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by

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Empirical Modelling of Norwegian Import Prices*

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Abstract

In this paper we investigate the formation of Norwegian import prices of manufactures over the period 1970(1)–1998(3), thereby extending the sample period used in the study by Naug and Nymoen (1996). If international goods markets are perfectly integrated and the law of one price holds, then for a small open economy we would expect import prices to be exogenously given in foreign currency and to fully respond to movements in the exchange rate. However, empirical studies of small open economies have shown that exchange rate changes are not fully reflected in import prices, and that domestic variables have significant effects on import prices. Applying both single-equation and multivariate cointegration analysis we find evidence of a long-run cointegrating relationship between Norwegian import prices, foreign export prices measured in domestic currency, domestic unit labour costs, and the domestic unemployment rate. Our results indicate that exchange rate pass-through is complete in the long run. In contrast, Naug and Nymoen (1996) report a long-run pass-through coefficient of 0.63.

Keywords: Import prices; Exchange rate pass-through; Equilibrium-correction models

JEL classification: C51, E31, F31

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1 Introduction

Understanding the behaviour of import prices is important for a small open economy like Norway. First, growth in import prices is an important determinant of domestic inflation. Approximately 25 percent of the official Norwegian consumption basket consists of imported goods, and another 15 percent is made up of goods whose prices are influenced by the prices of imports through foreign competition or imported inputs. Second, by affecting the terms of trade, the development of import prices influences the trade balance.

Of particular importance is understanding the relationship between import prices and nominal exchange rates. The degree to which changes in exchange rates are reflected in local currency import prices is referred to as the degree of exchange rate pass-through. If pass-through is complete, a one percent depreciation of the domestic currency will cause import prices to increase by one percent. Incomplete pass-through of exchange rate changes to import prices has important implications for macroeconomic policy. First, incomplete pass-through implies that exchange rate policy are less effective in inducing the adjustments in terms of trade needed to equilibrate trade imbalances, and second, the effects of an exchange rate depreciation on consumer price inflation are moderated.

A common assumption in macro models of small open economies is that the "law of one price" holds, and that import prices are exogenously determined in foreign currency on the world market. In this situation, exchange rate pass-through will be complete, and domestic market conditions will not influence import prices. However, empirical studies of small open economies like Sweden, Finland, and Australia have shown that import prices do not fully respond to changes in exchange rates, and that domestic variables have significant effects on import prices.[1] Theoretically, these results could be explained in models of imperfect competition and segmented markets. The phenomenon of exchange rate induced price discrimination is referred to as "pricing to market" (Goldberg and Knetter (1996)).

Naug and Nymoen (1996) investigate the formation of Norwegian import prices of manufactures over the period 1970(1)–1991(4). Using multivariate cointegration analysis they find evidence of a long-run cointegrating relationship between import prices, the exchange rate, foreign export prices, and domestic unit labour costs. The estimated long-run elasticity of import prices with respect to the exchange rate and foreign export prices is 0.63, and the long-run elasticity with respect to domestic unit labour costs is estimated to 0.37. Naug and Nymoen (op cit) also estimate a dynamic structural import price equation in which deviations from the cointegrating relationship enter as an equilibrium-correction mechanism. The import price equation contains significant

[1] See Alexius (1997), Kuismanen (1995), and Menon (1995b).

effects of variables proxying demand pressure in the domestic economy, and so the small open economy assumption is rejected in favour of the pricing to market hypothesis. The purpose of this dissertation is to investigate the robustness of the results reported in Naug and Nymoen (op cit) when the estimation period is extended to 1998(3).

The dissertation is organised as follows: In Section 2 we briefly review the theoretical and empirical literature on exchange rates and traded goods prices discussing topics such as exchange rate pass-through, the law of one price, and pricing to market. We also derive an import price equation which will serve as a starting point for the empirical analysis.

Section 3 reports results from an econometric analysis of the relationship between Norwegian import prices of manufactures, the exchange rate, foreign export prices, domestic unit labour costs, and the unemployment rate over the period 1970(1)–1998(3). Section 3.1 describes the data and investigates the time series properties of the variables. The unit root tests suggest that the series are non-stationary and integrated of order one. In Section 3. 2 we derive a dynamic conditional single equation model of import prices. The starting point is a general unrestricted equilibrium-correction model estimated by OLS. To obtain a simpler model that is easier to interpret, but which represents the data equally well, we simplify the general model by deleting insignificant variables and imposing restrictions on the parameters. Evaluating the final model by analysis of residuals and tests for parameter stability, we find that the model is well-specified with parameters that are relatively constant over the sample period.

From the single equation analysis we find evidence of a cointegrating relationship between the variables in the model. This is consistent with the evidence in Naug and Nymoen (1996). However, the hypothesis of long-run unit homogeneity in foreign and domestic prices measured in the same currency is no longer accepted by the data, and the long-run elasticities of import prices with respect to the exchange rate and foreign export prices are close to one which is larger than those reported in Naug and Nymoen (op cit). The magnitude of the long-run elasticities is confirmed in the multivariate cointegration analysis in Section 3. 3. The hypothesis that the exchange rate, foreign export prices, and domestic unit labour costs are weakly exogenous for the cointegration parameters is not rejected by the data, implying that single equation estimation of the cointegration parameters is efficient.

The data series used in the empirical analysis are taken from Norges Bank's RIMINI model database and are available on request from the author. The numerical results are obtained using PcGive and PcFiml version 9.20.[2]

[2]See Hendry and Doornik (1999) and Doornik and Hendry (1997).

2 Theory and Existing Evidence

In the following subsections we give a brief overview of the theoretical literature on exchange rates and import prices focusing on the law of one price, exchange rate pass-through and pricing to market. In Section 2.1.3 we derive a theoretical import price equation which will serve as the starting point for the econometric analysis in Section 3.

2.1 The law of one price and exchange rate pass-through

As pointed out by Goldberg and Knetter (1996, p.3), when discussing the relationship between exchange rates and goods prices it is important to distinguish between integrated and segmented markets. An integrated market is defined as a market in which geography does not have a systematic effect on the prices of identical goods. In a perfectly integrated market identical products sells for the same price everywhere, that is, the absolute version of the law of one price (LOP) holds. For any product i, the LOP in its absolute form states

$$P_i = EP_i^* \tag{2.1}$$

where P_i is the price of good i in domestic currency, P_i^* is the analogous foreign currency price, and E is the nominal exchange rate. The mechanism assumed to be enforcing the law of one price is arbitrage. If the LOP holds for all products between two countries, and assuming that the weights used in constructing each country's price level are the same, then absolute purchasing power parity (PPP) holds between these two countries.

For reasons such as transport costs, imperfect information, and government-imposed trade barriers, the absolute version of the LOP is unlikely to hold in practise. However, if the factors causing deviations from (2.1) give rise to a constant price differential between the two markets, a weaker version of the LOP, relative LOP, holds:

$$P_i = \alpha EP_i^* \tag{2.2}$$

where α is a constant. Letting lower-case letters denote natural logarithms and taking first differences we get

$$\Delta p_i = \Delta e + \Delta p_i^* \tag{2.3}$$

The relative version of the LOP thus examines the proportionate changes in the variables in (2.1). According to Goldberg and Knetter (1996, p.7), there are three main reasons why the empirical literature has focused on the relative version of the LOP. First, transport costs, tariffs and other trade barriers make arbitrage costly, implying that a complete equalisation of prices is unlikely. Second, the identical goods assumption is strong and likely to be violated in most datasets. Finally, information on P_i and P_i^*

usually come in the form of price indices relative to a base year, making the *levels* of P_i and P_i^* arbitrary.

The conventional definition of exchange rate pass-through (ERPT) is the percentage change in local currency import prices caused by a one percent change in the nominal exchange rate. If the exporting firm adjusts the home currency price to fully offset the exchange rate change, then pass-through will be zero. On the other hand, if the exporter leaves the home currency price unadjusted, then the exchange rate change will be fully reflected in the import price and pass-through is complete.

In his examination of the relationship between the law of one price and exchange rate pass-through, Menon (1995c, p.551) asks the question of whether incomplete pass-through always implies violation of the law of one price. To see that it does not, we consider a simple supply and demand model.[3] Demand and supply for an imported good is given by

$$Q_D = D(P) \tag{2.4}$$

$$Q_S = S(P^*) = S\left(\frac{P}{E}\right) \tag{2.5}$$

where Q_D and Q_S denote the quantity demanded and supplied of the imported good, P and P^* are the domestic and foreign currency prices of the good, and E is the exchange rate. Total differentiation of (2.4) and (2.5) yields

$$dQ_D = \frac{\partial D}{\partial P} dP \tag{2.6}$$

$$dQ_S = \frac{\partial S}{\partial P^*}\left\{\frac{1}{E}dP - \frac{P}{E^2}dE\right\} \tag{2.7}$$

Imposing the condition that dQ_D and dQ_S are equal in equilibrium we get

$$\text{ERPT} = \frac{dP}{dE}\frac{E}{P} = \frac{\frac{\partial S}{\partial P^*}}{\frac{\partial S}{\partial P^*} - E\frac{\partial D}{\partial P}} = \left(1 - \frac{\varepsilon_D}{\varepsilon_S}\right)^{-1} \tag{2.8}$$

where $\varepsilon_D = \frac{\partial D}{\partial P}\frac{P}{D}$ and $\varepsilon_S = \frac{\partial S}{\partial P^*}\frac{P^*}{S} = \frac{\partial S}{\partial P^*}\frac{P}{ES}$ are the elasticities of demand and supply.

A small open economy is assumed to be a price taker in world markets and hence, to face a perfectly elastic supply of exports (corresponding to $\varepsilon_S \to \infty$). From (2.8) it follows that for a small open economy, the law of one price implies that pass-through will be complete. Changes in the exchange rates of large economies, however, could alter world prices, and ensure the co-existence of incomplete pass-through and the law of one price. Thus, for large countries the simple supply and demand model is sufficient to explain a failure of import prices to move in proportion to changes in the exchange rate.

[3] The model is based on Menon (1995a).

There is also the question of the relationship between the degree of segmentation and the degree of competition in a market. In a perfectly competitive market price is equal to marginal cost and so a perfectly competitive market must be integrated. However, an integrated market is not necessarily perfectly competitive. In an imperfectly competitve market producers may charge a price above marginal cost, but if markets are integrated, arbitrage could still eliminate differences in the common currency price of goods across markets. Moreover, if the producer has constant marginal costs of production and charges a constant markup over cost, exchange rate pass-through will be complete. Thus, we conclude that the relationship between the law of one price, the degree of exchange rate pass-through, and the nature and degree of competition is ambiguous.

As noted by Rogoff (1996, p.652) the empirical support for the law of one price is weak. The LOP has been rejected in a large number of studies covering a broad range of product categories and countries. Moreover, the deviations from the law of one price appear to be highly correlated with exchange rate movements. Engel and Rogers (1995) find that the price differentials for consumer goods across cities in the United States and Canada are much larger and more volatile than the price differentials for the same goods across cities within the same country even when controlling for the distance between the cities.

2.2 Incomplete pass-through and pricing to market

The theory presented above suggests that when identical goods are traded in an integrated world market, arbitrage should eliminate differences in the common currency prices of goods across countries. Exchange rate pass-through is complete and domestic market conditions are of no importance in the determination of import prices in small open economies. In this subsection we turn to the case where markets are segmented. Following Goldberg and Knetter (1996, p.3) we say that a product market is geographically segmented if the location of the buyers and sellers has a significant influence on prices. Market segmentation may be due to transportation costs, trade barriers or imperfect information. Now, if markets are imperfectly competitive as well as segmented, then profit maximisation could imply price discrimination. If so, market conditions in the importing country could affect import prices, and pass-through may be less than complete even in a small open economy.

A concept that has received much attention in recent literature on exchange rates and traded goods prices is the concept of *pricing to market* (PTM). According to Krugman (1987, p.50) what is meant by PTM is that import prices respond “too little” to exchange rate appreciations or depreciations. However, as stressed by Krugman, PTM is not present whenever import prices fail to respond in proportion to the exchange rate change. Any effect of the exchange rate on *world* prices of the imported good should be excluded from a measure of PTM.

When markets are characterised by imperfect competition producers can charge a markup over marginal costs. How this markup varies in response to exchange rate changes will depend on factors such as the degree of market segmentation, the degree of product differentiation, the functional form of the demand curve, and the exact form of market organisation.[4] Hence, a wide range of responses to exchange rate changes is possible. The models considered below fall into two main categories, namely static models and dynamic models. In the static models the actual or expected duration of an exchange rate change does not affect the pricing decision of the producer, whereas in dynamic models the distinction between temporary and permanent exchange rate changes is crucial. The dynamic models typically predict that pass-through is incomplete in the short-run, but that prices respond fully to exchange rate changes in the long-run.

Following Krugman (1987, p.59) we first consider the case of monopolistic price discrimination. Suppose that a foreign monopolist can sell its product either in the foreign or in the domestic market. Arbitrage is assumed to be prohibitively costly. The optimal price is

$$P^* = \frac{\varepsilon^* C^*}{\varepsilon^* - 1} \qquad \text{and} \qquad \frac{P}{E} = \frac{\varepsilon C^*}{\varepsilon - 1} \tag{2.9}$$

where C^* is the constant marginal cost in foreign currency, and ε and ε^* are the elasticities of market demand in the domestic and the foreign markets respectively.[5] The question is now whether a change in E will produce a more or less than proportional change in P. From (2.9) we see that this depends on the shape of the demand curve. If the demand curve has constant elasticity, pass-through will be complete. In order to get incomplete pass-through the elasticity of demand must fall as the price of the good falls. Depending on the shape of the demand curve, then, price-discriminating monopoly can explain pricing to market. The model also serves to illustrate that violations of the law of one price do not necessarily imply incomplete pass-through.

