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Relative Prices and Monetary Policy Information Variables: Long Run Evidence From Finland

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Abstract

Under certain assumptions, the permanent income model yields the result that prices of different products share common stochastic trends. We construct four price series from the components of the consumer price index, combine this four variable system with various macroeconomic variables, such as broad money (M2), nominal exchange rates, import and export prices in domestic currency etc., with the ultimate hope of capturing empirical regularities between this price structure and the proposed macroeconomic information. It is found that broad money, money market and bond yield indices, nominal exchange rates and import prices contain long-run information about the price level. We also find that the common trends of price groups cointegrate with the broad monetary aggregate and the money market yield index, indicating that these variables might be the driving forces of underlying inflation.

Tiivistelmä

Pysyväistulohypoteesin mukaan eri hyödykkeiden ja palvelusten hinnoilla on tietyin oletuksin yhteinen stokastinen trendi. Tämä yhteinen hintatrendi, josta lyhyen aikavälin hintojen suhteellisen muutokset on puhdistettu pois voidaan tulkita erääksi pohjainflaation mittariksi. Tutkimuksessa disaggregoidaan kuluttajahintaindeksi neljään alaerään. Tämän jälkeen tarkastellaan kehittyvätkö nämä alaerät pitkällä aikavälillä yhdessä talouden eri makromuuttujien kanssa. Tarkasteltuja makromuuttujia ovat mm. laveat raha-aggregaatit, valuuttakurssit, ulkomaankaupan hinnat sekä erilaiset tuottoindeksit. Tulokseksi saadaan, että lavea raha, raha- ja joukkolainamarkkinoiden tuottoindeksit sekä tuontihinnat sisältävät informaatiota kuluttajahinnoista pitkällä aikavälillä. Lisätulokseksi saadaan, että rahamarkkinoiden tuottoindeksi ja lavean rahan määrä kehittyvät pitkällä aikavälillä yhdessä ehdotetun pohjainflaatiomittarin kanssa.

Contents

1 Introduction

More often than not, an empirical (macro)economist starts from the presumption that the best way to understand the determinants and dynamics of prices and inflation in an economy is to model the dynamic behaviour of some broad price index such as the CPI or GDP deflator. Moreover, in an open economy context the problem has typically been analysed either in terms of validity of the Purchasing Power Parity (PPP) or less stringently in terms of modelling deviations from the PPP, ie modelling the dynamic behaviour of the terms of trade. While certainly not without its merits, this type of modelling strategy may involve some loss of information, eg as to the precise nature of the (price) mechanism by which various shocks to the economy are transmitted to variations in the price level. In particular, modelling the behaviour of the relative price structure, either of the broad commodity aggregates or of the industrial sectors, in conjunction with the relevant macroeconomic information may produce valuable information regarding the adjustment of the economy to various macroeconomic shocks.

In this paper we are in fact proposing an empirical modelling strategy, whereby potentially useful information on the behaviour of the aggregate price level can be inferred from the behaviour of the component price indexes combined with various kinds of macroeconomic information, preferably those macroeconomic variables containing information relevant to the central bank and to monetary policy. To be more precise, we construct four price series from the components of the consumer price index, combine this four variable system with various macroeconomic variables such as broad money (M2), nominal exchange rates, import prices in domestic currency etc., with the ultimate hope of capturing empirical regularities between this particular price structure and the cited macroeconomic variables.

We must, of course, substantiate the concept of 'empirical regularities' in this particular context. There are several potentially useful approaches. First of all, since various price series are often shown to be unit root processes, as is true of many macroeconomic series, a natural question to ask is whether there exist any cointegration relationships between CPI component prices and other macroeconomic variables. If so, cointegration renders precise empirical content to the claim that there is a nontrivial relationship between CPI component prices and macroeconomic information. To allow for the possibility of cointegration relationships between component prices and other macroeconomic variables in our empirical model, we apply the FIML methodology proposed by Johansen to our data set. This has the added advantage, because of full information inference, of being able to produce the error correction representation of our k-variable ($k > 4$) system simultaneously with the long-run cointegration relationships. Furthermore Johansen's framework provides a natural testing methodology, by which we can test, eg for the stationarity of the component series, exclusion of any of the component series from the cointegration relationships and, what is often highly important, weak exogeneity of any of the variables in the system, within the same framework. In what follows we shall utilize these various aspects of the VAR methodology.

Of course, there need not exist any cointegration relationship between the component price series and the proposed macroeconomic variable. But an equally obvious statement is that this does not mean that there is no relationship between the two 'blocks' of variables. If there does not exist common comovement in the *level* of the series ("long-run" relationship), *changes in the level* of the series may be interrelated, though such interrelationships may take various, perhaps complex, forms. These short-run relationships between variables can arise from eg dynamic adjustment in the economy, and can thus be quite reasonably interpreted as 'disequilibrium' co-movements.

The paper is organized as follows. In the next section we describe the data and the econometric methodology applied. Section 3 then presents a summary of the main results from the cointegration analysis and exclusion and weak exogeneity tests as well as the results of the common trend analysis. Finally, section 4 concludes by discussing possible interpretations of the results and attempting to justify the cointegration of prices with a version of the permanent income model.

2 Data and Econometric Setup

The usual way of analysing temporal correlation between prices and indicators includes descriptions of cross-correlation and Granger non-causality. In such studies (see eg McNees 1989 or Bank of England 1993) prices are usually measured by the consumer price index (CPI). The set of economic variables often consists of monetary and credit aggregates, interest rates and the slope of the yield curve and variables related to the real economy and wage formation.

In a slightly more careful analysis the non-stationarity of the data is tested and the focus is on differenced data. Usually these kinds of studies omit the long run co-movement of the variables. In our setup the long run co-movement is the main issue and at this stage of the study we do not discuss causality. We raise the question of exogeneity (with respect to long run co-movement). In our case the information variable is weakly exogenous if the disturbances in the co-movement do not influence the short-run development of the information variable, ie only the prices adjust in the long-run.

We use a disaggregated price index. There are several possible ways to disaggregate consumer prices. Since we cannot disaggregate consumer prices by sectors, we are restricted to disaggregation by commodities. Disaggregation makes it possible to analyse co-movement among price subgroups, ie we can try to discover the common underlying stochastic forces which are driving the prices. In a sense we look for underlying inflation. This cannot be done with the aggregate price index.

According to Granger (1980), the aggregation might — in certain cases — yield processes that require fractional differencing to be $I(0)$. The aggregate price index might contain these features, ie there might be strong persistence in the inflation rate (see, eg Delgado and Robinson 1994). In using disaggregate indices we might be able to get rid of the some part of the persistence. This is, of course, not obvious, since the possibility of persistence applies to the price groups too.

