1994M9. The generated series include 1000 observations. logistic map(4) indicates that every 4:th observation is picked up from a generated series which originally include 30000 observations.

Table 5.

Autocorrelation tests for different residual transformations

Variable	Significance level of the Ljung-Box Q(60) statistic for residual transformation			First-order autocorrelation coefficients for residual transformations		
	u_t	$ u_t $	u_t^2	u_t	$ u_t $	u_t^2
ip	.000	.000	.000	-.007	.326**	.167**
bank	.000	.000	.000	-.026	.346**	.247**
tt	.000	.000	.000	.018	.189**	.038
fx	.430	.000	.159	.023	.362**	.088*
r	.000	.000	.000	-.010	.284**	.072*
cpi	.000	.000	.000	-.008	.397**	.308**
wpi	.001	.000	.000	-.006	.332**	.182**
credit	.000	.000	.000	-.013	.343**	.314**
M1	.000	.000	.000	-.007	.219**	.166**
M2	.000	.000	.000	-.008	.265**	.129**
sx	.000	.000	.000	.002	.253**	.161**
st	.000	.000	.000	-.018	.116**	.197**

* (**)= significant at the 5 (1) per cent level.

Table 6. NLS estimates for the ARFIMA(4,d,0) model

Variable	ip	bank	tt	fx	r	cpi
\hat{d}	-.0377 (.0327)	.8829 (.0483)	.4895 (.0872)	-.1098 (.0799)	.4523 (.1930)	.0873 (.0316)
$\hat{\alpha}_1$.6691 (.0462)	-.5650 (.0566)	.6249 (.0926)	1.1537 (.0866)	.4517 (.1933)	1.2517 (.0450)
$\hat{\alpha}_2$.1948 (.0404)	-.3134 (.0597)	.1327 (.0382)	-.0660 (.0682)	.2181 (.0366)	-.0383 (.0578)
$\hat{\alpha}_3$.1072 (.0378)	-.1826 (.0516)	.2744 (.0378)	-.1851 (.0467)	.1181 (.0478)	-.2125 (.0327)
$\hat{\alpha}_4$.0285 (.0274)	-.1255 (.0387)	-.2225 (.0374)	.0895 (.0284)	.0971 (.0606)	-.0013 (.0064)
$\hat{\delta}^2$.0031	.0906	4.8806	8.8599	.0957	.0002

Variable	wpi	credit	M1	M2	sx	st
\hat{d}	.1421 (.0374)	.0741 (.0159)	.0187 (.0667)	.1038 (.0319)	.0609 (.0501)	.9139 (.0556)
$\hat{\alpha}_1$	1.2378 (.0487)	.8519 (.0368)	.9777 (.0745)	1.2547 (.0438)	1.2339 (.0599)	-.0939 (.0627)
$\hat{\alpha}_2$	-.0132 (.0595)	.1378 (.0441)	.0096 (.0567)	.0860 (.0568)	-.2869 (.0627)	-.1506 (.0452)
$\hat{\alpha}_3$	-.2188 (.0309)	-.0005 (.0331)	.0053 (.0462)	-.3362 (.0301)	.0574 (.0385)	-.0512 (.0415)
$\hat{\alpha}_4$	-.0068 (.0060)	.0099 (.0088)	.0035 (.0363)	-.0047 (.0035)	-.0051 (.0188)	-.1211 (.0376)
$\hat{\delta}^2$.0002	.0001	.0008	.0001	.0025	.1268

Numbers inside parentheses are standard errors. $\hat{\delta}^2$ is the error variance.

Table 7. **ML and NLS estimates for the ARFIMA(1,d,0) model for AR(4) residuals**

Variable	ML estimates			NLS estimates			Geweke Porter Hudak
	\hat{d}	$\hat{\alpha}_1$	$\hat{\delta}^2$	\hat{d}	$\hat{\alpha}_1$	$\hat{\delta}^2$	
ip	-.0621 (.0429)	.0528 (.0551)	.0031	-.0631 (.0429)	.0544 (.0552)	.0031	.202 (.112)
bank	-.1001 (.0365)	.0654 (.0498)	.0926	-.1007 (.0365)	.0659 (.0499)	.0926	.238 (.121)
tt	-.0086 (.0461)	.0273 (.0582)	5.1148	-.0093 (.0458)	.0284 (.0579)	5.1113	-.118 (.152)
fx	.0393 (.0418)	-.0169 (.0530)	9.0370	.0593 (.0387)	-.0549 (.0494)	8.7906	.116 (.108)
r	.0370 (.0410)	-.0468 (.0520)	.0753	.0376 (.0412)	-.0474 (.0521)	.0752	-.142 (.138)
cpi	.0373 (.0362)	-.0461 (.0484)	.0002	.0378 (.0364)	-.0468 (.0485)	.0002	.386 (.144)
wpi	.0239 (.0357)	-.0296 (.0482)	.0002	.0249 (.0354)	-.0318 (.0479)	.0002	.168 (.162)
credit	.0420 (.0308)	-.0553 (.0443)	.0001	.0421 (.0309)	-.0554 (.0444)	.0001	.575 (.142)
M1	.0745 (.0330)	-.0845 (.0455)	.0008	.0747 (.0331)	-.0848 (.0456)	.0008	.402 (.153)
M2	.0676 (.0306)	-.0787 (.0439)	.0001	.0680 (.0307)	-.0791 (.0440)	.0002	.626 (.142)
sx	-.0069 (.0454)	.0084 (.0565)	.0025	-.0071 (.0457)	.0086 (.0568)	.0025	-.342 (.172)
st	-.2032 (.9450)	.1675 (.0581)	.1314	-.2054 (.0438)	.1726 (.0573)	.1311	-.167 (.120)

Numbers inside parentheses are standard errors. Geweke-Porter-Hudak denotes the respective differencing parameter estimate.

Table 8.

The rescaled range V statistics and the estimates of the Hurst exponent ($q=0$)

	V	H
ip	1.28	0.59
bank	1.36	0.49
tt	1.34	0.46
tx	1.66	0.51
r	1.73	0.60
cpi	1.69	0.80
wpi	1.95	0.83
credit	2.77	0.70
M1	2.27	0.72
M2	2.56	0.66
sx	0.83	0.34
st	1.14	0.44

The graphs of the R/S series are presented in Figure 8. On the basis of this figure, one may conclude whether the slope (i.e., the estimate of the Hurst exponent) is almost constant over all data points. The 5 per cent confidence interval for the V statistic is 0.85–1.75.

