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Import Price Formation and Pricing to Market: A Test on Norwegian Data

Discussion Papers

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Abstract:
This paper investigates the determinants of Norwegian import prices of manufactures over the period 1970(1) - 1991(4). Multivariate cointegration analysis establishes a long-run relationship between import prices, foreign prices, the exchange rate and domestic unit labour costs. Normalized on import prices, the long-run elasticities are 0.63 (foreign prices and the exchange rate) and 0.37 (domestic costs). Deviations from this relationship are highly significant in a structural import price equation, which also contains positive effects of growth in domestic demand and inflation, as well as a negative effect from the Norwegian unemployment rate. The estimated parameters appear reasonably stable within the sample.

Keywords: Import price formation, pricing to market, domestic effects, Johansen procedure, structural error correction model, super exogeneity.

JEL classification: C32, C51, C52, C22, D40, F41, L16.

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1 Introduction

Economic analysis of open economies should account for several causal relationships between prices of imports, the current account, inflation and the level of domestic economic activity. First, the trade balance is directly affected by import prices through the terms of trade effect. Next, domestic firms often face foreign competition, and relative prices between imported and domestically produced goods are therefore prime determinants of manufacturing output and import volumes. It is also a fact that price inflation in small open economies is severely affected by the growth in prices on imported goods. In the case of Norway, the weight of imports in the official consumer price index is approximately 1/5. Another 20 percent of the index is made up of goods which are indirectly influenced by the prices of imports, i.e. through import competition and imported material inputs. Finally, wages are likely to respond to increases in import prices, either directly or through increased consumer and producer prices. It follows that knowledge of import price formation is important for the understanding of wage-price inflation.

Conventional theory-models of small open economies pay little attention to the potential role of domestic factors in the determination of import prices. The assumption concerning import price formation in these models states that the "law of one price" holds: A given traded good is sold for the same price in all destinations when it is measured in a common currency unit. However, several empirical studies have produced results that contradict the usual small open economy assumption. There is evidence that foreign producers; *a)* adjust import prices to changes in prices or costs in the importing country; and *b)* do not completely pass through shocks in the exchange rate to the prices of imports, i.e. "prices to market" in the terminology of Krugman (1987), given these prices/costs. See e.g. Llewellyn (1974), Llewellyn and Pesaran (1976), Bond (1981), Barker (1987) and Menon (1995). These phenomena can be explained by theories of imperfect competition and price discrimination (Dornbusch (1987) and Krugman (1987)).

The purpose of the present study is to investigate econometrically the role played by foreign prices, the exchange rate and domestic factors in the formation of Norwegian import prices over the period 1970(1)-1991(4). We first analyze a VAR-model using cointegration techniques, and then develop a structural import price equation. The rest of the paper is organized as follows. In section 2, we derive hypotheses about long-run behaviour from a simple model of import pricing. Section 3 gives a brief description and discussion of the data series used in the empirical analysis. In section 4, we test the hypotheses derived in section 2, using the multivariate cointegration procedure proposed by Johansen (1988). Norwegian import prices of manufactures do not cointegrate with foreign prices and the exchange rate, but there is significant cointegration between these variables and domestic unit labour costs. Moreover, tests of weak exogeneity suggest that deviations from the estimated cointegrating vector is corrected by import prices only. Normalizing a restricted cointegrating vector on import prices yields long-run elasticities of 0.63 (world prices and the exchange rate) and 0.37 (domestic costs). In section 5, we formulate a structural error correction model of import prices, using the long-run estimates from the VAR-analysis in the error correction term. The structural model also contains positive effects of growth in domestic demand and inflation, as well as a negative coefficient for the Norwegian unemployment rate. In Section 6, we show that the estimated parameters of the structural model are reasonably stable. The equation is also found to be invariant towards the discrete changes in the exchange rate that occurred during the sample period. Section 7 concludes the paper.

2 Economic background

The theoretical framework adopted here draws on Krugman (1987), Dornbusch (1987) and Hooper and Mann (1989). It is assumed that producers of differentiated tradeable goods sell in several markets characterized by imperfect competition. These markets are segmented as a consequence of transportation costs, trade barriers and imperfect information. In this situation, profit maximization normally implies price discrimination; the prices charged reflect conditions in each particular market. To formalize, a representative foreign firm sets the price of exports to destination country i as a mark-up over its marginal costs:

$$(1) \qquad PX_i = \theta_i C^*, \quad i = 1, .., n.$$

PX_i denotes the export price of the good and C^* the foreign producer's marginal costs, both in the currency of the exporting country. n is the number of export markets and θ_i is the destination-specific mark-up. To obtain the import price in the currency of the importing country, PB_i, both sides of (1) are multiplied by the (bilateral) exchange rate, E_i:

$$(2) \qquad PB_i = E_i\, PX_i = E_i \theta_i C^*, \qquad \forall\, i.$$

The mark-up is taken to be a function of the price on competing goods produced in destination i, PH_i, relative to the import price and (a measure of) demand pressure in the importing country, DP_i:[1]

$$(3) \qquad \theta_i = K_i\, [PH_i/PB_i]^{\tau_{1i}}\, DP_i^{-\tau_{2i}}, \qquad \tau_{1i} \geq 0,\ \tau_{2i} \lesseqgtr 0, \qquad \forall\, i,$$

where K_i is a constant. Note that the sign of τ_{2i} is undetermined from theory. Solving (2) and (3) and taking logs yields:

$$(4) \qquad pb_i = k_i + (1 - \phi_i)c^* + (1 - \phi_i)e_i + \phi_i ph_i - \varphi_i dp_i, \qquad \forall\, i,$$

[1] We abstract from competition between foreign exportes in market i, which would imply additional relative prices in equation (3).

with $k_i = \ln K_i/(1+\tau_{1i})$, $\phi_i = \tau_{1i}/(1+\tau_{1i})$, and $\varphi_i = \tau_{2i}/(1+\tau_{1i})$. Here and in the following, lower case letters denote logarithmically transformed variables.

The coefficient $(1-\phi_i)$ measures the degree of pass-through of changes in marginal costs and the exchange rate to import prices, given the levels of ph_i and dp_i. Owing to the competitive pressures in the importing market changes in C^* and E_i are not fully passed trough to import prices as long as ph_i is unchanged, a phenomenon dubbed "pricing to market"—henceforth denoted PTM— by Krugman (1987). In the limiting case where $\phi_i = 0$, shifts in marginal costs and the exchange rate are completely passed through to import prices, and the weight of competing prices is zero in (4).

