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Trade liberalisation and import price behaviour: the case of textiles and wearing apparels

Abstract:
Previous studies on the relationship between exchange rates and traded goods prices typically find evidence of incomplete pass-through, usually explained by pricing-to-market behaviour. Although economic theory predicts that incomplete pass-through may also be linked to presence of non-tariff barriers to trade, variables reflecting such a link is rarely included in existing empirical models. In this paper, we estimate a pricing-to-market model for Norwegian import prices on textiles and wearing apparels, controlling explicitly for the removal of non-tariff barriers to trade and the shift in imports from high-cost to low-cost countries through a Törnqvist price index based measure of foreign prices. We show that this measure of foreign prices unlike standard measures used in the literature is likely to produce unbiased estimates of the degree of pass-through, and thereby also the extent of pricing-to-market behaviour. Finally, we demonstrate that the estimated import price equation is reasonably stable and exhibits no serious forecasting failures. These findings contradict the hypothesis that pass-through has changed alongside trade policy shifts during the second half of the 1990s and the monetary policy regime shift in 2001.

Keywords: Trade liberalisation, import prices, pricing-to-market, exchange rate pass-through, vector autoregressive models.

JEL classification: C22, C32, C43, E31

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1 Introduction

A key topic in monetary economics of interest for policy makers in general and inflation targeting central banks in particular is the responsiveness of prices of internationally traded goods to changes in nominal exchange rates. Empirical research on this area, known as the degree of exchange rate pass-through (henceforth pass-through), is abundant. Typically, existing studies find evidence of incomplete pass-through, which is often explained by pricing-to-market behaviour under conditions of imperfect competition and segmented markets, see e.g. Menon (1995a), Goldberg and Knetter (1997), Gil-Pareja (2003), Herzberg *et al.* (2003), Campa and Goldberg (2002, 2005) and Bugamelli and Tedeschi (2008). Also, empirical studies of small open economies confirm that import prices do not fully respond to changes in exchange rates and that domestic market conditions influence the price setting behaviour of foreign firms, see Menon (1995b, 1996), Naug and Nymoen (1996), Alexius (1997), Kenny and McGettigan (1998) and Doyle (2004).[1] These findings contradict the "law of one price" and have important implications for monetary policy as import prices apparently are *not* exogenously determined in foreign currency on the world market.

Most previous studies of small open economies have, however, in common that they analyse the degree and determinants of pass-through by means of relatively aggregated data. For instance, Naug and Nymoen (1996) investigate Norwegian import prices of total manufactures including different types of commodities, raw materials and quota-protected products (e.g. textiles, wearing apparels and footwear) over the sample period $1970Q1 - 1991Q4$. The general finding in studies using disaggregated, industry level data is that pass-through varies considerably across industries and product categories, which raises the question of possible aggregation bias in the pass-through estimates obtained on aggregated data. Moreover, previous studies of small open economies usually ignore the hypothesis that presence of non-tariff barriers to trade may affect pass-through, see e.g. Naug and Nymoen (1996) and Doyle (2004). Briefly speaking, the hypothesis known as the "Bhagwati hypothesis", see Bhagwati (1991), says that in the presence of quantity restraints on imports a small depreciation of the exchange rate is likely to be absorbed into the quota rents extracted by the exporter rather than being reflected in import prices. If the depreciation, on the other hand, is large enough to push import prices above the point where the quantity restraints are no longer binding, then pass-through will be positive, but incomplete. Menon (1996) estimates a range of price models for Australian imports of total manufactures and forty different product categories contained therein with strong and significant negative effects of quantity restraints on trade as a separate explanatory variable.

[1] Recently, Bache and Naug (2007) estimate a wide range of New Keynesian import price models for total Norwegian and UK manufactures and find that only for the UK do the results suggest a role for domestic prices or costs in explaining import prices.

In this paper, we estimate a model for Norwegian import prices on textiles and wearing apparels (henceforth clothing) that explicitly controls for the shift in imports from high-cost to low-cost countries and the gradual removal of non-tariff barriers to trade experienced in the clothing industry since the mid 1990s. The model is based on the pricing-to-market theory by Krugman (1987) and is estimated on quarterly time series data over the period $1986-2008$. By using data at the disaggregated industry level we may avoid aggregation bias typically associated with the previous studies on more aggregated data. We apply the cointegrated VAR framework to quantify the degree of pass-through and pricing-to-market, thereby paying attention to the time series properties of the variables involved.

The motivation of our study follows from the fact that low consumer price inflation observed over several years in Norway to a large extent can be attributed to a simultaneous fall in import prices on clothing. The development in import prices on clothing during the last two decades may partly be explained by traditional factors such as shifts in exchange rates, international prices (measured in foreign currency) and domestic market conditions. However, it should also be viewed in light of the massive trade liberalisation, which increased the imports of clothing from China and other low-cost countries at the expense of imports from high-cost countries, the euro area in particular. The significant deflationary effect on traded goods prices of shifts in the country composition of imports has become commonly known among economist as the *China effect* and is likely to be important when quantifying pass-through in regression models. That is, the gradual removal of quota restrictions on trade may in accordance with the Bhagwati hypothesis have pushed the estimate of pass-through upwards, an empirical question which we pursue in the present paper.

To answer this question we construct two different measures of foreign prices to be used in the estimation of the degree of pass-through. The first measure is based on the total differentiation of the Törnqvist price index. Accordingly, we are not only able to take account of *inflationary differences* as is standard in the literature, but also *varying import shares* and *differences in price levels* − i.e., the China effect − among trading partners when constructing the measure of foreign prices. The second measure of foreign prices is based on the often used geometric mean price index with constant import shares as weights, a measure which thereby fails to take account of the China effect. By comparing estimates of pass-through that come out of modelling the import price of clothing with the two alternative measures of foreign prices, we are able to shed some light on the potential problem of omitted variable bias in empirical tests of pricing-to-market.[2] Finally, regime

[2] To our knowledge, no previous studies have estimated pricing-to-market models with a Törnqvist price index based measure of foreign prices. Generally, there are few academic papers that examine the impact of increased imports from China and other low-cost countries on traded goods prices and overall inflation in developed countries. Nickell (2005) and Coille (2008) compute the China effect on traded goods prices with different operational indices and with no estimated models when analysing the impact of a changing trade pattern on import prices and overall consumer

shifts in monetary policy may, just like changes in trade policy, influence the degree of pass-through. The monetary policy regime in Norway switched from managed float to inflation targeting and a freely floating exchange rate in early 2001, which has brought about more volatile exchange rates facing foreign exporters. In light of Froot and Klemperer (1989), pass-through may have become significantly higher if producers have regarded the exchange rate appreciations after 2001 as more permanent than previous appreciations. We pursue this hypothesis by a forecasting exercise on the estimated import price model to examine its stability properties or lack thereof.

One important finding in this paper is that the China effect on traded goods prices is substantial in the clothing industry. Our calculations suggest that the shift in imports from high-cost to low-cost countries since the early 1990s on average has reduced the international price impulses on imports of clothing by around 2 percentage points per year. Controlling for these effects by means of the Törnqvist price index based measure of foreign prices, we estimate an import price model consistent with the pricing-to-market hypothesis. Specifically, the pass-through and pricing-to-market elasticities are significantly estimated to 0.45 and 0.55, respectively. We find that the use of the geometric mean based measure of foreign prices biases the estimates due to international price impulses being substantially overestimated. We also establish that the estimated dynamic equilibrium correction model (both in terms of its long and short run parts) is reasonably stable in-sample and exhibits no serious forecasting failures around the dates of the shifts in trade and monetary policies. That no serious structural breaks are detected may reflect that likely pass-through effects of changes in trade policy are controlled for through the Törnqvist price index based measure of foreign prices. Consequently, once the effect of shifts in imports towards low-cost countries is controlled for, we find little evidence that the slopes of the import price equation have changed alongside trade liberalisation.

The rest of the paper is organised as follows: Section 2 outlines the pricing-to-market theory for a small open economy and discusses the effects of non-tariff barriers to trade on pass-through. Section 3 presents the construction of the two alternative measures of foreign prices and the data used in the empirical analysis. Section 4 describes and reports results from the cointegrated VAR modelling, while Section 5 presents the estimated dynamic equilibrium correction model for import prices on clothing. Section 6 concludes.

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price inflation in the UK. Wheeler (2008) also calculates the China effect on traded goods prices following the operational route in Nickell (2005), but in addition estimates panel regressions of UK inflation by goods category on the level and growth of the import share from China as the main determinants. See also Høegh-Omdal and Wilhelmsen (2002) and Røstøen (2004) for discussions on the effects of trade liberalisation and shift in imports towards low-cost countries on clothing prices and on overall consumer price inflation in Norway.

2 The theoretical framework

The underlying theoretical model for the behaviour of import prices on clothing is based on the pricing-to-market theory by Krugman (1987). We thereby take into account that markets for clothing typically are characterised by imperfect competition between firms producing differentiated products. Furthermore, markets for clothing are segmented due to trade barriers, transportation costs and imperfect information.[3] Profit maximisation under such circumstances normally implies that foreign exporters can charge different markups over their marginal costs, and hence can charge different prices, depending on the conditions in each particular market. The following exposition of the pricing-to-market model and the relationship between pass-through and presence of (and removal of) non-tariff barriers to trade build on Naug and Nymoen (1996) and Menon (1996), respectively.