Table 9. Estimation results of a nonlinear AR model

	a_0	a_1	a_2	a_3	a_4	a_5	a_6	SEE	DW	F3
ip	.273 (2.81)	.105 (3.10)	.580 (10.5)	.201 (3.13)	.043 (2.23)	-.623 (2.15)	.223 (0.26)	.056	2.08	1.79
bank	.921 (3.77)	.071 (1.35)	.273 (3.41)	.146 (1.24)	.096 (2.88)	-.761 (2.45)	-.013 (0.34)	.318	2.24	17.80
tt	.216 (2.45)	.008 (1.47)	1.110 (12.7)	-.603 (4.49)	.333 (2.51)	-4.427 (2.82)	-5.617 (4.02)	.023	2.06	9.96
fx	.416 (4.05)	.002 (0.49)	.830 (8.83)	-.819 (4.98)	.648 (4.13)	-7.859 (4.34)	-1.701 (4.33)	.030	1.90	14.71
r	.005 (0.85)	.000 (0.16)	.774 (8.08)	.027 (0.17)	2.006 (0.86)	-59.76 (0.97)	181.12 (0.83)	.276	1.98	0.74
cpi	-.002 (1.44)	.033 (3.31)	1.344 (31.9)	-.340 (8.03)	-.002 (3.31)	.047 (2.42)	3.467 (1.34)	.013	2.14	10.65
wpi	-.164 (3.70)	.024 (2.58)	1.563 (39.2)	-.497 (11.5)	-.007 (3.37)	.019 (3.31)	-12.64 (3.39)	.013	2.17	12.70
credit	-.017 (0.93)	.019 (1.93)	1.486 (36.0)	-.480 (11.5)	-.001 (1.03)	.000 (0.42)	-106.8 (5.14)	.010	2.17	19.89
M1	-.390 (3.44)	.115 (4.00)	.809 (20.4)	.292 (6.39)	-.007 (3.74)	.008 (3.35)	5.225 (3.06)	.029	2.00	9.70
M2	-.118 (1.52)	.023 (1.25)	1.037 (25.5)	-.011 (0.25)	-.001 (1.63)	.001 (1.12)	-.511 (0.12)	.011	2.00	17.55
sx	.006 (0.96)	.140 (3.32)	1.256 (31.3)	-.281 (7.00)	.001 (0.62)	-.006 (0.25)	.240 (0.25)	.050	1.98	0.89
st	-.492 (7.44)	1.375 (7.45)	.736 (18.5)	.080 (1.92)	.021 (6.94)	-.129 (4.28)	-.011 (0.40)	.362	2.03	37.29

The estimating equation is of the form: $x_t = a_0 + a_1 t + a_2 x_{t-1} + a_3 x_{t-2} + a_4 (x_{t-1} x_{t-2}) + a_5 (x_{t-1}^3 x_{t-2}) + a_6 (x_{t-1} - x_{t-2})^3 + \mu_t$, where μ is the random term. If we restrict $a_4 = a_5 = a_6 = 0$, we end up with a standard linear model. F3 represents an F test statistic for this restriction. The corresponding 5 % (1 %) critical value(s) is 2.64 (3.86). ip denotes (log) industrial production, bank (log) bankruptcies, tt the terms of trade, fx the real exchange rate index, r yield on long-term government bonds, cpi the (log) consumer price index, wpi the (log) wholesale price index, credit the (log) banks' total credit supply, M1 the (log) narrow money, sx the (log) Unitas stock price index and st the turnover in stock exchange. The sample period is 1922M5-1996M9. Coefficient a_5 has been divided by 1000.

Table 10. **Some stability test results**

	Average lag length		Stability tests	
	I	II	Chow	Dummy test
ip	1.44	1.80	5.08	5.18
bank	2.12	2.12	0.68	0.30
tt	0.42	0.74	3.06	3.24
fx	0.88	0.89	3.05	3.03
r	0.79	1.01	1.32	1.55
cpi	0.30	0.83	8.14	10.10
wpi	0.33	0.46	3.64	3.77
credit	0.48	0.38	2.51	2.41
M1	0.68	1.52	8.92	11.75
sx	0.72	0.68	3.77	4.52
5 %	2.22	2.38
1 %	3.04	3.34

The average lag length is computed for the depression periods (I) and non-depression periods (II). Chow notes a Chow test statistic for the hypothesis that the coefficients of the AR(4) model are the same for these two subperiods. The dummy test denotes a F test for the multiplicative dummy* x_{t-1} -terms.

Table 11.

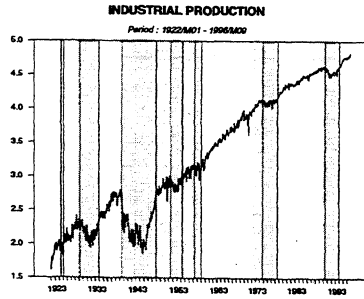
Number of significant Box-Ljung test statistics for the cross-correlation coefficients of different powers of AR(4) residuals

	u	u ²	u ³	u	u _t
ip	3(6)	5(8)	2(6)	10(10)	1
bank	5(7)	8(9)	8(9)	10(10)	3
tt	5(1)	5(1)	5(1)	11(3)	1
fx	7(6)	5(7)	2(5)	8(9)	2
r	5(4)	5(6)	3(3)	8(9)	1
cpi	10(8)	7(6)	6(5)	9(8)	5
wpi	8(9)	8(7)	7(6)	9(10)	4
credit	10(8)	8(8)	8(7)	10(10)	5
M1	4(7)	3(3)	1(2)	6(8)	4
M2	7(6)	4(3)	1(0)	7(7)	2
sx	7(5)	7(7)	6(5)	10(11)	6
st	6(6)	7(7)	4(4)	8(11)	1

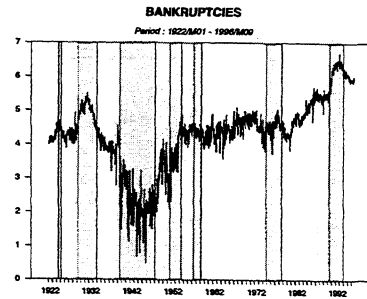
The first number indicates the number of significant Box-Ljung test statistics at the 5 per cent level of significance for the first 24 positive lags of the other variable. The second column (inside parentheses) indicates the corresponding number for the same number of leads. u indicates untransformed residuals, u² squared residuals, u³ third power of residuals, |u| absolute values of residuals and u_t contemporaneous values of residuals (these values are computed from simple correlation coefficients).

