CHRISTIAN C. STARCK

FOREIGN AND DOMESTIC SHOCKS AND FLUCTUATIONS IN THE FINNISH ECONOMY 1960 – 1988

Akademisk avhandling, som med tillstånd av undervisnings- och forskningsrådet vid Svenska handelshögskolan, för avläggande av ekonomie doktorsexamen, framlägges till offentlig granskning fredagen den 11 maj kl. 15 i högskolans auditorium 313.

Helsingfors 1990

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SUOMEN PANKKI
Helsinki 1990

B:44

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ISBN 951-686-241-1
ISSN 0357-4776
Bank of Finland Publications, Series B:44

OY TRIO-OFFSET AB
Helsinki 1990

ACKNOWLEDGEMENTS

The idea for this doctoral thesis took shape whilst I was at the University of Chicago during the academic year 1986 - 1987. Moreover, some of the techniques employed throughout this study were picked up on courses held at the Economics Department and at the Graduate School of Business at the University of Chicago. The bulk of the study was written at the Bank of Finland's Research Department during the fall of 1988 and during 1989 when I was temporarily released from my normal duties. The importance of the excellent working environment provided by the Research Department cannot be overstated, and the stay is deeply appreciated.

In the course of the work, I have benefitted from discussions with and comments from many colleagues both within and outside the Bank of Finland. In particular, I wish to express my gratitude to my examiners Helmut Lütkepohl and Lars-Erik Öller for pointing out some (minor) errors in the draft of this study. They also directed my attention to the clearness of the exposition in general and to many specific passages in particular. Special thanks are also due to Heikki Koskenkylä for numerous discussions and for his admirable endurance in keeping me concerned with the economics - as opposed to the econometrics - of my work. Pertti Haaparanta and Seppo Honkapohja supplied constructive criticism with regard to my theoretical model of a small open economy. In solving programming problems, the help of Thomas Doan, Søren Johansen and Katarina Juselius has been indispensable. Hans Christer Blomqvist, Pentti Pikkarainen, Matti Virén and Anders Vredin supplied useful comments on various parts of the study. I am grateful for the help from my colleagues, but remain solely responsible for the interpretation and implementation of their feedback.

During various phases of the work, I have received able assistance from a number of persons. Esa Jokivuolle carried out tests for the order of integration and drew the figures. Sampo Alhonsuo, Virpi Andersson, Tom Kokkola and Anne Kolehmainen have compiled some of the data and supplied me with figures during early phases of the work. Vuokko Varis drew the hand-drawn figures. Päivi Lindqvist typed most of the text, and occasional help from Marita Castrén, Tuula Naskali and Aila Raekoski is also appreciated. Malcolm Waters checked the language with great efficiency. Pirjo Föhr-Tolvanen edited the text for publication. Antero Arimo and Marja Hirvensalo-Niini organized the printing work. The Bank of Finland Library has - as usual - promptly delivered all the literature I have requested. Sincere thanks are due for all this help.

As my stay at the University of Chicago was of crucial importance in initiating this study, I wish to thank the Fulbright Foundation and the Yrjö Jahnsson Foundation for the financial support that enabled me to visit Chicago.

Helsinki, April 1990

Christian C. Starck

CONTENTS

1 INTRODUCTION

What are the main sources and characteristics of aggregate fluctuations in the Finnish economy? This question is fundamental, but it is tricky since there exists a number of fairly different views about the mechanisms that underly short- and long-run economic fluctuations. Nevertheless, quantitative empirical knowledge is needed to gain insight into the workings of the economy, to aid in the formulation and conduct of economic policy, and to direct attention to phenomena to be explained by means of economic theory. The overall aim of this study is to provide such quantitative empirical knowledge.

Earlier empirical studies addressing the sources and nature of aggregate variability in the Finnish economy have typically been partial econometric analyses or evaluations based on traditional econometric macro models. We will take a fresh look at this topic by drawing on recent developments in modern business cycle research and macroeconometric methodology. These developments include an emphasis on shocks, dynamics, and data-oriented model building. Consequently, we choose to carry out our analysis within the framework of structural vector autoregression models and allowing for multivariate cointegration. Our study is the first comprehensive analysis of a small open economy employing this vehicle of research.

The output of this study will consist of a fairly large body of empirical evidence, from which we hope to be able to extract some stylized facts about shocks impinging on the Finnish economy and the dynamics through which the shocks affect the economy. While such an aim is descriptive, it is far from being modest, since our findings could - provided they are credible - have far-reaching implications for empirical and theoretical model building as well as for economic policy. In any case, our hope is that using a new vehicle of research to address this topic will both deepen, and add to, our understanding of the workings of the Finnish economy.

2 THEORETICAL AND EMPIRICAL ANALYSIS OF ECONOMIC FLUCTUATIONS

The aim of this chapter is to present the subject of our study and to describe earlier related research. Against that background, we state and elaborate on the aim of our study, noting some limitations. The exposition will be kept very brief, since each of the different strands of the literature we shall draw on is voluminous and fairly well known. Rather than offer a complete account of the relevant literature, we seek merely to point out why our approach may be a useful complement to previous studies.

2.1 The Study of Economic Fluctuations

There has never been a consensus on the ultimate sources of aggregate economic variability. Indeed, it is not even completely clear what economic fluctuations are like in terms of time series properties. These statements hold despite research efforts spanning more than a century, as the search for the sources of economic fluctuations go back at least to Jevons and the description of economic fluctuations to Burns and Mitchell. Much of the recent work on economic fluctuations has been done under the auspices of the National Bureau of Economic Research, and good expositions of this work can be found in the Bureau's two conference volumes published in 1951 and 1986, respectively (NBER (1951), Gordon (1986)).

The study of economic fluctuations has enjoyed something of a revival during the last two decades, stimulated in part by the severe supply shocks and development of the equilibrium business cycle model in the 1970s. As Robert J. Gordon puts it, "There seems now to be little dispute that "Understanding Business Cycles," to use the title of a famous Lucas article, is the central preoccupation of theoretical and applied macroeconomics..." (Gordon (1986), p. 2). But what is the current state of the art, or more specifically, what are the possible approaches to the study of the main sources of aggregate fluctuations in the Finnish economy?

The approaches can be classified according to their explicit use of economic theory. At one end of such a continuum we have theoretical models. The Finnish case has not been analyzed expressly, but the analyses of Korkman (1980), Kähkönen (1982), Lempinen (1984) and Haaparanta (1986) contain elements applicable to small open economies. Theoretical analyses enable rigorous and succint treatment of the most essential issues, but stylized, simplified models (deliberately) lack in realism, and, above all, they do not give quantitative insight. Secondly, numerical simulation models have become increasingly popular. This approach has not been applied to Finland, but international examples include Nandakumar (1985) and Bhandari (1987). These models are appealing since they represent theoretical rigour, yet they are relatively complex. However, the heavy reliance on a priori knowledge and a relative lack of realism still limit their use as a source of quantitative knowledge.

In the realm of empirical analyses we can distinguish at least three approaches. Thus, traditional econometric macromodels represent a third vehicle for the analysis of the main sources of aggregate variability. Attempts in this vein include Halttunen (1980) and the Bank of Finland quarterly model (Tarkka & Willman (1985)). Also, Lindskog (1985) has studied this subject with reference to selected Nordic countries. Traditional econometric macromodels represent a fusion of economic theory and statistical observation, but it can be argued that they rely to a considerable extent on more or less implausible a priori, intuitive, or ad hoc restrictions. Their structures are often not easily understandable and the relative importance of various transmission mechanisms is usually fairly unclear. Furthermore, most models are characterized by substantial amounts of inertia, implying contributions to aggregate variability of shocks at implausible time horizons (10 - 20 years). Moreover, respecification and re-estimation to cope with structural change over time is cumbersome, to say the least.

Fourthly, partial models can be used to study the sources of aggregate fluctuations in Finland. Problems with omitted variables, identification and estimation may hamper the use of partial models, however. Finnish studies in this vein include Öller (1978), Blomqvist (1981), Koskenkylä (1985), Aurikko (1986) and Sukselainen (1986). Fifthly, at the other end of the continuum of approaches, purely atheoretical data descriptions can be used for the study of aggregate fluctuations. Comprehensive Finnish studies have not been carried out, but international examples include Burns & Mitchell (1946), the prologue of Frenkel & Razin (1987) and Greenwald & Stiglitz (1988). Such studies leave data free to speak as no a priori restrictions are used, but the approach represents a narrow and limited way of extracting information from observed economic time series.

The second of the two National Bureau of Economic Research conference volumes (Gordon (1986)) on research in economic fluctuations attests to the dominance of the last two approaches mentioned above. The methodological uniformity derives, on the theoretical side, from the breakthrough of the general equilibrium view forcefully advocated by Robert Lucas, Edward Prescott, Thomas Sargent and others. On the applied side the critiques put forward by Robert Lucas and Christopher Sims guide the choices (this is also attested to in the important volume from a conference sponsored by the Federal Reserve Bank of Minneapolis; see Sims (1977)). The sources of economic fluctuations are sought by making minimal but plausible identifying restrictions on models that focus on a small number of broad aggregates summarizing activity in the economy.

The common approach of the bulk of the papers in the National Bureau of Economic Research conference volume to the analysis of economic fluctuations is the vector autoregression (henceforth VAR) approach due to Christopher Sims (see Sims (1980a, 1980b, 1982)). The VAR approach blends ingredients of reduced form econometric analysis with atheoretical data description. Refinements of this approach incorporate elements of structural econometric models thus yielding what we henceforth will call structural VAR models (see Sims (1986)). The bulk of our analyses of economic fluctuations will be carried out using structural VAR models.

Thus, judging by the recent literature (to be further reviewed below), the structural VAR approach would seem to be suitable for the study of economic fluctuations in Finland. While representing quantitative economics its foundations are in a business cycle theory framework which puts a heavy emphasis on dynamics and shocks. Perhaps the most appealing feature of the structural VAR approach is that it strives to deliver macroeconomic models with useful descriptive characteristics without as much of a burden of maintained hypotheses as is usually imposed. Of course, any restrictions used are subject to the same criticism as in the case of standard structural models. Among the drawbacks of the structural VAR approach we note the heavy need for data, which forces one to work with small models. Furthermore, most VAR models, including those in the current study, are linear models.

Since the seminal paper on VARs by Sims (1980a), a large and steadily growing number of papers analyzing economic fluctuations using the VAR approach has been published. Since the vast majority of these studies have focused on the U.S. economy, the literature covering open economy models is considerably smaller.[1] Smaller still is the literature explicitly addressing the question of the main sources of aggregate fluctuations in a small open economy. When the focus is restricted to the variable of ultimate interest - output - the absence of VAR analyses is total, but descriptive analyses employing cross-spectral techniques, like the study by Gerlach (1988), can be found. To the best of our knowledge, only the papers by Burbidge & Harrison (1985), Genberg et al. (1987) and Kuszczak & Murray (1987) have so far used the VAR approach to deal with the same topic as the present study.

Gerlach (1988) sets out to characterize the dynamic interactions between monthly industrial production indices of nine OECD countries. The aim is to establish some stylized facts about output behavior in open economies so as to generate hypotheses for future work. The analysis covers the period 1963 - 1986, hence allowing for separate analyses of periods of fixed and flexible exchange rates.

[1]Examples of VAR models with open economy features include Choudhri (1983), Burbidge & Harrison (1984), Kunst & Neusser (1986), Leiderman & Razin (1986), Baum (1987), Englund & Vredin (1987), Genberg & Salemi (1987), Ahmed et al. (1988, 1989), Ambler (1989), Kugler (1989) and the Bank of Finland publication on VAR models. For VARIMA models of the Finnish economy, see Öller (1982, 1985).

The author draws the following conclusions. Output variability has increased since the breakdown of the Bretton Woods exchange rate arrangement. Output is generally more volatile the more open, but less volatile the bigger the economy is. Output movements have been correlated across countries under both exchange rate regimes, and comovements at business cycle and lower frequencies have increased following the breakdown of the Bretton Woods arrangement. Small open economies (Finland is not included in the analysis) constitute an exception to this; these economies are less affected than other economies by international developments, and the dependence on international output movements of these economies has decreased over time (sic).

The paper by Gerlach (1988) represents a thorough analysis of a topic of fundamental importance, and new and interesting evidence is generated. However, it may be argued that when applied to the current topic, the frequency domain approach is a relatively limited method for emptying the data of its information. The VAR approach would seem to offer most of the insights the frequency domain has to offer, in addition to presenting the information in a more revealing form. More specifically, impulse responses and decompositions of variance are capable of documenting the features offered by cross-spectral techniques, but the exposition is clearer both qualitatively (impulse responses display effects - including the direction - over time) and quantitatively (cumulative impulse responses can be calculated). In addition, the VAR approach enables formal assessments of causality and stability, it permits the researcher to let insights from economic theory have a bearing on the results, and it enables a characterization of the impulses underlying movements in the international economy.

Burbidge & Harrison (1985) set out to explore the interdependencies between the Canadian and the U.S. economies. To this end, a nine-variable VAR model based on monthly data from the period 1971 - 1983 is estimated. The authors reach the conclusion that the Canadian economy is strongly influenced by the U.S. economy. In particular, the U.S. short-term interest rate and money supply affect the Canadian interest rate to a considerable degree. There is also causality from U.S. prices to Canadian prices, and some from U.S. output to Canadian output.

The paper by Burbidge & Harrison (1985) is the first specific attempt to study aggregate variability in a small open economy using the VAR approach. However, several aspects of the analysis evoke criticism. Firstly, the model variables are treated as belonging to the trend stationary family of time series instead of the more likely difference stationary family. Hence, statistical inference as carried out by the authors is treacherous. Secondly, the likely existence of long-run equilibrium relationships between Canadian and U.S. variables is neglected resulting in a likely loss of efficiency of the estimates. Thirdly, ad hoc causal chain models are employed, yielding shocks with blurred interpretations. Fourthly, the rest of the world is assumed to consist of the U.S. economy alone, suggesting a possible omitted variables problem. The omission of an oil price variable raises similar doubts. Fifthly, long-run impacts are dealt with only casually leaving open questions about the actual degree of openness of the Canadian economy.

Genberg et al. (1987) focus on the Swiss economy using a seven variable VAR model and monthly data from the period 1964 -1981. They set out to answer the following questions: how sensitive are Swiss output, prices and interest rates to foreign shocks; has the advent of floating exchange rates served to insulate the domestic economy from foreign shocks; is Swiss money supply exogenous, and; what economic mechanisms are consistent with the empirical findings. The authors reach the following conclusions: foreign shocks explain most of the variation in Swiss variables both under the fixed and floating exchange rate regimes; Swiss money supply is endogenous, and; relative price changes are more important than changes in the domestic price level. However, although an interesting contribution, the paper by Genberg et al. (op.cit.) may be criticized on the same grounds as the paper by Burbidge & Harrison (1985).

Kuszczak & Murray (1987) use various nine-variable VAR models to study the dynamic behavior and interactions of the Canadian and U.S. economies. Their quarterly data cover the period 1964 -1984 and the following countries: Canada, France, Germany, Italy, Japan, the U.K., and the U.S. The aim is to measure the relative importance of foreign and domestic shocks and to compare the effectiveness of various policy actions under fixed and flexible exchange rate regimes. The main results are: foreign variables are of crucial importance for domestic variability; up to 30 per cent of the forecast variance of U.S. output and prices can be attributed to shocks in foreign variables, and; the time series behavior of key macroeconomic variables remained virtually unchanged following the move to flexible exchange rates.

The analysis of Kuszczak and Murray (op.cit.) can be criticized on grounds similar to the ones given above. Furthermore, the use of quarterly data may obscure the delicate dynamics of the fast-moving variables, and it leads to a degrees of freedom problem. Moreover, the use of seasonally adjusted data may lead to mistaken inferences about the strength and dynamics of the economic relationships. Lastly, since the estimated models do not yield estimates identifiable as policy shocks, it is not clear how the aim of comparing the effectiveness of policy actions under different exchange rate regimes can be fulfilled.

We emerge from the VAR literature with the finding that some empirical attempts at studying aggregate variability in small open economies have been made, but that these attempts suffer from many shortcomings. In addition, the previous studies have been of typical research paper size, thus being rather modest in scope. A reliable assessment necessarily requires that the previously mentioned shortcomings be overcome, and a study of the topic in the whole of its compass inevitably requires a rather extensive study. The current study attempts to constitute an analysis with these features for the case of the Finnish economy.

2.2 Aim and Limitations of this Study

The aim of this study is to empirically assess foreign and domestic shocks and their role in aggregate fluctuations in the Finnish economy. Thus, we seek to gain insight into the main sources of aggregate long-run and short-run variability in the Finnish economy and the empirical regularities characterizing the international and domestic transmission of economic shocks.

The aim entails generating stylized facts on economic shocks and their impact on the economy using historical data on key macroeconomic variables. Our methodology imposes as few prior beliefs as possible on the vehicle of research and thus on the results. The identification of shocks focuses on the distinction between foreign and domestic, demand and supply, and nominal and real shocks. The quantification of shocks concerns the size and time series properties, as well as the temporal allocation of the shocks. The transmission of shocks involves assessments of international as well as domestic channels of influence. The impact of shocks is revealed through the time profile and persistence of the impulse response.

For the sake of clarity, some limitations on the study caused by our choice of approach should be stated at this point. Our data-intensive approach forces us to focus on only a relatively small number of variables. Consequently, many potentially important variables will be missing from our models. Data availability limits our analyses to the period 1960 - 1988. Furthermore, we restrict the analysis of aggregate variability mainly to the frequency bounds corresponding to long- and short-run fluctuations. Fluctuations at the seasonal frequency will not be addressed in the same detail as fluctuations at the other frequencies.

3 METHODOLOGY FOR THE EMPIRICAL ANALYSIS OF ECONOMIC FLUCTUATIONS

The aim of this chapter is to present the macroeconometric framework that will be utilized throughout our study. While bits and pieces of this framework have been spelled out in articles and working papers, we feel that a reader who has not closely followed this branch of the time-series literature could benefit from a coherent and succint presentation of the framework. In particular, this applies to the roots and newest developments of the current approach. In addition, we describe in broad outline how the macroeconometric techniques will be implemented.

3.1 Dynamic Macroeconomics: The Frisch - Slutsky Approach

3.1.1 Propagation Mechanism and Impulses

One of the most important contributions of the VAR approach is that it formalizes the by now widely accepted general framework for the study of macroeconomic fluctuations - the framework introduced into economics by Ragnar Frisch and Eugen Slutsky. This framework was accepted as a common analytical framework in the important National Bureau of Economic Research conference volume from the Conference on Research in Business Cycles held in 1984 (see Gordon (1986)). In this section an explicitly articulated formulation of the Frisch - Slutsky framework will be presented to facilitate the interpretation of the general VAR model and to fix notation.

According to Frisch (1933), the first steps toward an analytical framework for the analysis of movements, cyclical or other, in an economic system were taken by the Swedish economist Knut Wicksell in the early 1900s. Wicksell seems to be the first to be aware of the distinction between the propagation mechanism and impulses in the analysis of economic fluctuations. He was also the first to formulate explicitly the theory that the driving force of economic fluctuations is erratic shocks. The mechanism by which such disturbances are transformed into cycles was the subject of three independent papers published in 1927 by Harold Hotelling, Eugen

Slutsky and G. Udny Yule. The most influential of these papers, Slutsky (1927), was later made available to a broader public through Slutsky (1937).

Slutsky experimented with series obtained by performing iterated differences and summations on random drawings, and established that some sort of swings will be produced by the accumulation of erratic shocks. However, Slutsky provided neither specific expressions nor general laws for what kind of fluctuations a given type of shock would cause. This, together with an economic interpretation, was accomplished in the influential paper by Frisch (1933).

In his search for a theoretical setup which could give a rational interpretation of the typical movements in macroeconomic time series, Frisch made a clear distinction between the structural properties of the economic system - the propagation mechanism - and the exogenous shocks hitting the economy - the impulses. Using a linear, three equation mixed difference and differential equation system, Frisch argued that by combining the continuous solution of a deterministic dynamic system and stochastic shocks it is possible to mimic observed macroeconomic fluctuations. Furthermore, he showed that the current state of the world can be thought of as the cumulation of the effects of shocks, the cumulation being made in accordance with a system of weights given by the structure of the economic system. Subsequently, it has become standard practice to interpret linear stochastic difference equations as solutions to linear-quadratic optimization problems (Hall (1978), Long & Plosser (1983)).

The nature of the shocks deserves some clarification. The original Frisch (1933) model is a highly stylized model comprising only real magnitudes driven by more or less undefined erratic shocks and discontinuous Schumpeterian innovations. The model is locally stable around a nongrowing time path. Not much is said about the type or the size of the shocks, but Frisch thought of the shocks as consisting of a continuous flow of white noise impulses. Both Frisch and Slutsky had a bias towards small shocks, although they did not rule out occasional big shocks.[1] Fluctuations in economic activity are seen as being caused by an accumulation of (mainly) small shocks - pieces of new information - each shock being fairly unimportant if viewed in isolation. More recently, this view has been forcefully restated by Lucas (1977).

We will adopt the Frisch - Slutsky - Lucas view of the nature of shocks, noting, however, that this view does not preclude shocks of different kinds (monetary, real, supply etc.) or occasional large shocks. Throughout the analysis we will refer to the shocks interchangeably as shocks, disturbances, impulses and innovations. We will not be referring to them as errors, however. This will serve to highlight the difference between the impulses we have in mind and the error terms in Cowles Commission-type representations of the economy. We think of impulses as pieces of new information, i.e. as unexpected movements in relevant economic variables ("news"). In the "shock models" of the Cowles Commission approach there is heavy emphasis on the propagation mechanism, while the "impulses" are mere unobserved random variables representing "shocks" in behavioral relationships and errors of measurement. While economic variables may contain noise (Black (1986)), this kind of randomness should be kept apart from measurement errors. The impulses we have in mind certainly may contain noise, but people react to it as to pieces of new information. Noise may thus be an inherent feature of information, whereas errors of measurement merely reflect data deficiencies.

[1]The Frisch - Slutsky framework does not, per se, rule out big shocks. However, it does rule out large effects of shocks, since - as pointed out by Blatt (1980) - the effects of shocks must remain small enough to allow a linear approximation of the model.

3.1.2 A Linear Time Series Interpretation of the Frisch - Slutsky Approach

The notion of the Frisch - Slutsky framework that the current state of the economy is a linear function of current and past white noise shocks indicates how deep the foundations of this approach, and consequently the VAR approach, are. The Wold decomposition of stochastic processes (Wold (1938)) provides the connection between the Frisch - Slutsky framework and time series analysis. In this section we will present a formal, linear time series interpretation of the Frisch - Slutsky framework, which will form the basis for the subsequent presentation of our research methodology.

Consider the nx1 dimensional sequence $\{\underset{\sim}{y}_t: t = 1, 2, 3, \ldots\}$ of random variables where the vector $\underset{\sim}{y}_t$ contains variables observed by economic agents at time t. Let $\{\underset{\sim}{\varepsilon}_t: t = 1, 2, 3, \ldots\}$ be the sequence of Nx1 dimensional random vectors used as building blocks for the sequence $\{\underset{\sim}{y}_t: t = 1, 2, 3, \ldots\}$. These building blocks - impulses - generate a sequence of information sets $\{I_t: t = 0, 1, 2, \ldots\}$ where I_0 is generated by any initial random vector $\underset{\sim}{y}_0$ (with a fixed probability distribution), and I_t is generated by $\underset{\sim}{y}_0, \underset{\sim}{\varepsilon}_1, \underset{\sim}{\varepsilon}_2, \ldots, \underset{\sim}{\varepsilon}_t$. The building block process is assumed to be a martingale difference sequence adapted to the sequence of information sets. Formally, the stochastic process $\{\underset{\sim}{y}_t: t = 1, 2, 3, \ldots\}$ is generated recursively using the initial random vector $\underset{\sim}{y}_0$ and the linear, time invariant law of motion

$$(3.1) \qquad \underset{\sim}{y}_{t+1} = \Phi\, \underset{\sim}{y}_t + \Xi\, \underset{\sim}{\varepsilon}_{t+1}, \qquad t = 0,1,2,\ldots,$$

where Φ is a nxn matrix representing the propagation mechanism and Ξ is a nxN matrix that allocates the impulses to different elements of the vector $\underset{\sim}{y}_{t+1}$. The forcing process is assumed to satisfy the following conditions

$$(3.2i) \qquad \underset{\sim}{\varepsilon}_t = \underset{\sim}{0}, \qquad t \leqslant 0$$

$$(3.2ii) \qquad E[\underset{\sim}{\varepsilon}_{t+1} \mid I_t] = \underset{\sim}{0}, \qquad t \geqslant 1$$

$$(3.2\text{iii}) \qquad E[\underset{\sim}{\varepsilon}_{t+1}\, \underset{\sim}{\varepsilon}'_{t+1} \,|\, I_t] = \Xi\Xi' = \Omega, \qquad t > 1$$

$$(3.2\text{iv}) \qquad E[\underset{\sim}{\varepsilon}_{t+1}\, \underset{\sim}{\varepsilon}'_{t+1+s} \,|\, I_t] = 0, \qquad s \neq 0$$

$$(3.2\text{v}) \qquad E[\underset{\sim}{y}_{t+1}\, \underset{\sim}{\varepsilon}'_{t+1+s} \,|\, I_t] = 0, \qquad s > 0$$

$$(3.2\text{vi}) \qquad E[\,|\underset{\sim}{\varepsilon}_{t+1}|^2 \,|\, I_t] < \infty \qquad \forall\, t$$

where $E[.]$ is the expectation operator and Ω is positive definite. We note that by specifying the matrices Φ and Ξ suitably, the Frisch - Slutsky law of motion (3.1) can encompass a great variety of time series processes including deterministic processes, higher order VAR processes, and VARMA processes. In other words, the Frisch - Slutsky framework can produce both deterministic and stochastic growth as well as random, seasonal, cyclical or secular fluctuations (see e.g. Starck (1989)). The dynamics can be uncoupled using the distinct eigenvalues of the matrix Φ and the Jordan decomposition

$$(3.3) \qquad \Phi = \Lambda^{-1}\phi\Lambda$$

where Λ is a nonsingular matrix and ϕ is a matrix constructed using the eigenvalues of Φ. More precisely, for each distinct (possibly complex) eigenvalue μ_j of the matrix Φ one can construct a matrix ϕ_j such that

$$(3.4) \qquad \phi_j = \begin{bmatrix} \mu_j & 1 & 0 & \cdots & 0 \\ 0 & \ddots & \ddots & \ddots & \vdots \\ & & \ddots & \ddots & 0 \\ \vdots & & & \ddots & 1 \\ 0 & \cdots & & 0 & \mu_j \end{bmatrix}$$

with the same dimensions as the number of eigenvalues of Φ which are equal to μ_j. Matrix ϕ is then a block diagonal matrix with ϕ_j in the j^{th} diagonal block. Transforming the state vector $\underset{\sim}{y}_t$ into

$$(3.5) \qquad \underset{\sim}{y}_t^* = \Lambda^{-1}\underset{\sim}{y}_t,$$

substituting (3.5) into (3.1), and using (3.3), we get

$$(3.6) \qquad \underset{\sim}{y}_{t+1}^* = \phi\, \underset{\sim}{y}_t^* + \Lambda^{-1}\Xi\, \underset{\sim}{\varepsilon}_{t+1},$$

which can be partitioned according to the diagonal blocks of ϕ, i.e. according to the distinct eigenvalues of Φ. Thus, the dynamics of the system (3.6) are uncoupled in terms of eigenvalues of the propagation mechanism Φ.

Expression (3.1) gives the recursive representation of the Frisch - Slutsky framework. An equivalent, nonrecursive representation, which also conveys information about the system's dynamics, may be expressed explicitly if we assume that the eigenvalues μ_i, $i = 0, 1, 2, \ldots, n$, of Φ all have moduli strictly less than unity. Define the lag operator L by $L^i(E[\underset{\sim}{y}_{t+1} \mid I_t]) = E[\underset{\sim}{y}_{t+1-i} \mid I_{t-i}]$, $i = \ldots, -2, -1, 0, 1, 2, \ldots$. Then the law of motion (3.1) implies

$$(3.7) \qquad \underset{\sim}{y}_t = A(L)\underset{\sim}{\varepsilon}_t + \Phi^t\underset{\sim}{y}_0$$

where $A(L) = \sum_{i=0}^{t-1} \Phi^i \Xi\, L^i$ is a matrix polynomial in the lag operator. Expression (3.7), the moving average representation of expression (3.1), shows that the current state of the world can be thought of as the cumulation of the effects of shocks.

Expression (3.7) also reveals the connection between the Frisch - Slutsky framework and linear time series analysis. The decomposition (3.7) of the stochastic process $\underset{\sim}{y}_t$ is the Wold decomposition (extended to the multivariate environment). This is seen by noting that the first expression on the right-hand side of (3.7) is a linearly regular process, the second expression on the right-hand side is a linearly deterministic process, and $\underset{\sim}{y}_t$ is the sum of these processes.[2]

In the rest of this chapter we will assume that the process $\underset{\sim}{y}_t$ is a regular linear stochastic process. The regularity assumption will hold if $\underset{\sim}{y}_0 = \underset{\sim}{0}$ or if the deterministic elements of $\underset{\sim}{y}_t$ have been removed before analysis. The model we will be working with then is

$$\underset{\sim}{y}_t = A(L)\underset{\sim}{\varepsilon}_t. \tag{3.8}$$

3.2 Dynamic Macroeconometrics: The Sims Approach

3.2.1 The Vector Autoregression Approach

The time series framework spelled out above became practically applicable in the univariate case through the contribution by Box & Jenkins (1970). Multivariate applications also became frequent in the 1970s. The vector autoregression methodology grew out of the fairly widespread view in the late 1970s that traditional interpretations of the Cowles Commission research program suffered from some critical defects (see Sims (1977) and Sargent (1979)). More generally, the emergence and subsequent popularity of the VAR approach undoubtedly also builds on the unease with the state of macroeconomic theory in the late 1970s and the early 1980s. In his seminal article Sims (1980a) dissented vigorously from the Cowles Commission approach, insisting that the standard identifying assumptions of that approach were "incredible" (op.cit., p. 33). He argued that the zero restrictions embodied in the Cowles Commission structural model

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[2]Let $\underset{\sim}{y}_t^{lr} = A(L)\underset{\sim}{\varepsilon}_t$. When $|\mu_i| < 1$, $i = 0, 1, 2, \ldots, n$, the sequence $\{\underset{\sim}{y}_n^{lr} : n = 1, 2, 3, \ldots\}$ satisfies $\lim_{n\to+\infty} \sup_{m\geq n} E[(\underset{\sim}{y}_n^{lr} - \underset{\sim}{y}_m^{lr})^2] = 0$. Thus $\underset{\sim}{y}_t^{lr}$ is a Cauchy sequence, i.e. a linearly regular process.

$$(3.9) \qquad W_{\underset{\sim}{w}}(L)\underset{\sim}{w}_t = W_{\underset{\sim}{\tilde{w}}}(L)\underset{\sim}{\tilde{w}}_t + \underset{\approx}{\xi}_t$$

where $W_{\underset{\sim}{w}}(L)$ and $W_{\underset{\sim}{\tilde{w}}}(L)$ are matrix polynomials, $\underset{\sim}{w}_t$ is a vector of endogenous variables, $\underset{\sim}{\tilde{w}}_t$ is a vector of exogenous variables and $\underset{\approx}{\xi}_t$ is a white noise error vector, are too restrictive. This may occur, inter alia, because any lagged component of $\underset{\sim}{w}_t$ or $\underset{\sim}{\tilde{w}}_t$ may influence the formation of expectations. For, if economic agents maximize lifetime utility subject to a budget constraint, decisions on consumption, portfolio allocation, labor supply etc. are determined by the same set of variables.[3] Thus, it is possible to argue that all the variables appearing in the Cowles Commission structural model (3.9) are in fact endogenous. It should be pointed out that this argument was by no means new; it had in effect been put forward earlier by Liu (1960). However, the proposed solution to the identification problem was new (Liu was against specifying that every variable is a function of all other variables).

Sims put forward the VAR methodology in the first place, not explicitly because of its deep connections with business cycle theory, but because he wanted to show how "... one can obtain macroeconomic models with useful descriptive characteristics ... without as much of a burden of maintained hypotheses as is usually imposed ..." (Sims (1980a), pp. 32 - 33). He describes VAR models as "... forthrightly descriptive statistical models that do nothing more than summarize correlations in a convenient way" (Sims (1986), p. 3). Indeed, nowadays "...most researchers would agree that vector autoregressions are a parsimonious and useful means of summarizing time series "facts"" (Holtz - Eakin et al. (1988), p. 1371).

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[3]Indirect evidence supporting the expectations argument may be obtained by comparing simulation results of Cowles Commission-type macroeconometric models, models with rational, explicitly forward-looking expectations, and VAR models. The major experiment undertaken in 1986 under the auspices of the Brookings Institution and comprising 11 leading large-scale multi-country econometric models and one VAR model furnishes exactly this kind of evidence. The Liverpool model and the VAR model show very similar results, and these results generally differ from the results of the conventional Cowles Commission-type models (Frankel (1987)).

The VAR approach differs in many respects from more traditional econometric approaches. The most distinct difference is that more weight is given to the a priori beliefs in the conventional approach. In other words, the VAR approach endeavors to impose as few as possible of the biases inherent in conventional structural macroeconometric models. In the VAR approach, the emphasis is on dynamic interactions, the effects of shocks and data-oriented model building, while much less interest usually is placed on estimation of individual elasticities or traditional hypothesis testing. Thus, the VAR approach is most naturally seen to be something of a complement to the traditional (Cowles Commission) approach rather than a substitute for the older approach. Among the newer approaches the VAR approach to some degree overlaps with the data-oriented approach of Hendry (1987).

By defining $\underset{\sim}{y}_t' = (\underset{\sim}{w}_t' \; \underset{\sim}{\tilde{\tilde{w}}}_t')$, where $\underset{\sim}{\tilde{\tilde{w}}}_t$ includes all components of $\underset{\sim}{\tilde{w}}_t$ that are not lagged variables, model (3.9) may be rewritten as

(3.10) $$\underset{\sim}{y}_t = B(L)\underset{\sim}{y}_t + \underset{\sim}{\xi}_t$$

where $B(L) = \sum_{i=0}^{l} \Gamma_i L^i$ is a matrix polynominal in the lag operator. Sims chooses to work with model (3.10) primarily in its vector autoregression form

(3.11) $$C(L)\underset{\sim}{y}_t = \underset{\sim}{\xi}_t$$

where $C(L) = I - B(L)$.[4] He then proceeds by choosing the number of lags l large enough for the model (3.11) to capture the dynamics of the system under consideration. Since even very low order univariate stochastic difference equations can reproduce fluctuations typical of aggregate macroeconomic variables, a small l should, in general, suffice. Having estimated model (3.11) it is inverted to its moving average form

[4]Zellner & Palm (1974) were the first to point out that every linear (or linearized) structural model can be expressed as a restricted VAR model (see also Zellner (1979)).

$$(3.12) \qquad \underset{\sim}{y}_t = D(L)\underset{\sim}{\xi}_t$$

where $D(L) = C^{-1}(L)$ is, in general, an infinite-order matrix polynomial in positive powers of L, and the polynomial C(L) is assumed to have eigenvalues with moduli strictly less than unity.

In section 3.1.2 we demonstrated the connection between the Frisch - Slutsky framework and the Wold decomposition theorem, and above we pointed out the relationship between the Cowles Commission approach and the VAR approach. By comparing the moving average representations (3.8) and (3.12) we can pin down the connection between all the four representations of the process $\underset{\sim}{y}_t$. Essentially, the Frisch - Slutsky, Cowles Commission and the VAR representations all are approximations of the (unknown) Wold representation of $\underset{\sim}{y}_t$.

First, note the similarities between (3.8) and (3.12). Both present the process $\underset{\sim}{y}_t$ as a well behaved sum of white noise building blocks. Convergence of the sums in the Frisch - Slutsky and Wold framework, as well as in the Cowles Commission and VAR frameworks, is achieved through the restriction that the eigenvalues of the propagation matrices (Φ and Γ_i, i = 0, 1, 2, ..., l, respectively) all have eigenvalues with moduli strictly less than unity. However, as pointed out in section 3.1.1, the <u>interpretation</u> of the impulses differs between the Cowles Commission approach and the other approaches. In the former approach they "...represent "shocks" in behavior relations (i.e., the aggregate effects on economic decisions of numerous variables that are not separately observed) and "errors" of measurement" (Hood & Koopmans (1953), p. xv). In the Cowles Commission approach "impulses" arise from the exogenous variables and the error terms in the system. In the latter characterizations the impulses are viewed as fundamental building blocks representing new information hitting the propagation mechanism.

Although only dealt with in an ad hoc way by Sims (1980a), one should also point out the differences in <u>identification</u> of the impulses in the Cowles Commission approach and the traditional VAR approach.

In the Cowles Commission model the "impulses" are identified through the a priori exclusion restrictions on the matrices $W_w(L)$ and $W_{\tilde{w}}(L)$ in expression (3.9). However, when estimating the unconstrained VAR model (3.11) one is not imposing any a priori restrictions on the matrices Γ_i, $i = 0, 1, 2, \ldots, l$ of (3.10). Model (3.11) is actually estimated from the reduced form

$$(3.13) \qquad \underset{\sim}{y}_t = E(L)\underset{\sim}{y}_{t-1} + \underset{\sim}{\zeta}_t$$

where $E(L) = (I - \Gamma_0)^{-1}B^*(L)$, $B^*(L) = \sum_{i=1}^{l} \Gamma_i L^{i-1}$, $\underset{\sim}{\zeta}_t = (I - \Gamma_0)^{-1}\underset{\sim}{\xi}_t$ and $E[\underset{\sim}{\zeta}_t \underset{\sim}{\zeta}_t'] = \Omega_\zeta = (I - \Gamma_0)^{-1}\Omega_\xi(I - \Gamma_0)^{-1'}$. This suggests that it can be treacherous to interpret impulse responses based on residuals from unrestricted VAR models as indicating causality between the elements of $\underset{\sim}{y}_t$. The reduced form disturbances $\underset{\sim}{\zeta}_t$ are linear combinations of the structural shocks $\underset{\sim}{\xi}_t$, and only if one knows the structure of the matrix Γ_0 describing contemporaneous relationships among the elements of $\underset{\sim}{y}_t$, can one give a precise interpretation of the shocks used in impulse responses.

More precisely, the reduced form counterpart to the vector moving average representation (3.12) is

$$(3.14) \qquad \underset{\sim}{y}_t = F(L)\underset{\sim}{\zeta}_t$$

where $F(L) = (I - E(L)L)^{-1} = (I - (I - \Gamma_0)^{-1}B^*(L)L)^{-1} = (I - \Gamma_0)D(L)$ and, in particular, $\underset{\sim}{\zeta}_t = (I - \Gamma_0)^{-1}\underset{\sim}{\xi}_t$. Thus, decompositions of variance and impulse responses - to which we turn in the next section - will, in general, be impossible to interpret without knowledge of the contemporaneous relationships of the structural model underlying the VAR representation. The identifiability problem was first pointed out by Sachs (1982) (see also Cooley & LeRoy (1985) and Leamer (1985a)).

The structural impulses $\underset{\sim}{\xi}_t$ are easily obtained from the estimated reduced form disturbances $\underset{\sim}{\zeta}_t$, as $(I - \Gamma_0)^{-1}$ is the orthogonal matrix of eigenvectors of Ω_ζ, and Ω_ξ is the diagonal matrix of corresponding eigenvalues. If the model is recursive, the decomposition Ω_ζ =

$(I - \Gamma_0)^{-1}\Omega_\xi(I - \Gamma_0)^{-1'}$ is unique, and $(I - \Gamma_0)^{-1}$ will be lower triangular with ones down its main diagonal. The structural impulses are given by $\underset{\sim}{\xi}_t = (I - \Gamma_0)\underset{\sim}{\zeta}_t$, and the matrix Γ_0 describing contemporaneous relationships among the elements of $\underset{\sim}{y}_t$ can be recovered from the estimate of $(I - \Gamma_0)^{-1}$ through elementary matrix operations.

3.2.2 Decompositions of Variance and Impulse Responses

Through decompositions of variance we assess the proportion of the forecast error of each component of $\underset{\sim}{y}_t$ that is attributable to each of the (orthogonalized) estimated impulses. Impulse responses are sequences of responses of one element of $\underset{\sim}{y}_t$ to an innovation in some component of $\underset{\sim}{y}_t$. When talking in terms of forecast errors, one should, however, point out that our aim is not to construct conditional forecasts, but rather to assess the relative importance of different kinds of impulses during a given period. Although VAR models are vulnerable to the Lucas critique, such an exercise is a valid application (Bernanke (1986)).

Assume that the structural impulses $\underset{\sim}{\xi}_t$ of the representations (3.9) - (3.12) are serially and mutually uncorrelated with the diagonal covariance matrix Ω_ξ. The assumption that Ω_ξ is diagonal reflects the notion that the shocks influencing each individual behavioral equation are uncorrelated, i.e. structural in the sense of being attributable to one particular component of $\underset{\sim}{y}_t$ solely. The sample covariance matrix corresponding to the estimated reduced form shocks $\underset{\sim}{\zeta}_t$ in the representations (3.13) - (3.14) is Ω_ζ. As our objective is to obtain estimates of the response of each component of $\underset{\sim}{y}_t$ to structural shocks we further define Ξ_ξ to be the unique positive diagonal matrix that satisfies $\Xi_\xi\Xi_\xi' = \Omega_\xi$. Then the structural representations and the fact that $\underset{\sim}{\xi}_t = (I - \Gamma_0)\zeta_t$ imply that orthogonalized and scaled innovations $\underset{\sim}{\eta}_t$ can be obtained by calculating

$$(3.15) \qquad \underset{\sim}{\eta}_t = \Xi_\xi^{-1}(I - \Gamma_0)\underset{\sim}{\zeta}_t.$$

To obtain the decompositions of variance, define an n-dimensional column vector $\underset{\sim}{\upsilon}_\kappa$ and an N-dimensional column vector $\underset{\sim}{\iota}_\tau$, both consisting of zeroes except for a one in position κ and τ, respectively. Thus $\sum_{\kappa=1}^{n} \underset{\sim}{\upsilon}_\kappa \underset{\sim}{\upsilon}'_\kappa = \sum_{\tau=1}^{N} \underset{\sim}{\iota}_\tau \underset{\sim}{\iota}'_\tau = I$ (if n = N). Let T denote the covariance matrix of forecast errors in $\underset{\sim}{y}_t$ (at some forecast horizon k). Then the forecast error variance of the κ^{th} component of $\underset{\sim}{y}_t$ is $\underset{\sim}{\upsilon}'_\kappa T \underset{\sim}{\upsilon}_\kappa$. Let T_τ denote the covariance matrix of forecast erros in $\underset{\sim}{y}_t$ due to the τ^{th} component of $\underset{\sim}{y}_t$ (at the forecast horizon k). We then define the relative forecast error variance $\lambda_{\kappa,\tau}$ as

$$(3.16) \qquad \lambda_{\kappa,\tau} = \frac{\underset{\sim}{\upsilon}'_\kappa T_\tau \underset{\sim}{\upsilon}_\kappa}{\underset{\sim}{\upsilon}'_\kappa T \underset{\sim}{\upsilon}_\kappa} .$$

Expression (3.16) shows what proportion of the variance of the κ^{th} component of $\underset{\sim}{y}_t$ can be attributed to the τ^{th} shock. In other words, λ tells us how much a shock to one component of $\underset{\sim}{y}_t$ is reflected in a particular component of $\underset{\sim}{y}_t$. Hence, the decomposition of variance gives indications about the main channels of influence in the model and about how exogenous the model variables are.[5]

Essentially the same information conveyed by λ of expression (3.16) is provided by the impact multipliers F(L) of expression (3.14). In terms of orthogonalized residuals the vector moving average representation (3.14) is

$$(3.17) \qquad \underset{\sim}{y}_t = G(L)\underset{\sim}{\eta}_t$$

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[5]We follow the jargon of the VAR literature, in which exogenous refers to the degree of influence one variable has on another in terms of relative forecast error variance. This jargon is not altogether satisfactory, since readers unfamiliar with the VAR literature may associate exogenous with the classical or standard meaning of this term. Hence, in the standard simultaneous equation model, a variable is either "exogenous" or "endogenous", and in a VAR model all variables are "endogenous". While we recognize this possible source of confusion, we hope that the meaning of exogenous and endogenous will be clear from the context.

where $G(L) = (I - (I - \Gamma_0)^{-1}B^*(L)L)^{-1}\Xi_\xi(I - \Gamma_0)^{-1}$. The sequence of matrices $\{G_i: i = 0, 1, 2, \ldots\}$ (which corresponds to the sequence of matrices $\{\Phi^i\Xi: i = 0, 1, 2, \ldots\}$ constructed from the model (3.1)) gives the response of any component of $\underset{\sim}{y}_t$ to a unit disturbance in any component of $\underset{\sim}{\xi}_t$. Thus, the sequence will, in contrast to a decomposition of variance, also give the direction of impact. Hence, the sequence should be useful in describing the dynamics of the propagation. From (3.17) it is evident that the key descriptive output of the VAR methodology cannot, in general, be interpreted without knowledge about the underlying economic structure. The elements of the sequence of matrices $\{G_i: i = 0, 1, 2, \ldots\}$ are referred to as impulse responses.

Impulse responses are a central part of the VAR methodology, and in particular this is so because they deliver a fundamental measure of shock persistence, namely the cumulative impulse response. This is defined as the sum of the coefficients of the moving average (matrix) polynomial in the lag operator of the first-differenced (vector) time series. Rewriting the vector moving average representation (3.17) in difference form yields

$$(3.18) \qquad (1 - L)\underset{\sim}{y}_t = H(L)\underset{\sim}{\eta}_t$$

where H(L) is, in general, an infinite-order matrix lag polynomial. The ultimate impact of a structural shock contained in $\underset{\sim}{\eta}_t$ on the <u>level</u> of a component of $\underset{\sim}{y}_t$ is given by the long-run multiplier matrix H(1) calculated as

$$(3.19) \qquad H(1) = \lim_{k \to +\infty} \left(\frac{\partial E[\underset{\sim}{y}_{t+k} \mid I_t]}{\partial \underset{\sim}{\eta}_t}\right) = \sum_{i=0}^{\infty} H_i.$$

Expectations about future values of $\underset{\sim}{y}_t$ can be formed using linear least squares prediction theory. Specifically, we can use the Wiener-Kolmogorov k-step ahead linear least squares prediction formula to calculate the rational conditional expectation of $\underset{\sim}{y}_t$, which in our case is

$$(3.20) \qquad E[\underset{\sim}{y}_{t+k} \mid I_t] = [(I - B(L))^{-1}/L^k]_+(I - B(L))\underset{\sim}{y}_t$$

where $[.]_+$ is the annihilation operator instructing us to ignore negative powers of L.

Letting γ denote an element of H(1), we note that for a trend stationary process we have $\gamma = 0$, while a pure random walk will yield $\gamma = 1$. It should furthermore be noted that, in general, $\gamma \neq [0,1]$, thus representing an absolute, not a relative, measure of the importance of a unit root component in $\underset{\sim}{y}_t$. The estimated cumulative impulse response is, of course, only an approximation of the (infinite horizon) impulse response function H(1), but the approximation can be made arbitrarily close to the true value. Thus, issues of persistence can be addressed in a convenient way through cumulative impulse responses.

3.2.3 Structural Vector Autoregression Models

In the preceding sections we have shown why unrestricted VAR models do not yield estimates of the true impulses. In general, we will not be able to interpret decompositions of variance or impulse responses based on such unrestricted models. In the language of the Frisch - Slutsky framework, some structure has to be imposed on the propagation mechanism for the impulses to be meaningfully identified from observations on $\underset{\sim}{y}_t$. As most of the issues involved here have already been eloquently dealt with by Koopmans (1947), we will not reproduce that discussion here. Rather, we will show how the "measurement without theory" aspect has been tackled within the VAR approach in recent years. All our VAR analyses will utilize these recent developments; none of the applications are of the (seemingly) astructural type originally introduced by Sims (1980a).

In principle, identification of structural impulses can be accomplished by restricting the contemporaneous propagation matrix and the contemporaneous covariance matrix of the impulses. Such restrictions yield impulses that have structural interpretations, in the sense of the shocks not being nonsensical linear combinations of a host of regression residuals but innovations pertinent to a certain variable or to a certain meaningful function of variables. However, some general caveats to the exact identifiability of innovations should be mentioned.

(i) The VAR models used in the subsequent analysis are fairly aggregated systems. They should be thought of as low dimension representations of the behavior of economic vector time series. Consequently, the estimated shocks may well be linear combinations of current and lagged values of larger groups of related, but distinct disturbances (for an example see King (1986)). However, given the short spans of data typically available, the degrees of freedom constraint in VAR models is severe enough to necessitate focusing on key variables only. The choice of variables to include into the vector $\underset{\sim}{y}_t$ will be dealt with at length in the subsequent analyses. At this point it may suffice to note that one is typically restricted to considering only those variables on which macroeconomics focuses the most interest (output, prices, money, the interest rate etc.).

(ii) Even at a high level of aggregation, one cannot rule out changes over time in the dynamic relationships between different components of $\underset{\sim}{y}_t$. Sims (1980a, 1982, 1986, 1987) argues at length that changes in the propagation mechanism can be expected to be small, but evidence on larger changes has been presented (Litterman (1984), Litterman & Weiss (1985) and Miller & Roberds (1987)). If parameter variation is stochastic, but not allowed for in estimation, some portion of the estimated impulses is merely a reflection of parameter variation.[6]

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[6]Parameter variation may be caused by omitted variables, proxy variables, incorrect funtional form, aggregation and changes in policy rules. If a substantial change in policy occurred, we would expect large changes in all parameters, but because we do not allow for parameter variation, we could observe large contemporaneous shocks in all variables. Thus, large contemporaneous shocks in all variables may be symptomatic of changes in policy rules or other structural relationships.

(iii) Our estimated disturbances will always to some extent reflect (unsystematic) measurement errors in the data.

It should be emphasized that in addition to the identifiability of the structural shocks, the exact interpretation or labelling of the shocks may be problematic in some cases. Thus e.g. a division of shocks into demand and supply shocks may be slippery. A technology shock may influence the supply of goods through fluctuations in work effort (at a given level of inputs), but it may also affect the demand for goods through its effect on wealth and the labor/leisure decision. Even the distinction between real and nominal shocks can be hard to make as in e.g. the case of real interest rate shocks, when the real rate of interest is calculated (in the conventional manner) as the nominal interest rate less the inflation rate, and the nominal interest rate is set by the policy authorities. With these caveats in mind we can proceed to the issue of how to identify the structural impulses.

The most straightforward example of identifying restrictions is the Wold causal chain (Wold (1953)). In fact, the recursive model was the only type of model in use in the early VAR analyses. The Wold causal chain is imposed upon the structural model (3.10), and hence upon the estimated VAR model (3.13), by specifying the matrix Γ_0 to be lower triangular. The reason why all the early VAR models were recursive stems from the use of the singular value decomposition of the sample covariance matrix Ω_ζ - introduced by Sims (1980a) - when orthogonalizing the estimated VAR residuals $\underset{\sim}{\zeta}_t$. The (Cholesky) decomposition allows one to write Ω_ζ = PP', where P is a lower triangular matrix, thus giving orthogonalized VAR residuals $\underset{\sim}{\eta}_t^*$ through the transformation $\underset{\sim}{\eta}_t^* = P^{-1}\underset{\sim}{\zeta}_t$. As in general, $\underset{\sim}{\eta}_t = \Xi_\xi^{-1}(I - \Gamma_0)\underset{\sim}{\zeta}_t$ (see section 3.2.2), and Ξ_ξ is diagonal, the orthogonalization method, in effect, triangularizes the whole model.

In some applications recursion may be a reasonable assumption (as in our model in chapter 5), but it will not be a suitable assumption in many cases. In the early VAR applications the recursivity of the model was often doubted, and the remedy was to evaluate the robustness of the estimation results by reporting results from different orderings of the variables. In many cases varying results were obtained. As pointed out by Bernanke (1986), it is disturbing that authors of the early VAR applications "knew" that their models were recursive, but that they did not know the causal chain of their models. In larger VAR models a recursivity assumption coupled with an unknown causal chain will be infeasible anyway because of the large number of different orderings that would have to be considered (there are n! different orderings in an n-variable VAR model).

Very little has been done to assess how costly it is to use a triangular VAR model when the structural model is not recursive. In a series of papers, Ray Fair has indirectly addressed this question. Fair (1988a) investigates how well a recursive structure approximates his structural model of the U.S. economy (Fair (1984)). This is done by comparing impulse responses generated by three recursive, eight-variable (linear) VAR models to impulse responses from a simultaneous, nonlinear structural model with 29 stochastic equations and over 100 predetermined variables. Fair's documentation reveals that, surprisingly enough, the linear recursive models do approximate the nonlinear simultaneous structure, but he also finds discrepancies between the structures. Fair interprets his findings to suggest that his model of the U.S. economy is not well approximated by recursive structures. Although scant, this evidence suggests that an unwillingness to impose other a priori restrictions than recursivity may be costly when one attempts to use a VAR model to uncover structural relationships in an economy.