As an example of how different market structures may affect the degree of pass-through, we turn to the case of n^* identical foreign firms and n identical domestic firms engaging in Cournot competition in the domestic market.[6] Domestic and foreign markets are again assumed to be segmented, and the domestic and the imported good are perfect substitutes. The profits of the domestic and foreign firms are

$$\pi = Px - Cx \qquad \text{and} \qquad \pi^* = Px^* - EC^*x^* \tag{2.10}$$

respectively, where x and x^* are the quantities produced, and C and C^* are the constant

[4]See Dornbusch (1987).

[5]The elasticity of market demand, D, is defined as $\varepsilon = -\frac{\partial D}{\partial P}\frac{P}{D}$.

[6]The exposition follows Menon (1996).

marginal costs in domestic and foreign currency. The inverse demand function is $P(X)$, where $X = nx + n^*x^*$. Each firm maximises profit taking the outputs of the other firms as given. The first-order conditions for each of the domestic and foreign firms are

$$P\left(1 - \frac{x}{\varepsilon X}\right) = C \quad \text{and} \quad P\left(1 - \frac{x^*}{\varepsilon X}\right) = EC^* \tag{2.11}$$

where ε is the constant elasticity of market demand. In equilibrium, the market shares of each firm must be such that these pricing rules are consistent. Adding up the first-order conditions for the n domestic and the n^* foreign firms we find the equilibrium market price:

$$P = \frac{\varepsilon(nC + n^*EC^*)}{\varepsilon(n + n^*) - 1} \tag{2.12}$$

The market price thus depends on the sum of marginal costs of all the firms, the elasticity of demand, and the total number of firms in the market.

Differentiating (2.12) with respect to E and P we get

$$ERPT = \frac{dP}{dE}\frac{E}{P} = \frac{n^*}{n^* + n} = \frac{1}{1 + \frac{n}{n^*}} \tag{2.13}$$

where for simplicity we have assumed that domestic and foreign marginal costs are equal in the same currency ($C = EC^*$). From (2.13) we see that the degree of pass-through is decreasing in the ratio of domestic to foreign firms. An equal number of domestic and foreign firms results in a pass-through elasticity of 0.5. In the limiting case where $n^* \to \infty$, pass-through is complete. Menon (1996, p.437) shows that pass-through is increasing in the total number of firms in the market, $N = n + n^*$, approaching 1 as N approaches infinity.

In the static models described above, pass-through does not depend on the perceived duration of the exchange rate change. This distinction between temporary and permanent exchange rate changes is crucial in the dynamic models we turn to next. Krugman (1987) considers a price-discriminating monopolist which faces costs of changing supply to the foreign market.[7] In order to expand sales the monopolist must expand the marketing and distribution "infrastructure". This expansion is costly, and more costly the more rapid the expansion. Absent the adjustment costs, it would be optimal for the monopolist to reduce prices following an exchange rate appreciation. However, if there is no capacity to meet the expanded demand, then even if the appreciation is perceived to be permanent, it may be optimal not to reduce prices immediately. Instead prices could fall gradually as the sales infrastructure is expanded. Eventually, pass-through would be complete. If the appreciation is perceived to be temporary, the initial expansion in the

[7]See Krugman (1987) for a formal exposition.

infrastructure is likely be smaller implying an even lower degree of pass-through. Thus, the model predicts that the degree of pass-through will depend both on how recently the exchange rate has changed and how persistent the change is perceived to be. Moreover, the model provides an explanation of why there could be asymmetries in the response to currency depreciations and appreciations.

A second explanation for PTM emphasising dynamic supply-side effects is given in Baldwin (1988).[8] Baldwin assumes that firms incur irreversible costs when entering a market. The costs could represent the cost of setting up a sales organisation or of establishing a brandname through advertising. The model suggests that there will be a band within which the exchange rate could fluctuate without inducing entry or exit in the market. The band is greater the higher the costs of entry. The idea is that when firms are faced with irreversible costs of entry and volatile exchange rates, they will adopt a "wait and see" strategy and will be reluctant to enter a market following a temporary or a "small" exchange rate change. The result is a low degree of pass-through when the exchange rate fluctuates within the band. However, if the exchange rate moves outside the band and entries and exits take place, the degree of pass-through could be close to one. The model thus suggests that large exchange rate movements could produce a structural break in estimated pass-through relationships.

Froot and Klemperer (1989) focus on dynamic demand-side effects. The authors study a two-period duopoly where firms' second period demands depend on their current market shares. This intertemporal dependence may arise because consumers face substantial costs of switching between brands, or because there are so-called positive network externalities i.e. the product becomes more valuable to a consumer if more consumers have purchased it previously. Froot and Klemperer (op cit) show that in this model both the magnitude and the *sign* of pass-through will depend on the perceived persistence of exchange rate movements.

In response to a temporary appreciation of the importing country's currency exporting firms could either increase or decrease prices. The ambiguity arises because a temporary appreciation increases the value of current, relative to future, profits expressed in the exporter's currency. During the appreciation the value of the importing country's currency is temporarily high, so that rather than investing in market share by setting a low price exporting firms will instead increase their profit margins in the current period. If the appreciation is perceived to be permanent there will not be such incentives to shift profits from the future to the current period. Thus, the fall in the destination price of imports after a temporary appreciation will be smaller than after a permanent appreciation, that is, the degree of exchange rate pass-through is higher when the exchange rate change is expected to be permanent.

[8]See Baldwin (1988) for a formal derivation. The exposition here is based on Menon (1995a).

The dynamic models help explain both why the pass-through of an exchange rate may be low and why there might be long lags in the transmission of exchange rate changes to destination currency prices. The general suggestion is that incomplete pass-through is more likely to be a short-run than a long-run phenomenon, and that import prices will respond more to exchange rate changes that are perceived to be permanent than changes that are expected to be temporary. Menon (1995a, 1996) adds to these explanations by describing how the increasing presence of non-tariff barriers and multinational companies may reduce pass-through.

The hypothesis that the presence of non-tariff barriers affects the degree of exchange rate pass-through has become known as the "Bhagwati hypothesis".[9] Menon (1996, p.437) describes the process by which quantity restrictions influence pass-through in a small open economy as follows: In the presence of import restraints a small depreciation is likely to be absorbed into the quota rents extracted by the exporter rather than be reflected in import prices. If the depreciation is large enough to push import prices above the point where the restraints are no longer binding pass-through will be positive but less than complete. Menon (1995a, p.204) also suggests that the increasing presence of multinational corporations in international markets could limit the transmission of exchange rate changes to import prices. Practices such as the use of internal exchange rates on intra-firm transactions and flexibility in the choice of currency denomination of contracts and in the timing of settlements are likely to result in a weakened link between exchange rates and import prices in national markets.

Menon (1995a) surveys the empirical literature on exchange rate pass-through. The studies covered by the survey differ with respect to country-coverage, data and methodology, but typically report incomplete pass-through and long lags in the transmission of exchange rate changes to prices. The significant differences in pass-through estimates for a given country point to sensitivity of the results to the choice of data and econometric methodology. Furthermore, the general finding in studies using disaggregated, industry-level data is that pass-through varies across industries and product categories, and this raises the concern of possible aggregation bias in the pass-through estimates obtained on aggregate data. Finally, several studies examining the stability of the estimated pass-through coefficient over time find evidence of structural breaks in the pass-through relationship.

2.3 An econometric import price equation

In this section we derive the import price equation which will serve as the starting point for the empirical analysis in Section 3. The analytical framework is the markup model

[9] The reference is Bhagwati (1991).

employed in several previous studies including Hooper and Mann (1989), Menon (1995b) and Naug and Nymoen (1996). As manufactured goods are typically differentiated and traded in segmented markets, the markup model seems to be appropriate for the purpose of studying import prices of manufactures (Menon (1995b, p.298)).

Assume that a representative foreign producer sets the price of exports (PX_i) to a particular importing country i as a destination specific markup (λ_i) over its marginal costs of production (C^*):[10]

$$PX_i = \lambda_i C^*, \qquad i = 1, 2, \ldots, n \tag{2.14}$$

PX and C^* are both measured in the exporting country's currency, and n is the number of export markets. The import price in the currency of the importing country (PB_i) is obtained by multiplying through by the nominal bilateral exchange rate (E_i):

$$PB_i = E_i PX_i = E_i \lambda_i C^*, \qquad \forall\ i \tag{2.15}$$

The markup is assumed to respond to competitive pressures and demand pressures in the importing country. Abstracting from competition between foreign exporters in market i, Naug and Nymoen (op cit) specify the markup as

$$\lambda_i = K_i \left[PH_i/PB_i\right]^{\mu_i} DP_i^{\eta_i}, \qquad \forall\ i \tag{2.16}$$

where K_i is a constant, PH_i is the price of import competing goods, and DP_i is a measure of demand pressure in market i. We expect both μ_i and η_i to be positive, although strictly speaking, the sign of η_i is undetermined from theory. Substituting (2.16) into (2.15) and using lower-case letters to denote natural logarithms we find:

$$pb_i = \kappa_i + (1 - \phi_i)(c^* + e_i) + \phi_i ph_i + \nu_i dp_i, \qquad \forall\ i \tag{2.17}$$

where $\kappa_i = \frac{\kappa_i}{1+\mu_i}$, $\phi_i = \frac{\mu_i}{1+\mu_i}$, and $\nu_i = \frac{\eta_i}{1+\mu_i}$. The pass-through coefficient, defined as the partial elasticity of the import price with respect to the exchange rate, is $(1 - \phi_i)$. As long as $\phi_i > 0$, pass-through is incomplete and there is a pricing to market effect due to the presence of PH_i in the import price equation. In the limiting case where $\phi_i = 0$ (corresponds to $\mu_i = 0$) changes in the exchange rate (and foreign costs) are passed through completely and the price of competing goods has no effect on import prices.

[10]The exposition is based on Naug and Nymoen (1996).

Since C^* is not directly observable we employ a geometric average of the export prices to the n markets (PX) which relates to C^* by

$$PX = \prod_{j=1}^{n} PX_j^{\omega_j} = \prod_{j=1}^{n} (\lambda_j C^*)^{\omega_j}, \qquad 0 \leq \omega_j < 1, \qquad \sum_{j=1}^{n} \omega_j = 1 \tag{2.18}$$

where ω_j is the weight of market j. Taking logarithms and substituting (2.18) into (2.17) yields

$$pb_i = \kappa_i + (1-\phi_i)(px + e_i) + \phi_i ph_i + \nu_i dp_i - (1-\phi_i) \sum_{j=1}^{n} \omega_j \ln \lambda_j, \quad \forall\ i \tag{2.19}$$

In the following we consider time series data for *one* market. Introducing a time subscript t and adding a stochastic disturbance term u_t we get

$$pb_t = \ \kappa' + (1-\phi)(px_t + e_t) + \phi ph_t + \nu dp_t + u_t \tag{2.20}$$

From (2.19) we see that using the average export price as a proxy for marginal costs implies that the foreign exporter's (unobserved) markups in all n markets are contained in the disturbance term.

A limitation of the model as specified above is that it is static and hence, does not allow for import prices to adjust gradually to changes in the explanatory variables. Following Naug and Nymoen (op cit) we therefore interpret (2.20) as a long-run cointegrating relationship. Import price determination in the short run will be explained by a dynamic model where changes in import prices depend on deviations from the long-run relationship and on current and lagged changes in the explanatory variables. A second limitation of the model is that it is a partial equilibrium model. The coefficients in (2.20) are all interpretable as partial elasticities. To get an estimate of the full effect of an exchange rate change on import prices, we should also take into account the effects of an exchange rate change, as well as the causes of that change, on the other explanatory variables in the model.

The model (2.20) imposes the same rate of pass-through of exchange rates and foreign costs as well as unit homogeneity in $e+px$ and ph. In practice, however, these restrictions need not hold. In the short-run, exchange rates are more variable than costs, and a reasonable conjecture is that exporters will be more willing to absorb into their markups changes in exchange rates than changes in costs, which are likely to be permanent. Moreover, as pointed out by Athukorala and Menon (1995), apart from purely economic reasons, the coefficient restrictions may not hold because incompatibility of the price proxies which may result from differences in aggregation levels and methods of data collection. Therefore, we do not impose the cross-coefficient restrictions implied by

(2.20) apriori. Finally, assuming that the variables are integrated of order one, the long-run version of LOP holds if pb, e, and, px cointegrate with cointegration parameters equal to one. .

Adopting the analytical framework presented above Naug and Nymoen (op cit) find evidence of a single cointegrating relationship between Norwegian import prices of manufactures, foreign export prices, the exchange rate, and domestic unit labour costs over the period 1970(1)–1991(4). The estimated cointegrating vector is

$$pb = const. + \underset{(0.08)}{0.63} px + 0.63e + 0.37ulc \tag{2.21}$$

where ulc denotes unit labour costs in domestic manufacturing. The unit labour cost variable is included as a proxy for the price of import competing goods. The authors thus find support for the hypothesis that import prices are homogenous of degree one in foreign and domestic prices measured in the same currency. The presence of the domestic cost variable in the cointegrating vector is interpreted as evidence of a long-run pricing to market effect. The significant effect of domestic costs on Norwegian import prices is consistent with the findings reported by von der Fehr (1987).