Figure 2. **Historical Finnish time series**

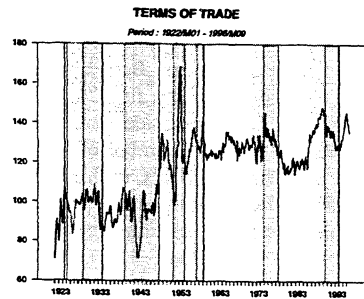
Industrial production



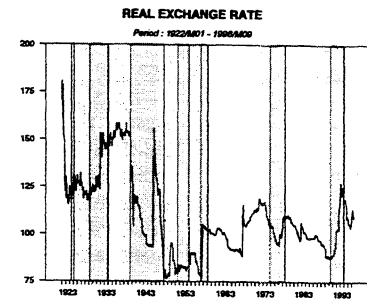
Bankruptcies



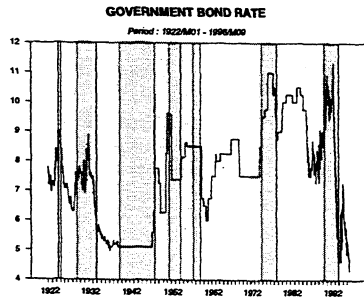
Terms of trade



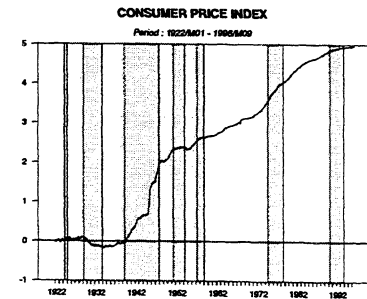
The real exchange rate index



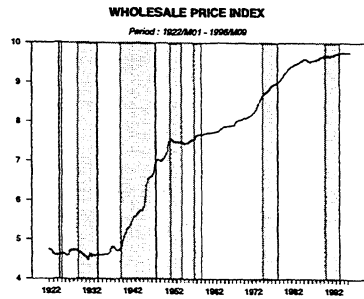
Yield on long-term government bonds



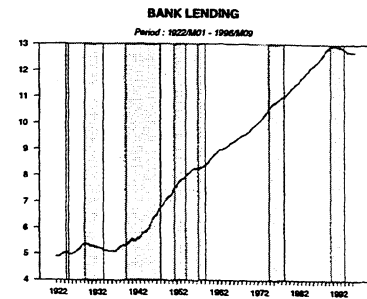
The consumer price index



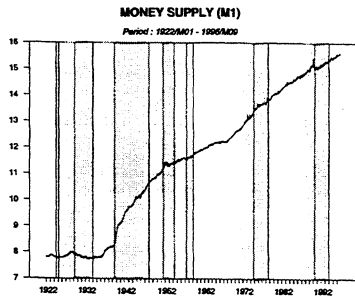
The wholesale price index



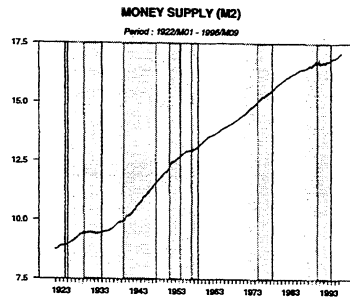
Bank's total credit supply



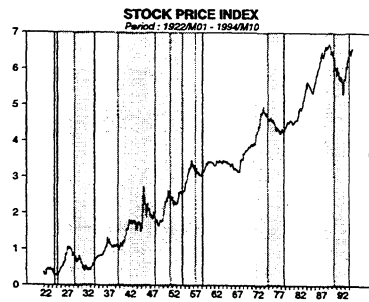
Narrow money (M1)



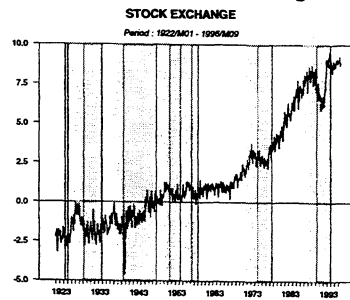
Money supply (M2)



The Unitas stock exchange index

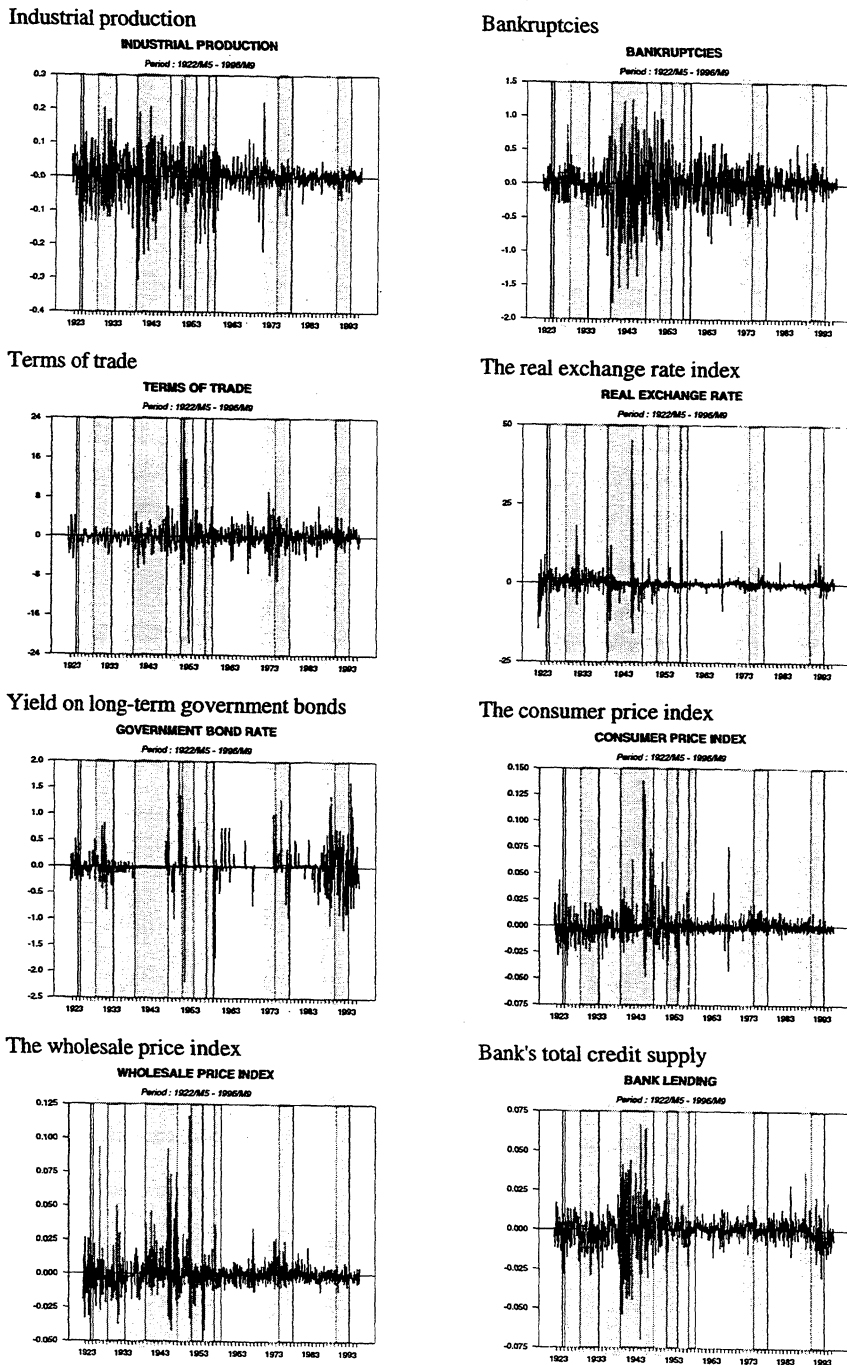


Turnover in Helsinki stock exchange

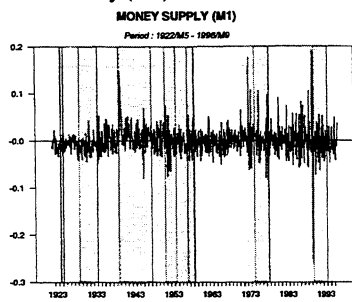


Nominal variables (except for the interest rate) are expressed in logs.

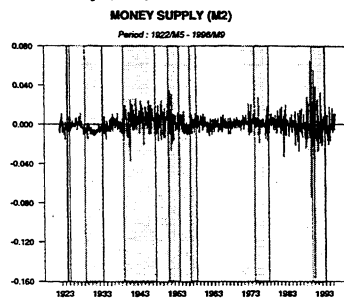
Figure 3. Time series of AR(4) residuals



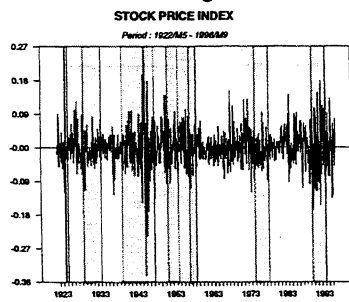
Narrow money (M1)



Broad money (M2)