The "law of one price" (LOP)—the conventional assumption concerning import pricing in theoretical models for small open economies—follows from a *special case* of (6). The absolute version of LOP holds when $\varphi_i = \phi_i = 0$ and $k_i = k > 0 \; \forall \; i$, which implies that $PX_i = PX$ in all destinations. The relative version of this law holds under the weaker conditions $\varphi_i = \phi_i = 0$, $\forall \; i$, which is consistent with price discriminating behaviour since $\theta_i = K_i$ may vary between markets.

Equation (4) is based on a model of foreign firms' behaviour, and identifies *one* mechanism whereby domestic factors can affect import prices. Another channel is represented by domestic importing firms acting as agents for foreign products. If importers emphasize domestic price and demand conditions in negotiations with foreign exporters, this is an additional rationale for the presence of ph_i and dp_i in equation (4).

Data series for foreign *marginal* costs, C^*, are not available, hence (4) cannot be estimated directly. An alternative route is to employ the average export price, PX, defined as:

$$(5) \qquad PX = \prod_{i=1}^{n} PX_i^{\omega_i}, \quad 0 < \omega_i < 1, \quad \sum_{i=1}^{n} \omega_i = 1,$$

where ω_i is the weight of market i. Using (2), (4) and (5), the import price equation can be expressed as:

$$(6)\, pb_i = k_i + (1-\phi_i)px + (1-\phi_i)e_i + \phi_i ph_i - \varphi_i dp_i - (1-\phi_i)\sum_{i=1}^{n}\omega_i \ln\theta_i, \quad \forall \; i.$$

We consider aggregate time series data for *one* destination, e.g. Norway. Hence, in the following we omit the index i and introduce the subscript t, denoting time. Aggregating over foreign suppliers and allowing stochastic variation yields the following empirical "counterpart" to (6):

$$(7) \qquad pb_t = const + (1-\phi)px_t + (1-\phi)e_t + \phi ph_t - \varphi dp_t + \zeta_t.$$

Note that, in addition to random variation, the ζ_t disturbance term subsumes the mark-ups of foreign exporters in all markets, the consequences of which are discussed below. Since price indices are used instead of disaggregated price levels in the empirical analysis, k_i is not identified in (7). Consequently, the absolute version of LOP is not testable within this framework. Rather, we investigate implications of absolute and relative LOP (whether $\varphi_i = \phi_i = 0$) for one country, Norway.

It is likely that import prices adjust *gradually* to changes in the explanatory variables in (7), so econometric import price equations ought to be specified dynamically, while equation (7) is static. However, (7) may be interpreted as a *long-run* relationship and will therefore serve as the starting point for the cointegration analysis in section 4.

It is reasonable to assume that the price series in (7) are non-stationary, $I(1)$, where $I(1)$ denotes "integrated of order one". An empirical implication of LOP is then that the logs of import prices, foreign prices and the exchange rate cointegrate with cointegration parameters equal to one, i.e. that $(pb - px - e)_t$ is stationary, $I(0)$. Since cointegration (or lack thereof) characterizes long term properties of data series, this is stated as the *long-run* version of LOP. Correspondingly, we define long-run purchasing power parity (PPP) as a situation where $(ph - px - e)_t \sim I(0)$.[2] Interestingly, the long-run versions of LOP and PPP may be consistent with PTM, $\phi > 0$, and the existence of domestic demand effects, $\varphi \neq 0$. To see this, it is useful to reformulate (7) as:

$$(pb - px - e)_t = const + \phi(ph - px - e)_t - \varphi dp_t + \zeta_t \tag{8}$$

where ζ_t is taken to be $I(0)$, since it measures the deviation from the long-run equilibrium. If $(pb - px - e)_t$ and $(ph - px - e)_t$ are $I(0)$, (8) is a balanced equation with $\phi > 0$ and $\varphi > 0$ under the plausible assumption $dp_t \sim I(0)$. However, since both LOP and PPP hold as long-run relationships, PTM is strictly a *short-run* phenomenon and therefore not accountable for any significant part of the variation in import prices. Another possibility is that $(pb-px-e)_t \sim I(1)$ and $(ph-px-e)_t \sim I(1)$, i.e. neither LOP nor PPP hold in the long-run, while $(pb - px - e)_t - \phi(ph - px - e)_t \sim I(0)$. In this case, PTM is a *long-run* phenomenon.

3 The data[3]

In the empirical analysis we use quarterly, seasonally unadjusted, data for the period 1968(1) – 1991(4). After allowing for lags, the sample period used for estimation is 1970(1) – 1991(4). Import prices, foreign prices and the exchange rate are measured by indices with 1990 as the base year. The import price index is an implicit deflator for manufactured products with Norwegian substitutes. Lawrence (1990) shows that, over the period 1980 – 1990, the growth in implicit deflators of U.S. non-oil imports

[2]See Johansen and Juselius (1992) for a similar interpretation of the PPP-hypothesis.

[3]Detailed data definitions are given in appendix A.

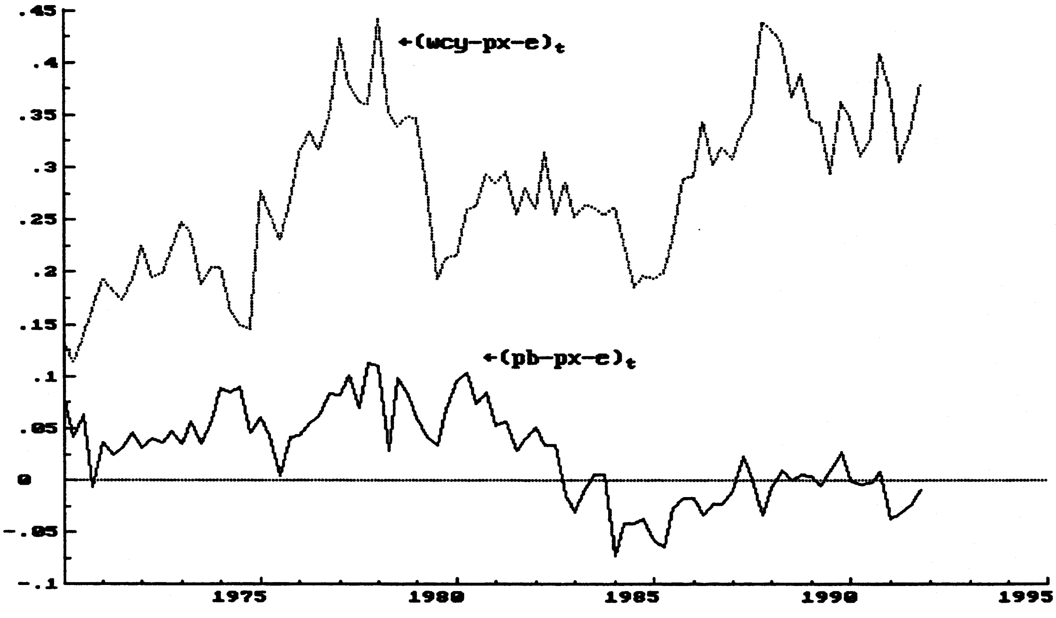

Figure 1: The log of relative prices, $(pb - px - e)_t$, and the log of competetiveness, $(wcy - px - e)_t$.

to a large extent hinges on whether computers are included or not. This is due to the increasing weight of computers in import volumes and the fall in computer prices in that period. In order to adjust for this, Lawrence (1990) excludes computers from the import price index, while Hooper and Mann (1989) use a fixed weight index in their analysis of U.S. import prices. Our import price index is not significantly affected if we exclude computers, and it shows more or less the same development over the sample period as a fixed-weight index of import prices.[4] Also, all the main results reported below continue to hold when the implicit deflator is replaced by this fixed-weight index.