2.1 Pricing-to-market

Consider a representative foreign firm producing a differentiated product of clothing exported to i segmented markets or countries ($i = 1, \ldots, n$). The product is assumed to be weakly separable from all other competing goods in the consumer's utility function. The demand faced by the firm in each export market may then be expressed as $X_i = X_i(PX_i/ER_i, PQ_i, DP_i)$, where PX_i is the firm's export price measured in the exporter's currency, ER_i is the bilateral exchange rate with country i, PQ_i is an index of prices on competing products, and DP_i represents other factors affecting demand (henceforth referred to as demand pressure). The profit of the firm is given by

$$\begin{aligned} \Pi(PX_i, ..., PX_n) &= \sum_{i=1}^{n} PX_i \cdot X_i\left(PX_i/ER_i, PQ_i, DP_i\right) \\ &\quad - C\left[\sum_{i=1}^{n} X_i\left(PX_i/ER_i, PQ_i, DP_i\right), W\right], \end{aligned} \tag{1}$$

where $C[\cdot]$ is the cost function depending on production and input prices (W). Time arguments are provisionally suppressed for simplicity. Profit maximisation generates the following first order conditions

$$PX_i = \lambda_i MC, \ i = 1, ..., n. \tag{2}$$

Hence, the foreign firm sets each export price as a markup (λ_i) on the common marginal costs (MC) measured in the currency of the exporter. Generally speaking, $\lambda_i = \eta_i/(\eta_i - 1)$, where $\eta_i = \eta_i(PX_i/ER_i), PQ_i, DP_i)$ is the elasticity of demand in market i. As every export price reflects conditions in each particular market,

[3] See Moe (2002) for a more thorough description of the Norwegian clothing market.

profit maximisation typically leads to price discrimination, and thus market specific markups. The import price (PI_i) measured in the currency of the importing country i is obtained by multiplying through (2) with the bilateral exchange rate ER_i.

$$PI_i = ER_i PX_i = ER_i \lambda_i MC,\ i = 1, ..., n. \tag{3}$$

Following Naug and Nymoen (1996), we abstract from competition between foreign firms in market i to simplify matters and specify the destination specific markup as $\lambda_i = K_i (PD/PI)_i^{\gamma_{1i}} DP_i^{\gamma_{2i}}$, where K_i is a constant, PD_i/PI_i is the price on competing goods *produced* in market i relative to the import price and DP_i is the demand pressure in the importing country. Economic theory predicts that $\gamma_{1i} \geq 0$, because higher prices on competing goods imply a potential for increasing markups. The sign of γ_{2i} is, however, undetermined from theory. An increase in the demand pressure may rise the scope for an increase in the markup, but may very well also increase economies of scale in production and distribution, and hence pave the way for a decrease in the markup. Substituting the expression for λ_i into (3) and using lower case letters to indicate natural logarithms, we obtain[4]

$$pi_i = \kappa_i + (1 - \psi_i)(mc + er_i) + \psi_i pd_i + \delta_i dp_i,\ i = 1, ..., n, \tag{4}$$

where $\kappa_i = \ln K_i/(1 - \gamma_{1i})$, $\psi_i = \gamma_{1i}/(1 - \gamma_{1i})$ and $\delta_i = \gamma_{2i}/(1 - \gamma_{1i})$. When $\psi_i > 0$ domestic prices (pd_i) matter for the determination of import prices, and changes in marginal costs and the exchange rate are not entirely passed through to import prices. This phenomenon is what Krugman (1987) labelled pricing-to-market. The degree of pass-through from mc and er_i to pi_i is given by the coefficient $(1 - \psi_i)$. In the special case when $\psi_i = 0$, the pass-through from mc and er_i is complete, and pd_i has no role in the determination of import prices. Conversely, $\psi_i = 1$ implies zero pass-through.

The law of one price (henceforth LOP) is the standard assumption of import pricing in theoretical models of small open economies, and follows as a special case of (4). As pointed out by Naug and Nymoen (1996), the absolute version of LOP requires full pass-through ($\psi_i = 0$), no effects from domestic demand pressure ($\delta_i = 0$) and the same markup ($\kappa_i = \kappa > 0$) in all countries, which implies that $PX_i = PX$ in all markets. The relative version of LOP, on the other hand, only requires ($\psi_i = \delta_i = 0$) in all countries. Hence, the relative version of LOP allows price discrimination through a varying constant (κ_i) − which under both versions of LOP equals the markup (λ_i) − across markets.

The pricing-to-market model outlined here is based on foreign firms' price setting behaviour and two channels through which domestic factors in the importing country may affect import prices on clothing, namely through competitive pressure (pd_i) and demand pressure (dp_i) in the importing country. Another motivation for

[4]In what follows, lower case letters indicate natural logarithms of a variable unless otherwise stated.

including pd_i and dp_i in the model would be when importers of clothing act as agents and find domestic factors important in price negotiations with foreign producers. The model implicitly assumes, on the other hand, that markets for clothing are segmented due to inter alia presence of non-tariff barriers to trade. As previously mentioned, such trade barriers may limit the degree of pass-through according to the Bhagwati hypothesis, an issue which we now turn to.

2.2 Non-tariff barriers to trade

Up until the so-called Uruguay Round starting in 1986, the clothing industry was among the most strictly regulated manufacturing sectors, both in terms of tariffs and quantity restrictions on trade. During the 1970s and 1980s, the Norwegian market for clothing was mainly regulated through the Multi-Fibre Agreement, an agreement that allowed importers to negotiate bilateral export restraint quotas with low-cost countries. As an example, Norway signed some 20 bilateral agreements in 1984 with countries in Asia and eastern Europe. The Uruguay Round, however, led to major changes in the trade policy and it was decided that quota regulations should be eliminated between 1995 and 2005. Norway was relatively quick in liberalising the quota system and the last quantity restrictions on trade with clothing were abolished in 1998. The removal of quotas has no doubt contributed significantly to further increase in imports of clothing from low-cost countries during the last ten to fifteen years. Substantial reduction over time in tariff rates on imports of clothing has likewise pulled in the same direction.[5]

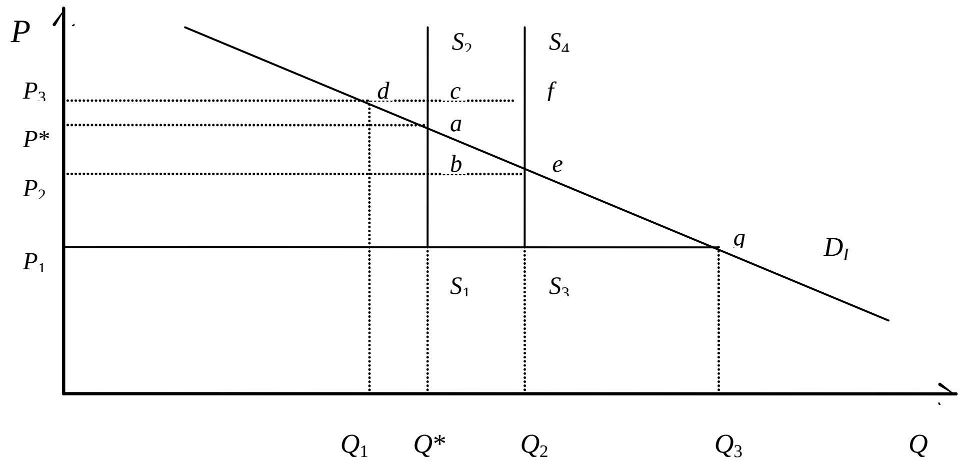
Here we shall focus on the link between non-tariff barriers to trade and pass-through as the empirical analysis is based on import prices of clothing exclusive tariffs. Because the effects that the quantity restrictions are likely to have on pass-through do not depend on particular market structures, we extend Menon's (1996) analysis and highlight the relationship between pass-through and gradual removal of non-tariff barriers to trade by means of a small country being a price taker with respect to its imports. Figure 1 illustrates the implications of the Bhagwati hypothesis for pass-through in the presence and removal of quantity restrictions on trade.

The demand curve for imports is represented by D_I, whereas the supply curve consists of the horizontal line P_1S_1 and the vertical line S_1S_2. The supply curve is perfectly elastic at P_1 (reflecting the small country assumption) and becomes perfectly inelastic when the quantity restrictions on trade are met at Q^*. The initial equilibrium is at point a with quantity Q^* and price P^*. At point a the seller is able to pull out $P_1S_1aP^*$ in quota rents due to the presence of quantity restrictions.

A small depreciation of the importing country's currency will shift the horizontal part of the supply curve upwards, while the vertical part is unchanged. For

[5] For instance, the average ordinary tariff rate was reduced from about 20 per cent in 1994 to 12 per cent in 2004. See Melchior (1993) and Høegh-Omdal and Wilhelmsen (2002) for summaries of clothing trade policies in Norway.

Figure 1: Pass-through with presence and removal of quota restrictions



example, a depreciation of the currency to P_2 will neither affect equilibrium quantity nor market price, but will reduce the quota rents to P_2baP^*. It follows that the depreciation is entirely absorbed into the quota rents and that pass-through is zero. However, if the depreciation is large enough to push the market price above P^* to say P_3, the horizontal supply curve (P_3c) will be taken to a level where the quantity restrictions are no longer binding. At the new equilibrium point d quantity falls below the quota limit to Q_1 and the market price increases from P^* to P_3. Hence, some part of the currency depreciation is now passed through to the import price. Specifically, the degree of pass-through in this situation equals the change in the market price relative to the magnitude of the currency depreciation, that is $(P_3 - P^*)/(P_3 - P_1) < 1$ as $P^* > P_1$.