The Unitas stock exchange index



Turnover in Helsinki stock exchange

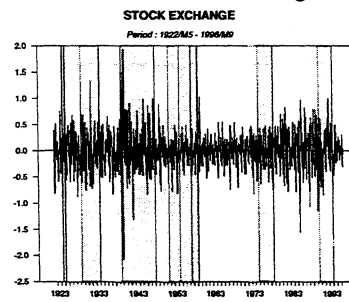
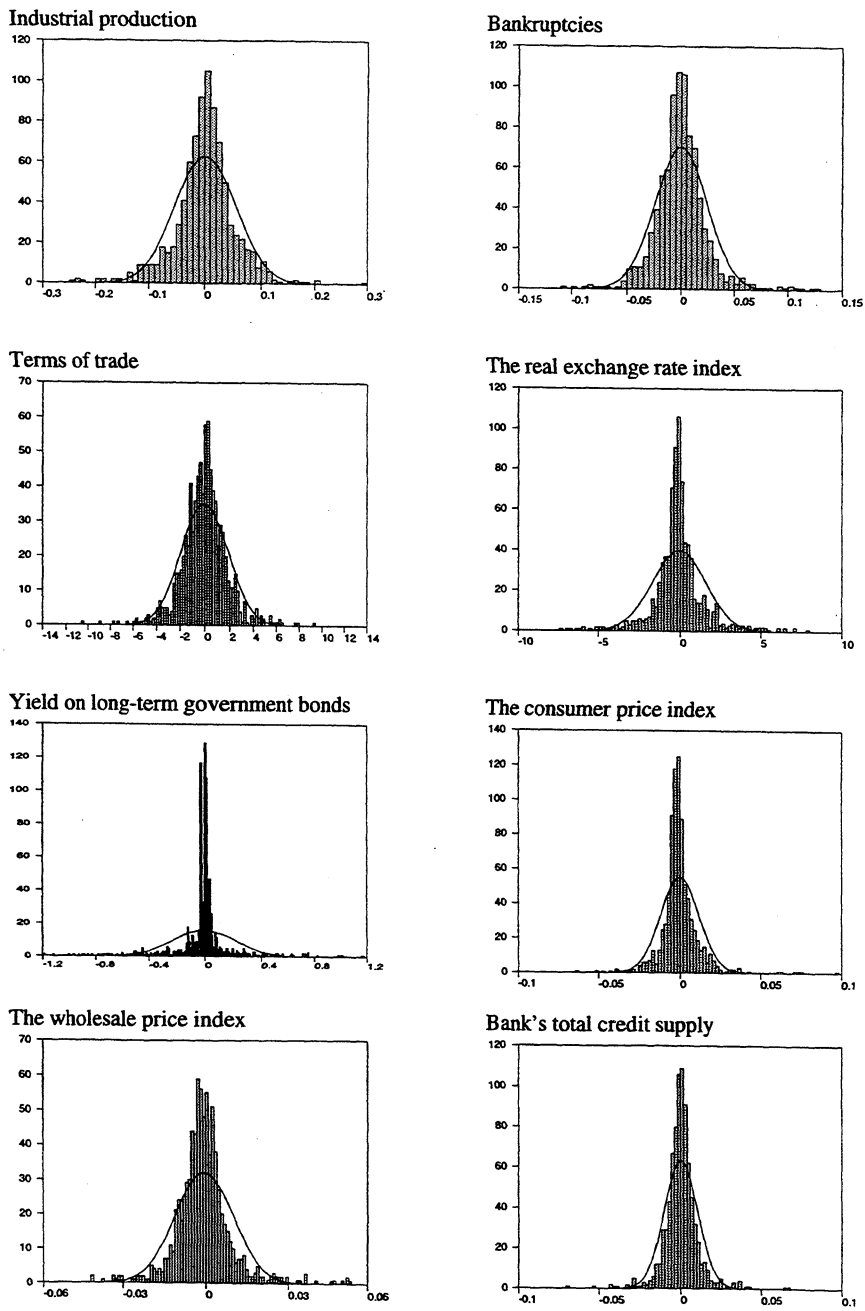
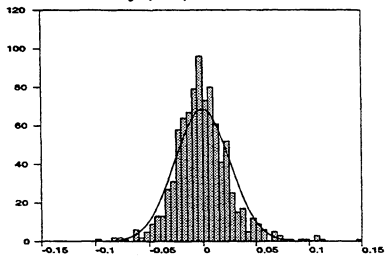


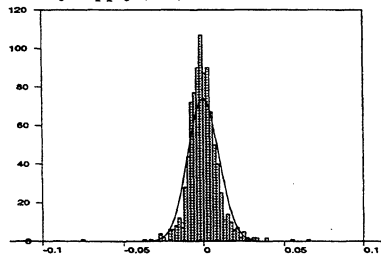
Figure 4. Frequency distribution of AR(4) residuals



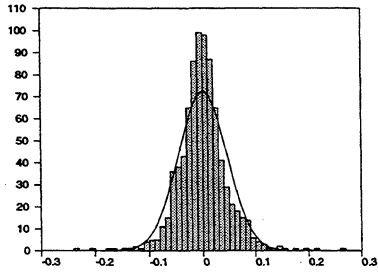
Narrow money (M1)



Money supply (M2)



The Unitas stock exchange index



Turnover in Helsinki stock exchange

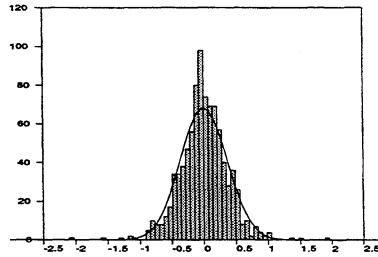
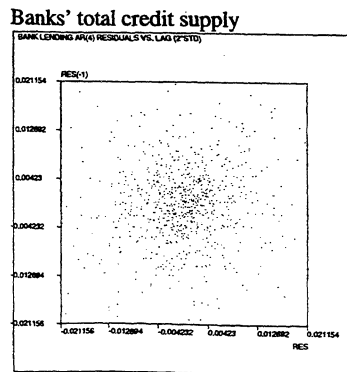
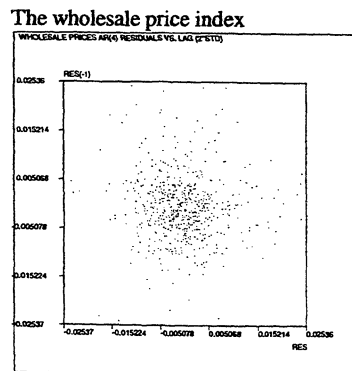
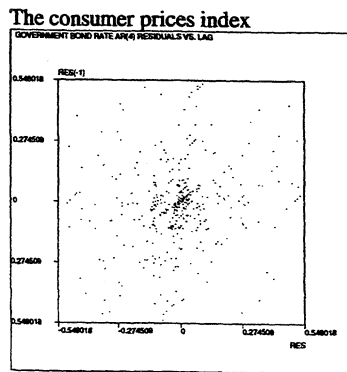
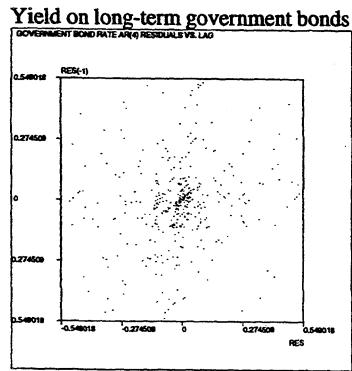
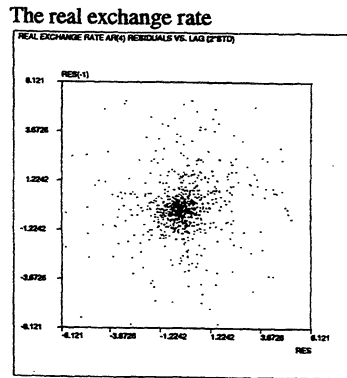
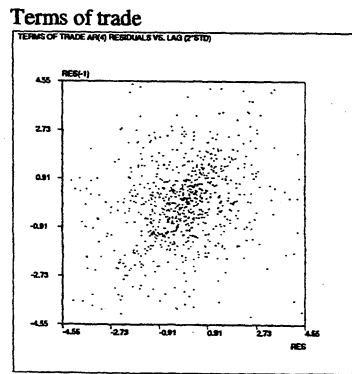
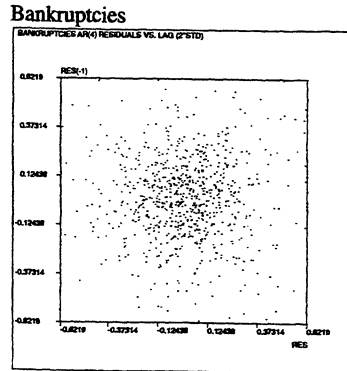
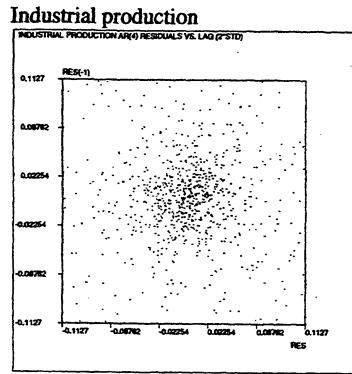
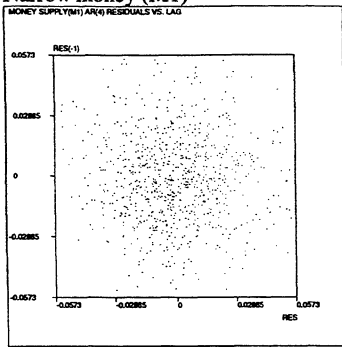