An estimated long-run pass-through coefficient of 0.63 is close to the estimates reported in other studies of the pass-through to import prices in small open economies. In a study of Finnish import prices for total imports, Kuismanen (1995) finds a long-run pass-through coefficient of 0.68. In Alexius (1997) the long-run pass-through to Swedish import prices of manufactured goods is estimated to be in the range $0.6 - 0.8$, while Menon (1995b) estimates the long-run elasticities of Australian import prices of manufactures to 0.66 (the exchange rate), 0.75 (foreign costs) and 0.37 (the price of import competing goods).

Naug and Nymoen (op cit) also estimate a dynamic structural import price equation where deviations from (2.21) enter as an equilibrium-correction mechanism. The estimation period is 1970(1)–1991(4). The equation contains significant effects of variables proxying demand pressure in the domestic economy. The authors report a positive effect from domestic inflation and growth in domestic absorption, and a negative effect from the unemployment rate. Naug (1996) extends the sample to 1994(4) and finds that the estimated short-run coefficients remain relatively constant.

3 Empirical Modelling of Norwegian Import Prices of Manufactures 1970(1)-1998(3)

In this section we present results from an econometric analysis of Norwegian import prices of manufactures over the period 1970(1)-1998(3). To start, Section 3.1 describes the data and investigates the time series properties of the variables. Section 3.2 presents estimation results for a single equation equilibrium-correction model of import prices. Then, in Section 3.3 we compare the results from the single equation analysis with the results from multivariate cointegration analysis, and test for weak exogeneity of the regressors with respect to the cointegration parameters.

3.1 The data

Throughout the analysis we use quarterly, seasonally unadjusted data for the period 1970(1)–1998(3). Allowing for lags, estimation is over 1971(2)–1998(3) unless otherwise mentioned. Import prices, the exchange rate, and foreign export prices are indices taking the value 1 in 1996. All data series are taken from Norges Bank's RIMINI model database. The sources of the original data are given in Appendix A.

3.1.1 Variable descriptions

Taking equation (2.20) as the starting point for the empirical analysis we need data for import prices, exchange rates, foreign export prices, domestic prices of import competing products, and an indicator of domestic demand pressure. We consider these in turn.

Import prices The import price series (denoted PB) is an implicit deflator for imports of manufactures with Norwegian substitutes. The products in the index are priced cif Norwegian port, that is, the prices include cost, insurance and freight, but exclude import duty. The implicit deflator is calculated by dividing the value of imports by the volume of imports. Implicit deflators of this kind are subject to well-known limitations such as not accounting for shifts in the quality of a product, and reflecting not only underlying price changes, but also changes in the composition of imports. For example, the implicit deflator gives increasing weight to computers over the same period during which prices of computers have fallen sharply. Because it gives a low weight to computers and because it abstracts from shifts in the commodity composition of imports, Hooper and Mann (1989) prefer to use a fixed-weight index. However, Naug and Nymoen (1996) constructed an index excluding computers and a fixed-weight import price index for Norwegian manufactured imports over the period 1968–1991 and found that both showed a development more or less like the implicit deflator.

Figure 1: Import prices of manufactured products 1970(1)–1998(3)

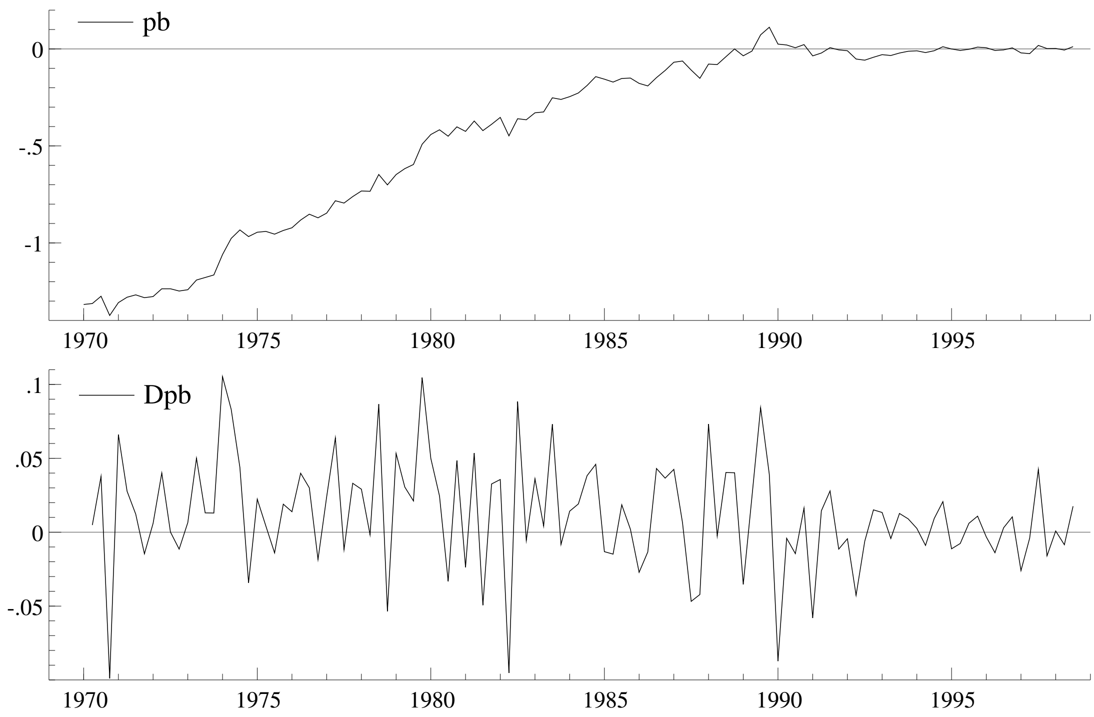


Figure 1 plots the log of the import price series (pb) and its quarterly growth rates (Δpb). In the period from 1970 to 1990 import prices increased steadily, while since 1990 growth in import prices has been markedly slower and the fluctuations from quarter to quarter appear to have been smaller. These are important features that our model should capture. Notice also the strong fluctuations in the index in 1989. Splitting the aggregate index into subindices according to the commodity classification in the quarterly national accounts, we find that these movements can be accounted for by fluctuations in the price of metals, which has an average quantity weight of about 9% in the aggregate index.

Exchange rates The exchange rate index (denoted E) is a nominal effective import weighted exchange rate for Norwegian Kroner (NOK). The weights reflect the relative importance of Norway's main trading partners. The 14 countries included in the index are the same as those that were included in Norway's official exchange rate basket in the period from August 1982 to October 1990. Each country's weight is the imports from this country as a share of total imports from the countries in the index. The weights are given in Table 1 and are calculated as the average import shares over the period 1978–87.

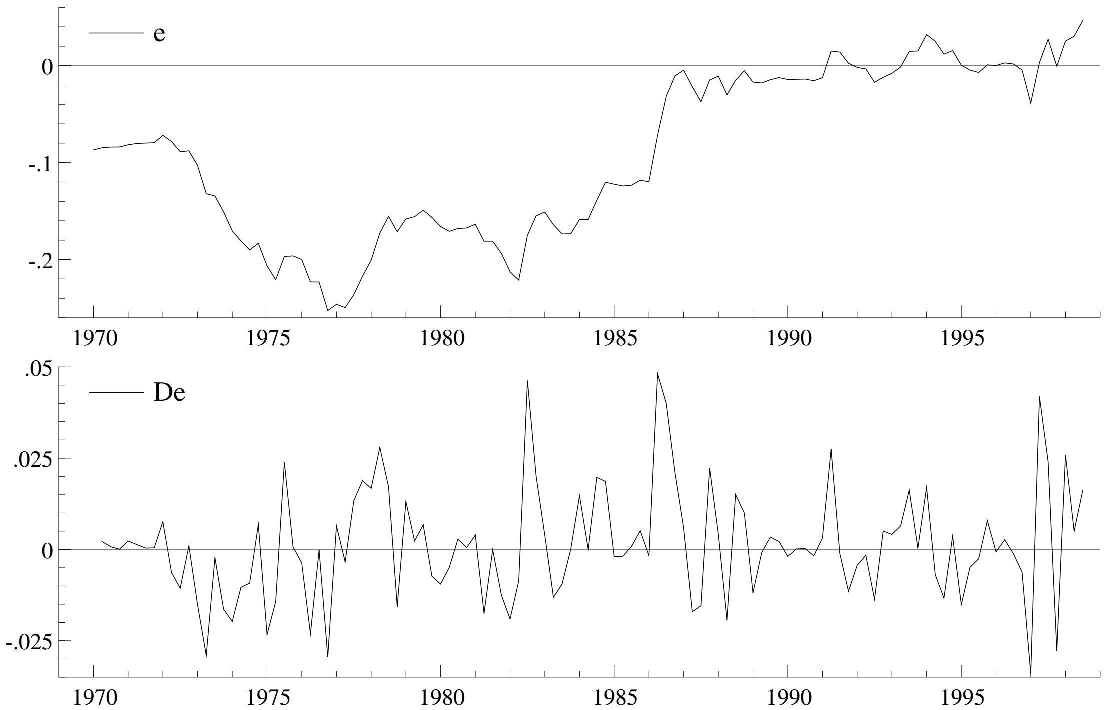
Note that using a trade-weighted exchange rate index is not necessarily optimal. As pointed out by Menon (1995a), to get a true representation of the extent of exchange rate fluctuations faced by exporters we should instead use a currency-contract-weighted exchange rate.

Table 1: Average import shares 1978–87. Total = 100.

Country	Import share
Sweden	20.6
Germany	17.5
Great Britain	13.4
USA	9.2
Denmark	7.7
Japan	6.3
Finland	4.9
France	4.4
Netherlands	4.1
Belgium	3.3
Italy	3.3
Canada	2.0
Switzerland	1.9
Austria	1.4

Figure 2 plots the log of the exchange rate (e) and the quarterly growth rates in the series (Δe). A rise in the index represents a depreciation of the NOK. The figure serves as a background for a brief overview exchange rate policy in Norway in the period 1970–1998. After the collapse of the Bretton Woods system in 1972, Norway joined the European snake arrangement leaving the NOK floating against the US dollar and other currencies outside the snake. After a 5 percent revaluation in November 1973, the NOK was devalued four times between 1976 and 1978 before Norway withdrew from the snake arrangement in 1978. In December 1978 a national currency basket with weights reflecting the relative importance Norway's trading partners was introduced. In the period 1979–1986 there were several small devaluations until in May 1986 the central value of the exchange rate was changed from 100 to 112. One reason for the adjustment was the sharp fall in the oil price early in 1986. The devalution marked the beginning of a period during which only small fluctuations around the fixed target value were allowed. In 1990 the weights in the currency basket were changed to ECU weights. In December 1992, after extensive speculation against the krone following the ERM crisis, the central bank was forced to let the NOK float. After a period of relative stability following the crisis, the exchange rate fluctuated widely in 1997 and 1998. The new guidelines for monetary policy were set out in the Exchange Rate Regulation of May 6, 1994. Monetary policy should be directed at maintaining a stable exchange rate against European currencies, but no fluctuation margins are specified.

Figure 2: Import weighted exchange rate 1970(1)–1998(3)

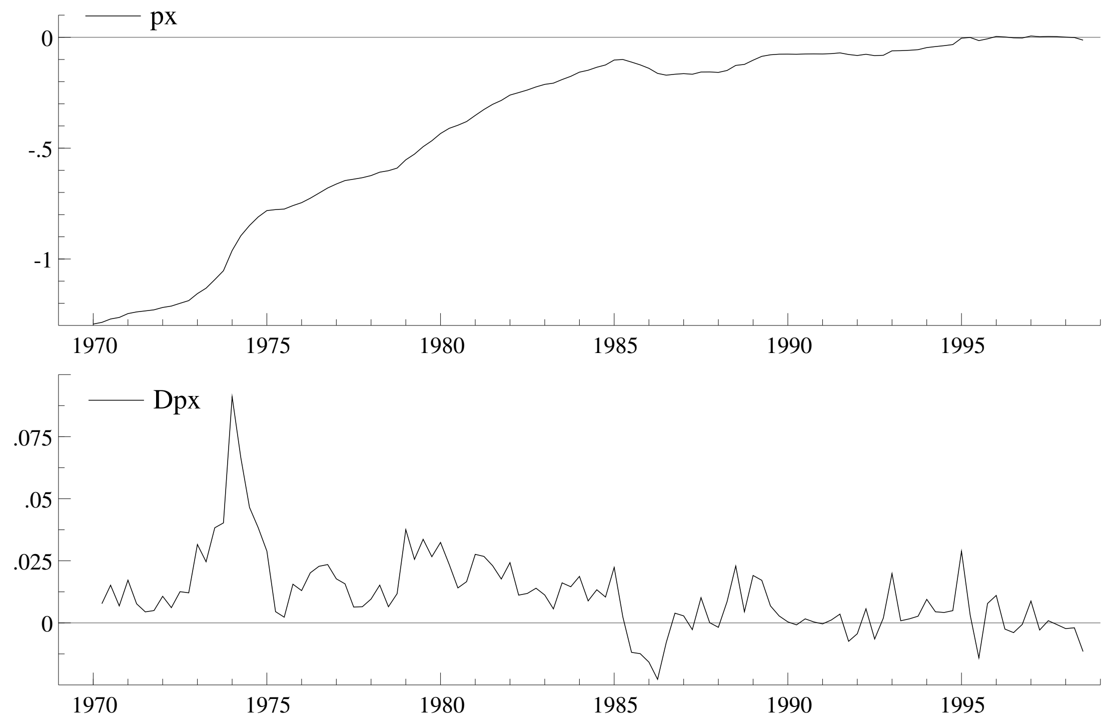


Foreign export prices Since foreign marginal costs are not directly observable, we follow Naug and Nymoen (1996) and use an import weighted foreign export price index (denoted PX) as a proxy variable. Previous studies have employed other proxies such as foreign producer price indices, foreign unit labour costs, and foreign consumer price indices. The import weighted foreign export price index is constructed with the same set of weights as the exchange rate index above. As pointed out in Section 2.3, using average foreign export prices as a proxy for marginal costs means that foreign exporters' markups in all markets are contained in the disturbance term in equation (2.20). Then for (2.20) to form a cointegrating relationship, we must assume that this measurement error is $I(0)$. Even if this assumption is satisfied, the measurement error may induce biases in the estimated cointegrating parameters in finite samples.