Suppose instead that trade liberalisation takes place in the sense that quantity restrictions on trade are effective at Q_2 rather than at Q^*. Consequently, the horizontal supply curve is represented by the line P_1S_3, whereas the vertical supply curve (which shifts to the right alongside the reduction in the quota restrictions) is represented by the line S_3S_4. The new initial equilibrium is at point e with quantity Q_2, price P_2 and quota rents $P_1S_3eP_2$. We notice that $P_1S_1aP^* > P_1S_3eP_2$. The possibilities to absorb currency depreciations into the quota rents are reduced in situation e compared to situation a as $P^* > P_2$. If a currency depreciation again pushes the market price to P_3, so that the horisontal supply curve shifts to the line P_3f, the equilibrium point d is still reached. However, both the quantity and the market price will change relatively more for a given currency shock when the initial equilibrium is at point e rather than at point a, where no reduction in the quota restrictions has yet taken place. In other words, a reduction in the quota restrictions from Q^* to Q_2 implies that pass-through to import prices will be higher, other things equal. To see this, we notice that the degree of pass-through in situation e equals $(P_3 - P_2)/(P_3 - P_1)$, which is greater than $(P_3 - P^*)/(P_3 - P_1)$ because $P_2 < P^*$. Pass-through is still incomplete in situation e as $P_2 > P_1$. Only when the quantity restrictions on trade are entirely removed, as in situation g in Figure 1, will pass-through be complete.

To summarise, a currency depreciation in the presence of non-tariff barriers to trade will generally reduce the quota rents first, hence absorbing much of its impact, before it is reflected in the market price. It is only when the depreciation is large enough to push the market price above the point where the quota restrictions are no longer binding that pass-through will be positive, but incomplete according to the Bhagwati hypothesis. Finally, if incomplete pass-through is inter alia linked to presence of non-tariff barriers to trade, gradual removal of such barriers will push pass-through upwards, other things equal.

3 From theory to empirics

In this section, we describe the operational route chosen to obtain an empirical counterpart of (4) that explicitly controls for the effects of shift in imports towards low-cost countries and the removal of quota restrictions on trade discussed above. Because the focus is on aggregated time series for one destination country, namely Norway, we first translate (4) into a testable empirical representation by replacing the index i with the subscript t to denote time. We further replace marginal costs, which are not directly observable, with either a measure of foreign prices that is based on (i) the Törnqvist price index (pf_t^T) with varying import shares as weights or (ii) the geometric mean price index (pf_t^G) with constant import shares as weights. Besides, we approximate domestic prices and demand pressure with variable unit costs (vc_t) and the unemployment rate (UR_t), respectively, and add a disturbance term (u_t) to (4). The following empirical representation of (4) emerges:

$$pi_t = const. + (1-\psi)(pf^i + er)_t + \psi vc_t + \delta UR_t + u_t, i = T, G. \tag{5}$$

We remark that the unemployment rate enters (5) without a logarithmic transformation. Accordingly, the markup underlying (5) is specified as $\lambda_t = K(VC/PI)_t^{\gamma_1}$ $\exp^{(\gamma_2 UR_t)}$.[6] As we use price indices and not price levels (which are not available)

[6] Naug and Nymoen (1996) emphasise that the use of a geometric mean of export prices proxying marginal costs induces measurement errors as the disturbance term contains the foreign producers' markups. The disturbance term is thus correlated with the export price measure and their pricing-to-market model similar to (5) only forms a cointegration relationship when the measurement errors are stationary. Similarly, the disturbance term in (5) contains domestic producers' markups when we replace domestic prices by variable unit costs. As pointed out by Naug and Nymoen (1996), these measurement errors may be correlated with the unemployment rate representing demand pressure. If markups of foreign firms are affected by domestic demand pressure, we expect that markups of domestic firms also are influenced. We therefore acknowledge that effects of demand pressure will be overestimated in (5) to the extent that u_t is correlated with UR_t.

in the empirical analysis, the absolute version of LOP is not testable from (5). The analysis of LOP is therefore limited to its relative version for Norway as the only market, i.e., whether $\psi = \delta = 0$ in (5). A testable implication of LOP is that $(pi - pf^i - er)_t$ is stationary or forms a long run cointegration relationship when the variables involved all are nonstationary. If imports and domestic products of clothing are close substitutes, we expect LOP to be a reasonable approximation, and pass-through to be nearly complete. The long run version of PPP implies similarly that $(vc - pf^i - er)_t$ is stationary. As noted by Naug and Nymoen (1996), the long run versions of LOP and PPP may be consistent with (5) rewritten as $(pi - pf^i - er)_t = const + \psi(vc - pf^i - er)_t + \delta UR_t + u_t$. We see that this equation is balanced when $\psi > 0$, $\delta \neq 0$ and $(pi - pf^i - er)_t$, $(vc - pf^i - er)_t$, UR_t and u_t all are stationary variables. If both LOP and PPP hold in the long run, then pricing-to-market is only a short run phenomenon and (5) predicts the existence of two cointegrating vectors relating the variables. On the other hand, if $(pi - pf^i - er)_t$ and $(vc - pf^i - er)_t$ are nonstationary, neither LOP nor PPP holds in the long run, and $(pi - pf^i - er)_t - \psi(vc - pf^i - er)_t$ is stationary. In this case, pricing-to-market is a long run phenomenon.

We also notice that (5) imposes the same coefficient on pf_t^i and er_t as well as unit homogeneity between pi_t, $(pf^i + er)_t$ and vc_t. In practice, however, these restrictions need not hold. Exchange rates are typically more volatile than costs, and foreign exporters may be more willing to absorb into their markups changes in exchange rates (which are likely to be permanent) than changes in costs. We test the parameter restrictions in the empirical analysis rather than imposing them from the outset.

3.1 The measure of foreign prices

After the resolution following the Uruguay Round to gradually dismantle the system of import quotas, the share of imports from China and other low-cost countries has increased steadily at the expense of imports from high-cost countries within the Euro area. The shift in imports towards countries with lower price levels (production costs) has contributed to reduced purchasing prices for Norwegian importers of clothing, the so-called China effect on traded goods prices. We attempt to take account of these effects by constructing a measure of foreign prices that is based on the total differentiation of the Törnqvist price index. Hence, we not only allow for *inflationary differences* as is common in related studies, but also *varying import shares* and *price level differences* among trading partners when constructing the measure of foreign prices. The fact that available data on foreign prices on clothing are *indices* and not levels makes the Törnqvist price index (PF^T) in our context equal to

$$PF^T(t,t-1) = \prod_{j=1}^{n} PF_{j,t}^{(s_{j,t-1}+s_{j,t})/2}, \tag{6}$$

where $s_{j,t-1}$ and $s_{j,t}$ are the value shares of imports from trading partner j in the base and comparison period, respectively, $0 \leq s_{j,h} < 1$ and $\sum_{j=1}^{n} s_{j,h} = 1$ $(h = t-1, t)$.[7] We observe that the aggregate foreign price $\left(PF^T\right)$ of clothing is a weighted geometric average of the foreign price indices (PF_j), the weights being the arithmetic means of the value import shares of the base and comparison period. In the following, we only consider two trading partners $(j = 1, 2)$ to simplify matters without loss of generality. Taking natural logarithms of (6), differentiating with respect to both $PF_{j,t}$, $s_{j,t}$ and $s_{j,t-1}$ and making use of the summing up condition of import shares, we obtain an expression for the percentage change in the aggregate foreign price in period t that reads as

$$\begin{aligned} \Delta pf_t^T &= 0.5\,(s_{1,t-1} + s_{1,t})\,\Delta pf_{1,t} + 0.5\,(s_{2,t-1} + s_{2,t})\,\Delta pf_{2,t} \\ &\quad + 0.5(s_{1,t} - s_{1,t-2})\,(pf_{1,t-1} - pf_{2,t-1}), \end{aligned} \tag{7}$$

where Δ indicates the first difference of a variable. By calculating Δpf_t^T in this way, we allow for inflationary and price level differences as well as varying import shares among the main Norwegian trading partners. The two first terms on the right hand side of (7) show that increasing inflation on clothing from each of the trading partners contribute to increasing international inflationary impulses faced by Norwegian importers. The larger the price increase and the larger the import share, the larger is the foreign inflationary impulse (measured in foreign currency) of pf_t^T. The last term on the right hand side of (7) constitutes the China effect. If the import share is changing in favour of a low-cost country, the last term becomes negative. The larger the change in the import share and the larger the difference in price levels, the larger is the foreign deflationary impulse of pf_t^T. We notice that the China effect is zero with constant import shares. Although the cross-country distribution of the China effect can be sensitive to the choice of numeraire country, the size of the *aggregated* China effect calculated from (7) is not affected.

Traditional measures of foreign prices (measured in foreign currency) on imported commodities typically fail to capture the price level effects of the switch in imports from high-cost to low-cost countries. The standard practise in related studies is to weight together some proxy for foreign prices by means of a geometric

[7] The Törnqvist price index is a discrete time approximation to the continuous time Divisia price index, see e.g. Balk (2008, p. 25). Also, the Törnqvist price index is defined as the geometric mean of the geometric Laspeyres and Paasche price indices, see Balk (2008, p. 72). Nickell (2005) and Wheeler (2008) use the geometric Paasche price index as a basis for calculating aggregate foreign export prices (measured in Sterling) on all products faced by UK importers.

mean price index with constant import shares as weights, see e.g. Naug and Nymoen (1996), Kenny and McGettigan (1998), Herzberg *et al.* (2003) and Campa and Goldberg (2005). The geometric mean price index in our context is equal to

$$PF^{G}(t) = \prod_{j=1}^{n} PF_{j,t}^{\bar{s}_j}, \tag{8}$$

where the exponent $\bar{s_j}$ now is the constant value share of imports from trading partner j, $0 \leq \bar{s_j} < 1$ and $\sum_{j=1}^{n} \bar{s_j} = 1$. Following Naug and Nymoen (1996), we set $\bar{s_j}$ equal to the average of each import share over the sample period. Again, taking natural logarithms of (8), differentiating with respect to $PF_{j,t}$ $(j = 1, 2)$ and making use of the summing up condition of import shares, we get the following expression for the percentage change in the aggregate foreign price in period t:

$$\Delta pf_t^G = \bar{s_1}\Delta pf_{1,t} + \left(1 - \bar{s_1}\right)\Delta pf_{2,t}. \tag{9}$$

We see from (9) that only inflationary differences among the trading partners are allowed for when constructing the aggregate measure of foreign prices. As the downward pressure on foreign prices from the China effect is potentially important, it is likely that international price impulses will be overestimated by (9). On this background, we expect that the estimate of the degree of pass-through will reflect an omitted variable bias when the geometric mean based measure of foreign prices is used instead of the Törnqvist price index based measure in a regression model for import prices of clothing, other things equal. One way to remedy this potential econometric problem may be to add a linear trend to approximate the price level term in (7). However, we thereby implicitly assume that the China effect has been constant over the sample period, a strict assumption to impose on the regression model from the outset. We argue in this paper that a more flexible and reliable approach is to allow the China effect, and thereby also the consistency of the degree of pass-through, to be entirely controlled for through the Törnqvist price index based measure of foreign prices.