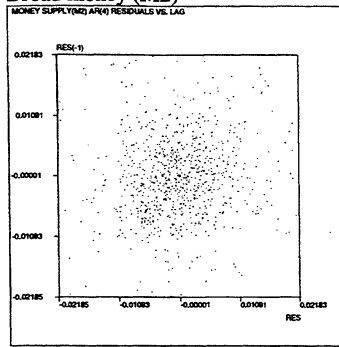
Figure 5. Two-dimensional plots of AR(4) residuals



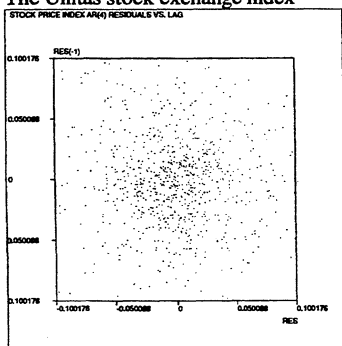
Narrow money (M1)



Broad money (M2)



The Unitas stock exchange index



Turnover in Helsinki stock exchange

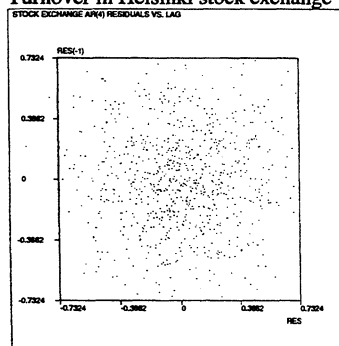
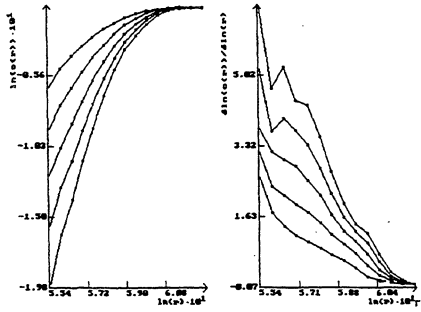
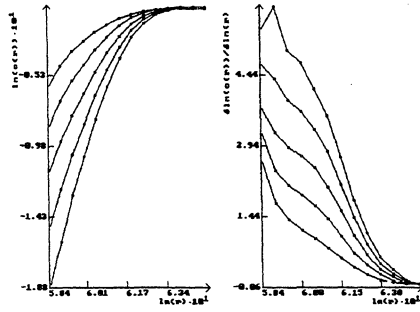


Figure 6. Correlation dimension estimates

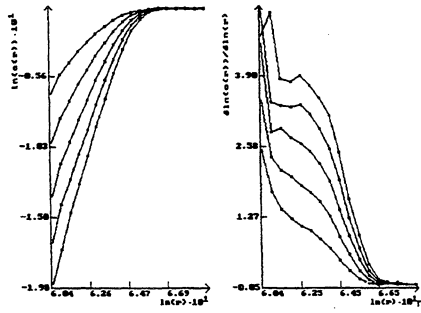
Industrial production



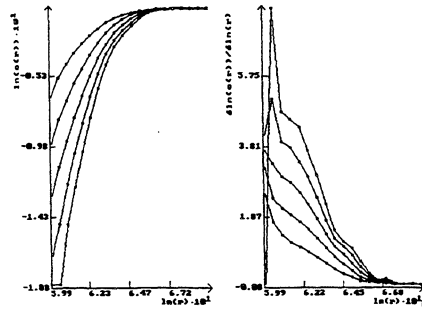
Bankruptcies



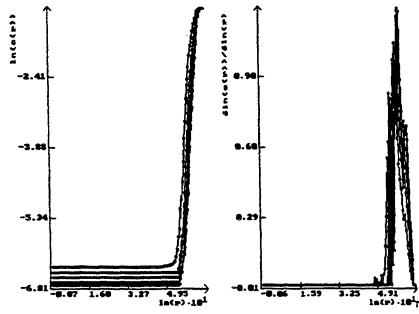
Terms of trade



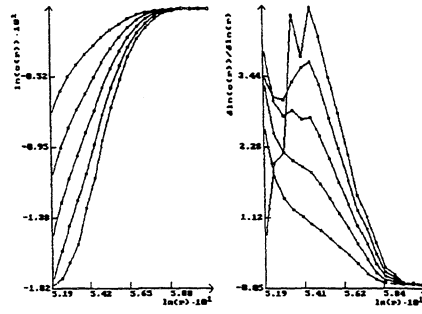
The real exchange rate



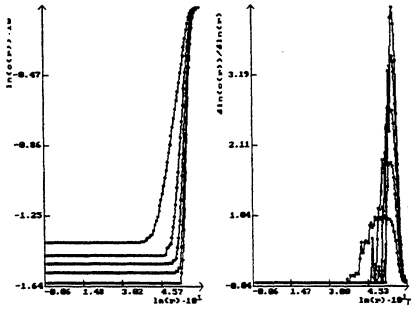
Yield on long-term government bonds



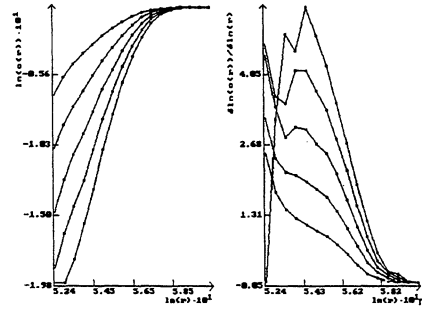
The consumer price index



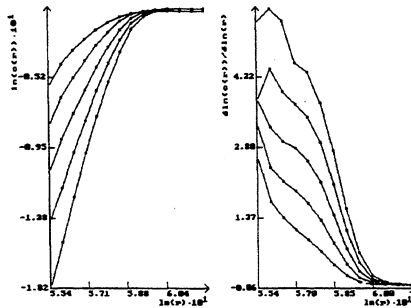
The wholesale price index



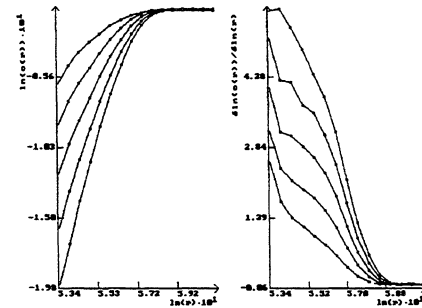
Bank's total credit supply



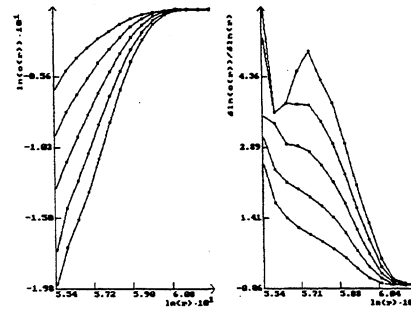
Narrow money (M1)



