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Large T and small N: A three-step approach to the identification of cointegrating relationships in time series models with a small cross-sectional dimension

by

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Large T and small N: A three-step approach to the identification of cointegrating relationships in time series models with a small cross-sectional dimension.*‡

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November 2004

Abstract

This paper addresses cointegration in small cross-sectional panel data models. In addition to dealing with cointegrating relationships within the cross-sectional dimension, the paper explicitly addresses the issue of cointegration between cross-sections. The approach is based upon a well-known distributional result for the trace test when some of the cointegrating vectors are a priori known, and advocates a three-step procedure for the identification of the cointegrating space when dealing with two-dimensional data. The first step of this procedure utilizes traditional techniques to identify

*I want particularly to thank Henrik Hansen in helping me out with the implementation of known cointegrating restrictions in Cats in Rats. In addition I am grateful for comments by Søren Johansen and Andreas Beyer and participants at the conference on the monetary transmission mechanism at Schaeffergaarden in Copenhagen.

†The analyses have been undertaken by using a combination of CATS in RATS (Hansen and Juselius (1995)) and PcFiml 9.20 (Doornik and Hendry (1999)). The I(2) tests have been undertaken by using Clara Jørgensen's I(2) procedure in Cats in Rats.

‡This paper forms part of my PhD thesis at the European University Institute, Florence. However, a substantial part of the paper is based on research undertaken in Norges Bank.

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the long-run relationships within each cross-sectional unit separately. In the second step these first step relationships are then treated as known when searching for potential long run relationships between units in a joint analysis comprising the whole cross-sectional dimension. The third step of the procedure then finally reestimate all free parameters of the identified long-run structure to get rid of a potential simultaneity bias as a result of a non-diagonal covariance matrix.

Identification of the long-run structures of Norwegian exports and international interest rate relationships are used as examples. Norwegian mainland exports have here been divided into two cross-sectional units; the traditional goods sector and the service sector. While in the study of international interest rate relationships the two sectors investigated are Germany and the US. The examples are used to address the more general issues of the degree of independence in capital markets and in goods markets of small open economies.

Keywords: Cointegration, Panel data, transmission mechanism, monopolistic competition, exports.
JEL: C32,C33,E43,F12,F41

Contents

1. Introduction

In the early 1990's, several studies developed the asymptotic properties and studied finite sample properties of unit-root tests on panel data as both time series and cross-section dimensions grow arbitrarily large [see e.g. Breitung and Meyer (1991), Quah (1994), and not least Levin and Lin (1992,1993)]. Their results showed that by using data varying not only along one dimension, but along two dimensions, the power of the unit root test in most cases increases dramatically against stationary alternatives. In the spirit of Engle and Granger (1987), these tests have recently been further extended to various tests for cointegration in a panel data framework by e.g. Pedroni (1996) and McCoskey and Kao (1998).[1] However, as a means of identifying cointegrating relationships in the multivariate case, with the possible existence of several cointegrating relationships, this method is far from sufficient. Hence there is a need to develop a multivariate system approach along the lines of Johansen (1988). Even though there is a lot of ongoing research aimed at meeting this requirement, a fully general system framework to deal with cointegration in the case of multivariate panel data has to my knowledge still not been developed. Also, even though the problem may have a general solution this most probably will turn out to be totally inadequate as a practical device for undertaking panel data cointegration analysis, as the level of complexity is almost unwieldy already when dealing with a small number of cross-sectional units.

The primary aim of this paper is to offer an easily accessible strategy for dealing with time-series models with a small cross sectional dimension and is therefore written in the spirit of developing a kind of *ad hoc* solution to a case that is less general than the general problem of panel data cointegration. It is based on the result in Horvath and Watson (1995), which gives the asymptotic distribution of the Wald test in vector autoregressive models when some cointegrating vectors are known, and advocates a two-step approach which first identifies the cointegrating relations in each cross-sectional unit separately and then uses these as

[1] The implied increased possibility of identifying cointegrating relationships in this setting has also initiated renewed interest in solving parity puzzles, like that of purchasing power parity (e.g. MacDonald (1996), Frankel and Rose (1996), Pedroni (1997)).

known when analyzing the cross-sections jointly in a second step.[2] The first step can be done in the traditional way by analyzing the cross-sections specific VARS. The second step implies interpreting the estimated cointegrating relations in the first step as representing known cointegrating relations, and then to use the distributional results of Horvath and Watson, as tabulated for the likelihood-ratio cointegration rank test in Paruolo (1999), to determine the cointegrating rank of the full system, given these. The contribution of this paper lies in the use of a result developed originally for a pure time-series model to help with the identification of cointegrating relations in the case where we also have to deal with a cross-sectional dimension. In addition to allowing for heterogeneous long-run cointegration relationships within each unit or sector and cross-sectional dependencies through error-correction terms and short-run effects, this approach explicitly takes into account the possibility of cointegration between sectors. Larsson and Lyhagen (1999) develop a framework where cointegrating relationships are only allowed for within each sector and as such therefore disregards the possibility of long-run cointegrating relationships between sectors. However, in contrast to an earlier paper, Larsson *et al* (1998), they explicitly allow for cross-sectional long-run effects through the potential inclusion of all sector-specific cointegrating relationships as equilibrium-correction terms in the equations of the panel data model. Larsson and Lyhagen also allow simultaneous modelling of the long run relationships within sectors, taking into account possible cross sectional dependencies in the error structure of the model. This suggests a third and final step in our procedure to identify cointegrating relations in time series with a small cross-sectional dimension. Namely after having gone through the two steps suggested earlier, I propose to re-estimate all free parameters of the identified cointegrating relationships in a VAR where the previously determined rank and structure of the cointegrating space are imposed.

The external part of most large-scale econometric models of small open economies has traditionally been modelled as one of monopolistic competition, implying a certain degree of monopoly power in the process of price determination. Within this framework it is therefore appropriate to ask whether the increased degree of openness during the eighties and nineties, have had a significant impact on the possibility to deviate from long-run relative purchasing power parity (RPP) in the process governing external trade prices. Another important issue in the wake of deregulation of capital markets and increased internationalization, is whether

[2] To improve upon readability I will hereafter use the concepts of a cross-sectional unit and a sector interchangeably, hoping that the context brings out the exact meaning.

the possibility of running independent monetary policies in Europe has been considerably weakened during the last decade. And if so, whether this has been accomplished through a stronger dependence on what is going on in international capital markets.

To provide examples of the suggested procedure and to analyze the political issues addressed above, this paper undertakes two independent analyses. The first seeks to identify the long-run structure of Norwegian exports between the first quarter of 1980 and the last quarter of 1998. To provide a cross-sectional time series data set, Norwegian mainland exports are divided into two sectors, the traditional goods sector and the service sector, respectively. The implications with regard to the identification of a RPP relationship are then compared with the results in Hammersland (1996), which based on an aggregate model of Norwegian exports, is not capable of identifying a RPP relationship and reveals significant signs of monopolistic power in the determination of prices over the period 1966 (4) to 1992 (4). The other study is a study of US and German interest rates and seeks to reveal the degree of European autonomy through identification of short- and long-term interest rate relationships over the period 1990 (1) to 1997 (12). The two sectors are naturally given by the two countries and the results of the analysis are compared with the results in Hammersland and Vikøren (1997) and Hammersland(2002a).

The reminder of the paper is organized as follows. Based on Horvath and Watson (1995), Section 2 introduces the model used to analyze the two-dimensional data set and gives a brief motivation for the choice of tables to be used when dealing with identification of cointegrating rank in the case of known cointegrating relationships. Section 3 then contains the two examples of using my suggested three-step procedure on two actual "panel" data sets. Before presenting the results, however, subsection 3.1.1 deduces theoretical hypotheses on long-term relations based upon the theory of monopolistic competition and Armington demand theory, extensively reviewed in Hammersland(2002d). Based on the theories of uncovered interest rate parity, UIP, and the expectation theory of the term structure, Section 3.2.1 does the same for the interest rate study. The results of the econometric analyses where we first identify the cointegrating relationships in the two sectors separately for finally ending up with a joint analysis of the full model, are then given for the two examples in 3.1.3 and 3.2.3, respectively. Section 4 concludes.

2. The Model

The general autoregressive I(1) model is given by:

$$\Delta X_t = a\beta' X_{t-1} + (\Upsilon, \mu) \begin{pmatrix} Z_t \\ d_t \end{pmatrix} + \epsilon_t \ , \tag{2.1}$$

where X_t and ϵ_t are both $p \times 1$ vectors, $Z_t = (\Delta X'_{t-1}, ..., \Delta X'_{t-k+1})'$ is $p(k-1) \times 1$, ϵ_t is assumed to be i.i.d.$N(0, \Omega)$ and d_t is a $q \times 1$ vector of deterministic terms like a constant term, trend and seasonal dummies. $\Upsilon = (\Gamma_1, ..., \Gamma_{k-1})$ is a $p \times p(k-1)$ matrix, μ $(p \times q)$ and α and β are both $p \times r$ matrices assumed to be of full rank, r, such that the $I(1)$ condition of $\alpha'_\perp(I_p - \sum_{i=1}^{k-1} \Gamma_i)\beta_\perp$ having full rank, $p - r$, is fulfilled when assuming that all the roots of the characteristic polynomial of X_t lie at one or outside the unit circle. For our purpose X_t consists of sector specific variables as well as variables that do not vary across the sectors. Thus, in the case of two sectors $X_t = \left(X'_{1,t}, X'_{2,t}, X'_{3,t}\right)'$, where $X_{i,t} = (X_{i1,t}, ..., X_{iN_i,t})'$ represents the N_i numbers of sector specific variables in sector i, $i = 1, 2$, and $X_{3,t} = (X_{31,t}, ..., X_{3N_3,t})'$ represents the N_3 numbers of common variables. We are going to look at the case where β can be partitioned into two submatrices, $\beta_1 = b$ and β_2, of dimensions $p \times s$ and $p \times m$ respectively. The first set of cointegrating vectors, b, represents the s a priori "known" cointegrating relationships that we are getting to "know" from a preliminary cointegration analysis undertaken at the sectoral level, while β_2 represents the $m = r - s$ remaining cointegrating vectors to be identified using the whole information set given the "known" relationships identified in the first step. r represents the total number of cointegrating vectors in the two-dimensional data set. Representing the estimated cointegrating vectors at the sectoral level by $\widehat{\beta}_1 = \left(\left(\widehat{\beta}'_{11}, 0', \widehat{\beta}'_{13}\right)', \left(0', \widehat{\beta}'_{22}, \widehat{\beta}'_{23}\right)'\right)$, this means that $\widehat{\beta}_1 = b$ and the level term in (2.1) can be given the equivalent representation of

$$\begin{aligned} \alpha\beta' X_{t-1} &= (\alpha_1, \alpha_2) \begin{pmatrix} b' \\ \beta'_2 \end{pmatrix} X_{t-1} = (\alpha_1, \alpha_2) \begin{pmatrix} \widehat{\beta}'_1 \\ \beta'_2 \end{pmatrix} X_{t-1} \\ &= (\alpha_1^*, \alpha_2^*, \alpha_2) \begin{pmatrix} \widehat{\beta}'_{11} & 0' & \widehat{\beta}'_{13} \\ 0' & \widehat{\beta}'_{22} & \widehat{\beta}'_{23} \\ \beta'_{31} & \beta'_{32} & \beta'_{33} \end{pmatrix} \begin{bmatrix} X_{1,.t-1} \\ X_{2,t-1} \\ X_{3,t-1} \end{bmatrix}, \end{aligned} \tag{2.2}$$

where we have partitioned the remaining cointegrating vectors to be estimated in the second step, β_2, in conformity with the partitioning of the variable vector, X_t. The argument for treating some cointegrating vectors as known when estimating the level matrix Π of the VAR, even though they strictly speaking have been estimated in a preliminary step, hinges on the super consistency property of the cointegrating vectors. This point may be clarified by looking at the asymptotic distribution of $T^{\frac{1}{2}}\left(\widehat{\Pi}-\Pi\right)=T^{\frac{1}{2}}\left(\widehat{\alpha}\widehat{\beta}'-\alpha\beta'\right)=\left(T^{\frac{1}{2}}\left(\widehat{\alpha}-\alpha\right)\right)\widehat{\beta}'+\alpha\left(T^{\frac{1}{2}}\left(\widehat{\beta}-\beta\right)'\right)$. While $\widehat{\beta}$ is superconsistent in the sense that $\widehat{\beta}-\beta\in o_p\left(T^{-\frac{1}{2}}\right)$, $\widehat{\alpha}$ converges to α at rate $T^{\frac{1}{2}}$, implying that $\widehat{\alpha}-\alpha\in O_p\left(T^{-\frac{1}{2}}\right)$. Thus the term $T^{\frac{1}{2}}\left(\widehat{\beta}-\beta\right)$ converges to zero while the first one,$\left(T^{\frac{1}{2}}\left(\widehat{\alpha}-\alpha\right)\right)\widehat{\beta}'$, converges to $N\left(0,\Omega\otimes\beta\Sigma_{\beta\beta}^{-1}\beta'\right)$, where $\Omega\otimes\Sigma_{\beta\beta}^{-1}$ is the variance of $T^{1/2}(\hat{\alpha}-\alpha)$, and the scaled distribution of $\widehat{\Pi}$ is asymptotically independent of the estimated cointegrating vectors. However this argument does not explain why we do not estimate all cointegrating vectors simultaneously in a pooled analysis at the outset. This is more an argument of feasibility as the problem of identifying all cointegrating vectors simultaneously becomes intractable when the possibility set increases. To reduce the dimension of the estimated CI space will therefore serve to enhance the interpretability as well as the identifiability of the system of cointegrating vectors. In the following I will resort to the simpler notation of denoting the estimated cointegrating vectors as b and the remaining unknown ones as β_2, keeping in mind the partitioning in (2.2) when interpreting the significance of the vectors.

Our analyses in the next section will be confined to the case where the deterministic term μd_t is decomposed as $\mu_1 d_{1t}+\mu_2 d_{2t}$ and μ_1 is restricted to lie in the α space such that $\mu_1=\alpha\kappa$, where $\kappa=\begin{pmatrix} k \\ \kappa_2 \end{pmatrix}$ is a $r=(s+m)\times q$ matrix. To accommodate these changes (2.1) must be transformed according to:

$$\Delta X_t=(\alpha_1,\alpha_2)\begin{pmatrix} b_1' \\ \beta_2' \end{pmatrix}X_{t-1}+\left(\Upsilon,(\alpha_1,\alpha_2)\begin{pmatrix} k \\ \kappa_2 \end{pmatrix},\mu_2\right)\begin{pmatrix} Z_t \\ d_{1t} \\ d_{2t} \end{pmatrix}+\epsilon_t$$

where α and κ have been decomposed conformably with the partitioning of $\beta=(b,\beta_2)$. This expression may equivalently be expressed as:

$$\Delta X_t=\alpha_2(\beta_2',\kappa_2)\begin{pmatrix} X_{t-1} \\ d_{1t} \end{pmatrix}+(\alpha_1,\Upsilon,\mu_2)\begin{pmatrix} b'X_{t-1}+kd_{1t} \\ Z_t \\ d_{2t} \end{pmatrix}+\epsilon_t \tag{2.3}$$

It is this set up of the model we are going to use in the determination of $m = r - s$, the number of cointegrating vectors, $b^* = (b', k)'$, beyond the known number of relationships following from the sectoral analysis, $\beta_2^{*\prime} = (\beta_2', \kappa_2)'$.

In Section 3, when identifying the cointegrating relationships in the first example, model (2.3) in addition to including unrestricted centralized seasonal dummies, is specified with a trend restricted to lie in the cointegration space and a non-restricted constant term, implying that $d_{1t} = t$ and $d_{2t} = (1, S_1, S_2, S_3)'$.[3] Therefore, the most appropriate critical values to use in identifying the cointegration rank are given by Table 5 in Paruolo (1999). In the second example where I study interest rate relationships, I have deliberately neglected a trend term and the constant term has been restricted to lie in the space spanned by the loading matrix α_2. In model (2.3) this is equivalent to $d_{1t} = 1$ and $d_{2t} = 0$. This implies that the correct critical values to use are given by Table 3 in Paruolo (1999). It is otherwise worth noting that in neither case has it been necessary to fall back on measures to improve diagnostics.[4]

3. Identification of cointegrating relations using times series data with a small cross-sectional dimension: two examples

This section provides two illustrative examples of how to use the advocated three-step procedure of this paper to analyze real data. In both cases I analyze data with a two dimensional structure where the cross sectional dimension is equal to two.[5] The first seeks to identify the structure of exports in small open economies by looking at Norwegian data. To be able to apply the suggested procedure, Norwegian mainland exports have been divided into two sectors, the traditional goods sector and the service sector, respectively. The second example concerns identification of international and domestic interest rate relationships and looks particularly at linkages between European and US long- and short-term interest rates as well as the degree of domestic control over the long end of the yield curve as reflected by a term structure relationship between short- and long-term domestic interest rates.