We believe that the main advantage of using disaggregated price index is that it allows us to study the richer interrelationships between information variables and prices in general. Dynamics and responses to shocks can be modelled with fewer constraints.

We refer economic variables as *information variables*, although the purpose of the study is to test whether they carry any long-run information about the price level. We have divided the variables into three blocks. The first block comprises financial variables, the second foreign trade prices and the third real variables.

In order to study the long run, the length of the data set is chosen to be as long as possible. Starting from early 1970s, the data period covers two large booms and recessions in the Finnish economy. Despite two major changes that took place in the economy during that period (financial deregulation in the 1980s and floating exchange rates since September 1992), we believe that this period is sufficiently homogeneous with respect to the inflation mechanisms. Such a long estimation period, however, restricts the choice of information variables.

Broad monetary and credit aggregates are traditionally used as inflation indicators. They are widely monitored by central banks. Short- and long-term interest rates are included in the form of yield indices, which are computed from monthly returns[1]. Share prices are also included as an indicators. The relationship between domestic prices and exchange rate developments plays an important role in Finnish aggregate price formation.

Export and import prices carry much information about prices in Finland. According to some studies (see eg Mellander *et al* 1990), the terms of trade has an important role in the dynamics of the Finnish economy. It is often said that inflation is a procyclical phenomenon. That can be examined by including GDP measures in the data set.

In the following two sections, we describe the data in details and explore the econometric model used to test for co-movement, exogeneity and driving forces.

2.1 Data

The data consist of aggregated main subgroups in the Finnish consumer price index and a set of macroeconomic variables. The data is monthly and covers the period January 1972 – December 1993. Thus the total number of obser-

[1]Since these indices are cumulants of interest earnings and price changes, they follow the same kind of trending behaviour as nominal macroeconomic variables.

vations is 252. The period is long enough to capture the boom and recession of the 1970s, the financial deregulation and boom of the 1980s and the deep recession in the 1990s.

We have divided the Finnish consumer price index into four subgroups. The first subgroup (`Food`) includes food, beverages and tobacco. The second subgroup (`Cloth`) consists mainly of imported commodities; it includes clothing and footwear and household equipment. Housing costs, including heating, lighting and rents, is the third subgroup (`House`). Unfortunately, housing prices are included only for the post-1985 period due to the historical definitions of the index. So, the housing subgroup uses mainly (regulated) rents for the pre-1985 period. Final subgroup (`Other`) includes other goods and services, with transportation being the main subgroup.

The consumer price index was revised in 1972, 1981, 1985 and 1990. The weights are presented in table 1. The interim years are approximated linearly. The weights of the first subgroup are declining and the importance of the final subgroup, consisting mainly of services, is increasing. The price subgroups are presented in the first panel of figure 1.

Table 1: **Weights by main subgroups**

Group	1972=100	1981=100	1985=100	1990=100
Food, beverages and tobacco	32.0	27.6	25.7	22.7
Clothing, footwear and household equipment	13.9	11.9	13.3	12.1
Housing	19.2	20.8	18.4	19.8
Transportation, health care, education and recreation, other goods and services	34.9	39.5	42.6	45.4

The monetary policy information variables are

- `M2`: broad money, M2, mill. FIM, log
- `M3`: broadest money, M3, mill. FIM, log
- `Flending`: credits, banks' markka lending, mill. FIM, log
- `MmY`: money market yield index[2], 1990=100
- `BondY`: bond yield index[3], 1990=100

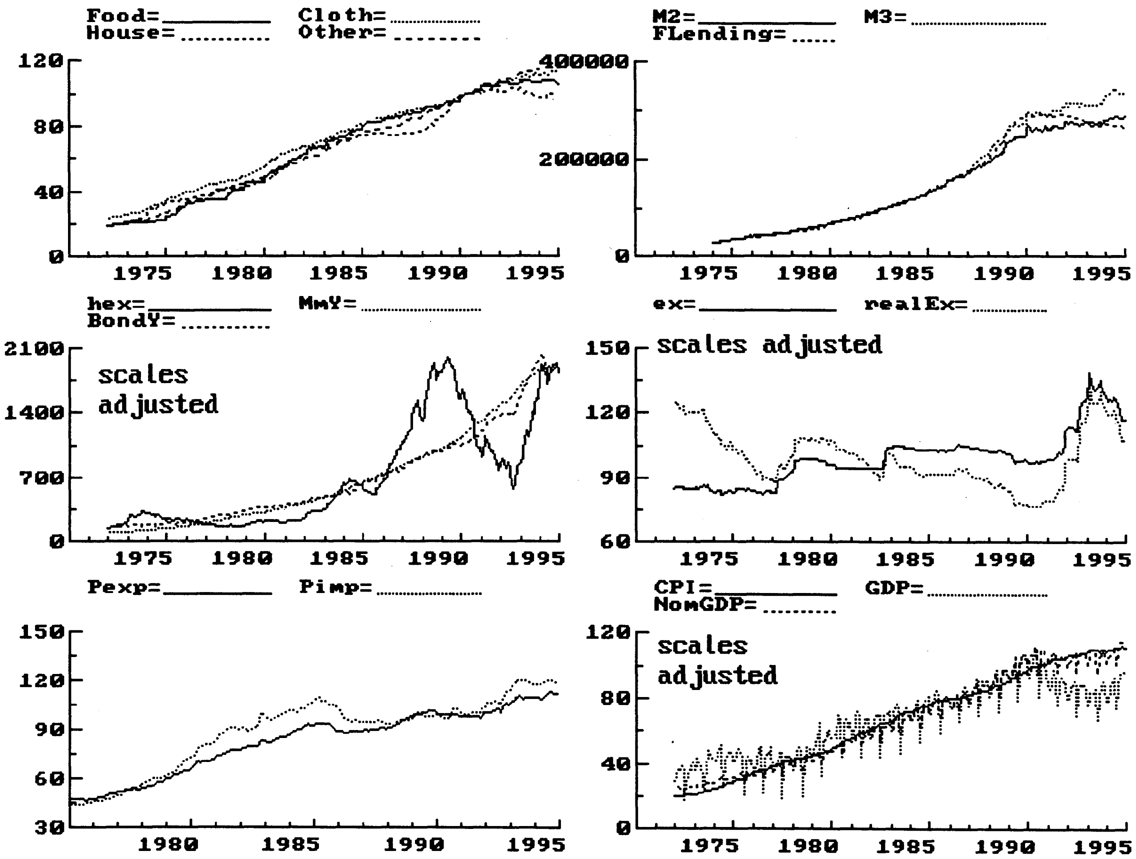
[2]The three-month money market rate is computed from the covered interest parity. The foreign exchange forward markets were controlled (and priced) by the Bank of Finland during the 1970s. In that period they might be referred as shadow money market rates.

