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OIL PRICE SHOCKS AND LONG RUN PRICE AND IMPORT DEMAND BEHAVIOR

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Abstract. The effect which the oil price time series has on the long run properties of Vector AutoRegressive (VAR) models for price levels and import demand is investigated. As the oil price variable is assumed to be weakly exogenous for the long run parameters, a cointegration testing procedure allowing for weakly exogenous variables is developed using a LU decomposition of the long run multiplier matrix. The likelihood based cointegration test statistics, Wald, Likelihood Ratio and Lagrange Multiplier, are constructed and their limiting distributions derived. Using these tests, we find that incorporating the oil price in a model for the domestic or import price level of seven industrialized countries decreases the long run memory of the inflation rate. Second, we find that the results for import demand can be classified with respect to the oil importing or exporting status of the specific country. The result for Japan is typical as its import price is not influenced by GNP in the long run, which is the case for all other countries.

Key words and phrases: Cointegration, weak exogeneity, import demand, oil price behaviour.

1. Introduction

The random walk model is nowadays one of the most well known models for the description of economic time series, in particular, price series. The attractiveness of the random walk model is due to a simple, albeit important, implication, that is: all future changes of the described time series are unpredictable and stochastic. So, the current level of the series reflects all available information. However, for several series one may argue that the random walk behavior is due to structural breaks, see Perron (1989). These structural breaks are assumed to be weakly exogenous, in the sense of Engle *et al.* (1983), for the long run parameters of the model and are therefore modelled separately. When structural breaks are not incorporated in the model the estimation results are biased towards the random walk hypothesis.

In import demand models use is made of import and domestic price series, see e.g. Magee (1975), Marquez (1991) and Orcutt (1950). These prices are strongly influenced by the level of the oil price. The OPEC cartel sets oil prices, in most cases, independent from the price levels in the different countries. It seems reasonable therefore to consider oil prices as weakly exogenous for the analysis of the long run behavior of import and domestic price series. In this paper we find that an univariate analysis of domestic and import prices, which neglects the weak exogeneity of the oil price, indicates that these series are (nearly) $I(2)$. This result is substantially changed when a multivariate cointegration analysis of the price series is performed where oil prices are treated as weakly exogenous.

Cointegration refers to the phenomenon that the joint behavior of several economic time series is more stable than one would expect from the analysis of each series individually. In recent years an increasing number of alternative models and estimation methods have been proposed for the analysis of multivariate cointegrated economic time series. The original idea was to approach the problem using single static regression models, e.g. Engle and Granger (1987), but a number of authors have investigated cointegration in a multivariate set-up. For example, a system of static regressions was proposed by Park (1990), while dynamic short run models are favored in the studies of Johansen (1991), Johansen and Juselius (1990) and also Boswijk (1992, 1994).

In this paper we start from an error correction representation of a cointegration model, and extend the likelihood based tests statistics, Wald, Likelihood Ratio (LR), Lagrange Multiplier (LM), of Kleibergen and van Dijk (1994) to test for the existence of long run relationships in the presence of weakly exogenous variables. These test statistics are developed using a LU decomposition of the long run multiplier Π, see Gantmacher (1959). Using this decomposition, the identification of the cointegrating vectors is solved in a purely parametric fashion and consequently the use of canonical correlations or principal components, which are inherent in the multivariate approaches of Stock and Watson (1988), Johansen (1991) and Johansen and Juselius (1990, 1992, 1994), is avoided. In our set-up, both the cointegrating space and the various individual cointegrating relationships are identified in an unique way, which is one of the major differences with some of the multivariate approaches described above.

Apart from the influence of the oil price on the long run behavior of price levels, we investigate its influence on the long run properties of an import demand model of seven industrialized countries, Belgium, France, Germany, Japan, Norway, the United Kingdom (UK) and the United States (US). We find that the results for import demand can be classified with respect to the oil importing or exporting status of the specific country. The result for Japan is typical as its import price is not influenced by GNP in the long run, which is the case for all other countries.

The structure of the paper is as follows. In the second section, we introduce a cointegration model incorporating weakly exogenous variables. The third section contains the derivation of the likelihood based test statistics, Wald, Likelihood

Ratio (LR) and Lagrange Multiplier (LM), and their limiting distributions. In the fourth section, the test statistics are applied to analyze the long run behavior of prices when a weakly exogenous variable, the oil price, is incorporated. In the fifth section, the oil price is incorporated as a weakly exogenous variable in a long run import demand study of imports in seven industrialized countries. Finally, the sixth section concludes.

2. Cointegration in systems with weakly exogenous variables

Cointegration is usually defined in closed form models where all variables depend upon one another, see for example Johansen (1991) and Kleibergen and van Dijk (1994). Many economic models, however, contain parameters which do not directly affect the values of certain variables. These variables are then said to be (weakly) exogenous for these specific parameters, see Engle *et al.* (1983). So, for the estimation of these parameters one can condition on these exogenous variables and treat them as given. In the following, we show the consequences of certain exogeneity relationships for systems with cointegrated variables. Before this, we briefly discuss the general closed form specification of cointegration models.

Consider a framework for the analysis of cointegration. Assume that the k dimensional series $x_t = (x_{1t}\ x_{2t}\ x_{3t})'$, $x_{1t} : 1 \times r$, $x_{2t} : 1 \times (k-r-q)$, $x_{3t} : 1 \times q$, is generated by a p-th order Vector AutoRegressive (VAR) model (see Lütkepohl (1991) for a general discussion of VARs),

$$\Pi(L)x_t = c + dt + \epsilon_t, \quad \text{for} \quad t = 1, \ldots, T, \tag{2.1}$$

$$\Pi(z) = I_k - \sum_{i=1}^{p} \Pi_i z^i, \tag{2.2}$$

where the disturbances ϵ_t, $t = 1, \ldots, T$, are identically and independently generated by a multivariate Gaussian distribution with zero mean and covariance matrix Ω, T is the sample size, and L is the usual lag operator such that $L^n x_t = x_{t-n}$.