Figure 3 plots the log of foreign export prices (px) and the quarterly growth rates (Δpx). From the figure we see that export prices increased markedly following the oil price shocks in 1973–74 and 1979. In 1985–86 turbulent oil markets contributed to a fall in export prices. Average growth in foreign export prices appears to have been significantly lower after 1985 than in the period before.

Domestic unit labour costs As a proxy for the domestic price of import competing goods we use domestic unit labour costs in manufacturing and construction (denoted ULC). Unit labour costs are defined as $ULC = WC/Z$ where WC is hourly wage costs and Z is value-added labour productivity in manufacturing and construction. Again

Figure 3: Import weighted foreign export prices 1970(1)–1998



there is a potential measurement problem. The markups of domestic producers of import competing goods are left in the disturbance term, and only if the measurement error is $I(0)$ will (2.20) form a cointegrating relationship. Figure 4 plots the log of unit labour costs (ulc) and the quarterly growth rate of the series (Δulc). The series has a strong positive trend and exhibits a marked seasonal pattern which stems from seasonal variations in value added labour productivity.

Unemployment rate We use the unemployment rate (denoted U), measured as the number of registered unemployed as a fraction of the total labour force, as an indicator for demand pressure in the domestic economy. As noted by Naug and Nymoen (1996) the unemployment rate is easily observable and may therefore be used by foreign producers to assess demand conditions in Norway when detailed market information is costly. Figure 5 plots the log of the unemployment rate (u) and the quarterly changes in the series (Δu). We see that the level of unemployment was low throughout the 1970s, then started to increase in the beginning of the 1980s before reaching an all time high in 1992–93. The quarter to quarter fluctuations in the series are markedly smaller after 1983 than in the preceding period.

Figure 4: Unit labour costs 1970(1)-1998(3)

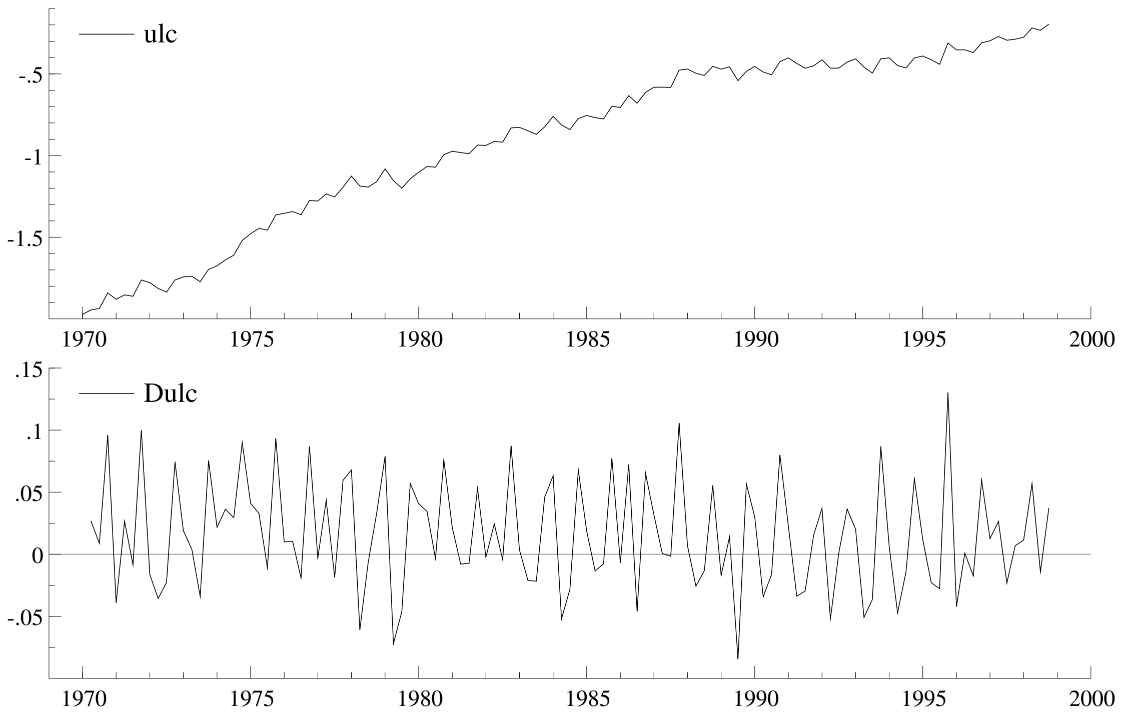
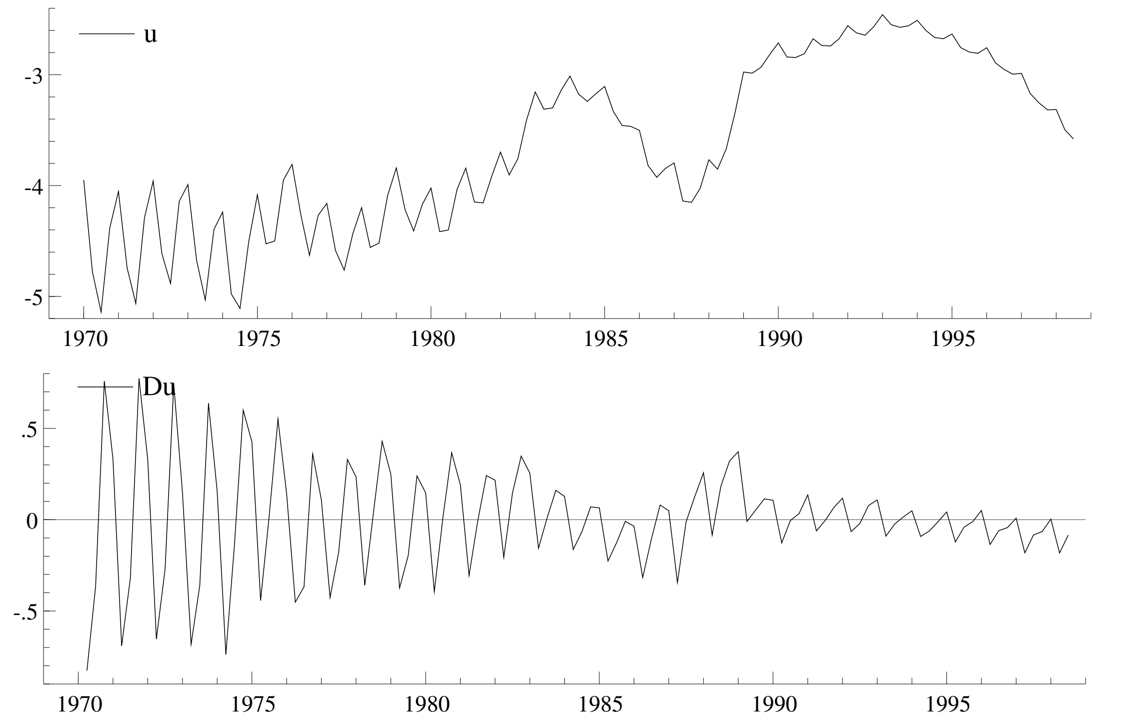


Figure 5: Unemployment rate 1970(1)–1998(3)



3.1.2 Unit root tests

Prior to modelling it is important to determine the order of integration for the variables of interest. In Table 2 we report ADF-tests for the levels and first-differences of the variables in the model.[11] As can been seen from the above figures import prices, foreign export prices, and unit labour costs all appear to be strongly trended. Thus, the appropriate alternative hypothesis seems to be that of trend stationarity, implying that the estimated model should include a deterministic trend. The test statistic for this specification is denoted τ_τ. For the unemployment rate and the exchange rate we also report the statistic τ_μ which is computed from a model estimated without a deterministic trend term. In each test we started out with 5 lags and then used a sequence of t-tests to determine the lag length. Misspesification tests were performed to ensure that the residuals were white noise. It proved difficult to obtain white noise residuals in the tests for u and px so the results from these tests should be interpreted with caution. In the table asterisks (*) and (**) denote rejection at the 5% and 1% critical values respectively. The Dickey-Fuller critical values are taken from Table B.6 in Hamilton (1994). The sample is 1971(4)–1998(3) for all series.

Table 2: Augmented Dickey Fuller tests 1971(4)-1998(3)

Variable	Lag-length	τ_u-statistic	τ_τ-statistic
pb	1	-	-0.460
px	5	-	-1.342
e	2	-0.424	-2.726
ulc	5	-	-1.818
u	5	-1.655	-1.915
Δpb	0	-11.490**	–
Δpx	5	-3.354*	–
Δe	2	-5.141**	–
Δulc	3	-17.347**	–
Δu	4	-3.535**	–

For the levels of the variables we fail to reject the null of a unit root in all cases, while for the first-differences the null is rejected at the 1% level for pb, e, ulc, and u and at the 5% level for px. These results suggest that all variables should be treated as non-stationary $I(1)$ series. The assumption that the unemployment rate is $I(1)$ may seem unreasonable, and previous studies such as Naug and Nymoen (1996) interpret the

[11]For a variable Y the ADF test statistic is the t ratio on ϕ_0 from the regression

$$\Delta Y_t = \phi_0 Y_{t-1} + \sum_{j=1}^{p} \phi_j \Delta Y_{t-j} + \xi + \omega t + u_t,$$

where p is the number of lags on ΔY , ξ is a constant term, t is a time trend and u_t is an error term which is assumed to be white noise. A unit root corresponds to $\phi_0 = 0$.

unemployment rate as an $I(0)$ series, but with possible structural breaks. Perron (1989) has shown that if the underlying process is stationary with a one time *structural break* in the trend or the constant term during the sample period, standard unit root tests cannot reject the null hypothesis of a unit root even asymptotically. Thus we cannot reject the interpretation of u as an $I(0)$ series with structural breaks solely on the basis of the tests above. However, Bjørnstad and Nymoen (1999) note that although the rate of unemployment is conceptually integrated of order zero with a bounded variance, the actual time series of the transformed rate of unemployment behaves as if it was $I(1)$ due to autocorrelation. In the subsequent analysis we follow Bjørnstad and Nymoen (1999) and treat u as an $I(1)$ variable.

3.2 A conditional single equation model

In this section we present estimation results for a single equation equilibrium-correction model of import prices. The formulation and interpretation of the general model is described in Section 3.2.1. The exposition is influenced by de Brouwer and Ericsson (1995). Estimation results are presented and interpreted in Section 3.2.2.