The series for foreign prices is an import weighted index containing implicit deflators for merchandise exports from the 14 countries accounting for the bulk of Norwegian imports of goods. The exchange rate index is constructed using the same set of weights.

We use unit labour costs in manufacturing, *WCY*, as a proxy for domestic prices on import competing products. Growth in domestic absorption, the rate of unemployment and the inflation rate are used as indicators of demand pressure or general market conditions in Norway. Growth in domestic absorption is a straight-

[4] The fixed weights were obtained by first splitting the aggregate index into 18 sub-indices according to the commodity classification in the quarterly national accounts. Next, these indices were weighted by the average import-share weights.

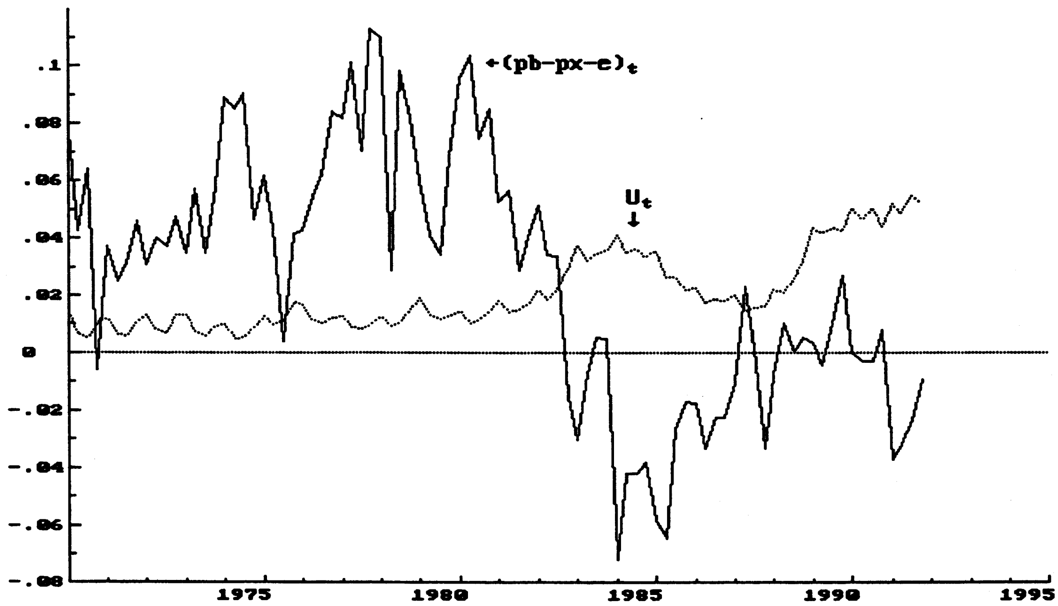

Figure 2: The log of relative prices, $(pb - px - e)_t$, and the rate of unemployment, U_t.

forward measure of demand pressure. Inflation and unemployment rates are easily observable, and may therefore be used by foreign firms to assess the cyclical stance of the Norwegian economy when it is considered costly to collect detailed market information.

The use of PX and WCY as proxies for C^* and PH may pose problems for the interpretation of the empirical findings. First, px_t is correlated with ζ_t in (7), since ζ_t contains the foreign producers' mark-ups in all export markets. This measurement error does not induce asymptotic bias in the estimated cointegrating parameters as long as it is $I(0)$. If the measurement error is $I(1)$, ζ_t is also $I(1)$, and (7) does not form a cointegration relationship. Tests of cointegration have zero power in this case, cointegration is effectively "hidden" (see Eitrheim (1991)). Second, replacing PH with WCY, leaves the mark-ups of domestic producers in the error term. This measurement error is presumably $I(0)$ and uncorrelated with WCY, but may well be correlated with the variables representing DP_t; if demand pressure in Norway affects the mark-ups of foreign firms exporting to Norway, we would also expect the mark-ups of domestic producers to respond. Hence, the importance of demand effects may be overstated, given the level of domestic prices.

Figure 1 shows relative prices, $(pb - px - e)_t$, together with domestic manufacturing unit labour costs relative to world prices measured in Norwegian currency,

$(wcy - px - e)_t$, which is a measure of competitiveness.[5] As can be seen from the figure, unit labour costs increased more than foreign export prices in Norwegian currency over the sample, and $(wcy - px - e)_t \sim I(0)$ does not seem to hold. Whether $(pb - px - e)_t$ is $I(0)$ or $I(1)$ is not obvious from Figure 1, but augmented Dickey-Fuller tests do not reject a null of $(pb - px - e)_t \sim I(1)$. Norwegian import prices increased relatively to world prices from 1970 to 1980, decreased considerably from 1981 to 1984, and then picked up from 1985 to 1987. The development of $(pb - px - e)_t$ from 1981 to 1987 may to some extent be explained by the significant movements in the U.S. dollar against other currencies. As predicted by the PTM theory, foreign firms seem to have increased their mark-ups in the U.S. market when the dollar appreciated from 1981 to 1985 and decreased their mark-ups during the subsequent depreciation between 1985 and 1987.[6] Figure 2 offers another story, consistent with the theoretical framework in section 2: Over the period 1981-1987, the development of $(pb - px - e)_t$ is matched by changes in domestic demand pressure, proxied by variations in the unemployment rate, U_t. We will pursue this possibility further in section 5 below.

The price series, unit labour costs, the exchange rate and domestic absorption all seem to be $I(1)$. For the rate of unemployment it is more difficult decide wether to adopt an $I(0)$ or $I(1)$ assumption. Figure 2 shows some evidence of mean-reversion in the unemployment rate, but the apparent upward tendency after 1988 might suggest a drift-component or possibly a structural break, resulting in a shift in the unconditional mean of the series. In the following we interpret u_t as an I(0) series, but with possible structural breaks. This assumption is supported by the results in Bjørnland (1995) and Johansen (1995).

