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by

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The degree of independence in European goods markets: An I(2) analysis of German and Norwegian trade data [*†]

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December 30, 2004

Abstract

It is almost common knowledge that foreign trade in Europe is characterized by an acceptance of prices set by the world market. Coupled with a constant profit share in domestic sectors this makes European exports vulnerable to vagaries of international demand and prices as well as to crowding out in the wake of shocks to supply. These circumstances have been used to legitimate special measures geared towards shielding the sector from adverse shocks and general preferential treatment in the past.

In fact econometric evidence is not totally at odds with this view. However, neither exports in a large European economy like Germany nor in a small open one, like Norway, are characterized by price taking behavior. On the contrary, both nations show strong evidence of monopolistic power in the process governing external prices, implying that supply shocks to a large extent can be passed on to prices. On

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

[†] The Norwegian part of this paper is based on research undertaken in Norges Bank.

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the other hand exports seem to be heavily subject to the vicissitudes of international trade, a feature compatible with exports determined by demand ex post for prices fixed ex ante.

Keywords: Polynomial Cointegration, Higher Order Non-Stationarity Monopolistic Competition, Exports
JEL: C32, F12, F14

1 Introduction

The notorious woes of private entrepreneurs in foreign trade is a characteristic features of quotidian media in our contemporary society. Almost daily we are reminded of how vulnerable foreign trade businesses are to the vagaries of international demand and prices and how important it is to avoid excessive domestic wage claims and to promote a culture of productivity growth to avoid a general crowding out of the external sector. Now that the project for monetary union is well under way, clearly founded on neo-liberal ideas of promoting so called "level playing fields" in just about all kinds of economic areas, this is perhaps even more so. An indication of this and the influence that this kind of movements might have on decision takers were demonstrated quite recently when the more than two centuries old Smithian threat of removing all special measures for privileging exports slipped off the agenda from a EU meeting. This is a threat that if realized, certainly could have profound negative effects on the segment of foreign trade that has traditionally received preferential treatment such as export guarantees, price subsidies etc. Therefore, that such a threat gave rise to a lot of opposition is no wonder. However, after having had to listen to all kinds of mercantilistic propaganda for centuries, it is legitimate to ask how much of it might represent the truth and what might only have had the effect of bewildering the wider populace. Even though theoretical models, like the Scandinavian (Aukrust 1970, 1977), give strong support for their hypothesis, in that price taking behavior in the trading sector coupled with constant profit shares and a domestic wage leading sector lead to crowding out in wake of excessive wage claims as well as an external sector that is vulnerable to the vicissitudes of the industrialized world, it is imperative to scrutinize these points of view by looking at what a more objective source can tell us. To confront prejudiced attitudes and theories used to support them with reality, this paper undertakes an in-depth analysis of European exports based on data and an interpretational framework that theoretically encompasses the predictions of the Scandinavian model. A main motivation in this context has also been to reveal the degree of monopolistic power in the process determining European exports and export prices and thus either to confirm or to reject the hypothesis of "the law of one price". In this context it will be of particular interest to find out whether the size of the economy might also play a significant role and thus whether small open economies in Europe are more susceptible to international influence than one of the so called "Big Five". To address this issue this paper looks at two European economies, the Norwegian as a representative of a typical small economy in an European context, and the German as a representative of one of the "Big Five". As opposed to the German analysis which is based on an aggregate analysis of data for the whole economy, the Norwegian study has been undertaken based upon data for two subsectors. This has mainly been done to compare with, and further elaborate on, the results in Hammersland (2004) and in this context to particularly scrutinize the indication in this paper of a possible common I(2) trend. However, it may also be given a rationale from the perspective of looking at the status with regard to international independence of still smaller entities.

As alluded to in the above, an important aspect of this study has been to reveal a potential occurrence of higher order non stationarity. However, the attitude has been rather relaxed in this respect insofar as there has been no intention of forcing I(2)-ness upon the data. Rather, in the lack of sufficient support the approach has been more to cling to a null of I(1) than to continue along the dimension of an artificially made supposition. Additionally it may be added that as the paper intends to reveal generic properties of the underlying data generating processes the legitimacy of undertaking an I(2) analysis is deemed less urgent. However, when this is said, it must also be stressed that an I(2) analysis may be an interesting exercise to carry out even in the case one might not feel confident about its premises. If nothing else, to compare with and eventually to support the outcome of an I(1) analysis. This more pragmatic view is the preferred when interpreting the results of the I(2) analysis for Germany in Section 4.

The paper is organized as follows. Section two gives a brief review of theory used to help with the interpretation and identification of long-run relationships. The choice of monopolistic competition as an encompassing framework has not been made only because its predictions encompass the ones of the Scandinavian model and thus is convenient from the perspective of explicitly testing the claims of private entrepreneurs, but also because the theory of monopolistic competition is particularly suited to unveil potential power in the process governing prices. Dependent on whether there is evidence of a second order stochastic trend or not, the section also goes one step further and presents alternative hypothetical scenarios based on theory. Section three describes the data and their properties. A particularly important feature of this section is to reveal the existence of potential common stochastic trends of a second order. The analysis of the data then follows in the next section, Section four. The last section, Section five, seeks to conclude.

2 An encompassing theory

As alluded to in the introduction the theory of monopolistic competition has been used as an encompassing framework to help with identification and interpretation of long-run cointegrating relations. To make the approach as general as possible theory has been recast in an ex ante ex post framework making the outcome dependent on whether exporters have complete knowledge of all variables or have to make their decisions relying on plans formulated on the basis of expected quantities. In a highly stylized case the situation of the monopolist may be depicted as in Figure 1 below.[1] Ex ante the producer does not know the exact position of the demand curve and has to base his or her plans with regard to prices as well as volume on an expected demand curve, denoted A^E in the figure. Assuming that our representative exporter is a profit maximizer and has perfect knowledge with regard to costs, he will therefore plan to produce the volume where his expected marginal income equals marginal costs, A^P, and set a price which is expected to clear the market, PA^R.

In case the ex post realized quantities perfectly match the expected ones, the export volume and price relationships may be given the following stochastic log-linear representations:

$$a_t = c + \alpha(pw_t - ulc_t) + \beta R_t + \varepsilon_t \tag{1}$$

$$pa_t = c + \varphi ulc_t + (1 - \varphi)pw_t + \rho R_t + \varepsilon_t \tag{2}$$

[1] For a comprehensive treatment the reader is referred to the mathematical exposition in Hammersland (2002).

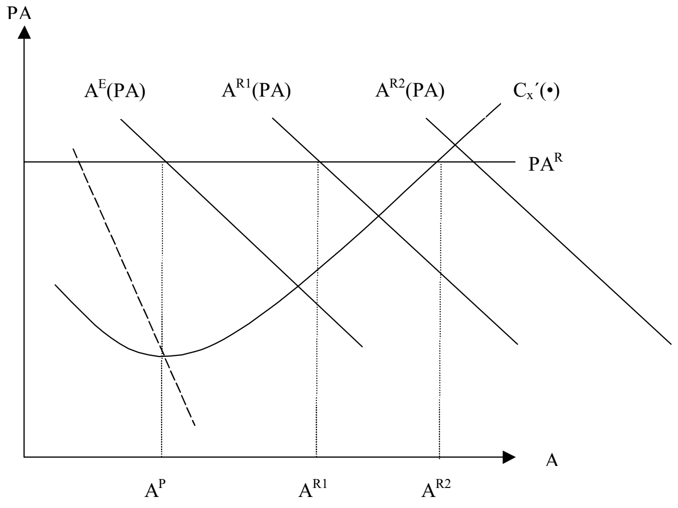

Figure 1: Monopolistic Competition and ex ante ex post determination of prices and volume.

where a_t and pa_t represent export volume and export prices while pw_t, ulc_t and R_t represent world market prices, unit labor costs and an indicator for "world" demand, respectively.[2]

However, a more likely scenario is that demand deviates significantly from the expected, ex post. The existence of long-term contracts, advertisements, price lists etc. may make it costly for the producer to deviate ex post from the ex ante decided price level. Thus, assuming that our representative monopolist is bound by its ex ante quoted price we will have to distinguish between two cases. In the first case demand is not sufficiently high to meet the volume that exporters want to produce for the fixed price ex ante. Our monopolist will therefore be rationed on the export market and the level of exports fully determined by ex post demand. In the second case ex post demand will exceed supply for the given price and exports will be given by supply. In the figure the first case is depicted by the intersection of the $A^{R1}(\cdot)$ demand curve with the horizontal curve representing the ex ante fixed export price, PA^R, while the second case is represented by the price taking level of production, A^{R2}, for which the marginal cost curve intersects with the fixed ex ante price line.[3] Following Armington (1968) assuming that demand is

[2] All variables are logarithmic transformations of the original series.

[3] Provided that the profit is positive the first case could equivalently be presented by a

specific to the producer, the demand for exports may be specified as a log linear function of the world demand indicator and the relative price ratio of export prices to world market prices. This gives us the following relationship:

$$a_t = c - \sigma(pa_t - pw_t) + \beta R_t + \varepsilon_t \tag{3}$$

In the case of a small open economy σ 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 again Hammersland (2002)), but it has also some intuitive appeal since the income effect of an increase in the export price of a small economy will be virtually negligible. Thus, the price elasticity will express the percentage change in the ratio of goods produced for export 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. It is important to note that this interpretation hinges on the fact that the economy is relatively small and that the income effect of an export price increase in a relatively big economy like i.e. Germany cannot be neglected.

2.1 Some I(1) scenarios

Economic theory contributes in an important way to our empirical analysis by providing suggestions for possible explanatory variables and also 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 nature 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 paper is therefore based on the export volume and price equations referred to above and consistency requires, in the I(1) case, that there are at least two cointegrating relationships such that all disturbances in (1), (2) and (3) are I(0), i.e. stationary variables. If we find support for two and only two cointegrating relationships, this will especially require that export

shift of the demand curve to the left. For ease of exposition this possibility has deliberately been left out in the figure.

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 world market prices to either export prices or unit labor costs. In the case of I(2) variables the analysis complicates somewhat as we in addition to have directly cointegrating vectors also might have relationships that cointegrate polynomially. This will be further dealt with in subsection 2.2 below.