3.2 Data[8]

We now describe the operational route based on available data for each one of the variables in (5) in more detail. The data are quarterly, seasonally unadjusted time series covering the period $1986Q1 - 2008Q1$. After allowing for lags, estimation is conducted over the period $1986Q4 - 2008Q1$ unless otherwise noted. We refer to the Appendix for details about the data definitions and sources.

The import price (pi) is an implicit deflator for imports of clothing with Norwegian substitutes. The products comprising the deflator are priced *cif* at the

[8]The data are available from the authors upon request.

Norwegian boarder. Hence, prices include costs of insurance and freight, but exclude tariffs. The deflator is a chained geometric mean price index calculated by weigthing together each one of the unit prices, which are based on the value and volume of each single imports, with the corresponding import share (measured in value) of each trading partner. Because the import shares are continously updated in accordance with the development in the country composition of clothing imports, the deflator reflects the shifts in imports from high-cost to low-cost countries over time. The import price on clothing is measured in the Norwegian currency and does not show directly the international price impulses as such. Also, the import price will reflect effects of changes in the exchange rate, whose degree of pass-through may depend on presence of (or removal of) non-tariff barriers to trade and pricing-to-market behaviour among foreign firms.

The operational route for the measure of foreign prices is not so clear-cut as the measure of import prices on clothing. To construct pf^T based on (7), we need data on import shares, export prices and price levels for each one of the main trading partners. According to the foreign trade statistics from Statistics Norway, which produces reliable time series of import shares by country, the main exporters of clothing to Norway studied here are China (CH), the euro area (EU), the United Kingdom (UK), Denmark (DK), Sweden (SW), Hong Kong (HK) and Turkey (TR). Together these countries covered nearly 80 per cent of Norwegian imports of clothing as an average over the sample period.[9] Because the euro area is treated as one country, we abstract from any import substitution from high-cost to low-cost countries within the monetary union.

It proved difficult to find long and consistent proxies for export prices when it comes to the low-cost countries China and Turkey. We therefore approximate Chinese export prices by connecting producer prices on clothing available from $1997Q1$ together with consumer prices on all products available from $1986Q1$. The fact that these two time series are highly correlated during the period $1997Q1 - 2008Q1$ may make consumer prices a fairly good proxy for producer prices of clothing during the first half of the sample period. Similarly, we connect Turkish export prices on clothing available from $2004Q1$ together with export prices on manufactures available from $1995Q1$ and import prices on all products available from $1986Q1$. We acknowledge that the Turkish export price measure may be a broad proxy, especially when import prices due to lack of more relevant data are used in the period $1986Q1 - 1994Q4$. However, pf^T is not much affected by the Turkish export price proxy during that subperiod as the import share is more or less constant around the level of 1.5 per cent up until 1995.

Price level differences among the trading partners should ideally be based on comparable price levels on clothing that reflect the level of production costs corrected

[9]The rest of exports of clothing to Norway came from countries with relatively small import shares during the 1980s and 1990s. Indeed, Bangladesh was represented by an import share of about 8 per cent in 2008, but is left out of the analysis due to lack of relevant price data.

for the level of productivity in each country. Because such data are not available we use purchasing power parity adjusted GDP figures. We do believe that there are price level differences in international trade that should be adjusted so that using purchasing power parity adjusted GDP amounts to taking differences in domestic costs into account in our model. Using data from IMF, the price level measure for each country from which Norway imports clothing appears by dividing nominal GDP by the purchasing power parity adjusted volume of GDP. Table 1 shows the average calculated price levels in each country as a share of the euro area price level over the period $1991-2008$.[10]

Table 1: Average price levels. $1991-2008$

DK	SW	UK	EU	HK	TR	CH
1.30	1.25	1.05	1.00	0.91	0.56	0.41

Sources: IMF and Statistics Norway

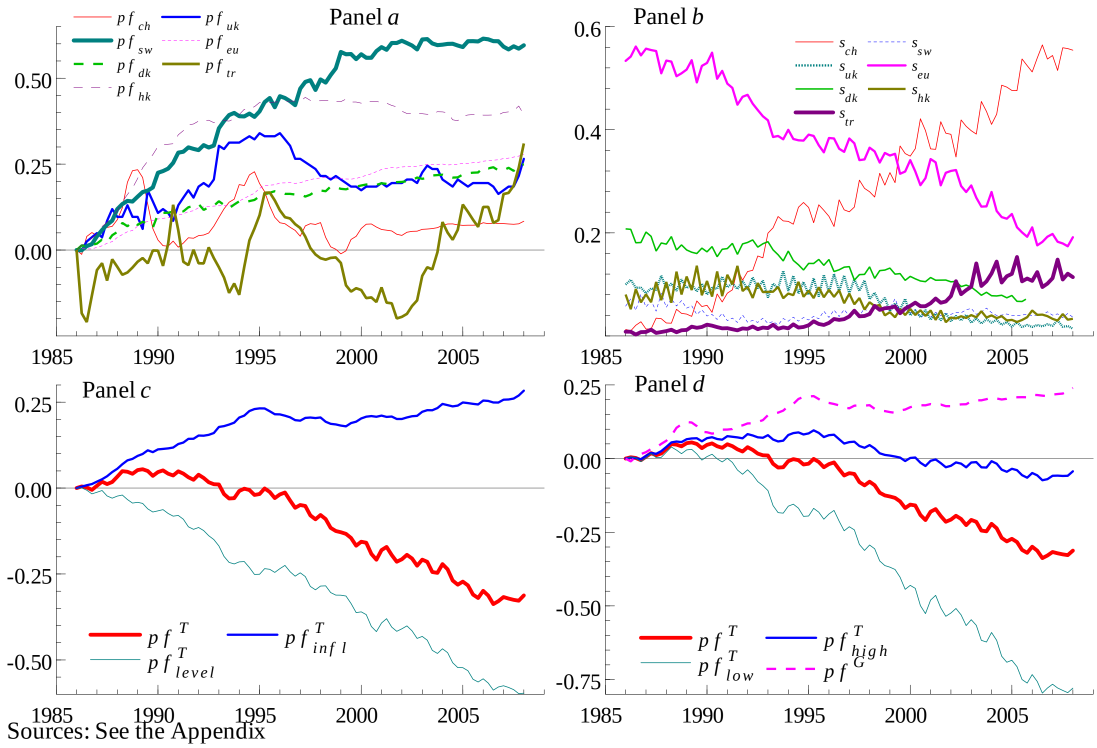
Our calculations show that the price level of Chinese products are between 30 per cent and 45 per cent of the price level on products in the high-cost countries Denmark, Sweden, United Kingdom, the euro area and Hong Kong. The corresponding figures for Turkish products are between 45 per cent and 60 per cent. Hence, both China and Turkey stand out as low-cost countries in our study. We recognise that the price levels in Table 1 are good proxies only to the extent that relative price levels on clothing are similar to relative GDP deflator levels across countries, an assumption that need not hold in practise. For instance, it may be the case that exporters of clothing from low-cost countries set their prices somewhat below the competitors' prices to gain market shares. Consequently, the price level of imports from low-cost countries may be higher than that calculated from the purchasing power parity adjusted GDP deflators. If this is indeed the case, the calculated development in pf^T based on the figures in Table 1 will overestimate the true negative price level impulses to the Norwegian economy. The calculated development in pf^T may, on the other hand, overestimate the true international price impulses as consumer prices, which also include markups on domestic costs of distribution not faced by Norwegian importers, approximate Chinese export prices of clothing in the first half of the sample period. We shed some light on the sensitivity of the development in pf^T, and thereby the sensitivity of the estimate of pass-through, when the price levels for China and Turkey in Table 1 are increased and decreased by 50 per cent, other things equal. Nevertheless, the price levels in Table 1 are used as benchmark to calibrate the respective export price indices when plugged into the price level term of (7).

Figure 2 displays the following time series: country specific export prices (pf_j)

[10] Data for purchasing power parity adjusted GDP are not available on a quarterly basis, and only from 1991 onwards for the euro area. Because the calculated price level series appear relatively stable we assume constant price levels equal to the average over the period $1991-2008$.

in panel a, country specific import shares (s_j)[11] in panel b, the Törnqvist price index based measure of foreign prices (pf^T) and its two components, the inflation effect $(pf^T_{\inf l})$ and the price level effect (pf^T_{level}), in panel c, and the geometric mean based measure of foreign prices (pf^G) and the Törnqvist price index based measure of foreign prices based on a 50 per cent increase (pf^T_{high}) and decrease (pf^T_{low}) in the price levels for China and Turkey in panel d. The price indices are normalised to unity in 1986Q1.