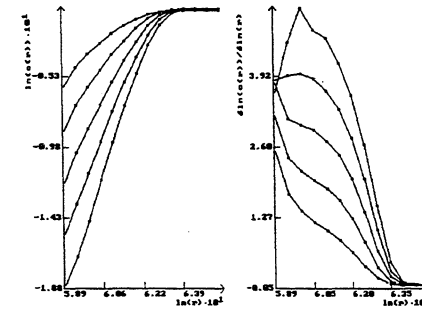
Money supply (M2)



The Unitas stock exchange index



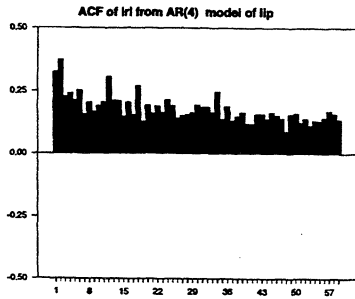
Turnover in Helsinki stock exchange



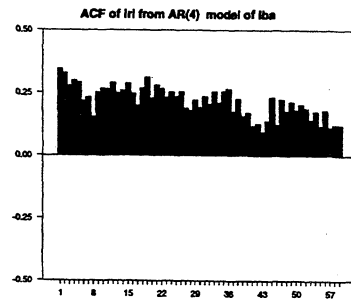
The x-axis in all figures is $10 \cdot \log(\epsilon)$, the y-axis in the left-hand- side figure is $10 \cdot \log(C_m(\epsilon))$ and in the right-hand-side figure $\partial \log(C_m(\epsilon)) / \partial \log(\epsilon)$.

Figure 7. Autocorrelations of absolute values of AR(4) residuals

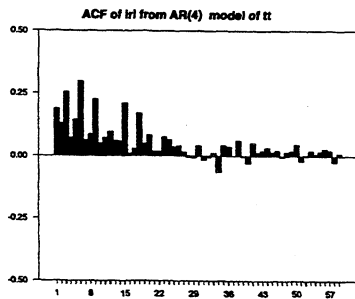
Industrial production



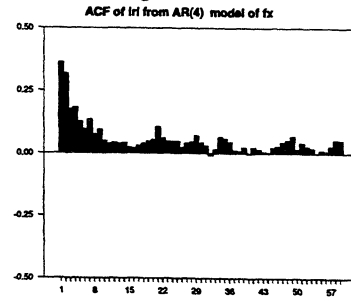
Bankruptcies



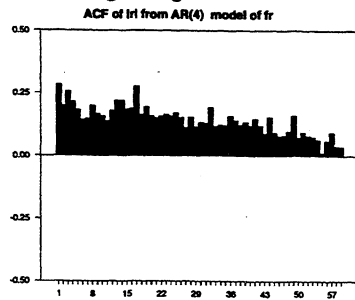
Terms of trade



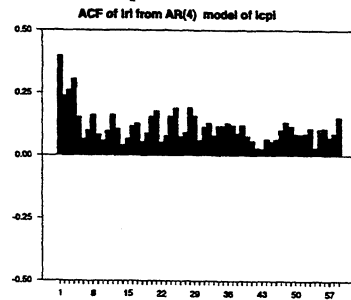
The real exchange rate index



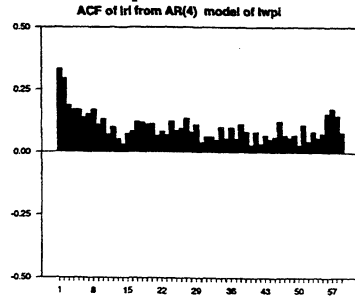
Yield on long-term government bonds



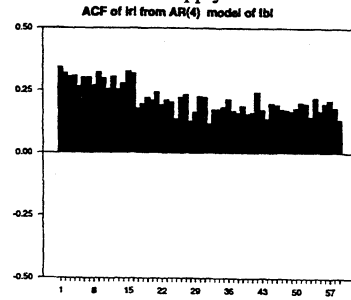
The consumer price index



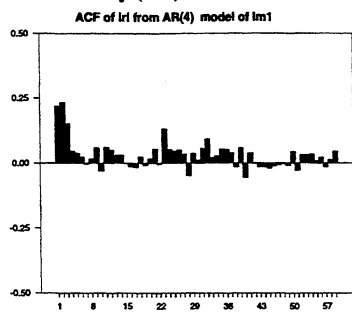
The wholesale price index



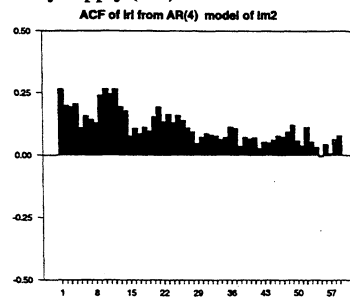
Bank's total credit supply



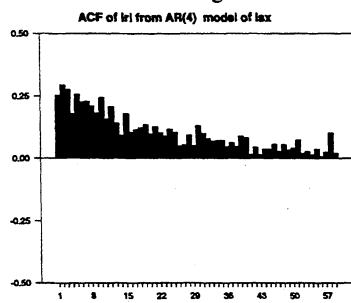
Narrow money (M1)



Money supply (M2)



The Unitas stock exchange index



Turnover in Helsinki stock exchange

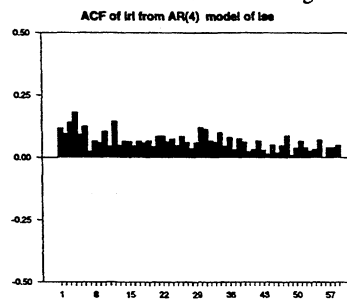
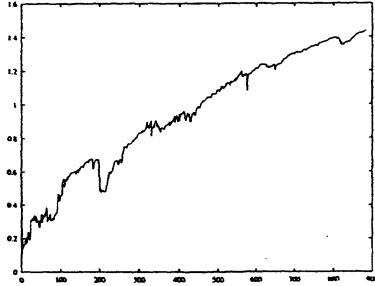
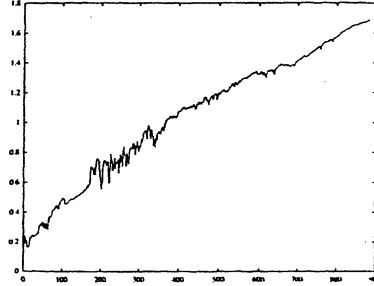