It is convenient to respecify the VAR model into an Error Correction Model (ECM), which contains only stationary components and allows the long run multiplier $\Pi = -\Pi(1)$ to be analyzed directly as equation system parameter,

$$\Gamma(L)\Delta x_t = \Pi x_{t-p} + c + dt + \epsilon_t, \tag{2.3}$$

where $\Gamma(L)$ is a $(p-1)$-th order matrix polynomial in L obtained by decomposing, $\Pi(z) = \Gamma(z)(1-z) - \Pi z^p$. It is known, see for example Johansen (1991), that if $(k-m)$ roots of the determinant equation, $|\Pi(z)| = 0$, are equal to 1, then the matrix of long run parameters is of reduced rank. Assuming roots equal to 1 or outside the unit circle, the long run behavior of the series is completely captured by the long run multiplier Π, see Johansen (1991).

In case of cointegration with r cointegrating vectors, we can specify the long run multiplier, which is then of reduced rank, as the product of the two $k \times r$ matrices α, β, $\Pi = \alpha\beta'$. Johansen (1991) uses canonical correlations to measure the relative distance between a model with an unrestricted multiplier and a model with a long run multiplier which accords with cointegration. In this paper we extend the approach from Kleibergen and van Dijk (1994), where a LU decomposition of the long run multiplier is used to measure this relative distance, by allowing for variables which are weakly exogenous for the long run parameters. We specify the long run multiplier as,

$$\Pi = \begin{pmatrix} \Pi_{11} & \Pi_{12} & \Pi_{13} \\ \Pi_{21} & \Pi_{22} & \Pi_{23} \\ \Pi_{31} & \Pi_{32} & \Pi_{33} \end{pmatrix} = \alpha\beta', \quad \text{where} \tag{2.4}$$

$$\alpha = \begin{pmatrix} \alpha_{11} & 0 & 0 \\ \alpha_{21} & \alpha_{22} & \alpha_{23} \\ \alpha_{31} & \alpha_{32} & \alpha_{33} \end{pmatrix} \quad \text{and} \quad \beta = \begin{pmatrix} I_r & 0 & 0 \\ -\beta_{22} & I_{k-r-q} & 0 \\ -\beta_{23} & 0 & I_q \end{pmatrix} \tag{2.5}$$

and assume that the variables contained in x_{3t} are weakly exogenous for the long run parameters of the equations of $(\Delta x_{1t}' \ \Delta x_{2t}')'$, for a definition of weak exogeneity, we refer to Engle *et al.* (1983). The assumption of weak exogeneity changes the specification of the long run multiplier Π, i.e., when the variables x_{3t} are not error correcting, such that α_{3i}, $i = 1, \ldots, 3$, are all equal to 0, the equation of Δx_3 has no influence on the long run parameters of the equations of $(\Delta x_{1t}' \ \Delta x_{2t}')'$, which are β_{22}, β_{23}, α_{22} and α_{23}. In this case x_{3t} is said to be weakly exogenous for the long run parameters, see Johansen (1992) and Urbain (1992*b*). Standard χ^2tests can be performed to test for this but in many cases, it is known a priori that x_3 is (at least) weakly exogenous for the long run parameters. Examples of this are given in Sections 4 and 5. When it is a priori known that certain variables are weakly exogenous, neglecting this information and using a closed form model will lead to a loss of power. It is therefore necessary to develop cointegration testing procedures which allow for weakly exogenous variables. Some of these methods are already documented in the literature, see for example Johansen (1992), but in the sequel we propose some new testing procedures which are based on a so-called LU decomposition of the long run multiplier Π, see also Kleibergen and van Dijk (1994).

When x_3 is weakly exogenous for the long run parameters, it is not necessary to estimate the complete ECM from equation (2.2) but only the first $k-q$ equations need to be estimated. The ECM then reads

$$\Psi(L)\left(\Delta\begin{pmatrix} x_{1t} \\ x_{2t} \\ x_{3t} \end{pmatrix}\right) = c_1 + d_1 t + \gamma\Delta x_{3t} + \Pi_1 x_{t-p} + \epsilon_{1t}, \tag{2.6}$$

where

$$\Psi(z) = I_{k-q} - \sum_{i=1}^{p} \Psi_i z^i, \qquad \Gamma(z) = \begin{pmatrix} \Psi(z) \\ \Gamma_3(z) \end{pmatrix}, \tag{2.7}$$

$$\Pi_1 = \begin{pmatrix} \Pi_{11} & \Pi_{12} & \Pi_{13} \\ \Pi_{21} & \Pi_{22} & \Pi_{23} \end{pmatrix}, \quad \text{and} \quad \Pi = \begin{pmatrix} \Pi_1 \\ \Pi_3 \end{pmatrix}. \tag{2.8}$$

A LU decomposition of the long run multiplier, Π_1, then reads

$$\Pi_1 = \alpha\beta', \tag{2.9}$$

$$\alpha = \begin{pmatrix} \alpha_{11} & 0 & 0 \\ \alpha_{21} & \alpha_{22} & \alpha_{23} \end{pmatrix}, \qquad \beta = \begin{pmatrix} I_r & 0 \\ -\beta_{22} & I_{k-r-q} \\ -\beta_{23} & 0 \end{pmatrix}. \tag{2.10}$$

Using this specification of the long run multiplier, cointegration amounts to $\alpha_{22} = 0$ and $\alpha_{23} = 0$. By varying r, the number of cointegrating vectors/unit roots can be tested for. Note that through the q weakly exogenous variables contained in x_3, we have implicitly incorporated q unit roots when we assume that the weakly exogenous variables are generated by an unit root process.

Given that we have a well defined hypothesis of cointegration, $H_0 : \alpha_{22} = 0, \alpha_{23} = 0$, we can construct the likelihood based test statistics, Wald, Likelihood Ratio (LR) and Lagrange Multiplier (LM), to test for cointegration using the parameters from equations (2.6)–(2.10). In the following section, we will discuss each of these three different statistics and their limiting distributions.