Figure 2: Time series for foreign prices and import shares



Sources: See the Appendix

We observe from panel a that the export prices of clothing from high-cost countries increased quite substantially during the first half of the sample period, possibly reflecting high economic growth and steady demand in their export markets. In the wake of the Asian financial crises, which started in Thailand in July 1997, high-cost countries generally faced reduced export possibilities and stronger price competition from the Asian countries with depreciated currencies. The price competition among trading partners was further amplified by increased presence of low-cost countries on international markets following the trade liberalisation after the Uruguay Round. Additionally, imports from China increased when the country joined the WTO in 2001 and the international economic downturn in 2002 gave rise

[11] We notice that the import shares sum to unity.

to reduced export possibilities for most high-cost countries in the successive years. Together these economic features generally may have led exporters of clothing in high-costs countries to lower their markups over costs during the second half of the sample period.

We see from panel b that the import share from China increased remarkably from a few per cent in 1986 to around 55 per cent in 2008. The import share from the euro area fell likewise from around 55 per cent in 1986 to around 20 per cent in 2008. After a substantial increase in the import share from the mid 1990s, Turkey supplied more than 10 per cent of Norwegian imports of clothing in 2008. Whereas the import share from Sweden was relatively stable around 5 per cent throughout the sample period, the import shares from United Kingdom and Denmark dropped by nearly 10 percentage points each during the period $1995-2008$. Hong Kong also experienced a lower import share by 5 percentage points during the same period. Overall, the shift in imports towards low-cost countries at the expence of high-cost countries was evident since the mid 1980s, but was intensified from the early 1990s and even more from around 1995 alongside the removal of the quota restrictions on trade.

We notice from panel c that pf^T shows a substantial fall in international export prices on clothing during the last two decades. According to our calculations, the shift in imports from high-cost to low-cost countries $-$ the China effect (pf^T_{level}) $-$ has on average pushed down international price impulses by around 2 percentage points each year since the early 1990s. During the second half of the 1980s, the price level effect was moderate, reflecting little substitution of imports towards low-cost countries due to strict trade regulations. The international price impulses were, however, pulled upwards and somewhat dominated by inflationary effects $(pf^T_{\inf l})$ up until 1995, before these effects became moderate and even negative in the late 1990s. Paralleling the period of trade liberalisation, the price level effects played a dominating role in the development of pf^T from 1995 onwards. Even though the last quota restriction was abandoned in 1998, the price level effect continued to pull down pf^T during the last decade, which indicates that trade liberalisation may have had long lasting effects on international export prices on clothing. Overall, our calculations indicate that pf^T was about 27 per cent lower in 2008 compared to 1986, which implies on average a yearly decrease of 1.2 percentage points. By way of comparison, Wheeler (2008) found that imports from China had a negative effect on aggregate export price inflation (measured in Sterling) on all products faced by UK importers, increasing gradually from zero in 1996 to one percentage point in 2004.

Finally, we observe from panel d that the development in pf^G is almost identical to the development in $pf^T_{\inf l}$ from panel c, which means that $\bar{s_j}$ in (9) is a good proxy for $0.5\,(s_{j,t-1} + s_{j,t})$ in (7).[12] More important though, as already illustrated in

[12]In fact, the results from the cointegration analysis below are independent of using pf^G or $pf^T_{\inf l}$

panel c, is the substantial differences in the calculated development in pf^T and pf^G. Because the latter fails to take account of the differences in price levels across trading partners it exhibits an overall international price increase of 27 per cent through the sample period rather than a price fall of the same magnitude. We believe that pf^T provides a better indication of the actual international price development faced by Norwegian importers of clothing given the significant change in the import pattern over time. That said, we also notice from panel d that the development in pf^T is rather sensitive to different assumptions made about the price levels for China and Turkey. Whereas a 50 per cent increase in the price levels for China and Turkey in Table 1 makes an international price fall of only 5 per cent (pf^T_{high}) from 1986 to 2008, a 50 per cent decrease in the same price levels produces a price fall of as much as 55 per cent (pf^T_{low}) in the same period. It remains to be seen (Section 4.1), however, whether this sensitivity in pf^T produces a serious sensitivity in the estimates of pass-through and pricing-to-market.

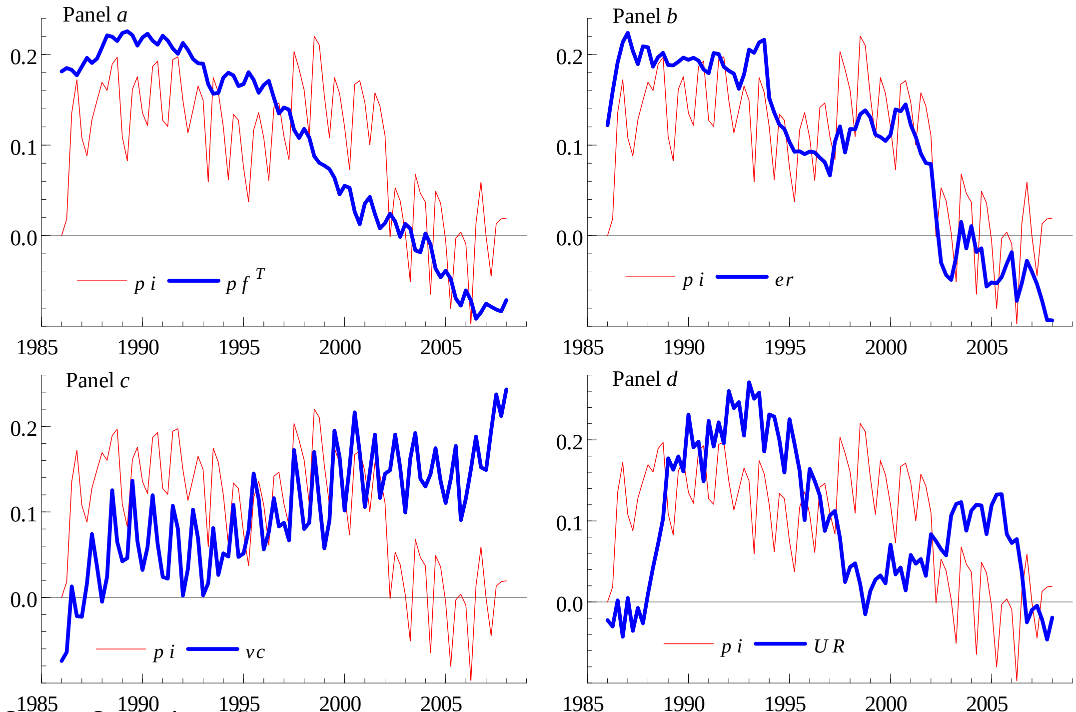
Figure 3 displays the time series for the import price of clothing (pi) together with the Törnqvist price index based measure of foreign prices (pf^T) in panel a, the exchange rate (er) in panel b, the domestic variable unit costs (vc) in panel c, and the unemployment rate (UR) in panel d. The exchange rate series is a chained geometric mean index whose construction parallels that of pf^T in the sense that the bilateral exchange rates between Norway and the seven trading partners are weighted together with their respective (variable) import shares as weights. Domestic variable unit costs are defined as the sum of costs of variable factor inputs relative to total production of clothing and the unemployment rate is measured as the number of unemployed as a fraction of the total labour force (according to the Labour Force Survey). The scale of pf^T, er, vc and UR are adjusted in Figure 3 to match that of pi, which is normalised to unity in 1986Q1.

It is evident that pi, pf^T and er all exhibit a clear downward trend throughout the sample period, whereas vc shows some upward trend. At the same time, import prices of clothing relative to foreign prices measured in Norwegian currency ($pi - pf^T - er$) increased from 1986 to 2008, which may be explained by the fact that variable unit costs relative to import prices ($vc - pi$) also increased in the same period. Although consistent with the pricing-to-market hypothesis, this cannot be the full explanation for the development in pi as ($pi - pf^T - er$) increased somewhat more than ($vc - pi$). As indicated by panel d, the development in pi may also partly be explained by the development in the domestic demand pressure (UR). Specifically, the apparent fall in pi during the first half of the 1990s and during the years between 1999 and 2006 coincides well with increased UR in the same periods. Likewise, the increase in pi during the second half of the 1990s matches rather closely with decreased UR, which suggests that $\delta < 0$ in (5).

That the two price series, the exchange rate series and the series for variable unit costs exhibit some trending behaviour with no apparent mean reversion points

as a proxy for foreign prices, other things equal.

Figure 3: Time series for pi_t, pf_t^T, er_t, vc_t and UR_t



Sources: See the Appendix

to nonstationary time series properties.[13] The unemployment rate, on the other hand, may be stationary by construction. However, we follow Bjørnstad and Nymoen (1999) in the subsequent analysis and treat UR as if it is nonstationary due to autocorrelation in the series. In any case, the price setting rule in (5) seems to be supported by the data.

4 The econometric procedure

Because the pricing-to-market theory predicts the possibility of multiple cointegrating vectors among the variables involved, we employ the Johansen (1995, p. 167) trace test for cointegration rank determination. We thus start with an unrestricted p-dimensional VAR of order k having the form

$$X_t = \sum_{i=1}^{k} \Pi_i X_{t-i} + \mu + \varpi t + \varepsilon_t, \;\; t = k+1, \ldots, T, \tag{10}$$

where X_t is a (p x 1) vector of modelled variables at time t, μ represents a (p x 1) vector of intercepts, ϖ is a (p x 1) coefficient vector of a linear deterministic trend t,

[13] Augmented Dickey Fuller tests indeed suggest that the time series are all $I(1)$.

$\Pi_1, \ldots, \Pi_k$ are (p x p) coefficient matrices of lagged level variables and $\varepsilon_{k+1}, \ldots, \varepsilon_T$ are independent Gaussian variables with expectation zero and (unrestricted) (p x p) covariance matrix Ω. The initial observations of $X_1, \ldots, X_k$ are kept fixed. The question now is how (10) can be reparameterised to a cointegrated VAR (henceforth CVAR) in which the pricing-to-market hypothesis can be formulated as a reduced rank restriction on the impact matrix $\Pi = -(I - \Pi_1 - \ldots - \Pi_k)$.