The "second generation" VAR models pioneered by Bernanke (1986), Blanchard (1986), Blanchard & Watson (1986) and Sims (1986) achieved identification by imposing more realistic restrictions on Γ_0 than triangularity. More specifically, high frequency fluctuations were restricted by specifying the contemporaneous relationships between the components of $\underset{\sim}{y}_t$ (and the covariance matrix Ω_ξ of the structural impulses) according to standard, sparse macroeconomic models.[7] Thus, orthogonal impulses are recovered without reliance on the Cholesky decomposition, and identification consequently achieved without having to render the model recursive. Such a sparse structural model will be used to identify the structural impulses in chapter 6.

Imposing constraints on contemporaneous relationships in the structural model underlying the reduced form VAR representation is, in principle, similar to the allocation of exclusion restrictions in the Cowles Commission approach. Thus, a VAR model with incredible high frequency restrictions will be open to the same type of criticism that has been raised against Cowles Commission-type models. However, the VAR specification may still be preferable, because the long-run dynamics of the structural model are left unconstrained in a typical "second generation" VAR representation. In any case, the descriptive power of the VAR methodology remains attractive.

Fair (1988b) has studied how well the estimated structural impulses identified by Bernanke (1986) and Blanchard & Watson (1986) resemble structural shocks estimated from the Fair (1984) model. Blanchard and Watson use a model with four variables, Bernanke includes six variables, and Fair operates with more than 100 variables. After a detailed comparison of the dates and magnitudes of large shocks, Fair concludes that the structural shocks identified by the VAR models in general are quite similar to the estimated shocks from the vastly more elaborate conventional model.

In "second generation" VAR models, identification of shocks can also be accomplished by the imposition of low frequency restrictions. Identification schemes involving low frequency restrictions have

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[7]High frequency fluctuations refer to the variability in a time series that corresponds to business cycle, or higher, frequencies, i.e. to short cycles. Low frequency variability refers to the spectral mass of a time series at frequencies close, or equal, to zero, i.e. to long cycles. Hence, high frequency restrictions affect short-run movements, and low frequency restrictions affect long-run movements, of a time series.

been used by King et al. (1987), Shapiro & Watson (1988) and Blanchard & Quah (1989). As in the cases of Wold causal chain or other high frequency restrictions, constraints are put on the matrices of the structural model (3.10) (and, in effect, on the parameter and covariance matrices of representations (3.11) - (3.14)). The long-run constraints derive from the steady state gain matrix H(1) of expression (3.19). The long-run multipliers of the moving average representation (3.18) translate into restrictions on the corresponding autoregressive representation (3.13), which is the model that is actually estimated. Long-run restrictions are employed in the analysis of chapter 7.

3.3 Research Strategy of this Study

First a VAR model of the form (3.13) is estimated, and the estimated disturbances are orthogonalized according to formula (3.15). The impulses are then scrutinized by examining their means, variances, kurtoses, skewness etc. and their relation to major historic economic events. Secondly, the model is inverted to yield a moving average representation of the form (3.17). The dynamics of the estimated system are then uncoupled through decompositions of variance according to formula (3.16) and through plots of the sequence $\{G_i : i = 0, 1, 2, \ldots\}$, and issues of persistence are addressed through the measure H(1) of expression (3.19). The question of changes over time in the model dynamics are addressed by repeating the analyses for subsamples. Time variation in model parameters could also have been handled through use of the Kalman filter (see Sims (1982)), but we refrain from that approach because of the computational burden. Fairly extensive sensitivity analyses are carried out throughout the study, because we strongly believe that "A fragile inference is not worth taking seriously" (Leamer (1985b), p. 308).[8]

The above procedure is applied in chapters 5 - 7 to three different types of structural VAR models. These models also represent

[8]All computations are carried out using RATS (Doan (1989)).

alternative strategies for identifying structural impulses. Firstly, a recursive model focusing on the variable of ultimate interest - output - is estimated. This model comprises the total industrial production of Finland, the output of the eight most important countries for Finnish exports and the output of the U.S. This model is designed for intensive and exclusive analysis of the transmission of output impulses from the U.S. economy to other bigger economies and from these countries to the Finnish economy. The model quantifies the influence of foreign output on the Finnish economy, and suggests the predominant routes by which foreign output fluctuations reach the Finnish economy.

Secondly, a structural VAR model is estimated where identification of the shocks has been achieved by imposing restrictions on the short-run interactions. The restrictions originate from a sparse theoretical dynamic model of a small open economy. The resulting class of models is closest to the conventional Cowles Commission models, and leaves the long-run dynamics to be determined by the data. In this model the domestic economy is complemented by including, inter alia, production, prices and money in the model. This requires aggregation of the international economy into very few variables. Special interest is focused on the long-run impact of policy variables and of foreign variables.

Thirdly, a structural VAR model with restrictions on the long-run behavior of the model is estimated. Hence, the model is designed to allow the data to speak freely about the short-run dynamics of the model. We use the same sparse theoretical model of a small open economy as a frame of reference as in the analysis with short-run restrictions, and base the identifying assumptions on the empirical evidence generated in that analysis. It should be pointed out that while the focus of interest in the second class of models is on issues of economic growth, long-run neutrality etc., and the focus of interest in the third class is on the short-run impact of shocks, it is the same shocks that cause both types of fluctuations.

All models will have in common possible long-run equilibrium relationships that emerge from the preliminary data analysis in chapter 4. If empirical evidence of cointegration among some subset of the key macroeconomic variables under consideration in this study is found, these long-run relations will be taken into account where appropriate. Structural VAR models containing cointegration relationships can be called "third generation" VAR models. Models of this kind have been employed by Blanchard (1986), Campbell & Shiller (1987), Walsh (1987), Shapiro & Watson (1988), Kugler (1989) and Lütkepohl & Reimers (1989a,b).

4 A PRELIMINARY DATA ANALYSIS

The aim of this chapter is to produce empirical evidence on the joint integrating properties of our data. This empirical evidence will constitute information about how the data should be stationarized, how seasonality should be addressed, and about possible long-run equilibria affecting the specification of the subsequent models. This information will also indicate which type of distribution theory should be used for tests of the VAR models. Of course, the joint integration properties of key international and Finnish macroeconomic variables may be interesting in their own right as well.

4.1 Choice and Description of the Data

Seasonally unadjusted data are employed throughout our study. This is because the use of seasonally adjusted data may lead to mistaken inferences about the strength and dynamics of the relationships in VAR models (see Sims (1974), Wallis (1974), Ghysels (1988) and Singleton (1988)). It is somewhat surprising - and worrying - that seasonally adjusted data have been employed in so many VAR studies, including the studies of Sims (1980a), Litterman & Weiss (1985), Bernanke (1986), Blanchard & Watson (1986), Kuszczak & Murray (1987) and Kugler (1989). While the hazards of using seasonally adjusted data have been known for as long as VAR analyses have been done, one may add that the more recent awareness of how the data should, in most cases, be stationarized, has attenuated the need for the use of seasonally unadjusted data (Singleton (op.cit.), p. 372).

Monthly data are used throughout our study. Like the pioneers of the study and measurement of business cycles Arthur Burns and Wesley Mitchell, one can argue that monthly data are basic, since only this kind of data permit observation of the dynamics in the essential detail (Burns & Mitchell (1946), pp. 80 - 81). In particular, monthly data are essential when the system includes fast-moving and jumping entities like prices, money, interest rates and the exchange rate. Furthermore, our approach necessitates the use of a fairly large number of observations, since VAR models are very data intensive and practically all of the distributional results for this class of models are asymptotic. These aspects are further attenuated by the need to be able to analyze subperiods separately and by the potential need to include many variables in the estimated models.[1]

At the risk of stating the obvious, we note that while the VAR approach necessitates the use of large data sets, the critical issue is the amount of genuine information - not the mere number of data points. This remark may be warranted, however, since it highlights the need to consider only actual, "true" observations and cautions against the use of synthetic, "invented" data. In other words, we should refrain from inventing monthly observations where none exist, because applying the data-oriented VAR approach to synthetic data would be tantamount to being methodologically self-defeating. While it certainly is regrettable that monthly observations on many macroeconomic variables do not exist, we will nevertheless be able to include those variables which have usually attracted the most interest (output, prices, money, interest rates etc.).

There are a total of 11 foreign variables and 10 Finnish variables in the data set. The foreign variables are: U.S. real total industrial production y^{*U}, two measures of intermediate economy real output y^{*E} and y^{*E*}, respectively, a broad measure of foreign real total industrial production y^*, foreign consumer prices p^*, a foreign short-term interest rate i^*, and the price of oil p^0. The variables y^{*E} and y^{*E*} are weighted aggregates of eight western economies (excluding the U.S. and the Finnish economy) with weights based on the size of the economy and its importance for Finnish exports, respectively. The variables y^* and p^* are weighted aggregates of the nine most important countries for Finnish exports. The variable i^* is weighted in accordance with the weights of the Finnish currency index. The foreign short-term interest rate is expressed both in nominal and in real terms, and the price of oil is operationalized in two different ways, expressed both in nominal and in real terms.

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[1]Monthly data have also been preferred by Sims (1980b), Burbidge & Harrison (1984, 1985), Litterman (1986), Genberg & Salemi (1987), Genberg et al. (1987), Ambler (1989) and Stock & Watson (1989).

The domestic variables are: real total industrial production y, consumer prices p, M1 m, a short-term interest rate i, net real borrowing requirement of the government g, credit advanced to the public c, the exchange rate e, and the terms of trade p/p*. The short-term interest rate and the exchange rate are expressed both in nominal and in real terms. In the subsequent analyses, all variables are expressed in natural logarithms, except the interest rates, which are expressed in per cent, and g, which is in natural units. Data sources and a detailed account of how the variables are constructed are presented in Appendix 1. Graphs of variables are presented in Appendix 2.

4.2 Joint Integration Properties of the Data

A purely trend stationary univariate time series exhibits transitory fluctuations around a deterministic trend, and the series can be rendered stationary by removing the deterministic trend. A purely difference stationary univariate time series shows no tendency to stabilize around some deterministic path, and this type of process can be stationarized by taking differences of the series (Nelson & Plosser (1982)). Nelson and Plosser (op.cit.) analyzed some key U.S. economic time series, and produced evidence in favor of the view that most of the series belong to the difference stationary family. They also pointed out some implications of such a finding for hypothesis testing and for business cycle theorizing.

Subsequent research has tended to support the findings of Nelson and Plosser (op.cit.), and the implications for statistical inference and business cycle theorizing have been elaborated upon. The following related results can be mentioned. If the components of $\underset{\sim}{y}_t$ include random walk components, and the VAR model is estimated in levels with deterministic trends, the estimated impulses $\hat{\underset{\sim}{\varepsilon}}_t$ will exhibit pseudo-cyclical and purely artifactual dynamic properties (Nelson & Kang (1981)). Low frequency movements in the estimated impulses will be exaggerated and high frequency movements will be attenuated (Chan et al. (1977)). In addition, the model will exhibit spurious correlation (Nelson & Kang (1984)). There is some evidence that causal relationships will be weakened or removed when such a misspecified model is employed (Kang (1985)). On the other hand, improper differencing will attenuate low frequency movements in the estimated impulses while exaggerating high frequency movements (Chan et al. (op.cit.)). In VAR analyses, sensitivity with respect to the detrending method has been encountered by Shapiro & Watson (1988) and Blanchard & Quah (1989).

Inference in multiple time series models with integrated processes has been studied by Phillips & Durlauf (1986), Sims et al. (1987) and Park & Phillips (1988). Theorem 3.2 of Phillips & Durlauf (op.cit.) shows that OLS regression of our basic time series representation (3.1) yields consistent estimates even if $\Phi = I$, or if assumptions (3.2iii) or (3.2iv) fail to hold. Specifically, their theorem 3.3 shows that even when the forcing sequence $\{\underset{\sim}{\varepsilon}_t: t = 1, 2, 3, \ldots\}$ is heteroskedastic and temporally dependent the OLS estimator of this innovation process will be consistent. A generalization of these results to inference in multiple time series models with near-integrated processes has been presented by Phillips (1988).

The situation when deterministic elements are introduced into model (3.1) has not yet been fully worked out. Sims et al. (1987) and Park & Phillips (1988) study a VAR(1) model, and generalize the analyses of Phillips & Durlauf (1986) by allowing for unit roots in each component of the vector process and a nonzero drift vector in the VAR model or in the integrated processes. The authors conclude that although the coefficient estimates are consistent (in fact, the OLS estimator for Φ of model (3.1) can converge at a rate of $O_p(T^{-1})$; see also Stock (1987)), the limiting distribution of the error variance-covariance matrix Ω is singular and nonnormal when more than one component of $\underset{\sim}{y}_t$ is dominated by a stochastic trend. Moreover, if model (3.1) contains a constant vector, a deterministic trend component, and more than one of the components of $\underset{\sim}{y}_t$ are random walks with drifts, then the estimators of the mean vector and of parameters on regressors other than the deterministic trend vector will be consistent, but the estimator of the deterministic trend coefficients will depend on the drift parameter. Again, all limiting distributions will be nonnormal.

Test statistics for VAR models with integrated processes and deterministic elements have been studied in more detail by Sims et al. (1987) and Park & Phillips (1988). Sims et al. show that the conventional test for lag length will have the usual asymptotic χ^2 distribution when the test concerns linear restrictions on coefficients on mean zero stationary regressors. Standard distribution theory even applies in certain cases if there is a drift or a deterministic trend in the regressors. In general, causality tests and tests for long-run neutrality in cointegrated systems have standard asymptotic distributions (Sims et al. (1987), Lütkepohl (1989), Lütkepohl & Reimers (1989a) and Reimers & Lütkepohl (1989)).

A Monte Carlo Study by Ljungqvist et al. (1988) on the distributions of such test statistics confirms that the nonstandard distributions differ substantially from the usual normal or χ^2 distributions. On the other hand, the asymptotic approximations to these nonnormal distributions are characterized as adequate in sample sizes of 100 and good in samples with 400 observations. A Monte Carlo study by Ohanian (1988) presents some evidence according to which the presence of an integrated regressor in a VAR model estimated in levels has important effects on block exogeneity tests and some moderate effects on decompositions of variance.

Moving on to issues involving common unit roots, i.e. cointegration, it will prove instructive to express the basic time series representation (3.1) in a slightly different form. The law of motion (3.1) can be written in the form

$$(4.1) \qquad (1 - L)\underset{\sim}{y}_{t+1} = -\Pi \, \underset{\sim}{y}_t + \Xi \, \underset{\sim}{\varepsilon}_{t+1}$$

where $\Pi = I - \Phi$. Notice that when model (3.1) is explicitly enlarged to higher order autoregressions, expression (4.1) emerges as a normal (assuming that $\Xi = I$) first differenced VAR model with the exception of the level term $-\Pi\underset{\sim}{y}_t$. Expression (4.1) can also be referred to as an error correction model (see Harvey (1981)).

The concept of cointegration formalizes the old notion that some linear combinations of time series variables appear nonstationary, whereas others appear to be stationary (see Sargan (1964), and Stock & Watson (1988) for some older references). The notion of cointegration, first introduced by Granger (1981), can be made precise through the specification of the coefficient matrix π of expression (4.1). If $0 < \text{rank}(\pi) = r < n$ there will exist nonzero nxr matrices F and G such that $\pi = FG'$. The cointegrating vectors $\underset{\sim}{\alpha}$ form the columns of G and have the property that the quantity $\underset{\sim}{\alpha}'\underset{\sim}{y}_t = z_t$ is stationary. Since the columns of G form the distinct cointegrating vectors there may exist up to r (up to a nonsingular rxr matrix) unique cointegrating vectors. The columns of the matrix F contain the loadings with which z_t enters model (4.1). If rank $(\pi) = n$ $\underset{\sim}{y}_t$ is stationary and if rank $(\pi) = 0$, model (4.1) reduces to a model for a stationary process $(1-L)\underset{\sim}{y}_t$.

Engle & Granger (1987) interpret the relationship $\underset{\sim}{\alpha}'\underset{\sim}{y}_t = 0$ as an equilibrium and consequently a quantity $\underset{\sim}{\alpha}'\underset{\sim}{y}_t = z_t$ as an equilibrium error. More generally, $\underset{\sim}{\alpha}'\underset{\sim}{y}_t$ need not be a stationary combination of the elements of $\underset{\sim}{y}_t$, but only a linear combination of lower order of integration than the individual elements of $\underset{\sim}{y}_t$. The equilibrium may be thought of as a hyperplane in n-dimensional real space R^n corresponding to an attractor set $A \subset R^n$ towards which the system (3.1) is moved by the attractor process z_t. In general, the attractor can be bounded as well as unbounded, and there may exist disjoint attractors with different strengths of attraction and with different ability to capture the system within itself (Granger (1987)).

However, we will restrict the present analysis to linear, static, constant, integer, zero frequency cointegration among first moments of variables of the same order of integration. Thus, e.g. cointegrated but nonstationary systems will not be dealt with. This is because, at present, implications of cointegration for VAR systems and exact asymptotics or Monte Carlo results for testing for cointegration are known for this case only.[2]

4.3 Empirical Findings

4.3.1 Tests for Integration

We begin by examining the degree and nature of integration among our series. The general technique is to isolate the unit roots as coefficients in finite autoregressive estimating equations. We follow Nelson & Plosser (1982), Stulz & Wasserfallen (1985), and others, in using the test statistics developed by Fuller (1976), Dickey & Fuller (1981) and Hasza & Fuller (1982). Distinction is made between time series containing seasonal elements and series which do not. The distinction is based on the autocorrelation and partial autocorrelation patterns at seasonal lags of the differenced series (see Appendix 3). No convincing signs of seasonal fluctuation were found in the series i^*, $i^* - \dot{p}^*$, e, e^r, p^{01}, p^{01r}, p^{02}, p^{02r}, and p/p^*. The series y^{*U}, y^{*E}, y^{*E*}, y, p, m, e, y^*, p^*, g, i and $i - \dot{p}$ were found to contain seasonal elements.

Since our tests for unit roots are sensitive to how the alternative hypothesis is specified (in applications, this has been encountered by Downes & Leon (1987), West (1987) and Rappoport & Reichlin (1989)), several different testing models are used. The models are shown in Table 4.1. For the non-seasonal time series three models implying three alternative hypotheses of different complexity were employed. N1 in Table 4.1 is the most restrictive model, with a pure driftless random walk under the null hypothesis. The order p of the

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[2]Nonlinear cointegration has been studied by Granger (1987) and dynamic cointegration by Salmon (1988). Time-varying cointegration appears in Granger (1986) and Osborn et al. (1988) and fractional integration in Diebold & Nerlove (1988). Cointegration at other frequencies than the zero frequency appears in Engle et al. (1989) and cointegration among higher moments in Escribano (1987). Cointegration of variables with different orders of integration has been studied by Davidson (1986) and Johansen (1988a).

correction for serial dependence in the test regression residual ν_t was set at p = 6 to ensure that the residual is white noise and to correct for possible moving average components in the variables under investigation (see Molinas (1986) on the former, and Said & Dickey (1984) and Schwert (1987, 1989) for elaborations on the latter). N2 and N3 add a constant and a trend to the test regressions, respectively.

TABLE 4.1

Models for tests for the order of integration

Non-seasonal (monthly) time series

(N1) $$x_t = a_1 x_{t-1} + a_2 v_{t-1} + \dots + a_p v_{t-p} + \nu_t$$

(N2) $$x_t = a_1 x_{t-1} + a_2 v_{t-1} + \dots + a_p v_{t-p} + b_0 + \nu_t$$

(N3) $$x_t = a_1 x_{t-1} + a_2 v_{t-1} + \dots + a_p v_{t-p} + b_0 + b_1 \mathrm{time}_t + \nu_t$$

Seasonal (monthly) time series

(S1) $$x_t = a_1 x_{t-1} + a_2 (1-L)L^{12} x_t + a_3 (1-L^{12})Lx_t + a_4 \tilde{v}_{t-1} + \dots + a_q \tilde{v}_{t-q} + \nu_t$$

(S2) $$x_t = a_1 x_{t-1} + a_2 (1-L)L^{12} x_t + a_3 (1-L^{12})Lx_t + a_4 \tilde{v}_{t-1} + \dots + a_q \tilde{v}_{t-q} + b_0 + b_1 \mathrm{time}_t + \sum_{i=1}^{11} h_i D_{it} + \nu_t$$

x_t = variable investigated, $v_t = (1-L)x_t$, $\tilde{v}_t = (1-L)(1-L^{12})x_t$, time = time trend (see Appendix 1), D_t = seasonal dummy variable (see Appendix 1) and ν_t = white noise residual.

Tests for non-seasonal and/or seasonal roots are performed using models S1 and S2. S1 is the seasonal counterpart of N1, and S2 is the most general seasonal model including a constant, a trend and seasonal dummy variables. The order q of the correction for serial dependence in the test regression residuals ν_t was set at q = 12 in order to whiten the residuals and to correct for the effects of possible moving average components in the seasonal series.

It should be strongly emphasized that the hazards of inference about integration are considerable when dealing with sample sizes like ours. Monte Carlo evidence by Dickey & Fuller (1979, 1981), Evans & Savin (1981, 1984), Banerjee et al. (1986) and Hakkio (1986) suggests that the power of unit root tests is exceedingly low so that large type II errors are a potential difficulty. The possibility that our series contain moving average components further distorts rejection rates. Monte Carlo evidence by Molinas (1986) and Schwert (1987, 1989) show that dramatic departures from the usual critical values can occur in these cases.

All tests are performed using the maximum amount of data, but the analyses are repeated for three subsamples. Primarily, we split the sample into what could be called Bretton Woods and post-Bretton Woods subperiods. In addition to representing data from two exchange rate regimes, the subperiods differ radically with respect to oil price volatility and corresponding overall volatility. Fairly widespread use of index clauses during the first subperiod, and a complete absence of de jure indexation during the second subperiod, furthermore stands out with respect to the Finnish economy. As the Bretton Woods arrangement was gradually abolished during 1972, and the first oil price shock occurred during the fall of 1973, we choose to split our sample at 1973M1. In addition, possible structural changes are investigated by also focusing on data from the 1980s only. The 1980s may differ from the 1970s since the EMS came into effect in 1979 and the FED adopted a new policy implementation in the fall of 1979. Thus, our analyses will be performed for data from the periods 1960M1 - 1988M8, 1960M1 - 1972M12, 1973M1 - 1988M8 and 1980M1 - 1988M8.

Results of the tests based on the full sample period for the order of integration of nonseasonal and seasonal variables are presented in Tables 4.2 and 4.3, respectively. Results for the subsample periods are presented in Tables A4.1 - 6 in Appendix 4. The results of the tests with nonseasonal variables can be summarized as follows. With one exception, all nonseasonal variables appear to be integrated of order one (I(1)). The foreign real rate of interest does not seem to display unit root nonstationarity, however. The foreign nominal interest rate appears to be I(1) albeit for some weak and mixed evidence in favor of I(0). Hence, we would expect the foreign price level to be I(2) (when using the (1 - L^3) filter to generate three-month inflation rates). The terms of trade are I(1) possibly with drift. All other variables are well approximated as I(1) variables without drift in their logarithmic form.

Since we cannot a priori rule out changes in the data generating processes of our variables during our sample period, a look at subsample evidence is necessary. However, subsample analyses lack in power even more than full sample analysis, and the subsample findings must consequently be considered with a healthy dose of caution. Summarizing, the subsample analyses documented in Tables A4.1 - 3 of Appendix 4 furnish the following picture. All variables are well approximated as I(1) variables during 1960 - 1972, except for the foreign real rate of interest which is I(0). In 1973 - 1988, we are unable to reject the hypothesis that the foreign real rate of interest contains a unit root. Moreover, the terms of trade seem to be I(0) with trendlike behavior or perhaps drift. There is also some evidence against a unit root in the oil price series, notably so in the case of the import price of energy series. During 1980 - 1988, all variables seem to be I(1), and the terms of trade I(1) with drift.

TABLE 4.2

Tests for the order of integration of nonseasonal variables using data from the period 1960M1 - 1988M8

	Data in levels				Data in first differences			
	t	F1	F2	F3	t	F1	F2	F3
i^*	0.998				0.404			
	(-0.61)				(-5.55)***			
	0.976	3.53			0.404	15.4***		
	(-2.64)*				(-5.54)***			
	0.965		3.34	5.00	0.401		10.2***	15.3***
	(-3.15)*				(-5.53)***			
$i^*-\dot{p}^*$	0.983				-0.880			
	(-1.27)				(-10.9)***			
	0.925	3.63			-0.880	58.9***		
	(-2.69)*				(-10.9)***			
	0.880		4.23*	6.35**	-0.881		39.2***	58.7***
	(-3.55)**				(-10.8)***			
e	1.000				0.255			
	(1.20)				(-6.39)***			
	0.996	2.27			0.201	22.0***		
	(-1.18)				(-6.64)***			
	0.969		2.98	2.87	0.196		14.8***	22.1***
	(-2.33)				(-6.66)***			
e^r	1.000				-0.118			
	(-0.40)				(-7.42)***			
	0.985	1.12			-0.126	27.6***		
	(-1.39)				(-7.42)***			
	0.985		0.90	1.19	-0.134		18.5***	27.8***
	(-1.37)				(-7.45)***			
p^{02}	1.000				0.456			
	(0.70)				(-6.11)***			
	0.997	1.29			0.437	19.5***		
	(-1.16)				(-6.24)***			
	0.993		1.01	0.88	0.425		13.3***	20.0***
	(-1.08)				(-6.32)***			
p^{02r}	1.000				0.348			
	(-0.40)				(-6.75)***			
	0.987	1.38			0.344	22.8***		
	(-1.60)				(-6.76)***			
	0.988		1.02	1.43	0.325		15.7***	23.6***
	(-1.36)				(-6.87)***			
p/p^*	1.000				0.366			
	(1.00)				(-5.56)***			
	0.997	4.16*			0.223	19.2***		
	(-1.23)				(-6.20)***			
	0.979		3.85	2.34	0.210		13.0***	19.5***
	(-2.03)				(-6.24)***			

The estimated first-order autoregressive parameter a_1 (see equations (N1) - (N3) of Table 4.1) is given in the column denoted t and the t value of this estimate in parentheses. The null hypotheses are: t; $a_1=1$, F1; $b_0=0$, $b_1=1$, F2; $b_0=b_1=0$, $a_1=1$ and F3; $a_1=1$, $b_1=0$ (see equations (N1) - (N3) of Table 4.1). The Dickey-Fuller tests are based on regressions with six lags of the differenced variable under investigation. Dummy variables taking the value 1 for the following variables (dates) and 0 otherwise have been employed in the test regressions: e^r (1967M11 - 1967M12) and p^{02}, p^{02r} (1971M11, 1974M8). Critical values for the tests are tabulated by Fuller (1976) and Dickey & Fuller (1981). Rejection at the 10,5 and 1 per cent level of significance is indicated by *, ** and ***, respectively.

TABLE 4.3

Tests for the order of integration of seasonal variables using data from the period 1960M1 - 1988M8

	Data in levels			Data in first differences				Data in levels			Data in first differences		
	t	F4	F5	t	F4	F5		t	F4	F5	t	F4	F5
y^{*U}	1.000	7.25***		0.883	23.2***		c	1.001	24.8***		0.958	27.5***	
	(2.00)			(-3.15)***				(2.63)			(-1.64)*		
	0.973		62.7***	0.439		98.9***		0.987		71.2***	0.493		104***
	(-2.90)			(-7.54)***				(-3.11)			(-6.88)***		
y^{*E}	1.000	5.61***		0.969	19.3***		y^{*}	1.000	6.81***		0.995	14.9***	
	(3.40)			(-2.11)**				(2.20)			(-0.32)		
	0.987		37.2***	0.306		104***		0.960		87.9***	0.523		149***
	(-2.04)			(-8.40)***				(-2.90)			(-7.57)***		
y^{*E*}	1.001	6.56***		0.988	13.5***		p^{*}	1.000	20.8***		0.946	23.2***	
	(2.65)			(-0.86)				(0.60)			(-1.89)*		
	0.960		42.0***	0.458		159***		0.995		103***	0.642		135***
	(-2.84)			(-7.77)***				(-2.29)			(-5.68)***		
y	1.001	5.09***		0.969	15.3***		i	0.992	41.0***		0.336	61.9***	
	(2.70)			(-1.38)				(-0.72)			(-6.01)***		
	0.955		58.3***	0.431		198***		0.867		97.6***	0.082		179***
	(-1.73)			(-7.41)***				(-3.42)*			(-9.30)***		
p	1.000	35.4***		0.868	36.7***		$i-\dot{p}$	0.964	33.6***		0.423	60.1***	
	(1.60)			(-2.78)***				(-1.50)			(-5.76)***		
	0.993		116***	0.369		148***		0.770		103***	0.072		163***
	(-1.69)			(-7.84)***				(-4.45)***			(-9.48)***		
m	1.003	12.1***		0.612	23.6***								
	(2.03)			(-4.90)***									
	0.958		57.7***	-0.031		243***							
	(-2.66)			(-10.4)***									

The estimated first-order autoregressive parameter a_1 (see equations (S1) - (S2) of Table 4.1) is given in the column t and the t value of this estimate in parentheses. The null hypotheses are: t; $a_1=1$, F4 and F5; $a_1=a_2=1$, $a_3=0$ (see equations (S1) - (S2) of Table 4.1). The Dickey-Fuller-Hasza tests are based on regressions with twelve lags of the differenced variable under investigation. Dummy variables taking the value 1 for the following variables (dates) and 0 otherwise have been employed in the test regressions: y^{*E*}, y^{*} (1980M5) and y (1971M2 - 1971M3). Critical values for the tests are tabulated by Fuller (1976) and Hasza & Fuller (1982). Rejection at the 10,5 and 1 per cent level of significance is indicated by *, ** and ***, respectively.

The tentative character of the above results should be emphasized; the low power of tests for unit roots is a widely recognized fact. This may also have showed up in the few mixed results found, in particular when comparing results from different subperiods. Furthermore, it is not clear how well these simple tests are suited for analyzing time series with complicated data generating processes, such as the exchange rate or the price of oil. While it is not possible to draw strong conclusions from our analyses, it appears that most of the variables are I(1). As there may be exceptions to this, such as the foreign real rate of interest (which may have been I(0) before 1973 but I(1) afterwards), sensitivity analyses seem warranted. As a by-product of the tests for the order of integration, we may note that the results suggest that the price of crude oil and the domestic price level have not been bivariately cointegrated with the foreign price level.

The results of the tests for unit roots in seasonal variables can be summarized easily, since the evidence is uniform and, relatively speaking, strong. All seasonal variables are I(1), with one, or possibly two, exceptions. The obvious exception is the (domestic) real interest rate, which, as its foreign counterpart, seems to be I(0). There is also some evidence that the nominal (domestic) interest rate could be I(0); one of the tests rejects the occurrence of a unit root practically at the five per cent level of significance. With respect to seasonal unit roots, the evidence is overwhelming; the occurrence of stochastic seasonality can be rejected in each and every case at the one per cent level of significance.

The analyses based on subsamples documented in Tables A4.4 - 6 of Appendix 4 tend to corroborate the full sample results. The fiscal policy stance proxy turned out to be I(1) with strong deterministic seasonal elements. Some signs of stochastic seasonality were recorded for y, y^{*E*} and g toward the end of the full sample period. These results may be suspect because of the short time span available in subsample analyses, however. The degree of integration for some of the output variables differs uncomfortably much across subsamples according to the tests, which may also attest to the perils of putting too much emphasis on these subsample results. A prime example of the hazards for inference of a short period is offered by the real rate of interest, which turned out to be I(0) when the maximum number of observations was analyzed, but appeared to be I(1) when only subsample information was allowed to bear on the results. In addition, some sensitivity of the results with respect to the alternative hypothesis could be detected. Hence, we again have to emphasize the tentative character of these results.

The seasonality of the output variable y deserves some further comments. Seasonal movements in Finnish real total industrial production are very pronounced, both in absolute terms and relative to seasonal movements in output in many other western economies. In fact, 89.9 per cent of the movements in the growth of Finnish real total industrial production can be attributed to deterministic seasonality during our sample period. Output growth slows down in June, and output takes a pronounced dip in July. In August and September output recuperates, but slows down somewhat again in December. Over our sample period, very little change in this seasonal pattern has occurred.

Despite the overall good accuracy of the fixed seasonal mean approximation, a close scrutiny of the seasonal movements of y reveals some signs of varying seasonal patterns. This shows up in the slowdown in July, but, in particular, it shows up in March and April. The nonconstant quasi seasonal movements in the early spring arise from the contract setting system in the Finnish labor market, which, in a rather unsystematic manner, has resulted in strikes in the spring. A fuller description of this phenomenon and the collective wage bargaining system in Finland is provided in Saikkonen & Teräsvirta (1985). We do not attempt to remove such stochastic "seasonals", however, since strikes represent one type of shock we will want to identify in the subsequent chapters. Also, in recent years labor market contracts have been made for two years ahead, removing this phenomenon from the realm of seasonality.

The empirical evidence on the degree of integration of key Finnish macroeconomic variables is in broad agreement with international and other Finnish evidence. Huizinga (1987) finds unit roots in y^{*U} and all the components of y^{*E}, y^{*E*} and y^*, and Gerlach (1988) and Wasserfallen (1988) reach the same conclusions using data from France, Germany, the U.K. and the U.S. Nelson & Plosser (1982) and Schwert (1987) find that U.S. counterparts to y, p, m, and i are I(1), and using a nonparametric approach Perron (1988) reaches similar conclusions. International evidence presented in Stulz & Wasserfallen (1985) and Wasserfallen (1986, 1988) gives the same picture, with the additional finding that international counterparts to e also are I(1).

Wallius (1983) examines Finnish data, and finds that y, p, and i are I(1). Johansen & Juselius (1989) find unit roots in p and m, but reject the null hypothesis of a unit root in i when i is measured by the Bank of Finland's call money rate of interest.[3] The evidence on seasonal unit roots in Stulz & Wasserfallen (1985) and Wasserfallen (1986) is mixed; the authors are not able to reject the null hypothesis of a seasonal unit root in about half of the cases they consider. Barsky & Miron (1989) find little evidence of stochastic seasonality in a large U.S. data set using unit root and spline function models.

4.3.2 Tests for Cointegration

We now move on to investigate whether some of the documented unit roots are common to two or more series. Since the seminal paper of Granger (1981) many tests for the null hypothesis of non-cointegration have been developed. The most widely used and studied tests are the residual based tests proposed by Engle & Granger (1987). These tests are closely related to the tests for unit roots used above. The residual based tests of Phillips (1987) are also frequently used (the augmented Dickey-Fuller test and the

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[3]Whether gross domestic product contains a unit root or not has not been empirically settled in the literature. An illustration of this is the ongoing debate concerning U.S. gross domestic product; see Christiano & Eichenbaum (1989) and the references cited therein.

Phillips Z_t test are asymptotically equivalent (Phillips & Ouliaris (1988a)).

An alternative strand of the testing for cointegration literature focuses on the matrix π in model (4.1) or, alternatively, on a diagonalization $P^{-1}\pi P = D$ of this matrix. Stock & Watson (1988) suggest testing for the number of common (stochastic) trends by examining how close to zero the largest eigenvalues of D are. Johansen (1988b, 1989) works directly with model (4.1), and constructs the maximum likelihood estimator of the space of cointegrating vectors $\underset{\sim}{\alpha}$ and likelihood ratio test statistics for the dimension r of this space. Essentially, an estimate of $\underset{\sim}{\alpha}$ is obtained by applying canonical analysis to the variables $(1-L)\underset{\sim}{y}_{t+1}$ and $\underset{\sim}{y}_t$. The common trends approach has been applied by Huizinga (1987), King et al. (1987), Giovannini (1988) and Shapiro & Watson (1988), and the maximum likelihood approach by Hall (1989) and Johansen & Juselius (1989).

A third approach to testing for cointegration has recently been proposed by Phillips & Ouliaris (1988b). They propose the use of principal components methods, and suggest testing whether or not the spectral density matrix Σ of the forcing process $\underset{\sim}{\varepsilon}_t$ has negligible eigenvalues. If there exists r distinct cointegrating vectors Σ will have r eigenvalues equal to zero. The authors also report some empirical applications.

We will employ the maximum likelihood approach due to Johansen (1988b, 1989). This procedure is chosen primarily to overcome the inherent major weakness of the two-step procedure of Engle & Granger (1987) involving the non-uniqueness of cointegration vectors when more than two variables are considered. More specifically, the two-step procedure would be applicable only if one knew a priori that none of the cointegration vectors involve precisely the same set of variables. Moreover, the two-step estimator is known to exhibit a considerable degree of small sample bias (Banerjee et al. (1986), Stock (1987)). Since cointegration implies causation, the need to consider higher dimensional systems than bivariate systems should be apparent (notwithstanding the fact that Granger (1981) speaks about bivariate relationships only and the vast bulk of empirical work beginning with Engle & Granger (1987) has dealt with pairwise relationships only).

When carrying out the tests for cointegration, we will follow the research strategy outlined in section 3.3 of chapter 3. This entails focusing on two categories of models, one comprising output variables only, and one consisting of both real and nominal foreign and domestic variables. With regard to the choices of variables for these models, we will anticipate the results of the analyses in chapters 5 - 7 in the sense that we take as given the basic variable vectors entering the two categories of models. Sensitivity analyses are carried out to strengthen the reliability of the choices. As in the case of the univariate tests for unit roots, the analyses are carried out for the full sample period as well as for the three subsample periods. The empirical results concerning subperiods will again be relegated to an appendix.

We begin by considering the output variables y^{*U}, y^{*E}, y^{*E*} and y entering the model of chapter 5. The empirical evidence for the full sample period is presented in Table 4.4, and the subperiod results are presented in Table A5.1 of Appendix 5. With regard to the full sample results, the tests for the number of cointegrating vectors unanimously indicate the presence of two long-run equilibrium relationships. Hence, international output developments during our sample period have been tied together by equilibrating forces guaranteeing that output in one country never drifts too far away from the output of other countries. Given the differences in the level of output across economies, this also implies that, on average, any one economy has not become richer <u>relative</u> to another economy during our sample period. These results on cointegration among output in different economies are thus consistent with the stylized facts on economic growth documented and analyzed by e.g. Lucas (1988).

TABLE 4.4

Maximum likelihood estimation results for a VAR model for the vector $\underset{\sim}{y}_t = (y_t^{*U}\ y_t^{*E}/y_t^{E*}\ y_t)'$ using data from the period 1960M1 - 1988M8

$\underset{\sim}{y}_t = (y_t^{*U}\ y_t^{*E}\ y_t)'$ $\qquad\qquad$ $\underset{\sim}{y}_t = (y_t^{*U}\ y_t^{E*}\ y_t)'$

Eigenvalues

(0.090 0.055 0.010) $\qquad$ (0.078 0.046 0.008)

Eigenvectors

$$\begin{bmatrix} -56.2 & -8.82 & 63.2 \\ 6.71 & -15.2 & -16.7 \\ 7.71 & 7.99 & -0.737 \end{bmatrix} \begin{bmatrix} -76.9 & 2.59 & 27.1 \\ 22.9 & -33.4 & -7.18 \\ 7.19 & 8.61 & -0.200 \end{bmatrix}$$

Loadings x 10^3

$$\begin{bmatrix} 0.122 & 0.180 & -0.242 \\ -1.83 & 1.36 & -0.183 \\ -8.86 & -8.62 & -0.747 \end{bmatrix} \begin{bmatrix} 0.237 & 0.071 & -0.213 \\ -3.66 & 1.24 & -0.133 \\ -10.9 & -5.58 & -0.713 \end{bmatrix}$$

Tests for the number of cointegrating vectors

H_2	-2ln(Q)	trace	maximum eigenvalue	-2ln(Q)
r<2	3.30	8.08	8.08	2.61
r<1	22.8	17.8	14.6	18.7
r=0	54.9	31.3	21.3	46.5

Normalized eigenvectors, loadings, and corresponding long-run coefficient matrices when the models are restricted to have two cointegrating vectors

$$\begin{bmatrix} -7.28 & -1.10 \\ 0.870 & -1.91 \\ 1.00 & 1.00 \end{bmatrix} \begin{bmatrix} 0.001 & 0.001 \\ -0.014 & 0.011 \\ -0.068 & -0.069 \end{bmatrix} \begin{bmatrix} -0.008 & -0.002 & 0.002 \\ 0.091 & -0.033 & -0.003 \\ 0.574 & 0.072 & -0.137 \end{bmatrix} \begin{bmatrix} -10.7 & 0.301 \\ 3.19 & -3.88 \\ 1.00 & 1.00 \end{bmatrix} \begin{bmatrix} 0.002 & 0.001 \\ -0.026 & 0.011 \\ -0.078 & -0.048 \end{bmatrix} \begin{bmatrix} -0.018 & 0.003 & 0.002 \\ 0.285 & -0.125 & -0.016 \\ 0.825 & -0.063 & -0.127 \end{bmatrix}$$

In addition to the variables in $\underset{\sim}{y}_t$, the model includes a constant, 11 seasonal dummy variables and a strike dummy taking the value 1 in 1971M2 - 1971M3 and 0 otherwise. The test statistic for the number of cointegrating vectors is denoted -2ln(Q), and critical values (95 per cent quantiles) have been tabulated by Johansen & Juselius (1989).

Scrutinizing the estimates of the cointegration, weighting, and long-run coefficient matrices presented in Table 4.4 reveals some interesting particulars about the cointegrating relationships. The long-run equilibrium relationships carry least weight in the case of the U.S. economy and most weight in the case of the Finnish economy. We take this to indicate that the U.S. economy has been the most closed and the Finnish economy the most open. There are differences in the results depending on which measure of the intermediate economy is used, but the general flavor of the results does not depend on this choice.

Turning to the subsample results presented in Table A5.1 of Appendix 5, the following findings emerge. During 1960 - 1972, the U.S. economy, the intermediate economies and the Finnish economy have grown on independent and diverging long-run paths. This absence of any long-run equilibrium implies that relative changes in real income took place. Our maximum likelihood estimates apart, the absence of a long-run equilibrium relation between our output series is also evident from a mere visual inspection of the series (see Figure 5.2 of chapter 5).

During the post-Bretton Woods period as a whole, and during the 1980s separately, one long-run equilibrium relationship tying the output of major western economies together is found. Visual inspection of the output series also supports the notion that output developments have been tied together from the early 1970s onward (see Figure 5.2 of chapter 5). Scrutinizing the estimated cointegrating and weighting vectors and the long-run coefficient matrices gives the same impression of the long-run equilibria as the full sample analysis. It also suggests that the importance of the international output equilibrium relationship has increased over time. Moreover, Finnish long-run output movements have become more dependent on output developments in other western economies than on corresponding developments in the U.S. economy.

Before turning to inference within the other basic frame of reference of our study, a few words about the number of cointegrating relationships across periods are in order. How should we interpret the fact that no cointegration was found during 1960 - 1972, one cointegrating relationship was detected during 1973 - 1988, and two cointegrating relationships were unraveled when the full 1960 - 1988 sample was used? Essentially, it should be recognized that it is bound to be difficult to base inference on the long-run behavior of the economic system on data covering less than 30 years. While the recent developments in inference on cointegration certainty in principle allows one to infer something about long-run equilibria from finite data sets, the power and small sample properties of the tests are largely unknown. Hence, our results are tentative rather than conclusive.[4]

Our tentative interpretation of the results on the number of cointegrating vectors is the following: during the 1960s and the early 1970s there were at best only weak forces making for similarity of long-run output developments in the countries of our sample. This changed with the gradual opening up of the economies, and it is also possible that the change in exchange rate regime rendered possible greater equilibration of international output developments. Whether there exists one or two cointegrating vectors describing the long-run equilibrium in place since the early 1970s is not clear; the full sample result - which should be the most powerful - supports the notion of two relationships.

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[4]The existence of outliers in the output series may raise caveats to our findings, since the Johansen procedure assumes normally distributed disturbances. However, following Phillips (1987), it seems likely that the Johansen tests and corresponding asymptotic distributions also apply in the case of nonnormal (and possibly autocorrelated) disturbances. Kunst (1989) presents some empirical evidence which supports this conjecture. In any case, outliers will not affect the consistency of the estimates of cointegrating vectors (but they may affect the estimates of weighting vectors).

We turn to the collection of variables entering the model of chapters 6 and 7. The estimation results using the full sample are presented in Table 4.5, and results for subsamples are shown in Table A5.2 of Appendix 5. Irrespective of the period analyzed, multivariate cointegration is found in the models consisting of real and nominal, foreign and domestic variables. When the full sample is analyzed, the long-run equilibrium does not stand out as a marked feature of the model, however. The first of two cointegrating vectors enters the model with essentially zero weight whereas the second cointegrating vector carries nonnegligible weight only in the equation for output.

The estimated cointegration vectors seem hard to interpret in terms of long-run elasticities (see the normalized eigenvectors). In fact, the parameter estimates may well be void of any interesting economic interpretation, as pointed out by Lütkepohl & Reimers (1989a,b). This is because we are working with a simultaneous, dynamic system of equations in which the other variables cannot be regarded as fixed. Furthermore, we have two relationships characterizing the equilibrium, thereby making it difficult to interpret the parameters of just one of the relationships.

A perhaps more revealing presentation of the information contained in the estimated cointegrating vectors is to plot the deviations z_t from the equilibrium paths. In Figure 4.1 we have plotted the equilibrium error corresponding to the second cointegrating vector - the one vector that enters the equation for output with a possibly nonnegligible weight - against time.[5] Since the devaluations of the Finnish markka in 1967M10 (23.8 per cent), 1977M4 (5.7 per cent), 1977M9 (3.0 per cent), 1978M2 (8.0 per cent), 1982M10 (10.0 per cent) and 1986M5 (2.0 per cent) have affected the deviations from the long-run equilibrium in a systematic way, we have added vertical lines at the devaluation dates in Figure 4.1.

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[5]In general, one would expect that the first cointegrating vector - which is most correlated with the endogenous variable - would be of most interest. However, in our particular case the Johansen procedure indicates the presence of a second cointegrating vector, and attaches higher weights to this vector.

TABLE 4.5

Maximum likelihood estimation results for a VAR model for the vector $\underset{\sim}{y}_t = (y_t\ p_t\ m_t\ i_t^*\ e_t\ p_t^*)'$ using data from the period 1960M1 - 1988M8

Eigenvalues

(0.296, 0.096, 0.046, 0.019, 0.014, 0.000)

Eigenvectors

$$\begin{bmatrix} 2.46 & -12.7 & 5.12 & 3.68 & -1.51 & -1.90 \\ -3.44 & -2.79 & 0.874 & -19.0 & 31.0 & 1.01 \\ -10.4 & 6.09 & -7.05 & 5.65 & 4.23 & 1.51 \\ -47.8 & 11.1 & -14.4 & -28.5 & -11.6 & -11.9 \\ -1.35 & 16.6 & 12.4 & -3.32 & -3.14 & -0.147 \\ 21.6 & -4.21 & 4.71 & 11.8 & -42.3 & -0.589 \end{bmatrix}$$

Loadings x 10^3

$$\begin{bmatrix} 2.61 & 12.7 & -0.554 & -0.846 & 0.843 & -0.443 \\ -1.47 & -0.113 & -0.346 & 0.355 & -0.284 & -0.042 \\ 0.098 & -1.84 & 5.16 & -0.777 & -1.17 & -0.302 \\ 0.325 & 0.339 & 0.344 & 0.440 & 0.220 & 0.004 \\ 1.03 & -1.59 & -0.537 & -0.180 & 0.756 & -0.155 \\ -1.49 & 0.004 & 0.115 & -0.040 & 0.102 & -0.001 \end{bmatrix}$$

Tests for the number of cointegrating vectors

H_2	-2ln(Q)	trace	maximum eigenvalue
$r \leq 5$	0.149	8.08	8.08
$r \leq 4$	5.07	17.8	14.6
$r \leq 3$	11.5	31.3	21.3
$r \leq 2$	27.7	48.4	27.3
$r \leq 1$	62.4	70.0	33.3
$r = 0$	182	n.a.	n.a.

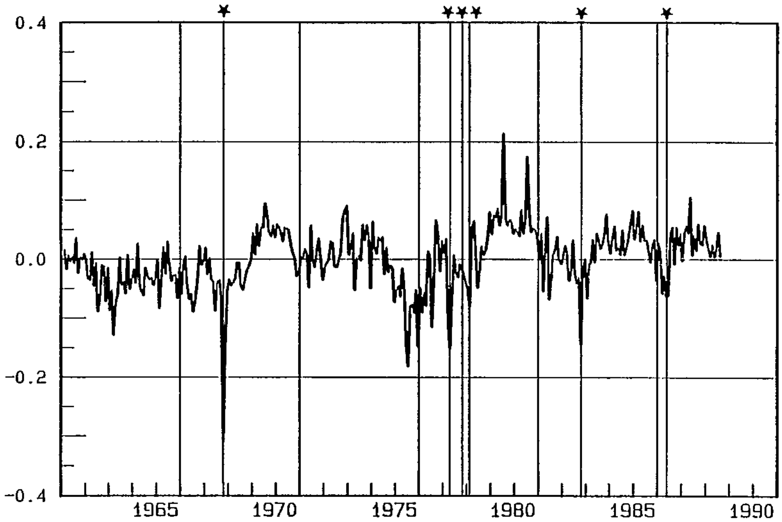
Normalized eigenvectors, loadings, and the corresponding long-run coefficient matrix when the model is restricted to have two cointegrating vectors

$$\begin{bmatrix} 1.00 & 1.00 \\ -1.40 & 0.220 \\ -4.21 & -0.481 \\ -19.4 & -0.877 \\ -0.549 & -1.31 \\ 8.78 & 0.333 \end{bmatrix} \begin{bmatrix} 0.006 & -0.161 \\ -0.004 & 0.001 \\ 0.000 & 0.023 \\ 0.001 & -0.004 \\ 0.003 & 0.020 \\ -0.004 & -0.000 \end{bmatrix} \begin{bmatrix} -0.155 & -0.045 & 0.050 & 0.017 & 0.208 & 0.003 \\ -0.002 & 0.005 & 0.015 & 0.069 & 0.000 & -0.031 \\ 0.024 & 0.005 & -0.012 & -0.025 & -0.031 & 0.010 \\ -0.003 & -0.002 & -0.001 & -0.001 & 0.005 & 0.006 \\ 0.023 & 0.001 & -0.020 & -0.067 & -0.028 & 0.029 \\ -0.004 & 0.005 & 0.015 & 0.071 & 0.002 & -0.032 \end{bmatrix}$$

In addition to the variables in $\underset{\sim}{y}_t$, the model includes a constant, 11 seasonal dummy variables and a strike dummy taking the value 1 in 1971M2 - 1971M3 and 0 otherwise. The test statistic for the number of cointegrating vectors is denoted -2ln(Q), and critical values (95 per cent quantiles) have been tabulated by Johansen & Juselius (1989).

FIGURE 4.1

Deviations from an estimated long-run equilibrium relationship ($y_t + 0.220p_t - 0.481m_t - 0.877i^*_t - 1.31e_t + 0.333p^*_t$) in the vector $\underset{\sim}{y}_t = (y_t\ p_t\ m_t\ i^*_t\ e_t\ p^*_t)'$ based on data from the period 1960M1 - 1988M8



Note: Vertical lines marked with a star indicate dates of devaluations of the Finnish markka.

A striking feature of the estimation result displayed in Figure 4.1 is how closely the deviations from the steady state follow the historical Finnish business cycle. We also note that devaluations have moved the economy back to equilibrium within one year after a devaluation. However, the economy has typically not settled down at the equilibrium, but it has, in a sense, overshot the equilibrium during the second and third year after a devaluation. After that, the economy has hovered about the equilibrium for an irregular period of time.

Turning to the subsample results documented in Table A5.2 of Appendix 5, the following remarks can be made. Two cointegrating vectors are found in the models for the periods 1960 - 1972 and 1980 - 1988, respectively. Three cointegrating vectors are found in the model for the period 1973 - 1988. During the Bretton Woods period, the foreign interest rate does not seem to have mattered for the domestic steady state. After the breakdown of the Bretton Woods arrangement, the long-run equilibrium appears to have had the greatest influence on the determination of output and money. Long-run equilirium constraints appear to have mattered only very little in the 1980s.

The full sample results remain practically unaltered when the basic model is varied. Hence, e.g. the inclusion of foreign output or the price of oil, or the exclusion of the foreign price level, still yielded two cointegrating vectors mattering most for the determination of domestic output. Subsample results differed more both in terms of the number of cointegrating vectors and in terms of estimated coefficients. Adding one variable to the basic models usually increased the number of cointegrating vectors, while removing a variable from the basic model usually reduced the number of cointegrating vectors. Multivariate cointegration was found in all cases, however.

Lastly, it may be of interest to compare the results on cointegration generated by the Johansen procedure with corresponding results produced with the Engle - Granger procedure and reported in Starck (1988a). In short, only very weak indications of bi- and multiple cointegration were detected using the two-step approach. DW-type tests signalled many bivariate cointegrating relationships, and ADF tests showed some signs of multiple cointegration, but these results did not stand up to closer scrutiny. A closer look showed low R^2s for the static cointegrating regressions, sensitivity of the results with respect to the choice of dependent variable in the static cointegrating regressions and the poor power of the DW test. The DW test is sensitive to the data generating process (Engle & Yoo (1987)) and to the occurrence of moving average components in the series (Molinas (1986)), and it has poor small sample performance (Banerjee et al. (1986)).