[3] The S_i's are the three centred seasonal dummies.

[4] The only exception is the use of seasonal dummies in the case of the export study.

[5] I have deliberately avoided using the term "panel data" as this concept usually is confined to the case of a large cross sectional dimension.

3.1. Example 1: Modelling of export volumes and export prices in a small open Economy: The Norwegian case.

As alluded to in the abstract, the analyses are partly motivated from the perspective of identifying the degree of independence in capital markets and in goods markets of small open economies. In the case of exports, a model of monopolistic competition is well suited for this purpose as it takes into account the possibility of monopolistic power in the process governing the determination of quantity and prices. Before presenting the results, therefore, a brief review of theory and its implications with regard to cointegration will be given.

3.1.1. Monopolistic competition and its implications for cointegration

Most models for the determination of export volumes are pure demand relationships based on Armington's theory of demand distinguished by place of production (Armington (1968)). They are often explained by models of monopolistic competition (Bruno (1979)) in which export prices are determined ex ante[6] and export volumes for fixed prices ex post. In doing so, it is common practice to assume a constant price elasticity in demand and constant returns to scale. However, when looking at the export price and export volume determination simultaneously, these assumptions may be mutually inconsistent with data and the possibility of developing a stable representation in the shape of an econometric model of the information contained in these. For example, under monopolistic competition these assumptions imply that there is no channel through which demand may affect prices. As these effects turn out to be significantly estimated in most econometric works, export price relationships are often implicitly based on either an assumption of decreasing returns to scale or that a non-constant price elasticity of demand creates cyclical movements in the mark-up. A non constant elasticity of demand will on the other hand necessarily imply an unstable Armington based model of the export volume which makes it inappropriate to assume production processes with constant returns to scale when combining Armington's theory with the theory of monopolistic competition. In this paper I have chosen to assume decreasing returns to scale, thus enabling us to model both price and volume determination under a consistent set of assumptions. For a mathematical exposition of both theories the reader is referred to Hammersland (2002d).

[6] The *ex ante* decision refers to a plan made before having complete knowledge of all variables affecting the decision-making, implying that the decision must be based on their expectations. The *ex post* decision, however, is made on the basis of complete knowledge of all variables.

An ex ante, ex post approach requires export prices to be completely fixed according to the ex ante plan while the export volume is allowed to depart from the same plan ex post. The existence of long-term contracts, advertisements, price lists etc. may motivate such a commitment of export prices to the plan. Thus, a representative producer will exercise price-taking behavior ex post and one may be faced with one of two possible situations. In the first case, the consumer will be rationed on the product market and the production will be determined by the price-taking level of production. In the second case, it is the producer that will be rationed on the product market and the export volume will be determined exclusively by real demand. Thus, we are in a situation in which prices are determined ex ante by the behavior of a monopolist facing decreasing returns to scale while at the same time the level of exports is determined exclusively by ex post demand.

The export price equation will thus be defined by the first order condition, price equal to marginal costs multiplied by a mark-up factor greater than one. In practice export prices, PA, may therefore be modelled as a log linear function of unit labor costs, WC/Y,[7] world market prices, PW and foreign real income R.

$$\ln(PA_t) = c + \phi(\ln(WC_t) - \ln(Y_t)) + (1-\phi)\ln(PW_t) + \rho\ln(R_t) + \epsilon_{1t} \tag{3.1}$$

The parameter ϕ is the partial elasticity of export prices to unit labor costs. From (3.1) it appears that the export prices are homogenous of degree one in unit labor costs and world market prices. ϵ_{1t} is a stochastic disturbance term for the export price equation.

I follow Armington(1968) and assume that demand is specific to the producer. Thus, the demand for exports, denoted A, may be specified as a log linear function of the foreign real income, R, and the relative price given by the ratio of export prices to world market prices.

$$\ln(A_t) = \mu + \beta\ln(R_t) - \sigma(\ln(PA_t) - \ln(PW_t)) + \epsilon_{2t} \tag{3.2}$$

The producers of small open economies generally have a very small market share, implying that the parameter of relative prices can be interpreted both as a relative price elasticity with regard to export demand and as the elasticity of substitution. This can be shown mathematically (see Hammersland (2002d)), but it also has some intuitive appeal since the income effect of an increase in the export

[7] WC and Y represents wage costs per man-hour and output per man-hour, respectively.

prices of a small open economy on foreign demand will be virtually negligible. The price elasticity expresses thus the percentage change in the ratio of the goods produced in the small open economy to foreign goods and an elasticity less than zero will imply a decreasing market share in real terms with regard to relative price changes. β bigger than or less than one will indicate whether the economy's market share is increasing or not when facing a growing world market. ϵ_{2t} is a stochastic disturbance term in the export volume equation. In both equations all prices are given in the currency of the small open economy.

Economic theory contributes in an important way to our empirical analysis by providing suggestions to possible explanatory variables and also to what kind of basic relationships we may expect to find between them. The interpretation of such relationships will however typically be as long-run relationships. Given the non-stationary properties of many of the relevant macro economic time series, such long-run relationships will be associated with the statistical concept of cointegration, which has the implication that an empirical long-run relation exists between the variables. To empirically substantiate economic theory, we will therefore have to require that the results of the cointegration analysis are consistent with theory. The cointegration analysis in this section is therefore based on the export-price and -volume equations in (3.1) and (3.2), respectively. Theory consistency requires that there are at least two cointegrating relationships and that both disturbance terms in (3.1) and (3.2) are I(0). If we find support for two and only two cointegrating relationships, this will especially require that export prices, unit labor costs and world market prices form a cointegrating linear combination, possibly with an additional demand effect from abroad. On the other hand we would also expect the export volume to be cointegrated with a linear combination of foreign real income and the relative price of export prices to world market prices.

To further develop the implications theory consistency may have with regard to cointegration, (3.1) may be reformulated as

$$\ln(PA_t) - \ln(PW_t) = c + \rho \ln(R_t) + \phi(\ln(WC_t) - \ln(Y_t) - \ln(PW_t)) + \epsilon_{1t}$$

First, let us assume that the logarithm of the ratio of unit labor costs to world market prices cointegrates. As theory consistency necessarily implies that $\epsilon_{1t} \sim I(0)$, this will then either imply RPP or for ρ different from 0 and $R \sim I(1)$, that the real exchange rate cointegrates with foreign real income. For ϕ different from 0, we see that the implication may also go in the other direction, as RPP in the case of $\rho = 0$ or $R \sim I(0)$, then would imply constant wage or profit share in the external sector. Further, looking at (3.2), we have that this, under the assumption of β different from 0 and $R \sim I(1)$, implies that real foreign income must cointegrate with the volume of exports.

Evidently, the imposition of theoretical restrictions still leaves us with lots of degrees of freedom to identify theoretically consistent long-run structures. A more heuristic interpretation with regard to what is consistent with regard to theory may in addition even further increase the possibility set, examples in this respect being removal of homogeneity restrictions, exclusion of variables etc. In the next section these issues are further investigated.

3.1.2. Data and time series properties

Before presenting the results of the cointegration analysis, I will first draw attention to a brief description of the empirical data set, herein undertaking a preliminary analysis with regard to time series properties of the individual data. Together with graphs of levels and first differences of all variables in the information set, all empirical results of these tests for stationarity, except for the I(2) analysis undertaken below, are placed in the appendix section of this paper.

The econometric analysis is based on quarterly seasonally unadjusted data over the period 1979: 2 to 1998: 2. The data set consists of observations on the following empirical proxies of the theoretical quantities:[8]

[8]From now on, I will stick to the convention of using small letters for variable names when in fact the variables are logarithmic transformations of the original series, the only exception being the foreign real demand indicator where capital R indicates the logarithm of foreign real demand.

$a1$	Export volume index of traditional goods
$a2$	Export volume index of services
$pa1$	Export price deflator on traditional goods
$pa2$	Export price deflator on services
pw	World market price index
R	Foreign demand indicator
ulc	Unit labor cost indicator

Before examining the long-run relationships between the variables, it is useful to first determine the orders of integration of the individual time series in the information set. In the appendix, I therefore first present the results of testing for stationarity within the multivariate framework based on the methodology developed by Johansen for estimation and identification of cointegrating relationships (Johansen (1988), (1995)). This test is conditional on the number of cointegrating relationships and differs in a very important respect from univariate Dickey-Fuller tests by testing the null of stationarity against a non-stationary alternative. These system-tests are superior to univariate testing for stationarity of individual time series. However, due to a generic bias towards these tests among time series econometricians, I have chosen also to present the results of Augmented Dickey Fuller tests. To avoid the problem of nuisance parameters in the DGP all these tests are made similar, implying the joint appearance of a trend and a constant term in the specification of the autoregressive equation. To get rid of as many anomalies as possible, I have also included seasonal dummies. Testing the null of $I(2)$ vs. the alternative of $I(1)$, however, has been done by only including a constant term in the equation to avoid the problem of having to deal with a quadratic trend under the alternative. A common problem with all these tests is the rather asymmetric treatment of the null and alternative concerning the status of nuisance parameters. However, this problem can easily be dealt with by undertaking a joint test of both the lagged level variable and the trend, and using Table 4.5 in Banerjee, Dolado, Galbraith and Hendry (1993), which gives the simulated critical values in finite samples for these F-type tests. However, to be able fully to address the issue of higher order integration, I have instead chosen to undertake a full analysis of the cointegrating indices based on the two-step I(2) procedure developed by Johansen (1995b).

The multivariate test statistics strongly suggest the rejection of the null of stationarity for most variables. However, it is worth noting that with regard to world market prices and export prices in the traditional goods sector we cannot reject the null of stationarity at conventional levels of significance. However, the overwhelmingly strong support for treating all variables in the information set as non-stationary I(1) variables based on univariate testing together with the fact that the significance probabilities of the multiple test statistics for these two variables are close to a nominal level of five per cent, indicates that we probably are not going to make too a serious mistake by treating prices as I(1). This is also indicated by the I(2) tests in Table 3.1 below, even though strictly speaking there is some evidence of an I(2) trend in the sample. These tests of I(2)-ness have been carried out by specifying a seven dimensional VAR of order three, where a drift term has been restricted to the cointegrating space and the constant restricted not to induce quadratic trends in the processes.[9] The test procedure starts from top left testing the null of seven common I(2) trends versus less than or equal to full rank and continues to the right until one reaches the last column which is the ordinary test of seven I(1) trends versus more than or equal to nil common trends. In the case where one rejects all nulls in the first row of seven common trends, one continues this stepwise testing from left towards right by moving down to the next row of six common trends. The number of cointegrating vectors, I(1) and I(2) trends are given by the first null that one cannot reject. In Table 3.1 below, this process of rejection does not end until the number of common trends are equal to three and the number of I(1) trends are identified to two which implies that the number of common I(2) trends are equal to $p-r-s$ = 7 - 4 - 2 = 1. However as the critical ten per cent level is equal to 49.69, the statistic is hardly significant to a level of ten percent. This could indicate that the cointegration indices are given by $r=4$, $s=3$ and $p-r-s=0$. If so, there are no common I(2) trends and the analysis can be undertaken by ordinary reduced rank analysis for times series integrated of order one. This would be in accordance with the conclusions made on behalf of multivariate testing and the univariate Dickey Fuller tests referred to above. Looking carefully at Table 3.1 it is also worth noting that to a level of slightly above ten percent we are in fact able to reject all combined nulls of more than one common trend and that some of these are I(2). However, the test of more

[9] To be able to fit the table in the text, the numbers have been rounded off to their nearest one decimal representation.

Table 3.1: The trace test of cointegrating indices

p-r	r	$S_{r,s}$							Q_r
7	0	**553.4**	**465.7**	**391.9**	**322**	**288.3**	**256.6**	**230.1**	**213.3**
		351.6	311.2	274.0	241.2	211.6	186.1	164.6	146.8
6	1		**445.0**	**359.2**	**286.7**	**224.9**	**191.2**	**166.2**	**149.8**
			269.2	233.8	202.8	174.9	151.3	130.9	115.4
5	2			**340.0**	**259.7**	**187.9**	**152.1**	**125.0**	**104.7**
				198.2	167.9	142.2	119.8	101.5	87.2
4	3				**237.2**	**165.1**	**104.4**	**78.1**	**66.3**
					137.0	113.0	92.2	75.3	62.8
3	4					**138.2**	**72.2**	**49.6***	**38.2***
						86.7	68.2	53.2	42.7
2	5						**58.2**	**36.8**	**21.2**
							47.6	34.4	25.4
1	6							**21.3**	**6.00***
								19.9	12.5
p-r-s		7	6	5	4	3	2	1	

[1] Table 3.1 is based upon a seven dimensional VAR of order three for the variables $a1$, $a2$, $pa1$, $pa2$, pw, ulc and R. A drift term has been restricted to lie in the cointegrating space and a constant is included such that it does not induce a quadratic trend in the process.
[2] The figure in italics under each test statistic is the 95 per cent fractile as tabulated by Paruolo(1996). The three preferred outcomes discussed in the text are marked with stars.

than or equal to one common I(1) trends versus less than or equal to full rank does not reject. This indicates that there could be as many as six cointegrating vectors and no I(2) trends in the information set.[10]

[10] This last hypothesis is also given some support by sectorial identification of the cointegration indices, as both sectors separately indicate the existence of three cointegrating vectors and no I(2) trends.

Table 3.2: Rank tests for the trading sector

System: a1, pa1, pw, R, ulc.
Deterministic part: Unrestricted constant, centered seasonals and restricted Trend.
VAR order: 3. Effective sample period: 1980 (1)-1998 (2).

Eigenvalues of Π: 0.9589 0.8244 0.7379 0.5543 0.4494

Max Eigenvalue Tests				Trace Eigenvalue Tests			
Null	Alternative	Statistics	95%	Null	Alternative	Statistics	95%
$r=0$	$r \leq 1$	59.18**	37.5	$r=0$	r≤5	142.7**	87.3
$r \leq 1$	$r \leq 2$	43.66**	31.5	$r \leq 1$	r≤5	83.55**	63.0
$r \leq 2$	$r \leq 3$	22.49	25.5	$r \leq 2$	r≤5	39.89	42.4
$r \leq 3$	$r \leq 4$	14.29	19.0	$r \leq 3$	r≤5	17.4	25.3
$r \leq 4$	$r \leq 5$	3.107	12.3	$r \leq 4$	r≤5	3.107	12.3

3.1.3. Cointegration Analysis

In estimating the two sectors the effective sample used for estimation has been from 1980 (1) to 1998 (2). In both sectors we have started out with a five dimensional VAR of order three. Both econometric models include a restricted trend term to avoid problems with regard to nuisance parameters when testing for the cointegration rank. Furthermore, constant terms and seasonal dummies have not been restricted to lie within the α-space.

The trading sector First, I want to draw attention to the diagnostics of the VAR for the traded sector given in Table A.3 of the appendix. Except for some hardly significant signs of autocorrelation and conditional heteroscedasticity in the processes governing export prices and the export volume, all single equation and system diagnostics are fine. Also looking at the recursively estimated Chow tests in Figure B.9 of appendix B, does not reveal any signs of parameter instability whatsoever. Thus, our VAR should be a good starting point for identification of cointegrating relationships.