The way the CVAR is formulated in our context depends on the exogeneity status or otherwise of the unemployment rate series. First, we shall consider the case when the unemployment rate series is endogenous in the system, hence (10) is a fifth-dimensional VAR in $X_t = (pi_t, pf_t^i, er_t, vc_t, UR_t)', i = T, G$. Once $X_t \sim I(1)$, then the first difference $\Delta X_t \sim I(0)$ implying either $\Pi = 0$ or Π has reduced rank such that $\Pi = \alpha\beta'$, where α and β are 5 x r matrices and $0 < r < 5$. Herein r denotes the rank order of Π. Assuming for notational simplicity that $k = 2$, the CVAR in this situation becomes

$$\Delta X_t = \Gamma_1 \Delta X_{t-1} + \alpha\beta' X_{t-1} + \mu + \delta t + \varepsilon_t, \tag{11}$$

where $\beta' X_{t-1}$ is an r x 1 vector of stationary cointegration relations among import prices, foreign prices, exchange rates, variable unit costs and the unemployment rate, and $\Gamma_1 = -\Pi_2$ is the (5 x 5) coefficient matrix of the lagged differenced variables. Next, we shall consider the case when the unemployment rate series is weakly exogenous for the long run parameters such that valid inference on β can be obtained from the fourth-dimensional system describing pi_t, pf_t^i, er_t and vc_t conditional on UR_t without loss of information, see Johansen (1992). Following Harbo *et al.* (1998), we may formulate the *partial* CVAR equivalent to (11) as (again assuming $k = 2$)

$$\Delta X_{1,t} = A_1 \Delta X_{2,t} + \Gamma_{1,1} \Delta X_{t-1} + \alpha_1 \beta' X_{t-1} + \mu_1 + \delta_1 t + \varepsilon_{1,t}, \tag{12}$$

with the corresponding marginal model given by $\Delta X_{2,t} = \Gamma_{1,2} \Delta X_{t-1} + \mu_2 + \delta_2 t + \varepsilon_{2,t}$ when $\{X_t\} = \{X_{1,t}, X_{2,t}\}$, $X_{1,t} = (pi_t, pf_t^i, er_t, vc_t)'$ and $X_{2,t} = UR_t$. It follows that the unemployment rate is included in the long-run part of (12) as a non-modelled variable. Because the number of relevant variables to be included in (10), and hence also the number of parameters to be estimated, is large relative to the number of observations in the available data set it would be useful to impose weak exogeneity on the unemployment rate. However, to know whether β can be estimated from (12) we first estimate the full system in (11) and test formally rather than assume the weak exogeneity status of the unemployment rate in that system.[14] We follow common practice and let inference about the rank of Π from the full system be based on unrestricted intercepts and a restricted linear trend. Likewise, dummies capturing seasonality in the data ($S1_t$, $S2_t$ and $S3_t$) enter the system unrestrictedly.

[14] Naug and Nymoen (1996) condition on the unemployment rate being stationary (with possible structural breaks) without testing formally for its exogeneity status.

Strictly speaking, the cointegration rank need not be determined from the partial system once it has been determined from the full system. Nevertheless, we re-determine the cointegration rank from (12) for the sake of comparison with the rank determination from (11). However, as noted by Harbo *et al.* (1998), the asymtotic distribution of the trace test statistic is influenced by conditioning on weakly exogenous variables and standard critical values are thus not valid. We therefore use the critical values in Table 2 in Harbo *et al.* (1998). Also, following the suggestions in Harbo *et al.* (1998) for partial systems, we restrict the linear trend to lie in the cointegration space for inference purposes only. Then, after having determined the rank order, we test whether the linear trend can be dropped from the cointegration relation (s) by a conventional χ^2-test. As in the full system, both the intercepts and the seasonals enter the partial system unrestrictedly. We now turn to the empirical findings from the cointegration analysis based on the econometric procedure outlined above.

4.1 Cointegration analysis based on PF_t^T

Irrespective of specifying a full fifth-dimensional VAR in $X_t = (pi_t, pf_t^T, er_t, vc_t, UR_t)'$ or a partial fourth-dimensional VAR in $X_{1,t} = (pi_t, pf_t^T, er_t, vc_t)'$ conditional on $X_{2,t} = UR_t$ being exogenous to the system, we find that $k = 3$ produces a model with no serious misspecification as indicated by standard diagnostic tests. Certainly, the estimated residuals of the UR_t-equation in the full system and thus also the estimated vector residuals are borderline cases (at conventional significance levels) with respect to suffering from autocorrelation. Such a potential problem may in itself be an argument for moving to a partial system to obtain even more satisfying residual properties in our case, see Juselius (2006, p. 198). Noticeably, no impulse dummies are required to mop up any outliers to obtain Gaussian residuals.[15] Table 2 reports trace test statistics for the sample period $1986Q4 - 2008Q1$, both in the case of the full system and the partial system with the Törnqvist price index based measure of foreign prices assuming $k = 3$.

We notice that the null hypothesis of no cointegration can be rejected at the 5 per cent significance level, whereas the hypothesis of at most one cointegrating relationship between import prices, foreign prices, exchange rates, domestic variable unit costs and demand pressure (proxied by the unemployment rate) cannot be rejected within the full CVAR. As shown below, choosing $r = 1$ gives a cointegrating vector with interpretable properties in line with the pricing-to-market hypothesis.

[15] A VAR of order 2 produces severe autocorrelation in the vector residuals and in the residuals of the pf_t^T-equation and the UR_t-equation of the full system. Results from the diagnostic tests of the VARs and other test results not reported, here and below, are available from the authors upon request. Naug and Nymoen(1996) include a set of dummy variables to account for outliers and structural breaks in the VAR. As noted by Franses and Lucas (1998), standard cointegration tests are sensitive to atypical events such as outliers and structural breaks.

Table 2: Tests for cointegration rank based on PF_t^T

Full CVAR system				Partial CVAR system			
r	λ_i	λ_{trace}	p-val$_{PcGive}$	r	λ_i	λ_{trace}	$5\%_{Harbo}$
$r = 0$	0.41	107.66	0.001	$r = 0$	0.40	94.57	71.7
$r \leq 1$	0.28	62.90	0.058	$r \leq 1$	0.24	50.72	49.6
$r \leq 2$	0.18	34.87	0.254	$r \leq 2$	0.18	26.99	30.5
$r \leq 3$	0.12	18.00	0.351	$r \leq 3$	0.11	10.36	15.2
$r \leq 4$	0.08	6.82	0.374				

Notes: Sample period: $1986Q4 - 2008Q1$. The underlying VARs are of order 3. The full CVAR consists of $X_t = (pi_t, pf_t^T, er_t, vc_t, UR_t)'$, whereas the partial CVAR consists of $X_{1,t} = (pi_t, pf_t^T, er_t, vc_t)'$ being endogenous and $X_{2,t} = UR_t$ being exogenous. Both systems include unrestricted constants and seasonals and a restricted linear trend. r denotes the cointegration rank, λ_i are the eigenvalues from the reduced rank regressions, λ_{trace} are the trace test statistics, p-val$_{PcGive}$ are the significance probabilities from PcGive and $5\%_{Harbo}$ are the critical values (5 per cent significance level) from Table 2 in Harbo *et al.* (1998).

Testing a zero restriction on the equilibrium correction coefficient of the unemployment rate under the assumption of $r = 1$ gives $\chi^2_{(1)} = 0.91$ with a p-value of 0.34. Hence, UR_t may be considered as weakly exogenous for the cointegrating parameters, whose estimates can then be efficiently estimated from the partial rather than the full system without loss of information. In so doing, we also obtain a more feasible VAR and save degrees of freedom. The formal tests in Table 2 support the hypothesis that $r = 1$ also in the case of the partial CVAR, albeit a borderline case at the 5 per cent significance level. Likelihood ratio tests (not shown) clearly reject the hypothesis that the modelled variables in $X_{1,t} = (pi_t, pf_t^T, er_t, vc_t)'$ as well as $X_{2,t} = UR_t$ are long run excludable from β with rank equal to unity. The linear trend, however, is strongly *in*significant with $\chi^2_{(1)} = 0.523$ and a p-value of 0.47. It is therefore excluded from the model in the following likelihood ratio tests about the pricing-to-market hypothesis, that is tests about α_1 and β in (12) assuming $r = 1$. Table 3 summarises results from these tests.

Table 3: Tests of the pricing-to-market hypothesis based on PF_t^T

Hypothesis	LR tests	p-value
H_1: $\alpha_{1(pi_t)} = 0$	$\chi^2_{(2)} = 20.34$	0.000
H_2: $\alpha_{1(pf_t^T)} = 0$	$\chi^2_{(2)} = 18.14$	0.001
H_3: $\alpha_{1(er_t)} = 0$	$\chi^2_{(2)} = 2.08$	0.354
H_4: $\alpha_{1(vc_t)} = 0$	$\chi^2_{(2)} = 0.53$	0.766
H_5: $(pi_t - \psi_1 pf_t^T - \psi_1 er_t - \psi_2 vc_t) \sim I(0)$	$\chi^2_{(2)} = 2.01$	0.367
H_6: $[pi_t - (1-\psi)(pf_t^T + er_t) - \psi vc_t] \sim I(0)$	$\chi^2_{(3)} = 2.07$	0.558
H_7: $(pi_t - pf_t^T - er_t) \sim I(0), \beta_{(vc_t)} = 0$	$\chi^2_{(4)} = 29.81$	0.000
H_8: $(vc_t - pf_t^T - er_t) \sim I(0), \beta_{(pi_t)} = 0$	$\chi^2_{(4)} = 33.71$	0.000

Notes: Sample period: $1986Q4 - 2008Q1$. All likelihood ratio (LR) tests are based on the partial CVAR with $r = 1$ and $\beta_{(trend)} = 0$ and with degrees of freedom in parenthesis.