[3]The bond yield for the pre-1986 period is based on the computation of bond yields quoted by the Helsinki Stock Exchange; see Alhosuo *et al* (1989).

- `ex`: nominal exchange rate, trade-weighted official index, log
- `realEx`: real exchange rate trade weighted consumer prices and official index, log
- `hex`: share prices, HEX index, 28.12.1990=1000, log
- `Pexp`: export prices, Finnish markka-dominated, index 1990=100, log
- `Pimp`: import prices, Finnish markka-dominated, index 1990=100, log
- `GDP`: GDP volume indicator, 1990=100, log
- `CPI`: consumer price index (aggregate), CPI, 1990=100, log

The information variables are presented in figure 1.

Figure 1: Price Subgroups and Information Variables



2.2 Econometric Setup

Voluminous empirical studies have not rejected the unit root hypothesis for behaviour of macroeconomic time series. It is generally useful to model macroeconomic times series as trending variables even if in the autoregressive representation the roots of the variables are near unity. Since time series behaviour of price levels and the possible monetary policy information variables can be approximated by an $I(1)$ process and we are interested in long-run relationships, we rely on the p-dimesional VAR(k) model with cointegration restrictions. The following statistical model is analyzed:

$$z_t = A_1 z_{t-1} + \ldots + A_k z_{t-k} + \mu + \Psi D_t + \varepsilon_t, \; t = 1, \ldots, T. \tag{1}$$

This can be written in the form:

$$\Delta z_t = \Pi z_{t-1} + \Gamma_1 \Delta z_{t-1} + \ldots + \Gamma_{k-1} \Delta z_{t-k+1} + \mu + \Psi D_t + \varepsilon_t, \; t = 1, \ldots, T. \tag{2}$$

Here $\Pi = -I + \sum_{i=1}^{k} A_i$ and $\Gamma_i = -(\sum_{j=i+1}^{k-1} A_j)$ Matrix Π has a reduced rank in cointegration, rank(Π)$< p$. Any reduced rank matrix can be presented as a product of two matrices. Thus, Π is partitioned as α and β'. Matrix β is called the cointegration matrix. In cointegration, α is called the matrix of loadings of the equilibrium errors of the linear combinations defined by β'. We also test whether we can restrict the presence of the constant to the cointegration vector. In such a case the variables might have a common deterministic linear trend. The vector of (centred) seasonal dummies, D_t, is also included.

The rank of the matrix Π determines the number of cointegration vectors. Two special cases might occur. When all p components of z_t are stationary, matrix Π has the full rank p. When no cointegration exists between the variables in z_t, the matrix Π has zero rank. Then all the variables are integrated order one, but there is no cointegration between them. In such a case, the simple VAR model in differences is the proper framework for empirical analysis.

Johansen (1988 and 1991) has proved that the ML estimators of α and β' can be obtained by solving the following eigenvalue problem

$$|\lambda S_{11} - S_{10} S_{00}^{-1} S_{01}| = 0$$

where S_{11} represents the residual moment matrix from the least squares regression of the levels of the variables on the lagged differences and the deterministic variables (ie levels corrected by short-run dynamics), S_{00} is the residual moment matrix from the least squares regression of the differences of the variables on the lagged differences and the deterministic variables (ie differences corrected by short-run dynamics) and S_{01} is the cross-product moment matrix of the residuals of the above regressions.

The solution of the eigenvalue problem above generates eigenvalues $\hat{\lambda}_1 > \ldots > \hat{\lambda}_p$ and eigenvectors $\hat{V} = [\hat{v}_1 \; \cdots \; \hat{v}_p]$, which are usually normalized as $\hat{V}' S_{11} \hat{V} = I_p$. The estimators of the cointegration vectors $\hat{\beta}$ are $\hat{\beta} = [\hat{v}_1 \cdots \hat{v}_r]$, the eigenvectors corresponding to the r largest eigenvalues. Johansen (1988, 233–236 and 1991) shows that the ML estimator of the space spanned by β is the space spanned by r canonical variates reflecting the r largest squared canonical correlations between the stationary and the non-stationary parts of the system.

It is important to note that we can estimate only the space spanned by β. We can give economic interpretation to the cointegration vectors only after identification. Until then they are a statistical property of the data.

Johansen (1988) derives two tests for the number of cointegration vectors. The null hypothesis can be formulated as

$$H_0 : \text{rank}(\Pi) \leq r.$$

The likelihood ratio test for the hypothesis above is

$$-2\log(Q) = -T \sum_{i=r+1}^{p} \log(1 - \hat{\lambda}_i^2),$$

where $\hat{\lambda}_{r+1}^2, \ldots, \hat{\lambda}_p^2$ are the $p - r$ smallest squared canonical correlations — eigenvalues. Now, the alternative hypothesis is that the number of cointegration vectors is larger than r. This test is called a *trace test.*

Johansen (1988) also derives likelihood ratio tests for different types of hypotheses concerning restrictions on the β-space and α. The test statistics are asymptotically χ^2 distributed. We are particularly interested in testing whether a component of z_t is stationary and whether a component of z_t can be excluded[4] from the cointegration space, ie that it does not cointegrate with the rest of the components of the vector. The hypotheses of interest are formalized respectively as follows

$$H_a : \beta \;=\; H_1\varphi, \qquad H_1\,(p \times s),\ \varphi\,(s \times r),\ r \leq s \leq p \tag{3}$$
$$H_b : \beta \;=\; (H_2, \psi), \qquad H_2\,(p \times r_1),\ \psi\,(p \times r_2),\ r = r_1 + r_2 \tag{4}$$

Sometimes the hypothesis of weak exogeneity is interesting. Johansen (1991) proposes a likelihood ratio test to test the hypothesis of weak exogeneity with respect to long-run parameters[5]. The test statistic is also asymptotically $\chi^2(r(p-m))$ distributed. The hypothesis can be formalized as follows:

$$H_c : A\alpha = 0, \quad \text{where } A\,(m \times p)$$

and m is the number of restrictions.

Cointegration vectors, β, define the direction in which the process is stationary, the orthogonal compelement, $\beta_\perp$, $(\beta'\beta_\perp = 0)$ defines the direction in which the process is non-stationary. The process is pushed along the attractor set $\text{sp}(\beta_\perp)$ by the common trends, $\alpha'_\perp \sum_{i=1}^{t} \varepsilon_i$. The common trends describe

[4]This is called an *exclusion test.*

[5]This is calle a *test of weak exogeneity.*

the "underlying" non-stationary stochastic trends. Juselius (1994, p. 175) defines common trends nicely: "Based on the idea that a driving trend should not be affected by equilibrium forces in the economy, it seems most natural to represent the common trends [...] as $\alpha'_{\perp} \sum_{i=1}^{t} \varepsilon_i$. This gives us linear combinations [...] which are orthogonal to the cointegration relations, $\beta' z_t$. This can easily be seen by premultiplying equation [here (2)] by $\alpha'_{\perp}$."