To further elaborate on the implications theory consistency may have for cointegration in the case where we are dealing with I(1) variables, (2) may be reformulated as

$$pa_t - pw_t = c + \rho R_t + \varphi(ulc_t - pw_t) + \epsilon_{1t}$$

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 relative purchasing power parity (RPP) or for ρ different from 0 and $R \sim I(1)$, that the real exchange rate cointegrates with the world demand indicator. 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.[4] Looking at the two alternative volume equations, we have that this, under the assumption of β differs 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 leaves us with lots of degrees of freedom to identify theoretically consistent long-run structures in the I(1) case. A more heuristic interpretation with regard to what is consistent and not together with the possibility of multicointegrating relationships in the case of stochastic I(2) trends in the data, may in addition increase the possibility set further, examples in this respect being removal of homogeneity restrictions, exclusion of variables etc. To particularly look at the

[4] The last assertion follows from the fact that the wage share is given by $\frac{WN}{(PA)Y}$, where W denotes the nominal wage level, N the number of wage takers and Y the level of production. PA denotes as before the export price. As unit labor costs, ULC, are given by $\frac{WN}{Y}$ and export prices cointegrate with world market prices, PW, we have that the wage share can be given the equivalent long-run representation $\frac{ULC}{PW}$. As the profit share is given by $\frac{(PA)Y-WN}{(PA)Y} = 1 - \frac{WN}{(PA)Y}$, which is equal to $1 - \frac{ULC}{PW}$ in the long run, my assertion should follow.

implications multicointegration may have for the identification scheme the next subsection presents an alternative scenario based on the assumption that export prices and unit labor costs are I(2) and cointegrate to I(1), that is that they are cointegrated $CI(2,1)$.

2.2 Some I(2) scenarios

The moving average representation of the VAR when dealing with I(2) variables is in the general case given by[5]:

$$X_t = C_2 \sum_{s=1}^{t} \sum_{i=1}^{s} (\varepsilon_i + \phi D_i) + C_1 \sum_{i=1}^{t} (\varepsilon_i + \phi D_i) + C(L)(\varepsilon_t + \phi D_t) + A + Bt, \quad (4)$$

where $C_2 = \beta_{\perp 2}(\alpha'_{\perp 2}\Theta\beta_{\perp 2})^{-1}\alpha'_{\perp 2}$, $C_1 = \bar{\beta}\bar{\alpha}'\ \Gamma C_2 +\ \bar{\beta}_{\perp 1}\bar{\alpha}'_{\perp 1}\ (I - \Theta C_2) +\ \bar{\beta}_{\perp 2}$ $C_{\beta'_{\perp 2}\Delta X_t}(1)$ and $\alpha_{\perp 1}$, $\alpha_{\perp 2}$, $\beta_{\perp 1}$, $\beta_{\perp 2}$ and Θ given respectively by $\alpha_{\perp 1} = \bar{\alpha}_\perp\ \xi$, $\beta_{\perp 1} = \bar{\beta}_\perp\ \eta$, $\alpha_{\perp 2} = \alpha_\perp \xi_\perp$, $\beta_{\perp 2} = \beta_\perp \eta_\perp$ and $\Theta = \Gamma\ \bar{\beta}\bar{\alpha}'\ \Gamma + \sum_{i=1}^{k-1} i\Gamma_i$. In the expressions for the different C matrices the shorthand notation $\bar{\kappa} = \kappa(\kappa'\kappa)^{-1}$ is used for κ being equal to respectively α, $\alpha_\perp$, $\alpha_{\perp 1}$, β, $\beta_\perp$, $\beta_{\perp 1}$ and $\beta_{\perp 2}$ and the matrices α, β, ξ and η are all defined in Footnote 5. $C_{\beta'_{\perp 2}\Delta X_t}(z) = C_{\beta'_{\perp 2}\Delta X_t}(1) + \widetilde{C}_{\beta'_{\perp 2}\Delta X_t}(z)\,(1-z)$ is a convergent power series for $|z| < 1 + \delta$ for some $\delta > 0$ and constitutes the stationary part in the moving average representation of $\beta'_{\perp 2}\Delta X_t$. D_t is a deterministic term and may constitute a constant term, trend and dummies of various kinds. The coefficients A and B depend on the initial conditions and satisfy $(\beta, \beta_{\perp 1})'\,B = 0$ and $\beta' A -\ \bar{\alpha}'\ \Gamma\ \bar{\beta}_{\perp 2}\ \beta'_{\perp 2} B = 0$. For a proof the reader is referred to Johansen (1995a). Assuming that $\beta'_{\perp 2} C_1 \alpha = 0$ such that $C_{\beta'_{\perp 2}\Delta X_t}(1) = C^*_{\beta'_{\perp 2}\Delta X_t.2}\,(1)\,\alpha'_{\perp 1} + C^*_{\beta'_{\perp 2}\Delta X_t.3}\,(1)\,\alpha'_{\perp 2}$, and defining the common $I(2)$ trends as $\sum_{s=1}^{t}\sum_{i=1}^{s} u_i^{I2} = \sum_{s=1}^{t}\sum_{i=1}^{s} (u_{1i}, .., u_{s_2 i})' = \sum_{s=1}^{t}\sum_{i=1}^{s} \alpha'_{\perp 2}\varepsilon_i$ and I(1) trends as $\sum_{i=1}^{t} u_i = \sum_{i=1}^{t} \left(u_i^{I2\prime}, u_i^{I1\prime}\right)' = \sum_{i=1}^{t} ((\alpha'_{\perp 2}\varepsilon_i)', (\alpha'_{\perp 1}\varepsilon_i)')'$, where $u_i^{I2} = (u_{1i}, .., u_{s_2 i})'$ are the s_2 linear combinations of the errors that

[5] The usual assumptions apply. That is that all unit roots of the characteristic polynomial, $\left|(1-z)I - \Pi z - \sum_{i=1}^{k-1} \Gamma_i (1-z) z^i\right|$ lie at one or outside the unit circle, that the matrices Π and $\alpha'_\perp \Gamma \beta_\perp$ have reduced rank such that $\Pi = \alpha'\beta$ and $\alpha'_\perp \Gamma \beta_\perp = \xi\eta'$ for matrices α and β of dimension $p \times r$ and ξ and η of dimension $p - r \times s$, respectively, all of full rank, and finally that the matrix $\xi'_\perp \alpha'_\perp (\Gamma\ \bar{\beta}\bar{\alpha}'\ \Gamma + \sum_{i=1}^{k-1} i\Gamma_i)\beta_\perp \eta_\perp$ is of full rank, where $\bar{\beta} = \beta(\beta'\beta)^{-1}$ and $\bar{\alpha} = \alpha(\alpha'\alpha)^{-1}$.

cumulate to I(2) trends and $u_i^{I1} = \left(u_{(s_2+1)i}, .., u_{(p-r)i} \right)$ are the corresponding $p - r - s_2 = s_1$ linear combinations that cumulate to an I(1) trend, (4) may alternatively be written as

$$X_t = \widetilde{\beta}_{\perp 2} \sum_{s=1}^{t} \sum_{i=1}^{s} u_{1i}^{I2} + [C_{11}, C_{12}] \sum_{i=1}^{t} \left(u_i^{I2\prime}, u_i^{I1\prime} \right)' + C(L)(\varepsilon_t) + A + Bt \ , \quad (5)$$

where we have deliberately suppressed the deterministic term, D_t, to make the representation more appropriate for a discussion of the cointegrating properties. In (5) $\widetilde{\beta}_{\perp 2} = \beta_{\perp 2}(\alpha'_{\perp 2} \Theta \beta_{\perp 2})^{-1}$ and C_{11} and C_{12} respectively equal to $\bar{\beta} \bar{\alpha}' \ \Gamma \ \widetilde{\beta}_{\perp 2} - \bar{\beta}_{\perp 1} \bar{\alpha}'_{\perp 1} \ \Theta \ \widetilde{\beta}_{\perp 2} + \overline{\beta}_{\perp 2} C^*_{\beta'_{\perp 2} \Delta X_t.3}(1)$ and $\bar{\beta}_{\perp 1} \ (\alpha'_{\perp 1} \alpha_{\perp 1})^{-1} + \overline{\beta}_{\perp 2} C^*_{\beta'_{\perp 2} \Delta X_t.2}(1)$.

To facilitate the econometric analysis and the economic interpretation of the subsequent empirical results of the I(2) analysis for Germany, equation (6) below presents a moving average representation similar to (5) of the five dimensional system for exports implicitly defined by equations (1) to (3) above.

$$\begin{bmatrix} a_t \\ pa_t \\ pw_t \\ ulc_t \\ R_t \\ \Delta pa_t \\ \Delta ulc_t \end{bmatrix} = \begin{bmatrix} 0 \\ \gamma_{21} \\ 0 \\ \gamma_{41} \\ 0 \\ 0 \\ 0 \end{bmatrix} \sum_{s=1}^{t} \sum_{i=1}^{s} u_{1i} + \begin{bmatrix} c_{11} & c_{12} & c_{12} \\ c_{21} & c_{22} & c_{23} \\ c_{31} & c_{32} & c_{33} \\ c_{41} & c_{42} & c_{43} \\ c_{51} & c_{52} & c_{53} \\ \gamma_{21} & 0 & 0 \\ \gamma_{41} & 0 & 0 \end{bmatrix} \begin{bmatrix} \sum_{i=1}^{t} u_{1i} \\ \sum_{i=1}^{t} u_{2i} \\ \sum_{i=1}^{t} u_{3i} \end{bmatrix} + X_0 \quad (6)$$

In (6), $X_0 = \widetilde{C}\ (L) + A + Bt$ in (5), and the presence of three common trends out of which one is of a second order, is made upon the anticipation of subsequent empirical results. (6) also implies that only export prices and unit labor costs are $I(2)$ and that a linear combination of the two reduces the order of non-stationarity from two to one. An implication of (6) is that neither export prices nor unit labor costs can separately enter into a cointegrating relationship. This implies particularly that (6) rules out the possibility of a purchasing power parity (PPP) relationship or perhaps more correctly denoted, the hypothesis of "one price," already at the outset. Assuming that exports ex post do not deviate from the ex ante planned level so that exports are given by a variant of (1), this seems either to imply that the concept of unit labor costs must be in real terms or that unit labor costs do not enter into the cointegrating relationship at all. The latter case could perhaps more likely be taken to mean that exports are determined exclusively by demand ex post without changes in own prices affecting the output. Assuming that the long-run export volume relationship constitutes a directly cointegrating relationship, this would therefore imply that the linear relation $a_t - \omega_1 pw_t - \omega_2 R_t$ is stationary. For this to be the case we must have that the vector $(1, 0, -\omega_1, 0, -\omega_2, 0, 0)$ is a cointegrating vector and hence that it is orthogonal to all the vectors in the matrices in (6). Thus we must have that

$$c_{1j} - \omega_1 c_{3j} - \omega_2 c_{5j} = 0, \;\; j = 1, 2, 3.$$

for weights ω_1 and ω_2. The first situation, however is slightly more elaborate and implies that the linear combination $a_t - \omega_1 pw_t - \omega_2 R_t + \omega_3(\gamma_{21} ulc_t - \gamma_{41} pa_t)$ is stationary. This implies that the vector

$$(1, -\omega_3\gamma_{41}, -\omega_1, \omega_3\gamma_{21}, -\omega_2, 0, 0)$$

must be a cointegrating one and hence that also this is orthogonal to the matrices in (6) so that

$$c_{1j} - \omega_3\gamma_{41} c_{2j} - \omega_1 c_{3j} + \omega_3\gamma_{21} c_{4j} - \omega_2 c_{5j} = 0, \;\; j = 1, 2, 3,$$