First, we observe that weak exogeneity of both import prices and foreign prices for the long run parameters is strongly rejected. By way of contrast, we may assume that exchange rates and domestic costs are weakly exogenous. The hypotheses of identical parameters of foreign prices and exchange rates (H_5) and of long run homogeneity as an additional restriction (H_6) are both accepted by the data. On the other hand, the hypotheses of long run versions of LOP (H_7) and PPP (H_8), as defined in Section 3, are clearly rejected by the data. Finally, imposing equal parameters of pf_t^T and er_t, long run homogeneity and weak exogeneity of er_t and vc_t yields $\chi^2_{(5)} = 6.21$ with a p-value of 0.29. Hence, we obtain the following restricted cointegrating vector (normalised on import prices)

$$pi_t = const. + 0.444 pf_t^T + \underset{(0.016)}{0.444} er_t + 0.556 vc_t - \underset{(0.334)}{2.009} UR_t, \tag{13}$$

with standard errors in parentheses. The associated vector of equilibrium correction coefficients is estimated to $\hat{\alpha}_1 = (-0.44, -0.18, 0, 0)'$. Because any deviations from (13), due to say a shock in the exchange rate, are mainly and significantly corrected through the adjustment of import prices we regard the estimated cointegrating vector as a long run import price equation for clothing consistent with the pricing-to-market hypothesis.[16] The pass-through and pricing-to-market elasticities are significantly estimated to 0.45 and 0.55, respectively. Also, the estimated import price equation includes strong and significant negative effects of the unemployment rate. According to the point estimate in (13), import prices will decrease by 2 per cent in the long run following a one *percentage point* increase in the unemployment rate, other things equal.[17] As such, decreases in domestic demand pressure (proxied by increases in the unemployment rate) cause prices of imports to fall.

Interestingly, Naug and Nymoen (1996) found the pass-through elasticity to be 0.63 based on data for Norwegian imports of total manufactures over the sample period $1970Q1 - 1991Q4$. As price setting behaviour typically varies across products and presence of non-tariff barriers to trade is not controlled for the estimate of pass-through in Naug and Nymoen (1996) is likely to be biased.[18] Our estimate of pass-through also differs somewhat from those found by Menon (1996) based on disaggregated Australian data over the sample period $1981Q3 - 1992Q2$. In that study, the estimates in most cases indicate incomplete pass-through, but with substantial variation across products. Particularly, pass-through is estimated to be less than 30 per cent for some of the quota protected textiles and wearing apparels studied. Menon (1996) partly views this finding in light of the Bhagwati hypothesis as

[16] Although significantly estimated, the adjustment coefficient for pf_t^T is only 40 per cent of that for pi_t.

[17] We notice that the point estimate of the unemployment rate in (13) is a semielasticity because UR_t is defined as a *rate* variable (with no logarithmic transformation).

[18] On average, imports of clothing constituted for around 10 per cent of total imports of manufactures during the 1970s and 1980s, see http://statbank.ssb.no/statistikkbanken/.

significant negative effects from a quantity restriction variable are among the most convincing results. That is, exchange rate changes have to some extent been prevented from being fully passed-through to import prices by the import premium associated with quotas in the Australian context. Our estimate of pass-through, which is somewhat higher than comparable estimates in Menon (1996), may also be viewed in light of the Bhagwati hypothesis. As pointed out above, the hypothesis implies increased pass-through when non-tariff barriers to trade are gradually removed, other things equal. However, once the China effect is included in the measure of foreign prices as in pf^T, it is likely that pass-through has not changed dramatically since the mid 1990s. Indeed, recursive estimates of the pass-through coefficient in (13) are reasonably stable in the years after 1995. Also, recursively estimated $\chi^2_{(5)}$ indicate that the restrictions in (13) are supported by the data throughout the second half of the sample period.

We complete the cointegration analysis based on pf^T by examining potential sensitivity in the estimate of pass-through based on different assumptions made about the price levels for China and Turkey. As already revealed from Figure 2 (panel d), the calculated development in pf^T is somewhat sensitive to a 50 per cent increase (pf^T_{high}) and decrease (pf^T_{low}) in the price levels for China and Turkey in Table 1. We obtain the following estimated cointegrating vectors with pf^T_{high} and pf^T_{low} replacing pf^T, all other modelling issues equal:

$$pi_t = const. + 0.604 pf^T_{high,t} + \underset{(0.019)}{0.604} er_t + 0.396 vc_t - \underset{(0.280)}{1.949} UR_t, \tag{14}$$

$$pi_t = const. + 0.306 pf^T_{low,t} + \underset{(0.015)}{0.306} er_t + 0.694 vc_t - \underset{(0.469)}{2.285} UR_t. \tag{15}$$

Similar to (13), we have imposed equal parameters of $pf^T_{i,t}$ ($i = high, low$) and er_t, long run homogeneity and weak exogeneity of er_t and vc_t in (14) and (15), which yields $\chi^2_{(5)} = 1.55$ and $\chi^2_{(5)} = 9.01$ with p-values of 0.91 and 0.11, respectively. We observe that the estimates of pass-through, and hence also the estimates of pricing-to-market, do not depend critically on the assumptions made about the price levels for China and Turkey. The estimate of pass-through increases and decreases by 33 per cent when pf^T_{high} and pf^T_{low} replace pf^T, which we consider as a rather moderate sensitivity in the estimate given the rather substantial magnitude of the shift in the price levels. We shed some further light on the sensitivity in the estimates of pass-through and pricing-to-market in addition to the potential problem of omitted variable bias in the subsequent cointegration analysis based on pf^G rather than pf^T used herein.

4.2 Cointegration analysis based on PF^G_t

As with the Törnqvist price index based measure of foreign prices, a lag length of three is sufficient to render residuals with no serious misspecification, neither in the full nor in the partial VAR. We remark once again that no impulse dummies are needed to achieve Gaussian residuals in the VARs. Table 4 reports trace test statistics based on the VARs of order three with the geometric mean based measure of foreign prices.

Table 4: Tests for cointegration rank based on PF_t^G

Full CVAR system				Partial CVAR system			
r	λ_i	λ_{trace}	p-val$_{PcGive}$	r	λ_i	λ_{trace}	$5\%_{Harbo}$
$r = 0$	0.32	89.51	0.042	$r = 0$	0.31	72.14	71.7
$r \leq 1$	0.23	56.95	0.166	$r \leq 1$	0.18	39.58	49.6
$r \leq 2$	0.15	34.60	0.266	$r \leq 2$	0.15	22.88	30.5
$r \leq 3$	0.15	20.24	0.218	$r \leq 3$	0.10	8.67	15.2
$r \leq 4$	0.07	6.06	0.464				

Notes: Sample period: $1986Q4 - 2008Q1$. The underlying VARs are of order 3. The full CVAR consists of $X_t = (pi_t, pf_t^G, er_t, vc_t, UR_t)'$, whereas the partial CVAR consists of $X_{1,t} = (pi_t, pf_t^G, er_t, vc_t)'$ being endogenous and $X_{2,t} = UR_t$ being exogenous. Both systems include unrestricted constants and seasonals and a restricted linear trend. r denotes the cointegration rank, λ_i are the eigenvalues from the reduced rank regressions, λ_{trace} are the trace test statistics, p-val$_{PcGive}$ are the significance probabilities from PcGive and $5\%_{Harbo}$ are the critical values (5 per cent significance level) from Table 2 in Harbo *et al.* (1998).

Again, we notice that the rank should be set to unity in the case of the full CVAR system at the 5 per cent significance level. Also, the unemployment rate is weakly exogenous for the long run parameters in that system under the assumption of $r = 1$, as indicated by $\chi^2_{(1)} = 0.002$ with a p-value of 0.97. Accordingly, we may once again move to further inference about the α and β matrices relying on the partial CVAR. The formal tests in Table 4 also indicate existence of a unique cointegration relationship with the partial system. Besides, the hypothesis of long run exclusion of pi_t, pf_t^G, er_t, vc_t and UR_t from β (with $r = 1$) is not supported by the data, a finding in line with the analysis above using the Törnqvist price index based measure of foreign prices. However, the linear trend is now needed in the cointegration space and cannot be omitted from the long run relation according to $\chi^2_{(1)} = 8.12$ and its p-value of 0.004. Consequently, it is *not* excluded from the reduced rank partial VAR underlying the tests about the pricing-to-market hypothesis reported in Table 5.