The sample period covers two "generations" of quarterly national accounts, pre- and post 1978(1), with differences in the seasonal pattern. The seasonality of the unemployment rate also seems to have changed around 1978, cf. Figure 2. To take account of the seasonality in the series and the changes thereof in 1978, we use three centered seasonal dummy variables and a step dummy, QBR_t,[7] interactively with the seasonal dummies.

4 Cointegration

The error correction representation of the vector autoregressive model (VAR) of order k is expressed as:

$$(9)\, \Delta \mathbf{x}_t = \boldsymbol{\mu} + \boldsymbol{\pi} \mathbf{x}_{t-1} + \sum_{i=1}^{k-1} \gamma_i \Delta \mathbf{x}_{t-i} + \boldsymbol{\kappa} \mathbf{D}_t + \boldsymbol{\epsilon}_t, \quad \boldsymbol{\epsilon}_t \sim NIID_p(\mathbf{0}, \boldsymbol{\Omega}). \quad t = 1, ..., T.$$

[5]In Figure 1, the scale of $(wcy - px - e)_t$ is adjusted to match that of $(pb - px - e)_t$.

[6]See e.g. Dornbusch (1987), Krugman (1987), Hooper and Mann (1989) and Lawrence (1990).

[7]QBR_t is one until 1977.4, zero thereafter.

Table 1: Mis-specification tests

Statistic	pb	px	e	wcy	VAR
$F_{AR\,1-5}(5,45)$	0.33	1.68	0.76	1.42	
$F_{ARCH\,1-4}(4,42)$	0.60	0.44	0.45	0.47	
$F_{HET\,x_i^2}(32,17)$	0.33	0.39	0.37	0.26	
$\chi^2_{ND}(2)$	3.12	0.25	2.49	2.10	
$F^V_{AR\,1-5}(80,108)$					1.10
$F^V_{HET\,x_i^2}(320,104)$					0.22
$\chi^{2V}_{ND}(8)$					6.99

where $\mathbf{x}_t$ is a $(p \times 1)$ vector of modelled variables. The vector $\mathbf{D}_t$ consists of deterministic variables (e.g. dummies for seasonality and interventions) and conditioning, stochastic, $I(0)$ variables.

Assuming $\mathbf{x}_t$ to be $I(1)$, absence of cointegration implies $\boldsymbol{\pi} = \mathbf{0}$. In the case of cointegration, $0 < r < p$, where r denotes the rank of $\boldsymbol{\pi}$, the number of independent cointegrating vectors. $\boldsymbol{\pi}$ can now be decomposed as $\boldsymbol{\alpha\beta}'$, where $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are $p \times r$ matrices. The $\boldsymbol{\beta}$ matrix is made up of the r vectors of cointegrating parameters, and $\boldsymbol{\alpha}$ contains the adjustment coefficients.

In the following, $\mathbf{x}_t$ in (9) consists of four variables (i.e. $p = 4$) that were discussed in the previous section: the import price index (pb_t), the world market price variable (px_t), the exchange rate (e_t) and domestic unit labour costs (wcy_t). We include the dummy variables described in section 4 and the log of the unemployment rate lagged one quarter, u_{t-1}, as conditioning variables. The intercept term is kept unrestricted in order to account for drift in the individual price series. Initially, we estimated a 5th order VAR based on this information set. However, although the residuals in the equation for Δpb_t were empirically white noise and normally distributed, this was not the case for the other equations in the system. This is not surprising since the included variables were chosen to explain Δpb_t, not Δpx_t, Δe_t and Δwcy_t. To deal with this problem, we included a set of dummy variables to account for outliers and structural breaks.

The results below are based on a 4th order VAR with 14 intervention dummies. We employ two impulse dummies (PX74.1 and PX80.4) to account for large increases in px_t following the oil crises in 1973 and 1979 and 7 dummy variables designed to capture the effects of discrete devaluations and revaluations on Δe_t. Finally, we include 5 impulse dummies to mop up outliers in the equation for Δwcy_t. A detailed description of all the dummy variables is given in appendix A.

Diagnostic tests for the preferred VAR-specification are reported in Table 1 where $F_{AR\,1-5}$ is an F-test (with degrees of freedom shown in brackets) for residual autocorrelation of order 5, see Harvey (1981); $F_{ARCH\,1-4}$ is the Engle (1982) F-test for 4th order ARCH in the residuals; $F_{HET\,x_i^2}$ is the F-test for residual heteroscedas-

Table 2: Johansen cointegration tests

Eigenvalues: 0.25, 0.187, 0.074, 0.012					
null	alternative	λ_{max}	5% critical value	λ_{trace}	5% critical value
$r=0$	$r=1$	25.30	27.07	51.39	47.21
$r\leq 1$	$r=2$	18.26	20.97	26.09	29.68
$r\leq 2$	$r=3$	6.81	14.07	7.83	15.41
$r\leq 3$	$r=4$	1.02	3.76	1.02	3.76
	The critial values are taken from Table 1 in Osterwald-Lenum (1992).				

ticity due to White (1980); and χ^2_{ND} is the normality test described in Doornik and Hansen (1994); corresponding system (vector) tests are denoted by ^{V} (Doornik and Hendry (1994a)). None of the mis-specification test statistics are significant at the 10% level, so we consider the estimated system as a satisfactory representation of the data generation process.

Table 2 contains the results from applying Johansen's (1988) cointegration procedure to the VAR. It shows the four eigenvalues, the maximal eigenvalue and trace-statistics (λ_{max} and λ_{trace}) and the 5% asymptotic critical values. We have a relatively short sample, and the VAR includes several conditioning variables, so the asymptotic critical values are only approximations. With this caveat in mind, there seems to be one cointegrating vector in the data: The λ_{max} and λ_{trace} statistics reject the null of no cointegration at 10% and 5% levels respectively, but the null of at most one cointegrating vector is not rejected by any of the statistics.

Normalizing the estimated cointegrating vector on import prices yields:

$$pb = const + \underset{(0.12)}{0.53}\, px + \underset{(0.14)}{0.52}\, e + \underset{(0.10)}{0.46}\, wcy, \tag{10}$$

with standard errors in brackets. The corresponding vector of adjustment coefficients is given by:

$$\hat{\alpha}' = (-0.33, -0.03, 0.05, 0.16). \tag{11}$$

The coefficients in (10) and (11) suggest that the estimated cointegrating vector can be interpreted as a long-run import price equation consistent with the PTM hypothesis. First, (11) indicates that disequilibrium in (10) for the most part is corrected through adjustment of import prices. Second, the elasticities in (10) are highly significant, and the restriction implied by (7) are almost statisfied numerically, albeit with a coefficient for domestic costs that may seem unreasonably large in magnitude.