Table 3.2 shows that both the trace- and maximum-eigenvalue tests strongly support the existence of two cointegrating vectors, implying three common trends in the system. As complex eigenvalues come in pairs and the first eigenvalue of the companion form is real and the next one is complex (see Table A.6), this should further substantiate the existence of three common trends. To rephrase the message, as long as we believe that the second root really is complex this implies that there will either be two or four cointegrating relationships. However, the trace statistics of a rank less than or equal to two versus less than or equal to full rank has got a p-value slightly above 10 per cent, the asymptotic upper ten per cent fractile being approximately equal to 39.08. Also, complex roots may be realizations of stochastic processes with expectation values lying on the real line. If this is the case, three cointegrating vectors and two common trends will not represent an inconsistency problem but an interesting hypothesis that one should be able to at least investigate and later test formally given the distribution governing the roots. With regard to this last possibility, one may come a step further by undertaking a graphical inspection of how the roots within the unit circle are affected by the imposition of unit roots as to erroneously impose a complex root to lie at one should show up through its complex conjugate assuming a real value lying significantly far away from the unit circle. Looking at the four graphs in Figure B.11 of the eigenvalues of the companion matrix of the trading sector, we see that the imposition of the first common trend seems to reduce the second complex root to two real roots. Whether this change constitutes a significant change or not has to be formally tested, but based on the results of the trace-test statistics and the fact that the two new real roots seem to lie fairly close to each other, a rejection of the null of correct imposition of one unit root, would be very surprising. The third unit root, however, is slightly more controversial and one may discuss whether the imposition of the third unit root is accepted or not by looking at what happens to the complex conjugates of the complex root when imposing one of them to lie at one. I have chosen to decide on two common trends, even though the real transformation of the not restricted complex conjugate does not lie that much further away from the other one, restricted to lie at one, than in the preceding case. The decision of three cointegrating vectors, has been made more to be able to identify relationships interpretable in light of the theory outlined in Section 3.1.1 than to be consistent with the outcome of statistical tests. However, this is not to say that this has been a decision without controversy.

Table 3.3 gives the result of the identification of three cointegrating vectors. Three economically meaningful relationships are here identified to a level of 0.1182 for the LR-test with three degrees of freedom. The first CI-vector is a pure demand

Table 3.3: Restricted cointegrating relationships for the trading sector

a1	=	const.	- 0.605 (pa1-pw)	+ 2.7R	
			(0.0808)	(0.1614)	
pa1	=	const.	+ 1.775 pw	+ 0.854ulc	- 0.0085Trend
			(0.2933)	(0.2098)	(0.001695)
a1	=	const	+ R	+0.0096Trend	
				(0.00078)	

relationship for Norwegian exports with the implication of a quasi elasticity of relative prices in demand of about -0.6 and a foreign income quasi elasticity of 2.7.[11] Even though the second CI-vector is not homogenous in prices and does include a trend, it may be interpreted heuristically as a monopolistic price setting rule. The third CI vector says that the ratio of exports to foreign real income cointegrates with a deterministic trend, implying a yearly growth rate of about 3.8. Figure B.1 in the appendix, showing the graphs of the three concentrated restricted cointegrating relationships, does not reveal any threatening signs of non-stationarity, though the non-concentrated series (not shown here) do indicate a potential problem with the period before 1985.[12] Furthermore, the recursively estimated eigenvalues of Figure B.2, show no ominous signs of instability.

The service sector The appendix, Table A.4, also gives us the diagnostics of the five-dimensional VAR of the service sector. All individual equation diagnostics are fine, except perhaps for some marginal indications of autocorrelation in

[11]Strictly speaking, these coefficients cannot be interpreted as elasticities as changes in the residuals of the marginal processes will work through the whole simultaneous system. A one per cent increase in the marginal process governing i.e. foreign income, may therefore even lead to a percentage decline in the process governing exports in the long-run. However, one may modify the meaning of elasticity such that it fulfills the outcome of a feasible hypothetical experiment, an experiment where we change the initial values such that the outcome of a one per cent change in a marginal process gives the percentage change of another variable in the long-run as implied by the traditional interpretation of elasticity.

[12]The difference between the concentrated and non-concentrated cointegrating relationships could be taken to indicate a potential problem with higher order common trends. For a further investigation of this possibility the reader is referred to Hammersland(2002c).

Table 3.4: Rank tests for the service sector

System: a2, pa2, pw, R, ulc.
Deterministic part: Unrestricted constant, centered seasonals and restricted. Trend
VAR order: 3. Sample period: 1980 (1)-1998 (2).

Eigenvalues of Π: 0.9356 0.8134 0.7389 0.6521 0.5104

Max Eigenvalue Tests				Trace Eigenvalue Tests			
Null	Alternative	Statistics	95%	Null	Alternative	Statistics	95%
r=0	r≤1	49.77**	37.5	r=0	r≤5	124**	87.3
r≤1	r≤2	31.64*	31.5	r≤1	r≤5	74.24**	63.0
r≤2	r≤3	22.39	25.5	r≤2	r≤5	42.6*	42.4
r≤3	r≤4	15.28	19.0	r≤3	r≤5	20.21	25.3
r≤4	r≤5	4.925	12.3	r≤4	r≤5	4.925	12.3

the equations for *pw* and *ulc*. However, this nearly significant problem of autocorrelation at the individual equation level contributes to distorting the system test statistic, which in fact is significant to 1 per cent. However, to be able to precisely identify the long-run structures, I have given priority to the task of keeping the systems to as low order as possible. In this case, it is my belief that what is gained by giving priority to a parsimonious dynamic specification goes far beyond what is lost in terms of efficiency due to autocorrelation. The graphs showing the stability tests, Figure B.10, all indicate that the system seems to be fine.

The trace test statistic in Table 3.4 indicates two common trends. However, the third cointegrating vector is only marginally significant to a level of five per cent, so the additional information given by the fact that the first root is real and the second is complex may lead us to conclude that there are as many as three common trends in the system (see Table A.6). However, taking a closer look at what happens when imposing the third unit root, Figure B.12, we clearly see that the second remaining complex root which is the one affected, cannot have been the outcome of a stochastic processes with expectation values lying on the real line. When imposing one of the complex cojugates to lie at one, the one that is not restated does not reduce to a real value that is anyway close to the unit circle. This leads me to conclude that there seems to be evidence of three cointegrating vectors in the service sector. The recursive eigenvalues shown in Figure B.2, do all seem to be fairly stable.

The identified system of cointegrating relationships presented in Table 3.5, gives us again three economically meaningful relationships. First, we have that the ratio of exports to foreign income is constant in the long run. Contrary to the results of the traditional goods sector this implies that the relative market share of products in the service sector does not show any tendency to grow in the long run. The two last relationships constitute again two different types of pricing behavior, the first resembling the price setting behavior of a competitive firm which sets its prices as a markup over unit labor costs, the only difference being a trend term which may catch up a non constant markup over time. The other reflects the behavior of a monopolistic competitor.

So far, we have separately identified the long-run properties of the two sectors. Contrary to prior beliefs with regard to the effects of increased internationalisation, my results strongly indicate that small open economies like the Norwegian, still seem to have a considerable degree of monopolistic power when setting their prices. This finding is in accordance with the results in Bowitz and Cappelen(1994) who in fact find unit labor costs to be the single most important explanatory variable in all of their preferred equations for different subsectors of the Norwegian economy and is an expression of the fact that small countries can be "big" in what they produce.[13] Furthermore, to a large extent both sectors' export volumes seem to be driven by demand, which is the case when agents accommodate demand ex post to fixed prices ex ante. However, instead of elaborating further on these results, I will now look at the possibility of identifying long-run cross-sectional linkages when considering the long-run structure of each sector as known and given by the identified relationships of this section.

The pooled sector The VAR of the pooled data is of dimension 7 and order 3. As before, the trend is restricted to lie in the cointegration space and the constant and seasonal dummies enter unrestrictedly. Table A.5 of the appendix

[13] The findings of Bowitz and Cappelen however contrast with the finding in a more recent study, Naug (2001), who claims that Norwegian exporters of raw materials have limited power to set their own prices. A study that confirms the evidence of Bowitz and Cappelen at an aggregate level, is Hammersland (1996), who finds significant signs of monopolistic power in the process governing prices of exports in a study undertaken on aggregate data for the Norwegian mainland economy.

Table 3.5: Restricted Cointegrating Relationships for the service sector

a2	=	const.	+ R	
pa2	=	const.	+ 0.506 ulc	+0.0025 Trend
			(0.0588)	(0.00046)
pa2	=	const.	+ 0.653 R	+0.362 ulc
			(0.0966)	(0.0660)

Table 3.6: Rank tests for the pooled sector conditional on 6 known cointegrating vectors.

System: a1, a2, pa1, pa2, pw, R, ulc.
Deterministic part: Restricted Trend, Unrestricted Constant and centered seasonals
VAR order: 3. Effective Sample period: 1980 (1)-1998 (2)
Number of known cointegrating vectors $s = 6$

Trace Eigenvalue Tests: -2ln(Qm)= -T(log(det($\Omega(p)$))-log(det($\Omega(r-s)$)))

Null	Alternative	Test Statistics	HW 95% Critical values
m=0	m$\leq$1	14.1784	>26.4

contains the single equation and system diagnostics and all test statistics are fine. The imposition of the six known cointegrating vectors from the first step of our analysis has been implemented by using the option restrictions of subsets in CATS, and the LR test for overidentifying restrictions is fine.[14] The result of the cointegration analysis using the critical values of the trace test in Table 5 of Paruolo (1999), are given in Table 3.6 and clearly indicates that we cannot reject the null of one stochastic trend.[15] This indicates that we may have identified all cointegrating relationships in the information set already in the first step of the identification scheme and at first sight could seem to imply that there is no cointegration across sectors. However, taking a closer look at the cointegrating linear combinations of Table 3.3 and Table 3.5 above, reveals that in the long run $a1_t = a2_t + 0.009Trend$. This implies that we by undertaking a sector specific analysis have been lucky enough to also identify cointegration across sectors. To study this phenomenon closer, I have therefore undertaken an analysis where I have only treated the two first cointegrating vectors of the trading sector together with the three identified in the service sector, as known in the second step.[16]

The trace test statistics in the case when s is equal to five are given in Table 3.7 and strongly support the existence of another cointegrating vector.[17] The LR test of whether this additional cointegrating relationship is equal to the one identified from the two sector-specific analyses, gives a χ^2 equal to 0.18, implying that the significance probability of the test statistic is close to 0.67. Thus, by imposing five known cointegrating vectors we have been able to identify a cointegrating relationship across sectors which is consistent with the outcome of the sector specific analysis. It is imperative to point out that this relationship would not have been identified if we in the analysis of the first sector had accepted the outcome of the trace test without looking at the eigenvalues of the companion form. The

[14] The LR test for the imposition of seven restrictions in each of the six known cointegrating equations is χ^2 with 12 degrees of freedom and is equal to 15.6, implying that the p-value is approximately equal to 0.21.

[15] Table five in Paruolo does not calculate critical values for more than s=5 known cointegrating vectors. This is the reason for the bigger than sign in front of the five percent critical value which is taken from the s = 5 column in Paruolo (1999).

[16] To a certain degree this was also suggested by the rank test of the trading sector as neither the Trace nor the Max eigenvalue tests gave support to a rank beyond two.

[17] Based on small sample simulations of just about all test statistics of this paper this result is seriously called into question in Hammersland(2002d).

Table 3.7: Rank tests for the pooled sector conditional on five known cointegrating relationships

System: a1, a2, pa1, pa2, pw, R, ulc.
Deterministic part: Restricted Trend, Unrestricted Constant and centered seasonals
VAR order: 3. Effective Sample period: 1980 (1)-1998 (2)
Number of known cointegrating vectors $s = 5$

Trace Eigenvalue Tests: -2ln(Qm)= -T(log(det($\Omega(p)$))-log(det($\Omega(r-s)$)))

Null	Alternative	Test Statistics	HW 95% Critical values
m=0	m≤2	52.116	44.5
m≤1	m≤2	8.816	26.5

Table 3.8: Restricted long-run relationship in the pooled analysis

a1	=	const.	+ a2	+0.009 Trend
LR -test:	$\chi^2(1)$	=	0.18[0.67]	

relationship together with the LR test statistics are given in Table 3.8 and implies that exports in the Norwegian trading sector are growing approximately 3.6% faster than exports of the service sector, but that there is a strong long-run link between them. This coincides well with the perceived view of the trading sector being the main origin for innovative productivity improvements of the economy.

The third step In the introduction I alluded to a possible third step in my procedure of reestimating all free parameters in the identified structure to account for a possible non-diagonal covariance matrix. The results of this re-estimation are given in Table 3.9 and reveal a very interesting change in the equation for the export prices of the trading sector. With regard to the other equations all coefficients are pretty much the same as before. The coefficient for the world market price in the export price equation of the trading sector, however, changed from being significantly positive to being insignificantly negative. The null restriction on this coefficient did not represent a binding restriction on the cointegrating space so the restriction was not testable. However, we see that the restriction of the Trend in the same equation is a testable hypothesis and the test statistic clearly indicates that the restriction is valid in the sense of not being rejected. The important point however, is that by imposing these two restrictions we seem to have identified another across sector cointegrating relationship. Equations number 2 and 4 of Table 3.9 implies namely that we in the long run have the following relationship:

$$pa1 = 1.37pa2 - 0.004trend$$

The test statistic for the imposition of a unit quasi elasticity is χ^2 with one degree of freedom and is extremely close to zero. This implies that we cannot reject a null of a unit quasi elasticity and the final relationship becomes:

$$pa1 = pa2 - 0.002trend$$

Thus, in addition to the identified strong link between exports of the two sectors, there seem to be a strong long-run relationship between export prices of the two sectors. The relationship implies that the yearly inflation rate is about 0.8 per cent higher in the service sector than in the trading sector and could again be explained by a more competitive environment in the trading sector.

3.2. Example 2: Identification of international and domestic interest rate relationships: The case of Germany and the US.

As clearly indicated by Figure 3.1 below, the spread between long-term US and German interest rates reveals an extraordinarily high degree of correlation between the two countries' long-term interest rates. The figure also indicates that there seem to have been a lack of a similar relationship between domestic short and long rates in Germany. Taken together with the empirical evidence of a one-way causality between long US and German interest rates, going from the US economy to the German (ref. Hammersland(2002b)), these observations suggest that long-term interest rates in Germany during the nineties have been influenced

Table 3.9: Restricted long-run relationships in the pooled analysis when all free parameters have been estimated freely.

Eq.:			Cointegrating relationships:		
1:	a1	=	const.	-0.535(pa1-pw)	+2.380R
2:	pa1	=	const.	+ 0.694ulc	
3:	a2	=	const.	+ R	
4:	pa2	=	const.	+ 0.506ulc	+0.003Trend
5:	pa2	=	const.	+ 0.707R	+ 0.329ulc
6:	a1	=	const.	+ a2	+0.008 Trend

LR-tests :

All overidentifying restrictions:	$\chi^2(4) = 2.99[0.56]$
Restriction on the trend term in eq. 2:	$\chi^2(1) = 0.73[0.39]$

more by what is going on in international capital markets than by policy run by an independent German central bank. This again could be taken to indicate a lack of independence in the conduct of monetary policy on part of the Bundesbank, particularly as bank lending in most Continental European countries is overwhelmingly linked to long-term interest rates and Europe as a whole constitutes a fairly closed economy, the last point making it less susceptible to effects coming through the exchange rate channel.[18] The aim of this analysis, however, is not so much to go into a detailed discussion about this as undertaking an alternative empirical analysis based on the approach suggested in this paper and the interested reader is referred to Hammersland (2002c). Notwithstanding, when discussing the possibility of a third cointegrating vector in the last part of this section, the issue will be forced upon us as this hypothesis radically affects the implications of the analysis. Thus whether there are two or three cointegrating vectors in the information set is not going to be a trivial decision which might be left to the stochastic outcome of test statistics alone. As the decision will have a central bearing on the outcome of the analysis it should be substantiated within the framework of prior beliefs, reliability of results and not least theory in conjuction with the results of the statistical analysis.

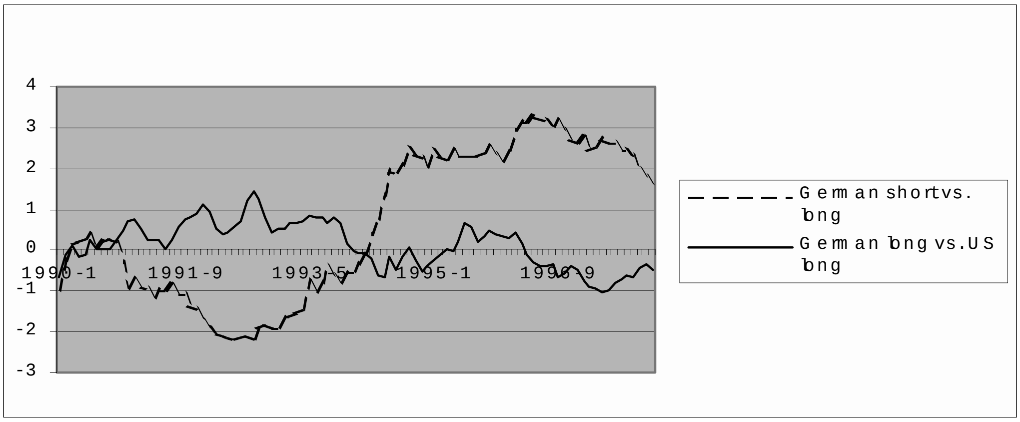
To study the degree of independence in European capital markets it is natural to base the analysis on different theories of arbitrage and especially to look at the long end of the market. To clarify matters further, I will therefore in the next subsection give a brief review of two dominating theories concerning the determination of long-term interest rates, the theory of uncovered interest rate parity (UIP) and the expectation theory of the term structure, respectively.