3. Likelihood based cointegration tests in systems with weakly exogenous variables

3.1 *Wald cointegration statistic*

Wald testing is performed in the unrestricted model, the model under the alternative hypothesis, and within this model, the relative distance with the restricted model is determined. This is also the case with the Wald cointegration statistic, which results from a two step estimation procedure of the unrestricted model. Before showing the two step procedure, it is convenient to slightly respecify the ECM,

$$\Delta y_t = c_1 + d_1 t + \Psi z_t + \Pi_1 x_{t-p} + \epsilon_{1t}, \tag{3.1}$$

where $y_t = (x_{1t}\ x_{2t})'$, $z_t = (x_t' \cdots x_{t-p+1}')'$, $\Psi = (\Psi_1' \cdots \Psi_{p-1}')' = \binom{\Phi_1}{\Phi_2}$, $\epsilon_{1t} = \binom{\epsilon_{11t}}{\epsilon_{12t}}$, $c_1 = \binom{c_{11}}{c_{12}}$, $d_1 = \binom{d_{11}}{d_{12}}$, $\Pi_1 = \begin{pmatrix} \Pi_{11} & \Pi_{12} & \Pi_{13} \\ \Pi_{21} & \Pi_{22} & \Pi_{23} \end{pmatrix}$. The two step procedure then becomes

1. Estimate

$$\begin{aligned} \Delta x_{1t} &= c_{11} + d_{11}t + \Phi_1 z_t + \Pi_{11} x_{t-p} + \epsilon_{11t} \\ &= \Phi_1 z_t + \Pi_{11}(x_{1t-p} + \Pi_{11}^{-1}\Pi_{12}x_{2t-p} \\ &\qquad + \Pi_{11}^{-1}\Pi_{13}x_{3t-p} + \Pi_{11}^{-1}c_{11} + \Pi_{11}^{-1}d_{11}t) + \epsilon_{11t} \\ &= \Phi_1 z_t + \alpha_{11}(x_{1t-p} - \beta_{22}' x_{2t-p} - \beta_{23}' x_{3t-p} - \mu_1 - \delta_1 t) + \epsilon_t. \end{aligned} \tag{3.2}$$

In this first step, estimates of the cointegrating vector, $(I_r \ \ -\hat{\beta}_{22}' \ \ -\hat{\beta}_{23}' \ \ -\hat{\mu}_1 \ \ -\hat{\delta}_1)' = (I_r \ \ -\hat{\Pi}_{12}'\hat{\Pi}_{11}' \ \ -\hat{\Pi}_{13}'\hat{\Pi}_{11}' \ \ -\hat{c}_{11}'\hat{\Pi}_{11}^{-1\prime} \ \ -\hat{d}_{11}'\hat{\Pi}_{11}^{-1\prime})'$, and the loading factor, $\hat{\alpha}_{11} = \hat{\Pi}_{11}$, are constructed. These estimates are used in the second step, to measure the relative distance with a cointegration model.

2. Estimate

$$\begin{aligned}
(3.3)\qquad \Delta x_{2t} &= c_{21} + d_{21}t + \Psi_2 z_t + (\Pi_{21}\ \Pi_{22}\ \Pi_{23})x_{t-p} + \epsilon_{12t} \\
&= \Phi_2 z_t + \alpha_{21}(x_{1t-p} - \hat{\beta}_{22}' x_{2t-p} - \hat{\beta}_{23}' x_{3t-p} - \hat{\mu}_1 - \hat{\delta}_1 t) \\
&\quad + (\Pi_{22} + \alpha_{21}\hat{\beta}_{22}')x_{2t-p} \\
&\quad + (\Pi_{22} + \alpha_{21}\hat{\beta}_{22}')x_{2t-p} + (\Pi_{23} + \alpha_{21}\hat{\beta}_{23}')x_{3t-p} \\
&\quad + (c_{21} + \alpha_{21}\hat{\mu}_1)x_{2t-p} + (d_{21} + \alpha_{21}\hat{\delta}_1)t + \epsilon_{12t} \\
&= \Phi_2 z_t + \alpha_{21}(x_{1t-p} - \beta_{22}' x_{2t-p} - \beta_{23}' x_{3t-p} - \mu_1 - \delta_1 t) \\
&\quad + \alpha_{22}x_{2t-p} + \alpha_{23}x_{3t-p} + \mu_2 + \delta_2 t + \epsilon_{12t}.
\end{aligned}$$

In this second step, the parameters α_{22} and α_{23} (and if one specifies cointegration including a certain specification of the deterministic components, μ_2 and δ_2) measure the distance with a cointegration model. The significance of these parameters can now be evaluated using a Wald statistic. By varying r, we can then test for the number of cointegrating relationships. We also notice that since we do not impose any unit roots, the limiting distribution of $\hat{\beta}_{22}$ will be biased, see Phillips (1991), contrary to the limiting distribution of $\hat{\beta}_{23}$, which is unbiased because of the weak exogeneity of x_3 for the cointegrating vectors. As a result of that, χ^2 tests cannot be performed on β_{22} but can on β_{23}. It is important to note that the construction of the cointegrating vectors does not restrict the analysis. Postmultiplying $(I_r\ \ -\beta_{22}'\ \ -\beta_{23}')'$ by $(\alpha_{11}\ \ \alpha_{12})$, any value in the space spanned by the cointegrating vectors can be obtained. This is due to the full rank of $(\beta_{22}'\ \beta_{23}')'$. However, the loading factor α_{11} is crucial because if α_{11} has a lower rank value, a different kind of limiting distribution of the Wald statistic results, see Kleibergen and van Dijk (1994).

In the two step procedure, an estimated value of $(\beta_{22}'\ \beta_{23}')'$, $(\hat{\beta}_{22}'\ \hat{\beta}_{23}')'$, is used in the second step. Because this estimator is superconsistent, when α_{11} has full rank, it does not affect the limiting distribution of the Wald-statistic but the functional form of the Wald-statistic differs from the standard F type tests to account for the uncertainty of $\hat{\beta}_2$. The functional form of the Wald statistic to test the hypothesis of cointegration, $H_0 : \alpha_{22} = 0, \alpha_{23} = 0$, against the alternative of stationarity, $H_1 : \alpha_{22} \neq 0, \alpha_{23} \neq 0$, reads

$$(3.4)\quad t_w = \operatorname{tr}\left[\left(\begin{pmatrix} -\hat{\alpha}_{21}\hat{\alpha}_{11}^{-1} \\ I_{k-r-q} \end{pmatrix}' \hat{\Omega} \begin{pmatrix} -\hat{\alpha}_{21}\hat{\alpha}_{11}^{-1} \\ I_{k-r-q} \end{pmatrix}\right)^{-1} \left(\sum_{t=1}^{T} \tilde{\epsilon}_{21t}\tilde{\epsilon}_{21t}' - \sum_{t=1}^{T} \hat{\epsilon}_{21t}\hat{\epsilon}_{21t}'\right)\right]$$

where $\tilde{\epsilon}_{21t}$ denotes the residuals under H_0 and $\hat{\epsilon}_{21t}$ under H_1, $\hat{\Omega} = \frac{1}{T}\sum_{t=1}^{T} \hat{\epsilon}_{1t}\hat{\epsilon}_{1t}'$. In the closed form VAR, the limiting distribution of this Wald-statistic for cointegration testing depends on the rank of α_{11}, see Kleibergen and van Dijk (1994). This also holds for the Wald-statistic in models with weakly exogenous variables since we cannot construct a superconsistent estimate of the cointegrating vector in that case. It is therefore important that the variables are ordered in such a way that α_{11}has a full rank value. The knowledge to order in a certain fashion can be derived from economic theory but the rank of α_{11} can also be tested using standard

χ^2 tests. It is also possible to include an additional estimation step in the first step of the two step procedure and through this additional step the dependence of the limiting distribution on the rank of α_{11} disappears, see Kleibergen (1994). The resulting cointegration test-statistic is, however, no longer a Wald-statistic in that case.