3 Estimation Results

Our data period covers the period January 1972 – December 1993 (except foreign trade prices, which start from January 1975). The (monthly) inflation rates are quite trend-like and one might suspect that comparable price indices might be $I(2)$ variables. We do not separately test that, but in the case of $I(2)$ness, our model with cointegration describes the linear combinations of the components of z_t, which bring the system from $I(2)$ down to $I(1)$. Thus, the cointegration parts of the conclusions would be still applicable.

3.1 Cointegration of Price Subgroups and Information Variable

We first estimate the model described in equation (2). The vector z_t contains four subgroups of consumer prices (food, beverages and tobacco; clothing and household equipment; housing; other goods and services) and one of the macroeconomic variables. We analyze the macroeconomic variables one by one together with the price subgroups. Neither the information criteria nor sequence of LR tests provide clear information about a suitable lag length of the original VAR model. We use the lag length 13 in every case, since we can thus obtain white noise residuals and since the only possible disadvantage of using 13 is that a huge number of parameters must be estimated. However, the number of observations is reasonable large and we do not lack for degrees of freedom. We estimate our systems with two possible specifications: first, we restrict the constant term to be present only in the cointegration relation. Second, we do not restrict the constant term and thus we allow for the presence of deterministic linear trends in the levels of the variables.

In the exclusion test, we test whether a variable (here the information variable) could be restricted from the cointegration space, ie that the price variables cointegrate without the information variable. The null hypothesis is that a variable is excluded. An information variable is weakly exogenous with respect to long-run parameters α and β if the inference can be conditioned on the variable of interest without losing any information. In model (2) this means that the related row of Π matrix is zero. The null hypothesis is that a variable is weakly exogenous with respect to long-run parameters.

According to the test statistics it is clear that none of variables concerned is stationary. A summary of the exclusion tests and weak exogeneity tests is given in table 2.

The exclusion of the information variable is accepted in the systems consisting the real exchange rate, real or nominal GDP or banks' markka lending. The results are fairly robust whether the constant is restricted or not. If we allow deterministic linear trend, banks' markka lending can no longer be removed from the cointegration space (p-value drops from 0.25 to 0.03). The exclusion of the nominal exchange rates is not robust subject to the treatment of the constant. The nominal exchange rates can also be excluded when deterministic linear trend is allowed.

Broad money clearly cointegrates with price subgroups, with or without deterministic linear trend. Moreover, it is weakly exogenous with respect to the long-run parameters, ie, the cointegrating prices-money vector does not influence broad money. That is, the error correction terms defined by the cointegration space do not enter the equation for money change. However, this does not mean that the causality is from money to prices, since the causal interrelationships might also exist in the short-run structure of the system. The outcome of the analysis of the broadest monetary aggregate, M3, is the same.

Both *money market yield index and bond yield index* clearly cointegrate with price subgroups. The price subgroups and yield indices share a common stochastic trend. The hypothesis of weak exogeneity is not rejected, which is not a very appealing result. One reason for this might be that the yield indices (containing the monthly price changes) are backward-looking indicators rather than forward-looking ones like interest rates. *Share prices* are also cointegrated, but in their case weak exogeneity is rejected, which roughly means that changes in the share prices are influenced by disequilibrium between price indices and share prices.

The *nominal exchange rate* cointegrates with price subgroups. The nominal exchange rate is weakly exogenous with the long-run parameters. The prices of foreign trade have traditionally had an important role in the dynamics of the Finnish economy. It is not surprising that *import prices* cointegrate with price vectors. Thus, the import prices and price subgroups do not diverge, but they have a common underlying driving force. Naturally, import prices are not influenced by the disequilibrium in consumer and import prices. The change in import prices will be passed to consumer prices in the long-run, and no permanent changes in the relative prices can be seen. According to the test results, weak exogeneity of *export prices* is rejected. That might be due to the endogeneity of the nominal exchange rates (the weak exogeneity test result is on the borderline).

Surprisingly, the economic activity variables do not cointegrate with price subgroups. Neither *nominal* nor *real GDP* enters the cointegration relation. They are not weakly exogenous with respect to the long-run parameters. The cointegration relation among price groups (ie relative prices) influences nominal or real GDP growth. The *real exchange rate* is not influenced by changes in relative consumer prices. However, the real exchange rate does not cointegrate with price subgroups. This is probably due to the fact that the price level is already taken into account in the definition of real exchange rate.

Finally, out of curiosity, we test whether the aggregate consumer price index shares a common stochastic trend with its own subgroups. Naturally it does. A confusing result is that in the case of non-restricted constant the consumer price index is exogenous with respect to the long-run parameters. So, the subgroups adjust — not the aggregate index itself — to disequilibrium between consumer prices and its subgroups.

Table 2: Tests of exclusion and weak exogeneity

Information variable	Marginal significance level, exclusion (as null)		Marginal significance level, weak exogeneity (as null)	
	no det. trend	trend allowed	no det. trend	trend allowed
Broad money, M2	0.017	0.004	0.14	0.80
Broadest money, M3	<0.001	0.001	0.24	0.61
Credits	0.25	0.03	0.06	0.29
Money market yield index	<0.001	<0.001	0.62	0.51
Bond yield index	<0.001	<0.001	0.51	0.98
Nominal exchange rate	0.02	0.22	0.04	0.04
Real exchange rate	0.79	0.14	0.36	0.22
Share prices	0.009	0.01	0.02	0.02
Export prices	0.007	<0.001	0.02	<0.001
Import prices	0.02	<0.001	0.75	0.56
Nominal GDP	0.19	0.15	<0.001	0.02
Real GDP	0.63	0.09	0.03	0.04
Consumer price index	<0.001	<0.001	0.01	0.17

Does PPP hold for the disaggregated prices? Next we take a closer look at the price subgroups system with import prices. Table 3 presents the results of the trace tests. The cointegration rank is estimated to be 4. However, the (yet) preliminary recursive analysis of the system indicates that $r = 3$ might be a possible choice as well.