3.2.1. Some theories of interest rate determination and their implications with regard to cointegration

The theory of uncovered interest parity is a relationship between foreign and domestic interest rates on assets of the same maturity and says that in a steady state the expected return of investing one unit of domestic currency must be

[18] As illustrated by Borio(1995), the share of outstanding debt bearing interest rates which were either predominantly fixed or indexed to long-term interest rates for six of the seven largest European economies, amounted in 1993 to more than 55% . The only country among the seven with a significantly lower share at that time was Italy. Recent evidence shows however that things have changed dramaticly in Italy since then and that the share of mortgages at fixed long-term interest rates has increased from 25 per cent in 1993 to more than 50 per cent in 1997 (European Mortgage Federation (1998)).

Figure 3.1: Interest rate spreads



the same whether one invests domestically or abroad. The long rates should therefore be equal to the corresponding foreign long rates plus the expected rate of depreciation of the home currency against the foreign currency.[19] Given a stochastic representation, this may be expressed as:

$$i_t = i_t^* + Dv + \epsilon_t \tag{3.3}$$

I have here assumed rational expectations such that $Dv = Dv^e + \varepsilon$.[20] Furthermore ϵ_t in (3.3) is assumed to be stationary, I(0), such that the spread between

[19] An important caveat in the following, is that the treatment below deliberately disregards the potential existence of disturbing risk and term premiums. As these probably are two of the most important reasons why econometricians have problems identifying long-run cointegrating arbitrage relationships between yields of different maturities as well as between yields of different countries of origin, as i.e. the UIP hypothesis, it is important to realize that they might apply in this study as well.

[20] Even in the case this highly disputed assumption is fulfilled to perfection, the so called peso problem might pose problems in small samples. This happens because rationality does not guarantee that the empirical mean of actual realignments coincides with the realignment expectations, particularly when the probability of observing small changes in the exchange rate within a band is high whereas the opposite is the case with regard to observing a realignment. Besides a non-zero risk premium, this is the most frequent explanation met in the literature to explain why the hypothesis of uncovered interest rate parity often is rejected in actual data sets. So also in this study as our sample not exactly is a big one.

domestic and foreign long term interest rates, $i_t - i_t^*$, cointegrates with the depreciation rate. It is worth noting that in the case of a stationary rate of depreciation the interest rate spread will be stationary as well.

The expectation theory of the term structure on the other hand is a relationship between interest rates of different degree of maturity and says that long rates should be equal to a weighted average of current and expected future short-term interest rates. Thus, the impact on long-term interest rates from a change in current short-term interest rates depends on how expected future short-term interest rates are affected. A rise in current short-term interest rates that is regarded as permanent will lead to a full pass-through from short-term to long-term interest rates. On the other hand, if an increase in the current short-term interest rate leads to a significant reduction in inflation expectations, long-term interest rates may even decline. In the case of a full pass through from the short to the long end of the market, the relationship can be given the following stochastic representation:

$$i_t^l = i_t^s + \epsilon_t \tag{3.4}$$

As above the noise term is assumed to be stationary, I(0), such that the spread between the long rate, i_t^l, and the short rate, i_t^s is stationary as well.

Taken at face value this implies that one should expect there to be at least two long-run relations. One that concerns the arbitrage across borders and another one representing a domestic arbitrage condition of bonds with different degree of maturity.

3.2.2. Data and time series properties

The econometric analysis is based on monthly observations of short- and long-term interest rates in Germany and the US together with the bilateral exchange rate between the two countries. The period I am looking at is from 1990 (1) to 1997 (12). More explicitly the data set consists of monthly observations on the following variables:

i^{GL} Effective interest rate on German Government bonds with ten years to maturity

i^{GS} Three months money market interest rate for Germany

i^{UL} Effective interest rate on US Government bonds with ten years to maturity

i^{US} Three months money market interest rate for the US

Dv The change in the bilateral exchange rate, German marks per US dollar.[21]

With regard to the time series properties of the data I refer to Hammersland (2002a). The results of multivariate tests and Dickey Fuller tests of stationarity herein, all indicate that interest rates are I(1), while the change in the bilateral exchange rate is stationary, I(0). However, to further substantiate the claim of no higher order of non-stationarity than of order one, I have run the data through Johansen's two-step procedure for estimation and identification of the cointegration indices (Johansen (1995b)). The I(2) test, fully described in Paruolo (1996) and Jorgensen, Kongsted and Rahbek (1999), has been based on the specification of a second order VAR of dimension five where the constant term is restricted to lie in the cointegration space and a possible drift term has been restricted such that it does not generate quadratic trends. In Table 3.10 below the joint test of the number of cointegrating vectors, r, and the number of I(1) trends, s, is denoted $S_{r,s}$. The test statistics are given in bold letters while the 95% fractiles simulated in Paruolo (1996), are given in italics below. As explained before the test procedure starts from top left testing the null of five common I(2) trends versus less than or equal to full rank and continues to the right until one reaches the last column which is the ordinary test of five I(1) trends versus more than or equal to nil common trends. In the case where one rejects all nulls in the first row of five common trends, one continues this stepwise testing from left towards right by moving down to the next row of four common trends. The number of

[21] I have chosen to use the first difference of the logarithm of the bilateral exchange rate knowing that it would have been more correct theoretically to use either the three months or the twelve months difference. In this respect the one month difference must be viewed as a compromise and as an indicator of what it after all is meant to proxy, the expected rate of change in the exchange rate.

Table 3.10: The trace test of cointegrating indices

p-r	r	$S_{r,s}$					Q(R)
5	0	**380.46**	**285.33**	**227.44**	**177.70**	**147.53**	**125.61**
		198.2	167.9	142.2	119.8	101.5	87.2
4	1		**222.77**	**165.43**	**119.16**	**89.12**	**68.08**
			137.0	113.0	92.2	75.3	62.8
3	2			**135.10**	**81.31**	**59.82**	**39.80***
				86.7	68.2	53.2	42.7
2	3				**58.34**	**36.07**	**17.61**
					47.6	34.4	25.4
1	4					**15.47**	**5.13**
						19.9	12.5
p-r-s		5	4	3	2	1	

[1)]Table 3.10 is based upon a five dimensional VAR of order two for the variables i^{GL}, i^{GS}, i^{UL}, i^{US} and DV. A constant is restricted to lie in the cointegration space and a possible drift term has been restricted not to generate quadratic trends.
[2)]The figure in italics under each test statistic is the 95 per cent fractile as tabulated by Paruolo(1996). The preferred outcome is marked with a star.

cointegrating vectors, I(1) and I(2) trends are given by the first null that one cannot reject. In our analysis this happens in the row of three common trends and in the column where all trends are I(1), clearly indicating that there are no I(2) trends in the data and that the number of cointegrating vectors is equal to two. The results are thus fully in line with the tests referred to above and confirms the finding of an order of non-stationarity not higher than one. Also, already at this stage the results give support to the finding in Hammersland (2002a) that there are no more than two cointegrating vectors among the variables in the information set. This finding however, will be further scrutinized in the last section in view of the outcome of the second step of the analysis.

Table 3.11: Rank tests for the German sector

System: i^{GL}, i^{GS}, Dv.
Deterministic part: Restricted constant and no trend
VAR order: 2. Effective sample period: 1990 (1)-1997 (12).

Max Eigenvalue Tests				Trace Eigenvalue Tests			
Null	Alternative	Statistics	90%	Null	Alternative	Statistics	90%
r=0	r≤1	41.84**	14.09	r=0	r≤3	56.53**	31.88
r≤1	r≤2	13.54*	10.29	r≤1	r≤3	14.70	17.79
r≤2	r≤3	1.24	7.50	r≤2	r≤3	1.24	7.50

3.2.3. Cointegration analysis

The two sectors we are looking at are Germany and the US. In analyzing the separate sectors I have in both cases started out with a three-dimensional VAR of order two where in addition to the country specific interest rates, I have also included the bilateral exchange rates. The econometric models do not include a trend and the constant terms are restricted to lie in the cointegrating space.[22] As these data have been extensively examined in Hammersland (2002a), I will here only comment on the diagnostics of the sector specific VARS to the extent that these deviate significantly from the full VAR diagnostics in this paper.

Germany Table 3.11 gives strong support to the existence of only one cointegrating vector, and as both tests are way off the ninety five percent critical values of 15.7 and 20, respectively, I have chosen to accept this outcome without resorting to a discussion of the eigenvalues of the companion matrix.[23] The unrestricted cointegrating linear combinations together with some LR-tests for overidentifying restrictions are given in Table 3.12. The first thing to note is the rejection of excluding the rate of depreciation, Dv_t. This indicates that the rate of depreciation must either be a stationary linear combination in itself, or that it in some way

[22]The inclusion of a restricted trend term to assure that the tests are similar, has been avoided on behalf of the a priori very unlikely finding of a trend-stationary interest rate relationship.

[23]The two first eigenvalues of the companion matrix are real and equal to 0.9971 and 0.9093, respectively. The trace test clearly shows that the second one is significantly close to the unit circle. The other eigenvalues are complex but their norms are all less than 0.43 so disregarding these should not represent any problem.

Table 3.12: The unrestricted cointegrating linear combinations and tests of restrictions on the cointegrating space for the German sector

$$\beta'(i^{GL}, i^{GS}, Dv, 1) = \beta_{11} i^{GL} + \beta_{12} i^{GS} + \beta_{13} Dv + \beta_{14}$$
$$= 0.591 i^{GL} - 0.096 i^{GS} + Dv - 0.036$$

Hypotheses:	LR-test, Rank=1.
$\beta_{13} = 0$	$\chi^2(1) = 28.74[0.00]$
$\beta_{11} = -\beta_{12}$, $\beta_{13} = \beta_{14} = 0.$	$\chi^2(3) = 41.25[0.00]$
$\beta_{11} = \beta_{12} = \beta_{14} = 0$, $\beta_{13} = 1.$	$\chi^2(3) = 01.75[0.63]$

Table 3.13: Rank tests for the US sector

System: i^{UL}, i^{US}, Dv.
Deterministic part: Restricted constant and no trend
VAR order: 2. Effective sample period: 1990 (1)-1997 (12).

Max Eigenvalue Tests				Trace Eigenvalue Tests			
Null	Alternative	Statistics	90%	Null	Alternative	Statistics	90%
r=0	r≤1	45.92**	14.09	r=0	r≤3	55.70**	31.88
r≤1	r≤2	6.99	10.29	r≤1	r≤3	9.78	17.79
r≤2	r≤3	2.79	7.50	r≤2	r≤3	2.79	7.50

cointegrates with the two interest rates. However, the significance probability for the test of a stationary depreciation rate is equal to 0.63, implying that we are way off rejecting the null of stationarity and in the following this is therefore going to be treated as one of the known cointegrating vectors in the pooled analysis to come.[24]

[24]The diagnostics of the VAR rejects normality and there are signs of ARCH effects too in the model. However, simulation studies have shown that the trace test is fairly robust against certain form of deviations from normality and ARCH effects do not in general seem to invalidate the analyses.

Table 3.14: The unrestricted cointegrating linear combinations and tests of restrictions on the cointegrating space for the US sector

$\beta'(i^{UL}, i^{US}, Dv, 1)$	$= \beta_{11} i^{UL} + \beta_{12} i^{US} + \beta_{13} Dv + \beta_{14}$
	$= 1.056 i^{GL} - 0.197 i^{GS} + Dv - 0.064$
Hypotheses:	LR-test, Rank=1.
$\beta_{13} = 0$	$\chi^2(1) = 38.52[0.00]$
$\beta_{11} = -\beta_{12}$, $\beta_{13} = \beta_{14} = 0.$	$\chi^2(3) = 44.62[0.00]$
$\beta_{11} = \beta_{12} = \beta_{14} = 0$, $\beta_{13} = 1.$	$\chi^2(3) = 06.33[0.10]$

The US As for Germany the trace test statistics of Table 3.13 clearly indicate that there is only one cointegrating vector among the variables in the information set.[25] Table 3.14 shows furthermore that the picture is very much the same as for the German sector. The stationarity of the depreciation rate cannot be rejected and there does not seem to be any support for a stationary interest spread or other linear combinations which includes the depreciation rate among the variables.

Pooled sector Based on the result that both country specific analyses gave the same outcome with regard to the identified cointegrating vector, the pooled analysis in this section will be contingent on one of the cointegrating relationships being the rate of depreciation. In the terminology of Section 2 this means that the unit vector with zeros for all coefficients except for the one for the depreciation rate will be treated as the known cointegrating vector.

Table 3.15 shows the trace eigenvalue tests when conditioning on the rate of depreciation as a known cointegrating relationship. The critical values are based on the distributional results in Horwath and Watson (1995) and the 95 per cent fractiles, as tabulated by Paruolo (1999), are given in the last column. The first thing to notice is that the test statistic strongly rejects the null of 4 common trends versus more than or equal to nil. This means that there is at least one

[25] The first eigenvalue of the companion form is complex with a norm equal to 0.9349. Disregarding the possibility that this root could be the outcome of a stochastic process with expectation value lying on the real line, this gives immediate support to an hypothesis of two common trends and thus only one cointegrating vector.

Table 3.15: Conditional rank tests for pooled sector

System: i^{GL}, i^{GS} i^{UL}, i^{US}, Dv
Deterministic part: Restricted constant and no trend
VAR order: 2. Effective Sample period: 1990 (1)-1997 (12)
Number of known cointegrating vectors $s = 1$

Trace Eigenvalue Tests: -2ln(Qm)= -T(log(det($\Omega(p)$)-log(det($\Omega(r-s)$)))

Null	Alternative	Test Statistics	HW 95% Critical values
$m = 0$	$m \leq 4$	72.18	59.0
$m \leq 1$	$m \leq 4$	43.43	40
$m \leq 2$	$m \leq 4$	21.06	24.1
$m \leq 3$	$m \leq 4$	6.67	12.1

cointegrating vector in the pooled information set beyond the one we found from the country specific analyses. Noteworthy, the statistics give also relatively strong evidence for the existence of a third cointegrating vector as the critical value to a level of one per cent of the test of the hypothesis of more than or equal to three common I(1) trends versus less than or equal to full rank is approximately equal to 46.2, implying that the test's significance probability lies somewhere between one and five per cent.[26] However, based on the results of former analyses and the outcome of the Johansen two-step procedure for identification of cointegrating indices in Section 3.2.2. which all give support to the hypothesis of only two cointegrating vectors in the full information set, I will first comment on the case with only one additional cointegrating vector to the one "known" from before. In this respect it will be of particular interest to find out whether this relationship can be given the interpretation of being the spread between German and US long-term interest rates. Then I will turn to the possibility of an additional third cointegrating vector and the consequences such an hypothesis might have with regard to the central conclusions of this paper concerning the ability of the Federal Reserve and the Bundesbank to control the long end of their respective capital markets.

[26] Based on the small sample Monte Carlo experiments in Hammersland (2002d) and the outcome of the unconditional cointegration analysis of the pooled information set in Hammersland (2002a) this seems to be a relatively robust finding.

The case of two cointegrating vectors The joint restriction of a stationary depreciation rate and interest rate spread gives a LR test statistic of 13.68. As this statistic is χ square with 8 degrees of freedom, this implies a significance probability of about 0.09. The test of a stationary spread conditional on the depreciation rate being stationary, is distributed $\chi^2(4)$ and the p-value of the test statistics is approximately equal to 0.05. In my view this should give sufficient support to the hypothesis that long rates in Germany and the US cointegrate and we have been able to identify a cointegrating relationship across the two countries by using the suggested two-step procedure. Before turning to the alternative analysis involving three cointegrating vectors, I will have a look at the finding in Hammersland and Vikøren (1997) that long-term German interest rates seem to cointegrate with a homogenous linear combination of domestic short and US long-term interest rates. The test statistic for this hypothesis is χ square with three degrees of freedom and gave a test statistic of 3.55. This implies that the null cannot be rejected even to a level as high as 30 per cent. However, we have already seen that the restriction implying a stationary spread is not rejected either. Thus German short interest rates do not seem to significantly enter the long-run relationship which could be taken to indicate reduced domestic control on part of the German central bank, the Bundesbank, with regard to the long end of the market.[27] Also the tests for exogeneity undertaken in Hammersland (2002a) imply that both US rates might be considered driven by two independent processes outside our information set. In particular this could be taken to mean that long-term interest rates in the US are beyond control of monetary authorities, a possible explanation being that these are driven by what is going on in international capital markets. However, such a finding also rejects the possibility of a causal relationship going in the opposite direction, that is that long rates would affect short rates through its capacity of being a variable entering the information set of a policy rule on part of the Fed. As these issues are of most importance to get scrutinized I will indulge in a more comprehensive discussion of these matters when analyzing the possibility of three cointegrating relationships in the next section to come.