where ω_3 is the weight given to the cointegrating $CI(2,1)$ linear combination of unit labor costs and export prices. Assuming a polynomially cointegrating export price relationship the most plausible candidate would be that the $CI(2,1)$ linear combination of export prices and unit labor costs cointegrates with world market prices, the indicator of foreign demand and growth in either export prices or unit labor costs. This implies that the linear combination $\omega_3(\gamma_{41} pa_t - \gamma_{21} ulc_t) - \omega_1 pw_t - \omega_2 R_t -\ \bar{\alpha}'\ \Gamma\ \bar{\beta}_{\perp 2}\ \beta'_{\perp 2} x$ is stationary, x representing either relative growth in export prices, Δpa_t, or in unit labor costs, Δulc_t. Thus the cointegrating vector is given by

$$(0, \omega_3\gamma_{41}, -\omega_1, -\omega_3\gamma_{21}, -\omega_2, -\ \bar{\alpha}'\ \Gamma\ \bar{\beta}_{\perp 2}\ \beta'_{\perp 2}, 0)$$

or

$$(0, \omega_3\gamma_{41}, -\omega_1, -\omega_3\gamma_{21}, -\omega_2, 0, -\ \bar{\alpha}'\ \Gamma\ \bar{\beta}_{\perp 2}\ \beta'_{\perp 2})$$

depending on which of the growth rates enters in the cointegrating relationship, and its orthogonality property implies that

$$\omega_3\gamma_{41}c_{2j} - \omega_3\gamma_{21}c_{4j} - \omega_1 c_{3j} - \omega_2 c_{5j} = 0,\ j = 2,3$$

and

$$\omega_3\gamma_{41}c_{21} - \omega_3\gamma_{21}c_{41} - \omega_1 c_{31} - \omega_2 c_{51} - \bar{\alpha}'\,\Gamma\,\bar{\beta}_{\perp 2}\,\beta'_{\perp 2}\gamma_{j1} = 0,\ j = 2 \text{ or } 4$$

Normalizing on export prices such that $\omega_3 = \frac{1}{\gamma_{41}}$, these conditions might equivalently be expressed as $c_{22} = \frac{\gamma_{21}}{\gamma_{41}}c_{42} + \omega_1 c_{32} + \omega_2 c_{52}$, $c_{23} = \frac{\gamma_{21}}{\gamma_{41}}c_{43} + \omega_1 c_{33} + \omega_2 c_{53}$ and that $c_{21} - \frac{\gamma_{21}}{\gamma_{41}}c_{41} - \omega_1 c_{31} - \omega_2 c_{51} - \bar{\alpha}'\,\Gamma\,\bar{\beta}_{\perp 2}\,\beta'_{\perp 2}\gamma_{j1} = 0$, for $j = 2$ or $j = 4$. Anticipating the result of the subsequent analysis it is particularly interesting to look at the case where $\gamma_{41} = \gamma_{21}$ and $\omega_1 = \omega_2 = 0$. That is, neither world market prices nor world demand enter into the multicointegrating relationship and the spread between unit labor costs and exports prices is cointegrated $C(2,1)$. If so, the implied restrictions must be that $c_{22} = c_{42}$, $c_{23} = c_{43}$ and finally that $c_{21} - c_{41} - \bar{\alpha}'\,\Gamma\,\bar{\beta}_2\,\beta'_2\gamma_{j1} = 0$.

3 Data and time series properties

Before presenting the results of the cointegration analysis, I will in this section 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. I will concentrate on commenting on the German data as data for Norwegian exports have been extensively commented on in Hammersland (2004).

The econometric analyses are based on quarterly seasonally unadjusted data for the period 1960 (1) to 1998 (4) in the case of Germany and for 1979 (2) to 1998 (2) for Norway.[6] The data set consists of observations on the following empirical proxies of the theoretical counterparts:[7]

[6] As unadjusted data were not available for unit labor costs of Germany these have been included sesonally adjusted. However, this is deemed less serious as unit labor costs should not show a strong seasonal pattern.

[7] In the whole paper 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 indicator of world demand where R also indicates a logaritmic transformation.

aj	Aggregate export volume for Germany (a), the trading sector ($j = 1$), and the service sector ($j = 2$) of Norway.
paj	Aggregate export prices for Germany (pa), the trading sector ($j = 1$), and the service sector ($j = 2$) of Norway.
pw	World market prices in domestic prices. A weighted average of GDP-deflators for the nine, respectively ten, most significant trading partners of Norway and Germany.
ulc	Unit labor costs
R	World demand indicator. A weighted average of GDP, respectively imports of the same trading partners as for world market prices of Norway and Germany.

The weights used to create the world demand, world price and an effective exchange rate, the last one used to convert world market prices into their domestic currency equivalents, have been the average share of total exports exported to individual trading partners.[8] Plots of all German series, both levels and first differences, are shown in the graphical part of the appendix. Based on graphical inspection there is scarce evidence of $I(2)$-ness in the data. The series closest in agreement with such a description perhaps are export prices and unit labor costs. However, to further investigate the issue of whether the data are $I(1)$ or $I(2)$ one will have to formally determine the orders of integration by thorough testing. In the appendix I therefore anticipate events somewhat by first presenting the results of testing for stationarity of order one within a multivariate framework based on the methodology developed by Johansen for estimation and identification of cointegrating relationships (Johansen (1988), (1995a)). The model used is the same as in Section 4 when identifying the cointegrating relationships when data are supposed to lie in the I(1) space. The tests are therefore

[8] In the German data the weights used are based on data on exports to individual countries for the period 1960-1980. It is a weakness that these weights are old and that for some countries they turn out to be higly unstable. Notably for the US where the weights show a significantly increasing tendency.

conditional on two cointegrating relationships and differ in a very important respect from univariate Dickey-Fuller tests by testing the null of stationarity against non-stationary alternatives. The test statistics are the LR-tests of restrictions on the cointegrating space and imply particularly testing the hypotheses that one of the cointegrating vectors consists of zero coefficients for all variables except for the coefficient of the variable we want to test for stationarity. In Table 9 the coefficient of the restricted trend has also been left unrestricted implying in fact that we are testing the null of trend-stationarity versus non-stationary alternatives. 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 in the same appendix, Table 10, chosen also to present the results of univariate Augmented Dickey-Fuller (ADF) tests. To avoid the problem of nuisance parameters in the DGP these univariate ADF-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 in addition included seasonal dummies. To avoid the problem of having to deal with a possible quadratic trend under the alternative, 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 be able to fully address the issue of higher order integration, however, I have finally chosen to present a full analysis of the cointegrating indices based on the multivariate two-step I(2) procedure developed by Johansen(1995b).

The multivariate test statistics for stationarity in Table 9 of the appendix strongly suggest rejection of the null of stationarity for all the variables. This is further confirmed by looking at the Augmented Dickey-Fuller tests of the subsequent table, Table 10, which are not able to reject the null of a non-stationary I(1) process for any of the variables. With regard to a possible second order trend the results of univariate testing are far from promising. All tests reject the null of a second order trend to a level far below one percent. On the other hand looking at the multivariate tests of the cointegration indices as reported in Table 1 below, the tests clearly indicate that there is a second order trend in the information set, however. These multivariate tests have been carried out by specifying a five dimensional VAR of order five, where a drift term has been restricted to lie in the cointegrating space and a constant restricted not to induce quadratic trends in the process. In addition to centered seasonal dummies the specification involves two unrestricted dummies out of which one implies a shift in the constant term and the other a transitory shift in levels.[9] Thus, our model does not contain intervention dummies that cumulate to trends in the DGP and therefore potentially might invalidate inference based on standard asymptotic tables (Johansen and Nielsen (1994)). We should therefore be able to proceed by using the asymptotic tables for the I(2) model of Paruolo (1996). The procedure starts testing from the upper left corner of Table 1 at a null of five common I(2) trends against the alternative of less than or equal to full rank. If this first test statistics is bigger than the 95 percent fractile given in italics under each statistics, the procedure continues towards the right reducing the number of common I(2) trends under the null by one. This process goes on to the end of the first row and proceeds similarly row-wise from left to right until the test statistics is insignificant to a level of five percent, in which case the cointegration indices are jointly identified by a rank equal to r, the number of I(1) trends under the null, s_1, and the number of I(2) trends given by $p - r - s_1 = s_2$. In Table 1, 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, implying that the number of common I(2) trends are equal to $p - r - s_1 = 5 - 2 - 2 = 1$. This finding is in accordance with the suggested scenario of Section 2 and will form the basis of the analysis to come.

4 The analyses

In Section two I presented the moving average representation of the I(2) model in the general case when no restrictions are imposed on the deterministic terms. In general, if X_t is integrated of order two and a linear regressor as well as a constant is included unrestrictedly in the model, this would allow for cubic as well as quadratic trends in the process governing the data while in the case X_t is integrated of order one an unrestricted trend term would generate a quadratic trend. As these peculiarities are not very likely to prevail in practice we will have to place restrictions on the deterministic components of the model. In the I(1) model this implies that the linear regressor has been restricted to lie in the cointegrating space while the constant, seasonals and dummies have been left unrestricted. This restriction has also the advantage of making inference similar with respect to the level

[9] The reader is referred to Section 4 and footnote 11 for a more complete presentation of the model.

Table 1: The trace test of cointegrating indices for German exports

p-r	r			$S_{r,s}$			Q(R)
5	0	**309.20**	**227.97**	**176.77**	**144.99**	**122.67**	**114.98**
		198.2	167.9	142.2	119.8	101.5	87.2
4	1		**187.73**	**126.90**	**95.56**	**78.71**	**68.89**
			137.0	113.0	92.2	75.3	62.8
3	2			**102.75**	**70.60**	**45.98** *	**42.53**
				86.7	68.2	53.2	42.7
2	3				**49.72**	**23.68**	**17.78**
					47.6	34.4	25.4
1	4					**28.66**	**5.32**
						19.9	12.5
p-r-s		5	4	3	2	1	

[1)]The figure in italics under each test statistic is the 95 per cent fractile as tabulated by Paruolo(1996). The preferred outcome of the sequential testing is marked by a star.