Formal support for this interpretation of the cointegrating vector is given in Tables 3 and 4, which report likelihood ratio statistics for tests of weak exogeneity and structural hypotheses, i.e. tests on $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$. The statistics are asymptotically distributed as χ^2 with degrees of freedom given in parenthesis. Weak exogeneity of

Table 3: Tests of weak exogeneity

Variable	pb_t	px_t	e_t	wcy_t
$\chi^2(1)$-statistic	6.56	0.32	1.33	1.94
p-value	(0.01)	(0.57)	(0.25)	(0.16)
See Johansen and Juselius (1990).				

Table 4: Tests of structural stationarity hypotheses

Hypothesis	Likelihood ratio	p-value
$H_1 : (pb_t - \kappa_1 px_t - \kappa_1 e_t - \kappa_2 wcy_t) \sim I(0)$	$\chi^2(1) = 0.01$	0.91
$H_2 : (pb - px - e)_t - \phi(wcy - px - e)_t \sim I(0)$	$\chi^2(2) = 0.04$	0.98
$H_3 : (pb - px - e)_t \sim I(0)$	$\chi^2(3) = 14.68$	0.00
$H_4 : (wcy - px - e)_t \sim I(0)$	$\chi^2(3) = 11.39$	0.01
See Johansen and Juselius (1990), (1992).		

import prices for the long-run parameters is strongly rejected in Table 3, but it seems safe to assume that foreign prices, the exchange rate and domestic costs are weakly exogenous.[8] The first hypothesis in Table 4 (H_1) tests whether the coefficients of foreign prices and the exchange rate are equal, whereas H_2 imposes the additional restriction of long-run unit homogeneity. It is seen that these hypotheses are easily accepted by the data. Moreover, the long-run versions of LOP (H_3) and PPP (H_4) (as defined in section 2) are firmly rejected. Finally, imposing H_2 and weak exogeneity of px_t, e_t and wcy_t gives $\chi^2(5) = 5.54$ (with a p-value of 0.35) and the following cointegrating vector:

$$pb = const + \underset{(0.08)}{0.63}\, px \; + \; 0.63e \; + \; 0.37wcy. \tag{12}$$

The elasticities in (12) seem more reasonable from an economic point of view than those in (10). The coefficient for domestic costs in (12) is also large in magnitude, but it is only 58 % of the coefficient for foreign prices (opposed to 87% in (10)).

5 A structural import price equation

The analysis in the previous section yields a long-run import price equation, (12), with plausible coefficients for world prices, the exchange rate and domestic costs. In this section, we focus on; a) the dynamic adjustment of import prices to changes in these variables; and b) the role of the $I(0)$ variables proxying demand pressure and general market conditions described in section 3. For this purpose, we formulate a structural import price equation using deviations from (12) as an error correction mechanism. Compared to section 4, we consider a wider information

[8]Testing the joint hypothesis of foreign prices, the exchange rate and domestic costs being weakly exogenous for the long-run parameters yields $\chi^2(3) = 4.61$ and a p-value of 0.20.

set that includes distributed lags on the Norwegian quarterly inflation rate, Δcpi_t, growth in domestic absorption, Δabs_t, and changes in the unemployment rate, Δu_t. The intervention dummies employed in the VAR analysis are not included since the effects on import prices of the specified interventions should be captured by the model if it is to be labelled "structural", see Hendry (1993). However, the validity of this assumption is investigated in section 6. The thrust of the modifications is a model of the form (13), where the error correction mechanism is defined as $ECM_t = (pb - px - e)_t - 0.37(wcy - px - e)_t$, and where the constant and the remaining dummies are suppressed to save space.

$$
\begin{aligned}
\Delta pb_t \;=\; & \textstyle\sum_{i=0}^4 \gamma_{1i}\Delta px_{t-i} + \sum_{i=0}^4 \gamma_{2i}\Delta e_{t-i} + \sum_{i=0}^4 \gamma_{3i}\Delta wcy_{t-i} + \\
& \textstyle\sum_{i=1}^4 \gamma_{4i}\Delta pb_{t-i} + \sum_{i=0}^4 \gamma_{5i}\Delta cpi_{t-i} + \sum_{i=0}^4 \gamma_{6i}\Delta abs_{t-i} + \\
& \textstyle\sum_{i=0}^4 \gamma_{7i}\Delta u_{t-i} + \rho_1 u_{t-1} + \rho_2 ECM_{t-1} + \varepsilon_t
\end{aligned} \tag{12}
$$

Equation (13) is useful in the sense that it encompasses a wide class of import price equations consistent with the cointegration analysis above, but it is also likely to represent an over-parameterization. We will therefore present a parsimonious model, (14), which is a valid simplification of (13). Equation (14) reports IV estimates with standard errors in brackets. The inflation rate is instrumented following Nymoen (1991), who found contemporaneous effects of import prices in an econometric model of Norwegian consumer prices. The conditioning on the two other contemporaneous variables, Δpx_t and Δe_t, is subject to a potential caveat: Although the results in table 3 clearly support weak exogeneity of px_t and e_t for the cointegrating vector, they need not be weakly exogenous for the other parameters in (13) and (14), see Urbain (1992) and Boswijk and Urbain (1994). In particular, the discussion of data issues in section 3 brought out that measurement errors endangers the orthogonality of Δpx_t to ε_t. However, tests of the significance of $\widehat{\Delta px}_t$ and $\widehat{\Delta e}_t$ from the VAR in both the general model (13) and the parsimonious model (14) (i.e. Wu-Hausman tests) showed that Δpx_t and Δe_t may be taken as weakly exogenous. Section 6 presents additional evidence on the exogeneity status of these variables.

$$
\begin{aligned}
\Delta(pb - px)_t = & -\underset{(0.013)}{0.027} + \underset{(0.082)}{0.222}\,\Delta_2 e_t - \underset{(0.068)}{0.229}\,\Delta pb_{t-1} \\
& + \underset{(0.179)}{0.675}\,\Delta_3\Delta cpi_t + \underset{(0.057)}{0.217}\,\Delta_3 abs_{t-1} - \underset{(0.005)}{0.024}\,u_{t-1} \\
& - \underset{(0.065)}{0.377}\,\{(pb - px - e)_{t-1} - 0.37(wcy - px - e)_{t-4}\} \\
& - \underset{(0.003)}{0.016}\,DUM_t + \hat{\varepsilon}_t
\end{aligned} \tag{13}
$$

$T = 88 \quad \{1970(1) - 1991(4)\} \quad \hat{\sigma}_{IV} = 1.73\% \quad DW = 1.97$

$\chi^2_{AR\ 1-5}(5) = 6.61 \quad F_{ARCH\ 1-4}(4,72) = 0.77 \quad F_{HET\ x_i^2}(14,65) = 0.49$