[27] However, in this respect one should note that the lack of a long-run relationship between short and long term domestic interest rates might be due to short-run dynamic interest rate effects being offset by an oppsite move in future interest rate expectations. If this is the case, there is nothing strange in the fact that one is unable to identify a long run relationship between domestic short- and long-term German interest rates. On the contrary it is exactly what to be expected if the monetary transmission mechanism works appropriately. For a critical discussion of this possibility the interested reader is referred to Hammersland (2002a and 2002b).

Table 3.16: The cointegration parameters, tests of restrictions on the cointegrating space and the identified system in case of three cointegrating vectors.

The cointegration parameters

$$\beta' \begin{bmatrix} i^{GL} \\ i^{UL} \\ i^{GS} \\ i^{US} \\ Dv \\ 1 \end{bmatrix} = \begin{matrix} \beta_{11} i^{GL} + \beta_{12} i^{UL} + \beta_{13} i^{GS} + \beta_{14} i^{US} + \beta_{15} Dv + \beta_{16} \\ \beta_{21} i^{GL} + \beta_{22} i^{UL} + \beta_{23} i^{GS} + \beta_{24} i^{US} + \beta_{25} Dv + \beta_{26} \\ \beta_{31} i^{GL} + \beta_{32} i^{UL} + \beta_{33} i^{GS} + \beta_{34} i^{US} + \beta_{35} Dv + \beta_{36} \end{matrix}$$

Hypotheses:	LR-test, Rank=3.
$\beta_{11} = \beta_{12} = \beta_{13} = \beta_{14} = \beta_{16} = 0,$ $\beta_{15} = 1.$	$\chi^2(3) = 04.00[0.26]$
$\beta_{23} = \beta_{24} = \beta_{15} = \beta_{26} = 0,$ $\beta_{21} = -\beta_{22}.$	$\chi^2(6) = 12.64[0.05]$
$\beta_{31} = -\beta_{33}, \beta_{32} = -\beta_{34}, \beta_{35} = 0.$	$\chi^2(8) = 13.24[0.10]$

The identified system of cointegrating relationships

$$\hat{\beta}' \left(i^{GL}, i^{UL}, i^{GS}, i^{US}, Dv, 1\right) = \begin{matrix} Dv_t \\ i^{GL} - i^{UL} \\ (i^{GL} - i^{GS}) + (i^{UL} - i^{US}) - 0.023 \end{matrix}$$

The case of three cointegrating vectors Table 3.16 gives the outcome of the analysis when accepting the existence of three cointegrating vectors. The cointegration space as implied by the identified structure in addition to the two relationships identified assuming three common trends, is spanned by a negative relationship between German and US domestic interest rate spreads. To reject the hypothesis of correctly imposed identifying restrictions for the system as a whole is not possible to a level below ten per cent. Furthermore, the p-value of the test of the identifying restrictions concerning the third vector conditional on a stationary rate of depreciation and a stationary spread between US and German long-term interest rates is close to 0.74. Thus the long-run relationship implied by the identifying restrictions on the third relationship constitutes a valid restriction on the cointegrating space. That this relationship between national spreads is a negative one might however seem rather puzzling as one intuitively would think of a hike in the US spread either through a hike in long term interest rates or a fall in short term interest rates, to cause a similar increase in German spreads through a redirection of funds on the part of investors. However, as already alluded to in the text, central banks seem to have a fairly tight control over the short end of the capital market. Based on the anti-inflationary reputation of the Bundesbank it is therefore not at all surprising that a fall in short-term US interest rates was not allowed to affect the corresponding German interest rates in a way that could seriously jeopardize its anti-inflationary reputation. Thus, the predominant effect of a fall in US short-term interest rates on the German spread would come via the effect it might have on domestic long-term interest rates, either directly via arbitrage or indirectly through changes in the bilateral exchange rate. Disregarding the possibility that the effect of a fall in short rates on expectations with regard to future domestic short-term interest rates might totally offset or even dominate the unequivocally positive portfolio effect, a fall in short rates will lead to a fall also in long-term interest rates. This will make investments in domestic assets with a long-term to maturity less favorable compared to foreign alternatives and lead to increased demand for foreign assets with a long term to maturity. As monetary authorities do not have the same possibility of controlling the long end of the market as the short, the outcome of this arbitrage activity would predominantly be determined by capital markets alone implying that long rates would fall. When a national central bank chooses to cut its policy rate, this will in addition have consequences for the bilateral exchange rates between the country that undertakes the cut and its trading partners, whose bilateral exchange rates naturally would have to appreciate.[28] The appreciation would surely contribute to reduce contemporary inflation as well as expectations with regard to future rates of inflation such that long-term interest rates would fall even more. Going back to our example of a fall in short-term US interest rates and its potential effect on the German economy, this would imply that prices on German assets with long term to maturity will rise and thus corresponding interest rates to fall. All in all therefore, there seem to be good reasons why we should see a negative relationship between domestic US and foreign German interest rate spreads when a widening of the US spread is due to a cut in the Federal funds rate. In the case the widening is due to rising international long-term interest rates, probably originating from shocks to international capital markets reflecting increased expectation of a future inflationary pressure, and short rates are informed by long rates through their capacity of informing policy rules on part of central banks, it is still possible to argue for a negative correlation between the domestic interest rate spreads of Germany and the US. This is due to the strong potentially offsetting effect that a hike in German policy rates might have had on expectations with regard to future inflation and thus future short term interest rates and long term interest rates just because of high credibility in its pursuit of an anti-inflationary policy stance. The anti-inflationary reputation of Germany further suggests that long-rates might have played a more significant role in forming a policy rule on part of the Bundesbank than on part of the Fed. An increase in international long-term interest rates might therefore in addition to affecting policy rates, have had the effect of increasing the spread between German and US short-term interest rates. As alluded to above this would bring about an immediate appreciation of the bilateral exchange rate and thus contribute to reduce inflationary expectations, future short-term interest rates and thus the long-term German interest rate. Whether the German spread will shrink will depend on whether the combined effect of the induced hike in the German policy rate and the appreciation that follows a potential widening of the spread between short-term international interest rates is sufficiently strong to offset the initial hike in long rates.

To be able to discuss the implication of the third cointegrating vector with regard to the central theme of this paper, that is the degree of independence in the

[28]The cut implies that the expected return of investing one unit of the domestic currency abroad will be higher than investing it domestically for a given expected rate of depreciation. For a given foreign interest rate and assuming uncovered interest parity, this implies that the bilateral exchange rate must appreciate to generate a higher expected rate of depreciation.

making of a monetary policy, one may give the identified long-run structure the equivalent representation of Table 3.17 below. From this formulation we clearly see that both US as German short-run interest rates enter into the cointegrating relationships. Thus there certainly seems to be a kind of link between domestic short and long rates even though at this stage it is too early to say anything about the direction of causality without making further inquiries into which processes might be characterized as exogenous and not. With regard to a potential relationship between short rates we have not been able to identify a relationship giving support to the hypothesis of uncovered interest parity(UIP).[29] Assuming therefore that Central banks are able to control the short end of the yield curve we may at least eliminate a direct causal long-run relationship between short rates. The important question to answer is therefore whether the link between domestic short- and long-term interest rates comes predominantly through the role played by domestic long-term interest rates as indicators of build ups in inflationary pressures and thus as explanatory variables informing a policy rule on part of Central banks, or through a pure term structure relationship between interest rates with different times to maturity. Tests on the loadings strongly indicate that US long-term interest rates seem to be determined by a process outside control of the monetary authorities, the Fed. However, the corresponding test of whether also long-term German interest rates are exogenous with regard to estimation of the long-run parameters, rejects to a level below one per cent. As the corresponding tests of the short-term interest rates of both countries also reject, this could be turned to account of domestic short-term interest rates being set in accordance with policy-rules informed by movements in domestic long-term interest rates of both countries.[30][31] Note that this eliminates the sort of argument used to argue for a negative relationship between spreads in the case of changes to domestic

[29] This result is in accordance with a great bunch of economic literature. Two references are MacDonald and Taylor (1992) and Froot and Thaler (1990). For a more recent account see MacDonald and Juselius (2002a,2002b).

[30] The test statistic of the joint null restriction on all loadings in the equation of long-term US interest rates conditional on the identified long-run structure of Table 3.16, is $\chi^2(3) = 0.208$ with a p-value of about 0.98. The statistics of the corresponding hypothesis test of German long and short rates and US short rates are respectively, 17.6 (0.0005), 9.56 (0.0188) and 11.88 (0.0078), all statistics being χ^2 with three degrees of freedom. The values in brackets are the corresponding statistics' p-values.

[31] This is also given support by the simultaneous equation modell identified in Hammersland (2002b), where the direction of contemporaneous short-run causality had to be reversed to get rid of some huge unexplained off-diagonal correlations in the covariance matrix of the structural model in Hammersland (2002a).

Table 3.17: Alternative representation of the identified structure of cointegrating relationships in case of three cointegrating vectors

	Dv_t
$\overset{\wedge}{\beta}{}' \, (i^{GL}, i^{UL}, i^{GS}, i^{US}, Dv, 1) =$	$i^{GL} - \frac{1}{2}(i^{US} + i^{GS} + 0.023)$
	$i^{UL} - \frac{1}{2}(i^{US} + i^{GS} + 0.023)$

short-run interest rates, as these changes are motivated from shocks to processes that inform the policy rule and not from shocks to exogenous processes governing short-term interest rates. With regard to the central theme of the chapter both the Fed and the Bundesbank seem to be able to set short-term interest rates at their discretion following policy rules informed by long-term interest rates. However, with regard to controlling the long-end of the market there does not seem to be a similar propensity. First, US long-term interest rates seem to be totally determined by processes outside the domain of US monetary authorities. The endogeneity status of the corresponding German interest rates and the fact that there is a long-run relationship between these and the long-term US interest rates suggest that also the German long-term interest rates mainly seem to be driven by the same forces that govern US long term interest rates.

4. Conclusion

In this chapter I have used the concept of known cointegrating vectors to come up with a suggested two-step procedure for how to deal with cointegration in the case of times series with a small cross-sectional dimension. The first step of this procedure implies getting to know the “known” cointegrating vectors by preliminary identification of long-run relationships along the cross-sectional dimension. Given these, the next step then implies identification of further long-run relationships across cross-sections by undertaking a joint analysis of all cross-sectional units simultaneously. The first step of this procedure uses ordinary reduced rank methodology to estimate the rank and to identify the CI-relationships. To identify the rank in the second step, however, we have to exploit the fact that these statistics will have the asymptotic distributions given in Horvath and Watson(1995) and simulated in Paruolo (1999).

To illustrate the procedure, I have undertaken two separate analyses. One where I estimate a two-sector model of exports for a small open economy on Norwegian data and another where I look at interest rate relationships between Germany and the US as well as relationships governing domestic interest rates within each individual country. In the first study we find no less than six theoretically consistent cointegrating relationships of which four represent sector specific long-run relationships. The other two are relationships between sectors and implies that there are strong ties between the export sectors of traditional goods and services in Norway. In particular, they imply that exports grow approximately at an annual rate of 3.6 per cent faster in the trading sector than in the service sector and that inflation seems to be approximately 0.8 percentage higher in the service sector than in the trading sector. This could be explained by a more competitive environment in the trading sector. Furthermore, the analysis does not give support to a long-run PPP relationship. On the contrary, my empirical results indicate that small open economies like the Norwegian, still have considerable market power in the export market. With regard to the interest rate study, the second step of the analysis clearly identified a relationship between German and US long-term interest rates. The first step of the analysis was not able, however, to reveal a similar relationship between short and long rates within each country. Taken at face value these findings could be taken to indicate that Europe in its conduct of monetary policy may have lost control to international capital markets. Although the picture changes somewhat with regard to how short-term interest rates are determined, the alternative analysis based on the existence of three cointegrating vectors does not change the central message that central banks do not seem to be able to control the long end of the capital market. However, the results clearly indicate that the exogeneity status given to US short-term interest rates based on two cointegrating vectors is incorrect. Short term US interest rates seem to be determined in accordance with a policy rule informed by long-term interest rates. Whether the Fed on basis of this can be said to have been successful with regard to controlling inflation is difficult to say and needs further investigation with an extended information set. However, the fact that the Fed seems to have lost control over the long end of the market, as implied by exogenous long-term interest rates, indicates that an important channel through which monetary policy could affect the real economy has been literally blocked. A plausible interpretation of this might be the increasing importance and dependence on international capital markets. With regard to Germany, there seems to be a similar causal relationship between short and long rates. However, in contrast to US long-term interest rates the corresponding German long-term interest rates are endogenous in the sense of being caused by movements to the US long-term interest rate. Again this finding substantiates the claim that the German Central Bank, the Bundesbank, during the nineties was not able to control the long end of the market, a conclusion that might have far-reaching implications given that bank lending in Continental Europe mainly is of a long-run character. However, as in the US the German Bundesbank seems to have been able to set their short-term interest rates independently in accordance with a policy rule. As for the US, to substantiate whether Bundesbank has succeeded in its endeavor of controlling inflation one will have to undertake further analysis on extended information sets.

All in all, in this chapter I have proposed a procedure of how to deal with cointegration when dealing with times series that vary along a small cross-sectional dimension. Even though the dimensions of the information sets in the two examples used to demonstrate the procedure have been too small to really demonstrate its full potential, its my view that the examples at least have served to illustrate that the suggested three-step procedure may be an useful device in helping out with the identification of cointegrating relationships when dealing with two-dimensional data.

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A. Tables

Table A.1:
Multivariate statistics for testing stationarity[1)]

	Variables						
	$a1$	$a2$	$pa1$	$pa2$	pw	R	ulc
$\chi^2\,(3)$	15.87**	16.19**	7.11	12.84**	6.87	11.6**	9.59*
	(0.001)	(0.001)	(0.07)	(0.005)	(0.076)	(0.009)	(0.022)

[1]The test statistics are the LR-tests of restrictions on the cointegration space within the Johansen framework. Specifically, these statistics test the restriction that one of the cointegrating vectors contains all zeros except for a unity corresponding to the coefficient of the variable we are testing for stationary. The test is conditional on the number of cointegrating vectors. In Table A.1, the statistics quoted are conditional on there being three CI-vectors and refer to the same VAR model that later is used to identify the long-run relationships. The figures in brackets under each Statistics are the tests' significance probabilities and * and ** denote rejection at.5% and 1% critical levels, respectively.

Table A.2:
ADF(N) Statistics for testing for a unit root.
Estimates of $|\widehat{\rho}-1|$in[1), 2)]

H_0	Variables						
	$a1$	$a2$	$pa1$	$pa2$	pw	R	ulc
$I\,(1)$	-2.11	-2.13	-1.884	-2.37	-0.68	-3.48	-2.57
	(0.22)	(0.35)	(0.080)	(0.167)	(0.027)	(0.127)	(0.11)
$I\,(2)$	-11.10** [3)]	-12.72**	-4.26**	-5.10**	-5.99**	-4.52**	-4.60**
	(2.708)	(2.992)	(0.929)	(1.272)	(1.41)	(0.755)	(1.624)

[1]For any variable x and a null hypothesis of I(1), the ADF statistics are testing a null hypothesis of a unit root in x against an alternative of a stationary root. For a null hypothesis of I(2), the statistics are testing a null hypothesis of an unit root in Δx against the alternative of a stationary root in Δx.
[2]For a given variable and the null hypothesis of I(1) and I(2), two values are reported. The N'th-order augmented Dickey-Fuller (1981) statistics, denoted ADF(N) and (in parentheses) the absolute value of the estimated coefficient on the lagged variable, where that coefficient should be equal to zero under the null hypothesis. Both a constant- and a trend-term together with seasonal dummies are included in the corresponding regressions when testing the null of I(1), whereas only a constant is spesified when testing for I(2). N varies across the variables for both tests and is equal to three for $a1$, $a2$, $pa1$ and R, two for $pa2$ and ulc, and four for pw in the test of I(1), whereas the corresponding lags in the second case are two for $a1$, $a2$ and R , and three for $pa1$, $pa2$, pw and ulc. The effective sample-periods have been.1980(1) -1998(2).
[3]Here and elsewhere in the paper, asterisks * and ** denote rejection of the null hypotheses at the 5% and 1% significance level, respectively. The critical values for the ADF statistics for testing I(1) are -3.47 at a level of 5% and -4.084 at a level of 1 % (MacKinnon (1991)).