While the results using the Engle - Granger procedure differ from the results using the Johansen procedure, the results are labelled highly tentative. Even though the maximum likelihood approach certainly should be expected to outperform the two-step approach, the results presented in this chapter should be considered tentative. This mainly stems from the fact that the small sample properties of the maximum likelihood approach are unknown. There may also be some power concerns due to the high dimensionality of some of our models. Moreover, the maximum likelihood procedure used above is not applicable if some of the data were I(2) (contrary to the evidence presented in section 4.3.1). Nevertheless, we regard the above comparative study as one piece of evidence cautioning against the use of the up to now widely applied two-step approach to cointegration inference.

In addition to Starck (1988a), the occurrence of multivariate cointegration in Finnish data has been studied by Johansen & Juselius (1989). Using the maximum likelihood approach and quarterly data, Johansen & Juselius (op.cit.) find three cointegration vectors in a system consisting of the variables GDP, Δp, m and i. By contrast, no multivariate cointegration is found by Giovannini (1988), who employs Stock - Watson tests in a system of international equivalents to y, p, m, i and e, or by Stock & Watson (1989), who analyze U.S. counterparts to y, p, m and i. Shapiro & Watson (1988) reach the same conclusion using the same approach on U.S. equivalents to y, Δp, i and labor supply, as, too, does Huizinga (1987) with a system consisting of international equivalents to y and e^r.

4.4 Conclusions

The aim of this chapter has been to provide empirical evidence on the joint integration properties of the data of our study. Such evidence is needed to indicate which are the proper stationarity inducing filters, how seasonality should be dealt with, what long-run equilibria affecting the specification of the subsequent models exist, and what distribution theory should be relied on when conducting tests of the subsequent models. The empirical evidence has intrinsic value from a descriptive point of view as well. Our analyses may also be interesting as results based on the Johansen procedure are contrasted with results based on the Engle - Granger procedure.

Univariate unit root properties of the data are addressed through Dickey - Fuller tests, and issues of common unit roots are examined in the first place by means of the Johansen procedure. All tests are performed for the full sample period and for three subperiods. With regard to the order and nature of integration, the following tentative results emerged. All but one variable belong to the difference stationary class of time series processes, with the order of integration in all these cases being one. Thus, differencing should be employed to induce stationarity of first moments. The (foreign and domestic) real rate of interest does not seem to display unit root nonstationarity. Seasonality was found to be essentially a deterministic phenomenon. Hence, seasonality should be accounted for by deterministic dummy variables.

With regard to common unit roots - cointegration - the following results were obtained. Multivariate cointegration was detected both in the model comprising only output variables and in the models comprising both real and nominal variables. Evidence of the long-run equilibria being characterized by more than one cointegrating vector was documented. The estimated weighting and long-run coefficient matrices suggested that steady state considerations may have affected developments to a very small extent, however. Devaluations of the Finnish markka were found to have affected deviations from the long-run equilibrium involving foreign and domestic real and nominal variables in an intuitively appealing way.

Some caveats to our choices and conclusions should be mentioned. The choice of monthly data implies that the subsequent analysis will be carried out partly with variables that may contain nonnegligible amounts of measurement error and noise. This may blur causal relationships. Secondly, the use of fixed seasonal means should be viewed as an approximation to reality, which in some cases may exhibit stochasticity. With regard to the unit root tests, the low power of these tests must be borne in mind. Hence, our results on the existence of unit roots are tentative. Finally, the maximum likelihood approach to testing for cointegration also relies on asymptotic results while having unknown small sample properties. Also, the power of these tests decreases with the dimension of the model, and we have partly been working with large models.

5 A WOLD CAUSAL CHAIN MODEL

The aim of this chapter is to disentangle how fluctuations in U.S. output, fluctuations in the output of other major western economies, and fluctuations in Finnish output are related, and to empirically analyze the long- and short-run character of this association. These issues are of relevance to the understanding of the global economy, to the formulation, conduct and effectiveness of economic policy, to the insulating properties of different exchange rate arrangements, and to output forecasting. We focus solely on the output of an economy as a whole, and operationalize this on the monthly level as real total industrial production. Industrial production constitutes only a part of gross national product, but from the point of view of our aim it represents precisely the interesting part, since it is the most business-cycle-susceptible part of gross national product.

5.1 Theoretical Considerations

The dominance of the U.S. economy has been recognized for many decades, and it still seems true that "when the U.S. economy sniffles the remainder of the western world sneezes". In other words, U.S. economic activity as an approximation seems to be causally prior in the Granger sense to economic activity in other economies.[1] Given that the U.S. economy has essentially been a closed one during our sample period and that its size is five times that of the German economy, the importance of the U.S. economy is paramount.

While output fluctuations in the U.S. are important for all economies, other economies also matter for small open economies.

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[1]This has been documented by Huth (1985) and Stulz & Wasserfallen (1985). See also the discussion and references in Layton (1987). With respect to interest rate linkages some signs of causal relationships from European interest rates to U.S. interest rates have been documented (Hartman (1984), Kirchgässer & Wolters (1987) and Kool & Tatom (1988)). However, the size of these possible reverse causal relationships, and in particular the size of their potential impact on the U.S. economy, should not be able to distort the one-way causality to any substantive extent.

These other economies also act as filters of the impulses from the U.S. economy. It is important to note that these filters will, in general, alter the size and character of the original impulse. They may also imply a sequence of secondary impulses. Furthermore, the other economies share with the U.S. economy the dominant role with respect to small open economies; the other economies are causal to the small open economy. For example, the Finnish economy is more open in terms of e.g. imports to GDP than the German economy or the U.S. economy, and the German economy is 11 times as big, and the U.S. economy 73 times as big, as the Finnish economy (figures refer to 1987). Of course, the qualifier "small" is given precisely to economies that essentially display no feedback on the world economy.

What emerges then is a hierarchial model of the world economy. In the theoretical literature a model of this type has been analyzed by Bhandari (1987) among others. Bhandari considers the relationship between economic size of a country and vulnerability to external disturbances. He presents numerical simulations of a three-country 20-equations model, and finds that under floating exchange rates a small open economy is better insulated against external disturbances than a medium-sized country. The problem - which Bhandari readily acknowledges - with such models is that they incorporate far too many restrictive assumptions to yield interesting results for any particular country. In the model of Bhandari (op.cit.) the economic structure in all three economies is identical down to parameter values, and the degree of openness of the two smaller economies is identical.[2] The realism of the model could also be affected by the complete absence of rigidities; all prices are fully flexible and continuous equilibrium prevails.

The relationship between output fluctuations in different countries has been studied empirically as well. The hierarchial three-economy

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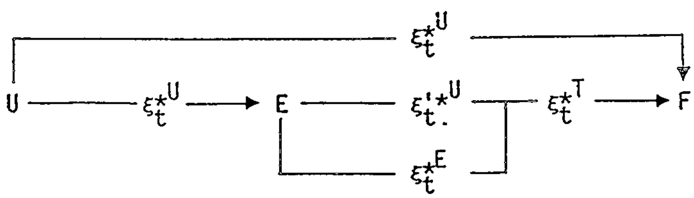
[2]Nandakumar (1985) demonstrates by numerical simulations of single-country models that the structure of a small open economy matters for macroeconomic adjustment to supply side shocks. Using a three-country, two-step hierarchial model, he also shows that it is important to take the global economy into account when analyzing the effects of shocks on a small open economy.

structure has not been employed, however. Typically, descriptive analyses in the time domain (Virén (1985) and Baxter & Stockman (1989)) and the frequency domain (Andresen & Everaert (1987) and Gerlach (1988)) and traditional time series studies (Huth (1985) and Layton (1987)) have been carried out. Bivariate VAR models have been employed in the study of the relationship between national and regional output fluctuations (Sherwood - Call (1988) and Wörgötter (1989)). VAR models designed primarily for other purposes but including the output of different countries have also been constructed (Ahmed et al. (1989)).

The (contemporaneous) structure of the model we will be using is illustrated in Figure 5.1.

FIGURE 5.1 Schematic model of the transmission of output impulses



In Figure 5.1 the U.S. economy is denoted U, the intermediate economy (more precisely, the aggregate of economies) E, and the Finnish economy F. An output impulse emerging from the U.S. at time t is denoted ξ_t^{*U}. It is defined as the unexpected part of output growth, and it is assumed to be expected mean zero white noise. This growth impulse is observed in E and F simultaneously, having a different impact depending on the openness, structure and size of the recipient economy. Let $\xi_t^{'*U}$ denote the unanticipated change in the growth of economy E brought about by the unanticipated change in the growth of economy U. In addition, economy E will be emitting a growth impulse ξ_t^{*E} which is due solely to unforeseen domestic factors specific to economy E. Thus, observed at any point in time, economy F is, in effect, hit by three distinct disturbances.

The impulse ξ_t^{*U} will most probably induce a sequence $\{\xi_{t+i}^{\prime *U}$: $i = 1, 2, 3, \ldots\}$ of further disturbances because of costs of forecasting and uncertainty about economic policy reactions in E to ξ_t^{*U}. Assuming that expectations are rational and that economic agents have symmetric loss functions, the sequence $\{\xi_{t+i}^{\prime *U}$: $i = 0, 1, 2, \ldots\}$ will consist of mean zero white noise disturbances. Observed at any point in time, economy E will then be emitting three distinct shocks; purely domestic shocks, shocks induced by the U.S. impulse and shocks due to domestic discretionary economic policy and the less-than-perfect forecastability of domestic output. Denote this composite shock ξ_t^{*T}.

The contemporaneousness of ξ_t^{*U} and $\xi_t^{\prime *U}$ arises as a consequence of forward-looking behavior of economic agents. As agents in F know the propagation mechanism of E, they can calculate the effect of ξ_t^{*U} on E the moment ξ_t^{*U} is observed (using (3.20)). The difference between the pre-shock and post-shock predicted output is $\xi_t^{\prime *U}$. Additional effects of ξ_t^{*U} will be observed because of the string of impulses $\{\xi_{t+i}^{\prime *U}$: $i = 1, 2, 3, \ldots\}$ emerging from E caused by ξ_t^{*U}. This follows from the fact that it is rational to forecast the output of E only to some degree of precision determined by nonzero costs of acquiring and processing information about the propagation mechanism of E. Moreover, it is rational to change one's prediction about economic activity in E only when the predicted effect on E is "big enough". The economic policy part of the propagation mechanism of country E will cause additional effects to the extent that economic policy is not rule-based, since unpredictable policy constitutes disturbances.[3]

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[3] In addition to a possible Peso problem the economic policy part of the propagation mechanism may cause an observational equivalence problem. Comovement of output in the three economies could in principle arise if the countries have the same policy objectives and if economic policy is effective in the desired way. Thus the comovement of output in E and in F could come from shared policy objectives rather than from structural relationships between E and F. However, it can be argued that neither of the two conditions required for policy-induced comovement are likely to be fulfilled.

The model described above can be stated formally using the VAR representation (3.13). Define $\underset{\sim}{y}_t = (y_t^{*U}\ y_t^{*E}\ y_t)'$ where y_t^{*U}, y_t^{*E} and y_t denote total output in economies U, E and F, respectively. Then $\underset{\sim}{\zeta}_t = (\zeta_t^{*U}\ \zeta_t^{*T}\ \zeta_t^{F})'$ is the reduced form representation of the unforeseen growth of economies U, E and F, respectively. Impose (contemporaneous) recursivity upon the representation by making Γ_0 lower triangular. The resulting model comprises the features brought up in the preceding discussion.

5.2 Empirical Evidence

5.2.1 Estimation Results

Data cover the following countries: Belgium, Finland, France, Germany, Italy, Japan, the Netherlands, Sweden, the U.K. and the U.S. We will for convenience be referring to this set of economies as the world economy. The choice of countries - in addition to Finland and the U.S. - reflects the size of the economy and the country's importance for Finnish exports. We also need to specify the intermediate economy in terms of weights.[4] We have chosen to employ two very different weighting schemes, one based on the size (GDP) of the economy and the other based on the country's importance for Finnish exports. The former aggregate is denoted y_t^{*E} and the latter y_t^{*E*}.[5]

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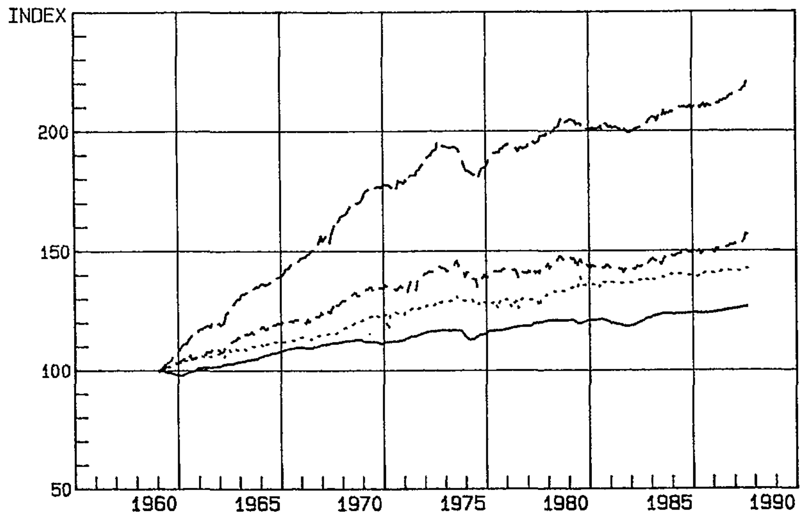
[4]The empirical implementation of our hierarchial three-country setup highlights why economy E is being considered an aggregate of economies. The U.S. output impulse will cause complex interactions between all countries, and, in general, the new equilibrium will imply changes in economic performance in all countries. If we want to isolate impulses emerging from one component of E only, the effects of all other economies would have to be controlled for. Even if the degrees of freedom problem could be overcome, we would have to specify how the "intermediate size" economies interact. That interaction is not well known, and specifying it would be worthy of a sizeable study of its own.

[5]In the spirit of the extreme bounds analysis of Leamer (1983) and Leamer & Leonard (1983), we choose schemes yielding markedly different rankings and having big differences in weights. More comprehensive weighting schemes, ranking countries e.g. in terms of their contribution to world trade and to international financial flows, typically yield more evenly distributed weights (Masson & Blundell - Wignall (1985)).

FIGURE 5.2 Seasonally adjusted real total industrial production in the U.S., aggregates of intermediate economies, and Finland 1960M1 - 1988M8 (1960M1 = 100)



Notes: ——— y_t^{*U} — — — y_t^{*E} —— —— —— y_t^{*E*} - - - - - - y_t

The time series of the variables y_t^{*U}, y_t^{*E}, y_t^{*E*} and y_t are graphed in Figure 5.2. Note that while seasonally adjusted series are plotted in the figure for expositionary clearness, the subsequent analysis is, as mentioned before, carried out using seasonally <u>un</u>adjusted data. The nonstationarity of the output series stands out in Figure 5.2, as does the tendency of the series to drift apart during the Bretton Woods era and move together during the post-Bretton Woods era. These findings are congruent with the findings of chapter 4 according to which the output series were not cointegrated during the former period but were cointegrated during the latter period.

The results in the previous chapter on the joint integrating properties of our output series imply that the VAR model should be estimated in differenced form including level (error correction) terms and deterministic seasonal dummy variables. During the Bretton

Woods subsample, no error correction terms need be included, and during the other subperiods one level term of each variable should be included. Sensitivity analyses of these specifications should be carried out, however.

We begin by determining how many lags are needed to capture the dynamics of the system. In the literature, models using monthly data have had everything from three to twelve lags. The lag length has usually been fixed on ad hoc grounds rather than as the outcome of formal tests for lag length. As we are not able to present credible theoretical justifications for any particular lag length, we let the data indicate it. We examine systems with one to twelve lags, and use the likelihood ratio test statistic of Sims (1980a) with a degrees of freedom correction. This test statistic is known to be fairly insensitive to lag structure, but some Monte Carlo results indicate that it may be biased against the null hypothesis of smaller models (Nickelsburg (1985)). Results are presented in Table 5.1.

Judging by the tests for lag lengths presented in Table 5.1, the data do not give altogether clear indications of how many lags to employ. Nevertheless, a choice can be made between two and nine lags. To facilitate this choice, the residuals from two and nine lag models were scrutinized, and the nine-lag models were found to yield cleaner residuals than the two-lag models. Thus subsequent analyses of the Wold causal chain model which utilize the full data set will be based on VAR models with nine lags of each variable. Sensitivity of the results with respect to the choice of lag length will be investigated, however.

TABLE 5.1

Tests for lag length in a VAR model in error correction form for the vector $\underset{\sim}{y}_t = (y_t^{*U}\ y_t^{*E}/y_t^{*E*}\ y_t)'$ using data from the period 1960M1 - 1988M8

Lag length	1	2	3	4	5	6	7	8	9	10	11	12
1												
2			0.659	0.108								
3		0.262							[0]			
4		0.003	0.001									
5		0.001	0.001	0.066								
6				0.003	0.007		0.016					
7						0.006		0.001				
8							0.001		0.002	0.008		
9			[0]				0.002	0.241		0.326		
10							0.002	0.171	0.207			
11												
12												

The test is the likelihood ratio test statistic of Sims (1980a) incorporating a degrees of freedom correction. The upper half of the table reports marginal significance levels for models involving the variable y^{*E}, and the lower half concerns models involving the variable y^{*E*}. [0] indicates that values smaller than 0.0005 were obtained. In addition to the components of $\underset{\sim}{y}_t$, the model includes a constant, 11 seasonal dummy variables and a strike dummy taking the value 1 in 1971M2 - 1971M3 and 0 otherwise.

Estimation results for our model of the transmission of output fluctuations are presented in Table 5.2. Estimation is carried out using first differences of model variables, and the models include error correction (level) terms, constant terms, seasonal dummy variables and a strike dummy. Both models are estimated by equationwise OLS using the full data set. We note that since the estimated parameters are complicated nonlinear functions of country-specific parameters, it is hard to give a structural interpretation of individual parameter estimates. Rather, we can note the location of significance of explanatory power.

TABLE 5.2

Estimation results for a VAR model in error correction form for the vector $\underset{\sim}{y}_t = (y_t^{*U}\ y_t^{*E}/y_t^{*E*}\ y_t)'$ using data from the period 1960M1 - 1988M8
1960M1 - 1988M8

Regressors/ Statistics	Regressands						
	y_t^{*U}	y_t^{*E}	y_t	:	y_t^{*U}	y_t^{*E*}	y_t
y^{*U}	0.498 (3.37)	0.467 (1.04)	-0.869 (-0.36)		0.454 (3.05)	2.52 (3.34)	0.955 (0.38)
y^{*E}/y^{*E*}	-0.101 (-1.33)	-0.107 (-0.46)	3.43 (2.76)		-0.060 (-0.69)	-2.05 (-4.62)	-0.148 (-0.10)
y	0.038 (1.51)	-0.006 (-0.08)	-3.09 (-7.57)		0.036 (1.37)	-0.080 (-0.59)	-2.66 (-5.95)
l	9	9	9		9	9	9
$\bar{R}^2$	0.831	0.975	0.943		0.831	0.953	0.940
Q	0.677	0.004	0.000		0.463	0.000	0.000
JB	0.395	0.000	0.000		0.223	0.000	0.000
CHOW(1973M1)	0.029	0.138	0.088		0.017	0.050	0.062
CHOW(1980M1)	0.128	0.158	0.068		0.129	0.132	0.071
EXOGENEITY	0.008	0.000	0.013		0.008	0.033	0.335
ζy^{*U}		0.310	-0.131			-0.090	-0.156
$\Omega_\zeta = \zeta y^{*E}/\zeta y^{*E*}$			0.022				0.544
ζy							

The models are estimated by equation-by-equation OLS from first-differenced data. In addition to the elements of $\underset{\sim}{y}_t$, the models include error correction (level) terms, a constant, 11 seasonal dummy variables and a strike dummy taking the value 1 in 1971M2 - 1971M3 and 0 otherwise. Estimation results for error correction terms and the deterministic elements have been suppressed. Estimates on the rows for y^{*U}, y^{*E}/y^{*E*} and y are the sums of the estimated coefficients on lags 1 through 9 of a variable. Numbers in parentheses are t statistics. The lag length is denoted l and $\bar{R}^2$ is the degrees-of-freedom-corrected squared multiple correlation coefficient. Q is the Ljung & Box (1978) statistic based on nine autocorrelations, JB is the Jarque & Bera (1980) statistic and CHOW(.) is the Chow test with the date of the supposed structural break given in parentheses. EXOGENEITY stands for the degrees-of-freedom-corrected likelihood ratio test statistic of Sims (1980a) for the null hypothesis that lags of other model variables do not belong in the equation for a particular variable. Marginal significance levels are reported for Q, JB, CHOW and EXOGENEITY. The residual correlation matrix is denoted Ω_ζ, and the ζ^i, $i = y_t^{*U}, y_t^{*E}, y_t^{*E*}, y_t$, are equationwise residuals. Twice the asymptotic standard error of the residual correlations is 0.110.

In general, we find explanatory power in the blocks of own lags of variables. Hence, domestic economic conditions appear to be the most important determinants of domestic output. U.S. output significantly influences intermediate economy output, which, in turn, significantly influences Finnish output. According to tests for exogeneity, we can reject the hypothesis that the output of the rest of the world does not affect U.S. and intermediate economy output, but such a hypothesis cannot be rejected with respect to Finnish output in the model comprising the variable y_t^{*E*}.

The equations from which inference is drawn have high explanatory power, but the residuals in the intermediate economy equation and in the equation for Finnish output are temporally dependent. In addition, the residuals of these equations are nonnormally distributed. Hence, the tests for significance in these equations should be viewed with care. Furthermore, the results may be open to criticism because our monthly output series presumably contain a nonnegligible amount of measurement error, and since the results are conditional, inter alia, on the use of a linear model.[6]

Some signs of parameter instability are found when comparison is made between fixed and floating exchange rate regime periods. The null hypothesis of constant coefficients can be rejected at the five per cent level in the equations for U.S. and intermediate economy (y_t^{*E*}) output. Parameter constancy (given constant error variances) can be rejected at the 10 per cent level in the equations for Finnish output, both when the breaking point is assumed to be 1973M1 and when it is assumed to be 1980M1.

One should caution against inferring that these statistically significant differences between subsamples necessarily imply also

[6]There exists some evidence of nonlinearities in macroeconomic time series. Industrial production has been studied by Brock & Sayers (1988), Luukkonen & Teräsvirta (1988) and Ashley & Patterson (1989). Brock & Sayers (op.cit.) and Ashley & Patterson (op.cit.) find signs of nonlinearities in U.S. data. Luukkonen & Teräsvirta (op.cit.) find evidence of nonlinearity in data from France, Germany, Japan and the U.S., but not in data from Finland, Sweden or the U.K.

economically significant differences. Namely, coefficients estimated from different subperiods may well differ by an amount that is statistically significant but economically trivial if the coefficients are estimated with sufficient precision. Also, a significant parameter change can generate only a small change in the endogenous variable. This is true, in particular, in a dynamic model involving many parameters because offsetting changes in coefficients can leave important properties of the overall relationships practically unaffected. The economic significance of the statistically significant instabilities will be evaluated in section 5.2.5.

5.2.2 Analysis of Structural Shocks

The estimated contemporaneous relationships are presented in Table 5.3.

TABLE 5.3
Contemporaneous structural relationships in a VAR model in error correction form for the vector $\underset{\sim}{y}_t = (y^{*U}_t \; y^{*E}_t / y^{*E*}_t \; y_t)'$ using data from the period 1960M1 - 1988M8

Model variables	Matrix of contemporaneous relationships
$\underset{\sim}{y}_t = \begin{bmatrix} y^{*U}_t \\ y^{*E}_t \\ y_t \end{bmatrix}$	$\Gamma_0 = \begin{bmatrix} 0 & 0 & 0 \\ 0.941 & 0 & 0 \\ -2.49 & 0.374 & 0 \end{bmatrix}$
$\underset{\sim}{y}_t = \begin{bmatrix} y^{*U}_t \\ y^{*E*}_t \\ y_t \end{bmatrix}$	$\Gamma_0 = \begin{bmatrix} 0 & & \\ -0.459 & 0 & 0 \\ -1.81 & 1.76 & 0 \end{bmatrix}$

The models are estimated by equation-by-equation unconstrained OLS from first differenced data. In addition to the elements of $\underset{\sim}{y}_t$, the models include a constant, 11 seasonal dummy variables and a strike dummy taking the value 1 in 1971M2 - 1971M3 and 0 otherwise.

The estimates of contemporaneous relationships indicate that the intermediate (GDP weighted) economy is linked very closely to the U.S. economy. Finnish output is modestly linked to the GDP-weighted intermediate economy, but strongly linked to the trade-weighted intermediate economy. The contemporaneous relation between Finnish and U.S. output is negative. Closer scrutiny of this relationship revealed that a negative relationship existed during the Bretton Woods subperiod whereas the relationship has been positive after the breakdown of the Bretton Woods arrangement (see also the subsequent analyses in section 5.2.5). See Öller (1978) for one of the earliest documentations of short lags in relationships between Finnish and foreign macroeconomic variables.

Empirical evidence of the structural shocks is presented in Table 5.4. As expected, the impulses have zero mean and are all individually temporally independent. Some signs of autocorrelation at the seasonal lag - notably so in the case of Finnish output - bear witness to the fact that the fixed seasonal mean approximation does not capture all of the seasonal movements in our output variables. No skewness is found in the output impulses emerging from the U.S. economy and the intermediate economy, but output shocks tend to be skewed to the left in the case of the Finnish economy. The distribution of the shocks emerging from the Finnish economy may be skewed to the left because of the large negative shocks in the 1970s. These shocks cause the estimated variance of the Finnish shocks to be quite large. The large shocks may show up in the kurtosis as well, as the distribution of the Finnish disturbances is leptocurtic. In fact, we can reject the hypothesis that the estimated impulses are normally distributed in all cases.

TABLE 5.4

Analysis of output innovations as identified using a lower triangular contemporaneous design matrix in a VAR model in error correction form for the vector $\underset{\sim}{y}_t = (y_t^{*U}\ y_t^{*E}/y_t^{*E*}\ y_t)'$ using data from the period 1960M1 - 1988M

	Innovation					
Statistic	ξ^{*U}	ξ^{*T}	ξ	ξ^{*U}	ξ^{*T*}	ξ
Mean	0.000	0.000	0.000	0.000	0.000	0.000
Variance	8.28	68.8	2150	8.27	212	1610
Skewness	-0.215	-0.054	-1.72***	-0.114	0.007	-0.802***
Kurtosis	1.32***	0.900***	10.2***	1.19***	5.71***	4.38***
Autocorrelation						
1	-0.015	-0.024	0.036	-0.004	0.015	0.055
2	-0.025	0.027	0.038	-0.012	-0.036	0.089
3	0.030	0.018	-0.053	0.041	0.098*	-0.046
4	0.045	-0.018	-0.027	0.012	-0.029	-0.021
5	-0.018	0.032	0.050	-0.090*	-0.018	0.030
6	-0.021	-0.067	-0.054	0.002	0.072	-0.066
7	-0.039	0.074	-0.019	-0.016	-0.009	-0.025
8	-0.054	0.010	-0.013	-0.045	-0.016	0.031
9	0.022	-0.039	-0.071	0.050	0.024	-0.025
10	-0.056	0.000	-0.002	-0.024	-0.024	0.048
11	0.064	-0.056	-0.076	0.041	0.087	-0.064
12	0.158***	0.185***	0.561***	0.187***	0.443***	0.437***
Significant positive shock						
	1961M4	1963M4	1961M7	1961M4	1961M7	1961M7
	1961M10	1963M9	1964M7	1964M11	1962M7	1964M7
	1964M11	1964M9	1965M7	1967M8	1963M7	1969M7
	1967M8	1967M9	1969M7	1967M11	1964M7	1970M7
	1970M12	1968M8	1970M7	1970M12		1971M4
	1978M4	1972M8	1971M4	1978M4		
	1981M7	1973M8	1971M7	1981M7		
	1983M7			1983M7		
	1984M1			1984M7		
	1984M7			1988M7		
	1988M7					
Significant negative shock						
	1961M2	1962M6	1967M6	1960M11	1961M8	1971M2
	1963M7	1962M7	1971M2	1963M7	1964M8	1971M3
	1974M11	1964M8	1971M3	1974M11	1972M7	1975M7
	1974M12	1965M7	1975M7	1974M12	1975M7	1976M7
	1980M4	1968M1	1976M7	1980M4	1976M7	1977M3
	1980M5	1968M5	1977M3	1980M5	1977M7	1977M7
		1974M10	1977M7		1978M7	1978M7
		1980M9	1978M7		1980M5	1979M7
		1981M8	1979M7			
		1982M8				

The models are estimated by equation-by-equation unconstrained OLS from first differenced data. In addition to the elements of $\underset{\sim}{y}_t$, the models include a constant, 11 seasonal dummy variables and a strike dummy taking the value 1 in 1971M2 - 1971M3 and 0 otherwise. Significant values at the 10,5 and 1 per cent level of significance are indicated by *, ** and ***, respectively. Listed shocks exceed the two standard error limit (the dummy variable strike dates are listed as significant negative shocks).

The vast majority of shocks are small and rather uninteresting if viewed in isolation. This finding agrees with the Frisch - Slutsky - Lucas view of the nature of shocks, according to which the impulses represent a steady flow of small pieces of new information ("news"). With reference to the literature on rules versus discretion in economic policy (see the papers in Campbell & Dougan (1986)), our finding that the shocks represent a steady flow of small pieces of new information constitutes a case favoring the use of rule-based economic policy.

A look at major historical economic events facilitates the following interpretations of the estimated large output disturbances. With regard to large positive shocks in U.S. real total output, we find that 1961M4 and 1961M10 may be associated with the unexpectedly strong recovery from the 1960 - 1961 recession. 1964M11 may possibly be associated with the tax cuts in 1964, and 1967M8 and 1967M11 with the Vietnam War upturn. 1970M12 may represent the rebound from the General Motors strike of 1970M10, and 1978M4 falls into the expansion of the latter half of the 1970s. The shocks in 1983 - 1984 may be associated with the boom in 1983 - 1985 insofar as it was unexpected. With regard to large negative shocks, 1960M11 may be associated with the 1960 - 1961 recession and 1974M11 and 1974M12 reflect the deep 1974 - 1975 recession. 1980M4 and 1980M5 reflect the "minirecession" in 1980.

It may be interesting to compare the large output shocks identified above with corresponding shocks identified by Blanchard & Watson (1986) and Blanchard & Quah (1989). Blanchard & Watson (op.cit.) use an IS-LM model augmented with monetary and fiscal policy reaction functions to identify supply, demand, money and fiscal shocks. Blanchard & Quah (op.cit.) impose low frequency restrictions on a bivariate model of output growth and unemployment, and interpret their structural shocks as supply and demand shocks. In general, the results of Blanchard & Watson (op.cit.) and our results coincide to some extent, whereas the results presented in Blanchard & Quah (op.cit.) differ from the other results. Only about one in ten large shocks identified in Blanchard & Quah (op.cit.) coincide with the other two sets of shocks. These findings suggest that the shocks estimated by Blanchard & Quah (op.cit.) represent linear combinations of more than two individual structural shocks.

Turning to large output shocks in the intermediate economy, we note that results differ between the two measures of the intermediate economy. There may also be a problem with seasonality in the case of the trade-weighted intermediate economy. One may nevertheless note that 1980M5 is due to labor market disputes in Sweden. Some of the estimated large shocks to total real Finnish output also seem to have identifiable economic content. 1971M4 can be associated with the recovery from the strike in the metal and engineering industry earlier in the year, and 1976M7 may be associated with strikes during 1976. Among the estimated large negative disturbances, 1971M2 and 1971M3 stand for the strike in the metal and engineering industry. 1977M3 and 1977M7 can be associated with the sharp, unexpected falls in output in 1977. We can also note the absence of large domestic output shocks in the 1980s.

5.2.3 Impulse Responses

In this section we compute short- and long-run impulse responses.[7] The confidence bounds for the impulse responses are not easily constructed analytically, because the impulse responses are nonlinear and convoluted functions of the coefficients of the VAR model (a relatively simple closed-form result has, nevertheless, recently been presented by Lütkepohl (1989)). Instead, the second moments of the impulse responses are computed using Bayesian methods and Monte Carlo integration. Zellner (1971) shows that the posterior distribution of the positive definite symmetric random coefficient matrix E(L) is an inverted Wishart, given an uninformative prior on the coefficients, the probability density function of $\underset{\sim}{y}_t$, and

[7]The level (error correction) terms in our first-differenced model are problematic from the point of view of inverting the model to yield the vector moving average representation. To this end, the estimated model is, following a suggestion by Mark W. Watson, reparametrized in level form and then inverted (see Doan (1989), p. 14-41).

normally distributed reduced form errors $\underset{\sim}{\zeta}_t$. Kloek & VanDijk (1978) outline a procedure for calculating the first two moments of the posterior distribution of the impulse responses. The integration procedure draws repeatedly from the inverted Wishart distribution and computes the moments that result from each drawing.[8]

Impulse responses with two standard deviation limits for systems involving y_t^{*E} and y_t^{*E*} are reported in Figures 5.3a - 5.3f and 5.4a - 5.4f, respectively. Responses refer to a one standard deviation, one period, positive innovation, and all estimates are based on the full sample period. Impulse responses refer to the (logarithmic) level of a model variable, and the responses are scaled to yield fractions of a standard deviation.

Turning first to the responses of output to a U.S. output shock the following picture emerges. The intermediate economies, weighted according to the size of the economy, display a positive contemporaneous relationship with U.S. output, and the U.S. output shock stimulates intermediate economy output, peaking three quarters after the shock. Some slight tapering off of the effect of the U.S. shock on the intermediate economy can be observed, and virtually all of the contribution to output variablity in the intermediate economy has ceased two years after the shock. The U.S. output shock raises intermediate economy output permanently in the long run. When the intermediate economy countries are weighted with Finnish trade weights, essentially the same dynamic pattern is observed, apart from a negative contemporaneous relationship and an overall smaller impact.

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[8]In the vast majority of VAR studies, confidence bounds have not been generated. We report results based on 1 000 drawings from the posterior distribution. Burbidge & Harrison (1984, 1985) and Genberg et al. (1987) use 100 drawings, but we found that results based on 100 drawings differed considerably from results based on larger numbers of drawings. More specifically, the second moments were systematically understated when a smaller number of drawings was used. Regarding the magnitude of the moments it should be noted that they may not be correct, since the estimated residuals $\underset{\sim}{\zeta}_t$ were not normally distributed in the intermediate economy and the Finnish economy equations (see Table 5.2).

FIGURE 5.3

Impulse responses for a VAR model in error correction form for the vector $\underset{\sim}{y}_t = (y_t^{*U}\ y_t^{*E}\ y_t)'$ using data from the period 1960M1 - 1988M8 (y_t^{*E} is GDP-weighted)

FIGURE 5.3a

RESPONSE OF INTERMEDIATE ECONOMY OUTPUT TO SHOCK IN U.S OUTPUT

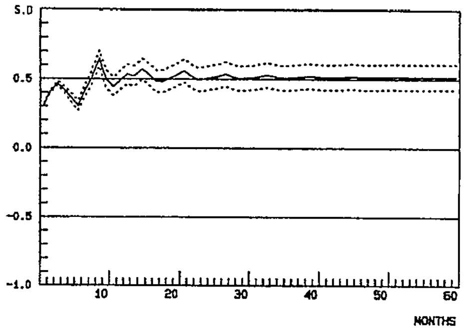


FIGURE 5.3b

RESPONSE OF FINNISH OUTPUT TO SHOCK IN U.S. OUTPUT

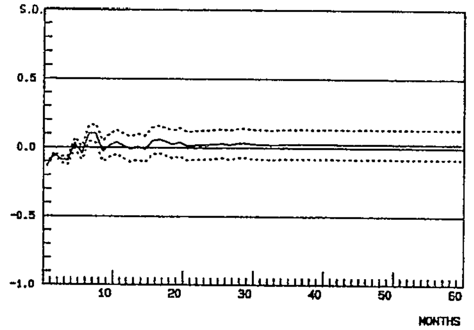


FIGURE 5.3c

RESPONSE OF U.S. OUTPUT TO SHOCK IN INTERMEDIATE ECONOMY OUTPUT

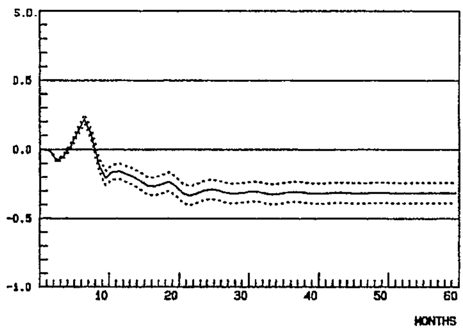


FIGURE 5.3d

RESPONSE OF FINNISH OUTPUT TO SHOCK IN INTERMEDIATE ECONOMY OUTPUT

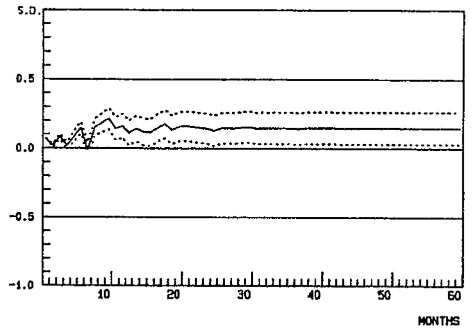


FIGURE 5.3e

RESPONSE OF U.S. OUTPUT TO SHOCK IN FINNISH OUTPUT

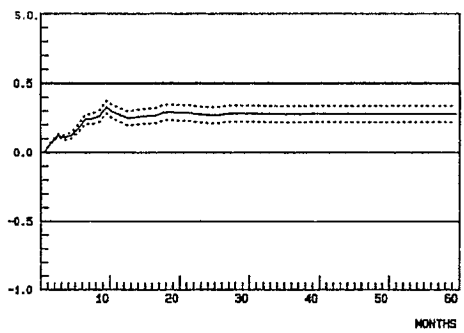
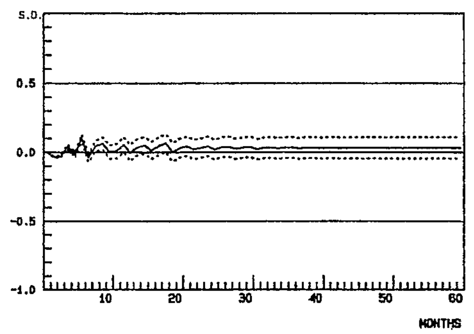


FIGURE 5.3f

RESPONSE OF INTERMEDIATE ECONOMY OUTPUT TO SHOCK IN FINNISH OUTPUT



Notes: ———————— impulse response
.... two standard deviation limit

FIGURE 5.4

Impulse responses for a VAR model in error correction form for the vector $\underset{\sim}{y}_t = (y_t^{*U}\ y_t^{*E*}\ y_t)'$ using data from the period 1960M1 - 1988M8 (y_t^{*E*} is trade-weighted)

FIGURE 5.4a

RESPONSE OF INTERMEDIATE ECONOMY OUTPUT TO SHOCK IN U.S. OUTPUT

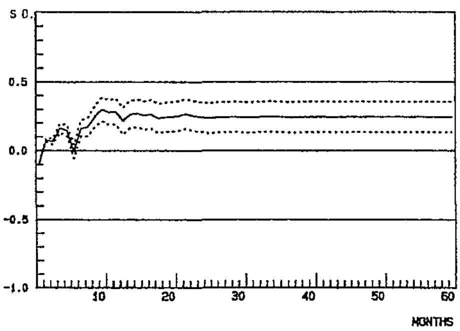


FIGURE 5.4b

RESPONSE OF FINNISH OUTPUT TO SHOCK IN U.S. OUTPUT

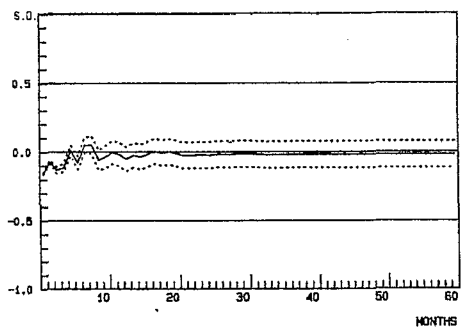


FIGURE 5.4c

RESPONSE OF U.S. OUTPUT TO SHOCK IN INTERMEDIATE ECONOMY OUTPUT

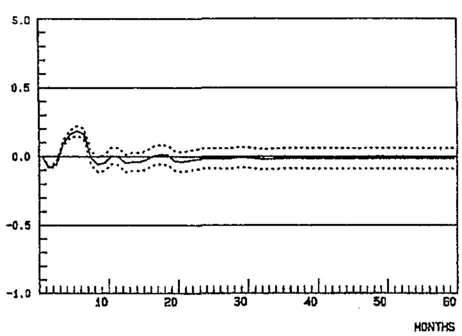


FIGURE 5.4d

RESPONSE OF FINNISH OUTPUT TO SHOCK IN INTERMEDIATE ECONOMY OUTPUT

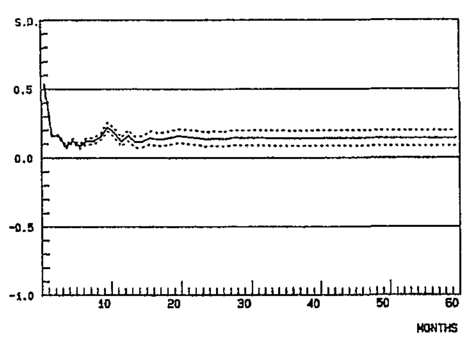


FIGURE 5.4e

RESPONSE OF U.S. OUTPUT TO SHOCK IN FINNISH OUTPUT

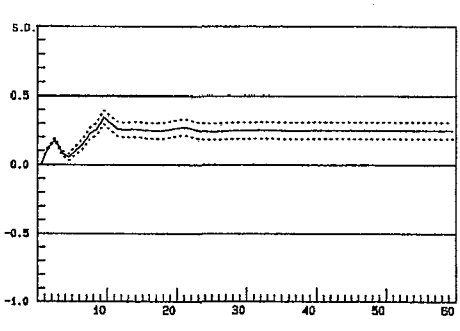
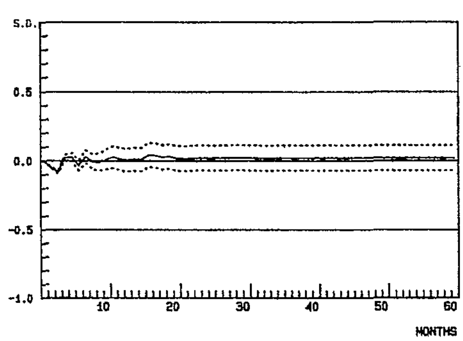


FIGURE 5.4f

RESPONSE OF INTERMEDIATE ECONOMY OUTPUT TO SHOCK IN FINNISH OUTPUT



Notes: ———————— impulse response
.... two standard deviation limit

The impact of a U.S. output shock on Finnish output emerges as very weak, if anything. The short-run impact is negative, and a small positive impact is recorded during the third quarter after the shock. The long-run impact is essentially zero. On average, the impact of a U.S. shock is clearly bigger on the intermediate economy than on the Finnish economy. We note that Gerlach (1988) - using cross-spectral techniques - finds that Norway and Sweden are also less affected than larger economies by foreign output movements.

The impacts of an output innovation in the intermediate economy on the rest of the world are graphed in Figures 5.3c - 5.3d and 5.4c - 5.4d. A shock to the GDP-weighted intermediate economy is seen to have brought about a small and short-lived stimulus to U.S. output half a year after the shock, but the long-run impact is negative. The same modest positive impact roughly half a year after a shock is found when Finnish trade weights are used, but now the long-run impact is zero. There is hardly any difference between the impulse responses of Finnish output to a shock in the intermediate economy depending on which weighting scheme is employed.

An intermediate economy shock has, historically, brought about a nonnegligible stimulus to Finnish output, which has peaked 10 months after the occurrence of the shock. In the long-run, Finnish output has been raised permanently by the foreign output innovation. The very short-run impact of the intermediate economy shock is different depending on how the intermediate economy is weighted; Finnish trade weights generate a sizeable contemporaneous relationship whereas GDP weights generate only a small contemporaneous correlation. The responses are dynamically well behaved with the dynamic effects having essentially worked themselves through the model in two years.

When U.S. output is subjected to a shock in Finnish output, a modest, positive impact is observed, which peaks some 10 months after the shock, and a positive long-run impact. This impulse response clearly seems spurious, and we suspect that it is a consequence of one (or a few) influential observations. When intermediate economy output is shocked by a Finnish shock, the impact is significantly positive, albeit very small, and the long-run impact is zero.

Apart from the spurious U.S. response, the impulse responses emerge as reasonable. The responses are dynamically well behaved and essentially of the expected sign and plausible magnitude. Intermediate economy output reacts more strongly than Finnish output to U.S. output shocks, intermediate economy output reacts more strongly to U.S. than to Finnish output shocks, and Finnish output reacts more strongly to intermediate economy output shocks than to U.S. output shocks. In other words, a hierarchial model of international output dynamics in which the U.S. calls the tune and Finland is "small and open" seems to be a reasonable approximation.

5.2.4 Decompositions of Variance

Decompositions of variance for systems involving $\underset{\sim}{y}_t^{E}$ and $\underset{\sim}{y}_t^{E*}$ are reported in Table 5.5. Decompositions of variance refer to a one standard deviation, one period, positive innovation, and all estimates are based on the full sample period. As all estimated λs converged within the 60-month horizon, we will be referring to 60-month horizon values as long-run values. We find that U.S. output is convincingly exogenous at all time horizons. The percentage of the expected 60-months-ahead squared prediction error in U.S. output caused by an innovation in the same variable is in excess of 95.0, whereas innovations in the other two output variables account for less than three percentage points each. Note, however, that less than half of U.S. foreign trade is with the other countries in $\underset{\sim}{y}_t$; hence it is possible that output developments in e.g. Canada - with which the U.S. conducts one fifth of its foreign trade - shows up in U.S. output developments.

TABLE 5.5

Variance decomposition for a VAR model in error correction form for the vector $\underset{\sim}{y}_t = (y_t^{*U}\ y_t^{*E}/y_t^{*E*}\ y_t)'$ using data from the period 1960M1 - 1988M8

Percentage of the expected k-step-ahead squared prediction error in variable	k	Innovation in					
		y*U	y*E	y :	y*U	y*E*	y
y*U	1	100	0.000	0.000	100	0.000	0.000
	2	99.8	0.003	0.187	99.2	0.259	0.516
	3	99.4	0.162	0.442	98.8	0.215	0.980
	6	99.1	0.230	0.676	98.8	0.645	0.553
	9	98.3	0.460	1.27	98.6	0.496	0.952
	12	97.6	0.639	1.81	98.1	0.357	1.58
	24	96.3	1.55	2.14	97.9	0.182	1.89
	60	95.1	2.55	2.37	97.9	0.077	2.03
y*E/y*E*	1	9.62	90.4	0.000	0.814	99.2	0.000
	2	16.3	83.6	0.075	1.27	98.5	0.227
	3	22.4	77.5	0.094	1.46	97.8	0.713
	6	22.0	77.6	0.360	4.10	95.2	0.711
	9	25.5	74.2	0.324	9.10	90.3	0.621
	12	26.1	73.6	0.271	17.1	82.4	0.532
	24	27.6	72.2	0.202	26.4	73.1	0.449
	60	28.4	71.4	0.145	32.8	66.9	0.346
y	1	1.71	0.438	97.8	2.45	28.3	69.2
	2	1.72	0.420	97.9	2.57	26.6	70.8
	3	2.27	0.994	96.7	3.82	27.0	69.2
	6	2.82	3.00	94.2	4.62	25.2	70.2
	9	3.90	6.31	89.8	4.70	26.0	69.3
	12	3.32	10.0	86.7	4.07	27.0	68.9
	24	2.59	15.2	82.3	3.05	26.8	70.2
	60	1.73	20.8	77.5	1.82	26.9	71.2

The models are estimated by equation-by-equation unconstrained OLS from first differenced data. In addition to the elements of $\underset{\sim}{y}_t$, the models include a constant, 11 seasonal dummy variables and a strike dummy taking the value 1 in 1971M2 - 1971M3 and 0 otherwise. Numbers across rows may not sum to 100.000 due to rounding errors.

The intermediate economy is also highly exogenous with respect to the Finnish economy, but not with respect to the U.S. economy. In the long run, shocks to the intermediate economy account for two thirds of the squared prediction error in intermediate economy output, while innovations in U.S. output account for the remaining third. In fact, when the GDP-weighted intermediate economy aggregate is used - which is the natural one to use when comparison is made with the U.S. - the importance of the U.S. economy is apparent at all forecasting horizons. When Finnish trade weights are used, the U.S. economy starts to exert a sizeable influence on the intermediate economy two to three quarters after the shock.

The impact of a U.S. output shock on Finnish output appears to be negligible for horizons up to five months, but it may account for a small portion of the expected squared forecast error at the six- to 12-month horizon. By contrast, intermediate economy output shocks account for nonnegligible portions of Finnish output developments. When the GDP-weighted aggregate y_t^{*E} is employed, the intermediate economy shock shows up in Finnish output about three quarters after the occurrence of the shock, and unexpected intermediate economy output movements account for one fifth of Finnish output movements in the long-run. Using the trade-weighted intermediate economy concept y_t^{*E*} shows a somewhat stronger influence of the intermediate economy on Finnish output developments; unexpected intermediate economy output developments account for one fourth of Finnish output variability both in the short and the long run.[9]

On the whole, the decompositions of variance give the following picture of fluctuations in total real output in the world economy. The U.S. economy, and to a lesser degree the intermediate economy and the Finnish economy, all seem to have grown largely dominated by movements in the respective domestic output. More precisely, U.S. output emerges as totally exogenous with respect to output in the rest of the world (as defined in the current study). The

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[9]These findings are robust to the use of the strike dummy to account for the strike in the metal and engineering industry in Finland in 1971.

intermediate economy is influenced by U.S. output developments, but not by Finnish output developments. Finnish output, in turn, is affected by intermediate economy output, but only arguably directly by U.S. output. Hence, what emerges is, notwithstanding some differences due to how the intermediate economy is parametrized, a hierarchial setup in which the smaller countries display the same order of magnitude of openness. Some signs of Finland being slightly less open than an average intermediate economy could be noted, however; foreign output developments account for some 29 - 33 per cent of the long-run intermediate economy output variability whereas the corresponding figures for Finland are 22 - 29 per cent.[10,11]

5.2.5 Analysis of Subperiods

Instead of reporting three new sets of Tables 5.1 - 5.5 and corresponding impulse responses, we have collected a selection of the key summary statistics in one table (Table 5.6), and will report impulse responses for reactions of Finnish output to U.S. and intermediate economy output shocks only.

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[10] When quoting percentage points, the most likely major uncertainty attached to these should be borne in mind. Work by Geweke (1984), Runkle (1987) and Diebold & Rudebusch (1989) suggests that confidence intervals attached to decompositions of variance from unrestricted VARs may be large.

[11] The analysis of sections 5.2.1 - 5.2.4 was repeated using lag lengths of six months and 12 months. Such a check for robustness of results seems warranted, inter alia, since the outliers may cause problems when testing for lag length. Results using six lag models were fairly similar to the ones reported in the main text, but all effects worked themselves through more rapidly than in the base case. When 12 lags were used most qualitative results carried through, but some nonsensical impulse responses bore witness of overparametrization. Presumably, the true signals were obscured by the noise from distant lags.