[2)]The multivariate tests have been carried out by specifying a five dimensional VAR for the variables a, pa, pw, ulc and R of order five. A drift term has been restricted to lie in the cointegrating space and a constant included in such a way that it does not induce quadratic trends in the process. In addition to centered seasonal dummies the specification also involves two unrestricted dummies of which one implies a shift in the constant term and the other a transitory shift in levels.

and linear trend parameters of the DGP. As for the I(2) models the constant regressor has in addition been restricted not to generate a quadratic trend and in such a way that it allows for linear trends in all linear combinations of X_t. Keeping this in mind and suppressing the deterministic terms, the general error correction form of the I(2) model

$$\Delta^2 X_t = \alpha\beta' X_{t-1} - \Gamma\Delta X_{t-1} + \sum_{i=1}^{k-2} \psi_i \Delta^2 X_{t-1} + \varepsilon_t$$

may be given the equivalent representation to be used as reference in the following when interpreting the results of the $I(2)$ analyses:[10]

$$\begin{aligned}\Delta^2 X_t &= \widetilde{\alpha}_1 \widetilde{\beta}_1' X_{t-1} + \widetilde{\alpha}_2 (\widetilde{\beta}_2' X_{t-1} + \widetilde{\delta}\beta_{\perp 2}' \Delta X_{t-1}) \\ &\quad + \widetilde{\kappa}\beta' \Delta X_{t-1} + \widetilde{\kappa}\beta_{\perp 1}' \Delta X_{t-1} + \sum_{1=1}^{k-2} \psi_i \Delta^2 X_{t-1} + \varepsilon_t\end{aligned} \qquad (7)$$

In (7) $\widetilde{\beta}_1' X_{t-1}$ denotes the directly cointegrating $CI(2,2)$ relationships while $\widetilde{\beta}_2' X_{t-1} + \widetilde{\delta}\beta_{\perp 2}' \Delta X_{t-1}$ are the multicointegrating relationships. Their respective loading matrices are denoted by $\widetilde{\alpha}_1$ and $\widetilde{\alpha}_2$. More details with regard to the specified VARS will be given in connection with the country specific analysis below.

4.1 Germany

The analysis of German exports is based on a five dimensional VAR of order five. In addition to the restrictions alluded to above with regard to deterministic regressors, the model has been specified with two dummies, respectively, D7334 and D741. While the first one effectuates a transitory shift in the levels of the series in the third quarter of 1973, the second imposes a permanent

[10]In (7) the different parameters as functions of the original parameters are given by: $\widetilde{\kappa} = [\alpha\psi\overline{\tau} + (\alpha(\alpha'\Omega^{-1}\alpha)^{-1}\alpha'\Omega^{-1}\overline{\alpha}_\perp + \Omega\alpha_\perp(\alpha_\perp'\Omega\alpha_\perp)^{-1})\kappa']$, $\widetilde{\beta}_1' = \delta_\perp'\rho'\tau'$, $\widetilde{\beta}_2 = \delta'\rho'\tau'$, $\widetilde{\alpha}_1 = \alpha\overline{\delta}_\perp$, $\widetilde{\alpha}_2 = \alpha\overline{\delta}$ and $\widetilde{\delta} = \delta'\delta$, where $\rho' = (I_{r\times r}, 0_{r\times s_1})$, $\tau = (\beta, \beta_{\perp 1})$, $\psi' = -\overline{\alpha}'\Gamma$, $\delta = \psi'\overline{\tau}_\perp$ and $\kappa' = -\alpha_\perp'\Gamma\overline{\tau}$. The result follows by straightforward use of the identity $\alpha(\alpha'\Omega^{-1}\alpha)^{-1}\alpha'\Omega^{-1} + \Omega\alpha_\perp(\alpha_\perp'\Omega\alpha_\perp)^{-1}\alpha_\perp' = I$ in combination with the orthogonal projection identities $P_\Theta + P_{\Theta_\perp} = \Theta\left(\Theta'\Theta\right)^{-1}\Theta' + \Theta_\perp\left(\Theta_\perp'\Theta_\perp\right)^{-1}\Theta_\perp' = I$ for $\Theta = \tau$, δ, α.

shift in the constant term in the first quarter of 1974.[11] The single equation diagnostics of the system are given in the upper part of Table 11 of the appendix and except for some hardly significant signs of conditional heteroscedasticity in the process governing German exports, they seem to fulfil most requirements for a congruent representation of a DGP. The system tests indicate though a marginal problem with regard to normality. However, this seems mainly to originate from excess kurtosis and based on the results of Gonzalo (1994) is therefore deemed less serious. Also, even though there seem to be some problems in the last part of the seventies/early eighties, the recursively estimated Chow tests of the appendix do not reveal any ominous signs of parameter instability. Thus our VAR should be a good starting point for identification and estimation of cointegrating relationships

The ordinary trace test statistics of Table 2 clearly identify the presence of at least two cointegrating vectors. We also note that two of the estimated eigenvalues of the $\Pi - I$ matrix are quite big consistent with three cointegrating vectors. However, the eigenvalues of the companion matrix indicate that imposing two unit roots leaves an unrestricted root with a large modulus, 0.965, in the model. Three unit roots are consistent with either two cointegrating vectors or in the case of three cointegrating vectors, that one of the three common trends is a trend of second order, I(2). As seen from the Table, imposing a third unit root still leaves us with an unrestricted eigenvalue with a relatively large modulus, 0.942. As a unit root in the characteristic polynomial that belongs to an $I(2)$ trend cannot be removed by lowering the number of cointegrating vectors, this is certainly an indication of $I(2)$-ness. However, as mentioned in the introduction the aim of this analysis is to unveil generic properties of the DGP and not to design a model that may better explain certain local phenomena. Before embarking on the more complicated $I(2)$ analysis I have therefore chosen first to undertake an analysis based on the more plausible presumption of $I(1)$. The outcome of this analysis will then function as a basis for comparison with other analyses and particularly with the one made on the assumption of a second order trend in the data.

[11] To be more specific this means that the two dummies are model specific. In the $I(1)$ case this means that the dummy, $D7334$, is given by (.....,0,1,-2,1,0........) such that when it is cumulated ones it assumes the values 1 in 1973 (3), -1 for 1973 (4) and nil otherwise. The dummy $D741$ is a blip dummy equalling 1 in 1974 (1) and nil otherwise and cumulates therefore to a level shift in 1974 (1). The corresponding dummies in the $I(2)$ models are given respectively by (....,0,1,-3,3,-1,0,....) and (....,0,1,-1,0,.....) such that when they are cumulated twice they assume the same values as for the cumulated I(1) case.

System: a, pa, pw, R, ulc.
Deterministic part: Restricted Trend, Unrestricted Constant, Centered Seasonals and the dummies D741 and D7334.
VAR order: 5. Effective Sample period: 1961 (2)-1998 (4)

Common trends imposed:	Modulus of the eigenvalues of the companion form and the estimated eigenvalues of the Π matrix:				
0	0.986	0.986	0.901	0.901	0.889
1	1	0.988	0.988	0.895	0.895
2	1	1	0.965	0.965	0.885
3	1	1	1	0.942	0.942
4	1	1	1	1	0.914
Eigenvalues of the Π matrix:	0.941	0.918	0.847	0.781	0.726

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

Null	Alternative	Test Statistics	95% Critical values
r=0	r$\leq$5	133 **	87.3
r$\leq$ 1	r$\leq$5	84.6**	63.0
r$\leq$ 2	r$\leq$5	47.3*	42.4
r$\leq$ 3	r$\leq$5	22.16	25.3
r$\leq$ 4	r$\leq$5	9.213	12.3

Table 2: Rank tests and modulus of eigenvalues of the companion form for the Germant system of exports

4.1.1 The I(1) case

The results of the $I(1)$ analysis based upon the presumption of two cointegrating vectors are given in Table 3 below. The first thing to notice is the lack of unit labor costs and the relatively strong effects of changes in international demand and prices in the volume equation. Furthermore, the export price equation is a pure markup relationship over unit labor costs and implies that effects from international conditions play a negligible role in the long run. If this is correct it means that German exporters when setting their prices, almost act as though the rest of the world does not exist and feel free to pass increases in unit labor costs on to prices without hardly squinting at international demand and price conditions. On the other hand, exports seem to depend heavily on international factors and even though the relationship is not of an Armington type, it suggests, as in the Norwegian case, an ex post interpretation of its origin. Why it is the nominal and not the real world market price denoted in units of the export price that enters in the equation is a conundrum. However, it may indicate that German exporters have some kind of money illusion.

With regard to the loadings the most puzzling artifact is perhaps the strongly significant positive error correction of the discrepancy of the aggregate export price from its long-run solution in the equation for unit labor costs. Even though such an effect may seem a little bit far fetched, it may be explained if expected (as opposed to unexpected) hikes in export prices are perceived as the result of excessive wage claims made by trade unions in their tripartite negotiations with the employers associations. Another puzzling effect though not significant, is the corresponding negative weight in the equation for foreign demand. Otherwise, we see that there is significant error correction in the export price equation from deviations of export prices from their long-run solution and a close to significant error correction in the volume equation from the deviation of exports from its equilibrium level.[12]

As alluded to above a typical sign of higher order common trends is that lowering the number of cointegrating vectors does not remove an additional unit root associated with $I(2)$-ness. Another sign is that the graphs of the concentrated and non-concentrated cointegration relations exhibit signif-

[12]In this context it is worth mentioning that a simultaneous reduction, incorporating the identifying restrictions of Table 3 together with zero restrictions on all loadings except for α_{11}, α_{21}, α_{22}, and α_{41}, is marginally significant to a level of five per-cent with a p-value equal to 0.0465.

Restricted cointegrated vectors in the $I(1)$ model

	a	pa	pw	ulc	R	$Trend$
$\beta'_{.1}$	1	0	-2.9425	0	-3.5469	0.0532
			(0.412)		(0.341)	(0.0065)
$\beta'_{.2}$	0	1	0	-1	0	0.0015
						(0.0002)

Loading matrix

$\alpha_{1.}$	-0.0382 (0.0255)	-0.0645 (0.0638)
$\alpha_{2.}$	-0.0037 (0.0037)	-0.0349 (0.0093)
$\alpha_{3.}$	0.0285 (0.0118)	0.0142 (0.0295)
$\alpha_{4.}$	-0.0456 (0.0104)	0.0834 (0.0261)
$\alpha_{5.}$	0.0182 (0.0192)	-0.0828 (0.0481)

LR-test, rank=2: Chi^2(4) = 1.1979 [0.8784]

Table 3: The restricted cointegrated vectors, loadings and test for the overidentifying restrictions in the German model of exports

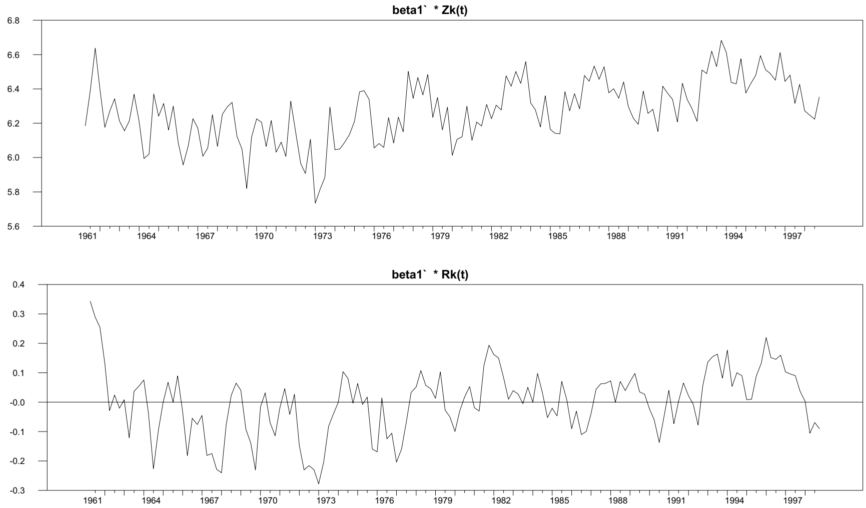

Figure 2: Concentrated, $beta1'Rk(t)$, and not-concentrated, $beta1'Zk(t)$, restricted cointegrating relation number 1

icantly different behavior, in particular if the former looks more stationary than the latter. Before proceeding to the $I(2)$analysis, it is therefore worth taking a closer look at Figures 2 and 3 which depict the graphs of the concentrated and not-concentrated restricted cointegrating relations where the upper panels contain the uncorrected relations, $\beta' X_t$, and the lower panels the cointegration relations corrected for short-run dynamics, $\beta' R_{1t}$. With regard to the first cointegrating relationship the two graphs do not seem to differ significantly. However, the second relationship reveals significant differences where the concentrated cointegrating relation looks considerably more stationary than the non-concentrated one. Coupled with the fact that we were not able to get rid of an extra unit root by lowering the number of cointegrating vectors, this clearly legitimates further investigation along the dimension of a potential second order common trend.