In a following section, the limiting distribution of the Wald-statistic is stated assuming that α_{11} has a full rank value.

3.2 *Likelihood ratio and Lagrange multiplier statistics*

In a similar way to the case of the Wald statistic, LM and LR tests for the hypothesis of cointegration can be derived, $H_0 : \alpha_{22} = 0, \alpha_{23} = 0$. These tests use the ML estimator of β_{22} and β_{23}, which can be obtained by means of canonical correlations (an eigenvalue problem) or numerical optimization, see also Section 2. The resulting estimators all have an unbiased mixed normal limiting distribution, see Phillips (1991), such that asymptotic χ^2 tests can be performed on these estimators.

The LR test reads

$$t_{LR} = T(\ln(|\tilde{\Omega}|/|\hat{\Omega}|)) \tag{3.5}$$

where $\tilde{\Omega}$, $\hat{\Omega}$ represent the estimated covariance matrix of the residuals, under the restriction of cointegration and under the alternative hypothesis of stationarity. Note that also in case of incorporating weakly exogenous variables, the likelihood ratio test exactly corresponds with Johansen's cointegration test, see also Johansen (1992).

As is usually the case with LM statistics, the LM statistic for cointegration testing results from an auxilliary regression.

$$\tilde{\epsilon}_{21t} = \alpha_{22}x_{2t-p} + \alpha_{23}x_{3t-p} + \mu_2 + \delta_2 t + u_{21t} \tag{3.6}$$

$$t_{LM} = T[\mathrm{tr}(\tilde{\Omega}_{22}^{-1}(\tilde{\Omega}_{22} - \hat{\Omega}_{22}))] \tag{3.7}$$

where $\tilde{\Omega}_{22} = \frac{1}{T}\sum_{t=1}^{T}\tilde{\epsilon}_{21t}\tilde{\epsilon}_{21t}'$, $\hat{\Omega}_{22} = \frac{1}{T}\sum_{t=1}^{T}u_{21t}u_{21t}'$ and $\tilde{\epsilon}_{21t}$ are the residuals estimated under the hypothesis of cointegration, $H_0 : \alpha_{22} = 0, \alpha_{23} = 0$. Note that the cointegrating vector estimator has to be estimated under the restriction of cointegration in this setting. The limiting distributions of the LR and LM statistics are identical and independent from α_{11}.

3.3 *Limiting distributions cointegration statistics*

In the former section, tests for cointegration in ECMs with weakly exogenous variables are defined. To be able to use these statistics, we need to know their limiting distributions. Assuming that α_{11} has full rank (which is only necessary for the Wald-statistic), the limiting distributions of the three different cointegration statistics are identical. Another assumption needs to be made concerning the weakly exogenous variables. Since no model is estimated for these variables, a Data Generating Process (DGP) has to be assumed for these variables. For the derivation of the limiting distributions, we therefore start with explicitly stating the DGPs of the different series.

$$(3.8)\quad \left(\Delta\begin{pmatrix} x_{1t} \\ x_{2t} \end{pmatrix}\right) = \begin{pmatrix} 0 \\ I_{k-r} \end{pmatrix}\mu_2 + \begin{pmatrix} 0 \\ I_{k-r} \end{pmatrix}\delta_2 t + \gamma\Delta x_{3t} + \sum_{i=1}^{p-1}\Psi_i\left(\Delta\begin{pmatrix} x_{1t-i} \\ x_{2t-i} \\ x_{3t-i} \end{pmatrix}\right)$$
$$+ \begin{pmatrix} \alpha_{11} \\ \alpha_{21} \end{pmatrix}(I_r \quad -\beta_{22}' \quad -\beta_{23}' \quad -\mu_1 \quad -\delta_1)\begin{pmatrix} x_{t-p} \\ 1 \\ t \end{pmatrix} + \epsilon_{1t}$$

$$(3.9)\quad \Delta x_{3t} = \mu_3 + \delta_3 t + \sum_{i=1}^{p-1}\Gamma_{3i}\left(\Delta\begin{pmatrix} x_{1t-i} \\ x_{2t-i} \\ x_{3t-i} \end{pmatrix}\right) + \epsilon_{3t}.$$

As shown in equation (3.9), the level variables do not appear in the equation for Δx_3, which explains the weak exogeneity of x_3 for the cointegrating vectors. The limiting distributions of the cointegration statistics will be discussed for different specifications of the deterministic components in equations (3.8) and (3.9).

The limiting distributions of the test-statistics are stated in the following theorem. Note that in case of the Wald statistic, the limiting distribution only corresponds with the limiting distributions stated in Theorem 1, when α_{11} has a full rank value.

THEOREM 1. *Assuming that the number of variables is k (including the weakly exogenous variables), the number of weakly exogenous variables is q and the number of cointegrating vectors is r, such that the number of unit roots is $k-r$, which exceeds or is equal to q ($k-r \geq q$), the following results*

(i) *When the DGP in equations* (3.8) *and* (3.9) *is such that $\delta_1 = \delta_2 = \delta_3 = 0$ and in the estimated model, $\delta_1 = \delta_2 = 0$, the cointegration statistics to test the hypothesis, $H_0 : \alpha_{22} = 0, \alpha_{23} = 0$, weakly converge to*

$$(3.10)\qquad \operatorname{tr}\left[\left(\int(\bar{S}_{k-r-1}\bar{\tau})'dS_{k-r-q}\right)'\left(\int(\bar{S}_{k-r-1}\bar{\tau})'(\bar{S}_{k-r-1}\bar{\tau})\right)^{-1}\right.$$
$$\left.\cdot\left(\int(\bar{S}_{k-r-1}\bar{\tau})'dS_{k-r-q}\right)\right].$$