TABLE 5.6

Selected summary statistics for VAR models for the vector $\underset{\sim}{y}_t = (y_t^{*U}\ y_t^{*E}/y_t^{*E*}\ y_t)'$ using data from subperiods

Time period		1960M1 - 1972M12						1973M1 - 1988M8						1980M1 - 1988M8					
Regressand		y^{*U}	y^{*E}	y :	y^{*U}	y^{*E*}	y :	y^{*U}	y^{*E}	y :	y^{*U}	y^{*E*}	y :	y^{*U}	y^{*E}	y :	y^{*U}	y^{*E*}	y
Statistic	k																		
l		7	7	7	7	7	7	3	3	3	3	3	3	4	4	4	4	4	4
$\bar{R}^2$		0.888	0.974	0.965	0.890	0.950	0.964	0.808	0.979	0.958	0.822	0.979	0.959	0.846	0.992	0.986	0.852	0.989	0.986
Q		0.884	0.838	0.034	0.872	0.573	0.021	0.410	0.000	0.000	0.729	0.003	0.000	0.681	0.240	0.383	0.191	0.478	0.810
EXOGENEITY		0.643	0.658	0.087	0.497	0.548	0.188	0.985	0.239	0.196	0.060	0.696	0.076	0.748	0.555	0.605	0.525	0.675	0.604
Percentage of the expected k-step ahead squared prediction error in variable																			
y^{*U}	1	100	0.000	0.000	100	0.000	0.000	100	0.000	0.000	100	0.000	0.000	100	0.000	0.000	100	0.000	0.000
	2	98.8	0.304	0.933	98.6	0.132	1.26	99.9	0.003	0.067	99.6	0.042	0.317	98.5	1.35	0.144	98.7	1.32	0.003
	3	98.9	0.511	0.581	99.1	0.096	0.833	99.8	0.002	0.168	99.5	0.031	0.498	97.3	2.61	0.083	98.8	1.16	0.009
	6	99.2	0.577	0.249	99.3	0.331	0.347	99.8	0.038	0.209	99.0	0.849	0.194	96.5	3.11	0.432	99.1	0.436	0.458
	9	99.5	0.366	0.149	99.6	0.198	0.231	99.7	0.051	0.231	98.8	1.10	0.120	96.2	3.29	0.467	99.3	0.283	0.410
	12	99.6	0.269	0.107	99.7	0.132	0.176	99.7	0.057	0.239	98.7	1.22	0.086	96.1	3.39	0.514	99.3	0.211	0.484
	24	99.6	0.335	0.043	99.9	0.053	0.091	99.7	0.064	0.251	98.6	1.38	0.041	95.9	3.54	0.580	99.3	0.126	0.550
	60	99.5	0.464	0.017	99.9	0.018	0.057	99.7	0.069	0.257	98.5	1.46	0.017	95.8	3.62	0.619	99.3	0.090	0.596
y^{*E}/y^{*E*}	1	0.748	99.3	0.000	0.081	99.9	0.000	11.3	88.7	0.000	11.1	88.9	0.000	1.75	98.2	0.000	2.15	97.9	0.000
	2	3.64	96.2	0.123	0.591	98.5	0.918	14.9	84.6	0.537	9.14	90.6	0.247	1.77	98.2	0.063	2.31	97.5	0.230
	3	4.32	95.5	0.162	3.17	96.0	0.860	19.6	79.9	0.474	9.10	90.1	0.777	1.74	97.4	0.848	2.02	97.1	0.873
	6	2.39	97.4	0.212	3.70	94.6	1.67	30.9	68.1	1.06	9.39	90.1	0.537	8.76	90.7	0.543	3.33	95.5	1.18
	9	1.62	98.2	0.192	3.55	94.2	2.30	35.0	63.9	1.13	9.30	90.2	0.451	18.7	80.9	0.383	3.86	95.1	1.05
	12	1.23	98.6	0.178	3.79	94.1	2.14	37.1	61.8	1.16	9.28	90.3	0.399	28.7	71.1	0.282	5.12	93.8	1.05
	24	0.668	99.2	0.140	3.34	94.9	1.73	40.2	58.6	1.20	9.25	90.4	0.311	54.2	45.7	0.149	9.25	89.9	0.866
	60	0.362	99.5	0.117	2.89	95.8	1.26	42.0	56.8	1.22	9.22	90.5	0.253	72.7	27.1	0.132	14.5	84.9	0.673
y	1	0.160	0.938	98.9	0.182	0.909	98.9	2.76	2.32	94.9	2.81	14.8	82.4	1.74	2.54	95.7	2.59	8.86	88.6
	2	0.324	0.728	98.9	0.429	1.37	98.2	1.98	1.63	96.4	2.08	12.8	85.1	2.14	2.25	95.6	2.52	9.54	87.9
	3	0.312	0.848	98.8	0.394	2.24	97.4	2.28	1.70	96.0	3.23	11.3	85.4	3.75	2.25	94.0	3.21	9.18	87.6
	6	2.59	2.13	95.3	2.45	6.37	91.2	4.22	1.18	94.6	7.27	8.27	84.5	8.11	2.09	89.8	5.49	10.7	83.8
	9	5.01	2.70	92.3	4.80	6.16	89.0	5.65	0.932	93.4	9.43	6.48	84.1	10.3	1.71	88.0	5.20	10.7	84.1
	12	6.78	2.35	90.9	6.44	5.32	88.2	6.53	0.783	92.7	10.7	5.53	83.8	11.6	1.46	86.9	4.57	11.4	84.0
	24	9.58	1.56	88.9	8.65	5.98	85.4	8.04	0.527	91.4	12.7	3.94	83.4	16.2	0.880	82.9	3.16	11.9	84.9
	60	13.7	0.774	85.6	12.3	5.84	81.9	9.06	0.353	90.6	14.0	2.89	83.1	21.6	0.405	78.0	1.71	12.5	85.8

The models are estimated by equation-by-equation unconstrained OLS from first differenced data. In addition to the elements of $\underset{\sim}{y}_t$, the models include a constant, 11 seasonal dummy variables and, where appropriate, a strike dummy taking the value 1 in 1971M2 - 1971M3 and 0 otherwise. The models for the time periods 1973M1 - 1988M8 and 1980M1 - 1988M8 include error correction (level) terms.

The lag length is denoted l and $\bar{R}^2$ is the degrees-of-freedom-corrected squared multiple correlation coefficient. Q is the Ljung & Box (1978) statistic based on four, five and three autocorrelations, respectively. EXOGENEITY stands for the degrees-of-freedom-corrected likelihood ratio test statistic of Sims (1980a) for the null hypothesis that lags of other model variables do not belong in the equation for a particular variable. Marginal significance levels are reported for Q and EXOGENEITY. In decompositions of variance, numbers across rows may not sum to 100.000 due to rounding errors.

Turning to Table 5.6 we can make the following observations. The number of lags needed to capture the dynamics of output differs across subperiods, and the number of lags also differs from full sample results. During the Bretton Woods subperiod, output developments have been largely independent. In the very short run, U.S. output shocks may have contributed a little to intermediate economy output variability, however. U.S. output developments contributed to Finnish output variability, the impact becoming nonnegligible three quarters after the occurrence of the shock. A very modest contribution of developments in trade-weighted intermediate economy output to developments in Finnish output also appears to have occurred two quarters after the shock. In sum, foreign output developments accounted for less than one fifth of the long-run movements of Finnish output during the Bretton Woods subperiod.

During the post Bretton Woods subperiod, U.S. output developments continue to be wholly exogenous. By contrast, intermediate economy output now reacts to U.S. output both in the short and the long run. More than one third of the variability of (GDP-weighted) intermediate economy output comes from movements in U.S. output at the one year horizon, and the corresponding long-run influence is in excess of two fifths. As in the case of the Bretton Woods subperiod, we do not find any signs of Finnish output developments affecting intermediate economy output (or U.S. output, for the matter).

Nearly one fifth of Finnish output variability comes from foreign output movements during the post-Bretton Woods subperiod. U.S. output developments start to show up in Finnish output half a year after the shock, and one seventh of Finnish output variability is caused by U.S. output in the long run. Intermediate economy output accounts for one seventh of Finnish output variability in the very short run, but the influence vanishes within two years after the occurrence of the intermediate economy shock. At least one tenth, but not more than one fifth, of Finnish output variability was caused by foreign output movements after the breakdown of the Bretton Woods arrangement.

When analyzing the 1980s only, U.S. output still appears to be unaffected by international output. The impact of a U.S. output shock shows up in intermediate economy output about half a year after the occurrence of the shock. In other words, during the EMS subperiod, intermediate economy output is insulated from developments in U.S. output in the very short run in contrast to the experience during the late 1970s. On the other hand, U.S. output appears to be of paramount importance for intermediate economy output in the 1980s; at the one year horizon nearly one third, at the two year horizon more than one half, and in the long run nearly three quarters of intermediate economy output variability can be attributed to U.S. output developments.

During the 1980s, Finnish output has been slightly less susceptible to foreign output developments than during most of the 1970s. U.S. output developments show up in Finnish output variability during the third and fourth quarters following a shock, and the intermediate economy accounts for roughly one tenth of Finnish output variability at all forecasting horizons. When the intermediate economy countries are weighted by GDP, their importance for Finnish output developments appears to be negligible, while U.S. output starts to show up in Finnish output variability half a year after a shock. U.S. output variability accounts for one fifth of the long-run variability of Finnish output in such a model. This means that at least one seventh, possibly even one fifth, of Finnish long-run output variability comes from foreign output developments in the 1980s.

The decompositions of variance attest to a causally prior role for U.S. output - a role which has increased during our sample period. The intermediate economy and the Finnish economy have become increasingly susceptible to U.S. output movements, and Finnish output has become somewhat more susceptible to intermediate economy output variability. Nevertheless, no more than one fifth of eventual Finnish output variability appears to originate in foreign output fluctuations in the 1980s. These subsample results tend to reinforce the picture of a hierarchial setup with regard to international output dynamics, but they also suggest that the intermediate economy may be more, and the Finnish economy somewhat less, open in the 1980s than the full sample results suggest.

While we have documented some signs of a decrease in the importance of U.S. output shocks (ξ_t^{*U}), and an increase in the importance of intermediate economy output shocks (ξ_t^{*T}), for Finnish output movements, it does not follow that developments in U.S. output have become less important for developments in Finnish output. The main reason is that the estimated intermediate economy impulse ξ_t^{*T} contains the (unexpected) reaction of the intermediate economy to the U.S. output shock ξ_t^{*U}. Also, we have documented an overall increase in the importance of shocks to U.S. output for developments in intermediate economy output. Thus, U.S. output seems to be an increasingly important determinant of Finnish output developments.

To further assess the importance of foreign output impulses for the Finnish economy, impulse responses based on our three subsamples were calculated. The responses of Finnish output to unexpected movements in U.S. output and intermediate economy output are displayed in Figures 5.5a - 5.5d. In general, results between subsamples, on the one hand, and subsamples and the whole sample, on the other hand, differ from each other. Responses of Finnish output to shocks in the output of the intermediate economy also differ depending on which weighting scheme is used in constructing the intermediate economy, but responses to U.S. output innovations are unaffected by this.

FIGURE 5.5

Impulse responses for VAR models for the vector $\underset{\sim}{y}_t = (y_t^{*U}\; y_t^{*E}/y_t^{*E*}\; y_t)'$ using data from subperiods

FIGURE 5.5a

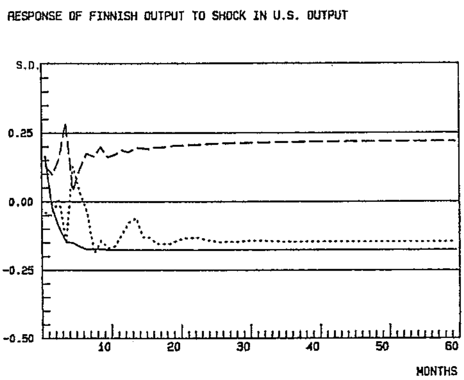


FIGURE 5.5b

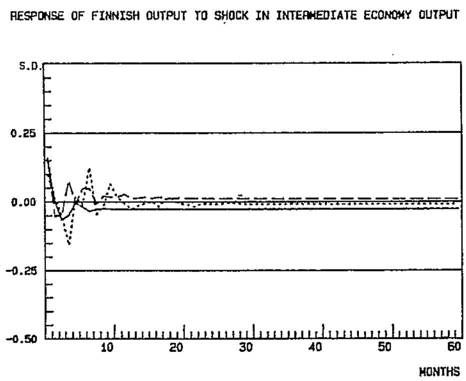


FIGURE 5.5c

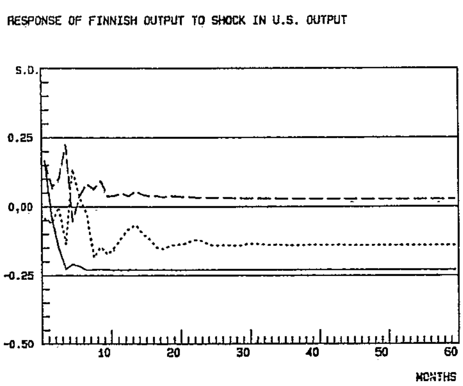
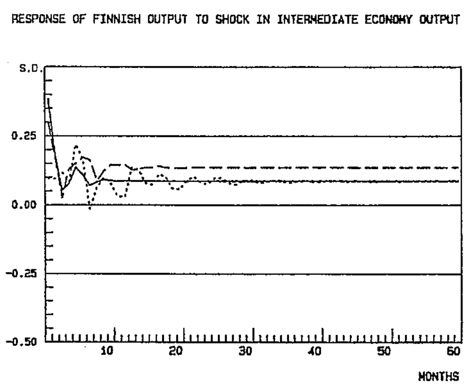


FIGURE 5.5d



Notes:

- - - - - 1960M1 - 1972M12

———— 1973M1 - 1988M8

— — — 1980M1 - 1988M8

Figures 5.5a - 5.5b refer to systems involving the GDP-weighted variable y_t^{*E}, and Figures 5.5c - 5.5d refer to systems involving the trade-weighted variable y_t^{*E*}.

With regard to the response of Finnish output to shocks to U.S. output, we can make the following observations. During the Bretton Woods era we find a sharp positive peak at the 5-month horizon, and the same response is observed in the analysis of data from the 1980s, although the magnitude of the response is bigger and the peak now occurs one month earlier. The contribution to Finnish output variability of the U.S. output shock dies out roughly within two years after the shock during the first subsample, while not much in the way of dynamics is found during the second subsample. During the third subsample, the long-run impact differs according to which intermediate economy aggregate is utilized. Using GDP weights, Finnish output is permanently raised by the U.S. output shock, but using trade weights no long-run impact is found. This parallels our decomposition of variance results.

Finnish output is positively affected contemporaneously by a shock to GDP-weighted intermediate economy output. The positive impact peaks seven months after the shock, but no long-run impact is found. Positive contemporaneous impacts and zero long-run impacts are, in fact, found during all subperiods. Positive impacts are also documented using the trade-weighted intermediate economy variable, but the impacts are bigger overall, and nonzero (positive) long-run effects emerge. Again, the findings parallel the results of the decompositions of variance. The impact of an intermediate economy shock on Finnish output seems to have increased somewhat over time, and the effects work themselves into Finnish output more rapidly than earlier. The subsample impulse responses suggest that changes in the dynamics of international output movements may have taken place. It may furthermore be mentioned that the impact of U.S. output movements on intermediate economy output differed across subsamples (graphs not shown); the impact has become stronger and more persistent over time.

Our empirical evidence confirms and quantifies the generally held belief that international economic linkages have increased over time.[12,13] However, the present findings differ to some extent from corresponding results based on seasonally adjusted data and disregarding the cointegration constraints presented in Starck (1988b). The differences are not dramatic, but the use of seasonally adjusted data implies twice as large a role for U.S. output developments in explaining Finnish output variability during the Bretton Woods subperiod than does the current analysis. Some decompositions of variance are distorted and the model dynamics are systematically understated. Hence, the analyses corroborate the claim that hazards for inference may appear if seasonally adjusted data are used and if cointegration relationships are neglected.

5.3 Conclusions

In this chapter we have implemented a new research strategy for the analysis of the extent to which movements in Finnish output are due to foreign output impulses. The empirical analyses were carried out using a hierarchial, three-country setup. The analysis yielded the following results.

Shocks to output have zero mean and are temporally independent, and they are normally small and insignificant when viewed one at a time.

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[12]One should bear in mind that we are analyzing aggregate openness, and we are not addressing the openness of different sectors of the economy. For analyses of such aspects, see Long & Plosser (1987), Stockman (1988), Krieger (1989) and Peisa (1989). Peisa (op.cit.), who analyzes output growth in the Finnish economy during 1961 - 1987, finds that sector-specific shocks do not provide a satisfactory explanation for fluctuations in aggregate output growth. He concludes that aggregate growth is driven mainly by economy-wide disturbances, especially at business cycle frequencies.

[13]Baxter & Stockman (1989) conclude that the correlation of international output fluctuations has decreased after the breakdown of the Bretton Woods arrangement. The converse is documented for the Finnish case. While these conclusions are correct in terms of a correlation coefficient, they bear witness to the hazards for inference of descriptive analyses that are too simple. In the light of our analysis, the conclusion of Baxter & Stockman (op.cit.) appears to be, if not outrightly false, at least dangerously uninformative.

Occasional large shocks were also documented, and many of these could be associated with historically exceptional events. The distribution of the estimated impulses is not normal, and in particular the distribution has tails which are too thick.

Finnish output is contemporaneously positively influenced by intermediate economy output movements and negatively linked to movements in U.S. output. Intermediate economy output is contemporaneously related to U.S. output in almost a one-to-one fashion. The effects of output movements in one country on output movements in another country in general peak during the third or fourth quarter following a shock, and dampen out fairly well within two years after the occurrence of a shock. Typically, the foreign shock causes a permanent change in the level of domestic output. The caveat to these results because inference is based on data which may contain a measurement error should be borne in mind, however.

U.S. output is found to be exogenous with respect to international output. U.S. output accounts for a sizeable portion of the variability of intermediate economy output, which, in turn, accounts for a sizeable portion of Finnish output variability. The susceptibility of intermediate economy output to U.S. output shocks appears to have increased over time, and during the 1980s U.S. shocks may account for nearly three quarters of the eventual movements of intermediate economy output. Foreign output developments account for one seventh to one fifth of the long-run variability of Finnish output during the 1980s. In terms of international output dynamics, the Finnish economy may not have been quite as open during the 1960s and the early 1970s as previously thought.

6 A BUSINESS CYCLE FLUCTUATION MODEL

The aim of this chapter is to provide empirical evidence on the long-run effects of shocks hitting the Finnish economy. Consequently, we achieve identification of shocks by imposing high frequency restrictions on our vector autoregression of foreign and domestic variables. By restricting high (and seasonal) frequency movements, we achieve the necessary identification of shocks and at the same time leave the data free to speak about low frequency movements. The results on low frequency movements will be further utilized in the next chapter in the identification of shocks when data are left free to speak about high frequency movements.

6.1 A Theoretical Model with High Frequency Restrictions

The choice of what variables to study - not to mention specifying how they interact - is inherently difficult. The classical paper by Koopmans (1947) makes it only too clear that the question admits of different answers in different cases, and that the answers depend, among other things, on the aim and scope of the study. To facilitate the choice of variables for the model and the identifying restrictions on the matrix Γ_0 of contemporaneous relationships, we will present and analyze a theoretical dynamic model of the Finnish economy. Although structural VAR studies (Bernanke (1986), Blanchard & Watson (1986), and others) have devoted practically no effort to motivating their choice of variables and models, we feel that a somewhat more lengthy discussion can be justified in a study of the current scope.

Firstly, economic theory should be consulted. Theoretical considerations may give us some hints as to how variables interact, but a unanimous indication as to which variables to include is not forthcoming. One easily ends up with a fairly large number of variables that potentially could be taken into account. Secondly, statistical inference may be helpful. Our preferred method is to make only limited use of tests for significance. We will test for the relevance of variables over and above a sparse set of core variables only when compelling empirical arguments for the inclusion of a variable can be advanced. It should be borne in mind that we want to document empirically regularities of major interest - not the potential influence of the myriad of factors that theoretically could matter for some aspect of some subset of economic activities. Thirdly, degrees of freedom considerations set upper bounds on the number of variables in a VAR model. Fourthly, there will be purely computational (memory) constraints limiting the scope of the analysis.

Using the theoretical model presented below, we claim that the variables output (y), the price level (p), the foreign interest rate (i^*), M1 (m) and the exchange rate (e) constitute a set of core variables for the study of a small open economy.[1] In the applied work, we will rely on statistical inference to determine whether or not to add a sixth variable to the model. Economic theory suggests that such a variable could be foreign output (y^*), foreign prices (p^*), a proxy for fiscal policy (g) or the price of oil (p^0).

The literature is full of basically similar models of an open economy, the core of which is the structure due to Dornbusch (1976), who built on the work of Mundell (1968) and Fleming (1962). We will follow this tradition, adding a few features that increase the realism when working at the monthly sampling frequency. The model is cast in continuous time and all variables of the final model except the interest rates are expressed in logarithms. We adopt the notation $\dot{x}(t) = dx(t)/dt$ and denote the expectation of a variable with the superscript e. Now, consider a small open economy that produces one good which can be either consumed domestically or

[1]Genberg et al. (1987) use the following variables: foreign output, prices and interest rate; domestic output, prices, the interest rate and M1. Burbidge & Harrison (1985) and Kuszczak & Murray (1987) use the following variables: foreign output, prices, the interest rate and M1; the exchange rate, domestic output, prices, the interest rate and M1. Models comprising the (real) exchange rate, domestic output, prices, the interest rate and M1 have also been chosen in some exploratory work with recursive models and Finnish data (see the Bank of Finland publication on VAR models). Closed economy models comprising output, prices, the interest rate and M1 have been employed by Sims (1980b), Litterman & Weiss (1985) and Stock & Watson (1989).

exported. The country imports a foreign good, but the international price of that good is assumed to be fixed, and it is an imperfect substitute for the domestic good. The economy is specified by the following equations:

(6.1i) $y^d(t) = \alpha_1[e(t) - p(t)] - \alpha_2[i(t) - \dot{p}^e(t)]$

(6.1ii) $m^d(t) = p(t) + \alpha_3 y(t) - \alpha_4 i(t)$

(6.1iii) $m^s(t) = \bar{m},\ m^d(t) = m^s(t)$

(6.1iv) $\dot{p}\ (t) = \alpha_5\{\alpha_1[e(t) - p(t)] - \alpha_2[i(t) - \dot{p}^e(t)] - y(t)\}$

(6.1v) $y^s(t) = \alpha_6[i(t) - \dot{p}^e(t)] + \alpha_7[p(t) - p^e(t)]$

(6.1vi) $i\ (t) = i^*(t) + \dot{e}^e(t)$

(6.1vii) $\dot{y}\ (t) = \alpha_8[y^d(t) - y^s(t)]$

where $\alpha_i > 0$, $i = 1, \ldots, 8$ are constant parameters and

where $y^d(t)$ = demand for domestic real output
$e\ (t)$ = nominal exchange rate
$p\ (t)$ = domestic general price level
$i\ (t)$ = domestic instantaneous nominal interest rate
$m^d(t)$ = demand for domestic nominal money
$y\ (t)$ = domestic real output
$\bar{m}$ = constant supply of domestic nominal money in natural units
$m^s(t)$ = supply of domestic nominal money
$y^s(t)$ = supply of domestic real output
$i^*(t)$ = foreign instantaneous nominal interest rate.

Equation (6.1i) is an IS curve specifying that demand for domestic output depends positively on the real exchange rate and inversely on the expected real interest rate. The latter is defined as the difference between the instantaneous (nonnegative) nominal interest rate and the instantaneous expected rate of inflation. The demand for money yielding the LM curve given by equation (6.1ii) is a positive function of real output and an inverse function of the interest rate. In expression (6.1iii), we assume that money supply is constant at the level $\bar{m}$, and in the subsequent exposition the constant $\bar{m}$ will be eliminated through log-linearization by normalizing on $e^{\bar{m}} = 1 \Rightarrow \bar{m} = 0$. Money demand continuously equals money supply. Equation (6.1iv) specifies a Walrasian adjustment process in that domestic prices are assumed to be sticky in the short run while responding to excess demand in the market for the domestic good.

Equation (6.1v) describes the supply side of the economy. The Lucas-type supply behavior depends positively on the expected real rate of interest and price shocks. Alternatively, this supply behavior could be viewed as the outcome of wage rate behavior where contract wages are set to reflect the expected rate of inflation. Supply does not necessarily equal demand in the short run. Equation (6.1vi) is the interest rate parity condition. The economy is assumed to be small in the world capital market, and hence faces a given foreign interest rate.

In equation (6.1vii) we assert that output essentially responds to variations in demand in the short run (recall that the model is formulated with an eye to application to monthly observations). Equation (6.1vii) is hence congruent with the Walras/Keynes/Phillips view that in the very short run money wages and prices are given. We think of the possibility that $y^d(t) \neq y^s(t)$ in the short run as arising from output responding to (temporary) demand shocks through changes in inventories, but other reasons can also be cited (see Tobin (1975)).

We will assume rational expectations, which amounts to perfect foresight in our framework. Thus $\dot{p}^e(t) = \dot{p}(t)$ and $\dot{e}^e(t) = \dot{e}(t)$. Since we are working in continuous time, the perfect foresight expression for the, at times, discrete exchange rate (jump) variable is

$$(6.2) \qquad \dot{e}(t) = \delta(t - \bar{t})dE$$

where $\delta(t - \bar{t})$ is the Dirac delta function, $\bar{t} = 0$ is the period a devaluation is expected to occur, and $dE = e(1) - e(0)$ where $e(0)$ and $e(1)$ are the pre- and post-devaluation exchange rate levels, respectively.[2] Although nominally fixed, the Finnish exchange rate has varied considerably during our sample period in addition to frequent devaluations; i.e. during the first half of the 1970s it moved in a virtually freely floating fashion and in the 1980s in a similar fashion but within prescribed fluctuation limits.

Now turn to an investigation of the general properties of the economy described by equations (6.1i) - (6.1vii) and the auxiliary equation (6.2). Define the vector $\underset{\sim}{x}(t) = (y(t)\ p(t)\ e(t))'$ of endogenous variables and the vector $\underset{\sim}{u}(t) = (0\ 0\ i^*(t))'$ of exogenous variables. The model can then be expressed in the state space form

$$(6.3) \qquad \dot{\underset{\sim}{x}}(t) = S\underset{\sim}{x}(t) + Z\underset{\sim}{u}(t)$$

where S and Z are parameter matrices. Matrix S describes the internal dynamics and matrix Z shows how any exogenous events hit the model. More specifically,

$$(6.4) \qquad S = \begin{bmatrix} \alpha_{12} & \alpha_{13} & \alpha_{14} \\ \alpha_{15} & \alpha_{16} & \alpha_{17} \\ \alpha_{18} & \alpha_{19} & 0 \end{bmatrix} \text{ and } Z = \begin{bmatrix} 0 & 0 & 0 \\ 0 & 0 & 0 \\ 0 & 0 & \alpha_{20} \end{bmatrix}$$

$$\begin{aligned}
\text{where } \alpha_{12} &= -\alpha_8\alpha_{10}[\alpha_3\alpha_9 + \alpha_{11}(1 + \alpha_2\alpha_3\alpha_9)] && < 0 \\
\alpha_{13} &= -\alpha_8\{\alpha_1 + \alpha_{10}[\alpha_9 + \alpha_{11}(1 + \alpha_2\alpha_9)]\} && < 0 \\
\alpha_{14} &= \alpha_1\alpha_8(1 + \alpha_{10}\alpha_{11}) && > 0 \\
\alpha_{15} &= -\alpha_{11}(1 + \alpha_2\alpha_3\alpha_9) && < 0 \\
\alpha_{16} &= -\alpha_{11}(\alpha_1 + \alpha_2\alpha_9) && < 0 \\
\alpha_{17} &= \alpha_1\alpha_{11} && > 0
\end{aligned}$$

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[2] The Dirac delta function is defined by $\delta(t - \bar{t}) = 0$, $t \neq \bar{t}$, and $\int\delta(t - \bar{t})dt = 1$ (integrating over $-\infty$ to $+\infty$), $t = \bar{t}$.

$$\alpha_{18} = \alpha_3\alpha_9 \qquad > 0$$

$$\alpha_{19} = \alpha_9 \qquad > 0$$

$$\alpha_{20} = -1 \qquad < 0$$

where $\alpha_9 = \alpha_4^{-1} \qquad > 0$

$$\alpha_{10} = \alpha_2 + \alpha_6 \qquad > 0$$

$$\alpha_{11} = \alpha_5(1 - \alpha_2\alpha_5)^{-1} \qquad > 0$$

and $\alpha_2 \neq \alpha_5^{-1}$. In general, many of the signs of the above parameters are ambiguous, but the signs displayed above are those that materialize for reasonable basic parameter values.[3]

To assess issues of uniqueness and stability, define $\tilde{\underset{\sim}{o}}(t) = \underset{\sim}{o}(t) - \bar{\underset{\sim}{o}}$ for $\underset{\sim}{o}(t) = (\underset{\sim}{x}(t), \underset{\sim}{u}(t))$ where $\bar{\underset{\sim}{o}}$ denotes a steady state value, and note that

$$(6.5) \qquad e(t) = H(t - \bar{t})dE + e(0)$$

where $H(t - \bar{t})$ is the Heaviside unit step function.[4] Then system (6.3) in deviation form is

$$(6.6) \qquad \dot{\tilde{\underset{\sim}{x}}}(t) = S\tilde{\underset{\sim}{x}}(t) + Z\tilde{\underset{\sim}{u}}(t)$$

and we note that the exchange rate is the only state variable in the model which is allowed to make discrete jumps. Upon Laplace expansion of the Jacobian, it is immediately seen that the Jacobian

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[3]An example of a set of reasonable basic parameter values is $\alpha_1 = 0.4$, $\alpha_2 = 0.2$, $\alpha_3 = 1$, $\alpha_4 = 4$, $\alpha_5 = 3$, $\alpha_6 = 0.3$, $\alpha_7 = 0.7$ and $\alpha_8 = 1$. In the calculations of the present section, we will be referring to this particular set of parameter values when talking about reasonable basic parameter values.

[4]The Heaviside unit step function is defined by $H(t - \bar{t}) = 0$, $t < \bar{t}$, and $H(t - \bar{t}) = 1$, $t > \bar{t}$. Note that $dH(t - \bar{t})/dt = \delta(t - \bar{t})$.

does not vanish. Hence, a sufficient condition for a locally unique solution is established.

Turning to the stability issue, we focus on the characteristic polynomial of the Jacobian

$$(6.7) \qquad ch(\mu) = \mu^3 + \alpha_{21}\mu^2 + \alpha_{22}\mu + \alpha_{23}$$

where
$$\begin{aligned} \alpha_{21} &= -(\alpha_{12} + \alpha_{16}) && > 0 \\ \alpha_{22} &= \alpha_{12}\alpha_{16} - \alpha_{13}\alpha_{15} - \alpha_{14}\alpha_{18} - \alpha_{17}\alpha_{19} && < 0 \\ \alpha_{23} &= \alpha_{14}(\alpha_{16}\alpha_{18} - \alpha_{15}\alpha_{19}) + \alpha_{17}(\alpha_{12}\alpha_{19} - \alpha_{13}\alpha_{18}) && < 0 \end{aligned}$$

where μ denotes an eigenvalue and the parameter signs follow from reasonable basic parameter values. Although the first principal minor (α_{21}) of the Jacobian is positive, the second principal minor ($\alpha_{21}\alpha_{22} - \alpha_{23}$) does not emerge as positive for reasonable basic parameter values, hence violating the Routh-Hurwitz conditions (see Samuelson (1963)). However, since α_{21} = -tr(S) > 0 and α_{23} = -det(S) < 0, tr(S) < 0 and det(S) > 0, implying that the model dynamics are governed by one nonnegative and two negative eigenvalues. Thus, the two roots with negative real parts imply that a convergent saddle-point solution will exist (locally).

It is natural (and necessary) to associate the unstable eigenvalue of our model with the jumping variable, i.e. with the exchange rate. In that case, after an initial jump, the paths of the state variables will be governed by stable processes with the dynamics determined by the two stable roots. Convergence to the steady state is monotonous if the two stable roots are real while the dynamics will be oscillatory if the roots are complex. Note that oscillatory convergence implies overshooting, and that the probability of overshooting increases with the modulus of the constant term (α_{23}) of the characteristic polynomial (6.7). As it happens, reasonable parameter values imply the existence of three distinct, real eigenvalues, but numerical calculations show that overshooting is more likely e.g. the more sticky prices are (the smaller α_5 is).

Having characterized the uniqueness and stability of our model, we proceed to the analysis of the steady state equilibrium and the equilibrium paths. Without loss of generality, we assume that the origin is the steady state of the system. Further define an equilibrium path as a set of continuous paths for $\tilde{y}(t)$, $\tilde{p}(t)$ and $\tilde{e}(t)$ that converge to their steady state levels, given initial conditions outside the steady state. In steady state equilibrium we have from (6.6) $\dot{\underset{\sim}{x}}(t) = \underset{\sim}{0}$, implying $\underset{\sim}{\bar{x}} = -S^{-1}Z\underset{\sim}{\bar{u}}$, or more specifically,

$$\begin{bmatrix} \bar{y} \\ \bar{p} \\ \bar{e} \end{bmatrix} = \begin{bmatrix} 0 & 0 & \alpha_{25} \\ 0 & 0 & \alpha_{26} \\ 0 & 0 & \alpha_{27} \end{bmatrix} \begin{bmatrix} 0 \\ 0 \\ \bar{i}^* \end{bmatrix} \tag{6.8}$$

where $\alpha_{25} = -\alpha_9^{-1}\alpha_{24}(\alpha_{12} - \alpha_3\alpha_{24} - \alpha_{14}\alpha_{15}\alpha_{17}^{-1}) \quad > 0$

$\alpha_{26} = \alpha_9^{-1}(1 - \alpha_{18}\alpha_{25}) \quad > 0$

$\alpha_{27} = -\alpha_{17}^{-1}(\alpha_{15}\alpha_{25} + \alpha_{16}\alpha_{26}) \quad > 0$

where $\alpha_{24} = \alpha_{13} - \alpha_{14}\alpha_{16}\alpha_{17}^{-1} \quad < 0$

where the parameter signs follow from reasonable basic parameter values. We may note that (6.8) implies cointegration among output, prices, the exchange rate and the foreign interest rate - precisely what was established empirically in the preliminary data analysis of chapter 4.

The comparative statics of the steady state are

(6.9i,ii,iii) $\quad \partial\bar{y}/\partial\bar{i}^* = \alpha_{25} > 0 \quad \partial\bar{p}/\partial\bar{i}^* = \alpha_{26} > 0 \quad \partial\bar{e}/\partial\bar{i}^* = \alpha_{27} > 0.$

The real variable in equilibrium depends only on the (foreign) interest rate. An increase in the foreign interest rate will, given the constant domestic interest rate, generate expectations of a devaluation, thus boosting domestic output. Likewise, the perfect foresight character of our model implies that the expectation of a devaluation shows up positively in the price level and in the exchange rate.

The transition paths to the steady state can be characterized through the eigenvalues and corresponding eigenvectors of the model.[5] Since the eigenvalues of the model are real and pairwise distinct (under reasonable parametrizations of the model),

$$(6.10) \qquad \tilde{\underset{\sim}{x}}(t) = \sum_{i=1}^{3} \omega_i \sigma_i e^{\mu_i t}$$

where ω_i, i = 1,2,3, are arbitrary scalars chosen to set the components of $\underset{\sim}{x}(t)$ equal to their initial steady state values and $\sigma_i = (\sigma_{1i}\ \sigma_{2i}\ \sigma_{3i})'$ is an eigenvector associated with eigenvalue μ_i. Furthermore, let μ_1 be the unstable root, μ_2 be the smaller stable root and μ_3 be the larger stable root. Then ruling out speculative bubbles, i.e. confining the model trajectories to a subspace spanned by eigenvectors corresponding to stable eigenvalues, requires setting $\omega_1 = 0$. Satisfying this necessary condition for convergence implies that the eigenvalues μ_2 and μ_3 will dictate the behavior of any equilibrium path.

Using the normalization $\sigma_{23} = 1$, the eigenvector rays corresponding to the eigenvalue μ_2 are

$$(6.11i) \qquad \sigma_{12}/\sigma_{32} = (\mu_2 - \alpha_{12})^{-1}\{\alpha_{14} + \alpha_{13}(\alpha_{15}\mu_2 + \alpha_{17}\alpha_{18})x [\alpha_{15}\alpha_{19} + \alpha_{18}(\mu_2 - \alpha_{16})]^{-1}\} < 0$$

$$(6.11ii) \qquad \sigma_{22}/\sigma_{32} = (\alpha_{15}\mu_2 + \alpha_{17}\alpha_{18})[\alpha_{15}\alpha_{19} + \alpha_{18}(\mu_2 - \alpha_{16})]^{-1} > 0$$

$$(6.11iii) \qquad \sigma_{12}/\sigma_{22} = (\mu_2 - \alpha_{12})^{-1}(\alpha_{15}\mu_2 + \alpha_{17}\alpha_{18})^{-1}[\alpha_{15}\alpha_{19} + \alpha_{18}(\mu_2 - \alpha_{16})]\{\alpha_{14} + \alpha_{13}(\alpha_{15}\mu_2 + \alpha_{17}\alpha_{18})x [\alpha_{15}\alpha_{19} + \alpha_{18}(\mu_2 - \alpha_{16})]^{-1}\} < 0.$$

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[5]This approach has been applied by Levin (1981) to a two equation system and by Calvo (1987) to a system with three equations.

Using the normalization $\sigma_{33} = 1$, the dominant eigenvector rays corresponding to the eigenvalue μ_3 are

$$\text{(6.12i)} \quad \sigma_{13}/\sigma_{33} = (\mu_3 - \alpha_{12})^{-1}\{\alpha_{14} + \alpha_{13}(\alpha_{15}\mu_3 + \alpha_{17}\alpha_{18})\text{x} \\ [\alpha_{15}\alpha_{19} + \alpha_{18}(\mu_3 - \alpha_{16})]^{-1}\} \qquad > 0$$

$$\text{(6.12ii)} \quad \sigma_{23}/\sigma_{33} = (\alpha_{15}\mu_3 + \alpha_{17}\alpha_{18})[\alpha_{15}\alpha_{19} + \alpha_{18}(\mu_3 - \alpha_{16})]^{-1} \qquad < 0$$

$$\text{(6.12iii)} \quad \sigma_{13}/\sigma_{23} = (\mu_3 - \alpha_{12})^{-1}(\alpha_{15}\mu_3 + \alpha_{17}\alpha_{18})^{-1}[\alpha_{15}\alpha_{19} + \\ \alpha_{18}(\mu_3 - \alpha_{16})]\{\alpha_{14} + \alpha_{13}(\alpha_{15}\mu_3 + \alpha_{17}\alpha_{18})\text{x} \\ [\alpha_{15}\alpha_{19} + \alpha_{18}(\mu_3 - \alpha_{16})]^{-1}\} \qquad < 0.$$

Reasonable basic parameter values give rise to the signs reported in (6.11i) - (6.11iii) and (6.12i) - (6.12iii), and they furthermore imply $|\sigma_{12}/\sigma_{22}| < |\sigma_{13}/\sigma_{23}|$. The elements of the eigenvectors are in any case nonzero, thus yielding well defined rays. We note that (6.11i) and (6.12i) should be mapped in the y - e dimension, (6.11ii) and (6.12ii) in the p - e dimension and (6.11iii) and (6.12iii) in the y - p dimension. Given our assumption that μ_1 is the unstable and μ_3 is the larger stable eigenvalue, it can be shown using (6.10) that

$$\text{(6.13i)} \quad \lim_{t\to+\infty} [\tilde{y}(t)/\tilde{e}(t)] = \sigma_{13}/\sigma_{33}$$

$$\text{(6.13ii)} \quad \lim_{t\to+\infty} [\tilde{p}(t)/\tilde{e}(t)] = \sigma_{23}/\sigma_{33}$$

$$\text{(6.13iii)} \quad \lim_{t\to+\infty} [\tilde{y}(t)/\tilde{p}(t)] = \sigma_{13}/\sigma_{23}$$

which will determine the equilibrium paths together with the rays $\dot{\tilde{y}}(t) = 0$, $\dot{\tilde{p}}(t) = 0$ and $\dot{\tilde{e}}(t) = 0$, the slopes of which are $\alpha_{12} < 0$, $\alpha_{16} < 0$ and 0, respectively. It may also be pointed out that the trajectories may not cross the rays given by expressions (6.11i) -

(6.11iii) corresponding to the smaller stable eigenvalue. The motion of the system is portrayed in the phase diagrams in Figure 6.1.

In Figure 6.1a the dynamics of the model are viewed in the y - e plane. The eigenvector ray - to which the ratio $\tilde{y}(t)/\tilde{e}(t)$ converges - is positively sloped whereas the gradient for the line of stationarity of the exchange rate - which determines the direction of movement on either side of that line - coincides with the e-axis. If the initial conditions are as at point A, the transition path will necessarily behave as depicted by the arrowed curve. A depreciation of the exchange rate is seen to have an expansionary effect on output. This conventional result comes about in an equally conventional manner through the demand side of the model.

Figure 6.1b portrays the equilibrium path in the price-exchange rate dimension. Here the dominant eigenvector ray is negatively sloped and the gradient for the line of stationarity for the exchange rate again coincides with the e-axis. Letting the initial conditions be as in point B, the arrowed trajectory traces out the effect of a price level rise on the exchange rate. During the transition to the steady state, the exchange rate is required to depreciate to counteract the reduction in demand. The price level rise, on the other hand, also reduces supply through the decrease in the real interest rate, but in our basic parameter constellation (see footnote 3 of this chapter) output demanded is more sensitive to changes in the real exchange rate than is output supplied to a change in the real rate of interest.

FIGURE 6.1 Equilibrium paths in a theoretical model of a small open economy

FIGURE 6.1a Dynamics in output-exchange rate space

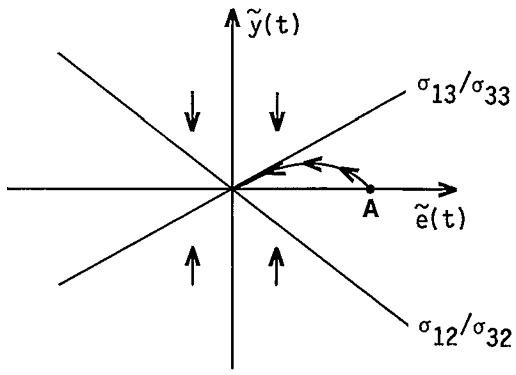


FIGURE 6.1b Dynamics in price-exchange rate space

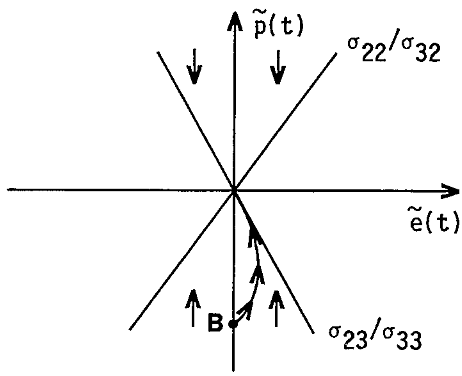
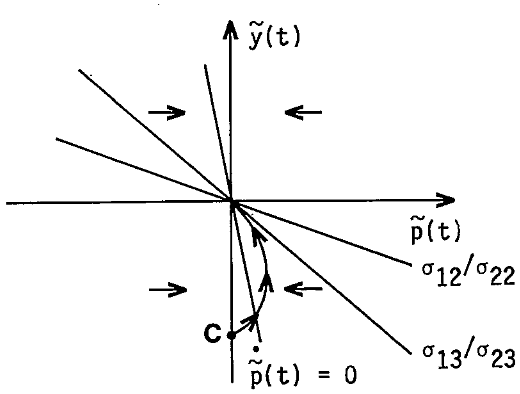


FIGURE 6.1c Dynamics in output-price space



In Figure 6.1c the motion of the system is depicted in the output-price dimension. Both the eigenvector rays and the gradient for the line of stationarity for prices are negatively sloped, but for a reasonable parametricization of the model the dominant eigenvector ray is more negatively sloped than the nondominant ray, while the stationarity line is more negatively sloped than both the eigenvector rays. From initial conditions, such as at point C, output will start to rise unambiguously, raising the price level during the transition to equilibrium. Such positive comovement between output and prices certainly enjoys empirical support, yet conventional Dornbuschian models tend to exhibit a negative relationship. The key difference in this respect would seem to be that in our model, output is not completely demand-determined. Rather, output reacts to demand <u>and</u> to supply.

We now proceed toward the empirical assessment by stating what restrictions our theoretical model might place on the matrix Γ_0 describing contemporaneous relationships in the structural VAR model. The rationale for focusing attention exclusively on contemporaneous interactions is that our theoretical model - like most models - does not provide guidance about lagged behavioral relationships; such restrictions could, in principle, be imposed through Bayesian methods.

With five variables, we are free to specify up to 10 elements of the matrix Γ_0 in order to retain an identified model. Recalling equations (6.1i) through (6.1vii) and representations (6.6) and (6.8), we can make the following observations. In our model, as in many other models, it will be hard to disentangle narrowly defined demand and supply shocks. However, what we can do is to separate money market shocks from asset market and supply shocks. It is also possible to isolate shocks to the price level. Moreover, the foreign interest rate is exogenous (at least contemporaneously). We specify the Finnish exchange rate to be exogenous in the short run reflecting the fact that the authorities can make discretionary adjustments in the exchange rate at will. The Finnish exchange rate is endogenous in the long run. The above considerations suggest the following set of contemporaneous relationships

$$\begin{array}{llll}
(6.14\text{i}) & y(t) & = \gamma_1 p(t) + \gamma_2 i^*(t) + \gamma_3 e(t) + & \xi^y(t) \\
(6.14\text{ii}) & p(t) & = \gamma_4 i^*(t) + & \xi^p(t) \\
(6.14\text{iii}) & m(t) & = \gamma_5 y(t) + \gamma_6 p(t) + \gamma_7 i^*(t) + & \xi^m(t) \\
(6.14\text{iv}) & i^*(t) & = & \xi^{i*}(t) \\
(6.14\text{v}) & e(t) & = & \xi^e(t)
\end{array}$$

where the $\xi(t)$s represent structural shocks. We can interpret $\xi^y(t)$ as a goods market or an aggregate supply shock, where the aggregate supply shock comprises technology and labor supply shocks. $\xi^p(t)$ can be interpreted as a price level shock and $\xi^m(t)$ as a money market shock. Clearly, $\xi^{i*}(t)$ is a shock to the foreign interest rate and $\xi^e(t)$ is a shock to the exchange rate. System (6.14i) - (6.14v) implies the following Γ_0 matrix

$$(6.15) \qquad \Gamma_0 = \begin{bmatrix} 0 & \gamma_1 & 0 & \gamma_2 & \gamma_3 \\ 0 & 0 & 0 & \gamma_4 & 0 \\ \gamma_5 & \gamma_6 & 0 & \gamma_7 & 0 \\ 0 & & \cdots & & 0 \\ 0 & & \cdots & & 0 \end{bmatrix}$$

which we will use in the next section to identify the structural shocks. When the model is augmented with one of the auxiliary variables, the modified matrix of contemporaneous relationships Γ_0^a, where a = y^*, p^*, g or p^0, can be specified as

$$
(6.16) \qquad \Gamma_0^a = \left[\begin{array}{ccccc|cccc}
 & & & & & \gamma_{11} & \gamma_{13} & \gamma_{17} & \gamma_{22} \\
 & & & & & \gamma_{12} & \gamma_{14} & \gamma_{18} & \gamma_{23} \\
 & & \Gamma_0 & & & 0 & \gamma_{15} & 0 & \gamma_{24} \\
 & & & & & 0 & \gamma_{16} & 0 & \gamma_{25} \\
 & & & & & 0 & 0 & \gamma_{19} & 0 \\
\hline
0 & . & . & . & 0 & & & & \\
0 & . & . & . & 0 & & & 0 & \\
\gamma_{20} & \gamma_{21} & 0 & 0 & 0 & & & & \\
0 & . & . & . & 0 & & & &
\end{array}\right]
$$

where the notation should be read to indicate that one variable in turn - in the order y^*, p^*, g, p^0 - has been grafted onto the model.

Foreign output is hypothesized to contemporaneously influence only domestic output and the domestic price level (among the core variables). Contemporaneous feedback from domestic variables to foreign output is precluded. The foreign price level exhibits instantaneous two-way causality with the foreign interest rate, and, in addition, influences domestic output, prices, and money. Fiscal policy contemporaneously affects, and is affected by, domestic output and prices, and the exchange rate is specified to react to fiscal policy. The equation for fiscal policy can be interpreted as a reaction function for the fiscal authorities. The price of oil is exogenous, and it affects all other variables except the exchange rate contemporaneously.

The model outlined above should be thought of as the basic, general framework within which the empirical analyses will be made. It is basic since we are restricted in the empirical analyses to using only a very small number of variables. The model is general because our methodology calls for the imposition of only a minimum amount of a priori structure. Generality applies in a theoretical sense as well. It includes rigidities and potentially allows a role for policy as Keynesian models typically do. On the other hand, the model has neoclassical features and displays neither long-run nonneutralities nor adaptive expectations formation, and it includes full output and price flexibility as a special case.

6.2 Empirical Evidence

6.2.1 Estimation Results

In the operationalization of the output variable we are, again, confined to the use of (real) industrial production. The countries comprising the aggregate of foreign output are Belgium, France, Germany, Italy, Japan, the Netherlands, Sweden, the U.K. and the U.S. We use moving weights based on the importance of the respective economy for Finnish exports (see Appendix 1 for details). We use consumer price indices as measures of the general price level; the foreign price level is constructed as a weighted sum of foreign price levels as in the case of output.

The money variable offers more material for discussion. Empirical business cycle research has almost invariably used (nominal) narrow money, i.e. M1.[6] When analyzing the Finnish economy some additional considerations related to credit rationing may arise. If liquidity is constrained, M1 may be a fairly uninteresting variable, whereas the availability of credit may be more important. Our output measure is stronger related to firms than to the more heavily rationed households, but the credit rationing issue cannot be dismissed on a priori grounds. Therefore, we will also report analysis based on some indicators of the availability of credit. Such indicators are total credit advanced to the public and the marginal interest rate on funds advanced by the central bank to commercial banks. The former indicator is most appropriate under conditions of widespread

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[6]Empirical evidence on money-income causality is mixed; see Stock & Watson (1989) and the references cited therein. No consensus has emerged, and given the importance of the question and the spirit of our approach, we will not exclude money from the model on a priori grounds. However, if money affects output, it most certainly is M1 we should be focusing on (see footnote 1 of this chapter).

credit rationing, while the latter allows for heterogenous credit conditions. These indicators will furthermore reflect the monetary policy stance.

In constructing an empirical interest rate, we follow earlier studies by considering short-term nominal interest rates. The foreign interest rate is a weighted average of foreign short-term interest rates, as were the output and price level variables. The robustness of the conclusions with respect to the choice of nominal as opposed to real interest rates will be evaluated. Similarly, we will follow the literature in using the nominal rather than the real exchange rate, but again supplying robustness analysis.

While the construction of monetary policy variables is relatively unproblematic, the same cannot be said about indicators of fiscal policy. This is because the construction of an indicator of fiscal policy is inherently difficult (see the calculations presented in Blanchard & Watson (1986)), and because of data availability issues. Theoretical difficulties in constructing an index of fiscal policy aside, we would want to consider measures such as government spending, government debt and various tax rates. However, monthly observations on these variables do not exist in the Finnish case. Instead, we will use the (real) central government net borrowing requirement as a proxy for the fiscal stance, since genuine realizations of this quantity are available for most of the sample period. Consequently, our conclusions regarding the contribution of fiscal policy to aggregate variability can be neither strong nor general.

Finally, the specification of the price of oil deserves some comments. Monthly observations on the price of oil are available from the late 1960s onwards only, but an import price index for mineral fuels (of which oil constitutes the major component) is available for the whole sample period. Both of these measures will be employed in the empirical analyses. Following the important VAR study by Hamilton (1983), we will use the nominal price of oil in our analyses. While a relative price is used in theoretical analyses, one may argue that it is the nominal oil price that tracks what historically are called oil shocks. However, the robustness of our conclusions to the oil price operationalization and the nominal versus real issue will be investigated.

Turning to estimation, we begin by analyzing how many lags are needed to capture the dynamics of the basic model. All variables are in logarithmic form, except i* which is expressed as a percentage and g which is in billions of FIM. In chapter 4 it was found that the variables are integrated of order one, and hence the model is estimated in first differences. Error correction terms are not included in the model at this stage of the analysis, since in chapter 4 they were found to carry essentially zero weight in all model equations. In addition, we include 11 seasonal dummy variables, a constant and a strike dummy for the strike in the metal and engineering industry in 1971M2 - 1971M3. The use of deterministic seasonals is based on the empirical results obtained in chapter 4, and the strike dummy is added because of a priori knowledge and concern about the effects of influential observations on inference.

We use the likelihood ratio test statistic of Sims (1980a) to test for system-wide significance of lags. Results for one to 12 lags are presented in Table 6.1. As it happens, the tests for lag length do not offer any unanimous indication of which order to use. A low order is admissible, but a relatively high order may also be congruent with the attained likelihoods (similar results were obtained when the foreign price level and the price of oil were added to the model). We choose to proceed with a lag length of order three, since this yields a parsimonious representation, yet allows for trending as well as cyclical behavior in univariate data generating processes (three monthly lags were also chosen by Burbidge & Harrison (1985)). The sensitivity of our conclusions with respect to the choice of lag length will be evaluated.

TABLE 6.1

Tests for lag length in a VAR model for the vector $\underset{\sim}{y}_t = (y_t\ p_t\ m_t\ i_t^*\ e_t)'$ using data from the period 1960M1 - 1988M8

Lag length	1	2	3	4	5	6	7	8	9	10	11	12
1												
2			0.074	0.022		0.001						
3		0.035		0.060	0.001	0.003		0.001	[0]	0.001		
4		0.004	0.024		0.001	0.007	0.001	0.002	0.001	0.001		
5						0.350	0.048	0.103	0.013	0.032		
6				0.001	0.183		0.022	0.079	0.023	0.010		
7					0.005	0.004		0.505	0.125	0.113		
8					0.007	0.008	0.270		0.048	0.057		
9			[0]				0.013	0.007		0.242		
10							0.004	0.003	0.064			
11												
12												

The test is the likelihood ratio test statistic of Sims (1980a). The upper half of the Table reports marginal significance levels using a degrees of freedom correction, and the lower half concerns testing without a degrees of freedom correction. [0] indicates that values smaller than 0.0005 were obtained. In addition to the components of $\underset{\sim}{y}_t$, the model includes a constant, 11 seasonal dummy variables and a strike dummy taking the value 1 in 1971M2 - 1971M3 and 0 otherwise.

To determine whether any one of the auxiliary variables belongs in the model, we make use of multivariate Granger-Sims causality tests. These tests have as their null hypothesis that the lags of one variable do not enter into the equations of the remaining variables. While we recognize that causality testing is treacherous, the current tests nevertheless offer a formal and objective way to test for the significance of augmentations to the basic model.

Granger-Sims tests are computed using the full sample as well as for the three subperiods. As a matter of sensitivity analysis, tests using alternative measures for the money variable, the exchange rate and the price of oil will be reported as well. The (nominal) money variable is alternatively operationalized as M1 and total credit advanced to the public c, the exchange rate as the nominal exchange rate e and the real exchange rate e^r, and the (nominal) price of oil as the price of oil p^{01} and the import energy price index p^{02}. Causality tests are performed for each of the components of the basic model as well as for the auxiliary variables. This serves to highlight the individual role of variables when different configurations of the total set of variables are used. The empirical evidence is presented in Table 6.2.