$F_{HET\ x_i x_j}(35,44) = 0.86 \quad \chi^2_{ND}(2) = 3.18 \quad \chi^2_{IV}(9) = 4.37$

Below equation (14) we report the IV residual standard error, denoted $\hat{\sigma}_{IV}$, the Durbin Watson statistic (DW), and several other mis-specification tests based on the residuals, $\hat{\varepsilon}_t$. In addition to χ^2_{ND}, $F_{ARCH\,1-4}$ and $F_{HET\,x_i^2}$ defined previously, we report $\chi^2_{AR\,1-5}$, which is the chi-square version of the LM test for 5th order residual autocorrelation; the $F_{HET\,x_ix_j}$ test of functional form mis-specification (see Doornik and Hendry (1994b)); and the Sargan (1964) test for validity of the instruments, χ^2_{IV}. None of the diagnostic tests are significant at the 10% level, and the validity of the over-identifying instruments is not rejected by the χ^2_{IV} test.[9]

According to our results, the pass-through from foreign prices (px_t) to the prices of Norwegian imports is approximately complete within the first quarter. Δpx_t therefore appears with an imposed unit coefficient in (14). Hence the short-run response of import prices to changes in foreign prices is considerably larger than the long-run effect in (14). However, since the equation includes explanatory variables which are not strongly exogenous, the multipliers based on (14) give only part of the story of how Norwegian import prices react to shocks. In the context of a macroeconometric model, the apparent overshooting with respect to changes in px_t will be mitigated by changes in wages. Although the long-run coefficient for world prices and the exchange rate is 0.63 in (14), the long-run pass through to import prices of shifts in these variables is unity when analysed within a macroeconometric model—assuming static homogeneity in all price and wage equations.

Turning to the exchange rate (e_t), the estimated impact elasticity of 0.22 is significantly *smaller* than the long-run coefficient of 0.63. The $\Delta_2 e_t$ formulation can be interpreted as a statistical smoothing of the exchange rate (by foreign firms) to extract changes which are more permanent. These results may reflect that the Norwegian exchange rate was fixed within an exchange rate band during most of the sample period: If there are costs to changing import prices, it will be rational to adjust slowly to (small) fluctuations in the exchange rate that are likely to be reversed. One problem with this explanation stems from the fact that the sample contains a number of discrete devaluations and one revaluation, cf. the "exchange rate" dummies in the VAR-model. These changes were probably perceived as permanent, in which case it may have been optimal to respond relatively fast. If indeed the short-run pass-trough depends on the specific type of currency change, the equation is subject to Lucas' (1976) criticism of econometric models being unstable in the face of policy interventions and changes in the environment, see Favero and Hendry (1992). However, the occurrence of policy instigated currency changes during the sample period makes it possible to test the empirical relevance of the Lucas critique, and this is done in section 6 below.

[9]The instruments for $\Delta_3\Delta cpi_t$ are: Δcpi_{t-3}, $\Delta cpi_{t-2} + \Delta cpi_{t-4}$, $(cpi - wcy)_{t-1}$, $(cpi - pb)_{t-1}$, cp_{t-1}, $j12_{t-1}$, $CPIDUM_t$, VAT_t, $Q2_t$ and $Q1_tQBR_t + Q1_t$. See the appendix for details.

The third quarter change in the quarterly rate of inflation ($\Delta_3\Delta cpi_t$) is highly significant in (14), with a coefficient of 0.675.[10] The explanation of this finding may be that information on growth in consumer prices is more easily available than information on the development in competing prices, and therefore the inflation rate is used as an indicator of market conditions in Norway. Next, the equation includes strongly significant lagged effects of the third quarter growth in domestic absorption ($\Delta_3 abs_{t-1}$) and the rate of unemployment (u_{t-1}), with coefficients of 0.22 and -0.024 respectively. The long-run coefficients of these variables are 1.73 (Δabs) and -0.064 (u). Accordingly, increases in domestic demand pressure results in price increases on imports.

An error correction mechanism measuring deviations from (12) enters (14) with a t-value of -5.82, hence reinforcing the results obtained in section 4. The estimated coefficient of -0.38 implies strong correction of disequilibria from the long-run relationship. Note that the error correction term is specified as $(pb - px - e)_{t-1} - 0.37(wcy - px - e)_{t-4}$, i.e. $(wcy - px - e)$ is at the *fourth* lag, a reparameterization which turned out practical because the within-year effects of unit labour costs were insignificant. The estimate of the cointegration parameter does not differ significantly from 0.37 when derived from a version of (14) that includes $(pb - px - e)_{t-1}$ and $(wcy - px - e)_{t-4}$ as separate variables instead of the ECM obtained from the Johansen analysis. Finally, DUM_t captures the effects of seasonality and the break in the seasonal pattern in some of the series after 1977.[11]

Our results are subject to caveats because of two important measurement problems discussed in section 3: a) unit labour costs is used as a proxy of prices on import-competing domestic products; and b) we employ foreign export prices instead of foreign marginal costs. In particular, the importance of the stationary domestic factors ($\Delta_3 abs_{t-1}$, u_{t-1} and $\Delta_3\Delta cpi_t$) may be overstated, since potential price increases on import-competing domestic goods following an increase in demand is not represented in the model. To shed light on the robustness issue, we constructed a price index of domestic products sold on the home market, PHD_t, and an index of foreign producer prices, PPX_t[12]. Equations very similar to (14) were estimated when PHD_t was substituted for WCY_t and when PPX_t was used instead of PX_t. The fit and stability were poorer than for (14), but the coefficients and significance of $\Delta_3\Delta cpi_t$, $\Delta_3 abs_{t-1}$ and u_{t-1} were only slightly affected. The results

[10] Nymoen (1991), using the same import price index as in our study, finds a short-run elasticity of 0.035 from import prices to Norwegian consumer prices. Employing this estimate and the coefficients from equation (14), the estimated short-run pass-through from foreign prices and the exchange rate to import prices is 1.02 and 0.25 respectively.

[11] i.e. the dummies are amalgamated into $DUM_t = 4Q1_t + Q2_t + Q3_t - Q1_tQBR_t - Q2_tQBR_t + Q3_tQBR_t$ in order to simplify the model and to facilitate recursive estimation.

[12] PPX_t ought to be substantially less influenced by fluctuations in the U.S. dollar exchange rate than PX_t, cf. the discussion of Figure 2.

for the alternative models also confirmed the role of domestic costs/prices in the cointegrating vector.