(ii) *When the DGP in equations* (3.8) *and* (3.9) *is such that $\mu_2 = \mu_3 = \delta_1 = \delta_2 = \delta_3 = 0$ and in the estimated model, $\delta_1 = \delta_2 = 0$, the cointegration statistics to test the hypothesis, $H_0 : \alpha_{22} = 0, \alpha_{23} = 0, \mu_2 = 0$, weakly converge to*

$$(3.11)\qquad \operatorname{tr}\left[\left(\int(S_{k-r}\iota)'dS_{k-r-q}\right)'\left(\int(S_{k-r}\iota)'(S_{k-r}\iota)\right)^{-1}\right.$$
$$\left.\cdot\left(\int(S_{k-r}\iota)'dS_{k-r-q}\right)\right].$$

(iii) *When the DGP in equations* (3.8) *and* (3.9) *is such that* $\delta_2 = \delta_3 = 0$, *the cointegration statistics to test the hypothesis*, $H_0 : \alpha_{22} = 0, \alpha_{23} = 0, \delta_2 = 0$, *weakly converge to*

$$\operatorname{tr}\left[\left(\int (S_{k-r}\tau)' dS_{k-r-q}\right)' \left(\int (S_{k-r}\tau)'(S_{k-r}\tau)\right)^{-1} \cdot \left(\int (S_{k-r}\tau)' dS_{k-r-q}\right)\right]. \tag{3.12}$$

(iv) *When the DGP in equations* (3.8) *and* (3.9) *is such that* $\mu_1 = \mu_2 = \mu_3 = \delta_1 = \delta_2 = \delta_3 = 0$ *and in the estimated model*, $\delta_1 = \delta_2 = 0$, *the cointegration statistics to test the hypothesis*, $H_0 : \alpha_{22} = 0, \alpha_{23} = 0$, *weakly converge to*

$$\operatorname{tr}\left[\left(\int \bar{S}'_{k-r} dS_{k-r-q}\right)' \left(\int \bar{S}'_{k-r}\bar{S}_{k-r}\right)^{-1} \left(\int \bar{S}'_{k-r} dS_{k-r-q}\right)\right]. \tag{3.13}$$

(v) *When the DGP in equations* (3.8) *and* (3.9) *is such that* $\delta_1 = \delta_2 = \delta_3 = 0$, *the cointegration statistics to test the hypothesis*, $H_0 : \alpha_{22} = 0, \alpha_{23} = 0$, *weakly converge to*

$$\operatorname{tr}\left[\left(\int \tilde{S}'_{k-r} dS_{k-r-q}\right)' \left(\int \tilde{S}'_{k-r}\tilde{S}_{k-r}\right)^{-1} \left(\int \tilde{S}'_{k-r} dS_{k-r-q}\right)\right] \tag{3.14}$$

where all integral signs are defined on the unit interval $(0, 1)$, S_i *represents an i-th dimensional Brownian motion defined on the unit interval*, d *is a derivative operator* (*the Brownian motion with respect to which the derivatives are taken corresponds with the other Brownian motion in the integral*); $\tau(t) = t$, $\iota(t) = 1$, $\bar{\tau}(t) = \tau(t) - \int \tau(t)dt$, $\bar{S}_{k-r-j} = S_{k-r-j} - \int S_{k-r-j}(t)dt$, $\tilde{S}_{k-r-j} = \bar{S}_{k-r-j} - \int \bar{\tau}(t)dt$, $0 \leq t \leq 1$.

PROOF. The results follow from the stochastic trend specification of the model in equations (3.8) and (3.9) using the theory derived in *a.o.* Phillips and Durlauf (1986), Johansen (1991) and Kleibergen and van Dijk (1994). For a proof see also Johansen (1992).

The limiting distributions stated in the above theorem differ from the limiting distributions of cointegration statistics in closed form ECMs since we have implicitly incorporated q unit roots through the q weakly exogenous variables. Given the limiting distributions stated in Theorem 1, it is also straightforward to construct limiting distributions for other specifications of the deterministic components.

4. Prices: additional unit roots through $I(2)$-ness or exogenous variables

For a lot of economic time series, it can be argued whether the "random walk behavior" of the analyzed series is the result of unit roots or structural breaks, see Perron (1989). These structural breaks are assumed to be weakly exogenous for the long run parameters of the analyzed model and are therefore modelled separately. When the structural breaks are not incorporated in the model, the estimated parameters are biased towards unit roots. In import demand studies, price levels are used and it is often argued that price levels are $I(2)$, implying random walk type of behavior for inflation rates. Inflation rates indeed possess certain long memory properties, which can be modelled by a random walk model. In the following, we argue that the $I(2)$-ness of prices or the long memory property of inflation rates is the result of an omitted variable. Price levels in different countries are strongly influenced by the value of the oil price. The OPEC oil cartel sets oil prices independently from the different price levels of the countries in the world. Oil prices can therefore be considered as weakly exogenous for the long run parameters of a model for the price level. When this property is neglected and prices are modelled univariately, we find long memory properties of inflation rates and for certain price levels we cannot reject the hypothesis that the price level is generated by an $I(2)$ process. These results crucially depend on omitting the oil price from the model, since the long memory property of inflation rates is substantially reduced when oil prices are incorporated in the model as a weakly exogenous variable.

For seven different countries, the United States (US), United Kingdom (UK), Germany, France, Belgium, Norway and Japan, the domestic and import price levels are analyzed. The data of these series are obtained from the IFS series of datastream and are quarterly data which range broadly from 1960 to 1991.