The estimated α and β vectors are presented in the table 4. The only natural candidate for a restriction with clear economic rationale is the hypothesis of purchasing power parity (PPP). Noting that import prices are the product of exchange rates and foreign prices, SP^*, we can write the PPP hypothesis as

$$H_{PPP} : \log(S) + \log(P^*) - \log \mathbf{P} = \text{ stationary},$$

Table 3: Determining the cointegration rank

H_0 :	Eigenvalue, λ	Trace test	Critical value, 95 % fractile
$r = 0$	0.2182	130.3	76.07
$r \leq 1$	0.1310	77.4	53.12
$r \leq 2$	0.0996	47.2	34.91
$r \leq 3$	0.0756	24.7	19.96
$r \leq 4$	0.0354	7.7	9.24

where $\log(\mathbf{P}) = \mathbf{p} = \beta_1 p_1 + \beta_2 p_2 + \beta_3 p_3 + \beta_4 p_4$ is the vector of price subgroups, $\log(P_i) = p_i$, $i = 1, \ldots, 4$. The hypothesis may then be written as $\sum_{i=1}^{4} \beta_i = -1$. According to our statistical model (2), the hypothesis may be formalized as in (4). Since our cointegration rank is greater than one, we cannot test that hypothesis[6] for a single cointegration vector. We test whether the whole cointegration space satisfies this restriction. According to the test result, this is not the case either when $r = 3$ or when $r = 4$ (p value < 0.001).

More promising results are based on extra restrictions on the β space. We impose restrictions on each vector of the three dimensional space. The first restriction zeros import prices, housing and other prices. The second one removes import prices and food prices. The modified PPP is restricted to the third vector. The estimated restricted space is reported in table 5. The restriction has substantially larger p-value (0.04). According to the graphics (figure 2) of the modified PPP relation the past devaluation and the floating markka have caused disequilibrium (in a historical perspective) of the PPP relation.

Surprisingly, many of the economic variables analyzed here contain long run information about prices[7]. There is long-run co-movement between price subgroups and broad money, short- and long-term interest rates (in the form of yield indices), terms of trade, nominal exchange rate, share prices and import prices (each alone with the price subgroups). Thus these variables can be called information variables of monetary policy. However, we have not analyzed the causal relationships.

[6]The hypothesis makes only one restriction on the cointegration space.

[7]This is exactly the opposite result of Ripatti (1995), which analyses the aggregate consumer price index and a set of macroeconomic variables and finds cointegration in very few cases.

Table 4: **Estimated cointegration space and loadings**

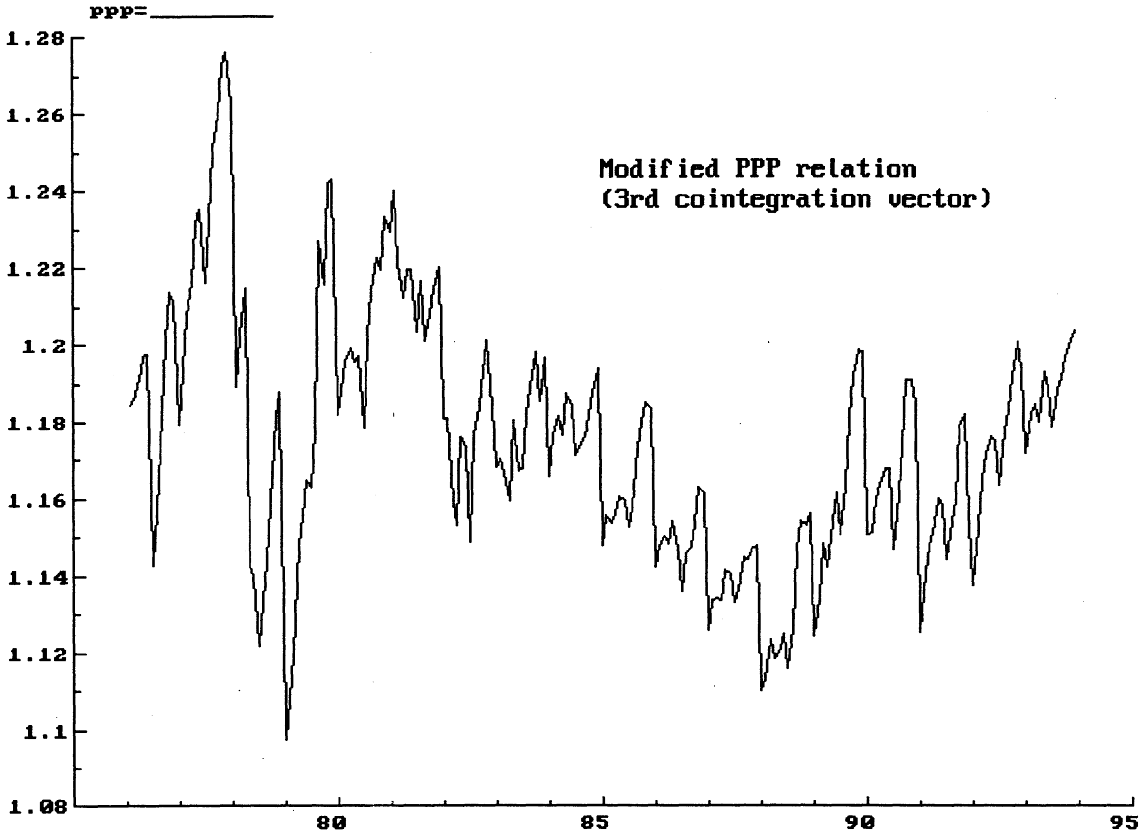
Cointegration space, β'						
Eigenvalue, λ	Import Prices	Food	Clothing	Housing	Other	Constant
0.2182	1	-6.408	6.832	1.078	-0.697	-8.341
0.1310	1	1.215	-7.324	-3.282	7.823	1.954
0.0996	1	-0.750	-5.063	0.445	4.026	1.589
0.0756	1	-3.072	2.000	-0.961	1.352	-1.461

Matrix of loadings, α, and t-values				
Equation				
Import prices	0.002	0.005	0.012	0.004
t-values	0.328	0.895	1.247	0.574
Food	0.012	-0.008	-0.002	0.024
t-values	1.836	1.522	0.216	3.913
Clothing	-0.017	-0.003	0.010	0.002
t-values	6.211	1.482	2.464	0.763
Housing	-0.022	0.013	-0.022	0.003
t-values	4.305	3.172	2.911	0.650
Other	-0.007	-0.006	-0.009	-0.000
t-values	3.037	3.448	2.702	0.061

Table 5: **Restricted cointegration space, β'**

Import Prices	Food	Clothing	Housing	Other	Constant
0	1	-1.430	0	0	2.094
0	0	1	0.985	-2.200	1.360
1	-2.084	-0.335	0.805	0.614	0.294

Figure 2: Modified Purchasing Power Parity



3.2 Information Variables as Common Trends of Price Subgroups

The same data set is used in the second stage of the empirical part. The presence of a deterministic linear trend and lag length are tested in the same way as in the previous section. First we estimate the price subgroup system (ie. we analyze the vector containing only the price subgroups) and test its cointegration rank. Then we extract the common trends (see section 2.2) from that system and continue the analysis with the common trends and information variables (one by one). We repeat the analysis with the constant restricted (no det. trend) and not restricted (det. lin. trend allowed).