3.2.1 Formulation and interpretation of the general model

The starting point for single equation modelling is an autoregressive distributed lag model in pb, px, e, ulc, and u.

$$\begin{aligned} pb_t = a_0 &+ \sum_{i=1}^{5} a_{1i} pb_{t-i} + \sum_{i=0}^{5} a_{2i} e_{t-i} + \sum_{i=0}^{5} a_{3i} px_{t-i} \\ &+ \sum_{i=0}^{5} a_{4i} ulc_{t-i} + \sum_{i=0}^{5} a_{5i} u_{t-i} + \sum_{i=1}^{3} a_{6i} S_{it} + v_t \end{aligned} \tag{3.1}$$

where v_t is a disturbance term assumed to be white noise, and S_{it} is a seasonal dummy taking the value 1 in quarter i and zero otherwise. As is common with quarterly data, the lag length is set to 5. Without loss of generality (3.1) may be reparameterised as an EqCM

$$\begin{aligned} \Delta pb_t = a_0 &+ a_{20}\Delta e_t + a_{30}\Delta px_t + a_{40}\Delta ulc_t + a_{50}\Delta u_t \\ &+ c_1 pb_{t-1} + c_2 e_{t-1} + c_3 px_{t-1} + c_4 ulc_{t-1} + c_5 u_{t-1} \\ &+ \sum_{i=1}^{4} b_{1i}\Delta pb_{t-i} + \sum_{i=1}^{4} b_{2i}\Delta e_{t-i} + \sum_{i=1}^{4} b_{3i}\Delta px_{t-i} \\ &+ \sum_{i=1}^{4} b_{4i}\Delta ulc_{t-i} + \sum_{i=1}^{4} b_{5i}\Delta u_{t-i} + \sum_{i=1}^{3} a_{6i} S_{it} + v_t \end{aligned} \tag{3.2}$$

where $c_1 = \sum_{i=1}^{5} a_{1i} - 1$, $c_j = \sum_{i=1}^{5} a_{ji}$, and $b_{ji} = -\sum_{k=i+1}^{5} a_{jk}$ for $j = 1,2,3,4,5$ and $i = 1,2,3,4$. The EqCM in (3.2) contains both a static levels model and a pure difference model of import prices as special cases. The non-stochastic static-state equilibrium can be found by setting the error term v_t and all growth rates to zero. Ignoring the seasonals and dropping time subscripts (3.2) can be solved for

$$pb = -\left(\frac{a_0}{c_1}\right) - \left(\frac{c_2}{c_1}\right) e - \left(\frac{c_3}{c_1}\right) px - \left(\frac{c_4}{c_1}\right) ulc - \left(\frac{c_5}{c_1}\right) u \tag{3.3}$$

The coefficient c_1 measures the feedback from disequilibrium in period $(t-1)$. This can be seen more clearly if we rewrite (3.2) so as to incorporate the long-run solution (3.3) directly

$$\begin{aligned}\Delta pb_t = {} & a_0 + a_{20}\Delta e_t + a_{30}\Delta px_t + a_{40}\Delta ulc_t + a_{50}\Delta u_t \\ & + c_1(pb - \gamma e - \delta px - \kappa ulc - \lambda u)_{t-1} \\ & + \sum_{i=1}^{4} b_{1i}\Delta pb_{t-i} + \sum_{i=1}^{4} b_{2i}\Delta e_{t-i} + \sum_{i=1}^{4} b_{3i}\Delta px_{t-i} \\ & + \sum_{i=1}^{4} b_{4i}\Delta ulc_{t-i} + \sum_{i=1}^{4} b_{5i}\Delta u_{t-i} + \sum_{i=1}^{3} a_{6i}S_{it} + v_t\end{aligned} \tag{3.4}$$

where $\gamma = -c_2/c_1$, $\delta = -c_3/c_1$, $\kappa = -c_4/c_1$, and $\lambda = -c_5/c_1$. For dynamic stability we require $c_1 < 0$ (subject to strong exogeneity of the regressors). Testing the null hypothesis H_0: $c_1 = 0$ is a way of testing for cointegration between the variables in the model. Under the null of no cointegration the test statistic $t = \widehat{c_1}/\sqrt{\widehat{var(c_1)}}$ has a nonstandard distribution for which appropriate critical values can be found in MacKinnon (1991). Kremers et al. (1992) have shown that this test will have higher power against the alternative of cointegration than the standard residual-based ADF test suggested by Engle and Granger (1987) unless certain common factor restrictions are satisfied.

An alternative is to move the levels terms in the EqCM to the longest lag:

$$\begin{aligned}\Delta pb_t = {} & a_0 + a_{20}\Delta px_t + a_{30}\Delta e_t + a_{40}\Delta ulc_t + a_{50}\Delta u_t \\ & + c_1 pb_{t-5} + c_2 px_{t-5} + c_3 e_{t-5} + c_4 ulc_{t-5} + c_5 u_{t-5} \\ & + \sum_{i=1}^{4} d_{1i}\Delta pb_{t-i} + \sum_{i=1}^{4} d_{2i}\Delta px_{t-i} + \sum_{i=1}^{4} d_{3i}\Delta e_{t-i} \\ & + \sum_{i=1}^{4} d_{4i}\Delta ulc_{t-i} + \sum_{i=1}^{4} d_{5i}\Delta u_{t-i} + \sum_{i=1}^{3} a_{6i}S_{it} + v_t\end{aligned} \tag{3.5}$$

with $d_{1i} = \sum_{k=1}^{i} a_{1i} - 1$ and $d_{ji} = \sum_{k=1}^{i} a_{jk}$ for $i = 1,2,3,4$. The coefficients of the levels terms are unaffected by this reparameterisation. However, as stressed by Bårdsen

(1992), the interpretation of the short-term dynamics depends on the dating of the levels terms. If the levels terms enter at the longest lag as in (3.5), the dynamic coefficients are easier to interpret than if the levels terms enter at the first lag as in (3.2), in which case all the lagged short-run coefficients change sign. However, the advantage of having the levels terms at the first lag is that it makes testing the lag length of the dynamics straightforward. In view of this, the following approach to estimating an equilibrium-correction model has been suggested:[12]

1. Date the levels terms to period $(t-1)$ and test the significance of the lag lengths.
2. Delete non-significant lags.
3. Test and impose restrictions on the long-run coefficients.
4. Move the individual levels in the equilibrium-correction term to the longest significant lag.
5. Simplify the short term dynamics in order to obtain a parsimonious model.

3.2.2 Estimation results

Preliminary analysis using the full sample 1971(2)–1998(3) to estimate the coefficients in (3.2) by OLS revealed large outliers in 1982(2), 1989(3), and 1989(4). Moreover, the seasonal dummies all proved to be statistically insignificant. In the subsequent analysis we exclude the seasonals and include the dummy variable *Dum* (which takes the value –1 in 1982(2), 1 in 1989(3), 1 in 1989(4), and zero otherwise). OLS estimates of the general unrestricted model are reported in Table 3. Estimated standard errors are in parentheses. The test statistics give no evidence of serious residual misspecification. This is substantiated by the graphs of the scaled residuals $\widehat{v}_t/\widehat{\sigma}$ and the actual and fitted values of Δpb_t in Figure 6.

Recursive estimation provides a useful tool for investigating constancy. We therefore reestimate the model by recursive least squares, starting from an initial sample of $M = 42$ observations and then increasing the sample sequentially from 42 to $T = 110$. Figure 7 shows the recursively computed 1-step residuals with corresponding ± 2 residual standard errors, the 1-step Chow test for each t, breakpoint F-tests ($N\downarrow$-step Chow tests) where the value at t tests for constancy from t to T, and forecast F-tests ($N\uparrow$-step Chow tests) where the value at t tests for constancy from M to t. The Chow tests are all scaled by their 1% critical values from the F-distribution.[13] Apart from the significant 1-step Chow test in 1988(1), the graphs give no strong indications of nonconstancy. Hence we

[12]See Bårdsen and Fisher (1999) p. 496 and Bårdsen (1992) p. 376-378.
[13]See Hendry and Doornik (1999) for details.

Table 3: OLS estimates of general unrestricted model for Δpb 1971(2)-1998(3).

$$
\begin{array}{rllll}
\Delta pb_t = & -\underset{}{0.084} & -\underset{(0.108)}{0.393}\Delta pb_{t-1} & -\underset{(0.110)}{0.088}\Delta pb_{t-2} & -\underset{(0.105)}{0.106}\Delta pb_{t-3} \\
& -\underset{(0.086)}{0.074}\Delta pb_{t-4} & +\underset{(0.177)}{0.283}\Delta e_t & +\underset{(0.195)}{0.540}\Delta e_{t-1} & +\underset{(0.209)}{0.021}\Delta e_{t-2} \\
& +\underset{(0.203)}{0.094}\Delta e_{t-3} & -\underset{(0.209)}{0.061}\Delta e_{t-4} & +\underset{(0.252)}{1.001}\Delta px_t & +\underset{(0.319)}{0.417}\Delta px_{t-1} \\
& +\underset{(0.332)}{0.095}\Delta px_{t-2} & -\underset{(0.322)}{0.046}\Delta px_{t-3} & -\underset{(0.294)}{0.293}\Delta px_{t-4} & -\underset{(0.097)}{0.102}\Delta ulc_t \\
& -\underset{(0.100)}{0.269}\Delta ulc_{t-1} & -\underset{(0.099)}{0.310}\Delta ulc_{t-2} & -\underset{(0.098)}{0.084}\Delta ulc_{t-3} & +\underset{(0.088)}{0.031}\Delta ulc_{t-4} \\
& +\underset{(0.030)}{0.007}\Delta u_t & +\underset{(0.020)}{0.010}\Delta u_{t-1} & +\underset{(0.020)}{0.022}\Delta u_{t-2} & -\underset{(0.018)}{0.022}\Delta u_{t-3} \\
& -\underset{(0.029)}{0.043}\Delta u_{t-4} & -\underset{(0.081)}{0.264}pb_{t-1} & +\underset{(0.119)}{0.228}e_{t-1} & +\underset{(0.119)}{0.384}px_{t-1} \\
& -\underset{(0.050)}{0.081}ulc_{t-1} & -\underset{(0.110)}{0.024}u_{t-1} & +\underset{(0.016)}{0.102}Dum_t & \\
& & & & \\
& T = 110 & \widehat{\sigma} = 0.023 & R^2 = 0.694 & DW = 1.93
\end{array}
$$

Model diagnostics

Test	Observed value	p-value
AR $1-5$	$F(5,74) = 1.547$	0.186
$ARCH$ 4	$F(4,71) = 1.023$	0.399
$Normality$	$\chi^2(2) = 4.370$	0.113
$Heteroscedasticity$ X^2	$F(60,18) = 0.659$	0.884
$RESET$	$F(1,78) = 0.079$	0.779

conclude that our general model is relatively well-specified and so forms a valid basis for further simplifications.

Ignoring the dummy the derived static long-run solution of the general unrestricted model is

$$pb = -\underset{(0.149)}{0.319} + \underset{(0.212)}{1.455}px + \underset{(0.300)}{0.866}e - \underset{(0.181)}{0.308}ulc - \underset{(0.041)}{0.089}u \tag{3.6}$$

The standard errors of the long-run coefficients are calculated using the Bårdsen (1989) formula. The Wald test statistic for a test of the joint significance of all the variables (excluding the constant but including the dummy) in the long-run solution, $\chi^2(5) = 1356.2$, has a p-value of zero. Thus, except for unit labour costs, all variables appear to enter significantly and with expected signs in the equilibrium-correction term. Conditional on the variables cointegrating, the t-statistics associated with the long-run parameters will be asymptotically normal and can be used to conduct valid inference. From Table 3 we find that the estimated coefficient on pb_{t-1} has a t-value of -3.256. Comparing this with the appropriate critical values in MacKinnon (1991) we cannot reject the null hypothesis of no cointegration.[14] However, due to the large number of insignificant regressors in the model the power of the cointegration test may be low.

[14] The asymptotic 5% and 10% critical values for a model with 5 variables and a constant term are -4.419 and -4.1327 respectively. See Table 7.2 in Banerjee et al. (1993).

Figure 6: Graphic analysis of general unrestricted model 1971(2)–1998(3)

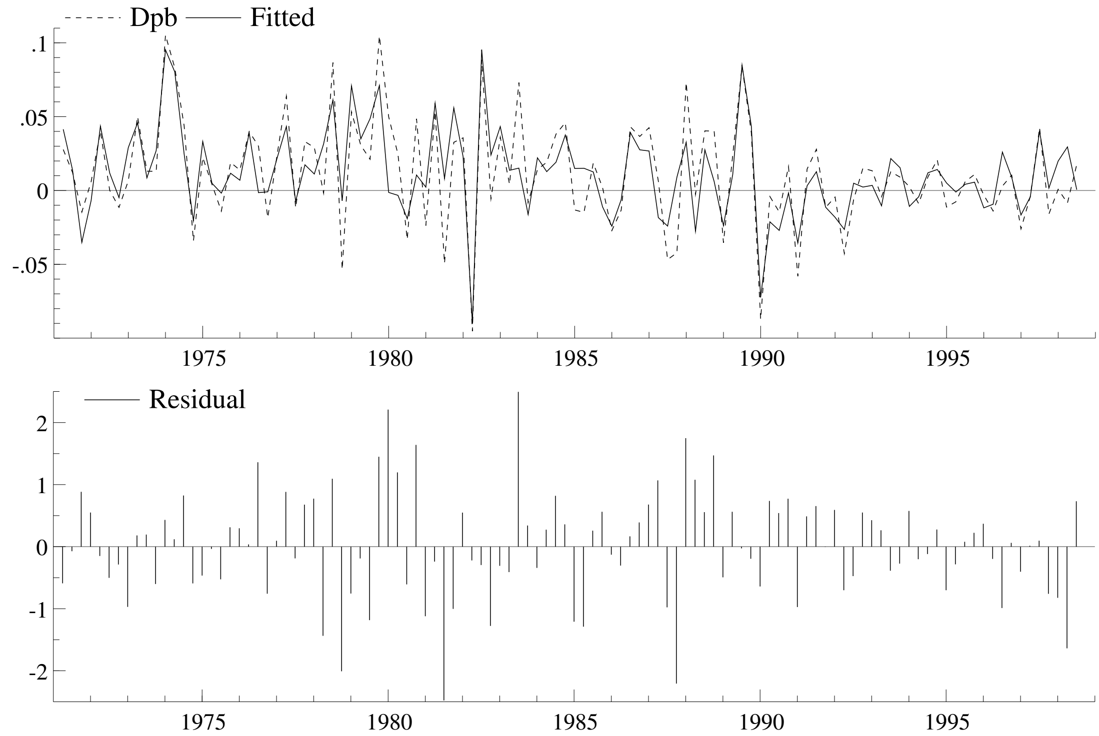
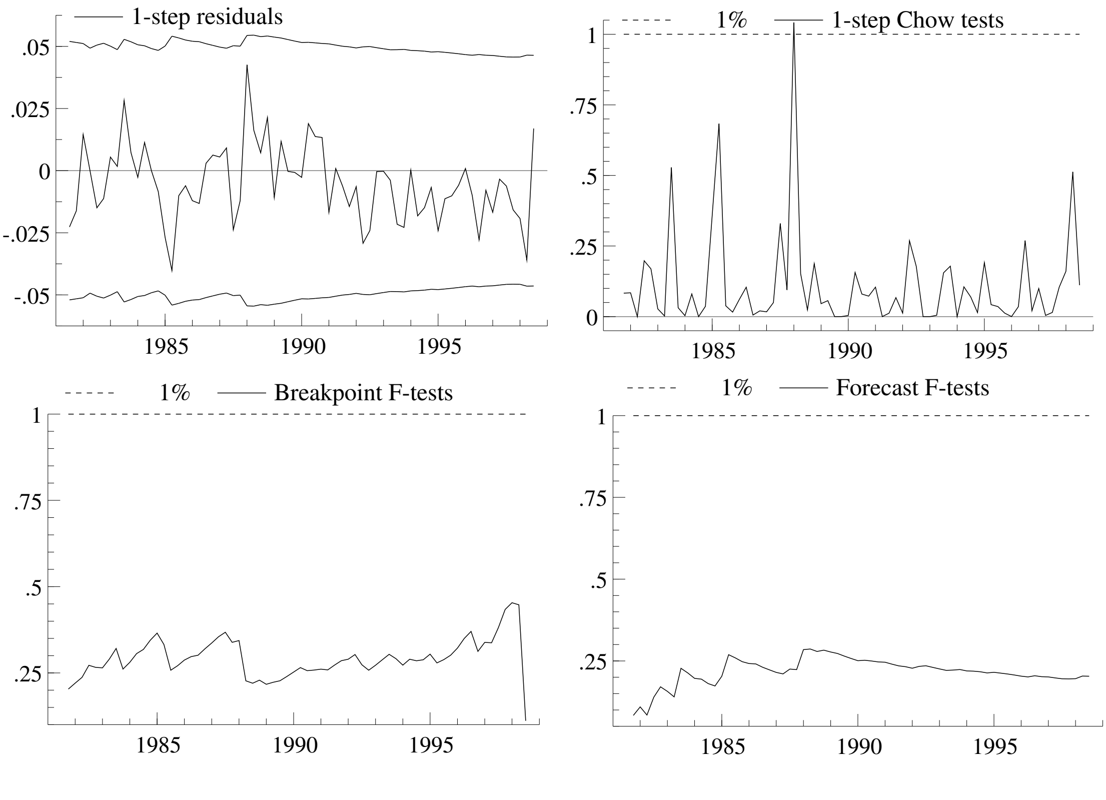