To investigate the stationarity issues, the following four models are estimated.

$$\Delta x_t = c + dt + \sum_{i=1}^{p-1} \gamma_i x_{t-i} + \pi_1 x_{t-1} + \sum_{i=1}^{3} \varphi_i SD_{it} + \epsilon_t \tag{4.1}$$

$$\Delta^2 x_t = c + \sum_{i=1}^{p-2} \gamma_i \Delta^2 x_{t-i} + \pi_2 \Delta x_{t-1} + \sum_{i=1}^{3} \varphi_i SD_{it} + \epsilon_t \tag{4.2}$$

$$\begin{aligned} \Delta x_t &= c + dt + \alpha \Delta y_t + \sum_{i=1}^{p-1} \gamma_i \Delta x_{t-i} + \sum_{i=1}^{p-1} \alpha_i \Delta y_{t-i} + \pi_3 \begin{pmatrix} x_{t-1} \\ y_{t-1} \end{pmatrix} \\ &\quad + \sum_{i=1}^{3} \varphi_i SD_{it} + \epsilon_t \end{aligned} \tag{4.3}$$

$$\begin{aligned} \Delta^2 x_t &= c + \alpha \Delta^2 y_t + \sum_{i=1}^{p-2} \gamma_i \Delta^2 x_{t-i} + \sum_{i=1}^{p-2} \alpha_i \Delta^2 y_{t-i} + \pi_4 \begin{pmatrix} \Delta x_{t-1} \\ \Delta y_{t-1} \end{pmatrix} \\ &\quad + \sum_{i=1}^{3} \varphi_i SD_{it} + \epsilon_t \end{aligned} \tag{4.4}$$

Table 1. Wald tests for long run memory properties of price indices.

		π_1	π_2	π_3	π_4
US	*pm*	4.45	11.4	5.86	16.1
	pd	4.32	10.2	7.31	21.4
	oil	2.62	34.3		
UK	*pm*	12.6	23.3	11.6	27.8
	pd	4.49	6.25	11.5	25.3
	oil	3.03	26.2		
Germany	*pm*	3.72	24.5	9.47	39.3
	pd	3.50	18.2	4.43	27.4
	oil	3.02	26.2		
France	*pm*	1.18	21.9	31.2	61.2
	pd	3.31	18.0	8.80	31.1
	oil	1.72	32.6		
Belgium	*pm*	2.85	16.2	12.3	63.8
	pd	0.86	10.7	6.45	39.4
	oil	1.90	24.5		
Japan	*pm*	3.49	20.1	25.1	48.1
	pd	2.62	14.7	3.99	26.7
	oil	1.42	29.4		
Norway	*pm*	4.58	15.0	9.15	36.5
	pd	6.10	10.3	11.3	22.2
	oil	1.52	26.4		
C.V.(5%)		11.9	8.17	14.0	10.8
C.V.(10%)				11.8	9.27

where x_t is the log of the price level in period t, y_t is the log of the oil price in period t denoted in the currency of the specific country, $p = 4$, and SD_i are centered seasonal dummies. To show the consequences of omitting the oil price from a model for price levels on the long run parameters of the model, we test for the significance of π_i, $i = 1, \ldots, 4$, in the models in the equations (4.1)–(4.2) using Wald tests (= squared Dickey-Fullers in case of π_1 and π_2). The models in equations (4.1) and (4.2) are univariate and only contain the price level while the models in equation (4.3)–(4.4) incorporate the oil price as an exogenous variable. The resulting statistics are stated in Table 1. The critical values to be used for π_3 and π_4 result from Theorem 1 ($q = 1$) (equation (3.14) refers to π_3 and equation (3.13) refers to π_4) and are constructed using Monte-Carlo simulation.

Table 1, where *pd* refers to the (log) (whole sale) domestic price of a specific country, *pm* to the (log) import price and *oil* to the (log) of the oil price in the currency of the specific country, shows some typical features of the domestic and import price levels.

Regarding the univariate models:

- The hypothesis of one unit root in the univariate model cannot be rejected for any of the analyzed price levels including the oil prices in the currencies of the different countries.

- The hypothesis of two unit roots is rejected for all different price levels at the 5% asymptotic significance level, except for the domestic price level in the UK. The strength of the rejection differs between the different series and is most pronounced for the oil prices.

When the oil prices, denoted in the currencies of the different countries, are incorporated as a weakly exogenous variable in the univariate model for domestic and import prices, the results change compared to the univariate models.

- For the oil exporting countries, Norway and the UK, cointegration between domestic prices and oil prices cannot be rejected.
- For all countries, except Germany, Norway and the US, the hypothesis of cointegration between oil prices and import prices cannot be rejected at the 10% significance level. So, the random walk behavior of these price levels is mainly caused by the random walk behavior of the oil price.
- When no cointegration in the level is imposed ($\pi_3 = 0$), the hypothesis of no cointegration ($\pi_4 = 0$) between the first differences of the price levels and the oil prices is rejected for all countries. Note that this rejection is much more pronounced than the rejection in case of no weakly exogenous variables (Although when this hypothesis is rejected, the prices can still be $I(2)$ since the nonzeroness of π_4 could indicate cointegration between the first differences of $I(2)$ variables. This is, however, rejected through the rejection of the hypothesis of two unit roots in the univariate models for the oil prices).
- Also for the domestic price level in the UK for which in a univariate setting the hypothesis of two unit roots could not be rejected ($I(2)$), the hypothesis that $\pi_4 = 0$ is rejected. So, the univariate analysis indicates that when we neglect the oil price, the domestic price in the UK can be characterized by a $I(2)$ process which is rejected by the model incorporating the oil price as a weakly exogenous variable.

Summarizing, the long memory property of the price levels reflected in their possible $I(2)$-ness is not the result of two unit roots in an univariate model of the specific price level, which leads to a random walk in the first differences, but the result of two unit roots in the model where oil prices are incorporated as a weakly exogenous variable (one unit root from the oil prices and one unit root from the model itself). So, the long memory property of the inflation rates is the result of two random walks but these are not added cumulatively (as in the univariate model) but independently (as in the model with the weakly exogenous variables).

5. Modelling aggregate imports

5.1 *Long run import demand models*

In order to investigate the applicability of cointegration models with weakly exogenous variables, we also consider the problem of estimating long run price and income effects in international trade (more precisely, in aggregate imports). The econometric modelling of aggregate trade flows has a long history in the economic and applied econometric literature, see the survey of Goldstein and Khan (1985).