Figure 8. Rescaled range estimates

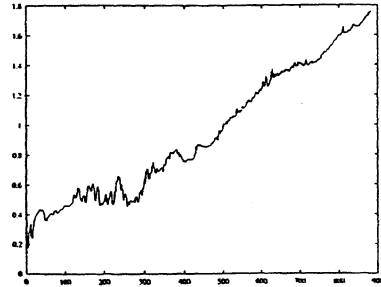
Industrial production



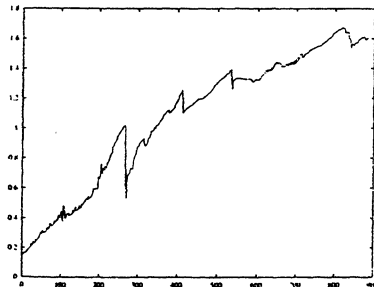
Bankruptcies



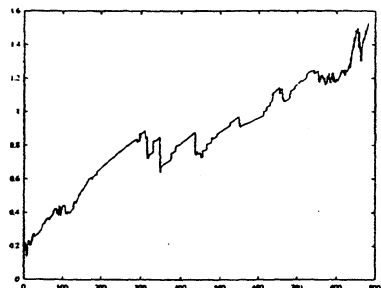
Terms of trade



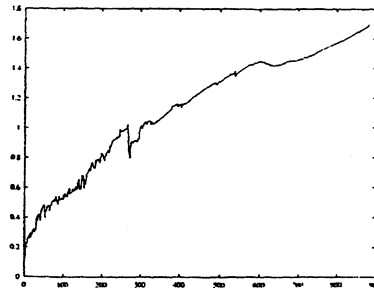
The real exchange rate index



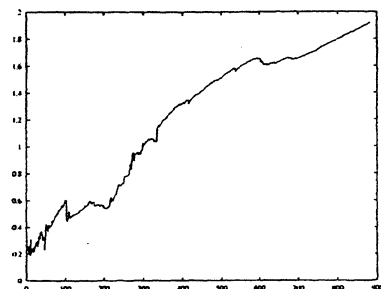
Yield on long-term government bonds



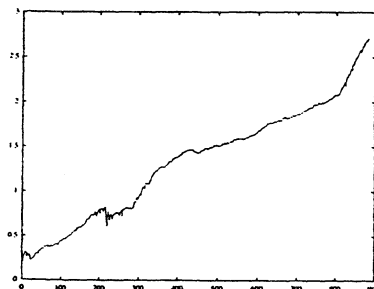
The consumer price index



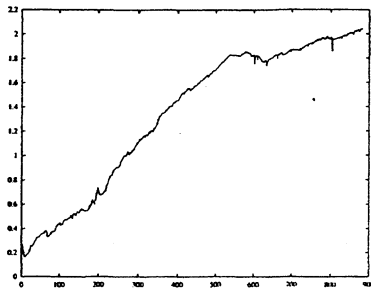
The wholesale price index



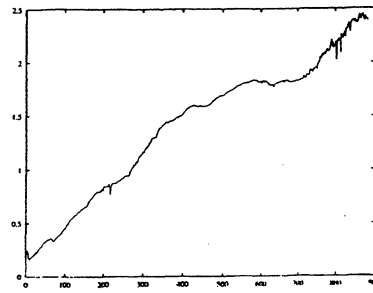
Bank's total credit supply



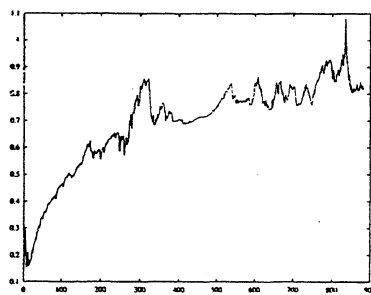
Narrow money (M1)



Broad money (M2)



The Uunitas stock exchange index



Turnover in Helsinki stock exchange

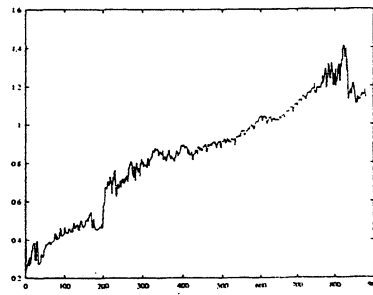
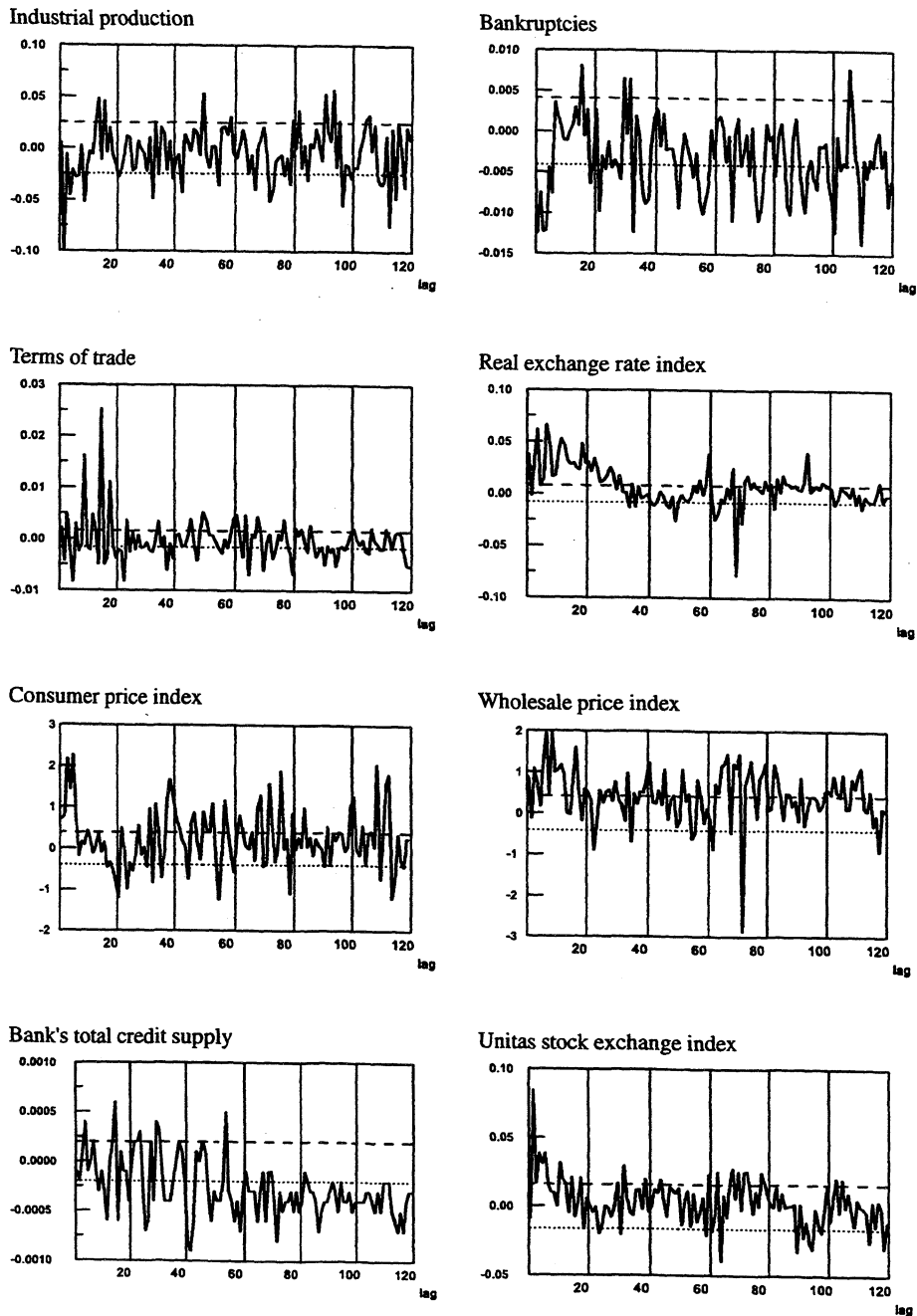