4.1.2 The I(2) case

Table 4 gives the unrestricted outcome of the $I(2)$ analysis when imposing three common trends of which one is of a second order. The first thing to notice is that the common $I(2)$ trend mainly seem to feed into export prices and

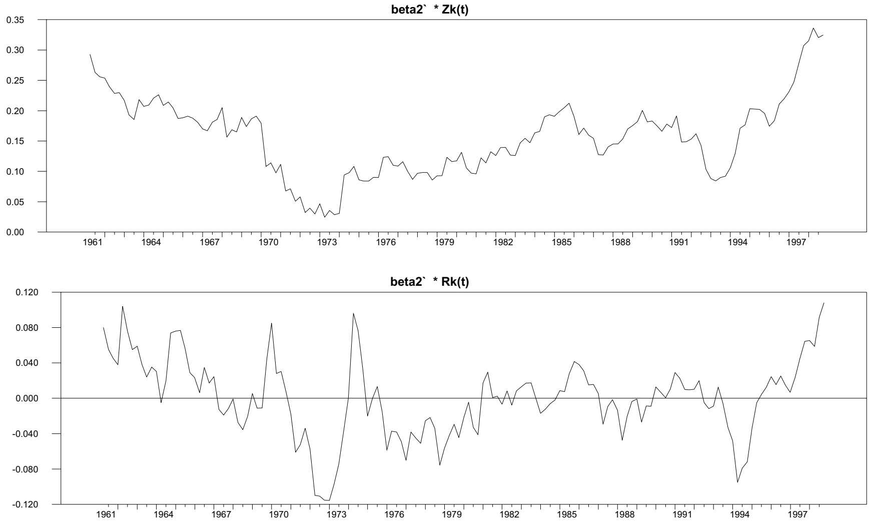

Figure 3: Concentrated, $beta2'Rk(t)$, and not-concentrated, $beta2'Zk(t)$, restricted cointegrating relation number 2

unit labor costs. The $I(2)$ trend itself originates from a linear combination of residuals in export prices, world market prices and unit labor costs and can thus be characterized as a purely nominal trend. In light of theory, the directly cointegrating relationship denoted in the table as a stationary linear combination of levels, is close to the outcome of monopolistic competition, the main difference being that there seems to be some measure of real unit labor costs that enters into the relationship. If so, this means that an increase in unit labor costs will only affect output insofar the increase also implies an increase in "real" terms. The multicointegrating relationship is interpretable as a monopolistic price setting rule. Getting rid of the export volume coefficient by multiplying the first directly cointegrating vector $\widetilde{\beta}_1$ by 0.339 and adding it to the second concerning the level part of the multicointegrating relationship, $\widetilde{\beta}_2$, gives namely the relationship:

$$pa_t = 1.48pw_t + 0.4396ulc_t + 1.156R_t - 0.023Trend$$

The polynomial part seems to be dominated by the coefficients of growth in prices and/or unit labor costs. However, following the suggested procedure in Rahbek et al (1999), these coefficient estimates can be made more interpretable by adding stationary relations to $\widetilde{\beta}_2' X_t + \widetilde{\delta}\beta_{\perp 2}' \Delta X_t$. That is, to combine $\widetilde{\beta}_2' X_t$ with linear combinations, $\upsilon' \Delta X_t$, where υ is a $p \times (p - r - s)$ matrix such that $\upsilon' \beta_{\perp 2}$ has full rank, see Johansen (1992). This may be clearer if we consider the alternative relation, $\widetilde{\beta}_2' X_t + \widetilde{\delta}^* \upsilon' \Delta X_t$, and rewrite it as in Rahbek et al, as

$$\widetilde{\beta}_2' X_t + \widetilde{\delta}^* \upsilon' \overline{\beta}_{\perp 2} \beta_{\perp 2}' \Delta X_t + \widetilde{\delta}^* \upsilon' (\overline{\beta}, \overline{\beta}_1)(\beta, \beta_1)' \Delta X_t$$

The last term is stationary as $(\beta, \beta_1)' X_t$ is $I(1)$, and the first terms define a polymomially cointegrating relation if $\widetilde{\delta}^* \upsilon' \overline{\beta}_{\perp 2} = \widetilde{\delta}$.

Focusing on the role of export prices, I will assume that $\upsilon = (0, 1, 0, 0, 0)'$ is a valid choice. Based on the estimated coefficients, this means that the loading to ν is $\widetilde{\delta}^* = \widetilde{\delta}\beta_{\perp 2}' \beta_{\perp 2} (\upsilon' \beta_{\perp 2})^{-1} = 16.175$ and the polynomially cointegrating relation is therefore given by:

$$pa_t - 1.48 pw_t - 0.4396 ulc_t - 1.156 R_t + 0.023 Trend + 16.175 \Delta pa_t. \qquad (8)$$

(8) is a dynamic export price relation and could readily constitute an error-correction model in an $I(1)$ framework. However, in this context (8) is a polynomially cointegrating relation and functions as an error correction term in the $I(2)$ model.

In the I(1) analysis we were able to identify a system where unit labor costs did not enter into the volume equation at the same time as export prices are pure markup relationships over unit labor costs. To investigate whether this also could be the case when having to deal with a second order trend I present below the results when these restrictions are imposed on the level part of the cointegrating vectors. Looking at Table 5, the first thing to notice is the LR-test for overidentifying restrictions which is not able to reject the null of correctly imposed restrictions as the p-value is equal to 0.24. Otherwise, the results are strikingly similar to the results of the $I(1)$ analysis, the only difference being that the second cointegrating relationship now is a dynamic export price equation given by:

$$pa_t - ulc_t + 10.926 \Delta pa_t + 0.0013 Trend. \qquad (9)$$

	a	pa	pw	ulc	R	$Trend$
	Stationary linearcombinations of levels					
$\widetilde{\beta}_1'$	1	-1.107	-1.618	0.656	-2.413	0.031
	Multicointegrating relations					
$\widetilde{\beta}_2'$	-0.339	1	-0.379	-0.497	0.096	0.004
$\widetilde{\delta}\beta_{\perp 2}'$	1.898	4.250	0.638	6.826	0.264	
	The $I(2)$ trend loadings					
$\beta_{\perp 2}'$	0.4712	1.055	0.1583	1.694	0.065	
	The common I(2) trend coefficients					
$\alpha_{\perp 2}'$	0	0.0506	-0.0285	-0.0266	0	
	The adjustment coefficients					
$\widetilde{\alpha}_1'$	-0.005	-0.002	0.033	-0.082	0.123	
$\widetilde{\alpha}_2'$	-0.145	-0.014	0.065	-0.033	-0.122	

Table 4: Unrestricted estimates in the I(2) model for Germany

	a	pa	pw	ulc	R	$Trend$
	Stationary linearcombinations of levels					
$\widetilde{\beta}_1'$	1	0	-2.81	0	-3.36	0.05
	Multicointegrating relations					
$\widetilde{\beta}_2'$	0	1	0	-1	0	0.0013
$\widetilde{\delta}^*$	0	10.926	0	0	0	
LR-test, rank=2: Chi^2(4) = 5.54 [0.24]						

Table 5: Restricted estimates in the I(2) model for Germany

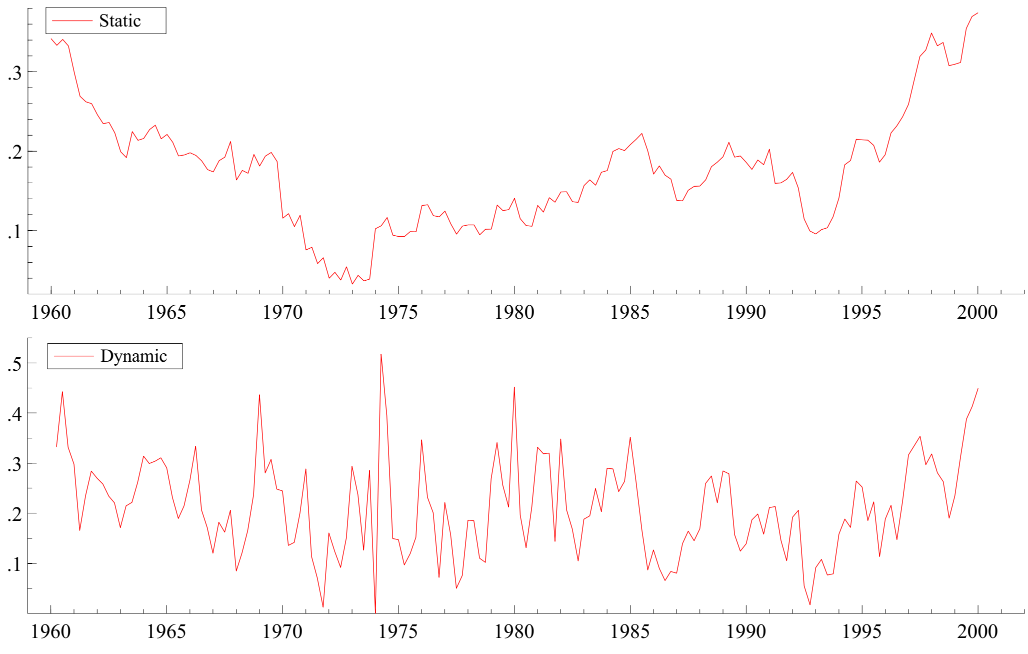

Figure 4: Static cointegrating relationship vs. multicointegrating relationship for the German economy.

Figure 4 compares the second cointegrating relation of the $I(1)$ case with the multicointegrating relationship (9). Even though different scales contribute to give a slightly biased impression, it is readily seen that the new long-run relationship appears much more stationary than the old static one.