One of the reasons for this large number of studies is the simplicity of the economic theoretical framework for the determination of prices and trade volumes, which is familiar from standard consumer demand or production theory. Also, the effectiveness of international trade policy depends on the size of price and activity effects in trade flows. So, policy makers show important interest in reliable estimates of these parameters. From an econometric point of view it has been surprising that, among all fields of applied econometrics, researchers dealing with the modelling of trade flows have been reluctant to integrate the recent advances in time series econometrics. Exceptions are, among others, the contributions by Gagnon (1988), Husted and Kollintzas (1987), Haynes and Stone (1985) and Clarida (1991), who develop theoretical based dynamic stochastic models for aggregate imports and Urbain (1990, 1992*a*) and Asseery and Peel (1991), who use cointegration theory for the empirical modelling of aggregate imports. We restrict our attention to total aggregate import flows so that the basic underlying economic theory can be quite simple. Traditionally the basic question is whether imports (and exports) are perfect substitutes for domestically produced goods. Since two-way trade is usually observed, i.e. imports, exports, domestic production and intra-industry trade, a perfect substitute model is ruled out for the modelling of international aggregate trade flows. Within the framework of the imperfect substitute model (and under the assumption of infinite supply elasticity), a prototypical *long run* import demand model reads as,

$$m_t = f(gnp_t, pm_t, pd_t), \tag{5.1}$$

where m_t is import volume, gnp_t some activity or demand variable (usually real domestic product), pd_t the domestic price of tradable goods and pm_t the import prices expressed in the domestic currency. Usually, price homogeneity is imposed so that price effects are captured using a relative price ratio defined as $pr_t = pm_t/pd_t$. Such a homogeneity assumption is, however, questionable, see Urbain (1993), since imports and domestic products are normally quite different in nature. To allow for heterogeneity between imports and domestic products, we incorporate an additional variable in the long run import demand model, i.e. the oil price denoted in the national currency. Since the oil price affects the economies of the different countries in an entirely different way, depending on whether they produce any oil or even export it, the oil price is well suited for incorporating heterogeneity effects in the long run import demand model. After correcting for the fluctuations in the domestic and import price level caused by the oil price, we can then still test for the homogeneity in that parts of domestic and import price levels which are not related to the oil price. We shall therefore not impose the restriction for price homogeneity and allow the prices to have nonsymmetric effects, even in the long run. Later on, the homogeneity assumption will be tested.

We report empirical results for Belgium, France, Germany, Norway, Japan, the UK and the U.S. over a sample which roughly covers the period 1960–1990. The motivation for choosing these countries stems from their rather different nature. This dataset contains some of the main developed economies in the world and both exists of some developed oil exporting as oil importing countries.

5.2 *The data*

The data used in the analysis are quarterly figures taken from the International Monetary Fund files and are

- import volume, denoted by m_t,
- import prices expressed in domestic currency, denoted by pm_t,
- domestic prices (wholesale price index), denoted by pd_t,
- real gross product (GDP or GNP deflated by pd), denoted by gnp_t,
- crude oil price expressed in the currency of the specific country, denoted by oil_t.

All variables are transformed to natural logarithms. The calculations have been performed using TSP. Since the data are quarterly time series, the problem of seasonality has to be accounted for. Applying univariate seasonal unit root tests, the presence of unit roots at seasonal frequencies is rejected for all series when centered seasonal dummies were used (complete results available upon request). Given the lack of support for the presence of seasonal unit roots, we use deterministic centered seasonal dummies in the subsequent multivariate models. For each series, the presence of an unit root at the zero frequency is, however, not rejected using Augmented Dickey-Fuller tests. In the previous section, the issue of unit roots in the first differences has been discussed and it was concluded that incorporation of the oil prices leads to strong rejection of unit roots in the first differences. So, our series seem well characterized as $I(1)$ processes displaying a single unit root in their autoregressive part and a non zero drift.

5.3 *Estimating long run import demand models using ECMs*

Resuming the statements made in the previous sections, cointegration is tested in the model

$$(5.2)\qquad \left(\Delta\begin{pmatrix} x_{1t} \\ x_{2t} \end{pmatrix}\right) = c + \gamma\Delta x_{3t} + \sum_{i=1}^{3}\phi_i SD_{it} + \sum_{i=1}^{g}\varphi_i D_{it} + \sum_{i=1}^{p-1}\Psi_i\left(\Delta\begin{pmatrix} x_{1t-i} \\ x_{2t-i} \\ x_{3t-i} \end{pmatrix}\right) + \Pi\begin{pmatrix} x_{1t-p} \\ x_{2t-p} \\ x_{3t-p} \end{pmatrix} + \epsilon_t$$

where SD_{it} represents the centered seasonal dummies, D_{it} represents dummies which are incorporated to account for certain large outliers or important events, which by no means could be explained by the data (for example the political instability in France in 1968). A constant term is incorporated in the model to capture linear trend components which also appear in the model through the incorporation of a trending weakly exogenous variable, the oil price.

Preliminary to a cointegration analysis, the lag length of the ECM must be determined. As shown in Boswijk and Franses (1992), the choice of the lag length used in the specification of an ECM like (30) can affect the number of cointegrating vectors found in the analysis. We therefore analyzed various candidate models and based our final choice of the lag length on the absence of serial correlation in the residuals as well as on the significance of the parameter estimates of the short run coefficients.

Although we can for example compute the various Wald test statistics for cointegration (for $r = 1, 2, 3, 4$) for various causal ordering of the variables, it is tempting to use some prior knowledge of the imperfect substitute model for the computation of these statistics. In particular, we expect the economic model underlying our ECM specification to have at most three long run relations: one involving all variables and being interpretable as a long run import demand model, and two other long run relationships, mainly interpretable as a long run relationships between the domestic and import price levels and the oil price. This suggests the following ordering of the variables, *m*, *pd*, *pm*, *gnp* (and *oil*), such that the problem of lower rank values of α_{11} for the Wald cointegration testing procedure, is accounted for.

A problem which we encountered when estimating the various VAR models for the different countries was the evidence of substantial residual non-normality and ARCH effects. Although this non-normality (as well as significant ARCH effects) is mainly due to a few outliers which are not picked by the variables selected in this study, a number of impulse dummies were required to "filter" out their effect.

The weak exogeneity of the oil price is also tested by estimating an ECM for the oil price, which incorporates the estimated cointegrating relationships. The significance of the loading factors is then tested using a LM test, which did not reject the hypothesis of weak exogeneity for any of the analyzed countries.

The results obtained from estimating the ECMs for the different countries can broadly be separated into three different parts:

- number of cointegrating vectors/unit roots
- the results of the oil importing/exporting status
- long run price homogeneity between import and domestic prices.

5.3.1 *Number of cointegrating vectors/unit roots*

The results on the number of cointegrating vectors/unit roots in the ECMs for the different countries are quite uniform and Table 2 shows the LR statistics:

- for all countries, the hypothesis of no or one cointegrating vector is rejected at the 5% asymptotic significance level (except for Belgium and Norway where the hypothesis of one cointegrating vector is rejected at the 15% asymptotic significance level).
- the hypothesis of two cointegrating vectors is for all countries rejected at approximately the 50% asymptotic significance level.