The intuition here is as follows. The consumer's budget constraint guarantees that there is cointegration between price subgroups. In the case of r cointegration vectors, there are also $p-r$ common trends in the system[8]. These common trends represent the stochastic trends which are driving the levels of price subgroups, ie. they could be called the underlying forces of inflation. Since we can estimate only the space of common trends, we are not able to indentify, ie. "name", them. But, if an information variable cointegrates with the estimated common "underlying inflation" trends, we might conclude that the information variable correlates with "underlying inflation" in the long-run and we can thus evaluate the importance of the information variable.

First we have to determine the cointegration rank of the system of four ($p = 4$) price subgroup variables. Table 6 reports the eigenvalues and trace test statistics for two cases. The first case contains test statistics for the system, where the constant is restricted to the cointegration relation, ie when there is no deterministic linear trend in the levels of the variables. In the second case the constant is not restricted. Following the test sequence suggested by Johansen (1994) and choosing the 95 per cent fractile, we end up with two cointegration vectors ($r = 2$) and no deterministic linear trend. If we choose the 90 per cent fractile, we find that the cointegration rank is three ($r = 3$). A closer look at the eigenvalues might lead to the idea of one cointegration vector, since the difference between the first and second eigenvalues is large compared to the other differences[9].

Since the determination of the cointegration rank of the price subgroup system is not so straightforward, we repeat the common trend analysis for different choices of cointegration rank. Since the dimension of the price subgroup system is four ($p = 4$), there are three common trends if the cointegration rank is one ($p - r = 4 - 1 = 3$), two common trends if the cointegration rank is two ($p-r = 4-2 = 2$), etc. Next we formulate one-by-onen systems which contain the common trends of price subgroups and each information variable. At this stage of the study we are only interested in whether an information variable cointegrates with common trends of price subgroups. So, it is adequate to test

[8] p is the dimension of the system.

[9] This is not a formal test.

Table 6: **Determination of cointegration rank of the price subgroup system**

H_0 :	no deterministic linear trend				deterministic linear trend			
	eigen-value	trace test	90 % fractile	95 % fractile	eigen-value	trace test	90 % fractile	95 % fractile
$r = 0$	0.17	85.97	49.65	53.12	0.17	82.52	43.95	47.21
$r \leq 1$	0.08	38.24	32.00	34.91	0.07	36.25	26.79	29.68
$r \leq 2$	0.05	18.05	17.85	19.96	0.05	16.88	13.33	15.41
$r \leq 3$	0.02	4.41	7.52	9.24	0.02	4.21	2.69	3.76

only the hypothesis $H_0 : r = 0$. The test statistics and critical values of the hypothesis are presented in table 7.

The results are quite interesting. Compared to the cointegration analysis of table 2, we find much less cointegration between "underlying inflation" and information variables. The most characteristic result is that the money market yield index appears to cointegrate with the common trends, whatever the number of common trends. If we suppose that the monthly money market return consists of expected inflation plus the real interest rate and that the difference between expected and actual inflation is stationary, then our empirical result could confirm that the real rate is stationary (given that the money market yield index and underlying inflation cointegrate with vector $[1 \ -1]'$). The bond yield index behaves in the same manner, but the evidence is weaker.

Nominal GDP seem to be very robust too. This might be due to its CPI component[10], since real GDP cointegrates with "underlying inflation" only when the constant is not restricted. CPI cointegrates with "underlying inflation" — as it must — in every case. In that case the difference between consumer prices and underlying inflation is stationary. We share this idea with Peisa (1994).

If we increase the number of common trends, the broad monetary aggregate, M2, seems to be the next robust one. When we have three common trends the candidates for third place are export and import prices. However, the interpretation is not so straightforward since we do not model the joint behaviour of the variables.

[10]The nominal GDP indicator is constructed here as GDP volume indicator $\times$ CPI.

Table 7: Testing the cointegration rank ($H_0: r = 0$)

Information variable	3 common trends		2 common trends		1 common trend	
	no det. trend	trend allowed	no det. trend	trend allowed	no det. trend	trend allowed
Broad money, M2	62.35	47.11	37.09	36.00	17.05	13.76
Broadest money, M3	60.28	44.40	33.76	31.25	15.06	11.02
Credits	51.42	45.21	27.03	33.36	11.05	9.64
Money market yield index	76.78	69.35	53.23	47.15	31.83	25.85
Bond yield index	58.08	40.03	37.05	17.93	23.35	5.30
Nominal exchange rate	44.40	39.67	24.52	26.06	7.47	8.94
Real exchange rate	47.51	40.35	28.23	30.37	13.08	12.48
Share prices	42.03	45.93	24.33	27.73	6.96	6.31
Export prices	48.68	61.78	27.30	31.17	13.12	8.66
Import prices	54.79	41.64	29.21	26.80	12.62	11.36
Nominal GDP	64.16	53.90	35.63	38.70	21.40	15.99
Real GDP	51.04	55.52	25.08	38.04	10.40	7.03
Consumer price index	64.61	69.54	39.21	53.33	17.53	18.29
90 % fractile	49.65	43.95	32.00	26.79	17.85	13.33
95 % fractile	53.12	47.21	34.91	29.69	19.96	15.41
99 % fractile	60.16	54.46	41.07	35.65	24.60	20.04

4 Discussion

Our analysis of price subgroups and information variables demonstrates that disaggregation of the price index might be usefull. With a vector of prices we can raise the question of common trends and we can discuss the determinants of underlying inflation. One consequence of using disaggregated prices is that we can raise the question of movements in relative prices. As an example, share prices do not cointegrate with aggregate CPI, but do cointegrate with disaggregated price subgroups. The richer price dynamics can be observed and more precise questions can be asked.

We found that relative prices cointegrate with broad money, the money market yield index, nominal exchange rates and, most significantly, with import prices. The reported analysis of the price system with import prices indicates that the PPP relation cannot be found in the strict sense and that the past three years are particularly difficult for analysing the behaviour of prices.

Our analysis has indicated that in Finland the driving inflationary forces for consumer prices might include the supply of broad money. Due to the fixed exchange rates and deregulation of capital movements, the central bank was not able to control the supply of money. Rather, the level of money holdings was demand determined during the first 20 years of the observation period. Under certain assumptions, the common trend analysis supports the outcome of the cointegration analysis (reported in table 2) in the sense that foreign trade prices also play a central role in Finnish price formation.

4.1 An Attempt to Justify the Existence of Cointegrating Vectors Theoretically

The empirical analysis conducted in the previous section suggests the existence of at least one cointegrating vector between the prices of different subgroups of consumer goods. This conclusion naturally puts us in a position where we have to try to seek further, theoretical arguments supporting our empirical results and possibly offering or even imposing more structure on the cointegrating space where our empirical vectors also lie. In this section we try to provide such theoretical priciples, in the form of a formal model, which will shed some light on the interpretation of the cointegrating vector(s) as reflecting an underlying demand equilibrium in an economy derived from a well specified intertemporal optimization problem of a representative consumer. Since the full model and its solution are presented in the appendix to this paper, we shall be fairly brief here giving only the main parts of the formalism involved.