Table A.3:
Individual equation and system diagnostics of the unrestricted VAR of the Trading sector[1]

Equation/Tests	AR 1-5 F[5,49]	ARCH 4 F[4,46]	Normality χ^2 (2)
$\Delta a1$	0.9517[0.4565]	2.9843[0.0285]*	2.9351[0.2305]
$\Delta pa1$	2.9069[0.0224]*	1.5551[0.2023]	0.2446[0.8849]
Δpw	1.8161[0.1271]	0.4855[0.7463]	0.1556[0.9252]
ΔR	0.9739[0.4431]	0.6538[0.6272]	0.3227[0.8510]
Δulc	1.7365[0.1439]	0.6161[0.6532]	0.6316[0.7292]

System tests:	AR 1-5[125,127]	VNormality χ^2(10)	VX^2 F[480,155]
Statistics:	1.2571[0.1001]	6.4595[0.7753]	0.31062[1.000]

[1]The Values shown in brackets are the individual test's significance probability. * and ** denote as usual rejection of the corresponding null at levels of 5 and 1 per cent, respectively. VNormality and VX^2 denotes the Vector tests of normality and heteroscedaticity. For an explanation of the various test statistics the reader is referred to Chapter 14 of the PcFiml manual (Doornik and Hendry (1999)).

Table A.4:
Individual equation and system diagnostics of the unrestricted VAR of the Service sector[1]

Equation/Tests	AR 1-5 F[5,49]	ARCH 4 F[4,46]	Normality χ^2 (2)
$\Delta a1$	1.5588[0.1810]	0.1525[0.9609]	1.1181[0.5717]
$\Delta pa1$	0.3199[0.8986]	0.5639[0.6900]	0.8459[0.6551]
Δpw	2.3861[0.0515]	0.4787[0.7511]	0.1229[0.9404]
ΔR	1.5309[0.1976]	0.1656[0.9548]	3.1257[0.2095]
Δulc	2.3399[0.0554]	0.1408[0.9662]	1.2142[0.5449]

System tests:	AR 1-5[125,127]	VNormality χ^2(10)	VX^2 F[480,155]
Statistics:	1.7455[0.0010]*	8.5836[0.5720]	0.359[1.000]

[1]The Values shown in brackets are the individual test's significance probability. * and ** denote as usual rejection of the corresponding null at levels of 5 and 1 per cent, respectively. VNormality and VX^2 denotes the Vector tests of normality and heteroscedaticity. For an explanation of the various test statistics the reader is referred to Chapter 14 of the PcFiml manual (Doornik and Hendry (1999)).

Table A.5:
Individual equation and system diagnostics of the unrestricted VAR of the Pooled data[1]

Equation/Tests	AR 1-5 F[5,43]	ARCH 4 F[4,40]	Normality χ^2 (2)
$\Delta a1$	1.3025[0.2808]	1.9082[0.1279]	3.4755[0.1759]
$\Delta a2$	2.2296[0.0685]	0.1409[0.9660]	0.8086[0.6674]
$\Delta pa1$	1.2499[0.3030]	0.2586[0.9027]	0.1110[0.9460]
$\Delta pa2$	0.5975[0.7020]	0.4717[0.7562]	1.0866[0.5808]
Δpw	1.0591[0.3962]	0.2464[0.9102]	0.1202[0.9417]
ΔR	2.1560[0.0768]	0.6230[0.6488]	0.7565[0.6851]
Δulc	1.9239[0.1101]	0.0361[0.9974]	1.6451[0.4393]

System tests:	VAR 1-5[245,60]	VNormality $\chi^2(14)$	VX^2 χ^2[1232]
Statistics:	1.8569[0.0026]**	6.5985[0.9491]	1172.9[0.8844]

[1]The Values shown in brackets are the individual test's significance probability. * and ** denote as usual rejection of the corresponding null at levels of 5 and 1 per cent, respectively. VNormality and VX^2 denotes the Vector tests of normality and heteroscedaticity. For an explanation of the various test statistics the reader is referred to Chapter 14 of the PcFiml manual (Doornik and Hendry (1999)).

Table A.6:
The unrestricted eigenvalues of the companion matrices

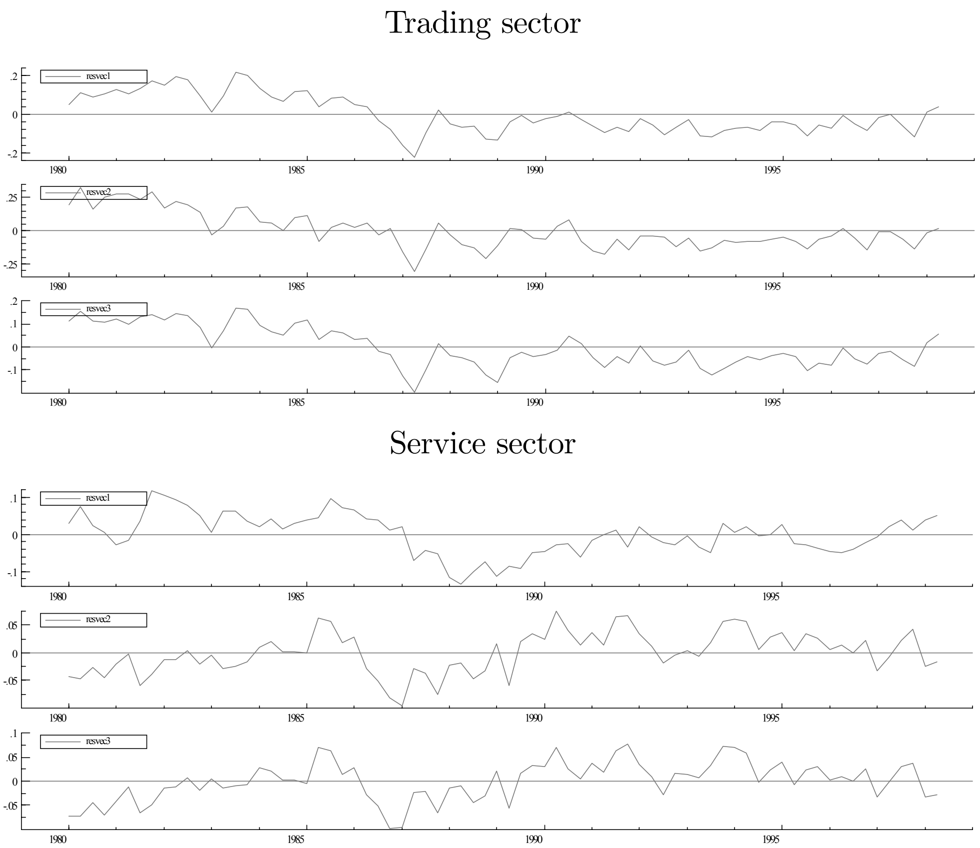
The trading sector

Eigenvalue number/Part of eigenvalue	real	complex	Modulus
1	0.9482	0.0000	0.9482
2	0.8024	±0.202	0.8275
3	-0.737	0.0000	0.7365
4	0.7272	±0.057	0.7995
5	-0.159	±0.679	0.6974
6	0.3674	±0.587	0.6921
7	0.1904	±0.452	0.4900
8	-0.3867	±0.232	0.4511
9	0.0494	0.0000	0.0494

Service sector

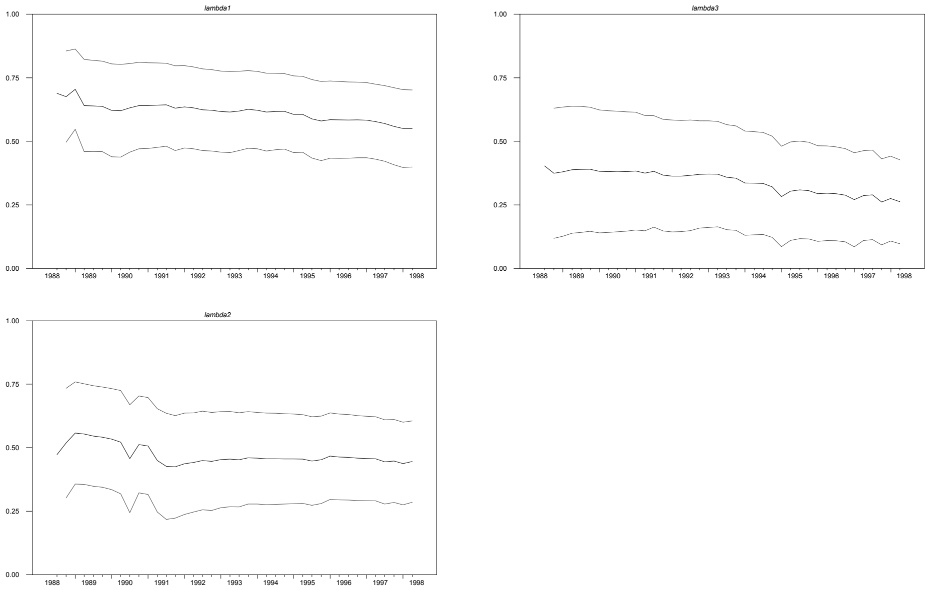
Eigenvalue number/Part of eigenvalue	real	complex	Modulus
1	0.9257	0.0000	0.9257
2	0.8204	±0.108	0.8275
3	-0.214	±0.693	0.7257
4	0.6737	±0.222	0.7094
5	-0.208	±0.615	0.6494
6	0.3473	±0.339	0.4856
7	-0.284	±0.345	0.4465
8	0.0599	±0.321	0.3264