Table 2. Likelihood ratio tests for the number of cointegrating vectors r.

r	US	UK	Germany	France	Belgium	Norway	Japan	C.V. (5%)
0	63.6	83.3	65.2	84.0	74.9	57.9	98.0	63.5
1	31.6	37.3	33.2	45.6	34.8	30.8	55.1	42.5
2	15.9	11.6	13.7	16.0	14.5	12.5	31.2	25.8
3	3.04	2.65	2.10	6.11	3.76	3.35	10.7	12.0

Table 3. Cointegrating relationship between import, GNP and oil price.

m	US	Japan	Germany	France	Belgium	Norway	UK
GNP	0.81*	0.99*	0.99*	0.65*	0.91*	1.30*	1.52*
oil	−0.11	0.003	−0.08	−0.15*	−0.22*	−0.08	0.09

Table 4. Cointegrating relationship between domestic price, GNP and oil price.

pd	US	Japan	Germany	France	Belgium	Norway	UK
GNP	0.5*	0.13*	0.34*	0.44*	0.33*	1.42*	1.80*
oil	0.18*	0.16*	0.11*	0.095	0.16*	0.10	0.57*

Table 5. Cointegrating relationship between import price, GNP and oil price.

pm	US	Japan	Germany	France	Belgium	Norway	UK
GNP	0.59*	−0.002	0.21*	0.41*	0.26	1.02*	1.37*
oil	0.25*	0.48*	0.23*	0.25*	0.42*	0.15*	0.66*

Given the uniformity of the results we choose for three cointegrating vectors given the asymptotic p-value of the cointegration tests for two cointegrating vectors (approximately 50%) and the interpretability of the results with three cointegrating vectors. Furthermore, in case of assuming two cointegrating vectors, these cointegrating vectors are almost an exact combination of the cointegration vectors which result when assuming three cointegrating vectors. Note that the assumption of three cointegrating vectors leads to two unit roots since we have incorporated a weakly exogenous variable which contains a unit root.

5.3.2 *The results of the oil importing/exporting status*

Assuming three cointegrating vectors and the variable ordering outlined in a previous section, the cointegrating vectors describe long run relationships between

- imports, GNP and the oil price
- domestic price, GNP and the oil price
- import price, GNP and the oil price.

Tables 3–5 contain a summary of the results. These tables contain the parameters estimates of the different elements of the cointegrating vector when we assume three cointegrating vectors (The estimates which are significant at the 5% asymptotic significance level are indicated with a^*).

It is typical that the resulting parameter estimates of the elements of the cointegrating vectors can be classified with respect to the oil importing/exporting status of the specific country. For all three relationships, it holds that GNP is a more important contributor to the cointegrating relationship when the specific country is an oil exporter. This result is shown in the relative value of the GNP

Table 6. Likelihood ratio tests for long run price homogeneity between domestic and import prices.

	US	UK	Germany	France	Belgium	Norway	Japan
	5.41	5.10	4.75	4.86	5.47	6.1	16.1
C.V. 5%	5.99	5.99	5.99	5.99	5.99	5.99	5.99

parameter with respect to the oil price parameter and the parameter value of the dependent variable in the cointegrating relationship, 1. Next to these two oil exporting countries, the UK and Norway, where the different parameter values can be attributed to their oil exportingness (for these countries Section 3 also shows that in a bivariate model, domestic price cointegrates with the oil price), the results for Japan are rather typical. For Japan, import price solely cointegrates with the oil price. So, in the long run the Japanese import prices are almost not influenced by GNP, which is the case for all the other countries. Consequently as domestic prices in Japan do depend on GNP in the long run, price homogeneity between domestic and import prices is rejected (see also the next section).

Note also that these conclusions regarding the dependence with respect to the oil price are only found when assuming three cointegrating relationships. This point also holds for the issues raised in the next section parts and this is, next to the asymptotic p value of 50% for a test for two cointegrating vectors, the reason why we prefer three cointegrating vectors (two unit roots).

5.3.3 *Long run price homogeneity between import and domestic prices*

Although oil prices are incorporated in the model to allow for heterogeneity effects in the long run import demand model, it is still possible to test for long run price homogeneity between import and domestic prices. This hypothesis can be tested by testing the equality of the parameters of the long run relationships of domestic and import prices. In Table 6, the value of the likelihood ratio tests for this hypothesis for the different countries are stated.

Table 6 shows that long run price homogeneity is essentially only rejected for Japan while for the other countries, the values of the test statistics lie in the neighborhood of their 5 or 10% asymptotic significance levels. This shows that only for Japan the whole sale price index really differs from the import price index. This again shows the difference between the Japanese economy and the other economies as in these other economies imports have a rather similar kind of composition as own production, which is shown by the non-rejection of price homogeneity, while this does not hold for Japan.

Note again that the hypothesis of price homogeneity cannot be rejected when assuming three cointegrating relationships. When assuming two cointegrating relationships, the hypothesis of price homogeneity is rejected for all countries, since the long run relationship of domestic prices has a significant coefficient for GNP, indicating that import and domestic prices are not homogeneous in the long run. Note also that when the hypothesis of price homogeneity cannot be rejected, the cointegrating relationship including imports can be interpreted as a long run import demand relationship.

6. Conclusions

The paper shows the applicability of cointegration techniques allowing for weakly exogenous variables. These techniques are applied to analyze the long run behavior of price levels and long run import demand in seven different countries. With respect to the first application, we find that the long run memory property of inflation rates crucially depends on omitting certain weakly exogenous variables, like the oil price. When this variable is incorporated as weakly exogenous to the long run parameters of the price levels, the long run memory properties of inflation rates decrease substantially. The possible $I(2)$-ness of prices is strongly rejected in this case but still the hypothesis of two unit roots cannot be rejected in most of the models for the price levels. The difference with the $I(2)$ model is, however, that the resulting random walks are not added cumulatively but independently because one of the random walks results from the weakly exogenous variable. The incorporation of the oil price in the import demand analysis leads to some typical results depending on the importing or exporting status with respect to oil (related) products of the specific country. For oil importing countries, the three long run equillibrium relationships depend to a much smaller extent on GNP than for the oil exporting countries. Furthermore, the import prices of Japan are almost independent for GNP, which does not hold for any of the other countries.

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