Think of a representative consumer maximizing a time-separable intertemporal utility function incorporating (trend-)stationary preference shocks. The stochastic utility maximum is acieved by selecting a flow of domestic and foreign (imported) goods subject to the usual flow constraint according to which the sum of the value of the goods (in terms of the next-period price level) and next-period non-human wealth equals the sum of the gross return on (accumulated) wealth and the exogenously given labour income. In the appendix we have restricted ourselves to the case of two domestic goods and one foreign good, but this does not mean we are sacrifying generality. The first order conditions for this programme are

$$\begin{aligned} \lambda_t P_{1t} &= u_1 \\ \lambda_t P_{2t} &= u_2 \\ \lambda_t P_{ft} &= u_f \\ \lambda_t &= \beta \mathrm{E}_t[\lambda_{t+1}(1+r_{t+1})], \end{aligned} \tag{5}$$

where P_1, P_2 and P_f are the real prices of domestic and foreign goods respectively, u_i denotes the partial derivative of the period utility fuction with respect to i, $(i = 1, 2, f)$, λ is the Lagrangean for the budget flow constraint and r the real rate of interest (from period t to $t + 1$).

Given a strongly separable period utility function of the CRRA form:

$$u(C_{1t}, C_{2t}, C_{ft}) = \frac{D_{1t}C_{1t}^{1-\alpha}}{1-\alpha} + \frac{D_{2t}C_{2t}^{1-\gamma}}{1-\gamma} + \frac{D_{ft}C_{ft}^{1-\eta}}{1-\eta},$$

where C_i denotes consumption of good i and the Ds are the (trend-)stationary preference shocks alluded to above, the F.O.C. in (5) can be written as

$$\begin{aligned} \lambda_t P_{1t} &= D_{1t}C_{1t}^{-\alpha} \\ \lambda_t P_{2t} &= D_{2t}C_{2t}^{-\gamma} \\ \lambda_t P_{ft} &= D_{ft}C_{ft}^{-\eta} \end{aligned} \tag{6}$$

Notice that according to equation (5), the marginal utility of wealth, λ_t, is approximately a martingale process if the real interest rate, r_t, is constant (and equal to the risk-free rate)[11]. Furthermore, if the variance in forecasting the marginal utility of wealth, λ_t, is small, $\log \lambda_t$ itself is well approximated by the following random walk:

$$\log \lambda_t = (\sigma - r) + \log \lambda_{t-1} + (\lambda_t - \mathrm{E}_{t-1}\lambda_t) / \mathrm{E}_{t-1}\lambda_t,$$

where σ is the pure rate of time preference, which, if not equal to the (constant) real interest rate is responsible for a linear trend in the marginal utility of wealth. Hence, by taking logs of (6), we conclude that the equations for log-consumption of the goods share a common trend, and that this trend can be identified with the marginal utility of wealth, $\log \lambda_t$ (cf. Clarida 1991, p. 5).

[11]The fact that the shadow price or, here, marginal utility of wealth, follows a martingale process is a typical outcome of intertemporal optimization problems and signifies the importance of matingales in economics. Furthermore Hall (1978), in deriving the stochastic implications of the permanent income hypothesis, also pointed out the approximate martingale property of marginal utility (of consumption).

The approximate random walk behaviour of the marginal utility of wealth and permanent income hypothesis has some interesting implications concerning cointegration of the consumption of the goods involved. First of all, although the three consumption variables share a common trend, they are not necessarily cointegrated; in fact, after straightforward calculation, (6) gives

$$c_{ft} - [(\alpha/\eta) + 1]c_{1t} + (\gamma/\alpha)c_{2t} = [(\alpha/\eta) + 1]p_{1t}/\alpha - p_{ft}/\eta + p_{2t}/\alpha + \xi_t,$$

where $\xi_t = d_{ft}/\eta + d_{2t} - [(\alpha/\eta) + 1]d_{1t}/\alpha \sim \mathrm{I}(0)$. Therefore, cointegration of the logs of the consumption of the goods results if the corresponding (real) prices are cointegrated! A further implication of the model presented here comes out by noting that not only do the three goods share the common trend of marginal utility of wealth, but every pair of goods constitutes a system driven by the common trend. Moreover, each pair of goods is cointegrated if the corresponding pair of (real) prices is cointegrated. Hence, our system of three consumer goods may involve up to $3 - 1 = 2$ cointegration vectors or that the dimension of the cointegration space is two, being generated by eg the price of the imported good cointegrating with the price of each domestic good. All other "cointegration relationships" or, more precisely, stationary relationships follow from the logic of taking linear combinations of I(0) variables[12]. A generalization to k (≥ 4) goods immediately suggests itself; a system of k consumption goods may incorporate up to $k - 1$ cointegration relationships corresponding to $k - 1$ cointegration relationships between the prices of these goods[13].

So in principle the model presented in the appendix is potentially useful in interpreting our data in the sense that according to it cointegration of the prices (of the subgroups of consumer goods) may actually be an indication of cointegration of the underlying consumer goods and that the system of k price series may, quite conceivably, incorporate more than just one cointegration relationship; in fact the dimension of the cointegration space may be as high as $k-1$. Now, given this basic microeconomic set up, the next question or problem we want to explore relates to what we call the macroeconomic information

[12]ie a linear combination of I(0) variables is I(0). Hence, if, for example, $ap_f + bp_1$ and $cp_f + dp_2$ are I(0), so is $\beta(ap_f + bp_1) + \phi(cp_f + dp_2)$. But this last linear combination is $(\beta a + \phi c)p_f + \beta bp_1 + \phi dp_d$. Thus, the triple is cointegrated if the relevant pairs are, and so on. Note further that the same linear combination of prices that eliminates the common trend of marginal utility of wealth also eliminates the price level, which is used to deflate the nominal prices of the various goods.

[13]Engsted and Tanggaard (1994) provide similar characterization of the cointegration relationships in the context of the term structure of interest rates; they show that for the expectations hypothesis of the term structure of interest rates to hold in a system of k unit-root yields (of different maturities), there must exist $k - 1$ cointegration relationships between the yields. The dimension of the cointegration space, $k - 1$, corresponds exactly to the idea that there are $k - 1$ pairs of yields one can form, with eg the first component of each pair being the same yield (eg the one for the shortest maturity). Once again other stationary linear combinations of the yields can be generated by combining these stationary pairs.

variables. Specifically, how can we motivate or even justify their presence in our cointegration analysis? The presentation of the model in the appendix suggests an approach to incorporating these information variables into the dynamic analysis of the consumer's intertemporal choice. The approach taken is really a short-cut, abstracting from considerations that may be considered relevant to the problem of a consumer's intertemporal choice, but the approach should be taken as a suggestion for a way of thinking about the interaction of microeconomic decisions and macroeconomic information.