Figure 7: Recursive analysis of general unrestricted model 1971(2)–1998(3)



The next step is to test and restrict the lag length of the dynamics in the model. Removing insignificant lags will improve the power of the cointegration test and increase the precision of the long-run estimates. Testing the lag length of each variable we find that the following variables are statistically insignificant: The third and fourth lags on Δpb, Δpx, Δe, Δulc, Δu; the second lag on Δpb, Δpx, Δe, Δu; the first lag on Δpx. Table 4 reports the OLS estimates of the reduced equation.

Table 4: OLS estimates of reduced model for Δpb 1971(2)-1998(3).

$$
\begin{aligned}
\Delta pb_t = \; & \underset{(0.029)}{-0.085} & & \underset{(0.080)}{-0.312}\Delta pb_{t-1} & & \underset{(0.169)}{+0.333}\Delta e_t & & \underset{(0.172)}{+0.537}\Delta e_{t-1} \\
& \underset{(0.191)}{+1.190}\Delta px_t & & \underset{(0.069)}{-0.086}\Delta ulc_t & & \underset{(0.071)}{-0.248}\Delta ulc_{t-1} & & \underset{(0.069)}{-0.242}\Delta ulc_{t-2} \\
& \underset{(0.012)}{-0.058}\Delta u_t & & \underset{(0.011)}{+0.032}\Delta u_{t-1} & & \underset{(0.063)}{-0.294}pb_{t-1} & & \underset{(0.095)}{+0.292}e_{t-1} \\
& \underset{(0.089)}{+0.397}px_{t-1} & & \underset{(0.041)}{-0.063}ulc_{t-1} & & \underset{(0.008)}{-0.026}u_{t-1} & & \underset{(0.014)}{+0.091}Dum_t
\end{aligned}
$$

$T = 110 \qquad \widehat{\sigma} = 0.023 \qquad R^2 = 0.639 \qquad DW = 2.03$

Model diagnostics

Test	Observed value	p-value
AR $1-5$	$F(5,89) = 0.456$	0.808
$ARCH$ 4	$F(4,86) = 2.259$	0.069
$Normality$	$\chi^2(2) = 1.701$	0.427
$Heteroscedasticity$ X^2	$F(30,63) = 1.487$	0.093
$RESET$	$F(1,93) = 0.000$	0.992

The F-statistic for testing the overall validity of the reductions is $F(15,79) = 0.953$ (with a p-value of 0.511), thus the imposed set of zero-restrictions is accepted by data. Except for weak indications of heteroscedasticity, the model diagnostics give no evidence of residual misspecification. The estimated coefficient on pb_{t-1} now has a t-value of -4.7, and the null hypothesis of no cointegration is rejected at the 5% level. The unrestricted static long-run solution of the reduced model takes the form

$$pb = \underset{(0.107)}{-0.289} + \underset{(0.230)}{0.991}e + \underset{(0.158)}{1.347}px - \underset{(0.137)}{0.213}ulc - \underset{(0.029)}{0.088}u \tag{3.7}$$

This is essentially the same long-run solution as that estimated for the general unrestricted model (see 3.6). In particular, the estimated coefficient on the exchange rate is close to 1, supporting the hypothesis of complete exchange rate pass-through. However, there seems to be little support for the hypothesis that LOP holds in the long run. The long-run effect of unit labour costs on import prices is not statistically significant, but the unemployment rate enters statistically significantly in the long-run solution. Moreover, the long-run elasticity of foreign prices exceeds 1.

Table 5 reports F-statistics for tests of restrictions on the long-run coefficients. The

Table 5: Wald tests for linear restrictions on the long-run coefficients

Hypothesis	Statistic	p-value
H_1: $\kappa = \lambda = 0$	$F(2,94) = 5.877$**	0.004
H_2: $\gamma = \delta$	$F(1,94) = 6.329$*	0.014
H_3: $\gamma = \delta = 1$	$F(2,94) = 6.379$**	0.003
H_4: $\gamma = \delta$ and $\kappa = 1 - \gamma$	$F(2,94) = 6.830$**	0.002
H_5: $\gamma = \delta = 1$ and $\kappa = 0$	$F(3,94) = 4.904$**	0.003

restrictions are: Absence of domestic macrovariables in the cointegration vector (H_1), equality of the coefficients of the exchange rate and foreign export prices (H_2), unit long-run coefficients on the exchange rate and foreign export prices (H_3), long-run unit homogeneity in domestic costs and foreign prices expressed in domestic currency (H_4), and finally, long run unit homogeneity and unit coefficients on the exchange rate and foreign export prices (H_5). Of these, only one hypothesis is not rejected at the 1% level, namely the restriction that the coefficients of the nominal exchange rate and foreign export prices are equal. It is also evident from the above that the long-run version of LOP (corresponding to the hypothesis that $\gamma = \delta = 1$ and $\kappa = \lambda = 0$) is not supported by the data.

The results so far suggest that the long-run elasticities of the exchange rate and foreign export prices are larger in magnitude than those reported by Naug and Nymoen (1996). The results also differ from theirs in that the coefficient on domestic costs is negative, and that the hypothesis of long-run unit homogeneity is rejected by the data. The presence of the unemployment rate in the long-run solution is a feature that our model shares with that of Naug and Nymoen (1996).

As explained in Section 2.1 3, the lack of support for long-run unit homogeneity may result from a problem with incompatibility of the price indices employed. Despite the evidence, however, we require that the long-run solution in the final model satisfies the restriction of long-run homogeneity. Since domestic unit labour costs enter with the opposite sign from what is expected from theory, we implement long-run homogeneity by imposing H_5: $\gamma = \delta = 1$ and $\kappa = 0$. The estimated equilibrium-correction term is

$$EqCM_t = pb_{t-2} - e_{t-2} - px_{t-1} + 0.09u_{t-2}$$

where we have moved the individual levels terms to the longest significant lags.

Replacing the levels terms in the reduced model with $EqCM_t$, the next step is to eliminate insignificant regressors and impose data-acceptable restrictions on the dynamic short-run coefficients to obtain the final equation. OLS estimates of the final model are given in Table 6. Recursive coefficient estimates and tests of parameter constancy are provided in Figure 8 and 9 respectively. There is no evidence of residual misspecification and the estimated coefficients appear to be relatively stable over the sample period although there is some indication of instability around 1988(1) and 1990(1). We note that current changes in the exchange rate, Δe_t, is only significant when the whole sample period is used for estimation. This casts doubts as to whether this variable should be included in the final model. However, dropping Δe_t from the model induces residual misspecification.

Table 6: OLS estimates of final model for Δpb 1971(2)-1998(3).

$$\begin{aligned}\Delta pb_t = & \underset{(0.009)}{-0.037} \quad \underset{(0.079)}{-0.523}\Delta pb_{t-1} \quad \underset{(0.163)}{+0.364}\Delta e_t \quad \underset{(0.166)}{+0.858}\Delta e_{t-1} \\ & \underset{(0.186)}{+1.217}\Delta px_t \quad \underset{(0.010)}{-0.057}\Delta u_t \quad \underset{(0.049)}{-0.203}\Delta_2 ulc_{t-1} \quad \underset{(0.014)}{+0.084}Dum_t \\ & \underset{0.037}{-0.166}EqCM_t\end{aligned}$$

$$T = 110 \qquad \widehat{\sigma} = 0.024 \qquad R^2 = 0.578 \qquad DW = 1.92$$

Model diagnostics

Test	Observed value	p-value
AR $1-5$	$F(5,96) = 0.811$	0.545
$ARCH$ 4	$F(4,93) = 1.390$	0.244
$Normality$	$\chi^2(2) = 1.777$	0.411
$Heteroscedasticity$ X^2	$F(16,84) = 1.038$	0.427
$RESET$	$F(1,100) = 0.006$	0.939

From Table 6 we see that the short run response of import prices to changes in foreign export prices is 1.2, and this is larger than the imposed long-run effect. The impact elasticity of the exchange rate is smaller than the impact elasticity of foreign export prices and is also smaller than the corresponding long-run elasticity. Moreover, the lags in the transmission of exchange rate changes to import prices is longer than the lags in the transmission of foreign export prices. This could be explained by the fact that in the short-run, exchange rates are more variable than costs, and that exporters are more willing to absorb into their markups changes in exchange rates which are likely to be reversed than changes in costs, which are more likely to be sustained. The fact that the Norwegian exchange was fixed within a specified band during most of the sample period, lends support to this interpretation. As noted by Naug and Nymoen (1996) there is a problem with this interpretation which stems from the fact that the sample period contains a number of discrete exchange rate changes which were probably perceived as permanent. Moreover, in Section 3.1 we concluded that the exchange rate could be treated as integrated of order one, which implies that all changes in the exchange rate series are permanent.[15]

[15]See W. Branson's comments to Hooper and Mann (1989).

The import price equation contains a significant negative effect of current changes in the unemployment rate, Δu_t. Our results thus indicate that increases in domestic demand pressure result in price increases on imports of manufactures. This is consistent with the findings in Naug and Nymoen (op cit). In addition to a significant negative effect from the unemployment rate, their model contains positive effects from domestic inflation and growth in domestic absorption.

The model also contains a significant negative effect from lagged changes in unit labour costs, $\Delta_2 ulc_{t-1}$. Giving a clear interpretation of this result is difficult, and attempts to reparameterise the model such that the coefficients were easier to interpret were not successful. The speed of adjustment towards the long-run equilibrium path, given by the coefficient of the equilibrium-correction term, is -0.17 and implies that the correction of disequilibria from the estimated long-run relationship is fairly slow.