4.2 Norway

In Section 3 I deliberately avoided commenting on the time series properties of the Norwegian data as these have been extensively commented on elsewhere. However, to briefly summarize the results, neither system tests nor univariate Dickey Fuller tests are able to question the inherent non-stationarity of the data. The only matter for concern thus seems to be whether this non-stationarity might be of a higher order or not. Table 6 below reproduces the results with regard to determination of cointegration indices of the pooled seven-dimensional analysis in Hammersland (2004), where the data in addition to including exports and export prices of both the service and trading sectors, consist of unit labor costs and indicators of world market prices and world demand.[13] As can be seen from the table, there is slight evidence of three common trends of which one is of a second order. However, the rejection of a common $I(2)$ trend is only marginally insignificant and if one is willing to reject at a level of close to ten per cent, the outcome could easily be accepting as many as six cointegrating vectors among the variables in the information set. If so, this would be totally in line with the outcome of the three-step analysis in Hammersland (2004) designed particularly to deal with identification of cointegrating vectors in the case of times series with a small cross-sectional dimension.

Looking at Table 7 below, there does not seem to be a problem of getting rid of a potential high additional unit root. Already, after imposing the first, the second largest has namely a modulus significantly below 0.9. Also, looking at the graphs of the concentrated and non-concentrated restricted cointegrating relations in the appendix does not reveal that the non-concentrated series exhibit a significantly different behavior from the concentrated ones. In total therefore, there seem to be little evidence of higher order non-stationarity and I have thus chosen to present the outcome of an

[13] To be able to fit the Table in the text, the numbers have been rounded off to their nearest integer representation.

Table 6: The trace test of cointegrating indices for the Norwegian pooled data

p-r	r				$S_{r,s}$				Q_r
7	0	**553**	**466**	**392**	**322**	**288**	**256**	**230**	**213**
		352	311	274	241	212	186	165	147
6	1		**445**	**359**	**287**	**225**	**191**	**166**	**150**
			269	234	203	175	151	131	115
5	2			**340**	**260**	**188**	**152**	**125**	**105**
				198	168	142	120	102	87
4	3				**237**	**165**	**104**	**78**	**66**
					137	113	92	75	63
3	4					**138**	**72**	**50***	**38***
						87	68	53	42
2	5						**58**	**37**	**21***
							47	34	25
1	6							**21**	**6***
								20	13
p-r-s		7	6	5	4	3	2	1	

1) Table 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 non-significant test statistics are marked with stars.

Common trends	Moduli					
0	0.9005	0.9005	0.8464	0.8464	0.8218	0.8218
1	1	0.8618	0.8618	0.8440	0.8440	0.7371
2	1	1	0.8761	0.8761	0.7312	0.7312
3	1	1	1	0.8014	0.7435	0.7435
4	1	1	1	1	0.7392	0.7392

Table 7: Moduli of the eigenvalues of the companion matrix under the imposition of common trends

analysis based on the existence of no less than six cointegrating vectors and a common trend of order one. For an account of the identification scheme the interested reader is again referred to Hammersland (2004).

The first point to notice is that the hypothesis of one price does not appear to be supported by the data. On the contrary, the empirical results of Table 8 indicate that small open economies like the Norwegian, still have a considerable degree of monopolistic power in the export market when setting their prices. Also, both sectors' export volumes seem to be totally driven by demand, which is the case when agents accommodate demand ex post for prices fixed ex ante. Finally, and perhaps even more interesting is the finding of strong long-run links across sectors, both with regard to the determination of prices and the determination of volume. Equation (5) for instance is a cointegrating relationship between the two sectors' export prices, saying that export prices in the service sector grow at approximately a yearly rate of 0.8 per cent faster than in the trading sector. This finding is completely in accordance with the perceived view of a more competitive environment in the trading sector. Likewise, equation (6) which is a cointegrating relationship between the two sectors' export volumes, implies that exports grow at a yearly rate of approximately 3.6 per cent faster in the trading sector than in the service sector. This also coincides well with another perceived view: namely that the trading sector is the main origin for innovative productivity improvements.

Eq.:			Cointegrating relationships:		
1:	a1	=	const.	-0.533(pa1-pw)	+2.366R
				(0.040)	(0.082)
2:	pa1	=	const.	+ 0.657ulc	
				(0.033)	
3:	a2	=	const.	+ R	
4:	pa2	=	const.	+ 0.474R	+0.542ulc
				(0.041)	0.033
5:	pa1	=	const.	+ pa2	- 0.002Trend
					(0.0005)
6:	a1	=	const.	+ a2	+0.008 Trend
					(0.0007)

LR-tests :

All overidentifying restrictions: $\chi^2(5) = 3.98[0.55]$

Table 8: Restricted long-run relationships in the pooled analysis when all parameters have been estimated freely.

5 Summary and Conclusions

In this paper I have tried to unveil the degree of independence in European goods markets. The results are rather divergent. On the one hand there is strong evidence of monopolistic power in the process governing external prices among European exporters. The perceived view that shocks to supply may more or less fully crowd out the foreign sector is therefore seriously called into question as the effects of wage hikes, intermediate price shocks etc. readily can be passed on to prices. Further, as this makes goods arbitrage ineffective, hypotheses like PPP and "the law of one price" are of course far from being supported by data. On the other hand, exports seem to be extremely vulnerable to changes in international demand and world market prices. This might be taken to indicate that agents accommodate demand ex post for prices fixed ex ante and if so helps explain entrepreneurs' cry for arrangements geared towards shielding the sector from the vicissitudes of international trade conditions.

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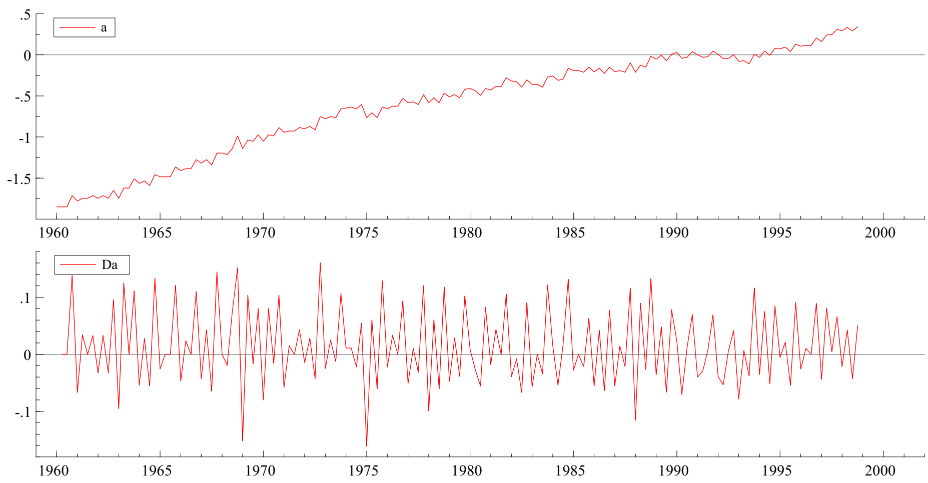
A Data and tests