So to begin with, the appendix suggests that macroeconomic information interacts with consumer's choice via the "exchange rate function" $s_t = s(R_t, r_t, gdp_t, m_t, \ldots)$[14], where s denotes the exchange rate (the markka price of foreign currency), while $R_t, r_t, gdp_t, m_t, \ldots$ signify long- and short-term interest rates, output and money supply respectively. Clearly, such an "exchange rate function" can result from a macroeconomic intervention rule followed by the monetary authority in its pursuit of macroeconomic stability (see eg Vilmunen 1992). The emphasis here should be on the idea that such a rule is a *macroeconomic* rule; hence, such a rule abstracts eg from all strategic aspects related to a possible monetary policy game between the private sector and monetary authority. Furthermore, such a rule corresponds to the monetary authority's broad objective of a stable macroeconomic environment and is taken as exogenously given by the consumers when making their decisions on utility maximizing consumption flows, as are the variables affecting the exchange rate via the rule. Hence, the "exchange rate function" is clearly a short-cut; it is a reduced form or summary of the macroeconomic environment facing the consumer. While simple, we find it a useful way of organizing our thoughts about possible channels through which macroeconomic information interacts with consumer's decisions and, hence, disaggregated prices. We must, however, emphasize the informal nature of our approach; we do not estimate an intervention rule supposedly followed by the monetary authority. No structure is imposed on the s-function, except (local) linearity so as to provide consistency with our linear estimation and testing methodology (VAR).

[14]Which, of course, affects the domestic price of imported goods, since $p_{ft} = q_{ft}s_t$, where q denotes the price of importables in international markets.

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A Permanent Income Model and Common Price Trends

The general idea is to use the permanent income model to derive cointegration restrictions on an import demand function and combine this representative agent structure with macroeconomic information, for example, on import prices, exchange rates, nominal GDP, monetary aggregates, etc. In general, we can achieve this by postulating a money *supply* or *monetary policy* rule, according to which exchange rates react to macroeconomic information through the policy rule. Thus the import prices in FIM, Q_t, can be written

$$Q_t = Q_t^* S_t = Q_t^* S_t(GDP, M, i, \ldots),$$

where Q_t^* is the exogenous world price of imports and S_t is the exchange rate function.

The basic problem is that of maximizing intertemporal preferences over the life of our consumer subject to a budget flow constraint:

$$\begin{aligned} \max \quad & E_0 \sum_{j=0}^{T} \beta^t u(C_{1t}, C_{2t}, C_{ft}) \\ \text{s.t.} \quad & P_{1t} \; C_{1t} + P_{2t}C_{2t} + P_{ft}C_{ft} + A_{t+1} = (1+r_t)A_t + Y_t, \end{aligned}$$

where

C_{1t}, C_{2t}	are home goods,
C_{ft}	foreign good,
P_{1t}, P_{2t}	real prices for home goods,
P_{ft}	real prices of the foreign good,
A_t	real (non human) wealth (capital),
r_t	real rate of interest,
Y_t	exogeneous flow of real labour income.

The first order condition yields

$$\lambda_t P_{1t} = u_1 \tag{7}$$
$$\lambda_t P_{2t} = u_2 \tag{8}$$
$$\lambda_t P_{ft} = u_f \tag{9}$$
$$\lambda_t = \beta E_t \left[\lambda_{t+1}(1+r_{t+1})\right]. \tag{10}$$

All the capital is consumed, ie $A_{t+1} = 0$. When we assume strongly separable preferences of the CRRA form,

$$u(C_{1t}, C_{2t}, C_{ft}) = \frac{D_{1t}C_{1t}^{1-\alpha}}{1-\alpha} + \frac{D_{2t}C_{2t}^{1-\gamma}}{1-\gamma} + \frac{D_{ft}C_{ft}^{1-\eta}}{1-\eta}, \tag{11}$$

where D_{1t}, D_{2t} and D_{ft} are (stationary) preference shocks, the equations (7) – (9) can be written as follows

$$\begin{aligned}
\lambda_t P_{1t} &= D_{1t} C_{1t}^{-\alpha} \quad (12)\\
\lambda_t P_{2t} &= D_{2t} C_{2t}^{-\gamma} \quad (13)\\
\lambda_t P_{ft} &= D_{ft} C_{ft}^{-\eta}. \quad (14)
\end{aligned}$$

A well know property of the permanent income hypothesis is that the marginal utility (MU) of consumption follows a martingale process (if we suppose that r_t is a constant risk-free rate of return). Above, the marginal utility of wealth, λ_t, follows a martingale, as do the marginal utilities of consuming the goods, once these marginal utilities are "deflated" by the relevant prices. Now, if the variance in *forecasting* the marginal utility of wealth, λ_t, is small, $\log \lambda_t$ itself is well approximated[15] by the following random walk:

$$\log \lambda_t \approx (\sigma - r) + \log \lambda_{t-1} + (\lambda_t - E_{t-1}\lambda_t)/E_{t-1}\lambda_t, \quad (15)$$

where real rate r is assumed to be constant and σ in $\beta = (1+\sigma)^{-1}$ is pure time preference. Moreover, note that (12) – (14) imply

$$\begin{aligned}
c_{ft} &= \frac{d_{ft}}{\eta} - \frac{1}{\eta} p_{ft} - \frac{1}{\eta} \log \lambda_t \quad (16)\\
c_{1t} &= \frac{d_{1t}}{\alpha} - \frac{1}{\alpha} p_{1t} - \frac{1}{\alpha} \log \lambda_t \quad (17)\\
c_{2t} &= \frac{d_{2t}}{\gamma} - \frac{1}{\gamma} p_{2t} - \frac{1}{\gamma} \log \lambda_t \quad (18)
\end{aligned}$$

The permanent income hypothesis implies that the logs of C_{1t}, C_{2t} and C_{ft} share a *common stochastic trend* and that this stochastic trend can be identified with the marginal utility of wealth, $\log \lambda_t$! Furthermore, the logs of the consumption of home goods and the foreign good are cointegrated if the log of their prices, p_{1t}, p_{2t} and q_t are cointegrated. Since prices, p_{it} $(i = 1, 2, f)$, are relative prices ie deflated eg the by consumer price index, $p_{it} = \log(\frac{\bar{P}_{it}}{P_t})$ $(i = 1, 2, f)$, the price deflator is also a common trend of prices.

[15]Note that $\log(1+r) \approx r$ and utilize equation (10).

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