B. Graphs

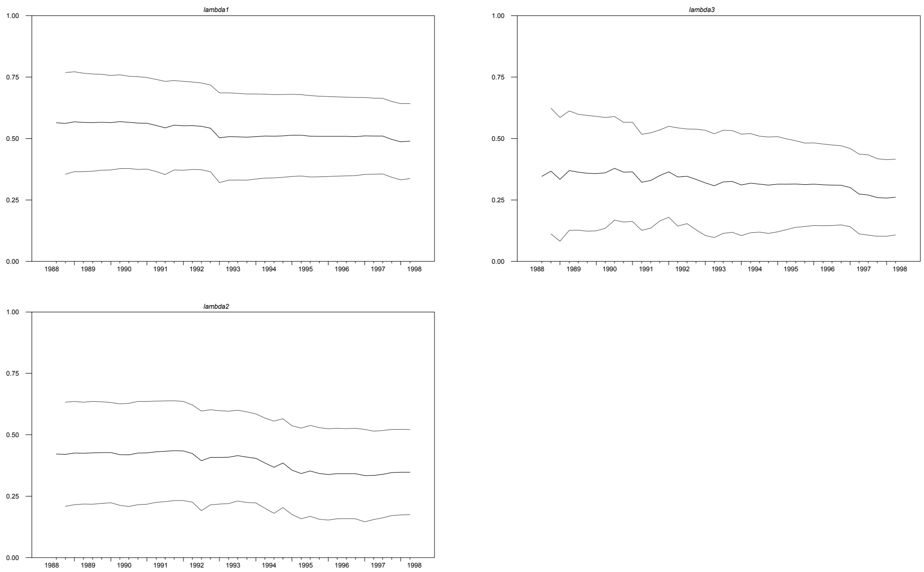


Figur B.1 Concentrated and Restricted cointegration relationships

Trading sector



Service sector



Figur B.2: Recursive eigenvalues

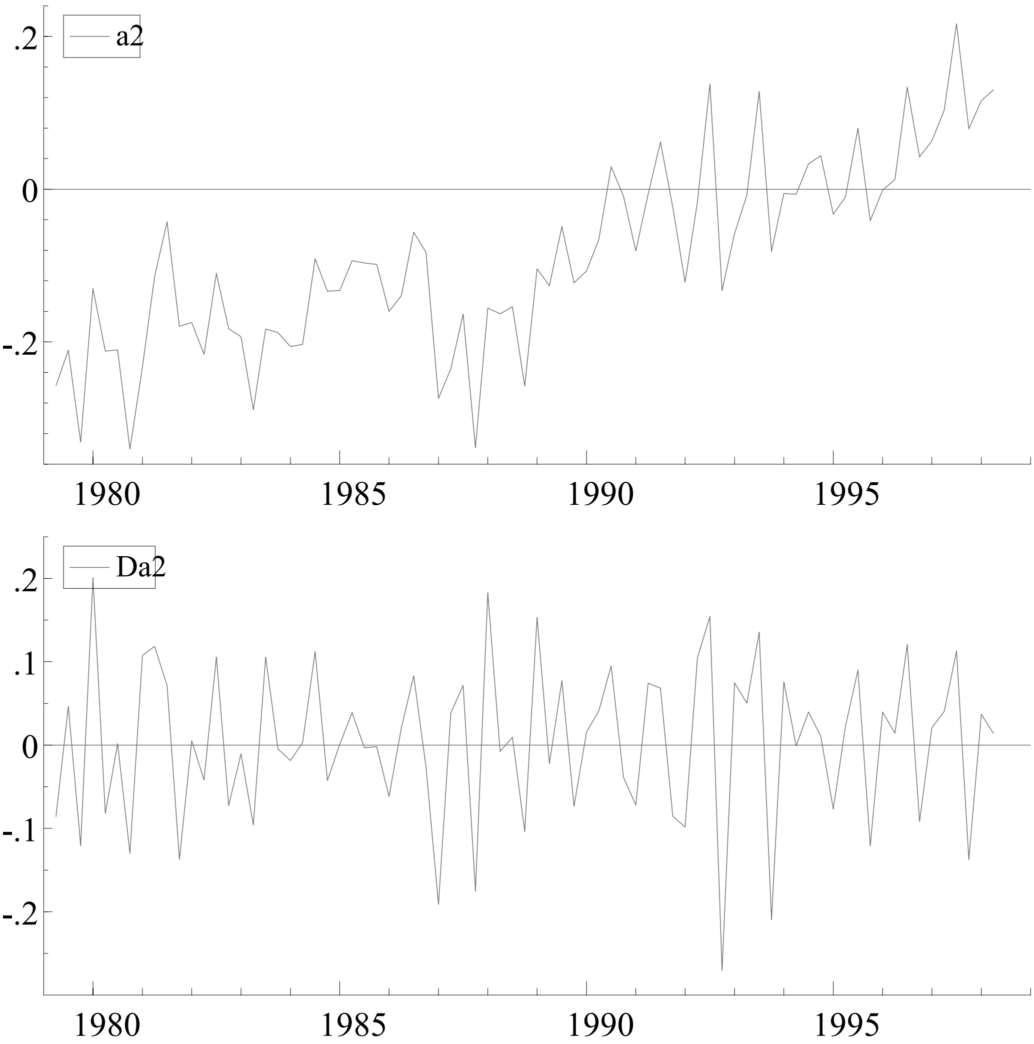


Figure B.3: Levels and first differences og exports in the service sector

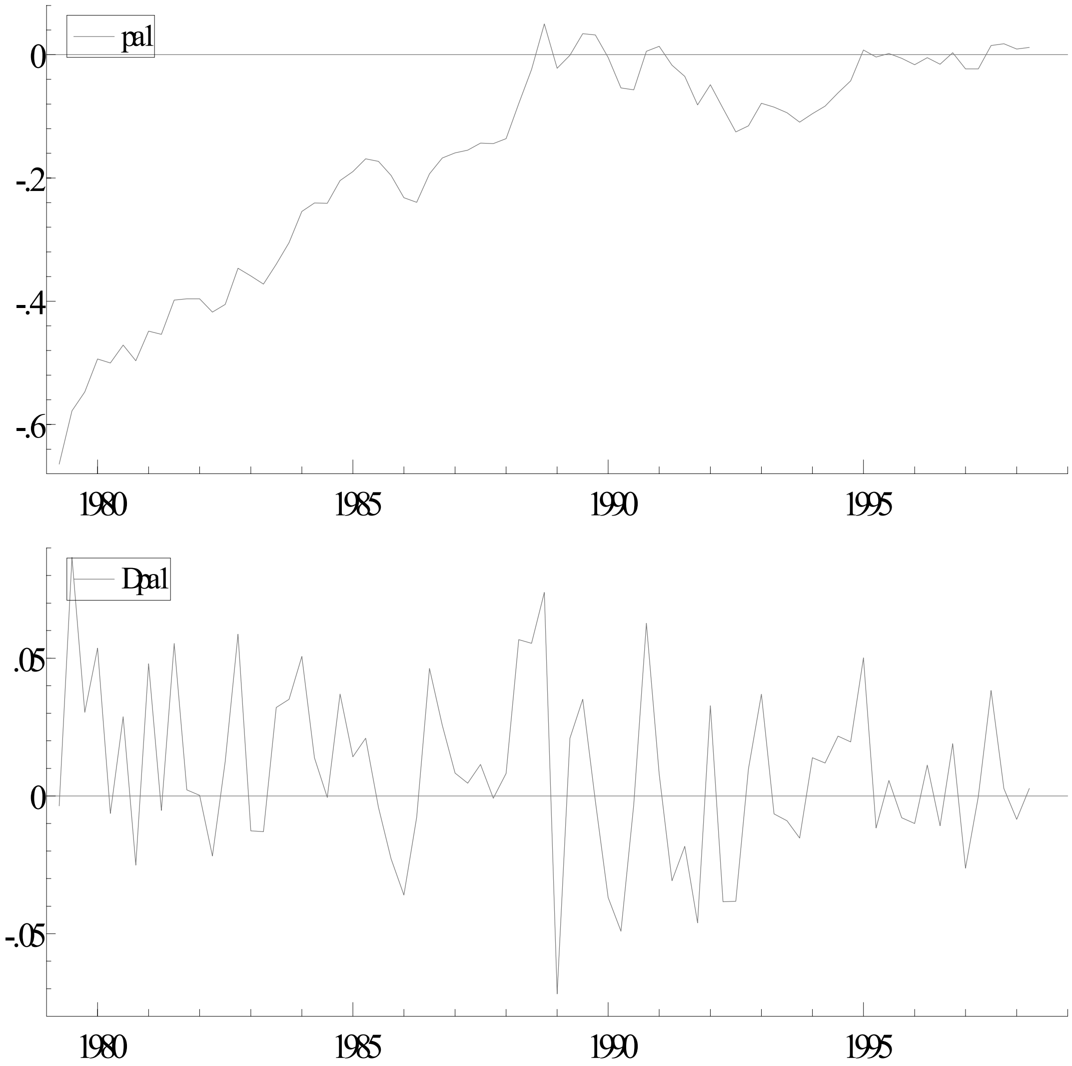


Figure B.4: Levels and first differences of export prices in trading sector

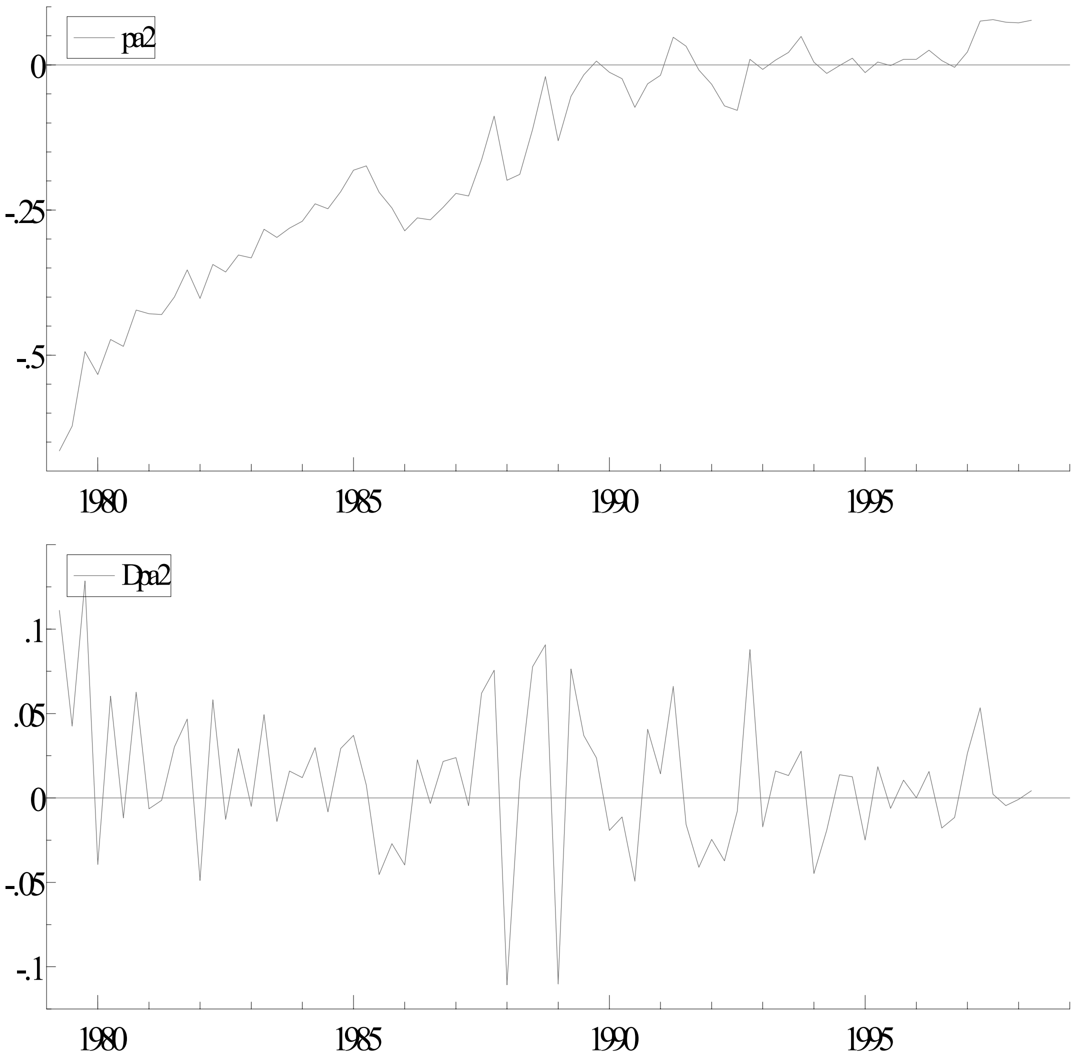


Figure B.5: Levels and first differences of export prices in the service sector

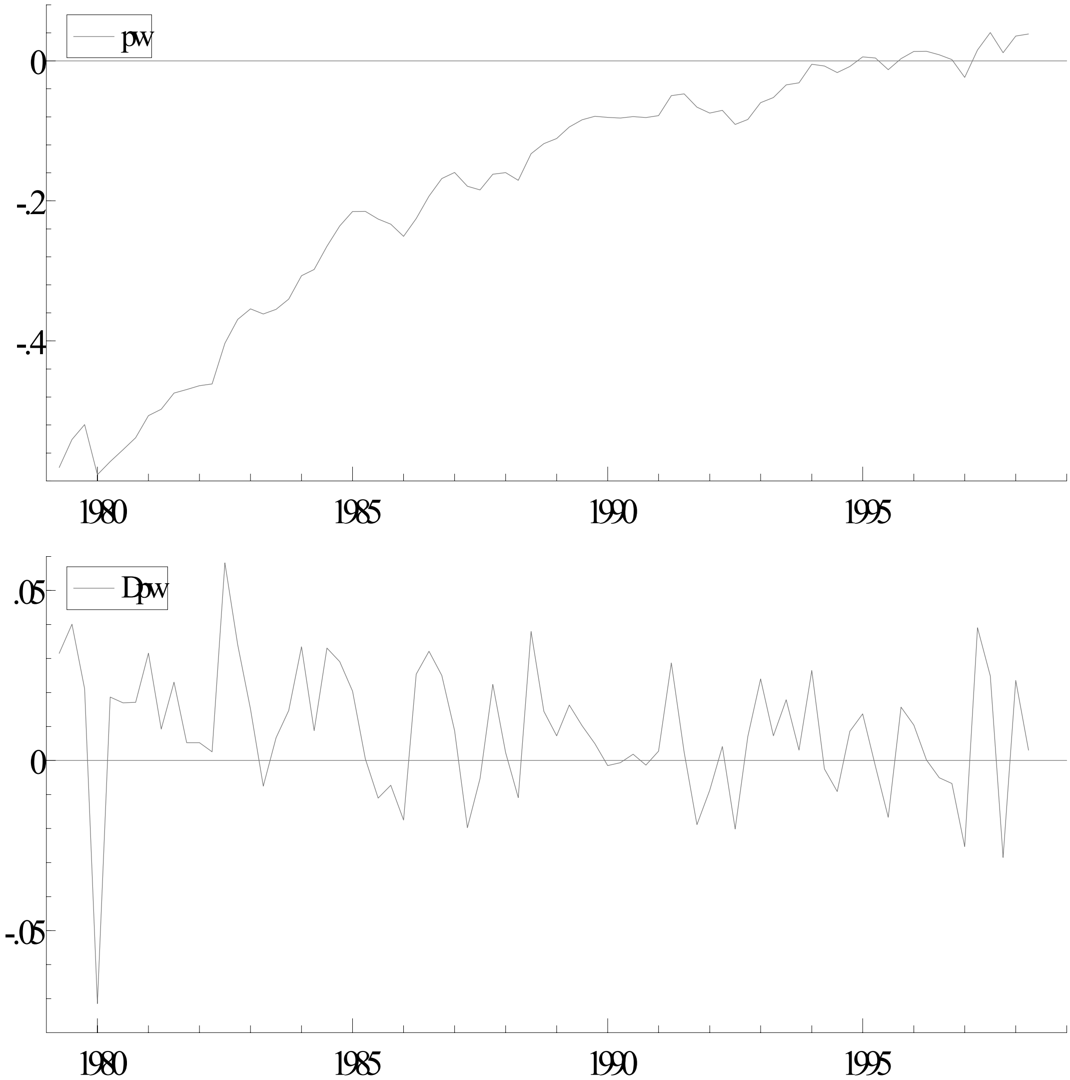


Figure B.6: Levels and first differences of world market prices

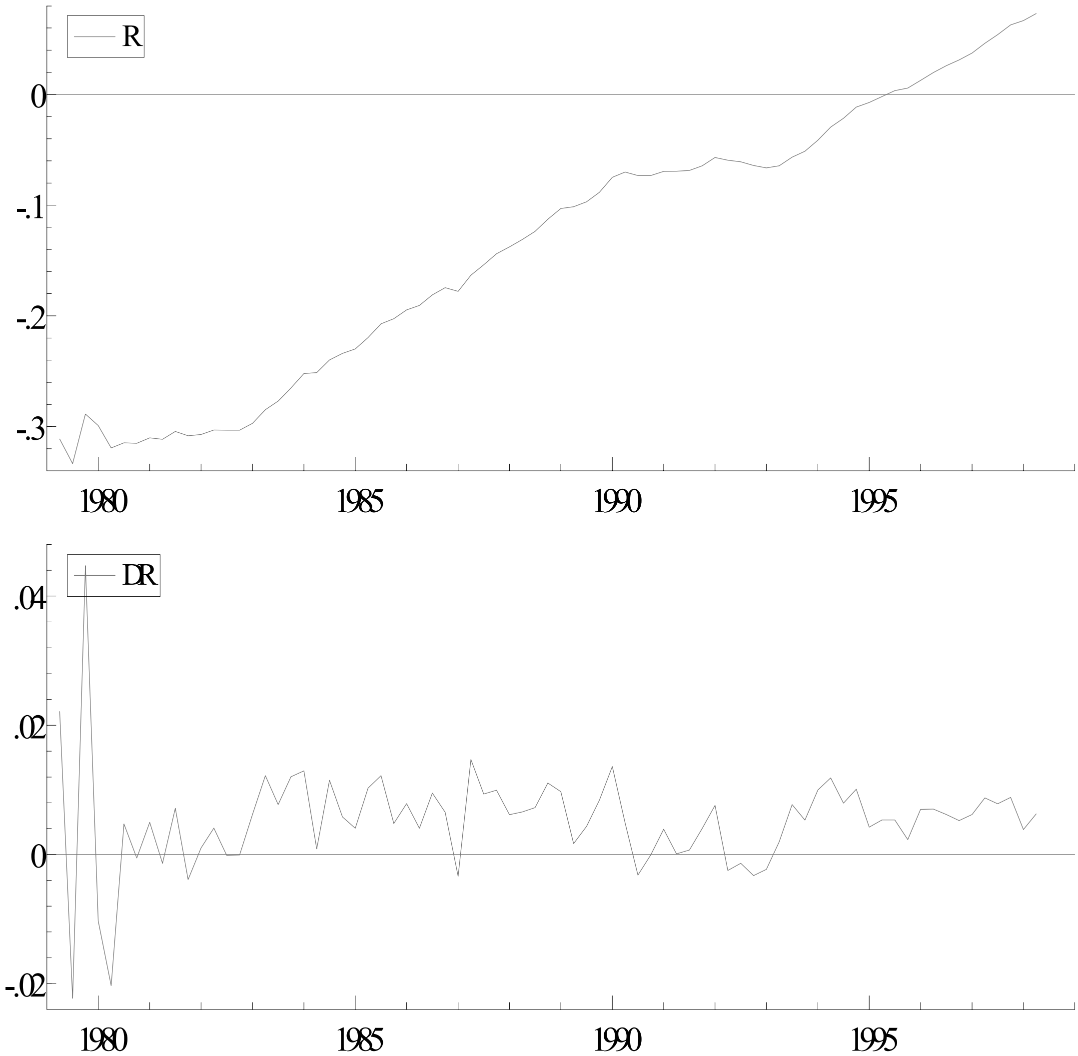


Figure B.7: Levels and first differences of the indicator of foreign real demand