German exports in levels and differences

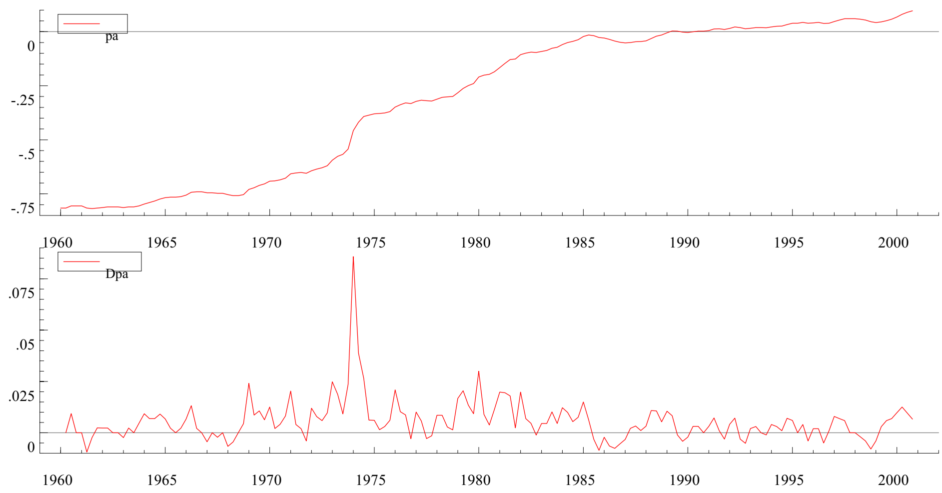


Figure 5: German export prices in levels and differences

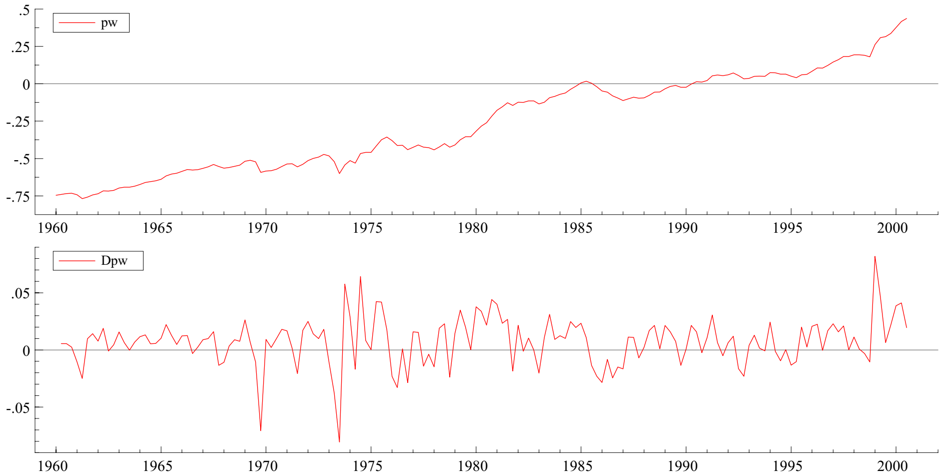


Figure 6: World market prices in levels and differences

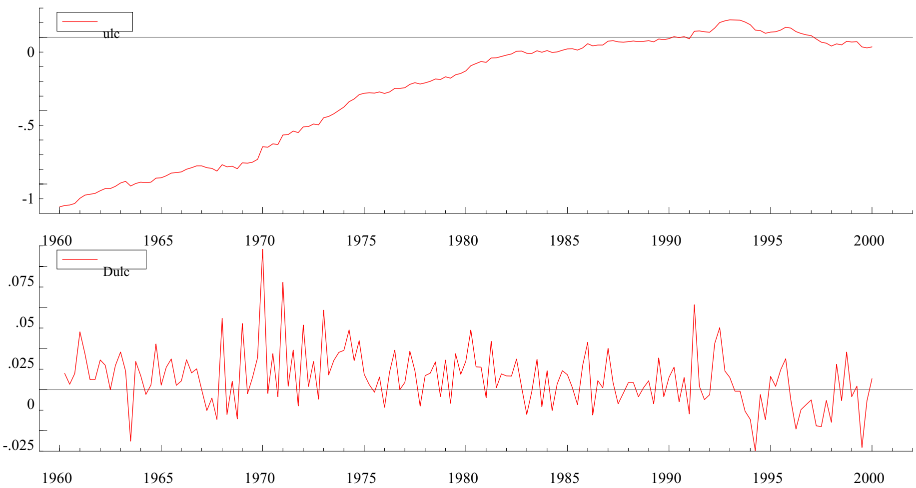


Figure 7: Unit labor costs in levels and differences

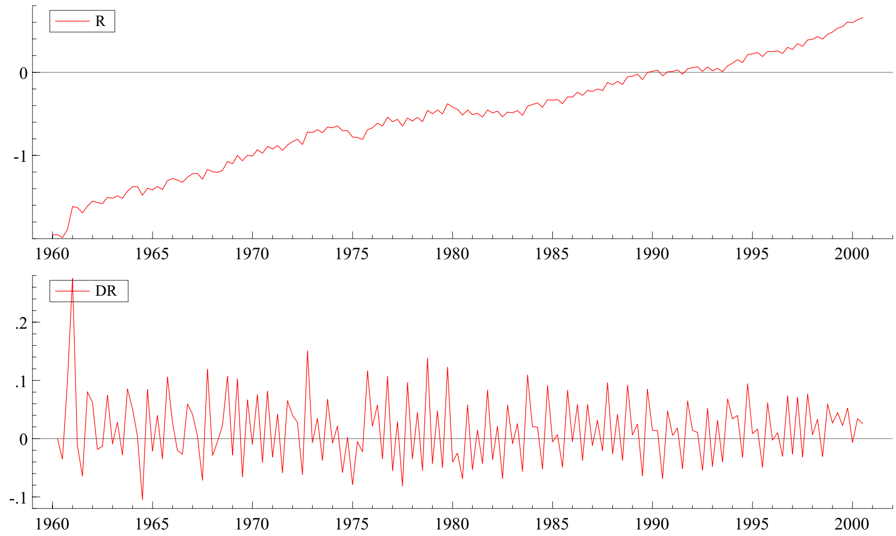


Figure 8: World demand in levels and differences

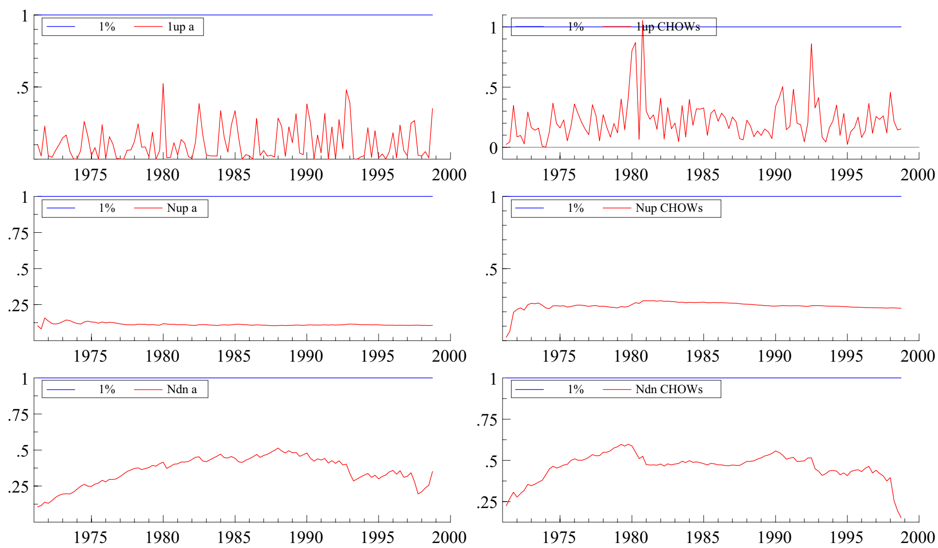


Figure 9: Recursively estimated Chow tests for German exports

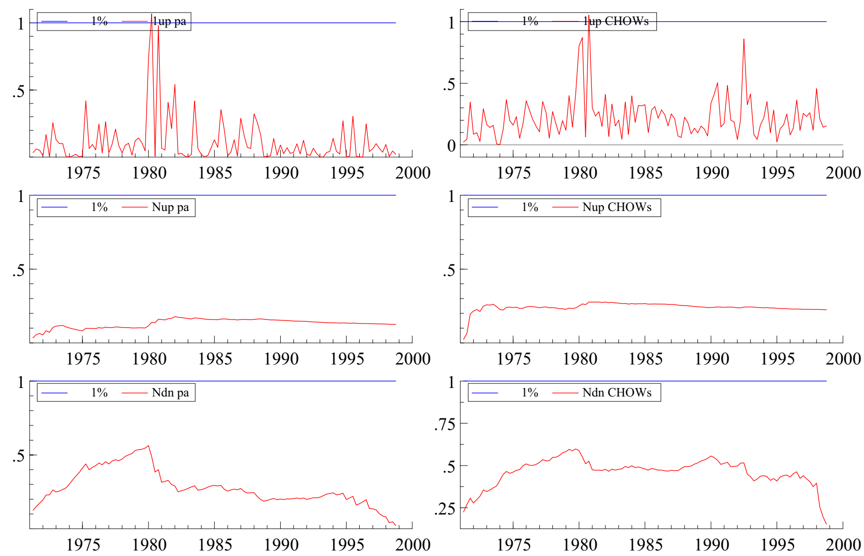


Figure 10: Recursively estimated Chow tests for German export prices

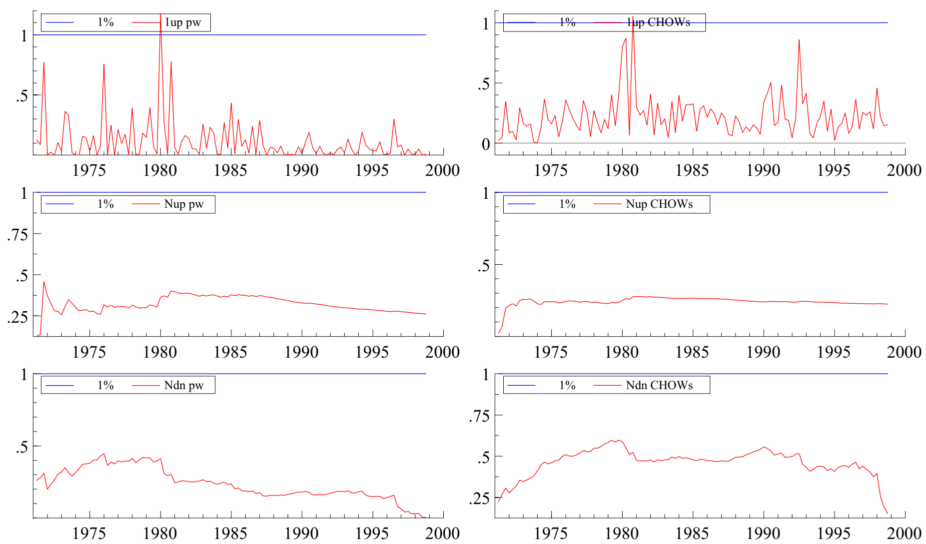


Figure 11: Recursively estimated Chow tests for the World market price equation.

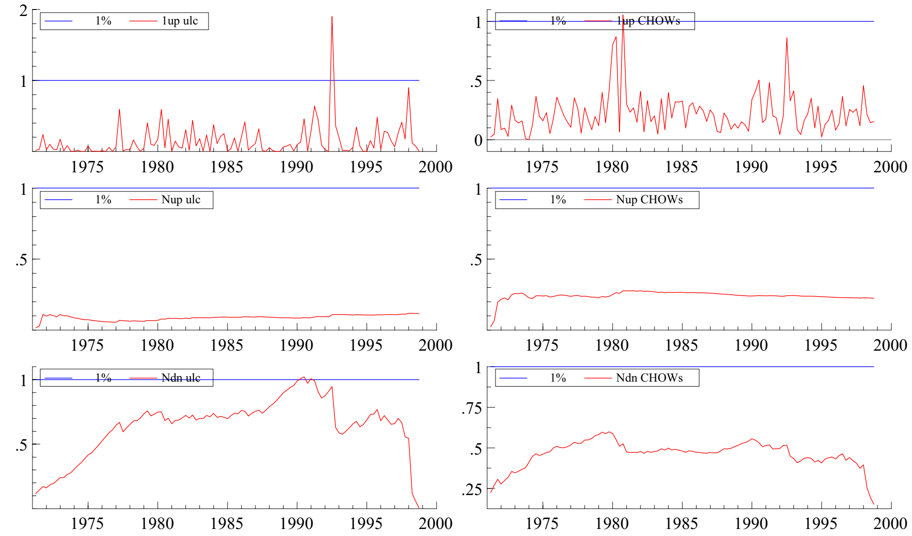


Figure 12: Recursively estimated Chow tests for parameter stability of the German unit labor costs equation

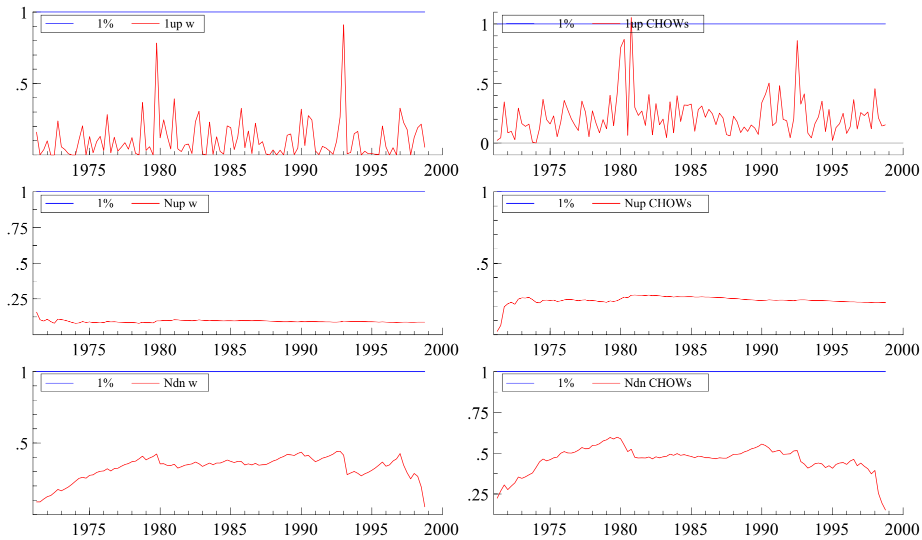


Figure 13: recursively estimated Chow-tests for parameter stability of the world demand equation for Germany

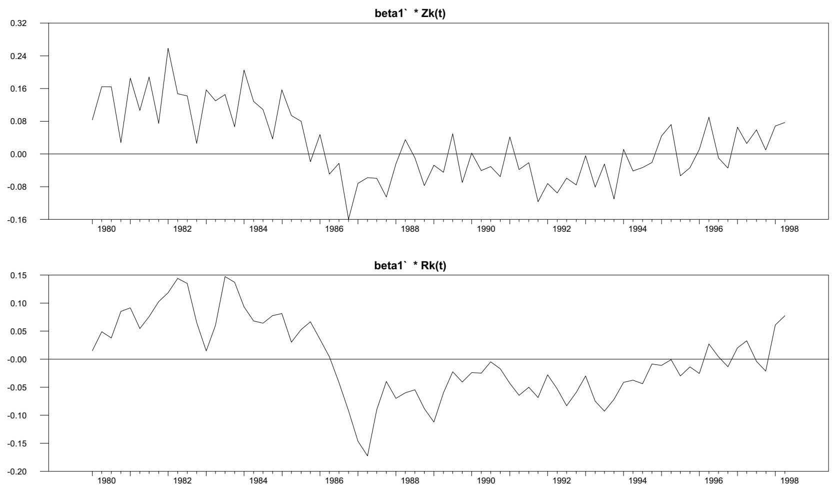


Figure 14: Concentrated, $beta1\prime Rk(t)$, and unconcentrated, $beta1\prime Z\prime k(t)$, restricted cointegrating relation number 1 in the Norwegian study.