TABLE 6.2

Multivariate causality tests

Time period	y	p	m/c	i*	e/e^r	y*	p*	g	p^{01}	p^{02}
			m		e					
1960M1 - 1988M8	0.834	0.190	0.666	0.039	0.238	0.145	0.002	n.a.	n.a.	0.264
1960M1 - 1972M12	0.733	0.913	0.460	0.763	0.002	0.567	0.930	n.a.	n.a.	0.001
1973M1 - 1988M8	0.694	0.271	0.360	0.007	0.157	0.076	0.001	0.070	0.334	0.525
1980M1 - 1988M8	0.088	0.271	0.808	0.152	0.070	0.487	0.078	0.261	0.942	0.938
			c		e					
1960M1 - 1988M8	0.981	0.334	0.173	0.138	0.253	0.107	0.004	n.a.	n.a.	0.396
1960M1 - 1972M12	0.624	0.985	0.000	0.965	0.004	0.939	0.909	n.a.	n.a.	0.049
1973M1 - 1988M8	0.696	0.172	0.600	0.015	0.126	0.125	0.001	0.270	0.342	0.478
1980M1 - 1988M8	0.097	0.044	0.735	0.981	0.024	0.419	0.085	0.775	0.941	0.988
			m		e^r					
1960M1 - 1988M8	0.924	0.038	0.690	0.056	0.019	0.140	0.010	n.a.	n.a.	0.252
1960M1 - 1972M12	0.777	0.883	0.541	0.862	0.002	0.569	0.935	n.a.	n.a.	0.005
1973M1 - 1988M8	0.633	0.086	0.236	0.027	0.023	0.199	0.001	0.059	0.376	0.541
1980M1 - 1988M8	0.084	0.141	0.790	0.139	0.015	0.538	0.194	0.165	0.945	0.947
			c		e^r					
1960M1 - 1988M8	0.996	0.143	0.250	0.175	0.022	0.093	0.017	n.a.	n.a.	0.392
1960M1 - 1972M12	0.677	0.974	0.000	0.984	0.004	0.951	0.904	n.a.	n.a.	0.038
1973M1 - 1988M8	0.663	0.101	0.651	0.059	0.012	0.173	0.001	0.249	0.369	0.457
1980M1 - 1988M8	0.106	0.031	0.703	0.064	0.001	0.346	0.181	0.591	0.952	0.989

Marginal significance levels for multivariate Granger-Sims causality tests are reported. Tests to the left of the vertical line refer to one variable conditional on the remaining four variables to the left of the vertical line. Tests to the right of the vertical line refer to one variable conditional on the five variables to the left of the vertical line. All models include a constant, 11 seasonal dummy variables and, where appropriate, a strike dummy taking the value 1 in 1971M2 - 1971M3 and 0 otherwise.

For the full sample, the following results emerge. Of the auxiliary variables, only the foreign price level p* seems to be statistically significant. In fact, the importance of this variable receives strong support from the data. As expected in view of the results of chapter 5, foreign output y* plays a statistically minor role. Only in the constellation with the real exchange rate e^r can the null hypothesis of no influence be rejected at the 10 per cent level. Perhaps somewhat surprisingly, the price of oil does not turn out to be statistically important. This conclusion also holds for the real import energy price, but because of data availability, the genuine price of oil variables could not be employed in these analyses. All results for the auxiliary variables are robust to the operationalization of the money and exchange rate variables.

With respect to the core variables, results differ according to the specification of the money and exchange rate variables. In our basic empirical counterpart, only the foreign interest rate i* is exogenous. This is a plausible result, and it is in fact the only exogeneity result one should expect for this set of variables. Using credit as the money variable has the effect of making the foreign interest rate an endogenous variable, and using the real exchange rate renders the domestic price level an exogenous variable. Neither of these results make economic sense in the current setting. Hence, we interpret the results for the full sample to indicate that subsequent analysis of the full sample should be primarily conducted with the basic model operationalization augmented with foreign prices.

Turning to subsample results, the following picture emerges. On the whole, results differ from full-sample results. Conclusions regarding auxiliary variables are now robust to the operationalization of the money and exchange rate variables only at the five per cent level of significance; some sensitivity is encountered if the 10 per cent norm is used. Results for the core variables again differ when alternative operationalizations of these variables are employed.

During the Bretton Woods era, the price of import energy is found to play a statistically significant role (the same holds for the real version of this variable). While this result may at first glance seem unexpected, corresponding results from other subperiods - on which we will elaborate below - may offer an explanation for this finding. Among the core variables, the exchange rate is clearly significant; the exogeneity of this variable reflects its use as a policy variable and its corresponding discrete movements. This observation holds whether one focuses on the nominal or on the real exchange rate.

Lastly, and not very surprisingly, the credit variable turns out to be the most relevant monetary variable during the 1960s and early

1970s.[7] While constituting somewhat crude evidence, this finding points to the prevalence of fairly extensive credit rationing in Finland during the 1960s and early 1970s. It is, moreover, in accordance with casual observation and the sparse empirical evidence that exists on this matter. Hence, we will operationalize the money variable as total credit advanced to the public in subsequent analyses of the Bretton Woods period, augmenting the basic model with the price of oil in these analyses.

During the post-Bretton Woods period only the foreign price level appears to be a statistically important auxiliary variable (at the five per cent level). However, judging by the 10 per cent norm, foreign output and the fiscal stance (in fact, also the domestic interest rate) play significant roles when the money variable is operationalized as M1. None of the four operationalizations of the price of oil appears to belong to the model even at the 30 per cent level of significance. These conclusions are robust to the operationalization of the money and exchange rate variables.

The different results across subperiods for the significance of the price of oil may be a consequence of the differing importance over time of the trade with COMECON countries (mainly the Soviet Union). The bilateral trade agreements require that the bilateral trade balance is basically zero, and since the main import item from the Soviet Union is oil, movements in the price of oil are intrinsically linked to trade developments. Specifically, when the price of oil rises, Finnish exports to the Soviet Union can rise, and conversely when the price of oil falls. Thus, a rise in the price of oil is in one sense "beneficial" for the Finnish economy. Moreover, trade with

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[7]Treating credit as an auxiliary variable and keeping narrow money in the basic configuration yields the same conclusion; this measure of monetary policy is significant during the 1960s and the early 1970s. Conversely, treating narrow money as an auxiliary variable confirms the endogeneity of that variable. The marginal interest rate on funds advanced by the central bank to commercial banks (expressed as a percentage) did not belong to the model whether substituting for the foreign interest rate or treated as an auxiliary variable. These results are robust to the operationalization of the exchange rate and the price of oil.

eastern countries has been countercyclical to trade with western countries during the 1970s and 1980s, thus smoothing output fluctuations (see Forsman & Haaparanta (1988) for further theoretical and empirical arguments in favor of this view). As Finnish trade with COMECON countries was modest during the first subsample and of considerable importance during the second subsample, the nonsignificance of the price of oil may have a plausible economic explanation.[8]

Among the core variables, the foreign interest rate stands out as the most important variable. The real, as opposed to nominal, exchange rate is also of statistical importance during the latter part of the 1970s and most of the 1980s. However, the use of the real exchange rate again tends to render the domestic price level exogenous. Credit is of negligible statistical importance during the post-Bretton Woods period. We emerge from the analysis of the 1973M1 - 1988M8 period with the conclusions that the basic parametrization should be favored but augmented with foreign prices. However, sensitivity analysis with respect to the importance of foreign output, the fiscal stance (and the domestic interest rate) and the operationalization of the exchange rate also seem warranted. Given the possible economic, as opposed to statistical, significance of the price of oil, different operationalizations of the oil price variable will also be considered.

The results for the third subperiod 1980M1 - 1988M8 should be interpreted with care because of the relatively small number of observations. Among the auxiliary variables, the foreign price level again stands out as being statistically significant; among the core variables support for the importance of the exchange rate is

[8]The insignificance of the price of oil has also been documented by Virén (1986) in a regression analysis of the period 1970 - 1983. The importance of the bilateral trade agreement for the issue at hand has been documented by Lehtinen(1988) through simulations of the Bank of Finland quarterly model. Both authors find a structural break after the first oil price shock, while none is found by the former author and some signs are found by the latter author in connection with the second oil price shock.

obtained. The unusually stable Finnish developments in output during the 1980s show up in an almost exogenous output characterization. On the whole, however, results are lacking in uniformity, which together with sample size considerations leads us to favor the use of the basic operationalization of the model primarily without augmenting variables in subsequent analyses.

As a summary of the conclusions from the tests for auxiliary variables and different operationalizations of variables, we recapitulate the basic variable vectors that will primarily be used in the subsequent analyses. The vectors for different samples are: 1960M1 - 1988M8; $\underset{\sim}{y} = (y\ p\ m\ i^*\ e\ p^*)'$, 1960M1 - 1972M12; $\underset{\sim}{y} = (y\ p\ c\ i^*\ e\ p^{\tilde{O}2})'$, 1973M1 - 1988M8; $\underset{\sim}{y} = (y\ p\ m\ i^*\ e\ p^*)'$, and 1980M1 - 1988M8; $\underset{\sim}{y} = (y\ p\ m\ i^*\ e)'$ (all vectors and their components refer to contemporaneous values).

In order to estimate the contemporaneous relationships (6.15) augmented with foreign prices according to (6.16), we first need to estimate the model in unconstrained form. Moreover, the unconstrained estimates may yield insights into the main channels of dynamics of the model, into issues of long-run neutralities and into the sensitivity of results with respect to the specification of contemporaneous relationships. Equationwise OLS estimates of the model based on the vector $\underset{\sim}{y}_t = (y_t\ p_t\ m_t\ i^*_t\ e_t\ p^*_t)'$ using data from the period 1960M1 - 1988M8 are presented in Table 6.3.

The parameter estimates are functions of the structural parameters, thus rendering interpretation of reduced form parameters impossible. We can, however, focus on the location of significant explanatory power in the model. Some care should be exercised in interpreting tests of significance since the estimated residuals are nonnormally distributed and in some cases temporally dependent, and parameter constancy can be rejected in some of the model equations.

TABLE 6.3

Estimation results for a VAR model for the vector $\underset{\sim}{y}_t = (y_t\ p_t\ m_t\ i_t^*\ e_t\ p_t^*)'$ using data from the period 1960M1 - 1988M8

Regressors/ Statistics	Regressands					
	y_t	p_t	m_t	i_t^*	e_t	p_t^*
y	-1.06	-0.008	-0.036	0.002	0.001	-0.007
	(-8.12)	(-0.56)	(-0.40)	(0.15)	(0.02)	(-0.93)
p	-0.744	0.305	0.027	-0.077	0.103	0.181
	(-0.96)	(3.56)	(0.05)	(-1.02)	(0.51)	(3.83)
m	-0.130	0.019	-0.542	-0.015	-0.013	0.291
	(-0.68)	(0.90)	(-4.14)	(-0.79)	(-0.27)	(2.50)
i*	0.350	0.178	0.251	0.270	-0.281	0.130
	(0.41)	(1.86)	(0.42)	(3.21)	(-1.24)	(2.46)
e	-0.284	0.096	0.010	0.009	0.191	0.011
	(-0.94)	(2.86)	(0.05)	(0.30)	(2.40)	(0.58)
p*	0.211	0.579	0.465	0.093	-0.250	0.543
	(0.19)	(4.62)	(0.60)	(0.85)	(-0.84)	(7.85)
$\bar{R}^2$	0.933	0.294	0.431	0.090	0.101	0.473
Q	0.000	0.305	0.000	0.005	0.999	0.029
JB	0.000	0.000	0.000	0.000	0.000	0.000
CHOW(1973M1)	0.000	0.003	0.115	0.188	0.941	0.001
CHOW(1980M1)	0.777	0.308	0.002	0.059	0.981	0.762

$$\Omega_\zeta = \begin{array}{c} \zeta^y \\ \zeta^p \\ \zeta^m \\ \zeta^{i^*} \\ \zeta^e \\ \zeta^{p^*} \end{array} \begin{bmatrix} & -0.021 & 0.005 & 0.155 & -0.010 & -0.024 \\ & & -0.032 & -0.027 & -0.019 & 0.151 \\ & & & -0.081 & 0.083 & 0.086 \\ & & & & -0.062 & 0.006 \\ & & & & & 0.005 \\ & & & & & \end{bmatrix}$$

The model is estimated by equation-by-equation unconstrained OLS from first-differenced data. In addition to the six elements of $\underset{\sim}{y}_t$, the model includes a constant, 11 seasonal dummy variables and a strike dummy taking the value 1 in 1971M2 - 1971M3 and 0 otherwise. Estimation results for the deterministic components have been suppressed. Estimates on the rows for y, p, m, i*, e and p* are the sums of the estimated coefficients on lags 1 through 3 of a variable, and below each sum a t statistic for a test that the sum is 0 is given. $\bar{R}^2$ is the degrees-of-freedom-corrected squared multiple correlation coefficient, Q is the Ljung & Box (1978) statistic based on six autocorrelations and JB is the Jarque & Bera (1980) statistic. CHOW(.) is the Chow test with the date of the supposed structural break given in parentheses. Marginal significance levels are reported for Q, JB and CHOW. Ω_ζ denotes the correlation matrix of estimated residuals. Twice the asymptotic standard error of the residual correlations is 0.108.

From Table 6.3 it is apparent that own-variable lags typically account for the bulk of the explanatory power. This holds in particular for m, i* and e. In the equation for y, some individual coefficients on the variables p and p* appear to be statistically significant (estimates not shown), but the temporal dependency and nonnormality of the estimated disturbances imply a caveat to these findings. As the sums of coefficients for the nominal variables in the equation of the real variable appear to be statistically indistinguishable from zero, we have an indication of the long-run neutrality of these nominal variables. It is premature to state this result as a conclusion, however, since we are here relying on astructural representations, possibly inefficient test statistics and only one variant of the basic model. The next sections present more careful analyses of these matters.

We find substantial dynamic interactions between the price, money, interest rate and exchange rate variables. The significance of the sums of coefficients of e and p* in the equation for p suggests that movements in the exchange rate and the foreign price level can cause permanent shifts in the domestic price level. Nonzero sums of own-variable lags give rise to the possibility of permanence in the effects of own-variable disturbances as well. The caveats mentioned above to conclusions regarding nonneutrality should, however, be borne in mind. Heteroskedasticity is not a problem judging by heteroskedasticity-consistent t values (White (1980); estimates not shown).

Stability can be rejected according to Chow tests in four or five of the six model equations. The most obvious rejection occurs in the equation for output when the structure is hypothetized to have changed in 1973. Strong rejections of the stability of the equations determining foreign and domestic prices with respect to the breaking point 1973 can also be documented. When the breaking point is in 1980, stability can be rejected in the equation for money, while stability is rejected in the equation for the foreign interest rate at a marginal significance level of 0.059. Given that slightly different models were found to be adequate across subsamples, the rejections of stability are natural, but it remains an empirical matter how much impulse responses and decompositions of variance differ across subsamples.

Finally, we note that the correlation matrix of estimated disturbances is nondiagonal even though monthly data are used. Two of the 15 unique off-diagonal elements differ from zero at the 1 per cent level of significance. The disturbances from the output and foreign interest rate equations as well as the domestic and foreign price disturbances are, respectively, mildly positively correlated (these findings are robust to the inclusion of an oil price variable). While being enough to blur the interpretation of impulse responses, decompositions of variance etc. from astructural VAR models, the relatively low number of significant off-diagonal elements and the low absolute size of these elements may, however, mitigate the problem in a certain sense. One could conjecture that impulse responses and decompositions of variance will not be overly sensitive to the specification of contemporaneous relationships. The accuracy of this conjecture will be documented below.

6.2.2 Analysis of Structural Shocks

To obtain estimates of structural shocks, we need to estimate the elements of the matrix of contemporaneous relationships Γ_0. If the matrix Ω_ζ of estimated reduced form disturbances was diagonal, we could estimate the elements of Γ_0 simply by applying OLS to individual model equations with unlagged variables included. When, as in the current case, Ω_ζ is nondiagonal, we can obtain estimates of Γ_0 and Ω_ξ by solving the likelihood-based problem

$$(6.17)\quad \min_{\{\Gamma_0,\Omega_\xi\}}\{2\ln(\det(I-\Gamma_0)) + \ln(\det(\Omega_\xi)) + \mathrm{tr}(\Omega_\xi^{-1}(I - \Gamma_0)\Omega_\zeta(I - \Gamma_0)')\}.$$

Assuming that Ω_ξ is diagonal, the first-order (necessary) conditions are quadratic, and the Newton-Raphson method can be employed to solve for the elements of Γ_0 and Ω_ξ. Estimates of Γ_0 for the augmented model using the full sample are displayed in Table 6.4 (for further reference, the table also contains estimates based on alternative specifications).

TABLE 6.4

Contemporaneous structural relationships in VAR models addressing nonstationarity in different ways using data from the period 1960M1 - 1988M8

Model variables	Matrix of contemporaneous relationships
Model estimated in first differences	

$$\underset{\sim}{y}_t = \begin{bmatrix} y_t \\ p_t \\ m_t \\ i^*_t \\ e_t \\ p^*_t \end{bmatrix} \qquad \Gamma_0 = \begin{bmatrix} 0 & -0.116 & 0 & 1.59 & -0.001 & -0.367 \\ 0 & 0 & 0 & -0.032 & 0 & 0.273 \\ 0.013 & -0.297 & 0 & -0.605 & 0 & 1.06 \\ 0 & 0 & 0 & 0 & 0 & 0.009 \\ 0 & & & \cdots & & 0 \\ 0 & & & \cdots & & 0 \end{bmatrix}$$

Model estimated in levels

$$\underset{\sim}{y}_t = \begin{bmatrix} y_t \\ p_t \\ m_t \\ i^*_t \\ e_t \\ p^*_t \end{bmatrix} \qquad \Gamma_0 = \begin{bmatrix} 0 & 0.017 & 0 & 1.46 & 0.096 & -0.180 \\ 0 & 0 & 0 & -0.015 & 0 & 0.234 \\ 0.013 & -0.172 & 0 & -0.632 & 0 & 1.03 \\ 0 & 0 & 0 & 0 & 0 & 0.050 \\ 0 & & & \cdots & & 0 \\ 0 & & & \cdots & & 0 \end{bmatrix}$$

Model estimated in error correction form

$$\underset{\sim}{y}_t = \begin{bmatrix} y_t \\ p_t \\ m_t \\ i^*_t \\ e_t \\ p^*_t \end{bmatrix} \qquad \Gamma_0 = \begin{bmatrix} 0 & 0.075 & 0 & 1.40 & 0.083 & -0.300 \\ 0 & 0 & 0 & -0.011 & 0 & 0.190 \\ 0.154 & -0.241 & 0 & -0.659 & 0 & 0.888 \\ 0 & 0 & 0 & 0 & 0 & 0.049 \\ 0 & & & \cdots & & 0 \\ 0 & & & \cdots & & 0 \end{bmatrix}$$

In addition to the components of $\underset{\sim}{y}_t$, the models include a constant, 11 seasonal dummy variables and a strike dummy taking the value 1 in 1971M2 - 1971M3 and 0 otherwise.

Inspection of the first estimate in Table 6.4 of the matrix of contemporaneous relationships in our augmented model reveals the following features. At the monthly level, output has historically reacted slightly negatively to changes in the domestic price level, and somewhat more negatively to changes in the foreign price level. In one interpretation, inflation has, on average, been harmful to real output developments in the short run. The dependence of output on the foreign interest rate is positive, while the contemporaneous impact of a change in the exchange rate is negligible. The former result suggests that historically output has been high in Finland when the interest rate has been high in the rest of the world. The nonsensitivity of output to within-month changes in the exchange rate may reflect adjustment lags or forward looking behavior by firms relying on the devaluation cycle, or both.

The domestic price level is contemporaneously positively related to foreign prices. This highlights the openness of the Finnish economy and the crucial, even contemporaneous, dependence of domestic variability on foreign variability. The contemporaneous dependence of the domestic price level on the foreign interest rate is practically nonexistent (this also holds for the domestic price level versus the domestic interest rate). The contemporaneous connection between the foreign interest rate and the foreign price level is positive albeit practically nonexistent.

Money reacts positively (albeit weakly) to income, negatively to the interest rate and in a roughly one-to-one fashion to the (foreign) price level. Oddly enough, the contemporaneous reaction of money to the domestic price level is estimated to be slightly negative. The weak contemporaneous connection between money and income may be a consequence of the credit rationing prevalent during most of the sample period.[9] The empirical findings are congruent with the theoretical model spelled out in section 6.1 (see expressions (6.9) and Figure 6.1). However, since we have not computed standard errors for the estimates of the elements of Γ_0, the importance of sampling variability remains unknown.

Empirical evidence on structural shocks as identified by our model during the full-sample period is presented in Table 6.5. The estimated large structural shocks are plotted against time in Figure 6.2. All shocks are zero mean entities in the first moment sense, but their variances differ considerably. Domestic goods market and aggregate supply shocks are by far the largest whereas money shocks qualify for second place. In a system augmented with foreign output, domestic goods market and aggregate supply shocks still come in first, but their variance is down by one fifth. Domestic price shocks are larger than foreign price shocks by a factor of three. Foreign interest rate shocks display very little variability; domestic interest rate/monetary policy shocks were found to be six times as volatile as their foreign counterparts when the model was augmented with i.

[9]The estimates of contemporaneous relationships are remarkably robust to alternative operationalizations and augmentations. The only substantial exception is the relation between output and foreign prices. The foreign price variable emerges with a sizeable positive coefficient if the model is augmented with foreign output or if stochastic seasonal filtering is used, and with a small and positive coefficient if six instead of three lags are used. Structural estimates based on a recursive design matrix turned out to be implausibly small, on the whole, in modulus. The importance of these findings will be examined as the analysis proceeds. Estimates of the variances of the structural shocks remain virtually unchanged under alterations in the model.

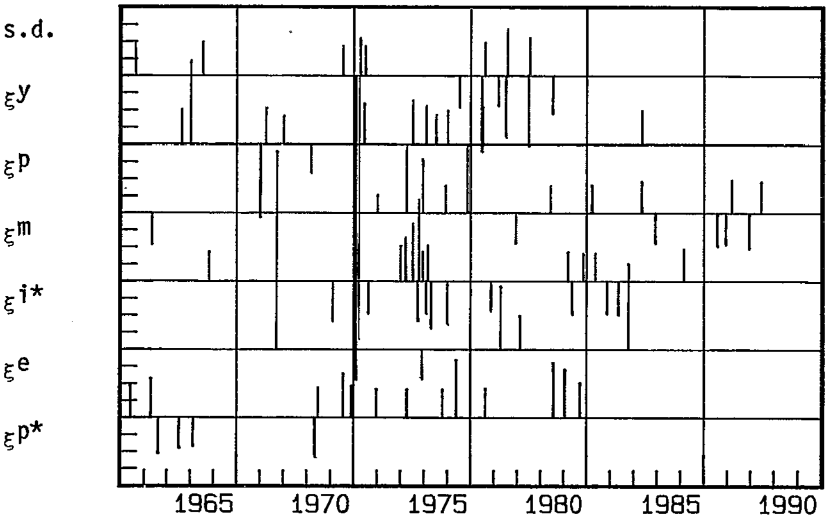
TABLE 6.5

Analysis of structural innovations in the vector $\underset{\sim}{y}_t = (y_t\ p_t\ m_t\ i^*_t\ e_t\ p^*_t)'$ as identified using high frequency restrictions during the period 1960M1 - 1988M8

Innovation Statistic		ξ^y	ξ^p	ξ^m	ξ^{i*}	ξ^e	ξ^{p*}
Mean x 10^3		0.000	0.000	0.000	0.000	0.000	0.000
Variance x 10^3		1.82	0.022	0.873	0.018	0.128	0.007
Skewness		-0.723***	0.570***	0.726***	0.624***	9.85***	0.791***
Kurtosis		5.61***	6.60***	3.24***	5.25***	1.33***	1.73***
Autocorrelation	1	0.021	-0.011	0.002	-0.007	0.007	-0.035
	2	-0.009	-0.040	-0.016	0.000	-0.010	-0.061
	3	-0.067	-0.024	-0.049	-0.032	-0.027	-0.114**
	4	-0.155***	-0.009	-0.070	0.122**	0.034	0.012
	5	-0.094*	0.081	-0.056	0.098*	-0.020	-0.040
	6	-0.045	0.064	0.026	0.115**	0.015	0.108**
	7	-0.099*	0.012	-0.132**	-0.033	-0.007	0.069
	8	-0.105*	-0.028	0.061	-0.004	-0.011	0.212***
	9	-0.004	0.038	0.024	0.029	-0.043	-0.093*
	10	0.013	-0.061	-0.007	-0.065	0.032	0.085
	11	-0.040	0.089	0.032	0.158***	0.024	0.043
	12	0.570***	0.071	0.209***	-0.027	-0.085	0.019
Significant positive shock		1960M7 1961M7 1964M7 1970M7 1971M4 1971M5 1976M8 1977M8 1978M8	1963M9 1964M1 1967M4 1968M1 1971M6 1973M7 1974M2 1974M7 1975M1 1976M7 1983M5	1973M4 1973M12 1974M12 1975M12 1979M6 1981M3 1983M5 1987M3 1988M6	1964M11 1973M1 1973M2 1973M7 1973M11 1973M12 1974M1 1974M3 1980M3 1980M11 1981M5 1985M2	1967M10 1977M4 1978M2 1982M10	1961M6 1962M4 1969M6 1970M7 1970M12 1971M12 1973M4 1974M10 1975M5 1976M8 1979M7 1980M1 1980M9
Significant negative shock		1971M2 1971M3 1975M7 1976M7 1977M3 1977M7 1978M7 1979M7	1967M1 1969M3 1972M1	1960M12 1962M6 1977M12 1983M12 1986M8 1986M12 1987M12	1970M2 1971M8 1973M10 1974M2 1974M4 1975M1 1976M11 1980M5 1981M11 1982M5	1973M11	1962M8 1963M7 1964M2 1969M5

Structural impulses are estimated from a first-differenced VAR model comprising the elements of $\underset{\sim}{y}_t$, a constant, 11 seasonal dummy variables and a strike dummy taking the value 1 in 1971M2 - 1971M3 and 0 otherwise. The model is estimated by equation-by-equation OLS using restrictions reported in Table 6.4. Significant values at the 10, 5 and 1 per cent level of significance are indicated by *, ** and ***, respectively. Listed shocks exceed the two standard error limit (the dummy variable strike dates are listed as significant negative shocks).

FIGURE 6.2 Plot of large structural innovations in the vector $\underset{\sim}{y}_t = (y_t\ p_t\ m_t\ i^*_t\ e_t\ p^*_t)'$ as identified using high frequency restrictions during the period 1960M1 - 1988M8



With reference to the optimal exchange rate regime literature (see Boyer (1978) and the surveys by Edison (1987) and Mathieson (1988)), the stylized facts about the structural shocks may be utilized in a discussion of the Finnish exchange rate arrangement. According to this literature, the degree of exchange rate flexibility is a function of the relative occurrence of real and nominal shocks. The more prevalent real shocks the more desirable a fixed exchange rate, and the more prevalent nominal shocks the more desirable a flexible exchange rate. Since neither real nor nominal shocks clearly seem to have dominated in terms of average frequency or size, neither a completely fixed nor a freely floating exchange rate would seem to be an optimal choice against this background.

All structural shocks have nonnormal distributions. The shocks ξ^y are skewed to the left and other shocks are skewed to the right. Moreover, the shocks ξ^{p^*} are platycurtic, the other shocks being leptocurtic. These findings are suggestive of the view that a portion of the structural shocks are sizeable and systematically of one sign. This conclusion has also been reached by Blanchard & Watson (1986) in their study of the U.S. economy. The structural shocks furthermore appear to contain temporal dependencies of a complicated nature; the first two autocorrelations of the shocks are of negligible size but other autocorrelations sometimes modestly exceed conventional standard error limits. Some autocorrelation at the seasonal lag is left in the equations for output and money.

Turning to individual estimated structural shocks we find that only a very small portion of the shocks can be classified as being large. Occasional large shocks are documented, however. When it comes to the sign of the large shocks one may note that price and exchange rate shocks have been predominantly positive. Given the difference in variance between domestic and foreign price shocks, this implies that historically domestic price shocks have constituted a major source of inflationary pressure on the Finnish economy.

Rather than to try to elaborate on the connection of each of the 91 estimated large structural shocks with historically exceptional events, we will draw attention to the temporal allocation of these shocks. During the sample period, one can distinguish some predominantly calm and some predominantly turbulent subperiods. With respect to the real variable, the 1960s and the 1980s emerge as relatively calm, whereas the years 1970 - 1971 and 1975 - 1979 were turbulent times. Domestic price shocks were largest in 1973 - 1974, and the major foreign inflationary impulses occurred mainly during the 1970s. Foreign interest rate shocks cluster around the years of the oil shocks. On the whole, the 1960s and the 1980s emerge as considerably calmer decades than the 1970s. The period 1973M1 - 1975M5 stands out as being particularly turbulent. These findings support the view that observed business cycles are not alike (cf. Blanchard & Watson (1986)).

One should also bear in mind that clusters might be symptomatic of changes in policy rules or other structural relationships pertinent to the workings of the economy (see footnote 6 of chapter 3). Of course, whether large economic shocks caused policy rules or other structural relationships to change or whether it was the other way around cannot be inferred from the mere occurrence of clusters. Our choice of Bretton Woods and post-Bretton Woods subperiods roughly corresponds to the implied periods of structural shifts.

6.2.3 Long-run Impact of Structural Shocks

We document impacts of structural shocks through cumulative impulse responses and long-run decompositions of variance. One standard error shocks are employed, and cumulative impulse responses are in terms of fractions of standard deviations. Decompositions of variance are displayed as percentage points. The long run is approximated as a 60-month horizon, since practically all computed statistics converged by then (exceptions are commented upon). Unfortunately, evaluation of the accuracy of cumulative impulse responses and decompositions of variance estimated from our structural VAR model is a nontrivial task, to say the least, and we are not able to report confidence bounds for these statistics. Evidence presented by Geweke (1984), Runkle (1987), Diebold & Rudebush (1989) and Lütkepohl & Reimers (1989a,b) unanimously suggests that the uncertainty associated with statistics like ours may be considerable.

When dealing specifically with the long-run properties of an economic system, special attention should be paid to what possible long-run constraints one is, or should be, imposing - explicitly or implicitly - on the estimated model (see section 4.2 of chapter 4). We will deal with the issue of long-run equilibrium relationships by reporting not only estimates based on the basic differenced specification but also estimates based on the model in levels and the model in differences augmented with error correction (level) terms. In the preliminary data analysis of chapter 4 the error correction terms were found to enter the model with essentially zero weights, however. It remains an empirical matter whether issues of cointegration are important or not for the conclusions. Empirical evidence on the long-run effects of structural shocks in the model for the full sample is summarized in Table 6.6.

TABLE 6.6

Long-run characteristics of VAR models for the vector $\underset{\sim}{y}_t = (y_t\ p_t\ m_t\ i_t^*\ e_t\ p_t^*)'$ using data from the period 1960M1 - 1988M8

Statistic	Variable	Innovation in y	p	m	i*	e	p*
		Model estimated in first differences					
Cumulative	y	0.483	-0.158	-0.011	0.014	0.009	-0.242
impulse	p	-0.084	2.09	0.014	-0.136	0.023	1.47
response to a	m	-0.041	0.424	0.664	-0.087	-0.049	0.746
one standard	i*	0.085	0.677	0.006	1.37	-0.217	1.06
error	e	-0.070	0.729	0.026	0.002	1.22	0.686
innovation	p*	-0.064	1.90	0.162	0.044	-0.177	3.69
Percentage of	y	89.0	2.94	1.10	2.67	0.328	3.94
the expected	p	0.481	76.5	0.781	3.59	3.52	15.1
long-run	m	1.79	0.947	92.6	2.91	0.940	0.846
squared	i*	0.119	0.679	1.19	97.0	0.648	0.382
prediction	e	0.504	0.611	0.108	1.05	97.0	0.680
error	p*	0.389	7.46	2.62	4.07	1.58	83.9
		Model estimated in levels					
Long-run	y	0.152	1.23	0.394	0.163	0.167	1.90
impulse	p	0.123	0.668	0.186	0.154	0.235	0.902
response to a	m	-0.033	1.09	0.130	-0.093	-0.016	1.42
one standard	i*	-0.030	0.589	0.004	0.028	0.007	1.12
error	e	0.189	1.92	0.556	0.259	0.183	3.25
innovation	p*	-0.045	0.023	-0.027	-0.080	-0.062	0.188
Percentage of	y	52.7	4.63	0.771	0.993	39.6	1.22
the expected	p	9.40	15.5	30.9	16.2	26.3	1.67
long-run	m	21.2	2.02	36.2	0.345	39.8	0.468
squared	i*	2.41	1.71	0.739	73.1	20.5	1.53
prediction	e	13.0	10.0	4.91	1.75	69.3	1.02
error	p*	9.44	1.96	22.6	21.0	37.3	7.77
		Model estimated in error correction form					
Cumulative	y	0.524	-0.083	-0.037	0.018	-0.087	-0.074
impulse	p	-0.028	1.39	-0.017	-0.030	0.110	0.312
response to a	m	-0.057	-0.025	0.640	-0.128	0.104	-0.007
one standard	i*	0.067	0.178	-0.063	1.42	-0.067	0.104
error	e	-0.116	0.677	-0.019	-0.048	1.33	0.338
innovation	p*	0.007	0.423	0.012	0.264	0.090	1.31
Percentage of	y	88.8	2.89	1.02	2.47	0.975	3.83
the expected	p	0.562	86.0	0.421	2.81	7.72	2.47
long-run	m	2.18	0.900	92.4	2.77	0.975	0.720
squared	i*	0.092	0.422	1.29	96.2	0.712	1.32
prediction	e	1.09	0.847	0.538	0.353	96.8	0.404
error	p*	0.438	1.77	0.562	0.719	1.54	95.0

In addition to the variables in $\underset{\sim}{y}_t$, the models comprise a constant, 11 seasonal dummy variables and a strike dummy taking the value 1 in 1971M2 - 1971M3 and 0 otherwise. The models are estimated by equation-by-equation OLS using the restrictions reported in Table 6.4. In all calculations, the long-run refers to a horizon of 60 months. Numbers across prediction error rows may not sum to 100.000 due to rounding errors.

The following considerations could be borne in mind when interpreting own-shock responses. A cumulative (or long-run) impulse response of zero implies that the effect of a shock is transitory. If the response is one, the effect of the shock is permanent and reflects the shock in a one-to-one fashion. If the response exceeds one, the response implies eventual overshooting. Since a considerable degree of uncertainty may be associated with the estimates, we will only refer to responses exceeding 0.400 in modulus as significant deviations from zero. Likewise, we will only consider decompositions of variance in excess of 10.0 per cent as important. Even more stringent rules of thumb should presumably be used when shorter samples are analyzed. It should be emphasized that the choice of the aforementioned values is entirely arbitrary, but one emerges from the study by Runkle (1987) with the feeling that these values may be too low if anything.

First consider the estimates based on the model in first differences. The cumulative long-run response of the real variable to shocks in nominal variables support the notion of neutrality, or, to put it slightly differently, no convincing signs of nonneutrality are found. In particular, the response of output to money, interest rate and exchange rate shocks is convincingly close to zero (money is superneutral). The impact of price shocks is, if anything, slightly negative, but it seems highly unlikely that the responses differ significantly from zero. With regard to the responses of nominal variables, price shocks - domestic and foreign - generate substantial responses in all (nominal) variables. The ultimate impact on the domestic price level of an unforeseen movement in the foreign price level is considerable. Domestic and foreign price developments are also to a considerable extent reflected in the exchange rate in the long run. Money reacts to price shocks with steady state gains between zero and one.

Half of an unforeseen movement in output is estimated to persist into the infinite future. In other words, when output grows unexpectedly, forecasts of (distant) future output should be adjusted upwards by half the amount of the shock. To put it yet another way, output contains both stochastic and deterministic trend components, and while mean reversion occurs, it is not complete. Thus interpreting a deviation of output from a deterministic trend as a boom or a recession will overestimate the deviation from the true trend which is, in part, stochastic. This conclusion has also been reached with respect to the output of other western economies (see Stock & Watson (1988) and the references cited therein). As the output shock ξ^y has the interpretation of a goods market shock or a shock to technology or labor supply, it may be argued that the stochastic part of the trend in output arises because of supply side, real shocks. The other variables are also seen to be dominated by random walk components, and in particular the price levels emerge as unstable.

Next, turn to estimates based on the model in level form. The estimate of the matrix of contemporaneous relationships - on which estimates of structural shocks rest - as identified from data in levels was presented in Table 6.4. Estimates differ very little whether first differences or levels are used. Only two sign reversals are encountered, but the moduli of these estimates are very small. When the model is estimated in levels, the contemporaneous relationship between output and domestic prices is positive (though essentially zero), and the contemporaneous relationship between output and the exchange rate is slightly positive. Both changes are compatible with commonsense economic reasoning, especially as the structural equation for output contains both demand and supply elements.

The estimated (nonaccumulated) long-run impulse responses as identified from level regressions requires closer scrutiny and discussion. Taken at face value, the results differ markedly from those obtained using the model in first differences. If in the true structure the variables are integrated (but not cointegrated), the cumulative impulse responses of the system in first differences should be roughly the same as nonaccumulated impulse responses of the corresponding system in levels (abstracting from estimation differences).

However, in our case only a few of the impulse responses obtained from level regressions converged. While the responses were similar to the corresponding results from the first-differenced system for the first 10 - 20 step ahead horizons, the responses displayed very long swings and some started going off towards infinity after a three year horizon. We take these nonsensical results to indicate that convergence may in practice not be as fast as the theoretical results of Sims et al. (1987) indicate, thus leaving spurious fits in the model. Consequently, our result documents a case against the recommendation by Sims et al. (op.cit.), Diebold & Nerlove (1988) and Doan (1989) to estimate VARs in levels when unit roots and possible cointegration relationships are present.

The preceding results have attested to the importance of explicitly addressing issues of unit roots in our analyses. We next turn to the question as to whether explicitly taking into account common unit roots has sizeable implications for the conclusions of our study. Whereas the preliminary data analysis in chapter 4 pointed toward the existence of cointegration among our model variables, maximum likelihood estimates of the corresponding long-run coefficient matrix generated an essentially zero matrix. Nevertheless, one of the cointegrating vectors may enter the equation for output with a nonzero weight (see Table 4.5 of chapter 4). Hence, we choose to impose this cointegrating constraint on the model as a matter of sensitivity analysis. This is done by adding the deviations from the maximum likelihood estimate of the corresponding cointegration relationship as a once lagged explanatory variable to the equations of the model in first differences.[10]

The estimate of the matrix of contemporaneous relationships when the cointegration constraint is imposed was displayed in Table 6.4.

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[10] It might be tempting to claim that the estimate of the long-run coefficient matrix (see Table 4.5 of chapter 4) implies long-run neutrality from nominal variables to the real variable. As pointed out by Lütkepohl & Reimers (1989a,b), such an interpretation would ignore the fact that we are considering a system of equations in which the other variables cannot be regarded as fixed. Hence, the reactions of the other model variables have to be taken into account, too. Thus, a direct interpretation of the results from the preliminary data analysis is likely to be difficult when it comes to issues of neutrality.

Results differ only a little with respect to the baseline case. Two differences - importantly enough both in the equation for the real variable - can be noticed, however. Both domestic prices and the exchange rate now affect contemporaneous real output positively, albeit with small coefficients. To see what impact these changes have on the cumulative impulse responses as compared to the estimates solely based on first differences, we turn to the evidence presented in Table 6.6.

Estimates of the cumulative impulse responses when a cointegration constraint is imposed do differ somewhat from the estimates from the model omitting the cointegration constraint. More precisely, the importance of the price level variables is downplayed. However, it is not obvious that the imposition of the cointegration constraint adds positively overall to the model. We find the result that the role of domestic prices for foreign price level variability is downplayed in accordance with prior beliefs. On the other hand, the findings that domestic prices are not eventually reflected in the nominal money stock, or that movements in the foreign price level account for hardly any of the low frequency variability of the domestic price level, are hard to reconcile with conventional economic thought. Hence, we will continue to work primarily with the first-differenced specification, but sensitivity analyses with respect to common unit roots will be made.

Estimated long-run decompositions of variance are also displayed in Table 6.6. The empirical evidence is fairly easily summarized, as results for the augmented model in differences and this model incorporating the cointegration constraint yield virtually identical results. In the long run, unforeseen movements in the nominal variables account for practically no portion of the unforeseen movements of the real variable. Goods market and aggregate supply shocks account for some nine tenths of the long-run variability of real output. With the exception of the foreign price level, no variable accounts for much of the unforeseen long-run movements of other variables. Shocks to the foreign price level account for a nonnegligible part of unexpected movements in the domestic price level in the long run, however.

As far as the long-run decompositions of variance estimated from level data are concerned, we regard these as dubious because of the reasons spelled out above. For example, the result that unforeseen movements in the Finnish exchange rate account for more than one third of the unforeseen long-run movements of foreign prices is spurious. Most of the other results in this case make equally little sense. This corroborates our claim that estimating VARs in levels when the model includes unit roots, some of which may be common to a set of model variables, may be inappropriate when working with finite samples.

Before moving on to analyses based on subsamples, a few words about the robustness of the above results to alternative operationalizations and augmentations of the model. The robustness was evaluated by calculating cumulative impulse responses and long-run decompositions of variance based on estimates of alternative variants of the model in first differences. Instead of reporting the ten additional sets of results, we will only highlight the main diverging results.

Adding foreign output to the model reveals that nearly one third of the long-run forecast error variance of domestic output has been determined by unexpected movements in foreign output. Consequently, domestic output emerges as less exogenous, with goods market and supply shocks accounting for roughly three fifths of the long-run variability of domestic output. Using six lags, reversion toward a deterministic output trend is stronger, with only some three tenths of the output shock persisting into the infinite future.[11] However, the use of six lags is not unproblematic, in part because an unforeseen rise in domestic prices has no long-run impact on the money stock. The same goes for the use of seasonal differences instead of

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[11] Taken together, our results suggest that the size of the random walk component of Finnish output lies in the range 0.3 - 0.6 with the best estimate being 0.5. Our results are in line with the multivariate estimates of 0.3 - 0.4 for the U.S. presented by Cochrane & Sbordone (1988). Univariate estimates, which may utilize less information than multivariate estimates, have yielded estimates in the range 0.4 - 1.9 (see Campbell & Mankiw (1988), Stock & Watson (1988a), and Hamilton (1989) and the references cited therein).

seasonal dummies, as it annihilates all links from foreign prices to other variables while producing causality from domestic to foreign prices.

Finally, it may be interesting to compare results using the structural model with the results that might have been entertained had an "astructural" (i.e. recursive) model been employed. While mechanical use of the Wold causal chain makes no sense in terms of implied economic relationships, this exercise serves to illustrate the sensitivity of our results to the choice of contemporaneous relationships. As it happens, the results turn out to be quantitatively somewhat different from, but qualitatively similar to, the baseline case. Judging by the cumulative impulse responses, the only major change is that the importance of the foreign interest rate for domestic output is overstated by a factor of 16. In terms of long-run decompositions of variance, the relationship between the domestic and the foreign price level is somewhat distorted. Domestic variables further appear somewhat more exogenous, and foreign variables somewhat less exogenous, when a lower triangular design is used. Thus our conjecture from section 6.2.1 that our results will not be overly sensitive to the specification of contemporaneous relationships is confirmed.

6.2.4 Analysis of Subperiods

The likelihood ratio test statistic of Sims (1980a) supports the use of three lags in all subsample analyses. Results for matrices of contemporaneous relationships are presented in Table 6.7. During the 1960s and the early 1970s, the contemporaneous relation between output and the domestic price level was negative. This could be interpreted to suggest that output has been determined primarily by aggregate demand (as opposed to aggregate supply) in the short run during this period. An increase in the price of oil has had a negative contemporaneous effect on output. The contemporaneous effect of the price of oil on the domestic price level is positive but very small. Credit advanced to the public has had a slightly positive contemporaneous impact on the domestic price level.

Turning to the contemporaneous structure as estimated from the period 1973M1 - 1988M8, we find that the relationship between output and prices is now positive. This could be taken to indicate that output has been dominated by aggregate supply (as opposed to aggregate demand) in the short run since the first oil shock. Output shocks have been both larger and more frequent than at other times.[12] Results from the last subperiod differ somewhat from previous results.

[12]The findings that the distribution of structural shocks is fat-tailed and that large shocks have from time to time tended to form clusters might suggest that the structural shocks follow autoregressive conditionally heteroskedastic (ARCH) processes. However, we think that this would be a false conclusion for two reasons. If the structural shocks follow ARCH-processes, these stochastic events follow processes governed by deterministic difference equations. Perhaps more importantly, the ARCH model implies that volatile periods should occur fairly regularly, but our evidence indicates that this has not been the case.

TABLE 6.7

Contemporaneous structural relationships in VAR models using data from subperiods

Model variables	Matrix of contemporaneous relationships

1960M1 - 1972M12

$$\underset{\sim}{y}_t = \begin{bmatrix} y_t \\ p_t \\ c_t \\ i_t^* \\ e_t \\ p_t^{02} \end{bmatrix} \qquad \Gamma_0 = \begin{bmatrix} 0 & -0.557 & 0 & 2.12 & 0.166 & -0.075 \\ 0 & 0 & 0 & -0.124 & 0 & 0.005 \\ -0.010 & 0.092 & 0 & 0.076 & 0 & -0.037 \\ 0 & 0 & 0 & 0 & 0 & -0.010 \\ 0 & & & \cdots & & 0 \\ 0 & & & \cdots & & 0 \end{bmatrix}$$

1973M1 - 1988M8

$$\underset{\sim}{y}_t = \begin{bmatrix} y_t \\ p_t \\ m_t \\ i_t^* \\ e_t \\ p_t^* \end{bmatrix} \qquad \Gamma_0 = \begin{bmatrix} 0 & 0.717 & 0 & 1.35 & 0.096 & -1.17 \\ 0 & 0 & 0 & -0.070 & 0 & 0.091 \\ 0.081 & -0.323 & 0 & -1.00 & 0 & 1.61 \\ 0 & 0 & 0 & 0 & 0 & 0.164 \\ 0 & & & \cdots & & 0 \\ 0 & & & \cdots & & 0 \end{bmatrix}$$

1980M1 - 1988M8

$$\underset{\sim}{y}_t = \begin{bmatrix} y_t \\ p_t \\ m_t \\ i_t^* \\ e_t \end{bmatrix} \qquad \Gamma_0 = \begin{bmatrix} 0 & 0.781 & 0 & 1.30 & -0.114 \\ 0 & 0 & 0 & -0.004 & 0 \\ 0.096 & -0.608 & 0 & 0.196 & 0 \\ 0 & & \cdots & & 0 \\ 0 & & \cdots & & 0 \end{bmatrix}$$

In addition to the components of $\underset{\sim}{y}_t$, the models include a constant, 11 seasonal dummy variables and, where appropriate, a strike dummy taking the value 1 in 1971M2 - 1971M3 and 0 otherwise.

The empirical evidence on cumulative impulse responses and long-run decompositions of variance for subperiods is summarized in Table 6.8. Turning first to the evidence for the period 1960M1 - 1972M12, the following features may be noted. The amount of credit advanced to the public has had no important long-run effect on the domestic price level. Interpreting the results for credit in terms of a reaction function for monetary policy shows that credit has been increased in the face of unexpected output growth and decreased in the face of unexpected bursts of inflation.

The results for the period 1973M1 - 1988M8 do not differ radically from the full-sample results. Moreover, augmenting the model with the measure of the fiscal stance left the basic estimates practically unchanged. Some evidence of neutrality of fiscal policy is thus documented. Augmentation with the genuine oil price variable suggests that average oil price shocks have had virtually no long-run impact on output and only a small effect on domestic prices. While our VAR model suggests that average oil shocks have had negligible long-run effects, the Bank of Finland quarterly model suggests that a permanent, albeit small, effect exists (Lehtinen (1988)). As this conventional model explicitly restricts the investment in new (energy-saving) technology induced by an oil shock to zero, while our model does not impose this restriction, it is possible that a nonnegligible offsetting effect is missing from the conventional model. An oil shock may induce a change in technology - a technology shock - which is not captured by the conventional model but taken into account by the VAR model.

Estimates based on data from the 1980s only differ more from the results for the full sample than other subsample results. Some anomalous results caution against taking the results drawn from the last subsample too literally, however. A possible explanation for the partly whimsical results is that estimates are based on very few observations. Thus erratic movements will bear on the results to a considerable extent.

TABLE 6.8

Long-run characteristics of VAR models using data from subperiods

Statistic	Variable	Innovation in					
		y	p	m/c	i*	e	p^*/p^{02}
				1960M1 - 1972M12			
Cumulative	y	0.541	0.224	-0.111	0.123	0.159	0.086
impulse	p	-0.062	1.32	0.012	0.041	0.108	0.021
response to a	c	0.423	-0.632	2.21	-0.323	0.063	-0.062
one standard	i*	0.242	-0.205	0.409	1.09	-0.123	0.154
error	e	0.001	0.884	-0.437	0.382	1.28	0.215
innovation	p^{02}	0.075	0.290	-0.341	0.138	0.225	0.700
Percentage of	y	77.0	2.14	4.98	5.35	2.03	8.52
the expected	p	2.24	72.3	8.36	1.01	14.2	1.98
long-run	c	0.948	1.15	92.0	1.90	2.12	1.84
squared	i*	0.727	0.392	4.50	88.4	4.84	1.13
prediction	e	1.25	0.500	0.356	0.527	95.5	1.83
error	p^{02}	0.957	0.446	3.43	1.39	6.38	87.4
				1973M1 - 1988M8			
Cumulative	y	0.514	-0.109	0.066	0.067	-0.074	-0.129
impulse	p	0.037	2.79	0.063	-0.277	-0.146	2.72
response to a	m	-0.023	0.587	0.611	-0.155	0.129	0.788
one standard	i*	0.093	0.977	-0.132	1.46	-0.352	1.59
error	e	-0.169	-0.010	-0.105	-0.114	1.37	0.049
innovation	p*	-0.104	2.93	0.147	0.124	-0.254	5.08
Percentage of	y	88.9	2.40	1.29	2.57	3.85	0.981
the expected	p	0.720	65.9	1.32	7.71	1.18	23.2
long-run	m	3.11	0.526	89.3	3.43	1.84	1.78
squared	i*	1.87	2.05	3.10	88.3	1.00	3.72
prediction	e	1.04	2.30	1.16	2.66	91.3	1.49
error	p*	0.143	16.2	2.37	4.83	0.663	75.8
				1980M1 - 1988M8			
Cumulative	y	0.419	0.870	-0.015	0.225	-0.040	
impulse	p	-0.076	2.71	-0.007	-0.039	-0.402	
response to a	m	-0.029	-0.465	0.410	0.049	-0.032	
one standard	i*	0.004	-0.079	-0.143	1.43	-0.218	
error	e	-0.066	1.16	0.140	-0.215	1.45	
innovation							
Percentage of	y	82.2	2.53	4.81	9.14	1.30	
the expected	p	8.40	70.0	1.56	7.51	12.5	
long-run	m	2.37	1.05	88.6	6.28	1.66	
squared	i*	5.30	0.317	3.97	89.1	1.29	
prediction	e	6.31	9.97	0.999	5.04	77.7	
error							

In addition to the variables in $\underset{\sim}{y}_t$, the models comprise a constant, 11 seasonal dummy variables and, where appropriate, a strike dummy taking the value 1 in 1971M2 - 1971M3 and 0 otherwise. The models are estimated by equation-by-equation OLS using the restrictions reported in Table 6.7. In all calculations, the long run refers to a horizon of 60 months. Numbers across prediction error rows may not sum to 100.000 due to rounding errors.

It should perhaps be emphasized that we are documenting the long-run effects of average, i.e. "typical", unforeseen oil price movements. The analysis of such oil price movements revealed no noteworthy long-run impacts, but it is quite possible, and even likely, that the "atypical", i.e. unusually large, oil shocks of 1973, 1979 and 1986 have had nonnegligible long-run impacts. In fact, our findings of structural breaks and differences in results across subsamples constitute indirect evidence of this. In other words, our subsample results may be symptomatic of profound structural changes induced by the large oil price shocks. Similarly, the exceptionally large unexpected oil price movements may have induced other, additional shocks (recall that we found a cluster of shocks in 1973 - 1975), but our framework allows only indirect study of such a phenomenon (see also footnote 1 of chapter 3).

International evidence obtained with VAR models of the effects of oil shocks add yet another dimension to our results. Pronounced differences in the responses of western economies to oil price shocks have been documented (Burbidge & Harrison (1984) and Mork et al. (1989), see also Virén (1986)). Hence, it does not appear to be self-evident what the actual outcome of an oil shock should be in some country in the long run. As in our study, differences in the response to the unusually large 1973, 1979 and 1986 oil shocks have also been documented (Hamilton (1983), Burbidge & Harrison (op.cit.), Roberds & Todd (1987), Mork (1987) and Mork et al. (op.cit.)). Thus Roberds & Todd (op.cit.), using the most elaborate VAR model constructed to date, generate only a modest and short-lived effect on U.S. output of the 1986 oil price shock.