Previous econometric work on Norwegian import prices include von der Fehr (1987) and Naug (1990). von der Fehr (1987) estimated disaggregated equations using annual data. In his equations for manufactured products with Norwegian substitutes, the long-run elasticity of domestic prices varies between zero and 0.7. The weighted average of these coefficients is approximately 0.5. In Naug (1990), the estimated long-run elasticity of domestic prices is 0.24 in an equation covering total imports, excluding crude-oil. None of these studies included variables proxying demand pressure or general market conditions in Norway. Interestingly, the estimates in (12) are close to those found by Hooper and Mann (1989) and Lawrence (1990) on U.S. import prices and Menon (1995) on Australian data. Hooper and Mann (1989, equation 10) report long-run elasticities of foreign unit labour costs and U.S. prices of 0.61 and 0.39 respectively. Lawrence (1990, table 7) finds long-run elasticities of 0.7 (foreign export prices) and 0.3 (U.S. domestic prices). The estimates in Menon (1995), which were obtained using the Johansen procedure, are 0.66 (foreign costs), 0.75 (the exchange rate) and 0.37 (domestic competing prices). In the UK-studies by Llewellyn (1974), Llewellyn and Pesaran (1976), Bond (1981) and Barker (1987), the weight of domestic prices/costs lie in the range $0.10 - 0.40$.

6 Stability, exogeneity and invariance

In this section we investigate to what extent the import price equation (14) can be viewed as a stable and invariant relationship. Stability and invariance are conceptually different. *Stability* refers to parameter constancy over time, whereas *invariance* refers to constancy across regimes and (policy) interventions. Parameter stability is a criterion of good model design, along with white noise residuals. Invariance is a much more elusive model property, and no econometric model can claim invariance to every conceivable regime shift or intervention. However, invariance against the interventions that occurred within the sample period *is* a testable property, and this section performs tests of invariance on equation (14).

Evidence on constancy of the structural import price equation over time is presented in Figures 3 and 4, which are based on recursive IV-estimation. Figure 3 shows the one-step residuals with ± 2 estimated equation standard errors. The standard error varies little and there are no obvious outliers. Figure 4 exhibits the recursive estimates (denoted $\beta(t)$) with twice their standard errors ($\pm 2SE(t)$) for four central parameters, namely those on $\Delta_3 \Delta cpi_t$, $\Delta_2 e_t$, u_{t-1} and the error correction term. All of the estimates are reasonably stable and significant (or nearly so) by 1978.

Turning to constancy across regimes and interventions, section 5 brought out the issue of potential instabilities in the import price equation arising from discrete

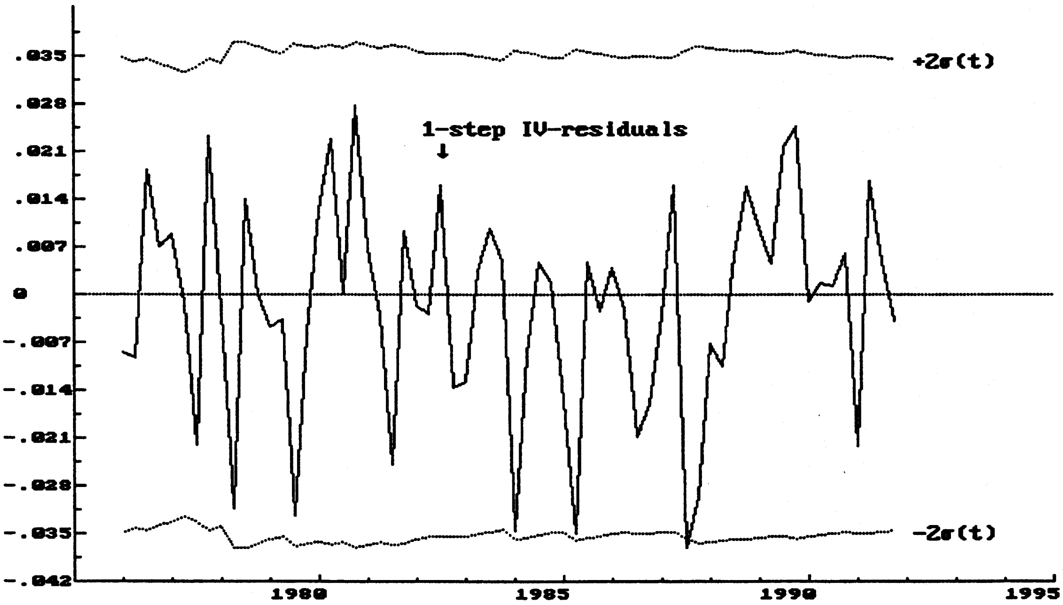

Figure 3: 1-step IV residuals with ±2 estimated standard errors.

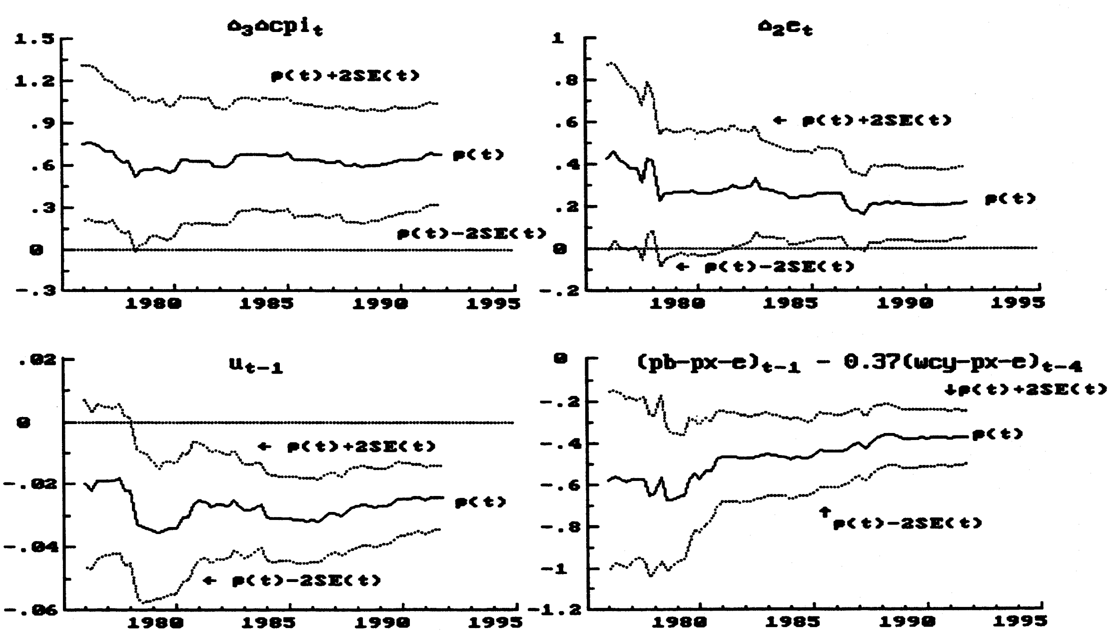

Figure 4: Recursive IV estimates with ±2 estimated standard errors.