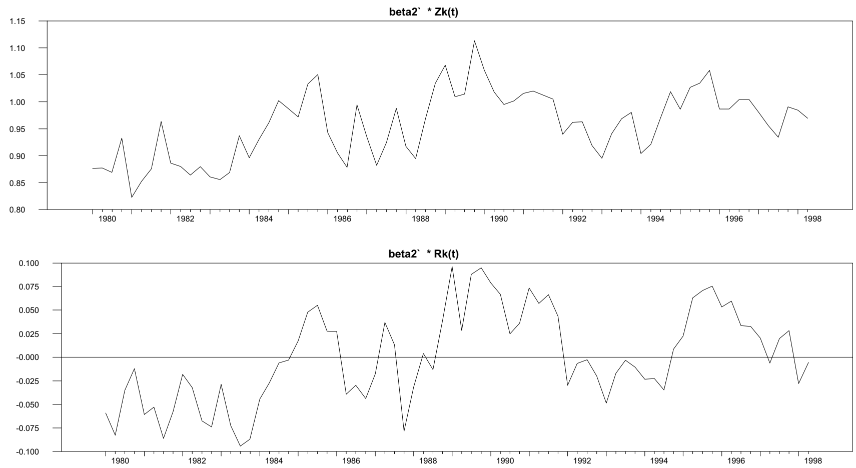


Figure 15: Concentrated, $beta2\prime Rk(t)$, and unconcentrated, $beta2\prime Z\prime k(t)$, restricted cointegrating relation number 2 in the Norwegian study.

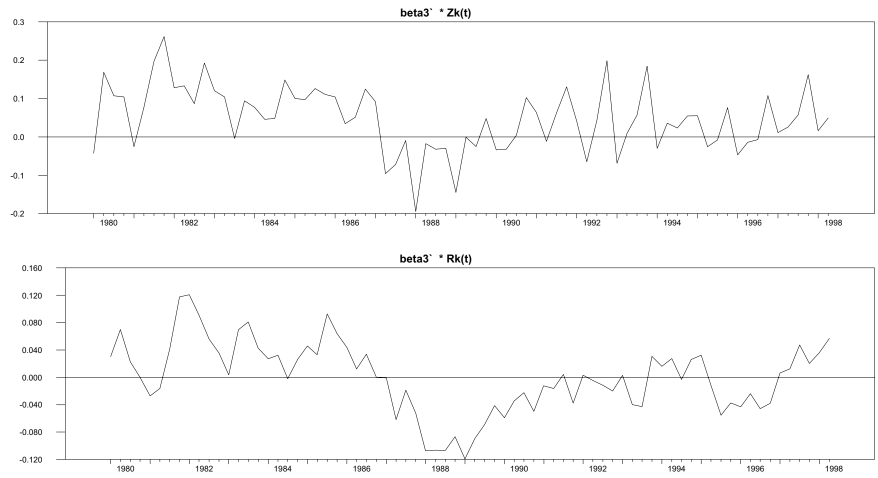


Figure 16: Concentrated, $beta3\prime Rk(t)$, and unconcentrated, $beta3\prime Z\prime k(t)$, restricted cointegrating relation number 3 in the Norwegian study.

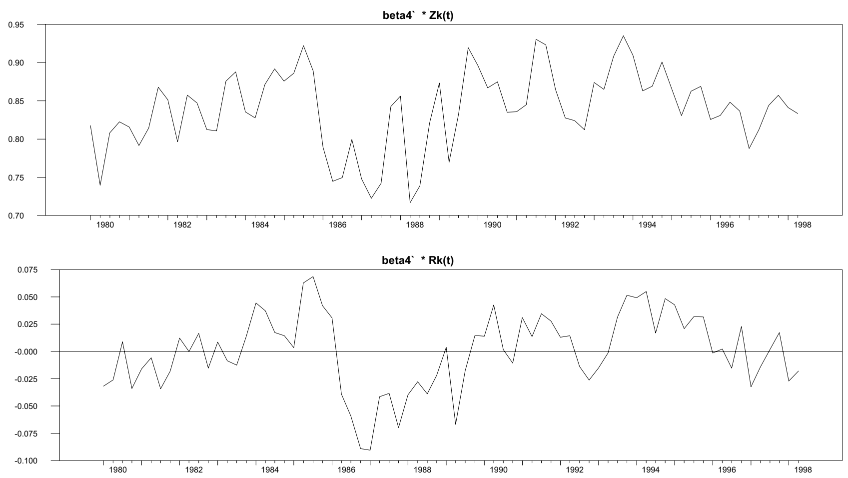


Figure 17: Concentrated, $beta4\prime Rk(t)$, and unconcentrated, $beta4\prime Z\prime k(t)$, restricted cointegrating relation number 4 in the Norwegian study.

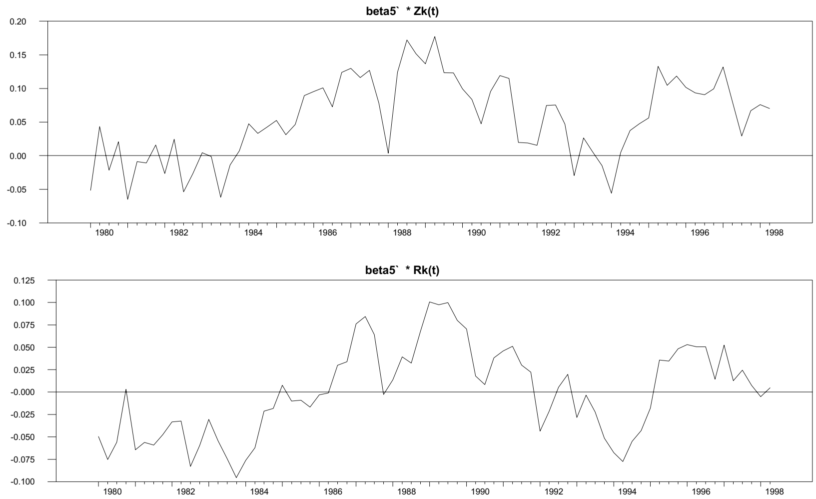


Figure 18: Concentrated, $beta5\prime Rk(t)$, and unconcentrated, $beta5\prime Z\prime k(t)$, restricted cointegrating relation number 5 in the Norwegian study.

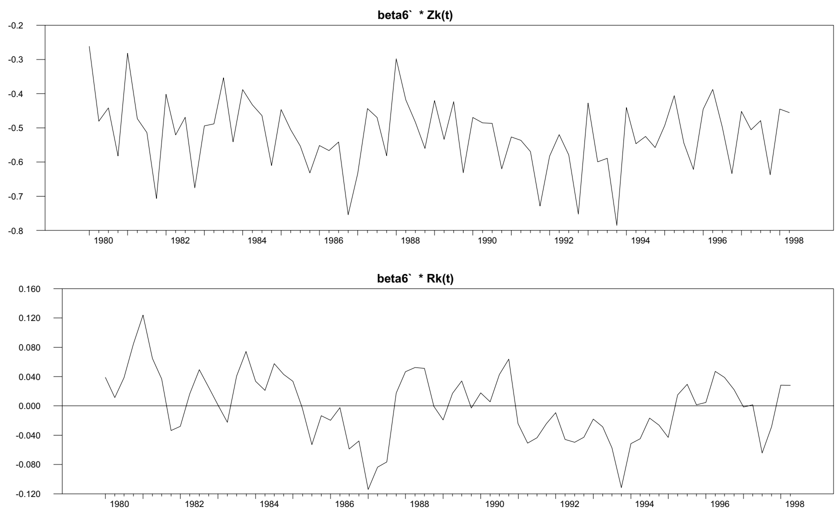


Figure 19: Concentrated, $beta6\prime Rk(t)$, and unconcentrated, $beta6\prime Z\prime k(t)$, restricted cointegrating relation number 6 in the Norwegian study.

Table 9:
Multivariate statistics for testing stationarity of the German data
Two cointegrating vectors and trend in CI space

	Variables				
	a	pa	pw	ulc	R
$\chi^2(3)$	20.045 [0.0002]**	17.23 [0.0006]**	20.668 [0.0001]**	14.345 [0.0025]**	16.751 [0.0008]**

[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 9, 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 in Section 4. The figures in brackets under each Statistics are the tests' significance probabilities and * and ** denote rejection at.5% and 1% critical levels, respectively.

Table 10:
ADF(N) Statistics for Testing for a unit Root in German data.
Estimates of $\left|\left(\hat{\rho}\right) - 1\right|$ in parenthesis

H_0	a	pa	pw	ulc	R
	Variable				
$I(1)$	-1.7928	-0.45905	-2.8923	-0.15304	-2.3491
	(0.04152)	(0.00337)	(0.079)	(0.00191)	(0.0635)
$I(2)$	-5.6386**	-4.2226**	-5.8016**	-3.9442**	-6.0913**
	(0.63109)	(0.39685)	(0.71758)	(0.56312)	(0.8318)

[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 hypotheses 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. 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 one for a and pa, three for pw, four for ulc and five for R when testing I(1) versus I(0), while two for pw, three for a, four for pa and R, and five for ulc when testing I(2)-ness.
[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 -3.44 at a level of 5% and -4.022 at a level of 1 % (Mac Kinnon (1991)) while the corresponding values are -2.881 and -3.475 when testing I(2).

Table 11:
Individual equation and system-diagnostics for the unrestricted VAR-model of German exports[1]

Equation/Tests	AR 1-5 F[5,114]	ARCH 4 F[4,46]	Normality $X_N^2(2)$
Δa	1.5477 [0.1807]	2.6367 [0.0377]*	0.3664 [0.8326]
Δpa	0.4368 [0.8220]	1.6502 [0.1667]	3.3688 [0.1856]
Δpw	0.9679 [0.4405]	0.8903 [0.4723]	4.3124 [0.1158]
ΔR	1.206 [0.3108]	0.2837 [0.8880]	1.8333 [0.3999]
Δulc	0.5425 [0.7437]	0.1414 [0.9664]	3.1946 [0.2024]
System tests:	VAR 1-5[125,447]	VNorm $X^2(10)$	V$X^2F(780, 822)$
	1.0905 [0.2622]	20.854 [0.022]*	0.61706 [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)).

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