Inclusion of foreign output yields some additional interesting insights although contributing little to the basic results. Since this augmentation - while not representing the result of a strict statistical model selection procedure - yields an interesting economic model, which, moreover, illustrates the degree of sensitivity typically encountered in these analyses, we have included the estimates as Appendix 6. In this appendix, estimates of contemporaneous relationships during subperiods are presented in Table A6.1, and cumulative impulse responses and long-run decompositions of variance in Table A6.2. These estimates can be compared to the results presented in the previous chapter, and specifically to the subsample estimates presented in Table 5.6. One may note that the inclusion of nominal variables accounts for five to 12 per cent of the long-run variability in domestic output, depending on which subperiod is considered.

During the period 1960M1 - 1972M12, a positive but very modest contemporaneous relationship between domestic and foreign output is found. In contrast to this, foreign output contemporaneously affected domestic output with a weight of 0.9 during the period 1973M1 - 1988M8. When this weight is estimated from the 1980s only, a value of 0.3 is obtained. We interpret this as an indication that while the openness of the Finnish economy has been increasing somewhat over time, the turbulent period of 1973 - 1975 caused big short-run movements in both foreign and domestic output which were similar enough to dominate the estimate of the weight.

Turning to cumulative impulse responses and long-run decompositions of variance that take foreign output into account, we note that results for the period 1960M1 - 1972M12 remain practically unchanged. During the period 1973M1 - 1988M8, the only change in the basic results is that domestic output emerges as more endogenous because foreign output now accounts for roughly one fifth of the long-run variability of domestic output. Dubious nonneutrality results with respect to the long-run impact of domestic variables on foreign output may attest to problems with the inclusion of foreign output in the model, however.

Estimates based on the period 1980M1 - 1988M8 very much resemble the estimates drawn from the post-Bretton Woods period, but we find that the importance of foreign output for domestic output is clearly downplayed. This supports the view that short-run movements at the time of the first oil shock tend to bias estimates of the degree of domestic output endogeneity upwards from what it really has been during recent years.

6.3 Conclusions

The aim of this chapter has been to provide empirical evidence on the long-run effects of shocks hitting the Finnish economy. We conclude that the structural shocks are zero mean entities in the first moment sense but that they are nonnormally distributed. Although seldom, large shocks of predominantly one sign have occurred, and these shocks have clustered around fairly short periods. The period 1973 - 1975 was extra ordinarily turbulent. Oil price shocks, strikes and exchange rate shocks have been the most sizeable. Domestic price shocks have been larger than foreign price shocks and the same holds for interest rate shocks.

With regard to the long-run impact of nominal shocks on Finnish real output, our results consistently point to the notion of neutrality. Stated slightly differently, no convincing signs of nonneutrality are encountered. This finding is robust, and it includes the effects of monetary policy as well as some features of fiscal policy. Evidence of the crucial importance of foreign price developments for domestic price developments is documented. These findings include a sizeable contemporaneous interaction and a roughly one-to-one long-run relation. The long-run impact of foreign output on domestic output is generally not substantial, though the developments immediately after the first oil shock constitute an exception here.

Half of an unforeseen movement in output is estimated to persist into the infinite future. In other words, when output grows unexpectedly, forecasts of output should be adjusted upwards by half the amount of the shock. This is an outcome of the fact that output contains both stochastic and deterministic trend components, implying that while mean reversion occurs, it is not complete. Thus interpreting a deviation of output from a deterministic trend as a boom or a recession will overestimate the deviation from the true trend which is, in part, stochastic. We argue that the stochastic part of the trend in output arises because of real supply shocks. Shocks to the price of oil have contributed to long-run aggregate movements in the Finnish economy to a lesser extent than in other economies. We argue that this is because of the bilateral trade arrangements with the Soviet Union.

7 A SECULAR FLUCTUATION MODEL

The aim of this chapter is to provide empirical evidence on the short-run effects of shocks hitting the Finnish economy. We use the theoretical model of the previous chapter as a frame of reference, and achieve identification of the shocks by imposing the low frequency restrictions implied by the empirical evidence of the previous chapter. By restricting low (and seasonal) frequency movements, we achieve identification of shocks and at the same time leave the data free to speak about high frequency movements.

7.1 An Empirical Model with Low Frequency Restrictions

In the present chapter the short run is defined as extending up to 24 months. However, one should note that the distinction between short- and long-run impacts of shocks is a matter of degree rather than of kind. This is because one must resort to arbitrary conventions in labelling what is the short run. Furthermore, recent empirical findings regarding the time series properties of macroeconomic variables - e.g. the evidence presented in the previous chapter - and theoretical work in the real business cycle vein suggest that shocks that move the economy in the short run may also affect the economy in the long run.[1]

It could also be emphasized that we are deriving our low frequency restrictions from empirical evidence, and not, like previous authors, from ad hoc or theoretical considerations.[2] What is gained by such a procedure is that the credibility of the identifying assumptions - which is typically subject to doubt - can be, and in

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[1]See the discussion and references in Shapiro & Watson (1988).

[2]King et al. (1987) build on a fully articulated real business cycle model. Blanchard & Quah (1989) implement an ad hoc restriction, which is also used as the main identifying restriction in a somewhat more articulated semi ad hoc model by Shapiro & Watson (1988). The model of Shapiro & Watson (op.cit.) is employed by Kugler (1989) in a study of the Swiss economy.

the present study has been, increased through demonstrations of the robustness of the empirical evidence from which the restrictions derive. One may also argue that basing the restrictions on empirical evidence is more in the markedly empirical spirit of the VAR approach than basing the restrictions on ad hoc or theoretical considerations.

In deriving the identifying low frequency restrictions, we will focus on the cumulative impulse responses presented in Table 6.6 for the model estimated from first differences of the data, and in the case of the relation between domestic and foreign prices also on the exogeneity results presented in Table 6.2.[3] Interpreting these estimation results in the same way as in the previous chapter yields the following matrix H(1) (see (3.19)) of long-run multipliers

$$(7.1) \qquad H(1) = \begin{bmatrix} \gamma_1 & 0 & 0 & 0 & 0 & 0 \\ 0 & \gamma_2 & 0 & 0 & 0 & \gamma_3 \\ 0 & \gamma_4 & \gamma_5 & 0 & 0 & \gamma_6 \\ 0 & \gamma_7 & 0 & \gamma_8 & 0 & \gamma_9 \\ 0 & \gamma_{10} & 0 & 0 & \gamma_{11} & \gamma_{12} \\ 0 & 0 & 0 & 0 & 0 & \gamma_{13} \end{bmatrix}$$

where all γs are expected to be positive. In matrix (7.1) nominal variables do not affect the real variable in the long run. The only difference between the exclusion restrictions implied by Table 6.6 and matrix (7.1) is that we have precluded long-run feedback from domestic prices to foreign prices. In doing so, we rely on the result presented in Table 6.2 according to which the foreign price level is exogenous to the domestic core variables including domestic

[3]The empirical evidence for the model estimated from data in levels and in error correction form is disregarded for the reasons spelled out in section 6.2.3 of chapter 6.

prices.[4] We subsequently allow the numerical values of the γs to be determined by the data; these parameters are not restricted to any particular values such as those of Table 6.6.

Upon rearrangement of the components of $\underset{\sim}{y}_t$, matrix H(1) can be made lower triangular. Assuming that the moving average coefficient matrix corresponding to the matrix of long-run multipliers is invertible, a VAR representation in first differences exists. In terms of autoregressive coefficients, the restrictions on moving average coefficients imply that the corresponding autoregressive coefficients each sum to zero. We follow Shapiro & Watson (1988) in imposing such constraints by differencing the corresponding variables once more, noting the changes in lag lengths. Imposing these constraints on our VAR representation yields the following near, or constrained, VAR representation

$$\text{(7.2i)} \qquad (1-L)a_{yy}(L)y_t = (1-L)^2a^*_{yp}(L)p_t + (1-L)^2a^*_{ym}(L)m_t + (1-L)^2a^*_{yi*}(L)i^*_t + (1-L)^2a^*_{ye}(L)e_t + (1-L)^2a^*_{yp^*}(L)p^*_t + \xi^y_t$$

$$\text{(7.2ii)} \qquad (1-L)a_{pp}(L)p_t = (1-L)^2a^*_{py}(L)y_t + (1-L)^2a^*_{pm}(L)m_t + (1-L)^2a^*_{pi*}(L)i^*_t + (1-L)^2a^*_{pe}(L)e_t + (1-L)a^{**}_{pp*}(L)p^*_t + a_{p\xi p*}\xi^{p*}_t + \xi^p_t$$

[4]Some feedback from domestic to foreign prices could be expected. This is because Sweden belongs to the aggregate of foreign countries, and if movements in Swedish prices can be reflected in Finnish prices, movements in Finnish prices can be reflected in Swedish prices. Nevertheless, the magnitude of the effect of domestic prices on foreign prices seems too high (Table 6.6). One explanation is that the more closely domestic prices follow foreign prices, the more difficult it becomes in practice to pin down what is a domestic price shock and what is a foreign price shock.

$$(7.2iii) \quad (1-L)a_{mm}(L)m_t = (1-L)^2a^*_{my}(L)y_t + (1-L)a^{**}_{mp}(L)p_t + (1-L)^2a^*_{mi*}(L)i^*_t + (1-L)^2a^*_{me}(L)e_t + (1-L)a^{**}_{mp*}(L)p^*_t + a_{m\xi p}\xi^p_t + a_{m\xi p*}\xi^{p*}_t + \xi^m_t$$

$$(7.2iv) \quad (1-L)a_{i*i*}(L)i^*_t = (1-L)^2a^*_{i*y}(L)y_t + (1-L)a^{**}_{i*p}(L)p_t + (1-L)^2a^*_{i*m}(L)m_t + (1-L)^2a^*_{i*e}(L)e_t + (1-L)a^{**}_{i*p*}(L)p^*_t + a_{i*\xi p}\xi^p_t + a_{i*\xi p}\xi^{p*}_t + \xi^{i*}_t$$

$$(7.2v) \quad (1-L)a_{ee}(L)e_t = (1-L)^2a^*_{ey}(L)y_t + (1-L)a^{**}_{ep}(L)p_t + (1-L)^2a^*_{em}(L)m_t + (1-L)^2a^*_{ei*}(L)i^*_t + (1-L)a^{**}_{ep*}(L)p^*_t + a_{e\xi y}\xi^p_t + a_{e\xi p*}\xi^{p*}_t + \xi^e_t$$

$$(7.2vi) \quad (1-L)a_{p*p*}(L)p^*_t = (1-L)^2a^*_{p*y}(L)y_t + (1-L)^2a^*_{p*p}(L)p_t + (1-L)^2a^*_{p*m}(L)m_t + (1-L)^2a^*_{p*i*}(L)i^*_t + (1-L)^2a^*_{p*e}(L)e_t + \xi^{p*}_t$$

where lags $a_{jj}(L)$ run from 0 to l, $a^*_{jj}(L)$ run from 0 to l-1, $a^{**}_{jj}(L)$ run from 1 to l, j = y,p,m,i*,e,p* and l is the lag length. In addition, the models comprise 11 seasonal dummy variables, a constant and a strike dummy in the equation for output to account for the strike in the metal and engineering industry in 1971M2 - 1971M3.

With an eye on estimation of the model (7.2i) - (7.2vi), we note that the occurrence of right-hand side current terms, in general, implies correlation between right-hand side variables and the structural shocks. Thus, one should consider the use of estimators such as 2SLS, 3SLS or system estimation methods. In the current case, two main arguments support the use of OLS, however. Anticipating the empirical results, we find that OLS estimation of our model comprising in excess of 300 parameters and involving the use of the Newton-Raphson method to account for the additional current shocks is very close to the limit of what our software can handle. Furthermore, since our model is inevitably misspecified in the sense that it excludes many variables of likely importance, OLS may be a superior technique of estimation compared to, inter alia, 2SLS (see Phillips (1983) and Amemiya (1985) and the references cited therein for discussions of the misspecification issue).

Hence, because of the computational constraint and the probably less pressing, but equally tangible, misspecification issue, we will proceed by estimating the model by OLS. The influence on impulse responses and decompositions of variance of the auxiliary shocks is taken into account by omitting the auxiliary disturbances in the estimation stage, and then transforming the model using the decomposition indicated by matrix (7.1) and employing the techniques described in section 3.2.1 of chapter 3 and section 6.2.2 of chapter 6 (see Shapiro & Watson (1988) for a discussion of alternative methods of accounting for the auxiliary disturbances when calculating impulse responses and decompositions of variance).

Before turning to estimation, we briefly comment on the long-run restrictions that will be imposed in subsample analyses. The restrictions are derived from Table 6.8 (and from Table A6.2). In our interpretation of the empirical results, the following matrices H(1) of long-run multipliers are obtained for the subperiods 1960M1 - 1972M12 ($\underset{\sim}{y} = (y\ p\ c\ i^*\ e\ p^{O2})'$), 1973M1 - 1988M8 ($\underset{\sim}{y} = (y\ p\ m\ i^*\ e\ p^*)'$) and 1980M1 - 1988M8 ($\underset{\sim}{y} = (y\ p\ m\ i^*\ e)'$), respectively (all vectors and their components refer to contemporaneous values)

$$(7.3) \qquad H(1) = \begin{bmatrix} \gamma_1 & 0 & 0 & 0 & 0 & 0 \\ 0 & \gamma_2 & 0 & 0 & 0 & 0 \\ \gamma_3 & \gamma_4 & \gamma_5 & 0 & 0 & 0 \\ 0 & 0 & \gamma_6 & \gamma_7 & 0 & 0 \\ 0 & \gamma_8 & \gamma_9 & 0 & \gamma_{10} & 0 \\ 0 & 0 & 0 & 0 & 0 & \gamma_{11} \end{bmatrix}$$

$$(7.4) \qquad H(1) = \begin{bmatrix} \gamma_1 & 0 & 0 & 0 & 0 & 0 \\ 0 & \gamma_2 & 0 & 0 & 0 & \gamma_3 \\ 0 & \gamma_4 & \gamma_5 & 0 & 0 & \gamma_6 \\ 0 & \gamma_7 & 0 & \gamma_8 & 0 & \gamma_9 \\ 0 & 0 & 0 & 0 & \gamma_{10} & 0 \\ 0 & 0 & 0 & 0 & 0 & \gamma_{11} \end{bmatrix}$$

$$(7.5) \qquad H(1) = \begin{bmatrix} \gamma_1 & \gamma_2 & 0 & 0 & 0 \\ 0 & \gamma_3 & 0 & 0 & 0 \\ 0 & \gamma_4 & \gamma_5 & 0 & 0 \\ 0 & 0 & 0 & \gamma_6 & 0 \\ 0 & \gamma_7 & 0 & 0 & \gamma_8 \end{bmatrix}$$

where all γs are expected to be positive, except for γ_4 and γ_9 in (7.3) and γ_4 in (7.5). The matrices (7.3) - (7.5) incorporate the finding that nominal variables do not affect the real variable in the long run, with one exception. The exception is γ_2 in (7.5), but the results concerning the last subperiod should not be taken too literally (see the discussion in section 6.2.4 of chapter 6). There is only one difference between the exclusion restrictions implied by Table 6.8 and those of matrices (7.3) - (7.5). This is that we have precluded long-run feedback from domestic prices to foreign prices in (7.4) (as in (7.1)). In doing so, we rely on the results presented in Table 6.2 according to which the foreign price level is exogenous to domestic prices. All restrictions are robust to the inclusion of foreign output in the models.

7.2 Empirical Evidence

7.2.1 Estimation Results

We begin by examining the question of lag length. Since the model equations contain different explanatory variables, it may be advantageous to allow for different lag lengths in different equations. To keep the tests of lag length within manageable limits, a common lag length within each equation is maintained. As it seems fair to say that testing for optimum lag length using finite data sets is an unresolved issue, we employ three criteria in guiding us to the equationwise value of l. These are the Akaike criterion (AIC), an Akaike criterion with the minimum penalty term suggested by Hannan & Quinn (1979) (HQIC), and the Schwarz criterion (SIC). We expect AIC to provide a lower, and SIC an upper, bound for l, while HQIC may suggest the direction within a possible allowable set of values for l.[5] Results for tests based on one to 18 lags are presented in Table 7.1.

[5]AIC overestimates the lag length asymptotically with positive probability, whereas SIC is consistent in the sense that it selects the correct lag length with probability one. HQIC imposes a heavier penalty for additional lags than AIC but imposes a less severe penalty than SIC.

TABLE 7.1

Tests for lag length in a near VAR model for the vector $\underset{\sim}{y}_t = (y_t\ p_t\ m_t\ i^*_t\ e_t\ p^*_t)'$ using data from the period 1960M1 - 1988M8

Equation	(7.2i)			(7.2ii)			(7.2iii)			(7.2iv)			(7.2v)			(7.2vi)		
Regressand		Δy			Δp			Δm			Δi^*			Δe			Δp^*	
Lag length Criterionx10^5	AIC	HQIC	SIC	AIC	HQIC	SIC	AIC	HQIC	SIC	AIC	HQIC	SIC	AIC	HQIC	SIC	AIC	HQIC	SIC
1	244	269	308	2.86	3.16	3.61	109	120	137	2.13	2.36	2.69	16.0	17.7	20.2	1.02	1.13	1.29
2	221	249	292	2.81	3.17	3.72	106	119	139	2.15	2.42	2.84	16.2	18.2	21.3	0.941	1.06	1.24
3	226	260	311	2.81	3.24	3.87	108	124	149	2.20	2.53	3.03	16.6	19.1	22.8	0.910	1.05	1.25
4	228	267	326	2.85	3.34	4.08	109	128	156	2.22	2.61	3.19	16.9	19.9	24.3	0.843	0.990	1.21
5	222	265	330	2.75	3.28	4.09	111	132	165	2.27	2.71	3.38	17.4	20.8	25.9	0.845	1.01	1.26
6	225	274	347	2.81	3.42	4.33	115	139	177	2.23	2.71	3.43	17.0	20.6	26.1	0.844	1.03	1.30
7	230	284	365	2.88	3.56	4.59	113	140	180	2.29	2.83	3.65	17.4	21.5	27.7	0.824	1.02	1.31
8	224	281	367	2.90	3.64	4.76	115	145	190	2.34	2.93	3.83	17.7	22.3	29.1	0.790	0.991	1.29
9	223	285	377	2.93	3.73	4.95	119	152	201	2.42	3.08	4.08	17.4	22.1	29.3	0.799	1.02	1.35
10	230	297	399	3.00	3.88	5.20	123	159	213	2.47	3.19	4.28	17.9	23.2	31.1	0.806	1.04	1.40
11	176	230	312	2.96	3.88	5.26	126	166	225	2.45	3.22	4.36	18.4	24.1	32.7	0.808	1.06	1.44
12	147	195	267	3.02	4.01	5.51	121	160	220	2.47	3.29	4.51	18.7	24.8	34.0	0.811	1.08	1.48
13	146	196	272	3.05	4.10	5.68	124	167	231	2.54	3.41	4.73	19.1	25.7	35.6	0.825	1.11	1.54
14	147	201	281	3.14	4.28	5.99	126	172	240	2.56	3.48	4.87	19.8	26.9	37.7	0.829	1.13	1.58
15	147	203	287	3.19	4.39	6.20	127	176	248	2.63	3.62	5.11	20.3	28.0	39.5	0.850	1.17	1.65
16	145	202	288	3.29	4.58	6.53	130	181	258	2.69	3.73	5.33	21.0	29.2	41.6	0.857	1.19	1.70
17	147	207	298	3.35	4.71	6.76	131	185	266	2.73	3.85	5.53	21.6	30.3	43.6	0.882	1.24	1.78
18	140	199	288	3.31	4.71	6.82	134	191	276	2.74	3.89	5.64	22.4	31.8	46.1	0.893	1.27	1.84

AIC is the Akaike criterion, HQIC is an Akaike criterion with a penalty term suggested by Hannan & Quinn (1979) and SIC is the Schwarz criterion. In addition to the components of $\underset{\sim}{y}_t$, the equations include a constant, 11 seasonal dummy variables and a strike dummy taking the value 1 in 1971M2 - 1971M3 and 0 otherwise. For notational simplicity we write $1 - L = \Delta$.

In equation (7.2i) for output, both HQIC and SIC pick out l = 12, whereas AIC indicates an even higher lag length. We suspect that the need to include lags as high as the seasonal lag may arise because of the quasi-seasonal fluctuations in the early spring caused by labor market disputes of varying severity (see section 4.3.1 of chapter 4 for a discussion of this phenomenon). In equation (7.2ii) determining the price level, the choice is between l = 1 or l = 5. The latter lag length was favored over the former one, which left the estimated equation plagued with severely autocorrelated disturbances. In equation (7.2iii) determining M1, a very low lag length would seem admissible, but this was disregarded because of the failure to account for serial correlation in the disturbances in favor of l = 7 hinted by AIC. Similar arguments lead to the choice of l = 6 in equation (7.2iv) for the foreign interest rate and l = 8 in equation (7.2vi) for the foreign price level. In equation (7.2v) for the exchange rate l = 1 was appropriate.

The chosen lag lengths imply a model with more than 300 estimating parameters. Given the available amount of data, this number of parameters poses no problem when it comes to the amount of information from which the parameters are to be estimated. The size of the model nevertheless precludes, by a large margin, the use e.g. of systemwide estimation of the model given the computer facilities at our disposal. Thus, as asserted previously, estimation of the model will be carried out using OLS.[6] Since the choices of lag length were not unanimous in every case, sensitivity analyses with respect to the choice of lag length will be performed.

Proceeding to estimation of the model, we note that the values of the parameters of the model are of no immediate interest. Therefore,

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[6]The size of the model renders the calculation of standard errors for impulse responses and decompositions of variance very cumbersome. Standard methods would require use of derivatives with respect to the model parameters, and, as pointed out by Sims (1986), the approximate Fisher information matrix need not be very accurate. Bootstrap methods would entail re-estimation of the model based on pseudo-histories requiring computing times in excess of 180 hours (on a personal computer for 1000 bootstrap replications). Consequently, we refrain from computing standard errors.

we have condensed the output of the estimation to the summary statistics presented in Table 7.2. Compared to the model in the previous chapter, the current specification performs marginally better in terms of explanatory power and whiteness of disturbances. The disturbances are, again, found to be nonnormally distributed. Likewise, we can reject the null hypothesis of constant coefficients in half of the model equations when the structure is hypothetized to change in 1973M1. The strongest rejection of stability again occurs in the equation for output. Hence impulse responses and decompositions of variance for output based on the full sample should be interpreted as weighted averages of two different patterns. When the breaking point is 1980M1, a rejection of stability can again be detected only in the equation for money.

TABLE 7.2

Selected summary statistics for a near VAR model for the vector $\underset{\sim}{y}_t = (y_t\ p_t\ m_t\ i^*_t\ e_t\ p^*_t)'$ using data from the period 1960M1 - 1988M8

Statistic	Regressand					
	Δy_t	Δp_t	Δm_t	Δi^*_t	Δe_t	Δp^*_t
l	12	5	7	6	1	8
$\bar{R}^2$	0.963	0.307	0.443	0.116	0.099	0.525
Q	0.138	0.606	0.315	0.035	0.999	0.976
JB	0.000	0.000	0.000	0.000	0.000	0.000
CHOW(1973M1)	0.018	0.092	0.019	0.217	0.688	0.043
CHOW(1980M1)	0.996	0.643	0.023	0.116	0.999	0.882

The model (7.2i) - (7.2vi) is estimated by equation-by-equation unconstrained OLS. In addition to the six elements of $\underset{\sim}{y}_t$, the model includes a constant, 11 seasonal dummy variables and in (7.2i) a strike dummy taking the value 1 in 1971M2 - 1971M3 and 0 otherwise. l indicates the lag length, and $\bar{R}^2$ is the degrees-of-freedom-corrected squared multiple correlation coefficient. Q is the Ljung & Box (1978) statistic based on nine autocorrelations, and JB is the Jarque & Bera (1980) statistic. CHOW(.) is the Chow test with the date of the supposed structural break given in parentheses. Marginal significance levels are reported for Q, JB and CHOW(.). For notational simplicity we write 1 - L = Δ.

Having estimated the model, we have a second set of structural disturbances on which to base impulse responses and decompositions of variance. The shocks identified in this chapter show broadly the same characteristics as the one identified in the previous chapter. Results are shown in Appendix 7.

7.2.2 Impulse Responses

Impulse responses based on data from the full-sample period 1960M1 - 1988M8 for a time horizon up to 24 months are presented in Figures 7.1a - 7.1h. All responses are the result of an unexpected, one standard deviation, one period, positive shock to the logarithm of a variable. Since in estimation the data are in the form of differences of logarithmic variables, cumulative responses are displayed in order to enable an interpretation of the responses in terms of logarithmic levels. Since some of the variables are measured in different units, the impulse responses are standardized in terms of fractions of standard deviations (of the residual of the responding variable). Hence, the numerical values of the responses cannot be given any economic interpretation.

As a general observation we note that most of the responses are more ragged than textbook trajectories or responses produced by tightly restricted conventional macroeconometric models. Our results are in line with other VAR studies carried out using stationary data.[7] The raggedness may simply reflect the estimation variability. The somewhat unreliable data may be the underlying reason for this. In the subsequent discussion of estimated impulse responses, we will focus on the qualitative picture emerging when erratic movements are disregarded. A second general observation is that the short-run dynamics dampen out fairly well within two years after a shock.

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[7]See Bernanke & Powell (1986), Kuszczak & Murray (1987), Walsh (1987), Ahmed et al. (1988) and McMillin & Koray (1989). Note that ragged impulses in these studies appear despite the use of quarterly data and efficient estimation procedures.

FIGURE 7.1 Impulse responses for a near VAR model incorporating low frequency restrictions for the vector $\underset{\sim}{y}_t = (y_t\ p_t\ m_t\ i_t^*\ e_t\ p_t^*)'$ using data from the period 1960M1 - 1988M8

FIGURE 7.1a

RESPONSE OF OUTPUT TO SHOCK IN PRICES

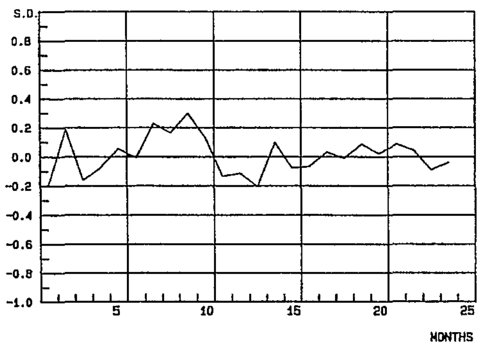

FIGURE 7.1b

RESPONSE OF OUTPUT TO SHOCK IN MONEY

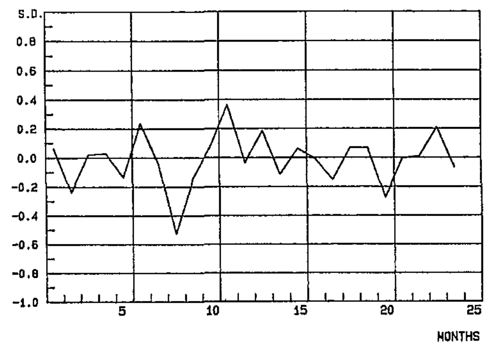

FIGURE 7.1c

RESPONSE OF OUTPUT TO SHOCK IN THE FOREIGN INTEREST RATE

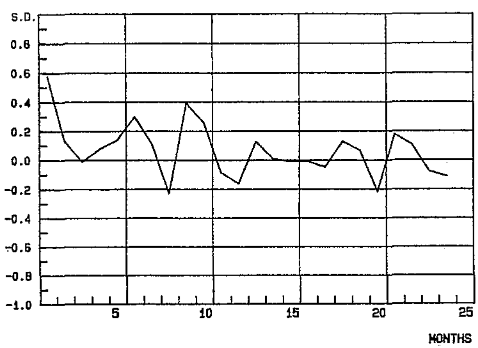

FIGURE 7.1d

RESPONSE OF OUTPUT TO SHOCK IN THE EXCHANGE RATE

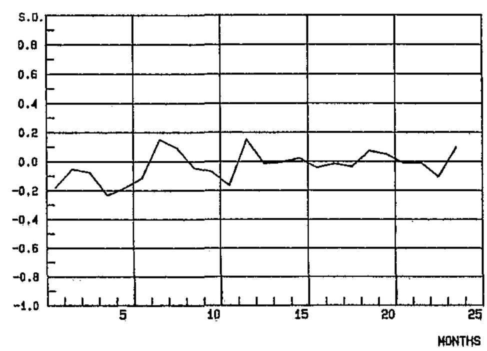

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FIGURE 7.1e

RESPONSE OF OUTPUT TO SHOCK IN FOREIGN PRICES

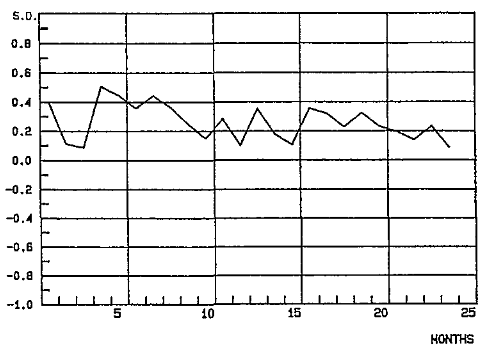


FIGURE 7.1f

RESPONSE OF PRICES TO SHOCK IN FOREIGN PRICES

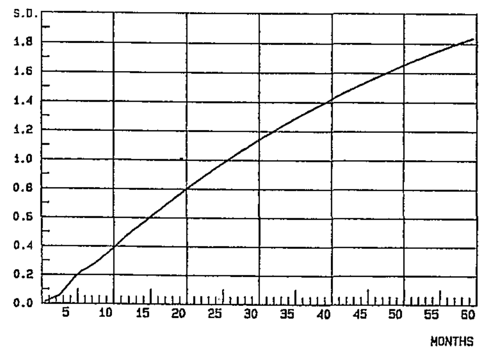


FIGURE 7.1g

RESPONSE OF PRICES TO SHOCK IN OUTPUT

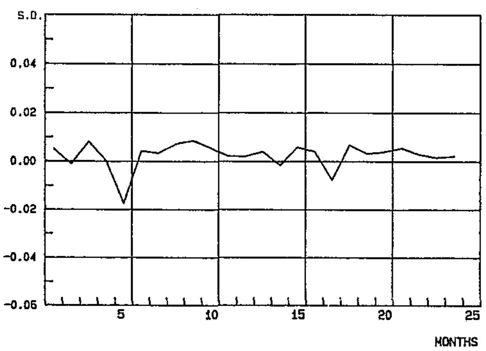
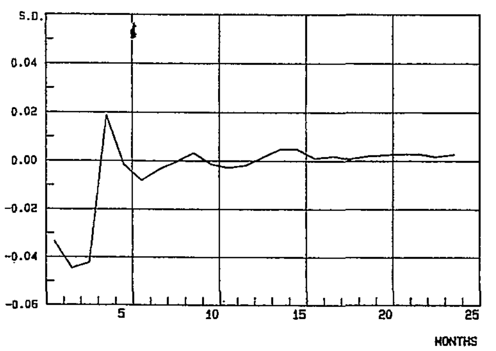


FIGURE 7.1h

RESPONSE OF PRICES TO SHOCK IN THE EXCHANGE RATE



This is in agreement with earlier evidence documented by Öller (1982) in a study of the Finnish economy carried out with a VARIMA model. The response of domestic prices to a shock in foreign prices constitutes a major exception in this respect, however.

The response of real output to an unforeseen upward movement in the domestic price level (Figure 7.1a) appears on the whole to be neutral during the first two quarters after a shock and positive after three quarters. One year after the shock the effect is temporarily contractionary, but output picks up again during the seventh quarter after the shock. The positive response of output to a price shock is what the Lucas supply function employed in our theoretical frame of reference leads us to expect (see expression (6.1v)). Two years after a price shock output has returned more or less to its pre-shock level.

The response of output to a money shock (Figure 7.1b) is slightly more volatile than the response to a price shock, but the short-run impact also seems to be neutral. Output reacts modestly during the first two quarters after the shock, becomes pronouncedly negative during the third quarter, but stands out as markedly positive during the fourth quarter after the shock. The positive effect vanishes a year after the shock, and after one and a half years the response centers, after some oscillation, around zero. The estimation result suggests that money is nonneutral in the short run.

The foreign interest rate affects output (Figure 7.1c) positively during the first three quarters following a shock. After that the impact fluctuates around the pre-shock level. The positive short-run impact may at first sight be unexpected, but we think that the following interpretation of the estimation result has something to it. Since i^* is the only genuinely foreign variable in the model apart from p^*, the interest rate may serve as a proxy for many foreign variables or for foreign economic conditions in general. In particular, it may be a proxy for export demand, as current account deficits abroad imply a higher i^*. Moreover, the positive impact of i^* on y is what our theoretical frame of reference leads us to expect (see expression (6.9i)).

The estimated effect on output of the unexpected part of a devaluation is displayed in Figure 7.1d. The impact is negative during the first six months, with the bulk of the growth stimulus being felt in the third quarter after the devaluation. A second positive peak occurs at the 12-month horizon. A devaluation still causes variability in output toward the latter part of the two-year horizon, but it is not clear whether the effect is positive or negative on average. In any case, we do not register long-lived effects of a devaluation on real total industrial production. We should bear in mind, however, that we are looking at the average importance for output fluctuations of fluctuations in the exchange rate. The temporal importance will be examined at the end of the section focusing on subsample information.

The impact of an unexpected devaluation on the level of output may also testify to the difficulties of modeling the exchange rate, and hence of the difficulty of constructing an exchange rate shock. From a time series modeling point of view, the exchange rate displays problematic behavior with its periods of constancy and discrete jumps mixed with periods of more continuous variability (see Appendix 2). Consequently, equation (7.2v) describing the behavior of the exchange rate explains only one tenth of the movements in the exchange rate (cf. Table 7.2). The rest - nine tenths - of the movements constitutes the unexpected part, but it is not obvious how well these unexpected movements coincide with the true (unobservable) unexpected movements in the exchange rate, i.e. with those movements that actually have impinged on output movements. In view of the ten-year devaluation cycle experienced in Finland, it is possible that a far smaller portion than nine tenths of the movements in the exchange rate may be inherently unexpected historically. Nevertheless, our impulse response closely resembles the empirical evidence from 12 countries presented by Edwards (1986) using conventional regression techniques.

The effect on domestic output of a foreign price shock is portrayed in Figure 7.1e. The impact is positive over the whole two-year horizon. Domestic output is stimulated most during the second quarter after a shock, and there is a systematic tapering off of the effect. The positive short-run effect on output should be interpretable in terms of increased relative competitiveness for domestic firms. Sizewise, shocks to the foreign price level cause bigger fluctuations in domestic output than do shocks to domestic prices. This could be taken as a corroboration of our finding in the previous chapter that the Finnish economy is vulnerable to foreign price developments.

In Figure 7.1f the impact of a foreign price shock on the domestic price level is displayed. Two features of the response are apparent; the impact is slow and smooth. We have increased the horizon of analysis in Figure 7.1f well beyond the short run to illustrate how long it has taken historically for a foreign price shock to work itself into the Finnish economy. The inflationary pressure peaks at four to five months after the shock, and a local maximum can be documented about a year after the foreign shock. After that, the inflationary pressure declines slowly resulting in smoothly rising domestic prices, the impact levelling out only very slowly. We interpret the estimated response as an indication that "domestic" inflation may have deep roots in foreign price movements.

The response of the domestic price level to a shock in domestic output is graphed in Figure 7.1g. The response is persistently positive over the two-year horizon, with the biggest impact in the third quarter after a shock. As the last impulse response estimated from the full-sample period, we turn to the response of the domestic price level to a devaluation (Figure 7.1h). An unforeseen devaluation is deflationary the first quarter after a shock, but causes a sharp positive peak in the price level four months after the devaluation. Following the inflationary burst, the price level fluctuates upwards. The contribution of a devaluation to price level variability is estimated to cease after one and a half years, leaving prices permanently higher than before the devaluation. The above discussion about the possible drawbacks of our measure of exchange rate shocks should be borne in mind.

Some words about the robustness of the above results to alternative operationalizations and augmentations are warranted. The robustness was evaluated using the same alternative operationalizations and augmentations as in the analysis of the long-run impact of shocks as estimated from the full sample (see section 6.2.3 of chapter 6). In addition, the effects on impulse responses of different long-run identifying restrictions were evaluated. Instead of reporting the nearly 200 alternative figures, we comment below only on those cases which yield very different results from those reported in Figures 7.1a - 7.1h.

Estimating the model in level form again turned out to be unsatisfactory because of some unstable impulse responses. Using the real as opposed to nominal exchange rate shifted the bulk of the positive stimulus on growth toward the fourth and fifth quarters after a devaluation. Adding the (nominal or real) price of oil to the model enhanced the growth stimulus of a devaluation at the six-month horizon. Had one worked with a conventional VAR model, the difficulties in interpreting the impulse responses would have been pressing. Thus e.g. a positive "foreign price shock" was seen to have a persistent deflationary impact on domestic prices (sic).

Using different lag lengths mattered to some extent for the results. The most noticeable differences occurred in the responses of output, where the responses became smoother and more rapid the fewer the lags used. Using fewer than nine lags in the equation for output shifted the positive effect on growth of a devaluation toward the first half year following a devaluation. However, since the tests for lag length suggested the use of 12 lags in the equation for output, the evidence produced by employing fewer lags should perhaps not be given too much weight. Assuming that seasonality is a stochastic phenomenon enhanced the growth stimulus of a devaluation at the one-year horizon. This finding may be given more weight, since the quasi-seasonal behavior of output caused by labor market disputes during the early spring is an established fact. Allowing the domestic price level to have a long-run effect on the foreign price level or the exchange rate to have a long-run effect on output hardly changed the results at all.

Before turning to decompositions of variance as estimated from the full sample, it may be of interest to briefly return to the relationship between the above findings and previous results based on heavily restricted, conventional macroeconometric models of the Finnish economy. Because of the qualitative similarity of the numerous results documented for conventional Finnish macroeconometric models, it will suffice to take only a few representative studies as a point of reference.

Among early conventional models, Halttunen (1980) and Halttunen & Korkman (1981) report responses of the domestic price level to a devaluation that are markedly smooth, peak at the six-year horizon and still contribute to price level variability after 18 years (the longest horizon reported being 18 years). The response of output is also smooth and persistent, but the main stimulus of a devaluation occurs at the one-year horizon. Halttunen (op.cit.) estimates that a foreign output shock affects domestic output within one year, but his model also produces a very sustained response of the domestic price level which does not peak within 10 years. Halttunen & Korkman (op.cit.) generate corresponding output responses, and find a smooth effect on domestic output peaking at the three-year horizon and still affecting output at the 18-year horizon. The corresponding effect on the domestic price level peaks at a 11-year horizon.

Representing more recent work, Männistö et al. (1989) examine the effects of a devaluation using the Bank of Finland quarterly model. This model generates an output response peaking at the two- to three-year horizon and continuing to affect output variability 10 years after the devaluation (the longest horizon reported being 10 years). The response of prices peaks at a six-year horizon. Using a small disequilibrium model of the Finnish economy, Aurikko (1986) generates markedly more ragged responses of output and prices to a devaluation. In that study output reacts somewhat unsystematically, while prices peak at the three- to four-year horizon. In both cases, the effects of a devaluation are felt in the variability of output even at the 12-year horizon (the longest reported horizon being 12 years). Conventional macroeconometric models of the newest vintage incorporating rational expectations also, on the whole, produce smooth and relatively slow reactions to disturbances (Aurikko (1988) and Lahti (1989)).

It is evident that our results, in particular the results concerning output, differ markedly from the results generated by Cowles Commission-type models of the Finnish economy. While the direction of impact is generally the same in the two types of models, our results point in some cases toward substantially more rapid responses to shocks.

7.2.3 Decompositions of Variance

The decompositions of variance based on the full-sample period are reported in Table 7.3. Three quarters of the variability of output is accounted for by the inherent dynamics of the output process itself at the one- to two-year horizon. Three quarters after the occurrence of a money shock or a foreign interest rate shock, nonnegligible portions of the variability of output can be attributed to these factors as well. Shocks to domestic and foreign prices and exchange rate shocks do not account for important portions of the variability of the real variable.

Only the foreign price level seems to have made a substantial contribution to domestic price level variability. The impact becomes nonnegligible about a year after the occurrence of a foreign price shock, and it is strong at the two-year horizon. Given the stretched out impulse response of a foreign price shock on the domestic price level documented in the previous section, this reinforces the impression of the importance of foreign inflationary pressures. The foreign variables and the policy variable (the exchange rate) are essentially exogenous as expected. In particular, the domestic price level accounts for hardly any of the variability of the foreign price level.

TABLE 7.3

Variance decomposition for a near VAR model for the vector $\underset{\sim}{y}_t = (y_t\ p_t\ m_t\ i^*_t\ e_t\ p^*_t)'$ using data from the period 1960M1 - 1988M8

Percentage of the expected k-step-ahead squared prediction error in variable	k	Innovation in y	p	m	i*	e	p*
y	1	94.0	0.400	0.036	3.51	0.341	1.68
	2	90.6	1.58	0.791	4.53	0.413	2.05
	3	89.4	2.50	1.30	4.48	0.398	1.96
	6	86.2	2.63	2.56	4.60	0.634	3.37
	9	78.9	2.96	5.62	8.05	1.23	3.26
	12	75.5	3.44	7.19	8.48	1.88	3.47
	24	75.5	3.34	7.74	8.01	1.82	3.57
p	1	0.012	99.1	0.187	0.126	0.525	0.068
	2	0.028	97.1	0.192	1.85	0.565	0.307
	3	0.064	96.0	0.370	2.60	0.548	0.461
	6	0.388	88.9	0.619	3.89	2.11	4.08
	9	0.385	87.6	0.690	3.90	2.07	5.35
	12	0.383	85.5	0.708	3.87	2.03	7.48
	24	0.520	80.0	0.686	3.63	1.91	13.3
m	1	0.002	0.001	99.3	0.426	0.249	0.057
	2	0.099	0.654	96.9	0.845	1.29	0.190
	3	0.386	0.657	95.7	1.77	1.28	0.207
	6	0.908	1.46	92.4	2.16	2.16	0.938
	9	2.00	1.72	90.5	2.48	2.22	1.10
	12	2.09	1.80	90.2	2.55	2.24	1.13
	24	2.89	1.91	89.0	2.65	2.25	1.29
i*	1	0.625	0.119	0.076	98.5	0.384	0.271
	2	0.598	0.189	0.408	97.4	0.874	0.489
	3	0.786	0.329	0.405	96.6	0.936	0.991
	6	1.15	0.611	1.08	95.0	1.16	1.04
	9	1.27	0.644	1.19	94.1	1.26	1.51
	12	1.28	0.780	1.22	93.9	1.30	1.54
	24	1.57	0.815	1.27	92.9	1.31	2.09
e	1	0.014	0.137	0.000	0.057	99.8	0.022
	2	0.035	0.127	0.000	0.058	99.7	0.080
	3	0.034	0.127	0.000	0.064	99.7	0.087
	6	0.036	0.128	0.004	0.069	99.7	0.094
	9	0.037	0.130	0.005	0.071	99.7	0.103
	12	0.037	0.130	0.007	0.072	99.6	0.117
	24	0.054	0.131	0.010	0.074	99.6	0.146
p*	1	0.051	0.066	0.227	0.009	0.000	99.6
	2	0.064	0.199	0.389	0.415	0.050	98.9
	3	0.205	0.194	2.47	0.657	0.074	96.4
	6	0.667	0.984	2.94	3.04	0.484	91.9
	9	0.813	2.39	3.23	2.69	0.426	90.5
	12	0.846	2.35	3.53	2.60	0.558	90.1
	24	0.946	2.03	3.07	2.24	0.504	91.2

The model (7.2i) - (7.2vi) is estimated by equation-by-equation unconstrained OLS. In addition to the six elements of $\underset{\sim}{y}_t$, the model includes a constant, 11 seasonal dummy variables and in (7.2i) a strike dummy taking the value 1 in 1971M2 - 1971M3 and 0 otherwise. Numbers across rows may not sum to 100.000 due to rounding erros.

With respect to the robustness of the decompositions of variance the following observations can be made. The suggestion that money has mattered for output variability is robust to all but one alteration. Only treating seasonality as a stochastic phenomenon nullifies the contribution of money. Using the domestic interest rate instead of the foreign interest rate or as an auxiliary variable changes the basic results only a little. Roughly one tenth of the variability of output at the one- and two-year horizons is accounted for by the (more or less exogenous) domestic interest rate. Hence, monetary policy has had effects historically in Finland. The price of oil has, on average, accounted for some eight to nine per cent of the variability of the foreign price level but for only four to five per cent of the variability of domestic output. Only a little of domestic output variability comes from foreign output movements.

7.2.4 Analysis of Subperiods

Summary statistics for the estimated subsample models are collected in Table 7.4. The adequate number of lags differs from the full-sample results and across subperiods. Explanatory power is fairly high during the first subsample, drops during the second subsample, but picks up during the last subsample. This could be an outcome of the fact that large shocks occurred mainly during the second subperiod while the third subperiod has been extraordinarily calm.

Subsample impulse responses are portrayed in Figures 7.2a - 7.2k. Again, most responses appear more ragged than textbook trajectories or impulse responses generated by Cowles Commission-models. Also, the short-run dynamics appears, on the whole, to converge almost fully within the two-year horizon. Exceptions to this are the responses of output during the period 1973M1 - 1988M8. In addition, these responses are, on the whole, more volatile than other responses, and they are also more volatile than the corresponding responses estimated from the first and the third subsamples. All in all, the estimated short-run impulse responses generally differ across subperiods, and they also generally differ from full-sample responses.

TABLE 7.4

Selected summary statistics for near VAR models using data from subperiods

Statistic	Time period / Regressand	1960M1 - 1972M12					
		Δy_t	Δp_t	Δc_t	Δi^*_t	Δe_t	Δp^{02}_t
l		7	6	8	6	7	6
$\bar{R}^2$		0.962	0.972	0.932	0.309	0.159	0.437
Q		0.002	0.012	0.361	0.152	0.878	0.207

Statistic	Time period / Regressand	1973M1 - 1988M8					
		Δy_t	Δp_t	Δm_t	Δi^*_t	Δe_t	Δp^*_t
l		12	6	2	2	1	4
$\bar{R}^2$		0.974	0.559	0.400	0.122	0.011	0.543
Q		0.100	0.074	0.061	0.090	0.735	0.228

Statistic	Time period / Regressand	1980M1 - 1988M8				
		Δy_t	Δp_t	Δm_t	Δi^*_t	Δe_t
l		8	4	2	2	5
$\bar{R}^2$		0.983	0.601	0.499	0.038	0.304
Q		0.612	0.029	0.067	0.001	0.051

In addition to the variables in the models, the models comprise a constant, 11 seasonal dummy variables and, where appropriate, a strike dummy taking the value 1 in 1971M2 - 1971M3 and 0 otherwise. The models are estimated by equation-by-equation OLS using the restrictions reported in expressions (7.3) - (7.5), respectively. l indicates the lag length, $\bar{R}^2$ is the degrees-of-freedom-corrected squared multiple correlation coefficient, and Q is the Ljung & Box (1978) statistic based on four, four and three autocorrelations, respectively. Marginal significance levels are reported for Q. For notational simplicity we write $1 - L = \Delta$.

FIGURE 7.2 Impulse responses for near VAR models incorporating low frequency restrictions using data from subperiods

FIGURE 7.2a

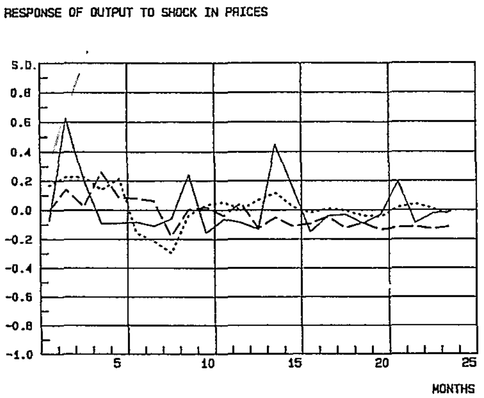


FIGURE 7.2b

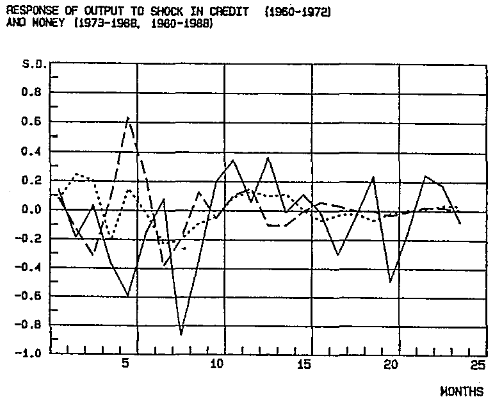


FIGURE 7.2c

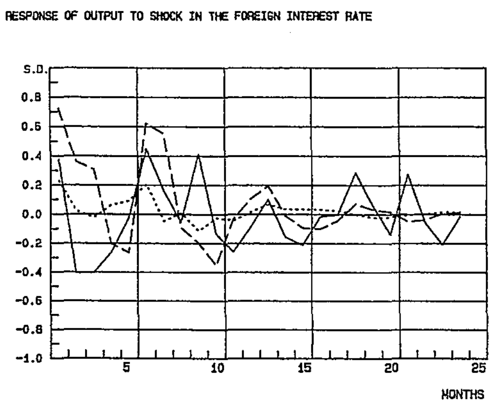
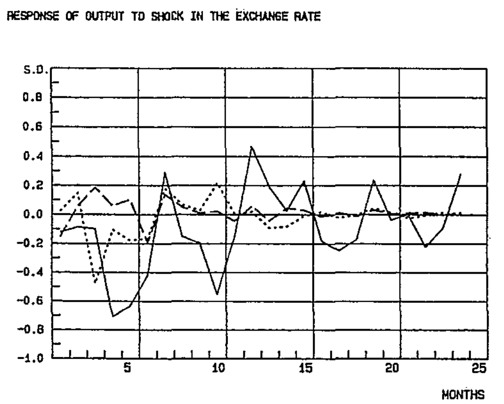


FIGURE 7.2d



Notes: --------- 1960M1 - 1972M12
——————— 1973M1 - 1988M8
— — — — — 1980M1 - 1988M8

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FIGURE 7.2e

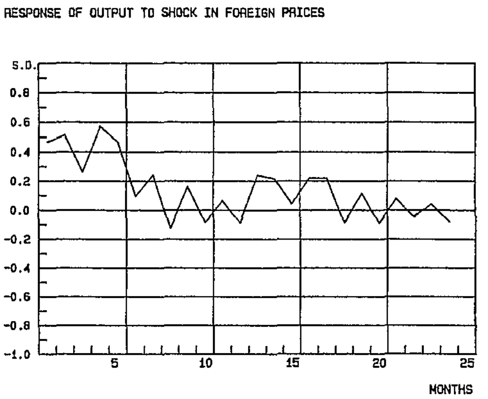


FIGURE 7.2f

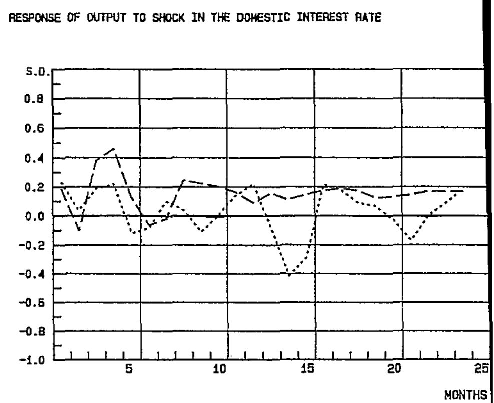


FIGURE 7.2g

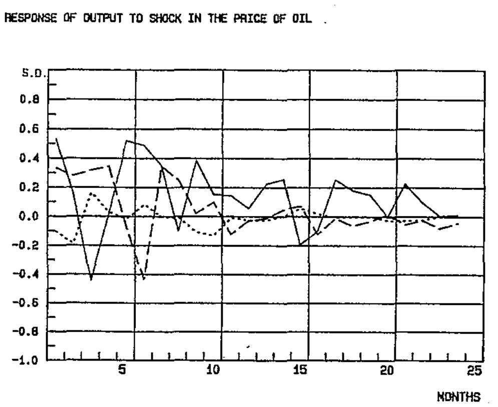
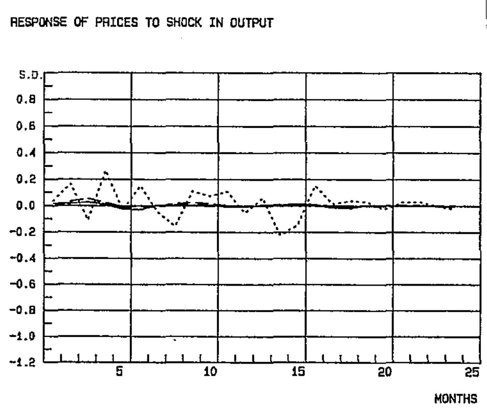


FIGURE 7.2h



Notes: --------- 1960M1 - 1972M12
——————— 1973M1 - 1988M8
— — — — — 1980M1 - 1988M8

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FIGURE 7.2i

RESPONSE OF PRICES TO SHOCK IN THE EXCHANGE RATE

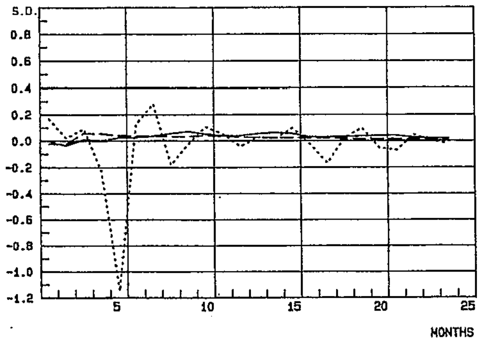


FIGURE 7.2j

RESPONSE OF PRICES TO SHOCK IN FOREIGN PRICES

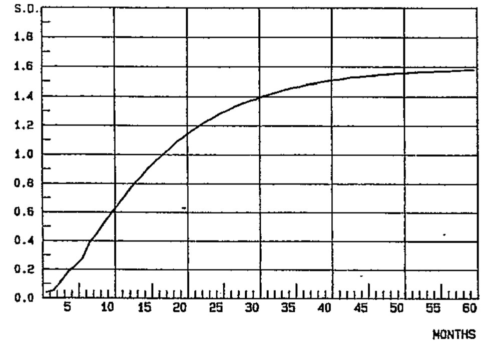
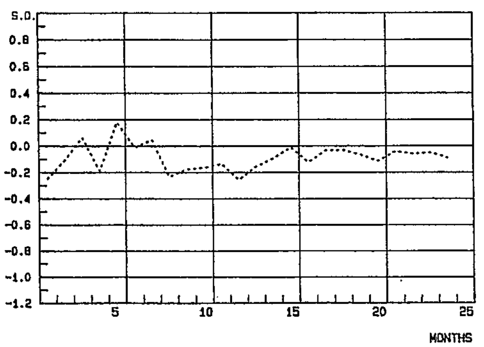


FIGURE 7.2k

RESPONSE OF PRICES TO SHOCK IN THE PRICE OF OIL



Notes: --------- 1960M1 - 1972M12

——————— 1973M1 - 1988M8

– – – – – – 1980M1 - 1988M8

The result that impulse responses for output estimated from the full post-Bretton Woods period are more volatile and less convergent than responses estimated from other periods may be interpreted in the following way. Recalling that the first two years of the post-Bretton Woods subperiod witnessed an extraordinary amount of severe shocks, and that the economic structure was found to be nonconstant across the two long subperiods, our findings point toward sustained structural change during the years 1973 - 1980. In other words, the turbulent years 1973 - 1975 may have initiated slow but profound structural changes that took many years to complete. During those years, the economy was more vulnerable to shocks than usual, and impulse responses based on data from these times reflect both the increased vulnerability to shocks and the ongoing structural change.