exchange rate changes. In particular, the adjustment of import prices may be faster to discrete devaluations or revaluations than to fluctuations in the exchange rate around the central parity, owing to agents having different expectations about the permanence of these two sorts of currency changes. Phrased differently, there is reason to suspect that the exchange rate is not super exogenous for the parameters in (14), see Engle, Hendry, and Richard (1983), and that the Lucas critique applies. To test the null of super exogeneity, we follow the approach of Engle and Hendry (1993) and make use of the (marginal) model of Δe_t in the VAR from section 4. This exchange rate equation is stable over the sample as a result of the included seven dummy variables that capture the effects of discrete changes in the exchange rate. If the import price equation is invariant towards these interventions, the dummies ought to be insignificant if added to (14). The joint Wald test of these variables having zero coefficients in the import price equation yielded $\chi^2(7) = 4.32$, which is far from significant as the p-value is 0.74. To test the more specific hypothesis of an invariant short-run pass through, we added the intervention-variables interactively with $\Delta_2 e_t$ in (14). The Wald statistic now became $\chi^2(7) = 8.50$, with a p-value of 0.29. Thus we do not find significant evidence against the hypothesis that the import price equation (14) is invariant to whether exchange rate changes are a result of discrete devaluations/revaluations or of fluctuations within the exchange rate band.

We also tested the super exogeneity of the other contemporaneous conditioning variable, Δpx_t, using the two "oil price" dummies that proved necessary for deriving a stable equation for Δpx_t in the VAR. Individually and jointly the dummies were insignificant, the joint Wald test giving $\chi^2(2) = 3.61$ [0.16], where the p-value is in square brackets.

In conclusion, we are unable to reject that the exchange rate and foreign prices are super exogenous for the parameters in (14), with respect to the specified interventions captured by the relevant dummies in the VAR. This finding implies that the Lucas critique is refuted: The model remains constant in the face of policy interventions and structural breaks in the marginal processes of e_t and px_t. Hence our results are in line with Ericsson and Irons (1995), who show that there is virtually no evidence demonstrating the empirical applicability of the Lucas critique in the literature up to 1990.

7 Conclusion

We find that a well specified import price equation on Norwegian data can be based on a cointegrating vector between import prices, foreign prices, the exchange rate and domestic unit labour costs. Normalized on import prices, the long-run elasticities were 0.63 (foreign prices and the exchange rate) and 0.37 (domestic costs). Deviations from this relationship were highly significant in a structural import price equation. Furthermore, the model contained a negative effect of the Norwegian unemployment rate and positive effects of growth in domestic absorption and inflation—all of which were strongly significant and numerically important. The thrust of this evidence is that the data rejects the conventional model of import price determination for small open economies. Instead, the formation of Norwegian import prices seems to be well represented by the pricing to market hypothesis.

The paper also investigated the stability of the estimated import price equation, and we tested for invariance with respect to interventions in the exchange rate and foreign prices. The estimates were found to be reasonably stable, and the hypothesis of invariance to the specified interventions was not rejected. In particular, the parameters of the structural model were not significantly affected by the discrete currency changes that occurred within the sample.

The insight that import prices respond to changes in domestic costs and market conditions is important for modelling and forecasting of inflation. Moreover, the inflationary/deflationary effects of demand policies will be underestimated if the endogeneity of import prices is not taken into account. Conversely, policies that aim at reducing the growth in domestic costs (e.g. reductions of payroll taxes or wage controls) will depress price growth more than is usually assumed.

A Data definitions

- ABS = Domestic absorption in fixed 1990 market prices, million Norwegian kroner, and defined as $ABS = CP+CO+JP+JO$; CP = Private consumption expenditure, CO = Government consumption expenditure, JP = Private gross investments in fixed capital excluding oil and shipping, JO = Government gross investments in fixed capital. Source: Quarterly National Accounts (QNA).
- CPI = The official consumer price index. 1990 = 1. Source: Statistics Norway.
- $CPIDUM$ = Dummy for the effect on consumer prices of introduction and abandonment of price regulations. 1 in 1971.1, 1971.2, 1976.4, 1979.1. -1 in 1975.1, 1980.1, 1981.1, 1982.1. Zero otherwise.
- E = Import weighted exchange rate index. 1990 = 1. The construction of the index parallels that of PX below. Source: Central Bank of Norway databank of economic time series.
- $EE73.4, E76.4, EE77.3, EE78.1, E82.3, EE84.3$ and $EE86.2$ = Intervention dummies used to take account of outliers (which can be explained by discrete changes in the exchange rate) in the equation for Δe_t in the VAR. $Ejj.i$ equals one in $19jj.i$, zero otherwise. $EEjj.i$ equals one in $19jj.i$ and the following quarter, zero otherwise.
- $J12$ = 12 quarter moving average of gross investments in manufacturing and construction and service production. Source: QNA.

- PB = Implicit deflator for imports of manufactures with Norwegian substitutes. $1990 = 1$. Source: QNA.
- PX = Index of foreign export prices. $1990 = 1$. The index includes total merchandise exports of Norway's main trading partners; the 14 countries with largest weights in Norway's imports of goods. The weight of each country is the imports from this country as a share of the imports from the fourteen countries included in the index. The weights are calculated as the average shares over the period 1978-1987. Source: IMF.
- $PX74.1$ and $PX80.4$ = Impulse dummies used to take account of outliers in the equation for Δpx_t in the VAR. $PXjj.i$ equals one in $19jj.i$, zero otherwise.
- Qi = Centered dummy for quarter i, equals 0.75 in quarter i, -0.25 otherwise. $i = 1, 2, 3$.
- QBR = step dummy for structural break in the seasonal pattern of the series from the quarterly national accounts. Equals 1 until 1977.4, zero thereafter.
- U = Registered unemployment as a percentage of the "labour force", where the labour force is calculated as employed wage earners plus the number of registered unemployed. Source: Central Bank of Norway.
- VAT = Dummy for the introduction of VAT. Equals 1 in 1970.1, zero otherwise.
- WCY = Nominal manufacturing unit labour costs, defined as WC/Y; WC = Wage costs per man-hour in manufacturing; Y = Value added per man-hour in manufacturing at fixed 1990 factor costs. Source: QNA.
- $WCY78.1, WCY79.3, WCY86.1, WCY86.2$ and $WCY87.4$ = Impulse dummies used to take account of outliers in the equation for Δwcy_t in the VAR. $WCjj.i$ equals one in $19jj.i$, zero otherwise.

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