Figure 9.1 Ramsey irreversibility test statistics

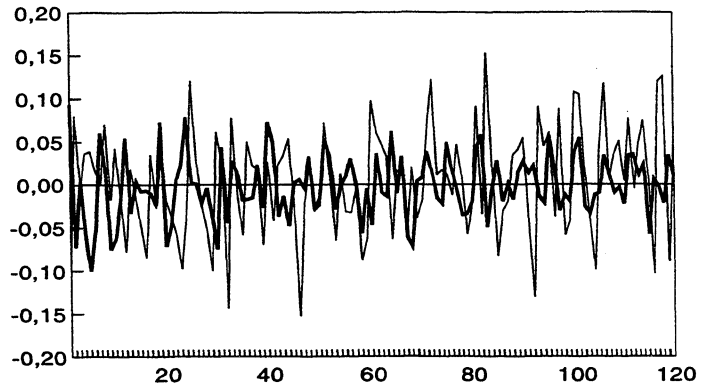


The solid line denotes the $G_{2,1}^k$ test statistic; above and below it are the corresponding 5 per cent confidence limits.

Figure 9.2

Ramsey irreversibility test statistics for ip and sx

$G_{3,1}^k$ -statistic for ip (thin line) and sx (bold line)



$G_{2,1}^k$ -statistic for ip (thin line) and sx (bold line)

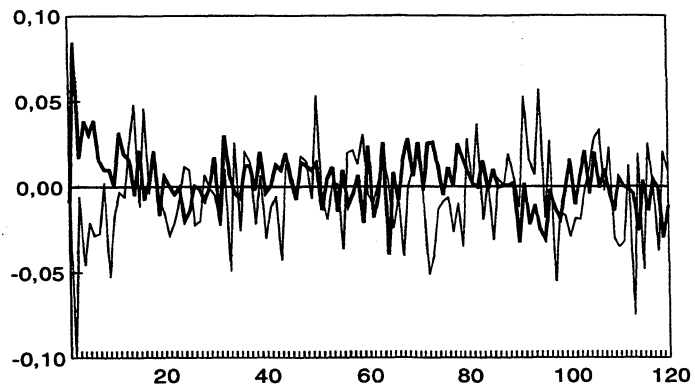
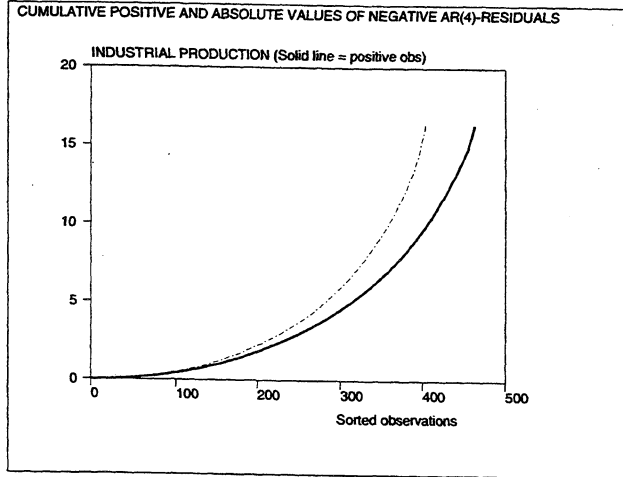


Figure 10.

Residuals for industrial production and stock prices

Positive AR(4) residuals of industrial production
Absolute values of negative AR(4) residuals of industrial production

———
- - - - -



Positive AR(4) residuals of stock prices
Absolute values of negative AR(4) residuals of stock prices

———
- - - - -

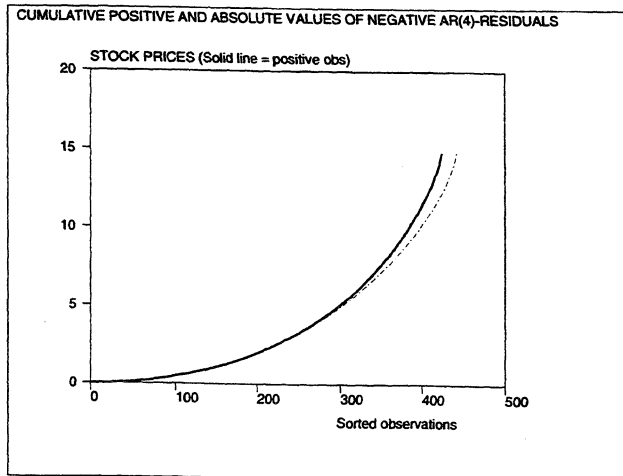
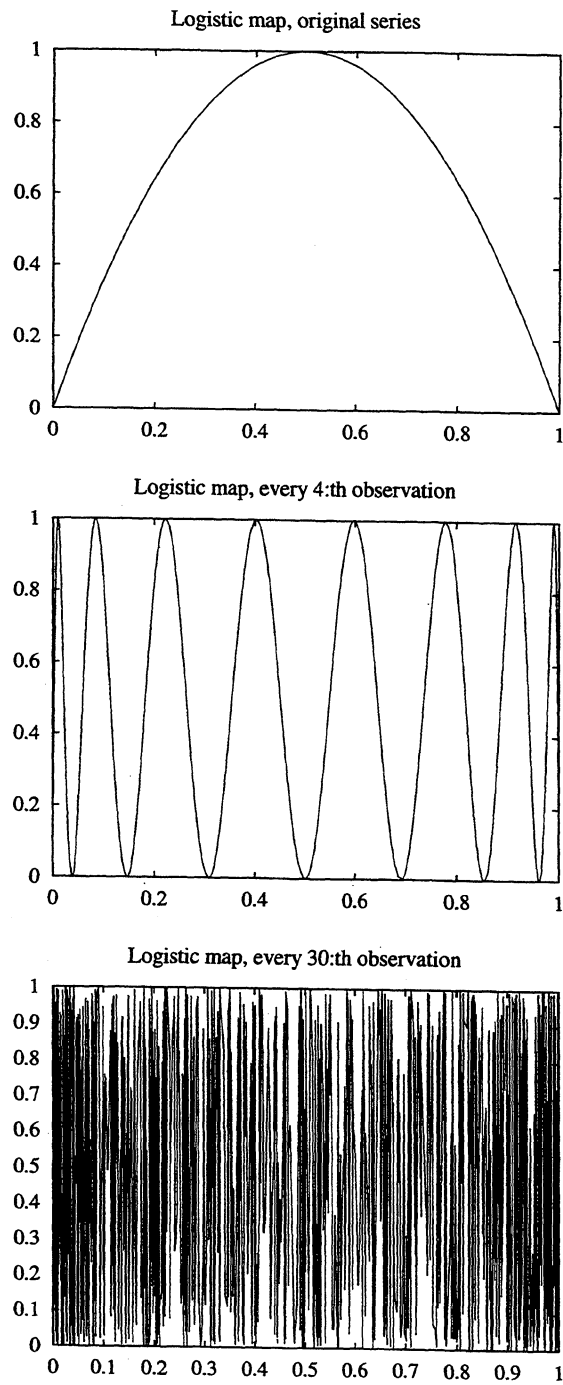


Figure 11. **Effect of sampling on two-dimensional plots**



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