Proceeding to the responses of output to shocks of various kinds (Figures 7.2a - 7.2g), we begin by scrutinizing the estimates based on the subperiod 1960M1 - 1972M12. The real variable is initially affected positively by a price shock, negatively at the six-month horizon, and again positively at the nine-month horizon. After that, the response dampens out in smooth oscillations. The response is far from neutral. An unforeseen increase in credit advanced to the public fosters output growth during the first quarter while curbing output in the third quarter after a shock. One year after a shock the impact on the level of output is positive, and after the fifth quarter the effect dies out. We interpret this as a short-run nonneutrality originating in the credit rationing prevalent during the 1960s and the early 1970s. Also, the finding may attest to potency of monetary policy during this period.

The response of output to a foreign interest rate shock is positive during the first two quarters, whereafter the effect oscillates smoothly, converging to the pre-shock level. A somewhat more pronounced impact is documented for the effects of a devaluation on the real variable. What essentially emerges is a negative stimulus to growth during the second quarter after the devaluation, and a positive stimulus to growth during the third quarter. The net effect is zero at the two-year horizon. This impact is not altogether different from the one estimated using the full sample.

The responses of the domestic price level to shocks of various kinds are reported in Figures 7.2h - 7.2k. A striking feature of the response to an output shock is how much more volatile the response appears to have been during the first subsample period than during other times. This may reflect the relative closedness of the Finnish economy during the 1960s and the early 1970s. A similar finding with a similar explanation is documented in the response to a devaluation. The devaluation is estimated to have exerted an inflationary impact during the first quarter after the shock, but a sizeable deflationary effect is found during the second quarter after the devaluation. The domestic price level is again higher than before the devaluation seven months after the shock, but the response is erratic thereafter. It is likely that our difficulties in constructing an exchange rate shock show up in this impulse response.

Moving on to impulse responses estimated from the 1973M1 - 1988M8 subsample, we begin by scrutinizing the responses of domestic output to shocks of various kinds. The response of output to a domestic price level innovation is more volatile than during the first subsample period. While the response reveals a nonneutral impact of a nominal variable on the real variable, any systematic impact is hard to detect. The effect of a money shock also emerges as more volatile than the full-sample estimate, but the response is not very systematic. Money emerges as markedly nonneutral in the short run while stimulating output most at a horizon of 10 to 13 months and continuing to show up in output movements throughout the two-year horizon.

A foreign interest rate shock is estimated to have exerted a considerable depressive impact on domestic growth during the second subsample at the two- to four-month horizon. On the other hand, the impact stimulated output six months after the shock. The rest of the estimated impact is erratic and nonconvergent. We take this to mean that while the foreign interest rate has historically contributed to domestic output variability, the impact has, on the whole, been somewhat fuzzy, possibly proxying for many foreign variables (see the discussion in section 7.2.2). Likewise, the effects of a devaluation on output emerge as unbelievably erratic, yet qualitatively resembling the full-sample estimate. Again, we suspect that the difficulties in modeling this variable explain the estimation result.

The impact of a foreign price shock resembles the full-sample impact; positive during the first two quarters and during the fifth and sixth quarter following a shock. However, the impact is essentially zero during the fourth quarter after a shock, and output returns to its pre-shock level one and a half years after the shock. The impact is thus weaker than when the Bretton Woods period experience is allowed to affect the estimate. The response of the domestic price level to an unforeseen upward movement in domestic output is weaker than during the first subperiod, whereas the domestic price level is raised permanently by a devaluation during the second subperiod.

Responses based on data from the 1980s only show the following features. Output has initially responded positively to a price shock, converging to zero after six months, and turning persistently negative thereafter. In other words, the decrease in inflationary pressures during the 1980s may have contributed to a strengthening of output growth. Or, to put it in yet another way: rises in the price level may serve to stimulate output at a one-year horizon, but the impact in the long run is detrimental. Money again seems to affect the real variable to a considerable extent. The positive response peaks five to six months after the shock, and the effects dampen out roughly within one and a half years after the shock. According to the response, it would be hard to claim anything other than that money matters.

If money matters, so does the foreign interest rate. Apart from an outlying positive peak at the six-month horizon, the impact on output growth is negative during the two to 10 months following the disturbance. However, the impact on the level of output is mixed, possibly reflecting the fact that the foreign interest rate proxies for foreign economic conditions at large. The effect of a devaluation is positive. The devaluations in 1982 and the small exchange rate adjustment of 1986 were not the independent cause of any sizeable inflationary pressure. The subsample impulse responses are, on the whole, robust to the same augmentations and different operationalizations as in the corresponding section of the previous chapter.

Subsample decompositions of variance are reported in Tables 7.5 - 7.7. We begin by looking at Table 7.5, which reports results for the subperiod 1960M1 - 1972M12. On the whole, some differences exist between these results and the full-sample results. The exchange rate has accounted for nontrivial portions of output variability during the first subperiod. Credit may also have contributed weakly to output variability. The effects of monetary policy in terms of the supply of credit show up weakly six months after a policy shock while the effect of a devaluation could be felt already three months after the devaluation. A year after the devaluation, the exchange rate adjustment accounts for one tenth of the variability of output. The foreign interest rate is of no importance for domestic output movements during the 1960s and the early part of the 1970s.

Credit and the exchange rate account for nonnegligible portions of domestic price movements. Unexpected movements in credit show up in price level movements three quarters after a shock, while the effects of a devaluation were felt strongly already two quarters after the devaluation. At the one-year horizon, movements in credit and the exchange rate account for nearly three tenths of the variability of the domestic price level. The supply of credit has reacted mainly to domestic output developments and foreign economic conditions. As regards the exchange rate, some degree of feedback between exchange rate movements and domestic prices may have existed. The foreign interest rate and the price of oil are largely exogenous. The price of oil does not appear to have contributed noteworthily, on average, to short-run variability during the first subperiod.

TABLE 7.5

Variance decomposition for a near VAR model for the vector $\underset{\sim}{y}_t = (y_t\ p_t\ c_t\ i^*_t\ e_t\ p^{02}_t)'$ using data from the period 1960M1 - 1972M12

Percentage of the expected k-step-ahead squared prediction error in variable	k	Innovation in y	p	c	i*	e	p^{02}
y	1	97.9	0.592	0.112	1.15	0.013	0.234
	2	96.6	0.543	0.611	1.68	0.275	0.311
	3	88.9	0.478	0.567	1.51	6.34	2.17
	6	80.5	2.60	5.15	1.66	7.77	2.37
	9	76.9	3.46	5.55	2.64	9.00	2.45
	12	75.8	3.47	5.68	2.71	9.81	2.58
	24	74.9	3.76	5.94	2.74	10.0	2.63
p	1	0.013	94.2	0.721	4.12	0.293	0.662
	2	0.159	86.5	3.41	8.78	0.425	0.715
	3	0.781	86.2	3.36	8.26	0.441	0.909
	6	2.31	67.4	4.32	7.08	16.7	2.21
	9	2.71	59.9	12.7	6.23	16.1	2.40
	12	2.84	59.4	12.6	6.68	16.1	2.47
	24	3.79	58.3	13.0	6.47	15.8	2.60
c	1	0.158	0.089	97.0	2.14	0.552	0.092
	2	1.03	0.074	95.1	1.51	0.402	1.90
	3	1.43	0.336	92.3	3.33	0.711	1.85
	6	3.36	2.27	81.7	7.50	0.990	4.21
	9	6.23	3.24	76.5	8.02	1.01	4.98
	12	9.99	3.18	72.5	8.08	1.21	5.04
	24	10.2	3.31	71.9	8.06	1.35	5.17
i*	1	2.50	2.19	0.030	95.1	0.101	0.039
	2	2.64	2.29	2.54	91.5	0.605	0.441
	3	2.48	2.11	2.88	90.0	1.03	1.47
	6	2.52	6.77	3.39	82.8	2.65	1.87
	9	3.16	6.44	6.41	79.3	2.75	1.93
	12	3.29	6.68	7.50	77.4	3.29	1.87
	24	3.99	6.95	8.15	75.8	3.23	1.91
e	1	0.042	0.213	0.029	0.009	97.6	2.10
	2	0.175	0.287	0.193	0.462	96.8	2.12
	3	0.189	0.280	0.478	0.518	94.5	4.02
	6	0.897	1.17	1.30	1.15	91.2	4.31
	9	1.04	6.81	2.11	1.62	84.4	4.04
	12	1.15	7.04	2.66	1.63	83.3	4.26
	24	1.62	7.29	3.45	1.71	81.6	4.31
p^{02}	1	0.013	0.472	0.454	0.585	3.96	94.5
	2	5.18	1.01	0.395	0.829	4.02	88.6
	3	6.37	0.989	1.51	0.904	3.70	86.5
	6	6.12	1.28	2.79	2.77	3.87	83.2
	9	5.97	1.60	3.09	2.49	3.59	83.3
	12	5.85	1.68	3.62	2.42	3.46	83.0
	24	5.93	1.63	3.83	2.31	3.31	83.0

The model is estimated by equation-by-equation OLS using the restrictions reported in expression (7.3). In addition to the six elements of $\underset{\sim}{y}_t$, the model includes a constant, 11 seasonal dummy variables and in the equation for y_t a strike dummy taking the value 1 in 1971M2 - 1971M3 and 0 otherwise. Numbers across rows may not

Decompositions of variance based on the subperiod 1973M1 - 1988M8 are reported in Table 7.6. Again, some differences can be detected between, on the one hand, these results and full-sample results, and, on the other hand, between these results and the first subsample results. Output emerges as substantially more endogenous during the latter part of the 1970s and the 1980s than during the 1960s and the early part of the 1970s. Money, exchange rate and foreign interest rate shocks, and possibly also domestic and foreign price shocks, all account for nonnegligible parts of the variability of output during the second subperiod. Taken together, these shocks account for half the variability of output at the one- and two-year horizons. The largest individual influence appears to come from money, which accounts for more than one sixth of output movements at these horizons. The influence of money and the exchange rate show up at the three-quarter horizon, while the influence of the foreign interest rate is already felt two months after a shock. Output reacts more quickly to domestic than to foreign price developments.

The domestic price level emerges as being dependent on movements in the other model variables during the second subperiod. At the one-year horizon more than two fifths, and at the two-year horizon one half, of the movements in the domestic price level are caused by other variables. More specifically, the other notable contributions to domestic price level variability come from the foreign price level and the foreign interest rate. Unforeseen movements in the foreign price level start showing up in the domestic price level in the second quarter after a shock, and nearly one third of the unforeseen movements in the domestic pricel level are caused by foreign prices at the two-year horizon. Foreign interest rate shocks contribute to domestic price level variability two months after a disturbance. All other variables are more or less exogenous.

Moving on to the evidence based on data from the 1980s only, the following emerges. Six months after a shock, the foreign interest rate accounts for one quarter, and money shocks for one tenth, of the unforeseen movements in output. The foreign interest rate accounts for sizeable portions of domestic price level variability. The exchange rate has reacted to domestic price movements and to a lesser extent to foreign interest rate movements.

TABLE 7.6

Variance decomposition for a near VAR model for the vector $\underset{\sim}{y}_t = (y_t\ p_t\ m_t\ i_t^*\ e_t\ p_t^*)'$ using data from the period 1973M1 - 1988M8

Percentage of the expected k-step-ahead squared prediction error in variable	k	Innovation in y	p	m	i*	e	p*
y	1	93.4	0.092	0.314	2.31	0.250	3.60
	2	80.0	6.15	1.58	9.31	0.208	2.75
	3	78.4	7.65	2.00	8.52	0.193	3.24
	6	67.7	7.33	5.56	10.1	4.20	5.09
	9	54.2	6.59	13.8	10.8	8.74	5.95
	12	48.3	6.94	14.9	11.8	12.2	5.94
	24	48.1	7.67	16.0	10.8	11.6	5.84
p	1	0.309	95.2	0.056	3.19	0.004	1.27
	2	0.276	84.1	1.39	12.0	0.967	1.25
	3	0.343	79.2	1.71	12.2	2.68	3.79
	6	1.26	69.3	2.79	15.2	2.77	8.67
	9	1.78	60.0	3.78	13.3	2.68	18.5
	12	1.71	56.2	3.65	12.4	2.79	23.3
	24	2.24	49.5	3.42	11.0	3.08	30.8
m	1	0.084	0.025	98.5	0.334	0.330	0.744
	2	0.307	0.022	97.7	0.294	0.788	0.875
	3	0.336	0.206	97.2	0.556	0.785	0.880
	6	0.368	0.228	97.0	0.663	0.835	0.904
	9	0.377	0.259	96.8	0.697	0.863	0.979
	12	0.378	0.279	96.7	0.734	0.873	0.991
	24	0.583	0.317	96.4	0.768	0.926	1.03
i*	1	0.005	0.047	0.002	97.7	2.08	0.210
	2	0.131	0.476	0.013	96.0	1.99	1.44
	3	0.301	2.76	0.191	91.7	3.05	2.01
	6	0.328	3.05	0.251	90.6	3.54	2.26
	9	0.380	3.28	0.407	89.9	3.60	2.39
	12	0.380	3.33	0.413	89.7	3.66	2.47
	24	0.551	3.36	0.479	88.9	3.73	2.93
e	1	0.050	0.013	0.240	0.258	99.4	0.008
	2	0.110	0.053	0.454	0.264	99.1	0.011
	3	0.111	0.056	0.460	0.304	99.1	0.018
	6	0.117	0.059	0.494	0.320	99.0	0.023
	9	0.118	0.062	0.527	0.324	98.9	0.027
	12	0.119	0.066	0.533	0.334	98.9	0.030
	24	0.182	0.077	0.557	0.346	98.8	0.039
p*	1	0.014	0.359	0.007	0.001	0.201	99.4
	2	0.102	0.891	0.383	0.111	0.873	97.6
	3	0.093	1.56	1.18	0.128	2.52	94.5
	6	0.112	1.51	1.38	1.37	2.89	92.7
	9	0.110	1.38	1.34	1.25	2.71	93.2
	12	0.110	1.31	1.28	1.21	2.60	93.5
	24	0.160	1.24	1.21	1.16	2.50	93.7

The model is estimated by equation-by-equation OLS using the restrictions reported in expression (7.4). In addition to the six elements of $\underset{\sim}{y}_t$, the model includes a constant and 11 seasonal dummy variables. Numbers across rows may not sum to 100.000 due to rounding errors.

TABLE 7.7

Variance decomposition for a near VAR model for the vector $\underset{\sim}{y}_t = (y_t\ p_t\ m_t\ i^*_t\ e_t)'$ using data from the period 1980M1 - 1988M8

Percentage of the expected k-step-ahead squared prediction error in variable	k	Innovation in				
		y	p	m	i*	e
y	1	82.3	0.002	0.205	16.8	0.756
	2	83.0	0.442	1.07	14.0	1.42
	3	81.8	0.723	1.81	13.9	1.76
	6	61.4	1.73	10.1	24.1	2.67
	9	53.5	2.57	14.5	25.7	3.74
	12	52.4	2.62	14.7	26.5	3.79
	24	51.2	3.14	15.2	26.5	3.95
p	1	0.104	97.3	0.561	0.911	1.09
	2	2.16	77.0	0.432	19.6	0.842
	3	2.30	71.5	1.59	16.5	8.13
	6	3.60	62.4	1.84	25.4	6.74
	9	3.81	62.2	2.99	24.5	6.41
	12	3.86	61.3	3.40	25.2	6.27
	24	3.88	60.9	3.52	25.5	6.16
m	1	0.225	0.000	97.4	2.21	0.158
	2	0.245	0.006	96.4	3.17	0.216
	3	0.981	0.065	95.3	3.37	0.239
	6	2.26	0.100	93.8	3.61	0.270
	9	2.57	0.136	92.8	4.08	0.438
	12	2.58	0.208	92.6	4.20	0.448
	24	2.61	0.241	92.4	4.24	0.464
i*	1	0.389	0.010	0.278	99.0	0.301
	2	0.409	0.095	0.546	98.6	0.307
	3	1.11	0.139	2.58	95.6	0.567
	6	1.41	0.229	2.70	95.1	0.574
	9	1.48	0.236	3.25	94.4	0.642
	12	1.52	0.270	3.30	94.2	0.656
	24	1.54	0.286	3.34	94.2	0.664
e	1	0.005	0.001	0.005	0.119	99.9
	2	0.056	20.6	0.348	4.45	74.5
	3	0.595	30.8	0.309	4.02	64.3
	6	2.88	30.0	0.833	8.97	57.3
	9	3.15	28.9	2.14	9.90	56.0
	12	3.16	28.1	3.48	11.3	53.9
	24	3.29	28.2	3.65	11.8	53.0

The model is estimated by equation-by-equation OLS using the restrictions reported in expression (7.5). In addition to the five elements of $\underset{\sim}{y}_t$, the model includes a constant and 11 seasonal dummy variables. Numbers across rows may not sum to 100.000 due to rounding errors.

Before leaving conventional subsample decompositions of variance, some comments on the robustness of the results are in order. With regard to the period 1960M1 - 1972M12, augmenting the model with the domestic interest rate, foreign output or foreign prices all nullify the weak contribution of credit that was documented in Table 7.5. The auxiliary variables themselves do not add substantially to the model, and the other basic results also remain practically unaltered. We thus conclude that while credit advanced to the public has had a nonneutral impact on the real variable during the 1960s and the early 1970s, the quantitative impact on output has been on the whole minimal. These findings are in line with the results of Raatikainen & Takala (1985), who using causality tests conclude that the supply of credit has not been causally prior to real output.

Augmenting the model for the period 1973M1 - 1988M8 with the domestic interest rate reveals that it has shown up weakly in domestic output movements with a lag of one quarter. Foreign output accounts for just over one tenth of domestic output variability in the short run. The inclusion furthermore downplays the roles of money and the foreign interest rate. We thus again register a caveat to the view that monetary factors are of quantitative importance for output fluctuations. Our results resemble the early international VAR findings according to which the importance of money vanishes when sparse models are enlarged. Fiscal policy may have accounted for more than one tenth of domestic output variability in the short run. This effect is both stronger and starts to operate more rapidly than do unforeseen movements in the domestic interest rate. The proxy for fiscal policy is, on the whole, exogenous.

Five to 14 per cent of the short-run movements in domestic output have been caused by oil price fluctuations during the period beginning with the first oil shock. Movements in the import price index for energy yield the lower estimated percentages while genuine oil price movements yield estimates in excess of 10 per cent. Basing the estimate on a period essentially containing only oil price rises (1973M1 - 1985M12) attributes one fifth of output variability to the oil price rises in the very short run, and one quarter of output variability at the one-year horizon. Hence, while average unforeseen oil price movements have been fairly unimportant for domestic output variability, the exceptionally large oil shocks have had profound impacts on output movements. Again, the distinction between nominal and real oil prices is not quantitatively important.

Turning to sensitivity analyses for the period 1980M1 - 1988M8, the following emerges. Money matters for output developments, and this is robust to all our alterations of the analysis. One tenth to one fifth of output variability comes from fluctuations in the stock of narrow money.[8] At the one-year horizon, the domestic interest rate accounts for approximately one tenth of the variability of output, while one seventh of the fluctuations in output can be attributed to the foreign interest rate. The domestic interest rate, in turn, is found to be influenced most by the foreign interest rate. A nonnegligible impact occurs with a one-month lag, and roughly one quarter of the variability of the domestic interest rate is accounted for by the foreign interest rate at the one-year horizon. Output, domestic prices and money also deliver nonnegligible contributions to domestic interest rate variability in the short run.

Controlling for foreign output movements again reveals an effect of just over one tenth for foreign output for domestic short-run output variability. Only one quarter of the variability of domestic output is attributed to the inherent dynamics of the output process itself during the 1980s. The fiscal policy variable accounts for one sixth of output variability at the one-year horizon. Foreign price movements account for one sixth of the variability of the domestic price level at the two-year horizon.

Some 12 to 22 per cent of the short-run movements in domestic output have been caused by oil price fluctuations during the 1980s. One

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[8]A proponent of the real business cycle model might appeal to the notion of "reverse causation", and argue that the observed correlation between money and output comes from the response of M1 to technology shocks which are the true causes of output fluctuations. We would not subscribe to such an interpretation, however, since it can be argued that technology shocks are controlled for in our analysis through the innovations in output.

explanation for the increased role of the price of oil may lie in the different importance of the bilateral trade with the COMECON countries (see the discussion in section 6.2.1 of chapter 6). Movements in the import price index for energy yield higher estimated percentages while genuine oil price movements yield estimates smaller than 16 per cent. This is the reverse of the finding based on the full post-Bretton Woods period. It may be a consequence of the fact that substitution away from oil-intensive production has occurred after the first oil shock.[9] Again, the distinction between nominal and real oil prices is not quantitatively important.

We conclude the subsample analysis by utilizing both impulse responses and decompositions of variance to go beyond the division of the full sample into three subsamples. Shortening the time span of the analysis even further, it may be of interest to assess the role that different shocks played in specific historical episodes. This can be accomplished by calculating moving weighted averages of short-run structural forecast errors in output. The weights are given by the estimated cumulative impulse responses for output, and the forecast horizon is set at two years in order to conform to our definition of the short run. Hence, the estimates have the interpretation of components of two-year ahead forecast errors in output. The basic full sample specification is used, and the empirical evidence is presented in Figure 7.3.

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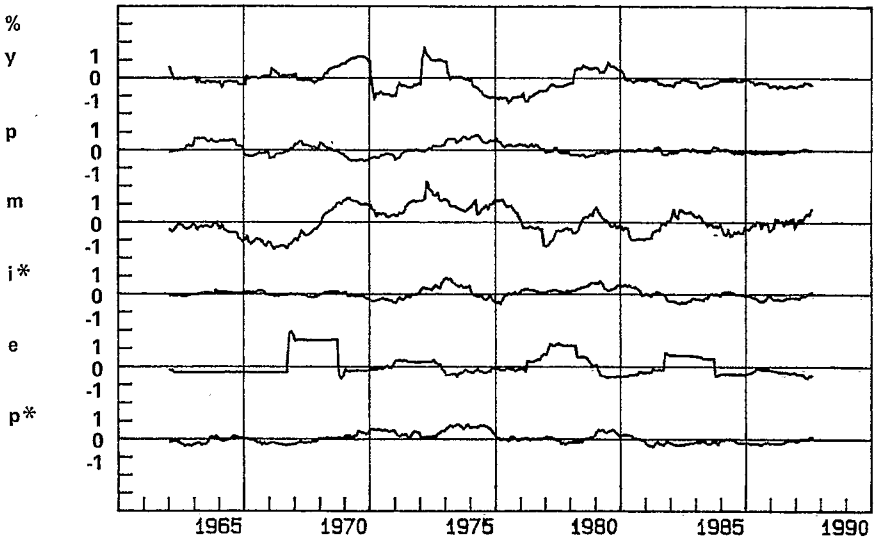
[9]Between 1973 and 1986, the proportions of energy consumed by Finnish industry changed as follows: oil products declined from 45 per cent to 22 per cent, electricity increased from 20 per cent to 27 per cent, and natural gas (which was new on the market) increased by nine per cent (The Price of Energy as a Factor Influencing the Investment Decisions of Industrial Companies (1988)). In addition, Ilmakunnas & Törmä (1989) present some evidence of structural change in factor substitution in Finnish manufacturing.

FIGURE 7.3

Components of forecast error for output as identified from the vector $\underset{\sim}{y}_t = (y_t\ p_t\ m_t\ i^*_t\ e_t\ p^*_t)'$ using low frequency restrictions and data from the period 1960M1 - 1988M8



We first turn to output developments, and notice inherently strong growth in the late 1960s, followed by a decline in 1970, which culminated in the major strike in early 1971. Output barely has time to recover from the strike before the prolonged and sizeable downturn following the first oil price shock set in. Some improvement in output developments is found during the latter half of the 1970s, but the second oil price shock again cuts back output growth somewhat. A striking feature of the remainder of the estimate is the long and very steady growth of the 1980s. Our estimate does not reveal any inherent cyclicality in output.

Movements in the domestic price level, which, on average, explained only a fraction of the short-run variability of output, may have spurred growth somewhat, mainly on three occasions during our sample period. Two instances occur during the 1960s, and the third longer, but less forceful, occurs in the years following the first oil price shock. According to the estimate, the burst of inflation in the mid-1970s may have served to counteract the economic slowdown. Developments in the price level may have contributed to the curbing of output growth in 1969 and in 1970. The contribution of price level movements to output growth ceased during the 1980s.

The stock of narrow money was estimated to have caused, on average, eight per cent of the variability of output. Figure 7.3 shows that the influence of money on output has varied considerably with both substantial positive and markedly negative contributions. Thus the upturn during the late 1960s and early 1970s may be attributable to monetary developments. Likewise, monetary factors seem to have contributed to the rebound from the metal and engineering strike in the early 1970s and to output growth during the last two years of our sample. Negative contributions to growth occur e.g. during the mid-1960s, in 1977 - 1978 and in 1980 - 1981.

The foreign interest rate, which according to our estimate also caused, on average, eight per cent of the movements in domestic output, was of negligible importance during the 1960s. It may have contributed to the slowdown in 1970, and it was most likely one factor curbing output developments following the first oil price shock. On the whole, foreign interest rate movements have contributed more to output fluctuations after the breakdown of the Bretton Woods than before. Foreign price level movements, on the other hand, are estimated to have contributed to domestic output variability mainly during the 1970s. The influence of the exchange rate is mainly limited to the years following devaluations. Such periods stand out following the devaluations in 1967, 1977, 1978 and 1982. During these times, the exchange rate has contributed considerably (positively) to output movements. The small exchange rate adjustment in 1986 induced no growth, however.

Before concluding the analysis of short-run fluctuations, it may be interesting to contrast the findings for Finland with those of other small open economies. Keeping within the realm of VAR analyses, this amounts to a comparison with Burbidge & Harrison (1985) and Kuszczak

& Murray (1987) for Canada and Genberg et al. (1987) and Kugler (1989) for Switzerland. Using index models, Krieger (1989) also presents related results for Canada. The overall impression is that the Finnish economy has been the most closed and the Swiss economy the least closed. There are considerable differences in the estimated degree of openness for these three small open economies. While foreign factors, on average, have accounted for one fifth of the variability of Finnish output at the two-year horizon, the corresponding figures are in excess of one half for Canada and three quarters for Switzerland. The corresponding values for the domestic price level are one sixth, one half and two thirds.

Moreover, while the Finnish economy appears to have become more open over time, the opposite emerges in the case of the Canadian economy. Movements in Swiss output have likewise become less determined by foreign influences whereas Swiss prices have become more closely tied to foreign developments. While a direct comparison of the results is not altogether straightforward, and the analyses for Canada and Switzerland can be criticized (see section 2.1 of chapter 2), the comparison tends to reinforce the impression that the Finnish economy may not have been as open historically as perhaps thought previously. The comparison also raises a question about the empirical content of the label "small open".

7.3 Conclusions

The aim of this chapter has been to provide empirical evidence on the short-run effects of shocks hitting the Finnish economy. This was carried out using a VAR model incorporating low frequency restrictions. The model derives from the theoretical and empirical analyses of the previous chapter. The identifying restrictions derive from the empirical evidence presented in the previous chapter - not from ad hoc or theoretical considerations. What is gained by basing the restrictions on empirical evidence is that the credibility of the identifying assumptions - which is typically subject to doubt - can be, and in the current study has been, increased through demonstrations of the robustness of the empirical evidence from which the restrictions derive.

The results can be summarized in the following way. The effects of shocks tend to peak within a one-year horizon, and impacts dampen out within two years after a shock. The response of the domestic price level to a shock in foreign prices is very slow, however. The direction of impact is, in the majority of cases, in accordance with what conventional wisdom leads one to expect. Money and prices are found to be nonneutral in the short run. While these results are robust up to an unknown estimation uncertainty, they indicate that changes in short-run dynamics have taken place within our sample period. In general, the responses of the economy to unanticipated events have become more rapid and pronounced. In addition, the turbulent period 1973 - 1975 has initiated structural changes which contributed to temporarily high vulnerability to shocks during the latter parts of the 1970s.

Finnish output emerges as decreasingly exogenous over time, with roughly one fifth of the short-run variability attributable to inherent output dynamics, two fifths to other domestic factors and the reminder to foreign factors during the 1980s. In particular, foreign output, the foreign interest rate, the price of oil and the fiscal policy stance have begun to contribute nonnegligibly to short-run output variability since the breakdown of Bretton Woods. Furthermore, money and domestic interest rate developments emerge as quantitatively important in the 1980s. Hence, both monetary and fiscal policy have contributed to short-run real variability. Movements in the exchange rate made important contributions to real variability during fairly short periods following devaluations. Caution in interpreting the empirical evidence is, however, mandatory because of the possibly large estimation uncertainty. Nevertheless, our study challenges conventional macroeconometric models of the Finnish economy through the documentation of nonconstant and fairly rapid short-run dynamics.

8 SUMMARY AND CONCLUDING REMARKS

In this study we set out to empirically assess the main sources and characteristics of aggregate fluctuations in the Finnish economy. This entails generating stylized facts of the effects over time of different kinds of foreign and domestic shocks using historical data on key macroeconomic variables. Seasonally unadjusted monthly data from the period 1960 - 1988 were employed. Empirical regularities were documented using structural vector autoregressions taking into account multivariate cointegration. This data-oriented approach imposes as few prior beliefs as possible, and it puts emphasis on shocks and dynamics. The approach yields shocks that are structural in the sense of being pertinent to a certain variable or to a meaningful function of variables.

The study was set up in the following way. After some introductory words, a presentation of the topic and of earlier research was given. We then introduced the macroeconometric framework utilized throughout the study. With regard to the empirical study, the analysis was divided into four parts. In the first part we studied the joint integration properties of our data. We assessed univariate time series properties through tests for unit roots, and multivariate aspects through maximum likelihood methods. The analysis served to provide information about how the data should be stationarized, how seasonality should be addressed, what distribution theory was appropriate when testing in the subsequent models, and about long-run equilibria affecting the specification of the subsequent models. The joint integration properties of key foreign and Finnish macroeconomic variables may also be interesting in their own right.

The second empirical part of the study focused on the variable of ultimate interest, i.e. output. We employed a Wold causal chain model to analyze international output dynamics comprising Finnish output and the output of nine other OECD countries. The remaining two empirical parts focused on the sources and nature of Finnish real and nominal economic variability both in the long run and the short run. While evidence was presented to support the view that the same shocks are responsible for fluctuations at both ends of the spectrum, we devoted separate chapters to the study of long- and short-run variability in order to let the data speak as freely as possible about both low and high frequency movements. Sensitivity analyses with respect to, inter alia, the choice and operationalization of variables and the estimation period were made throughout the empirical chapters.

The empirical results generated in the current study are numerous, to say the least, and quite detailed. The multitude of findings serves to give a comprehensive picture of the topic and aids in checking the sensitivity and consistency of our results. On the other hand, the wealth of details hampers the summarizing of our results. Hence, the reader interested in detailed findings is encouraged to consult the summaries concluding each of the four empirical chapters. In the current exposition, only the main findings will be recapitulated. We now turn to these main results.

With respect to the joint integration properties of our data, the following tentative conclusions were reached. In a fairly comprehensive vector of key foreign and domestic macroeconomic variables, all but one variable is integrated of order one at the zero frequency. Thus e.g. output, prices, money, interest rates and the exchange rate belong to the difference stationary class of time series. The exception is the real rate of interest, which does not display unit root nonstationarity. Seasonality was found to be essentially a deterministic phenomenon. Trivariate cointegration among the output of the U.S., two alternative aggregates of middle-sized economies, and the Finnish economy was detected. Similarly, Finnish output and key nominal foreign and domestic variables have appeared in two long-run equilibrium relationships. Our analysis constitutes an empirical example of the superiority of the maximum likelihood approach to testing for cointegration in comparison with the two-step approach.

The analysis of international output dynamics yielded the following conclusions. U.S. economic activity essentially calls the tune, middle-sized economies influence mainly each other and small economies, and the Finnish economy reacts to, but does not influence, other economies. Finnish output is contemporaneously influenced by foreign output. The effects of output movements in one country on output in another country in general peak during the third or fourth quarter following a shock, and most contributions to output variability have vanished two years after a shock. In general, foreign output shocks have permanent effects on the level of domestic output. Finland has, in terms of output developments, been relatively closed, historically.

In our analysis of the long-run impact of real and nominal shocks the following conclusions were reached. No sizeable long-run impact of nominal shocks, including economic policy, on Finnish real output was found. Half of an unforeseen movement in domestic output persists into the infinite future, however. Thus the common practice of measuring the size of a boom or a recession as the deviation from a deterministic trend overstates the relevant deviation. It is argued that the stochastic part of the trend in output arises mainly from supply side shocks. Domestic long-run output developments are only weakly related to foreign growth impulses, but foreign price developments are eventually crucial for domestic prices. The exceptionally large shocks of 1973 - 1975 have had permanent effects in the sense that they triggered structural changes that took years to complete and which altered the short-run dynamics of the economy. However, average movements in the price of oil have contributed to long-run movements in the Finnish economy to a lesser extent than in many other economies. It is argued that this is because of the bilateral trade agreement between Finland and the Soviet Union.

In the analysis of the short-run impact of shocks the following conclusions were reached. The response of the Finnish economy to shocks has typically not been as slow as indicated by previous estimates. The response tends to peak within a one-year horizon, and the impact settles down fairly well within two years after the shock. The effect on the domestic price level of a shock in foreign prices is slow, however. Money and prices are nonneutral in the short run, but their quantitative impact on the real variable has been small, historically. By contrast, both monetary and fiscal policy have attributed to short-run variability, at least during the 1980s.

Changes in Finnish short-run dynamics have most likely taken place during our sample period. Domestic output has become increasingly susceptible to shocks, with roughly one fifth of the short-run variability attributable to inherent output dynamics, two fifths to other domestic factors and the reminder to foreign factors in the 1980s. More specifically, foreign output, the foreign interest rate, the price of oil and some manifestations of fiscal policy have begun to contribute nonnegligibly to output variability since the end of the Bretton Woods era. Money and the domestic interest rate influenced output in the 1980s. Domestic prices have also become more endogenous over time, mainly reacting to shocks to foreign prices and the foreign interest rate. Our findings caution against the common practice of basing inference about current short-run dynamics on data from the 1960s and the early 1970s.

The shocks driving the economic system are zero mean entities that are temporally independent, but they are nonnormally distributed. Most shocks are small, but occasionally large shocks predominantly of one sign have occurred, and such large shocks of various kinds have typically hit the economy in clusters during periods of two to three years of duration. The years 1973 - 1975 stand out as particularly turbulent, while the 1980s emerge as remarkably calm. Short-run variability in the Finnish economy is the result of the occurrence of different kinds of shocks, both foreign and domestic, supply and demand, and real and nominal.

Lastly, we wish to state some general cautionary words about the reliability of the results. Our results are empirical, and there is always room for a healthy dose of scepticism about empirical evidence - our study can be no exception to this. The small number of variables entering our models is a drawback from a conventional macroeconometric point of view. Likewise, the absence of assessments of sampling uncertainty implies a caveat to our results. Last but not least, out statements about causation are based on correlation, but correlation is merely a necessary, not a sufficient, condition for causation.

LIST OF SYMBOLS

In some cases, the same symbol has been used in different contexts in different meanings, but the meaning of a symbol has been spelled out in each context so as to avoid confusion. Conventional mathematical symbols are not included in the list of symbols. In the main text, a tilde under a symbol denotes a vector. A bold-face letter denotes a matrix. A dot above a variable denotes its time derivative. A bar above a variable denotes a fixed value. The superscript e denotes the expectation of a variable. Symbols of data variables are listed in Appendix 1.

a	unit root test regression coefficient
b	unit root test regression coefficient
d	number of deterministic variables
f	indicator index
h	unit root test regression coefficient
i	indicator index
j	indicator index
k	number of steps ahead forecasted
l	length of autoregression
m	indicator index
n	number of variables
p	order of correction for serial dependence in unit root tests for a nonseasonal time series
q	order of correction for serial dependence in unit root tests for a seasonal time series
r	number of cointegration relationships
s	indicator index
t	time index
u	exogenous variable
v	variable in unit root test regression
w	endogenous variable
$\tilde{w}$	exogenous variable
$\tilde{\tilde{w}}$	exogenous, nonlagged variable
x	variable in unit root test regression, endogenous variable
y	generic symbol for variable
z	equilibrium error

A	attractor set
B	matrix
C	matrix
D	matrix
E	matrix, intermediate economy
F	matrix, Finland
G	matrix
H	matrix, Heaviside unit step function
I	information set
N	number of shocks
P	matrix
S	matrix
T	number of observations
U	U.S.A.
Z	matrix
α	cointegration coefficient, parameter
γ	parameter
δ	Dirac delta function
ε	generic symbol for shock
ζ	VAR residual
η	orthogonalized and scaled VAR residual
ι	indicator variable
κ	indicator variable
λ	relative forecast error variance
μ	eigenvalue
ν	unit root test regression residual
ξ	structural shock
o	deviation form variable
σ	eigenvector, or element of eigenvector
τ	indicator variable
υ	indicator variable
ϕ	matrix in Jordan decomposition of Φ
ω	weight
Γ	matrix of VAR coefficients
Λ	matrix in Jordan decomposition of Φ
Ξ	matrix allocating shocks to variables

Π	long-run coefficient matrix
Σ	spectral density matrix
Υ	covariance matrix of forecast errors
Φ	matrix representing propagation mechanism
Ω	covariance matrix of shocks

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APPENDIX 1

Data Sources and Construction of Variables

The main data sources are the OECD (Main Economic Indicators) (OECD), the Bank of Finland (BF) and the Central Statistical Office of Finland (CSOF). The data consist of seasonally unadjusted monthly observations from 1960M1 - 1988M8. Some variables cover a shorter period. The data are available from the author upon request.

U.S. total industrial production y^{*U}

Volume index of U.S. total industrial production, 1980 = 100, OECD.

Intermediate economy 1: total industrial production y^{*E}

Volume index of weighted average of the total industrial production of Belgium, France, Germany, Italy, Japan, the Netherlands, Sweden and the U.K., 1980 = 100. Weighting is by annual real GDP in USD, OECD.

Intermediate economy 2: total industrial production y^{*E*}

Volume index of weighted average of the total industrial production of Belgium, France, Germany, Italy, Japan, the Netherlands, Sweden and the U.K., 1980 = 100. Weighting is by monthly Bank of Finland currency index weights, BF, OECD.

Total industrial production y

Volume index of Finnish total industrial production, 1980 = 100, BF.

Consumer prices p

Index 1981 = 100, CSOF.

M1 m

Billion FIM at current prices, BF.

Credit advanced to the public c

Billion FIM at current prices, BF.

Foreign short-term interest rate i*

International weighted interest rate 1960M1 - 1981M12 in Belgium, Germany, the Netherlands, Switzerland, the U.K. and the U.S., OECD.
International weighted interest rate 1982M1 - 1986M10 of eight currencies, BF.
International weighted interest rate 1986M11 - 1988M8 of twelve currencies, BF.
Weighting is by monthly Bank of Finland currency index weights, BF.
The interest rates have been linked to the international weighted interest rate of twelve currencies to adjust for changes in levels. All interest rates refer to three-month paper. The interest rates are nominal per annum end-of-month observations expressed as percentage points.

Real foreign short-term interest rate $i^* - \dot{p}^*$

Per annum end-of-month observations expressed as percentage points using i* and p* and a three-month inflation rate, BF, OECD.

Exchange rate e

FIM/USD exchange rate 1960M1 - 1970M12, BF.
Bank of Finland currency index 1971M1 - 1988M8, 1982 = 100, BF.
The FIM/USD exchange rate has been linked to the Bank of Finland currency index.

Real exchange rate e^r

Index employing p, e and consumer prices from France, Germany, the Netherlands, Sweden, the U.K. and the U.S., 1982 = 100. Weighting is by monthly Bank of Finland currency index weights, BF.

Foreign total industrial production y*

Volume of weighted average of the total industrial production of Belgium, France, Germany, Italy, Japan, the Netherlands, Sweden, the U.K. and the U.S., index 1980 = 100. Weighting is by monthly Bank of Finland currency index weights, BF, OECD.

Foreign consumer prices p*

Index of weighted average of consumer prices in Belgium, Denmark, France, Germany, Italy, Japan, the Netherlands, Sweden, Switzerland, the U.K. and the U.S., 1980 = 100. Weighting is by monthly Bank of Finland currency index weights, BF, OECD.

Fiscal policy stance g

Central government net borrowing requirement deflated by p 1968M1 - 1988M8, billion FIM, Finnish Ministry of Finance, CSOF.

Oil price p^{01}

Index of imported crude oil at current prices 1969M1 - 1988M8, 1980 = 100, BF.

Real oil price p^{01r}

Index employing p^{01} and p^* 1969M1 - 1988M8, 1980 = 100, BF, OECD.

Import price of mineral fuels p^{02}

Index 1949 = 100, BF.

Real import price of mineral fuels p^{02r}

Index employing p^{02} and p^*, 1949 = 100, BF, OECD.

Short-term interest rate i

Marginal interest rate on credit advanced by the Bank of Finland to commercial banks 1960M1 - 1982M2, Saarinen (1986).
Interest rate on new certificates of deposit 1982M3 - 1986M12, BF.
Three-month HELIBOR 1987M1 - 1988M8, BF.
The interest rates are nominal per annum end-of-month observations expressed as percentage points.

Real short-term interest rate $i - \dot{p}$

Per annum end-of-month observations expressed as percentage points using i and p and a three-month inflation rate, Saarinen (1986), BF, CSOF.

Terms of trade p/p*

Index employing p and p*, 1980 = 100, BF, OECD.

Time trend time

Linear time trend: 0.01, 0.02, 0.03, ...

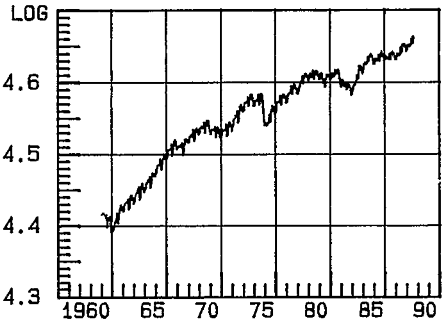
Seasonal dummy variables D_i

12 dummy variables with i:th variable taking the value 1 in the i:th month and the value 0 otherwise.

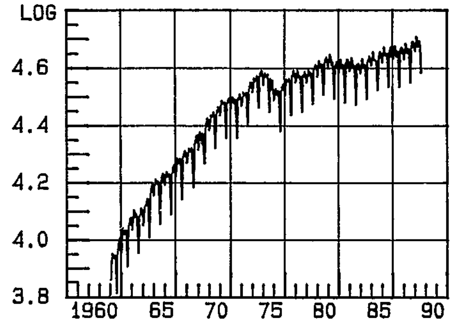
APPENDIX 2

Graphs of Variables

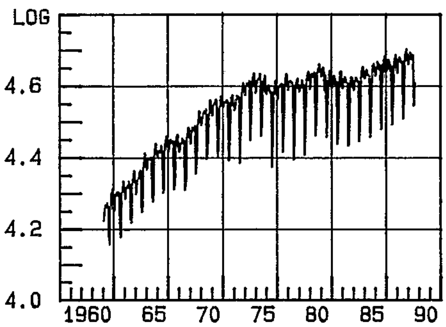
U.S. TOTAL INDUSTRIAL PRODUCTION



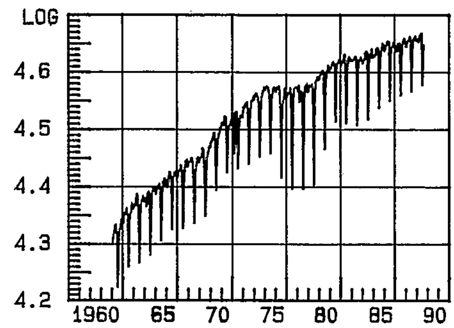
INTERMEDIATE ECONOMY 1:
TOTAL INDUSTRIAL PRODUCTION



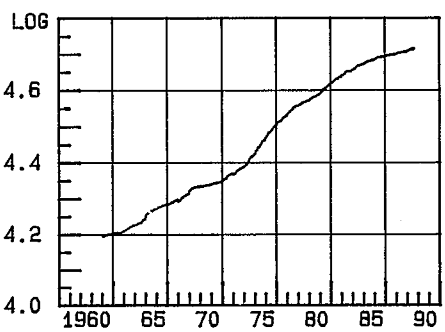
INTERMEDIATE ECONOMY 2:
TOTAL INDUSTRIAL PRODUCTION