Figure 8: Recursive OLS estimates of final model 1971(2)–1998(3)

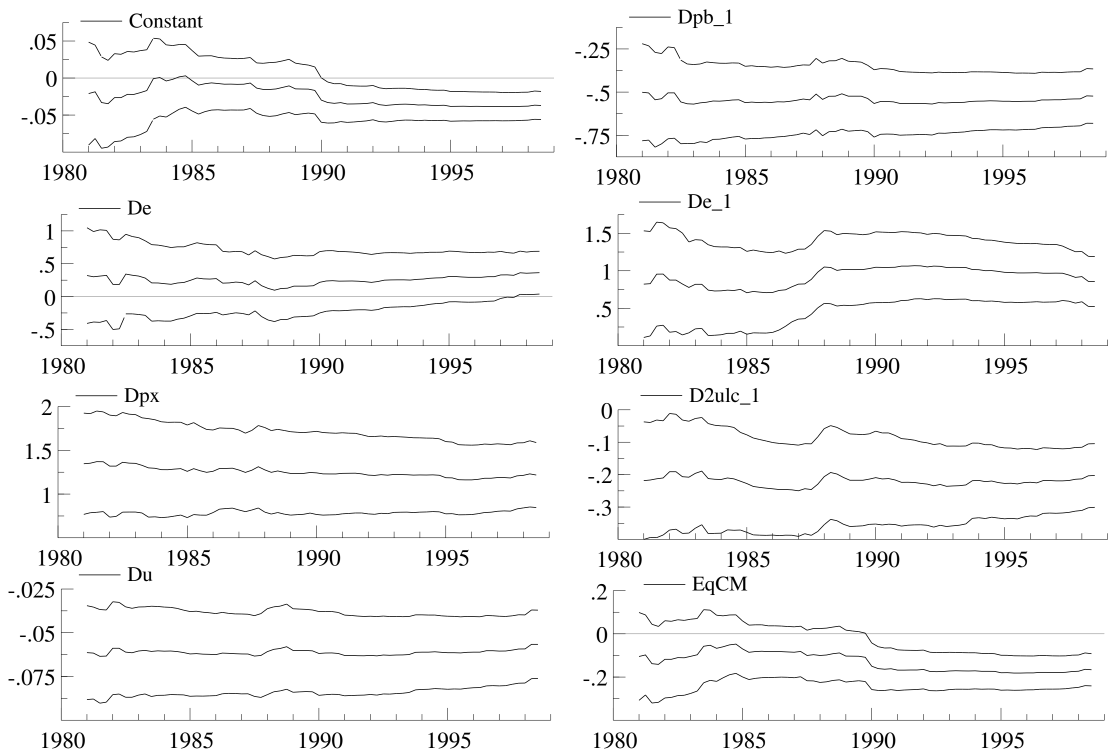
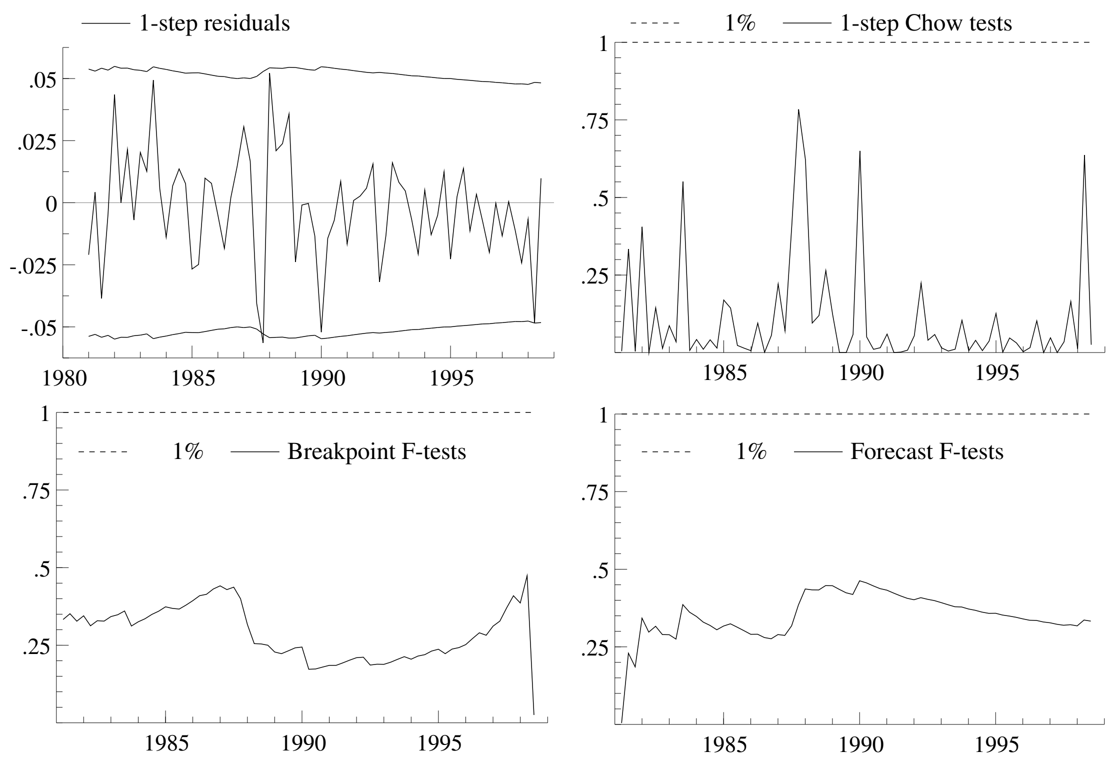


Figure 9: Recursive analysis of final model 1971(2)–1998(3)



3.3 Multivariate cointegration analysis

In this section the results from the single equation analysis are compared with the results from applying the Johansen (1988) full information maximum likelihood (FIML) procedure. For the single equation approach to yield efficient estimates of the long-run coefficients we require that the regressors are weakly exogenous for the cointegration parameters.[16] Moreover, with n variables in the model there may be up to $n-1$ distinct cointegration vectors, and when estimating a single equation we can only obtain an estimate of a linear combination of these. Taking a multivariate approach we are able to determine the number of cointegration vectors empirically as well as testing the assumptions of weak exogeneity implicit in the single equation analysis. If weak exogeneity is absent, we must choose between efficient (but more complicated) inference from the system analysis and inefficient inference from the conditional model.

Consider an n-dimensional vector equilibrium-correction model (VEqCM) of the type

$$\Delta \mathbf{Y}_t = \mathbf{\Pi}_0 \mathbf{Y}_{t-1} + \sum\nolimits_{j=1}^{p-1} \mathbf{\Pi}_j \Delta \mathbf{Y}_{t-j} + \mathbf{\Phi} \mathbf{D}_t + \boldsymbol{\varepsilon}_t, \qquad \boldsymbol{\varepsilon}_t \sim \mathbf{IN}(\mathbf{0}_n, \mathbf{\Sigma}) \tag{3.8}$$

where $\mathbf{Y}_t = (Y_{1t}, Y_{2t}, \ldots, Y_{nt})'$ is an $(n \times 1)$ vector of $I(1)$ variables, $\boldsymbol{\varepsilon}_t = (\varepsilon_{1t,} \varepsilon_{2t,} \ldots, \varepsilon_{nt})'$ is an $(n \times 1)$ vector of independently and normally distributed disturbances and $\mathbf{D}_t$ is a vector of deterministic variables. If the variables in $\mathbf{Y}_t$ are cointegrated $\mathbf{\Pi}_0$ can be factored into $\boldsymbol{\alpha}\boldsymbol{\beta}'$ where both $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are $(n \times r)$ matrices of rank r. Hence, cointegration implies that the matrix $\mathbf{\Pi}_0$ has reduced rank $r < n$. The columns of $\boldsymbol{\beta}$ contain the coefficients in the r cointegrating vectors such that the linear combinations $\boldsymbol{\beta}'\mathbf{Y}_t$ are $I(0)$. The matrix $\boldsymbol{\alpha}$ is a matrix of "loading coefficients" giving the weight attached to each cointegrating vector for all n equations.

Testing for cointegration thus amounts to determining the rank of $\mathbf{\Pi}_0$. The Johansen FIML procedure enables empirical determination of the cointegrating rank from (3.8). Application of Johansen's procedure provides n eigenvalues $\widehat{\lambda}_1 > \widehat{\lambda}_2 > \ldots > \widehat{\lambda}_n$. The estimates of $\boldsymbol{\beta}$ are obtained as the eigenvectors corresponding to the r largest eigenvalues.

[16] Consider two variables x_t and y_t with joint density $f(\cdot)$. The joint density can be factored into a conditional density for y_t given x_t and a marginal density for x_t as follows

$$f(y_t, x_t\ ; \theta) = \ g(y_t \mid x_t; \lambda_1) \times h(x_t\ ; \lambda_2)$$

The concept of weak exogeneity is defined relative to the *parameters of interest* ψ: x_t is weakly exogenous for ψ if

1. $\psi = \psi(\lambda_1)$; that is, ψ is a function of λ_1 alone
2. λ_1 and λ_2 are variation free

These conditions ensure that ψ neither directly (condition 1) nor indirectly (condition 2) depends on the parameters of the marginal model. Weak exogeneity is a sufficient condition for efficient inference on ψ from the conditional model. (The definition is taken from Ericsson et al. (1998))

A test of the null hypothesis that there are at most r cointegration vectors can be based on the *trace statistic*

$$\eta_r = -T \sum_{i=r+1}^{n} \ln(1 - \widehat{\lambda}_i), \qquad r = 0, 1, 2, \ldots, n-1 \tag{3.9}$$

where T is the number of observations. Under the hypothesis that there are r cointegrating relationships, the distribution of η_r is nonstandard. Asymptotic critical values are tabulated by Osterwald-Lenum (1992). The appropriate critical values depend on whether a trend and/or a constant are included in the model and whether these are restricted to lie in the cointegration space. Testing is sequential $\eta_0, \eta_1, \ldots, \eta_{n-1}$, and the cointegrating rank is selected as zero if η_0 is not significant and $r+1$ if the last significant statistic is η_r.

When the full system is large, we are often restricted to making inferences on the basis of a conditional model only. As shown in Harbo et al. (1998) making inference on cointegrating rank from a conditional model is not straightforward. For the asymptotic distribution of the test statistics to be free of nuisance parameters, the conditional model should include a restricted highest-order deterministic term, and asymptotic inference should be based on the critical values provided by Harbo et al. (1998). Hence, if there is the possibility of a linear but not a quadratic trend in the variables, inference on cointegrating rank should be made from a model which includes an unrestricted constant term and a restricted trend term. After having determined the cointegrating rank we can test whether the linear trend in the cointegrating relations can be dropped by a conventional χ^2-test.

3.3.1 Formulation and estimation of the VAR

The starting point for multivariate cointegration analysis is a congruent unrestricted vector autoregressive model. Initially, we estimate a VAR for pb, e, px, and ulc with five lags on each variable and three centered seasonal dummies.[17] An unrestricted constant term is included to allow for a linear trend in the levels of the variables. The rate of unemployment lagged two quarters, u_{t-2}, is assumed to be weakly exogenous for the cointegrating relations and is included in the long-run part of the system as a non-modelled variable. For inference purposes, following the analysis in Harbo et al. (1998), we also add a restricted linear trend. Finally, two conditioning variables are taken from the conditional single equation model above: Dum_t and Δu_t.

The results from estimating the fifth-order VAR by OLS over the full sample 1971(2)–1998(3) strongly indicate that the system is misspecified. The diagnostic tests reveal

[17] Centered seasonal dummies sum to zero over time and thus do not affect the asymptotic distributions of the tests for cointegrating rank. See Harris (1995).

significant non-normality and autocorrelation in the equation for px, and significant $ARCH$ effects in the equation for ulc. The vector normality test has a p-value of zero. Moreover, inspection of the residuals reveals a large outlier in the equation for px in 1974(1), and the recursively estimated 1-step residuals indicate nonconstancies in the equation for e in 1986 and 1997.

To mop up the outliers and induce constancy in the equations we use 5 impulse dummies: One to account for the increase in px following the oil crisis in 1973 ($PX74q1$), one to allow for the devaluation of the krone in May 1986 ($E86q2$) and three to account for the strong fluctuations in the exchange rate in 1997 ($E97q1$, $E97q2$, and $E97q4$).[18] The F-statistic for the null hypothesis that the fifth lag of the variables is zero is $F(16, 226) = 1.1683$ (with a p-value of 0.295) and indicates that it is statistically acceptable to simplify the system to a fourth-order VAR. Reducing the lag-order further induces residual misspesification.

Estimation of a fourth-order VAR with the five impulse dummies entering yields a more satisfactory representation of the system. Table 7 reports the residual standard errors and misspesification tests for the four equations individually and for the system.[19] The sample period is 1971(2)–1998(3). None of the tests are significant at the 5% level. Furthermore, the 1-step residuals shown in Figure 10 all lie within their respective ± 2 standard error bands. Thus, there are no strong indications of residual misspecification or parameter nonconstancies. The fourth-order VAR will form the basis of the cointegration analysis in the next subsections.