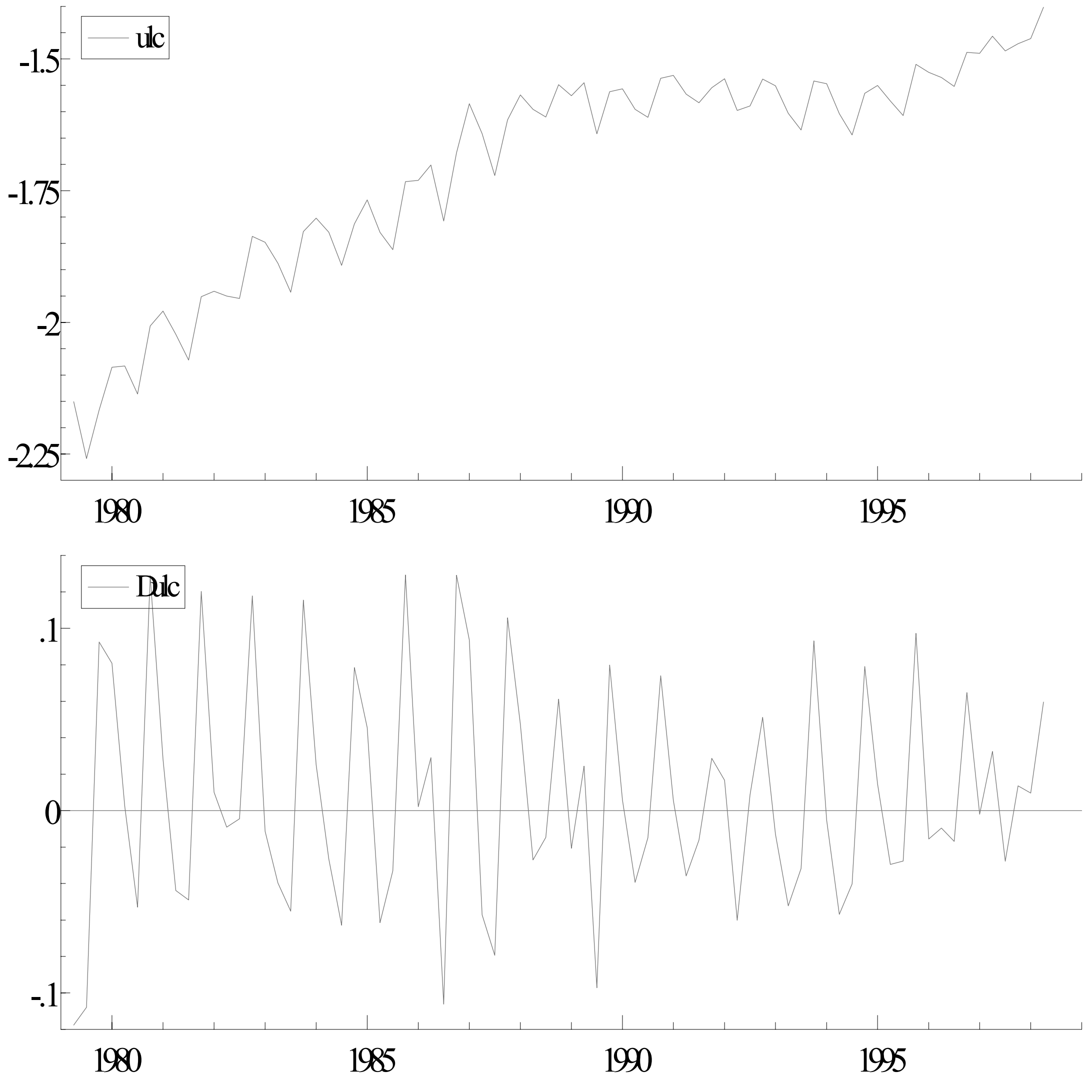


Figure B.8: Levels and first differences of unit labor costs

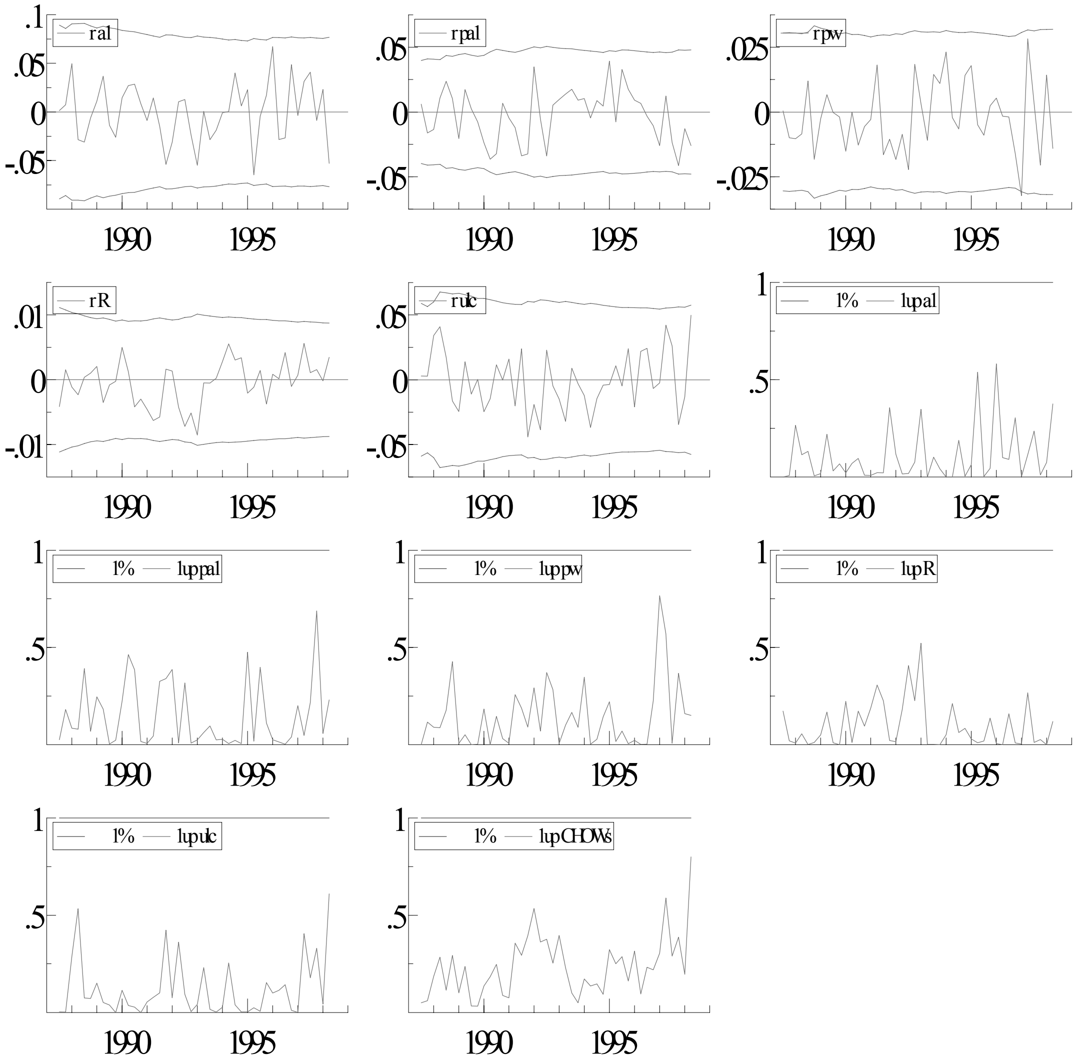


Figure B.9: Trading sector: Recursive 1-step residuals ±2 standard errors and 1-step Chow tests with one per cent critical values.

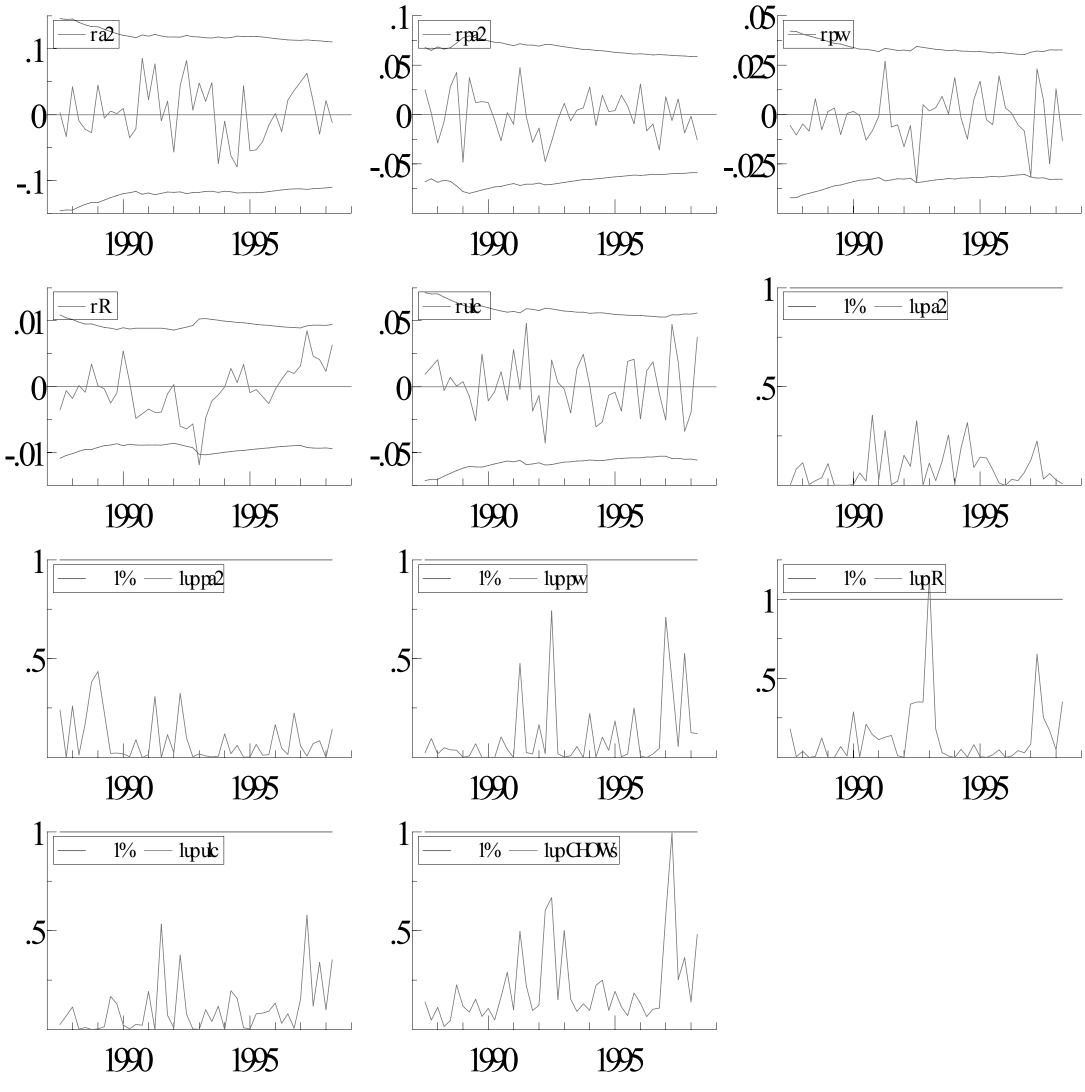


Figure B.10: Service sector: Recursive 1-step residuals ±2 standard errors and 1-step Chow tests with one per cent critical values.

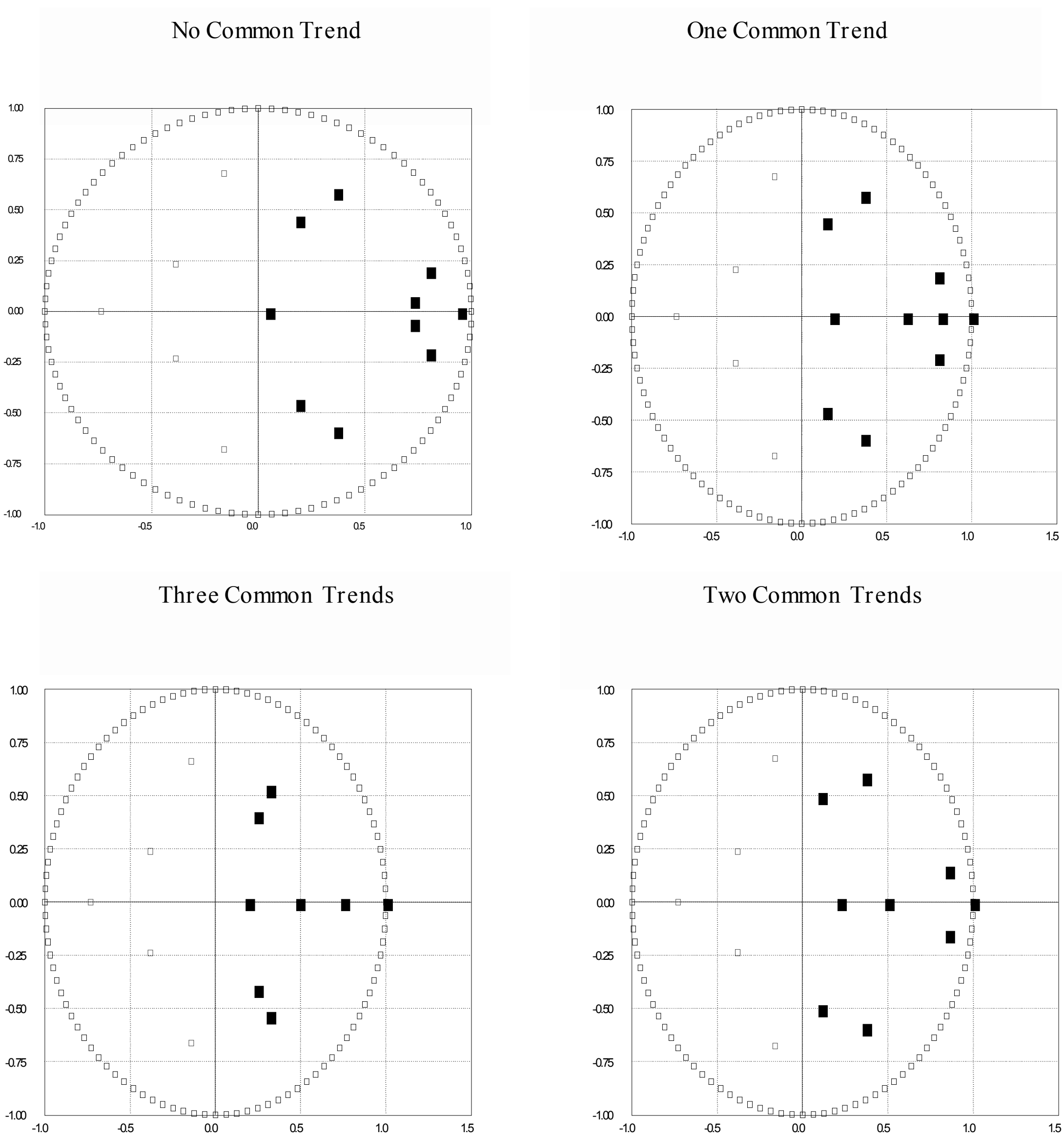


Figure B.11: Eigenvalues of the companion matrix in the trading sector and imposition of unit roots.

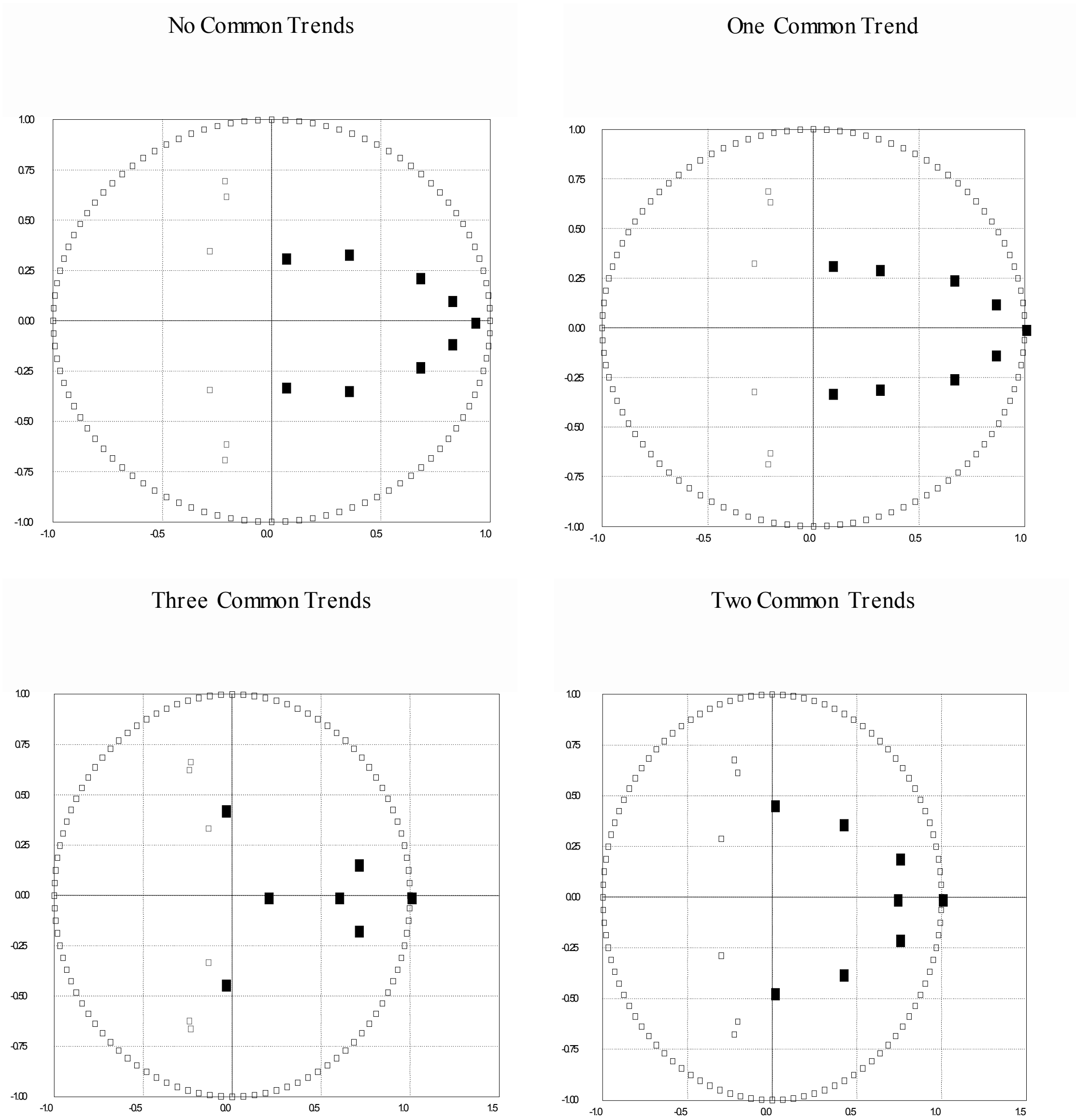


Figure B.12: Eigenvalues of the companion matrix in the service sector and imposition of unit roots.

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