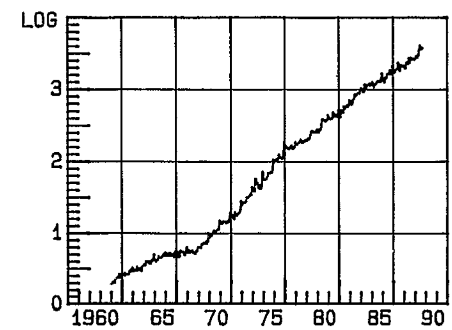
TOTAL INDUSTRIAL PRODUCTION



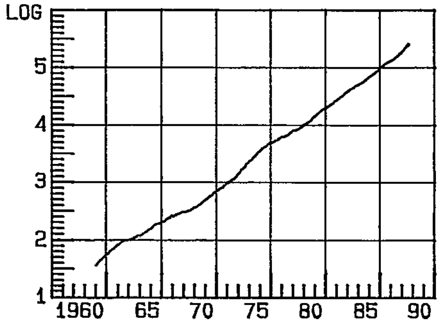
CONSUMER PRICES



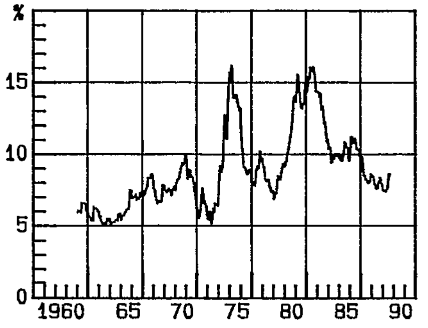
M1



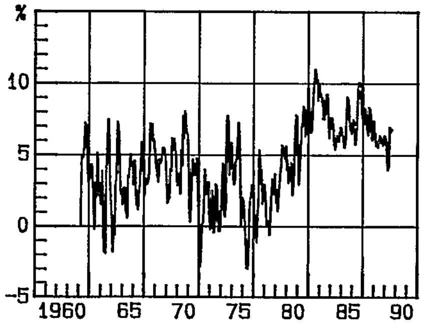
CREDIT ADVANCED TO THE PUBLIC



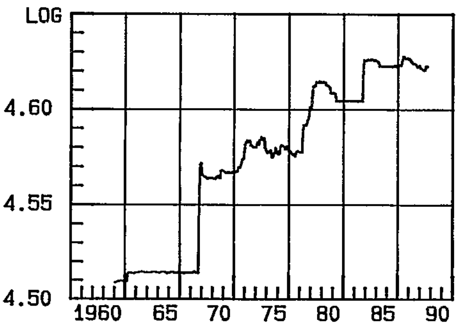
FOREIGN SHORT-TERM INTEREST RATE



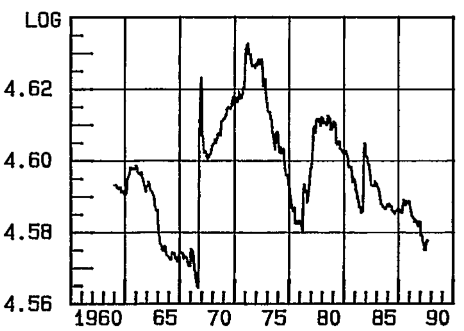
REAL FOREIGN SHORT-TERM
INTEREST RATE



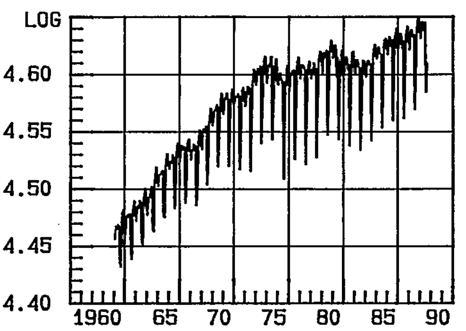
EXCHANGE RATE



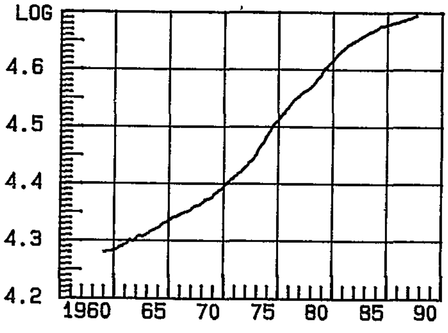
REAL EXCHANGE RATE



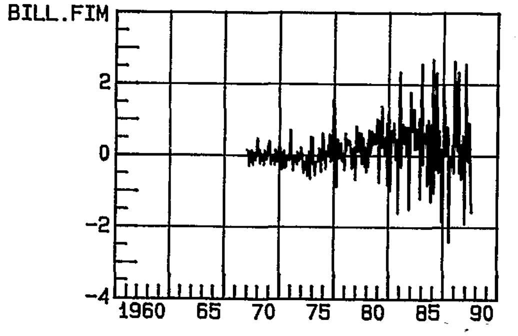
FOREIGN TOTAL INDUSTRIAL PRODUCTION



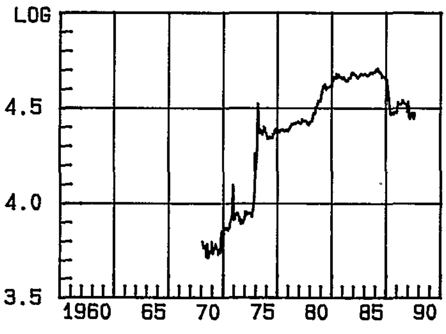
FOREIGN CONSUMER PRICES



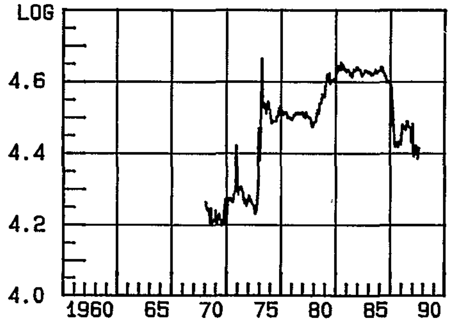
FISCAL POLICY STANCE



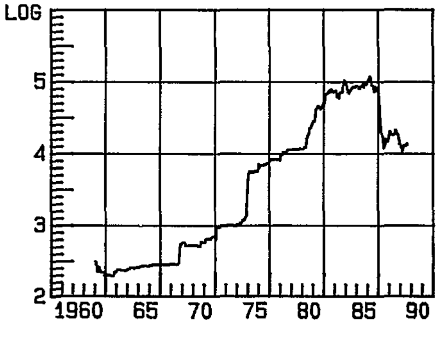
OIL PRICE



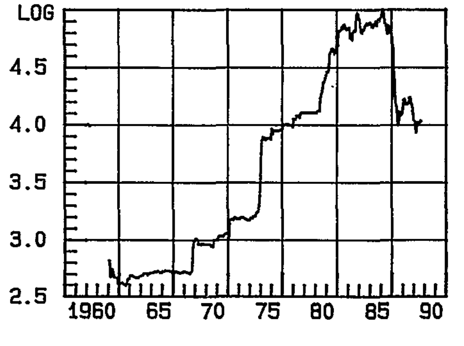
REAL OIL PRICE



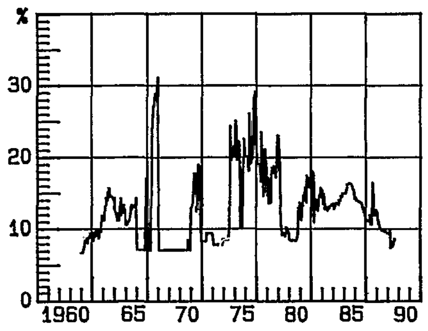
IMPORT PRICE OF MINERAL FUELS



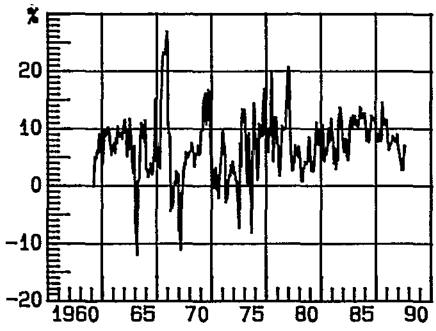
REAL IMPORT PRICE OF MINERAL FUELS



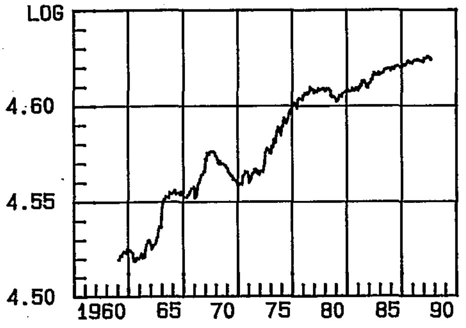
SHORT-TERM INTEREST RATE



REAL SHORT-TERM INTEREST RATE



TERMS OF TRADE



Autocorrelations and Partial Autocorrelations of Differenced Variables

TABLE A3.1

Autocorrelations of differenced variables

Lag	Variable																				
	Δy^{*U}	Δy^{*E}	Δy^{*E*}	Δy	Δp	Δm	Δc	Δi^*	$\Delta(i^*-\dot{p}^*)$	Δe	Δe^r	Δy^*	Δp^*	Δg	Δp^{01}	Δp^{01r}	Δp^{02}	Δp^{02r}	Δi	$\Delta(i-\dot{p})$	$\Delta(p/p^*)$
1	-0.106	-0.084	-0.206	-0.332	0.215	-0.339	0.311	0.234	0.232	0.292	0.237	-0.133	0.370	-0.522	-0.133	-0.193	0.381	0.339	-0.029	-0.029	0.036
2	-0.217	-0.324	-0.274	-0.134	0.280	-0.106	0.216	-0.004	-0.005	-0.030	0.005	-0.277	0.380	-0.006	-0.096	-0.146	0.183	0.145	-0.089	-0.088	0.153
3	-0.125	-0.112	-0.031	-0.052	0.218	0.074	0.176	0.055	0.050	-0.093	-0.238	-0.105	0.380	-0.023	0.449	0.416	0.165	0.117	-0.085	-0.084	-0.048
4	0.059	-0.143	-0.073	-0.017	0.212	-0.095	0.211	0.167	0.164	-0.007	-0.032	-0.148	0.284	0.138	-0.063	-0.058	0.136	0.105	-0.031	-0.032	0.045
5	0.075	0.071	0.049	0.023	0.203	-0.107	0.236	0.137	0.139	0.006	-0.017	0.063	0.298	0.010	-0.055	-0.084	0.051	0.024	-0.029	-0.029	0.055
6	0.044	0.195	0.108	0.027	0.245	0.308	0.310	0.062	0.059	0.004	-0.019	0.210	0.320	-0.183	0.034	0.024	0.017	-0.024	0.025	0.025	0.046
7	0.066	0.070	0.058	0.017	0.183	-0.159	0.197	0.000	-0.002	0.020	-0.008	0.065	0.318	0.004	0.010	0.023	-0.029	-0.085	-0.049	-0.049	-0.001
8	0.034	-0.144	-0.075	-0.015	0.161	0.013	0.083	0.031	0.027	-0.005	-0.015	-0.149	0.352	0.087	-0.062	-0.095	0.048	0.014	-0.128	-0.129	0.074
9	-0.149	-0.113	-0.024	-0.046	0.249	0.016	0.055	0.053	0.050	-0.030	0.049	-0.107	0.295	0.137	0.014	0.006	0.053	0.006	0.022	0.021	0.062
10	-0.245	-0.331	-0.276	-0.128	0.141	-0.092	0.089	-0.034	-0.033	0.011	0.054	-0.285	0.386	-0.158	0.000	-0.015	0.025	-0.009	-0.108	-0.108	0.028
11	-0.148	-0.068	-0.220	-0.330	0.214	-0.170	0.153	0.107	0.106	0.011	0.061	-0.120	0.348	-0.293	-0.012	-0.035	0.014	-0.031	-0.007	-0.008	0.037
12	0.856	0.991	0.981	0.978	0.228	0.539	0.422	0.087	0.094	-0.015	0.016	0.994	0.493	0.564	-0.008	-0.007	0.107	0.060	0.154	0.156	0.058
13	-0.165	-0.091	-0.203	-0.325	0.157	-0.185	0.145	-0.028	-0.025	-0.009	0.006	-0.140	0.292	-0.146	-0.035	-0.040	-0.039	-0.073	0.039	0.040	0.045
14	-0.275	-0.323	-0.277	-0.128	0.138	-0.029	0.040	-0.124	-0.122	-0.014	0.026	-0.277	0.283	-0.213	0.013	-0.013	-0.001	-0.029	-0.102	-0.101	0.015
15	-0.155	-0.111	-0.033	-0.054	0.137	0.013	0.011	-0.154	-0.156	-0.028	0.002	-0.106	0.297	0.113	-0.030	-0.015	-0.035	-0.050	-0.012	-0.012	-0.114
16	0.023	-0.139	-0.072	-0.019	0.112	-0.077	0.022	-0.025	-0.028	-0.034	0.002	-0.143	0.224	-0.017	-0.028	-0.032	-0.037	-0.049	-0.082	-0.083	0.004
17	0.070	0.063	0.043	0.023	0.087	-0.082	0.041	0.038	0.038	-0.020	-0.003	0.056	0.234	0.195	0.024	0.016	0.032	0.029	-0.004	-0.005	-0.071
18	0.051	0.201	0.108	0.028	0.184	0.261	0.044	-0.067	-0.070	-0.014	-0.011	0.215	0.273	-0.316	0.011	0.023	-0.004	-0.019	0.003	0.002	0.150
19	0.038	0.068	0.062	0.020	0.051	-0.116	0.046	-0.115	-0.115	-0.005	-0.011	0.063	0.188	0.101	0.020	0.021	0.038	0.052	-0.003	-0.003	-0.042
20	0.021	-0.150	-0.084	-0.014	0.092	-0.024	-0.074	-0.096	-0.097	-0.027	0.017	-0.155	0.150	-0.007	0.033	0.021	0.078	0.086	0.007	0.006	0.018
21	-0.144	-0.112	-0.025	-0.053	0.142	-0.017	-0.107	-0.019	-0.017	-0.029	-0.003	-0.106	0.273	0.176	0.051	0.057	0.009	0.002	0.008	0.007	-0.037
22	-0.248	-0.335	-0.271	-0.122	0.109	-0.123	-0.083	0.014	0.018	-0.049	0.006	-0.290	0.220	-0.124	0.030	0.039	0.006	0.002	-0.028	-0.029	-0.011
23	-0.156	-0.057	-0.218	-0.328	0.082	-0.085	0.026	-0.168	-0.166	-0.018	0.034	-0.108	0.229	-0.307	-0.002	-0.014	0.037	0.004	-0.044	-0.042	-0.043
24	0.847	0.992	0.972	0.970	0.196	0.478	0.193	-0.052	-0.047	0.049	0.030	0.998	0.419	0.481	0.026	0.035	0.019	0.014	0.134	0.135	0.048

All autocorrelations are based on data from the time period 1960M1-1988M8, except the autocorrelations for the variables Δg, Δp^{01} and Δp^{01r} which are based on data from the time period 1973M1 - 1988M8. The approximate two standard error critical value is 0.106 for the longer time period and 0.144 for the shorter time period. For notational simplicity we write $1 - L = \Delta$.

TABLE A3.2

Partial autocorrelations of differenced variables

Lag	Δy^{*U}	Δy^{*E}	Δy^{*E*}	Δy	Δp	Δm	Δc	Δi^*	$\Delta(i^*-\dot{p}^*)$	Δe
1	-0.106	-0.084	-0.206	-0.332	0.215	-0.339	0.311	0.234	0.232	0.292
2	-0.231	-0.334	-0.331	-0.275	0.245	-0.249	0.133	-0.062	-0.063	-0.126
3	-0.190	-0.200	-0.204	-0.243	0.134	-0.067	0.086	0.074	0.070	-0.051
4	-0.042	-0.349	-0.284	-0.225	0.105	-0.139	0.130	0.144	0.143	0.038
5	0.006	-0.179	-0.182	-0.181	0.090	-0.230	0.133	0.073	0.076	-0.016
6	0.045	-0.056	-0.079	-0.133	0.134	0.176	0.198	0.025	0.022	0.001
7	0.121	0.004	0.025	-0.088	0.048	-0.019	0.020	-0.026	-0.025	0.024
8	0.119	-0.123	-0.019	-0.078	0.012	0.033	-0.078	0.012	0.008	-0.021
9	-0.073	-0.115	0.029	-0.119	0.132	-0.017	-0.057	0.011	0.010	-0.023
10	-0.267	-0.573	-0.352	-0.308	-0.004	-0.071	-0.008	-0.073	-0.071	0.034
11	-0.368	-0.909	-0.711	-0.961	0.073	-0.264	0.061	0.142	0.141	-0.009
12	0.806	0.891	0.940	0.547	0.101	0.393	0.371	0.023	0.034	-0.021
13	-0.338	0.481	-0.291	-0.044	-0.002	0.170	-0.071	-0.056	-0.056	0.009
14	-0.220	-0.082	-0.107	0.187	-0.021	0.126	-0.099	-0.111	-0.109	-0.018
15	-0.068	-0.290	0.107	0.176	-0.015	0.006	-0.082	-0.151	-0.153	-0.025
16	-0.140	0.058	-0.085	0.204	-0.016	0.027	-0.103	0.002	-0.005	-0.019
17	0.068	-0.446	0.065	0.101	-0.046	-0.013	-0.076	0.033	0.031	-0.011
18	0.086	-0.185	-0.039	0.154	0.073	0.044	-0.117	-0.036	-0.041	-0.014
19	-0.146	0.273	-0.220	0.353	-0.059	0.056	0.040	-0.018	-0.016	0.000
20	0.007	-0.379	-0.156	0.370	-0.029	-0.065	-0.013	-0.062	-0.060	-0.034
21	0.007	-0.435	-0.131	0.106	0.062	-0.089	-0.029	0.021	0.021	-0.016
22	-0.121	1.19	0.229	0.077	0.024	-0.176	-0.021	0.016	0.020	-0.041
23	-0.026	1.91	0.352	-0.221	-0.028	-0.124	0.063	-0.179	-0.181	0.002
24	0.386	-1.61	0.014	-0.007	0.114	0.206	0.127	0.080	0.082	0.050

Lag	Δe^r	Δy^*	Δp^*	Δg	Δp^{01}	Δp^{01r}	Δp^{02}	Δp^{02r}	Δi	$\Delta(i-\dot{p})$	$\Delta(p/p^*)$
1	0.237	-0.133	0.370	-0.522	-0.133	-0.193	0.381	0.339	-0.029	-0.029	0.036
2	-0.054	-0.300	0.282	-0.384	-0.115	-0.190	0.044	0.034	-0.089	-0.089	0.152
3	-0.241	-0.216	0.220	-0.368	0.433	0.372	0.096	0.066	-0.091	-0.090	-0.060
4	0.092	-0.346	0.054	-0.146	0.043	0.080	0.045	0.048	-0.046	-0.047	0.026
5	-0.031	-0.213	0.085	0.077	0.009	0.021	-0.037	-0.039	-0.049	-0.049	0.071
6	-0.081	-0.038	0.121	-0.089	-0.213	-0.180	-0.015	-0.040	0.007	0.007	0.029
7	0.036	0.010	0.114	-0.265	0.000	-0.022	-0.053	-0.083	-0.063	-0.063	-0.019
8	-0.028	-0.115	0.134	-0.296	-0.055	-0.091	0.083	0.076	-0.142	-0.143	0.072
9	0.038	-0.121	0.028	0.008	0.109	0.058	0.024	-0.004	-0.003	-0.004	0.062
10	0.045	-0.497	0.164	0.129	-0.019	-0.045	-0.001	0.000	-0.151	-0.152	-0.006
11	0.027	-0.909	0.087	-0.518	0.048	0.026	-0.003	-0.025	-0.053	-0.053	0.024
12	0.016	0.940	0.297	-0.070	-0.074	-0.047	0.101	0.079	0.114	0.116	0.060
13	0.020	0.520	-0.076	0.317	-0.034	-0.029	-0.141	-0.135	0.011	0.011	0.023
14	0.044	-0.559	-0.050	-0.031	-0.028	-0.050	0.047	0.027	-0.096	-0.095	-0.014
15	-0.009	-1.39	-0.012	0.099	0.007	-0.010	-0.058	-0.038	-0.021	-0.022	-0.127
16	0.012	0.212	-0.036	-0.021	-0.005	-0.031	-0.008	-0.022	-0.116	-0.116	0.007
17	0.020	-1.93	-0.030	0.140	0.033	0.034	0.079	0.082	-0.040	-0.041	-0.057
18	-0.022	0.125	-0.009	-0.049	0.031	0.035	-0.042	-0.053	-0.072	-0.073	0.132
19	-0.002	0.806	-0.106	0.024	0.052	0.069	0.087	0.107	-0.036	-0.036	-0.035
20	0.027	-0.893	-0.190	0.076	0.021	0.016	0.017	0.015	0.010	0.008	-0.024
21	-0.030	-1.90	0.117	-0.096	0.042	0.049	-0.051	-0.050	-0.033	-0.035	-0.012
22	0.008	-3.95	-0.051	-0.000	0.011	0.022	0.006	-0.003	-0.043	-0.044	-0.023
23	0.045	2.51	0.006	-0.178	-0.020	-0.010	0.026	-0.008	-0.065	-0.064	-0.037
24	-0.004	1.08	0.234	-0.039	-0.019	-0.010	-0.017	0.010	0.049	0.049	0.061

All partial autocorrelations are based on data from the time period 1960M1 - 1988M8, except the partial autocorrelations for the variables Δg, Δp^{01} and Δp^{01r} which are based on data from the time period 1973M1 - 1988M8. The approximate two standard error critical value is 0.106 for the longer time period and 0.144 for the shorter time period. For notational simplicity we write $1 - L = \Delta$.

TABLE A4.1

Tests for the order of integration of nonseasonal variables using data from the period 1960M1 - 1972M12

	Data in levels				Data in first differences			
	t	F1	F2	F3	t	F1	F2	F3
i^*	0.999				-0.042			
	(-0.32)				(-5.39)***			
	0.957	1.63			-0.042	14.4***		
	(-1.81)				(-5.37)***			
	0.945		1.29	1.93	-0.051		9.61***	14.4***
	(-1.94)				(-5.36)***			
$i^*-\dot{p}^*$	0.943				-1.147			
	(-1.74)*				(-8.39)***			
	0.746	5.28**			-1.158	35.2***		
	(-3.21)**				(-8.40)***			
	0.739		3.61	5.28	-1.158		23.4***	35.0***
	(-3.25)*				(-8.37)***			
e	1.000				0.189			
	(0.90)				(-4.49)***			
	0.995	1.06			0.111	11.1***		
	(-0.47)				(-4.71)***			
	0.948		2.28	2.44	0.096		7.45***	11.2***
	(-2.14)				(-4.73)***			
e^r	1.000				-0.221			
	(0.60)				(-3.82)***			
	0.999	0.38			-0.230	7.68***		
	(-0.08)				(-3.85)***			
	0.975		1.42	1.75	-0.360		5.74**	8.32**
	(-1.16)				(-4.08)***			

	Data in levels				Data in first differences			
	t	F1	F2	F3	t	F1	F2	F3
p^{02}	1.000				-0.012			
	(1.10)				(-4.32)***			
	0.995	1.38			-0.186	11.5***		
	(-0.34)				(-4.79)***			
	0.880		3.82	4.34	-0.205		7.68***	11.5***
	(-2.89)				(-4.73)***			
p^{02r}	1.000				-0.158			
	(-0.70)				(-4.37)***			
	0.971	0.91			-0.192	9.76***		
	(-1.06)				(-4.42)***			
	0.902		2.29	3.08	-0.210		6.83***	10.2***
	(-2.39)				(-4.47)***			
p/p^*	1.000				0.336			
	(0.80)				(-4.00)***			
	0.989	1.70			0.260	8.82***		
	(-1.41)				(-4.20)***			
	0.979		1.27	1.20	0.224		6.20**	9.29***
	(-1.25)				(-4.31)***			

The estimated first-order autoregressive parameter a_1 (see equations (N1) - (N3) of Table 4.1) is given in the column denoted t and the t value of this estimate below in parentheses. The null hypotheses are: t; $a_1=1$, F1; $b_0=0$, $b_1=1$, F2; $b_0=b_1=0$, $a_1=1$ and F3; $a_1=1$, $b_1=0$ (see equations (N1) - (N3) of Table 4.1). The Dickey-Fuller tests are based on regressions with six lags of the differenced variable under investigation. Dummy variables taking the value 1 for the following variables (dates) and 0 otherwise have been employed, where appropriate, in the test regressions: e^r (1967M11 - 1967M12) and p^{02}, p^{02r} (1971M11, 1974M8). Critical values for the tests are tabulated by Fuller (1976) and Dickey & Fuller (1981). Rejection at the 10,5 and 1 per cent level of significance is indicated by *, ** and ***, respectively.

TABLE A4.2

Tests for the order of integration of nonseasonal variables using data from the period 1973M1 - 1988M8

	Data in levels				Data in first differences			
	t	F1	F2	F3	t	F1	F2	F3
i^*	0.997				0.367			
	(-0.92)				(-4.57)***			
	0.962	3.14			0.364	10.5***		
	(-2.48)				(-4.57)***			
	0.961		2.11	3.09	0.372		7.00***	10.5***
	(-2.48)				(-4.45)***			
$i^*-\dot{p}^*$	0.993				-0.636			
	(-0.55)				(-7.14)***			
	0.956	1.17			-0.638	25.3***		
	(-1.52)				(-7.12)***			
	0.917		1.47	2.19	-0.637		16.8***	25.2***
	(-2.09)				(-7.07)***			
e	1.000				0.447			
	(0.60)				(-3.65)***			
	0.991	1.04			0.418	6.99***		
	(-1.12)				(-3.74)***			
	0.956		2.12	2.76	0.415		4.65*	6.97**
	(-2.34)				(-3.74)**			
e^r	1.000				0.341			
	(-0.80)				(-4.03)***			
	0.965	3.66			0.296	8.57***		
	(-2.43)				(-4.13)***			
	0.947		3.09	3.90	0.292		5.70**	8.52***
	(-2.73)				(-4.12)***			
p^{01}	1.000				0.190			
	(0.84)				(-4.71)***			
	0.948	7.53***			0.154	11.6***		
	(-3.73)***				(-4.82)***			
	0.948		4.99**	6.96**	-0.058		10.3***	15.4***
	(-2.76)				(-5.54)***			

	Data in levels				Data in first differences			
	t	F1	F2	F3	t	F1	F2	F3
p^{01r}	1.000				0.020			
	(0.32)				(-4.95)***			
	0.930	5.15**			0.015	12.3***		
	(-3.18)**				(-4.96)***			
	0.939		4.07	6.03*	-0.176		10.3***	15.5***
	(-2.63)				(-5.56)***			
p^{02}	1.000				0.510			
	(0.80)				(-4.29)***			
	0.978	4.40*			0.496	9.47***		
	(-2.88)*				(-4.35)***			
	0.983		3.20	4.51	0.338		8.67***	13.0***
	(-1.82)				(-5.10)***			
p^{02r}	1.000				0.416			
	(-0.20)				(-4.80)***			
	0.984	0.95			0.414	11.5***		
	(-1.37)				(-4.79)***			
	0.978		2.93	4.38	0.293		9.64***	14.4***
	(-1.93)				(-5.37)***			
p/p^*	1.000				0.288			
	(1.20)				(-4.53)***			
	0.976	9.16***			0.004	13.9***		
	(-3.21)**				(-5.23)***			
	0.946		7.40***	7.06**	-0.180		10.5***	15.6***
	(-3.10)				(-5.52)***			

The estimated first-order autoregressive parameter a_1 (see equations (N1) - (N3) of Table 4.1) is given in the column denoted t and the t value of this estimate below in parentheses. The null hypotheses are: t; $a_1=1$, F1; $b_0=0$, $b_1=1$, F2; $b_0=b_1=0$, $a_1=1$ and F3; $a_1=1$, $b_1=0$ (see equations (N1) - (N3) of Table 4.1). The Dickey-Fuller tests are based on regressions with six lags of the differenced variable under investigation. Dummy variables taking the value 1 for the following variables (dates) and 0 otherwise have been employed, where appropriate, in the test regressions: e^r (1967M11 - 1967M12) and p^{02}, p^{02r} (1971M11, 1974M8). Critical values for the tests are tabulated by Fuller (1976) and Dickey & Fuller (1981). Rejection at the 10,5 and 1 per cent level of significance is indicated by *, ** and ***, respectively.

TABLE A4.3

Tests for the order of integration of nonseasonal variables using data from the period 1980M1 - 1988M8

	Data in levels				Data in first differences					Data in levels				Data in first differences			
	t	F1	F2	F3	t	F1	F2	F3		t	F1	F2	F3	t	F1	F2	F3
i*	0.996				0.285				p^{01r}	1.000				0.293			
	(-1.10)				(-3.25)***					(-0.88)				(-2.90)***			
	0.984	0.90			0.207	5.64**				0.993	0.45			0.230	4.60*		
	(-1.00)				(-3.33)**					(-0.30)				(-3.03)**			
	0.918		1.70	2.15	0.207		3.74	5.53*		0.920		2.61	3.48	0.110		3.60	5.40
	(-2.06)				(-3.32)*					(-2.23)				(-3.28)*			
$i^*-\dot{p}^*$	0.994				-0.610				p^{02}	1.000				0.394			
	(-0.57)				(-4.66)***					(-1.10)				(-3.89)***			
	0.890	1.29			-0.614	10.9***				0.992	0.42			0.371	7.79***		
	(-1.57)				(-4.65)***					(-0.53)				(-3.95)***			
	0.814		1.97	2.90	-0.659		7.23***	10.8***		0.957		2.19	2.99	0.240		6.32**	9.48***
	(-2.30)				(-4.56)***					(-2.05)				(-4.35)***			
e	1.000				0.203				p^{02r}	1.000				0.389			
	(0.60)				(-3.44)***					(-1.15)				(-3.96)***			
	0.962	1.78			0.170	6.20**				1.000	0.77			0.318	8.80***		
	(-1.72)				(-3.52)***					(-0.03)				(-4.19)***			
	0.937		1.44	1.86	0.132		4.34*	6.51**		0.951		2.62	3.12	0.210		6.87***	10.3***
	(-1.75)				(-3.60)**					(-2.13)				(-4.54)***			
e^r	1.000				-0.057				p/p*	1.000				-0.030			
	(-1.40)				(-4.26)***					(1.40)				(-4.17)***			
	0.952	2.36			-0.210	10.4***				0.979	6.39**			-0.768	16.2***		
	(-1.61)				(-4.56)***					(1.58)				(-5.69)***			
	0.773		4.52*	5.63*	-0.213		6.87***	10.3***		0.828		5.93**	3.56	-0.852		11.3***	16.9***
	(-3.34)*				(-4.54)***					(-2.38)				(-5.81)***			
p^{01}	1.000				0.333												
	(-0.55)				(-2.98)***												
	0.981	0.45			0.309	4.58*											
	(-0.77)				(-3.03)**												
	0.936		2.34	3.35	0.179		3.64	5.46									
	(-2.13)				(-3.30)*												

The estimated first-order autoregressive parameter a_1 (see equations (N1) - (N3) of Table 4.1) is given in the column denoted t and the t value of this estimate below in parentheses. The null hypotheses are: t; $a_1=1$, F1; $b_0=0$, $b_1=1$, F2; $b_0=b_1=0$, $a_1=1$ and F3; $a_1=1$, $b_1=0$ (see equations (N1) - (N3) of Table 4.1). The Dickey-Fuller tests are based on regressions with six lags of the differenced variable under investigation. Dummy variables taking the value 1 for the following variables (dates) and 0 otherwise have been employed, where appropriate, in the test regressions: e^r (1967M11 - 1967M12) and p^{02}, p^{02r} (1971M11, 1974M8). Critical values for the tests are tabulated by Fuller (1976) and Dickey & Fuller (1981). Rejection at the 10,5 and 1 per cent level of significance is indicated by *, ** and ***, respectively.

TABLE A4.4

Tests for the order of integration of seasonal variables using data from the period 1960M1 - 1972M12

	Data in levels			Data in first differences				Data in levels			Data in first differences		
	t	F4	F5	t	F4	F5		t	F4	F5	t	F4	F5
y^{*U}	1.000	1.29		0.935	9.83***		c	1.003	11.8***		0.929	9.33***	
	(1.50)			(-1.22)				(3.19)			(-1.41)		
	0.967		31.6***	-0.078		42.6***		0.998		16.9**	0.249		23.3***
	(-1.58)			(-5.11)***				(-0.06)			(-4.85)***		
y^{*E}	1.002	4.28***		0.959	7.70***		y^*	1.001	8.81***		1.048	7.76***	
	(2.65)			(-1.65)*				(3.10)			(1.82)		
	0.928		16.4**	0.157		21.3***		0.710		6.92	0.731		3.21
	(-1.04)			(-4.38)***				(-2.80)			(-1.80)		
y^{*E*}	1.002	6.24***		1.016	5.61***		p^*	1.000	5.80***		0.955	6.31***	
	(3.90)			(0.63)				(0.70)			(-0.67)		
	0.723		7.18	0.600		4.40		1.042		19.0***	-0.172		24.4***
	(-2.70)			(-2.48)				(1.42)			(-5.59)***		
y	1.002	3.73**		0.975	7.93***		i	0.987	14.4***		0.448	20.7***	
	(3.24)			(-0.63)				(-0.58)			(-3.56)***		
	0.654		14.7*	0.033		19.9***		0.738		19.8***	0.249		22.7***
	(-2.34)			(-3.55)**				(-2.27)			(-3.99)**		
p	1.000	21.7***		0.747	19.8***		$i-\dot{p}$	0.935	13.4***		0.309	23.1***	
	(3.10)			(-2.10)**				(-1.36)			(-3.90)***		
	0.908		30.4***	-0.546		39.1***		0.592		22.6***	-0.104		29.1***
	(-2.32)			(-6.18)***				(-3.24)*			(-4.97)***		
m	1.017	3.54**		0.806	6.39***								
	(2.26)			(-1.63)*									
	0.985		9.65	-0.370		22.1***							
	(-0.38)			(-6.33)***									

The estimated first-order autoregressive parameter a_1 (see equations (S1) - (S2) of Table 4.1) is given in the column denoted t and the t value of this estimate below in parentheses. The null hypotheses are: t; $a_1=1$, F4 and F5; $a_1=a_2=1$, $a_3=0$ (see equations (S1) - (S2) of Table 4.1). The Dickey-Fuller-Hasza tests are based on regressions with twelve lags of the differenced variable under investigation. Dummy variables taking the value 1 for the following variables (dates) and 0 otherwise have been employed, where appropriate, in the test regressions: y^{*E*}, y^* (1980M5) and y (1971M2 - 1971M3). Critical values for the tests are tabulated by Fuller (1976) and Hasza & Fuller (1982). Rejection at the 10, 5 and 1 per cent level of significance is indicated by *, ** and ***, respectively.

TABLE A4.5

Tests for the order of integration of seasonal variables using data from the period 1973M1 - 1988M8

	Data in levels			Data in first differences				Data in levels			Data in first differences		
	t	F4	F5	t	F4	F5		t	F4	F5	t	F4	F5
y^{*U}	1.000	9.19***		0.774	20.0***		c	1.002	16.2***		0.955	13.6***	
	(2.40)			(-4.10)***				(3.75)			(-1.44)		
	0.928		28.1***	0.218		40.9***		0.971		20.1***	0.324		29.6***
	(-2.70)			(-6.03)***				(-1.42)			(-5.19)***		
y^{*E}	1.000	6.41***		0.984	11.5***		y^*	1.000	4.28***		0.981	9.93***	
	(3.80)			(-0.90)				(1.40)			(-1.04)		
	0.876		22.6***	0.091		34.9***		0.854		16.7**	-0.062		28.3***
	(-2.54)			(-6.73)***				(-2.44)			(-6.33)***		
y^{*E*}	1.000	4.24***		0.986	8.42***		p^*	1.000	14.2***		0.911	18.4***	
	(1.70)			(-0.78)				(0.20)			(-2.64)***		
	0.849		16.2**	-0.080		26.1***		1.000		30.4***	0.239		41.4***
	(-2.20)			(-6.18)***				(-0.04)			(-4.64)***		
y	1.000	2.95**		0.969	7.86***		g	0.910	10.2***		0.774	9.18***	
	(1.80)			(-1.14)				(-0.90)			(-2.05)**		
	0.758		13.4	0.609		14.8*		0.660		15.5**	0.496		12.6
	(-3.20)*			(-3.33)*				(-2.29)			(-3.48)**		
p	1.000	20.4***		0.851	17.6***		i	0.988	37.9***		0.105	51.2***	
	(0.80)			(-3.42)***				(-1.20)			(-6.01)***		
	0.994		28.9***	0.199		33.0***		0.819		43.2***	-0.054		54.7***
	(-1.12)			(-5.81)***				(-2.52)			(-6.11)***		
m	1.006	11.9***		0.501	17.0***		$i-\dot{p}$	0.985	23.9***		0.516	39.0***	
	(2.79)			(-4.72)***				(-0.69)			(-4.58)***		
	0.736		25.4***	-0.168		30.6***		0.761		29.4***	0.344		39.1***
	(-3.26)*			(-6.94)***				(-2.68)			(-4.96)***		

The estimated first-order autoregressive parameter a_1 (see equations (S1) - (S2) of Table 4.1) is given in the column denoted t and the t value of this estimate below in parentheses. The null hypotheses are: t; $a_1=1$, F4 and F5; $a_1=a_2=1$, $a_3=0$ (see equations (S1) - (S2) of Table 4.1). The Dickey-Fuller-Hasza tests are based on regressions with twelve lags of the differenced variable under investigation. Dummy variables taking the value 1 for the following variables (dates) and 0 otherwise have been employed, where appropriate, in the test regressions: y^{*E*}, y^* (1980M5) and y (1971M2 - 1971M3). Critical values for the tests are tabulated by Fuller (1976) and Hasza & Fuller (1982). Rejection at the 10,5 and 1 per cent level of significance is indicated by *, ** and ***, respectively.

TABLE A4.6

Tests for the order of integration of seasonal variables using data from the period 1980M1 - 1988M8

	Data in levels			Data in first differences				Data in levels			Data in first differences		
	t	F4	F5	t	F4	F5		t	F4	F5	t	F4	F5
y*U	1.000	5.39***		0.846	11.7***		c	1.004	11.9***		0.993	6.66***	
	(2.00)			(-2.30)**				(3.01)			(-0.17)		
	0.823		16.3**	0.037		28.1***		1.069		11.6	-0.067		15.2**
	(-2.14)			(-4.01)**				(0.87)			(-4.18)***		
y*E	1.000	2.00		0.961	4.05**		y*	1.000	3.38**		0.973	3.88**	
	(1.45)			(-1.79)*				(2.00)			(-1.22)		
	0.600		16.4**	-0.723		17.8**		0.520		13.7*	-0.382		11.5
	(-3.29)*			(-5.82)***				(-3.22)*			(-3.69)**		
y*E*	1.000	2.51*		0.968	3.77**		p*	1.000	9.77***		0.820	7.79***	
	(1.60)			(-1.53)				(0.80)			(-2.42)**		
	0.645		10.1	-0.147		10.0		0.957		22.4***	-0.329		22.7***
	(-2.29)			(-3.18)*				(-2.49)			(-4.30)***		
y	1.001	4.68***		0.957	7.94***		g	0.916	4.93***		0.819	4.46***	
	(2.03)			(-1.58)				(-0.65)			(-1.30)		
	0.524		7.38	0.539		9.30		0.109		7.20	0.022		5.58
	(-1.95)			(-1.70)				(-2.92)			(-2.99)		
p	1.000	9.35***		0.688	10.8***		i	0.990	15.6***		-0.055	16.7***	
	(1.00)			(-3.38)***				(-1.13)			(-4.14)***		
	0.965		19.7***	-0.458		24.4***		0.866		19.1***	-0.308		19.5***
	(-1.75)			(-4.36)***				(-1.08)			(-4.18)***		
m	1.011	4.82***		0.550	4.29***		$i-\dot{p}$	0.989	6.36***		0.523	10.1***	
	(2.56)			(-2.50)**				(-0.52)			(-2.64)***		
	0.386		13.9*	-1.229		17.2**		0.971		26.4***	-0.446		31.2***
	(-2.81)			(-6.06)***				(-0.19)			(-4.52)***		

The estimated first-order autoregressive parameter a_1 (see equations (S1) - (S2) of Table 4.1) is given in the column denoted t and the t value of this estimate below in parentheses. The null hypotheses are: t; $a_1=1$, F4 and F5; $a_1=a_2=1$, $a_3=0$ (see equations (S1) - (S2) of Table 4.1). The Dickey-Fuller-Hasza tests are based on regressions with twelve lags of the differenced variable under investigation. Dummy variables taking the value 1 for the following variables (dates) and 0 otherwise have been employed, where appropriate, in the test regressions: y*E*, y* (1980M5) and y (1971M2 - 1971M3). Critical values for the tests are tabulated by Fuller (1976) and Hasza & Fuller (1982). Rejection at the 10,5 and 1 per cent level of significance is indicated by *, ** and ***, respectively.

APPENDIX 5

Tests for Cointegration Using Data from Subperiods

TABLE A5.1

Maximum likelihood estimation results for a VAR model for the vector $\underset{\sim}{y}_t = (y_t^{*U}\ y_t^{*E}/y_t^{*E*}\ y_t)'$ using data from subperiods

	$\underset{\sim}{y}_t = (y_t^{*U}\ y_t^{*E}\ y_t)'$			$\underset{\sim}{y}_t = (y_t^{*U}\ y_t^{*E*}\ y_t)'$
	Tests for the number of cointegrating vectors			
H_2	$-2\ln(Q)$	trace	maximum eigenvalue	$-2\ln(Q)$
		1960M1 - 1972M12		
$r\leq 2$	0.028	8.08	8.08	0.422
$r\leq 1$	4.33	17.8	14.6	5.98
$r=0$	17.4	31.3	21.3	13.6
		1973M1 - 1988M8		
$r\leq 2$	0.042	8.08	8.08	0.237
$r\leq 1$	16.9	17.8	14.6	7.51
$r=0$	38.0	31.3	21.3	30.9
		1980M1 - 1988M8		
$r\leq 2$	0.043	8.08	8.08	0.023
$r\leq 1$	8.94	17.8	14.6	9.98
$r=0$	41.8	31.3	21.3	35.5

Normalized eigenvectors, loadings, and corresponding long-run coefficient matrices when the models are restricted according to the outcomes of the tests for the number of cointegrating vectors

1960M1 - 1972M12

- -

1973M1 - 1988M8

$$\begin{bmatrix} -27.1 \\ 12.5 \\ 1.00 \end{bmatrix} \begin{bmatrix} 0.000 \\ -0.010 \\ -0.041 \end{bmatrix} \begin{bmatrix} -0.012 & 0.006 & 0.000 \\ 0.264 & -0.122 & -0.010 \\ 1.12 & -0.519 & -0.041 \end{bmatrix} \begin{bmatrix} -12.0 \\ 6.03 \\ 1.00 \end{bmatrix} \begin{bmatrix} 0.000 \\ -0.021 \\ -0.104 \end{bmatrix} \begin{bmatrix} -0.005 & 0.003 & 0.000 \\ 0.254 & -0.128 & -0.021 \\ 1.25 & -0.628 & -0.104 \end{bmatrix}$$

1980M1 - 1988M8

$$\begin{bmatrix} 1.22 \\ -3.26 \\ 1.00 \end{bmatrix} \begin{bmatrix} 0.013 \\ 0.072 \\ -0.157 \end{bmatrix} \begin{bmatrix} 0.016 & -0.042 & 0.013 \\ 0.088 & -0.236 & 0.072 \\ -0.192 & 0.514 & -0.157 \end{bmatrix} \begin{bmatrix} 0.101 \\ -2.59 \\ 1.00 \end{bmatrix} \begin{bmatrix} 0.012 \\ 0.097 \\ -0.127 \end{bmatrix} \begin{bmatrix} 0.001 & -0.030 & 0.012 \\ 0.010 & -0.250 & 0.097 \\ -0.013 & 0.330 & -0.127 \end{bmatrix}$$

In addition to the variables in $\underset{\sim}{y}_t$, the models include a constant, 11 seasonal dummy variables and, where appropriate, a strike dummy taking the value 1 in 1971M2 - 1971M3 and 0 otherwise. The test statistic for the number of cointegrating vectors is denoted $-2\ln(Q)$, and critical values (95 per cent quantiles) have been tabulated by Johansen & Juselius (1989).

TABLE A5.2

Maximum likelihood estimation results for VAR models using data from subperiods

Tests for the number of cointegrating vectors

H_2	$-2\ln(Q)$	trace	maximum eigenvalue
1960M1 - 1972M12 ($\underset{\sim}{y}_t = (y_t\ p_t\ c_t\ i_t^*\ e_t\ p_t^{02})'$)			
r<5	2.00	8.08	8.08
r<4	8.42	17.8	14.6
r<3	20.7	31.3	21.3
r<2	43.2	48.4	27.3
r<1	94.4	70.0	33.3
r<0	160	n.a.	n.a.
1973M1 - 1988M8 ($\underset{\sim}{y}_t = (y_t\ p_t\ m_t\ i_t^*\ e_t\ p_t^*)'$)			
r<5	1.80	8.08	8.08
r<4	10.5	17.8	14.6
r<3	25.9	31.3	21.3
r<2	58.6	48.4	27.3
r<1	105	70.0	33.3
r=0	194	n.a.	n.a.
1980M1 - 1988M8 ($\underset{\sim}{y}_t = (y_t\ p_t\ m_t\ i_t^*\ e_t)'$)			
r<4	2.09	8.08	8.08
r<3	10.1	17.8	14.6
r<2	27.1	31.3	21.3
r<1	52.4	48.4	27.3
r=0	104	70.0	33.3

Normalized eigenvectors, loadings, and corresponding long-run coefficient matrices when the models are restricted according to the outcomes of the tests for the number of cointegrating vectors

1960M1 - 1972M12 ($\underset{\sim}{y}_t = (y_t\ p_t\ c_t\ i_t^*\ e_t\ p_t^{02})'$)

$$\begin{bmatrix} 1.00 & 1.00 \\ 0.752 & -0.697 \\ -0.467 & 1.54 \\ 121 & 23.8 \\ 0.005 & 0.069 \\ -2.21 & -1.69 \end{bmatrix} \begin{bmatrix} -0.138 & -0.204 \\ -0.141 & 0.114 \\ 0.040 & -0.078 \\ -0.000 & 0.000 \\ -0.197 & -0.054 \\ -0.011 & 0.008 \end{bmatrix} \begin{bmatrix} -0.342 & 0.039 & -0.250 & -21.5 & -0.015 & 0.650 \\ -0.028 & -0.186 & 0.241 & -14.6 & 0.007 & 0.121 \\ -0.038 & 0.085 & -0.139 & 2.97 & -0.005 & 0.043 \\ -0.000 & -0.000 & 0.000 & -0.037 & 0.000 & 0.001 \\ -0.251 & -0.111 & 0.009 & -25.1 & -0.005 & 0.527 \\ -0.003 & -0.013 & 0.017 & -1.11 & 0.000 & 0.011 \end{bmatrix}$$

1973M1 - 1988M8 ($\underset{\sim}{y}_t = (y_t\ p_t\ m_t\ i_t^*\ e_t\ p_t^*)'$)

$$\begin{bmatrix} 1.00 & 1.00 & 1.00 \\ -0.271 & 1.15 & 31.7 \\ -1.08 & -0.580 & 3.37 \\ -5.35 & -0.828 & 29.9 \\ 0.193 & -0.335 & 6.59 \\ 1.75 & -0.924 & -43.2 \end{bmatrix} \begin{bmatrix} -0.055 & -0.378 & -0.002 \\ -0.021 & 0.003 & -0.002 \\ 0.020 & 0.168 & -0.002 \\ 0.009 & -0.002 & -0.001 \\ -0.001 & -0.008 & -0.003 \\ -0.012 & 0.009 & 0.001 \end{bmatrix} \begin{bmatrix} -0.435 & -0.489 & 0.271 & 0.543 & 0.102 & 0.343 \\ -0.020 & -0.069 & 0.012 & 0.034 & -0.021 & 0.068 \\ 0.186 & 0.124 & -0.126 & -0.306 & -0.066 & -0.032 \\ 0.005 & -0.050 & -0.013 & -0.088 & -0.007 & 0.079 \\ -0.011 & -0.091 & -0.003 & -0.067 & -0.015 & 0.117 \\ -0.002 & 0.037 & 0.010 & 0.079 & -0.001 & -0.062 \end{bmatrix}$$

1980M1 - 1988M8 ($\underset{\sim}{y}_t = (y_t\ p_t\ m_t\ i_t^*\ e_t)'$)

$$\begin{bmatrix} 1.00 & 1.00 \\ -5.93 & -0.133 \\ 0.778 & -0.267 \\ -6.09 & -0.385 \\ 0.474 & 0.307 \end{bmatrix} \begin{bmatrix} 0.003 & 0.336 \\ 0.005 & -0.020 \\ -0.003 & -0.008 \\ -0.000 & -0.042 \\ 0.003 & 0.096 \end{bmatrix} \begin{bmatrix} -0.333 & 0.027 & 0.092 & 0.111 & -0.102 \\ 0.025 & -0.032 & -0.001 & -0.038 & 0.009 \\ 0.005 & 0.017 & -0.004 & 0.015 & 0.001 \\ 0.042 & -0.006 & -0.011 & -0.016 & 0.013 \\ -0.093 & -0.005 & 0.028 & 0.019 & -0.028 \end{bmatrix}$$

In addition to the variables in $\underset{\sim}{y}_t$, the models include a constant, 11 seasonal dummy variables and, where appropriate, a strike dummy taking the value 1 in 1971M2 - 1971M3 and 0 otherwise. The test statistic for the number of cointegrating vectors is denoted $-2\ln(Q)$, and critical values (95 per cent quantiles) have been tabulated by Johansen & Juselius (1989).

APPENDIX 6

Estimation Results for an Extended Business Cycle Fluctuation Model

TABLE A6.1

Contemporaneous structural relationships in VAR models using data from subperiods

Model variables Matrix of contemporaneous relationships

1960M1 - 1972M12

$$\underset{\sim}{y}_t = \begin{bmatrix} y_t \\ p_t \\ l_t \\ i_t^* \\ e_t \\ p_t^{02} \\ y_t^* \end{bmatrix} \qquad \Gamma_0 = \begin{bmatrix} 0 & -0.537 & 0 & 2.30 & 0.147 & -0.054 & 0.026 \\ 0 & 0 & 0 & -0.126 & 0 & 0 & -0.001 \\ -0.007 & 0.088 & 0 & 0.055 & 0 & 0 & 0.012 \\ 0 & & & \ldots & & & 0 \\ 0 & & & \ldots & & & 0 \\ 0 & & & \ldots & & & 0 \\ 0 & 0 & 0 & 0.091 & 0 & 0.261 & 0 \end{bmatrix}$$

1973M1 - 1988M8

$$\underset{\sim}{y}_t = \begin{bmatrix} y_t \\ p_t \\ m_t \\ i_t^* \\ e_t \\ p_t^* \\ y_t^* \end{bmatrix} \qquad \Gamma_0 = \begin{bmatrix} 0 & 1.19 & 0 & 0.994 & -0.026 & -0.346 & 0.898 \\ 0 & 0 & 0 & -0.062 & 0 & 0.090 & 0 \\ 0.097 & -0.216 & 0 & -0.960 & 0 & 1.45 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0.181 & 0.060 \\ 0 & & & \ldots & & & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & -0.014 \\ 0 & & & \ldots & & & 0 \end{bmatrix}$$

1980M1 - 1988M8

$$\underset{\sim}{y}_t = \begin{bmatrix} y_t \\ p_t \\ m_t \\ i_t^* \\ e_t \\ y_t^* \end{bmatrix} \qquad \Gamma_0 = \begin{bmatrix} 0 & 1.01 & 0 & 1.15 & -0.099 & 0.349 \\ 0 & 0 & 0 & -0.009 & 0 & 0 \\ 0.076 & -0.484 & 0 & 0.119 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0.053 \\ 0 & & & \ldots & & \\ 0 & & & \ldots & & 0 \end{bmatrix}$$

In addition to the components of $\underset{\sim}{y}_t$, the models include a constant, 11 seasonal dummy variables and, where appropriate, a strike dummy taking the value 1 in 1971M2 - 1971M3 and 0 otherwise.

TABLE A6.2

Long-run characteristics of extended VAR models using data from subperiods

Statistic	Variable	Innovation in						
		y	p	m/c	i*	e	p*/p02	y*
				1960M1 - 1972M12				
Cumulative	y	0.545	0.227	-0.096	0.109	0.161	0.093	-0.021
impulse	p	-0.063	1.33	-0.011	0.039	0.112	0.028	-0.016
response	c	0.426	-0.638	2.25	-0.322	0.053	-0.055	0.039
to a one	i*	0.247	-0.212	0.425	1.10	-0.134	0.155	0.037
standard	e	-0.005	0.892	-0.449	0.375	1.28	0.222	0.016
error	p02	0.133	0.207	-0.108	0.154	0.219	0.712	-0.074
innovation	y*	0.032	-0.108	0.168	0.017	-0.008	-0.075	0.476
Percentage	y	75.1	2.24	5.20	6.08	1.71	7.42	2.24
of the	p	2.25	72.1	8.21	0.973	14.3	1.64	0.549
expected	c	0.713	1.08	92.2	2.03	2.23	0.628	1.09
long-run	i*	0.562	0.468	4.41	87.7	4.62	0.649	1.56
squared	e	1.19	0.514	0.339	0.642	94.8	1.76	0.728
prediction	p02	0.857	0.471	3.31	1.62	6.37	86.4	0.942
error	y*	2.29	2.61	0.730	3.47	1.04	4.03	85.8
				1973M1 - 1988M8				
Cumulative	y	0.510	-0.287	0.066	0.088	-0.019	-0.372	0.002
impulse	p	0.048	2.98	0.068	-0.254	-0.190	2.85	-0.228
response	m	-0.035	0.713	0.615	-0.153	0.108	0.907	-0.045
to a one	i*	0.085	0.872	-0.136	1.43	-0.317	1.43	-0.059
standard	e	-0.180	-0.063	-0.101	-0.141	1.40	-0.017	-0.051
error	p*	-0.093	3.03	0.144	0.178	-0.275	5.13	-0.469
innovation	y*	0.071	0.843	0.006	0.499	-0.294	1.04	0.537
Percentage	y	70.7	3.37	1.05	1.15	4.27	1.08	18.4
of the	p	0.669	63.0	1.77	7.74	1.03	22.2	3.57
expected	m	3.95	0.701	86.4	3.78	1.57	1.68	1.87
long-run	i*	1.34	2.15	3.22	81.7	1.23	3.80	6.60
squared	e	0.853	2.29	1.12	2.58	90.6	1.47	1.09
prediction	p*	0.386	16.2	2.59	3.70	0.667	73.7	2.76
error	y*	2.18	0.774	0.386	3.04	1.51	3.10	89.0
				1980M1 - 1988M8				
Cumulative	y	0.410	0.862	-0.035	0.196	-0.033		-0.030
impulse	p	-0.065	2.76	-0.001	-0.035	-0.381		-0.391
response	m	-0.010	-0.494	0.420	0.030	-0.002		0.067
to a one	i*	-0.000	-0.083	-0.146	1.38	-0.167		0.076
standard	e	-0.075	1.17	0.133	-0.203	1.41		-0.080
error	y*	0.040	0.025	-0.022	0.462	-0.297		0.578
innovation								
Percentage	y	73.3	3.59	5.23	7.38	1.12		9.32
of the	p	7.56	68.5	1.63	6.17	12.5		3.62
expected	m	1.58	0.960	79.7	4.83	2.02		10.9
long-run	i*	4.64	0.275	3.82	83.8	1.29		6.18
squared	e	5.67	9.65	1.28	3.99	75.6		3.79
prediction	y*	1.05	8.06	2.88	4.19	0.557		83.3
error								

In addition to the variables in $\underset{\sim}{y}_t$, the models comprise a constant, 11 seasonal dummy variables and, where appropriate, as strike dummy taking the value 1 in 1971M2 - 1971M3 and 0 otherwise. The models are estimated by equation-by-equation OLS using the restrictions reported in Table A6.1. In all calculations, the long run refers to a horizon of 60 months. Numbers across prediction error rows may not sum to 100.000 due to rounding errors.

APPENDIX 7

Analysis of Structural Innovations as Identified Using Low Frequency Restrictions

TABLE A7.1

Analysis of structural innovations in the vector $\underset{\sim}{y}_t = (y_t\ p_t\ m_t\ i_t^*\ e_t\ p_t^*)'$ as identified using low frequency restrictions during the period 1960M1 - 1988M8

Statistic		Innovation					
		ξ^y	ξ^p	ξ^m	ξ^{i*}	ξ^e	ξ^{p*}
Mean x 10^3		0.000	0.000	0.000	0.000	0.000	0.000
Variance x 10^3		0.858	0.022	0.810	0.016	0.137	0.006
Skewness		-0.058	0.345**	0.557***	0.312**	9.68***	0.624***
Kurtosis		2.98***	7.06***	2.72***	3.80***	129***	1.84***
Autocorrelation	1	0.051	-0.009	-0.010	0.008	0.034	0.013
	2	0.085	-0.036	0.002	-0.004	-0.097*	0.003
	3	0.048	-0.002	0.018	0.001	-0.084	-0.006
	4	0.023	-0.021	-0.014	-0.006	0.027	-0.031
	5	0.064	-0.027	0.035	-0.005	0.001	0.031
	6	0.002	0.044	-0.027	0.026	0.001	0.024
	7	0.041	0.000	-0.003	-0.056	0.017	0.034
	8	0.041	-0.010	0.005	0.024	-0.004	-0.019
	9	0.068	0.025	0.063	0.018	-0.031	-0.059
	10	0.049	-0.049	0.034	-0.053	0.040	0.066
	11	0.132**	0.045	0.043	0.161***	0.020	0.035
	12	0.033	0.056	0.156***	-0.034	-0.087	0.022
Significant positive shock		1964M7	1963M9	1973M4	1964M11	1967M10	1961M6
		1969M7	1964M1	1973M12	1973M1	1977M4	1961M8
		1971M4	1967M4	1975M12	1973M2	1978M2	1962M4
		1971M6	1968M1	1983M5	1973M7	1982M10	1963M2
		1972M2	1971M6	1987M3	1973M11		1969M6
		1972M3	1973M7	1987M8	1974M3		1970M1
		1978M3	1974M2	1988M6	1980M11		1970M7
		1980M7	1976M7		1985M1		1970M12
			1982M5		1985M2		1972M9
			1982M12				1973M11
			1983M5				1974M10
							1975M5
							1976M8
							1979M7
							1980M1
Significant negative shock		1971M2	1967M1	1962M6	1970M2	1968M1	1962M8
		1971M3	1968M5	1966M3	1973M10		1963M7
		1972M4	1969M3	1973M10	1974M4		1969M5
		1975M5	1972M1	1977M12	1974M5		1971M9
		1975M7	1974M6	1986M8	1975M1		1975M12
		1976M7	1975M6	1986M12	1976M11		
		1977M3		1987M12	1980M5		
		1977M7			1981M11		
		1978M7			1982M5		
					1984M11		

Structural impulses are estimated from the (near) VAR model (7.2i) - (7.2vi) comprising the elements of $\underset{\sim}{y}_t$, a constant, 11 seasonal dummy variables and in (7.2i) a strike dummy taking the value 1 in 1971M2 - 1971M3 and 0 otherwise (the strike dates are listed as significant negative shocks). The model is estimated by equation-by-equation OLS using restrictions reported in expression (7.1). Significant values at the 10,5 and 1 per cent level of significance are indicated by *, ** and ***, respectively. Listed shocks exceed the two standard error limit (the dummy variable strike dates are listed as significant negative shocks).

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