Table 7: Diagnostics for conditional fourth-order VAR 1971(2)–1998(3)

Single equation tests

$Variable$	pb	e	px	ulc
$\widehat{\sigma}$	0.023	0.013	0.008	0.023
$AR\ 1-5\ F(5, 76)$	1.168 [0.150]	0.846 [0.521]	1.130 [0.352]	0.378 [0.863]
$Normality\ \chi^2(2)$	3.565 [0.168]	2.598 [0.273]	0.234 [0.890]	0.152 [0.927]
$ARCH\ 4\ F(4, 73)$	1.703 [0.159]	0.549 [0.701]	1.119 [0.354]	2.107 [0.089]
$Hetero\ F(36, 44)$	0.945 [0.566]	0.494 [0.984]	0.650 [0.907]	0.806 [0.746]

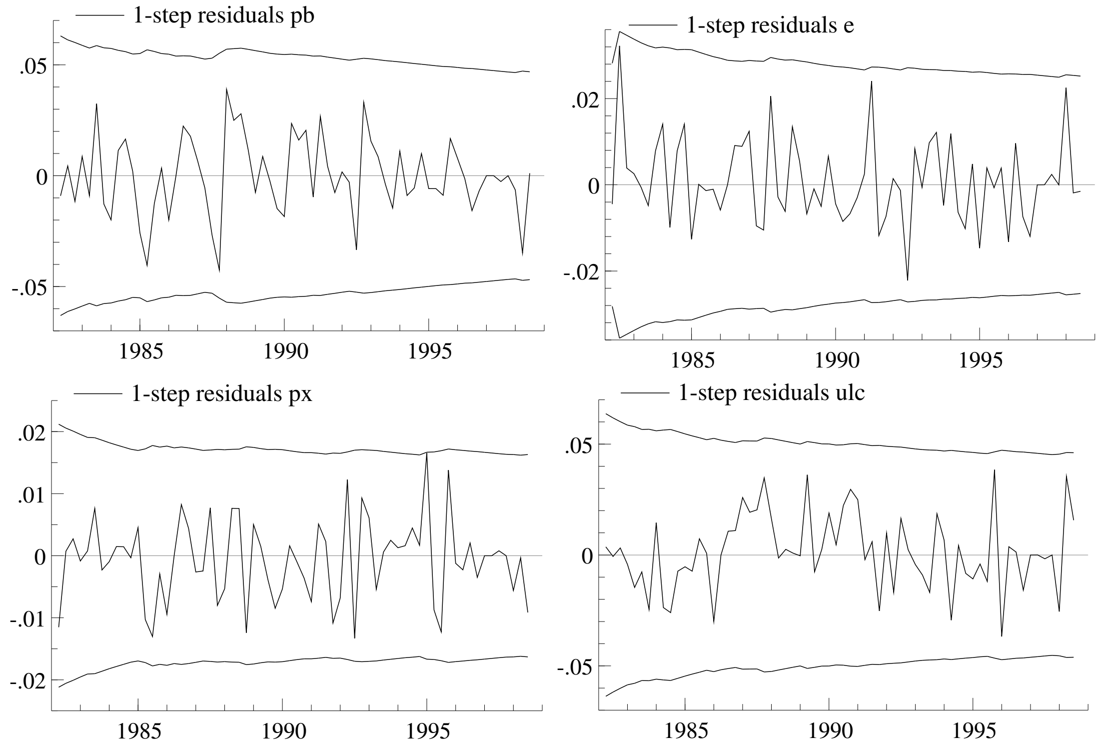
System tests

$Vector\ AR\ 1-5\ F(80, 231)$	1.213 [0.137]
$Vector\ normality\ \chi^2(2)$	9.428 [0.308]
$Vector\ heteroscedasticity\ F(360, 366)$	0.605 [1.000]

[18] The dummies are defined as follows: $PX74q1 = 1$ in 1974(1), 0 otherwise, $E86q2 = 1$ in 1986(2), 0 otherwise, $E97q1 = 1$ in 1997(1), 0 otherwise, $E97q2 = 1$ in 1997(2), 0 otherwise.

[19] See Doornik and Hendry (1997) for details and references.

Figure 10: 1-step residuals ±2 standard errors for fourth-order VAR 1971(2)–1998(3)



3.3.2 Determining cointegration rank

The next step is to determine the dimension of the cointegrating space. Table 8 reports the results from applying the Johansen procedure to the fourth-order VAR. It shows the four eigenvalues, the trace-statistics and the asymptotic 5% critical values taken from Table 2 in Harbo et al. (1998). Note that since the VAR includes several impulse dummies and a non-modelled differenced variable as conditioning variables, the reported critical values are only indicative. The trace statistic strongly rejects the null hypothesis of no cointegration ($r = 0$) in favour of at least one cointegrating vector, whereas the null of at most one cointegrating vector ($r \leq 1$) is not rejected at the 5% level. On the basis of these tests, we conclude that there is a single cointegrating vector. Table 8 also reports estimates of the eigenvectors $\boldsymbol{\beta}'$ and the adjustment coefficients $\boldsymbol{\alpha}$. The $\boldsymbol{\beta}'$ matrix is presented in normalised form, having one element of each row set equal to 1.

Next, we test for the absence of the linear trend in the cointegrating relations. The test is a conventional likelihood ratio (LR) test. Imposing the restriction $r = 1$, the model is in $I(0)$ space and the LR-statistic will be asymptotically distributed as $\chi^2(1)$. The observed value of the statistic is $\chi^2(1) = 7.129$ with a p-value of 0.008. Thus, the null hypothesis is rejected at the 1% significance level and the linear trend is retained in the model.

Table 8: Multivariate cointegration analysis

Eigenvalues			
0.332	0.153	0.119	0.090

Cointegration test

Null hypothesis	Trace-statistic	5% critical value
$r = 0$	86.92	71.7
$r \leq 1$	42.57	49.6
$r \leq 2$	24.34	30.5
$r \leq 3$	10.35	15.2

Normalised eigenvectors

pb	e	px	ulc	u_{t-2}	$trend$
1.000	−1.147	−1.105	−0.196	0.049	0.004
−1.736	1.000	0.648	1.8314	0.276	−0.018
−0.646	1.449	1.000	−0.123	0.009	−0.004
1.079	−0.126	−2.132	1.000	0.135	−0.011

Adjustment coefficients

pb	−0.488	0.026	−0.044	−0.004
px	0.009	0.006	−0.122	−0.013
e	−0.043	0.002	−0.046	0.013
ulc	−0.128	−0.093	−0.031	−0.006

Normalising the estimated cointegrating vector on import prices we find:

$$pb = +\underset{(0.137)}{1.147}e + \underset{(0.107)}{1.105}px + \underset{(0.117)}{0.196}ulc - \underset{(0.017)}{0.049}u_{-2} - \underset{(0.001)}{0.004}trend \tag{3.10}$$

with standard errors in parentheses. Giving a precise economic interpretation of the presence of a deterministic trend in the cointegrating vector is difficult. The trend coefficient is −0.004 which implies that in the long-run equilibrium (3.10) import prices decrease by 1.6% per annum after the influence of the other variables in the cointegrating vector is taken into account. The trend term might perhaps capture the effect of the continued growth in world trade and the increasing competition between producers in international markets.

Table 9: Testing the significance of a given variable in the cointegrating vector

Variables	pb	e	px	ulc	u	$trend$
$\chi^2(1)$-statistic	16.369	11.700	15.97	1.981	6.151	7.129
p-value	0.000	0.001	0.000	0.159	0.013	0.008

Table 9 reports LR-tests for the significance of each variable in the cointegrating vector. Except for the unit labour cost variable, all the variables are strongly significant. This confirms the results from the single equation analysis. Moreover, seeing that the estimated coefficient of unit labour costs is positive, all the coefficients have their expected signs. The long-run elasticities of exchange rates and foreign export prices are close to one, again confirming the estimates derived from the single equation model. The numerical value of the coefficient on the unemployment rate has dropped from 0.09 to 0.04, but the coefficient is still significant and indicates that domestic market conditions affect import prices even in the long run.

The adjustment coefficients in $\boldsymbol{\alpha}$ measure the feedback from disequilibrium in the long-run relationship onto the variables in the VAR. The vector of adjustment coefficients corresponding to (3.10) is given by

$$\widehat{\boldsymbol{\alpha}}' = [\underset{(0.080)}{-0.489}, \ \underset{(0.045)}{0.009}, \ \underset{(0.029)}{-0.043}, \ \underset{(0.084)}{-0.128}] \tag{3.11}$$

where, specifically, -0.489 is the estimated adjustment coefficient for the import price equation. A coefficient of -0.489 implies a relatively rapid correction of disequilibria in the cointegrating vector. The estimated adjustment coefficient is larger than what we found in the single equation analysis.

3.3.3 Testing cointegration restrictions

Having identified the single cointegration vector it is still of interest to test hypotheses on $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$. Testing for weak exogeneity of a given variable for the cointegrating vector amounts to testing whether the corresponding row of $\boldsymbol{\alpha}$ is zero. Table 10 reports likelihood ratio statistics for tests of weak exogeneity. The hypothesis that individually and jointly, exchange rates, foreign export prices, and unit labour costs are weakly exogenous for the cointegrating vector is supported by the data. Moreover, weak exogeneity of import prices is strongly rejected. These results imply that the cointegrating vector enters only the equation for import prices, and that single equation estimation of the long-run parameters is efficient. However, since we have included a restricted deterministic trend in the multivariate cointegration analysis, the results obtained from the single equation and the system analysis are not directly comparable.

Table 10: Weak exogeneity tests

Variable	pb	e	px	ulc	Joint test	$\{px, e, ulc\}$
$\chi^2(1)$-statistic	24.231**	0.035	2.230	1.872	$\chi^2(3)$-statistic	4.289
p-value	0.000	0.852	0.135	0.171	p-value	0.232

Table 11 reports statistics for tests of overidentifying restrictions on the cointegration vector. The hypotheses are formulated as restrictions on the equation $pb = \gamma e + \delta px + \kappa ulc + \lambda u + \mu trend$, and the degrees of freedom in the χ^2-distributions is equal to the number of independent restrictions to be tested. Not surprisingly, the hypothesis that domestic macroeconomic conditions do not affect import prices in the long-run ($\kappa = \lambda = 0$) is strongly rejected by the data. The hypothesis that the coefficients on foreign export prices and exchange rates are equal ($\gamma = \delta$) is accepted, as is the hypothesis that these coefficients are both equal to 1 ($\gamma = \delta = 1$). The statistic for testing the restriction of long-run unit homogeneity ($\gamma = \delta$ and $\kappa = 1 - \gamma$) has a p-value of zero and thus is not supported by the data. Note that, while data accepts that the long-run elasticities of the exchange rate and foreign export costs are equal to one, imposing this restriction joint with the hypothesis of long-run unit homogeneity ($\gamma = \delta = 1$ and $\kappa = 0$) is rejected. This casts doubts on the reasonableness of the equilibrium-correction term in the final single equation model of import prices in the previous section.

Table 11: Test of overidentifying restrictions on the cointegrating vector

Hypothesis	Statistic	p-value
H_1: $\kappa = \lambda = 0$	$\chi^2(2) = 18.393$**	0.000
H_2: $\gamma = \delta$	$\chi^2(1) = 0.123$	0.726
H_3: $\gamma = \delta = 1$	$\chi^2(2) = 1.050$	0.592
H_4: $\gamma = \delta$ and $\kappa = 1 - \gamma$	$\chi^2(2) = 15.453$**	0.000
H_5: $\gamma = \delta = 1$ and $\kappa = 0$	$\chi^2(3) = 16.284$**	0.001

Finally, imposing weak exogeneity of e, px, and ulc jointly with the restriction that the coefficients on exchange rates and foreign export prices are equal to 1 we get $\chi^2(5) = 5.49$ with a p-value of 0.359. The estimated cointegrating vector is found to be

$$pb = e + px + \underset{(0.039)}{0.298}ulc - \underset{(0.014)}{0.026}u_{-2} - \underset{(0.001)}{0.004}trend \tag{3.12}$$

and the corresponding estimate of the feedback coefficient is -0.501 with a standard error of 0.082.

3.4 Concluding remarks

In this section we have presented results from an econometric analysis of Norwegian import prices of manufactures over the period 1970(1)–1998(3). The purpose was to investigate the robustness of the results in Naug and Nymoen (1996) who found evidence of a cointegrating relationship between import prices, the exchange rate, foreign export prices, and domestic unit labour costs in a similar study covering the period 1970(1)–1991(4).

In addition to the extension of the sample period, several factors may contribute to differences in the results. These factors include revisions in the data and differences in the construction of the variables.[20] It is well-known from the empirical literature on exchange rates and traded goods prices that the selection of data can significantly affect the analysis. Another difference occurs in the treatment of the unemployment rate in the cointegration analysis. While Naug and Nymoen (op cit) treat the unemployment rate as stationary, the evidence in this and other studies[21] suggests that the transformed rate of unemployment rate behaves as if it were an $I(1)$ variable. Based on an untested assumption of weak exogeneity, we therefore include the unemployment rate as a non-modelled variable in the long-run part of the VEqCM. Then, for the asymptotic distribution of the cointegration test statistics to be free of nuisance parameters, we also add a restricted deterministic trend to the model. After having determined the cointegrating rank, the significance test on the trend coefficient leads us to retain the trend in the cointegrating vector. The presence of a deterministic trend in the cointegrating vector is one important difference between the results in this study and those of Naug and Nymoen (op cit).

The results from both the single equation and the system analysis lead us to conclude that there is a single cointegrating relationship between the variables in the model. This is consistent with the results in Naug and Nymoen (op cit), as is the significant effect of the unemployment rate in the long-run solution. The result that increases in domestic demand pressure lead to increases in import prices thus appears to be robust. What is not clear from the economic theory discussed in Section 2, however, is how this apparently robust effect should be interpreted in relation to the pricing to market hypothesis.

The hypothesis that there are significant pricing to market effects in Norwegian import prices is not supported by the data in the present study. Both the single equation estimates and the estimates obtained using the Johansen procedure suggest that the long-run pass-through of changes in exchange rates and foreign export prices is complete, and this conclusion is not altered if the deterministic trend is dropped from the model prior to the cointegration analysis. However, the presence of a deterministic trend in the cointegrating vector and the fact that we fail to find a long run equilibrium with homogeneity indicate that there is scope for further modelling.

[20] In particular, we employ data for unit labour costs in manufacturing and construction, while Naug and Nymoen (1996) employ data for unit labour costs in manufacturing only.

[21] See Bjørnstad and Nymoen (1999).

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A Variable definitions and sources

Symbol	RIMINI	Definition and source
PB	PBI	Deflator of imports of manufactures. 1996=1. Source: Quarterly National Accounts (Statistics Norway)
E	$PBVAL$	Effective import weighted value of the NOK. 1996=1. Source: Norges Bank's databank of economic time series
PX	$PBPRIS$	Import weighted foreign export prices. 1996=1. Source: International Financial Statistics (IMF)
WC	$WCIBA$	Hourly wage cost in manufacturing and construction. Source: QNA (Statistics Norway)
Z	$ZYIBA$	Value added labour productivity in manufacturing and construction. Fixed 1996-prices. Source: Norwegian National Accounts (Statistics Norway)
U	$UTOT2$	"Total" unemployment rate, fraction of total labour force. Source: Norges Bank's databank of economic time series

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