Overall, the test results in Table 5 are similar to those in Table 3. We remark though that H_2 now is not rejected by the data, indicating that pf_t^G just like er_t and vc_t is exogenous for the parameters of interest. Hence, imposing the hypotheses H_2, H_3, H_4 and H_6 gives $\chi^2_{(5)} = 4.28$ and a p-value of 0.51, and the following restricted cointegrating vector (normalised on import prices)

$$pi_t = const. + 0.601 pf_t^G + \underset{(0.101)}{0.601} er_t + 0.399 vc_t - \underset{(0.45)}{1.95} UR_t - \underset{(0.00047)}{0.00207} t, \tag{16}$$

with standard errors in parenthesis. The adjustment coefficient of import prices is now significantly estimated to -0.42, which is almost identical to the corresponding

Table 5: Tests of the pricing-to-market hypothesis based on PF_t^G

Hypothesis	LR tests	p-value
H_1: $\alpha_{1(pi_t)} = 0$	$\chi^2_{(1)} = 11.79$	0.001
H_2: $\alpha_{1(pf_t^G)} = 0$	$\chi^2_{(1)} = 2.48$	0.115
H_3: $\alpha_{1(er_t)} = 0$	$\chi^2_{(1)} = 0.09$	0.766
H_4: $\alpha_{1(vc_t)} = 0$	$\chi^2_{(1)} = 0.47$	0.492
H_5: $(pi_t - \psi_1 pf_t^G - \psi_1 er_t - \psi_2 vc_t) \sim I(0)$	$\chi^2_{(1)} = 1.34$	0.246
H_6: $[pi_t - (1-\psi)(pf_t^G + er_t) - \psi vc_t] \sim I(0)$	$\chi^2_{(2)} = 2.14$	0.343
H_7: $(pi_t - pf_t^G - er_t) \sim I(0), \beta_{(vc_t)} = 0$	$\chi^2_{(3)} = 11.25$	0.011
H_8: $(vc_t - pf_t^G - er_t) \sim I(0), \beta_{(pi_t)} = 0$	$\chi^2_{(3)} = 18.15$	0.000

Notes: Sample period: $1986Q4 - 2008Q1$. All likelihood ratio (LR) tests are based on the partial CVAR with $r = 1$ and with degrees of freedom in parenthesis.

estimate obtained with the Törnqvist price index based measure of foreign prices. More important though, when comparing (13) and (16), are the somewhat different estimates of long run pass-through and pricing-to-market that come out of the modelling with the two alternative measures of foreign prices. Another important difference between the two estimated cointegrating vectors is the linear trend, which enters significantly in (16) and *not* in (13).

One possible interpretation of these findings is that the effects of shifts in imports from high-cost to low-cost countries on internationally traded good prices (and thereby on the degree of pass-through) are likely to be controlled for through the linear trend in (16), effects which are *not* explicitly picked up by pf_t^G alone. As seen from Figure 2 (panel c), the calculated price level term of pf_t^T (the China effect) drifts downwards during the entire sample period and may accordingly behave like a deterministic linear trend in a regression model. Indeed, the linear trend enters significantly in (16) with a negative sign consistent with the a priori beliefs about the China effect on internationally traded goods prices. The estimate implies that the shift in imports towards low-cost countries has depressed import prices of clothing by around 0.8 percentage points yearly $(-0.00207 \cdot 400)$ since 1986, approximately equal to the yearly average of around 0.9 calculated by means of the pass-through estimate of 0.45 from (13) and the 2 percentage points yearly decrease in the price level term pf_{level}^T. However, the fact that pf_{level}^T exhibits some non-linearities with apparent cycles, especially around the years of 1995 and 2000, may make a linear trend a poor proxy for the true China effect on international price impulses faced by Norwegian importers as such.[19] For this reason, we suspect the estimates of pass-through and pricing-to-market in (16) to be somewhat biased, albeit not that

[19] Bache and Naug (2007) approximate effects of shift in imports from high-cost to low cost countries by detrending the variables prior to estimation through a linear trend. They acknowledge though that a linear trend may be a poor proxy as it implicitly assumes constant effects of trade liberalisation over the entire sample period.

critically, compared to those in (13). We also find that pass-through is more or less complete and thus that pricing-to-market behaviour is absent when the trend variable is dropped from (16), results which are very unlikely given the facts about the clothing industry outlined in Section 2. Although not accepted by the data[20], these findings point to the likely problem of an omitted variable bias if shifts in imports towards low-cost counries and trade liberalisation effects are *not* explicitly controlled for in the import price equation for clothing.

We have seen that using a Törnqvist price index based measure of foreign prices is a flexible approach that may overcome this potential econometric problem in our empirical case, all other modelling issues equal. Based on our findings, we also believe it is a more reliable approach than using a geometric mean based measure of foreign prices together with a linear trend (which from the outset is a strict assumption) in the regression model to quantify pass-through consistently.

5 A dynamic import price model

As noted in the introduction, the degree of pass-through may just like trade policy be linked to monetary policy and the nature and magnitude of exchange rate changes. According to Froot and Klemperer (1989), foreign firms are likely to price more aggressively in the domestic market to gain higher market shares when the currency of the importing country is expected to be permanently stronger. Conversely, when a currency appreciation is believed to be temporary, foreign firms will behave less aggressively in their price setting. To the extent that increased exchange rate volatility has led foreign exporters to believe the appreciations after 2001 to be more permanent in nature than previous appreciations, we should expect that pass-through has become significantly higher following the introduction of inflation targeting. We test this hypothesis in the following by examining stability properties (or lack thereof) of an estimated dynamic equilibrium correction model (henceforth $EqCM$) consistent with the cointegration findings in the previous section.[21] Our point of departure is a *general* $EqCM$ model (with the constant, the seasonals and the same lag length used in the reduced rank partial VAR) written as

$$
\begin{aligned}
(17)\ \Delta pi_t \quad = \quad & const. + \sum_{i=1}^{2} \varphi_{1,i} \Delta pi_{t-i} + \sum_{i=0}^{2} \varphi_{2,i} \Delta (pf^T + er)_{t-i} + \sum_{i=0}^{2} \varphi_{3,i} \Delta vc_{t-i} \\
& + \sum_{i=0}^{2} \varphi_{4,i} \Delta UR_{t-i} + \eta EqCM_{t-1} + \eta_1 S1_t + \eta_2 S2_t + \eta_3 S3_t.
\end{aligned}
$$

The general model contains impact effects and two lags of the first difference (denoted Δ) of vc_t, the sum of pf_t^T and er_t, and UR_t. We notice that $\Delta(pf^T + er)_t$

[20] The $\chi^2_{(6)} = 20.51$ with a p-value of 0.002.

[21] Stability of the pass-through relationship is also studied in the literature in light of theories of hysteresis in import prices, see e.g. Athukorala and Menon (1995).

is denominated in the Norwegian currency, a short run restriction imposed from the outset in line with the same long run restriction (i.e., equal parameters of pf_t^T and er_t) accepted by the data. Also, the first difference of pi_t is included in (17) with two lags, whereas the $EqCM$ term [which is defined in accordance with (13)] is lagged one period. Simplifications from the general to the specific model is performed using PcGets, see Hendry and Krolzig (2001). PcGets picks the following *specific* model in our case together with diagnostic tests[22] and the estimated standard errors below the point estimates (sample period: $1986Q4 - 2008Q1$):

$$
\begin{aligned}
(18)\Delta pi_t \quad = \quad & -\underset{(0.055)}{0.336}\Delta pi_{t-1} + \underset{(0.088)}{0.479}\Delta(pf^T + er)_t \\
& -\underset{(0.619)}{2.102}\Delta UR_t - \underset{(0.627)}{2.032}\Delta UR_{t-1} \\
& -\underset{(0.067)}{0.385}[pi_{t-1} - 0.44(pf^T + er)_{t-1} - 0.56vc_{t-3} + 2.01UR_{t-2}] \\
& +\underset{(0.024)}{0.188} - \underset{(0.009)}{0.065}S1_t - \underset{(0.007)}{0.109}S2_t
\end{aligned}
$$

Diagnostic tests:
AR_{1-5}: $F(5, 73) = 1.58$ [0.18], $ARCH_{1-4}$: $F(4, 70) = 1.82$ [0.14],
$NORM$: $\chi^2_{(2)} = 4.97$ [0.08], HET: $F(12, 65) = 1.59$ [0.12],
$RESET$: $F(1, 77) = 1.98$ [0.16].

Several interesting aspects are inherent in the specific model. First, the equilibrium correction term enters significantly in (18), whose estimated coefficient of -0.39 implies rapid adjustment of import prices of clothing towards the long run equilibrium level in the event of a shock in either foreign prices, exchange rates, domestic costs or demand pressure. We observe that the equilibrium correction term is specified with three and two lags on domestic costs and demand pressure, respectively, a reparameterisation that turned out useful to obtain reasonable short run dynamic properties. The estimated short run pass-through elasticity is somewhat greater than its long run counterpart.[23] Accordingly, import prices respond quickly and with some overshooting to shocks in either foreign prices (denominated in foreign currency) or exchange rates. However, the specific model also contains

[22] AR_{1-5} is Harvey's (1981) test for until 5th order residual autocorrelation; $ARCH_{1-4}$ is the Engle (1982) test for until 4th order autoregressive conditional heteroskedasticity in the residuals; $NORM$ is the normality test outlined in Doornik and Hansen (1994), HET is a test for residual heteroskedasticity due to White (1980) and $RESET$ is the Ramsey (1969) test for functional form misspecification. The numbers in square brackets are p-values.

[23] Both Δpf_t^T and Δer_t enter *in*significantly as separate explanatory variables in (18), a finding which supports the hypothesis of equal short run impact effects on import prices from changes in these variables. Moreover, the *residuals* from the equations for Δpf_t^T and Δer_t in the partial VAR are not significant when added to (18). Hence, Δpf_t^T and Δer_t may be regarded as weakly exogenous for the short run parameters in the specific model, whose estimates are consistently estimated by OLS, see Urbain (1992).

significant and negative short run autoregressive effects from Δpi_{t-1}, which make the adjustment process of import prices somewhat less smooth. That is, the first quarter adjustment of import prices following a shock in say the exchange rate is corrected somewhat in the opposite direction in the next quarter due to the autoregressive effects before the adjustment process continues steadily towards the long run equilibrium level. Altogether, pass-through is almost complete within one to three quarters according to (18). The rather fast speed of adjustment of import prices identified here may reflect the fact that the exchange rate was fairly volatile during most of the sample period, cf. Figure 3 (panel *b*). If there are costs related to changing import prices, it will be rational to respond relatively fast to large fluctuations in the exchange rate that are *not* likely to be reversed in the near future. We notice further from (18) that ΔUR_t and ΔUR_{t-1} enter the model with more or less identical effects on import prices, effects which also are almost identical to the long run counterpart. Hence, foreign firms seem to absorb quickly, but with some smoothing, into their markups changes in the unemployment rate which are normally of a somewhat permanent nature.

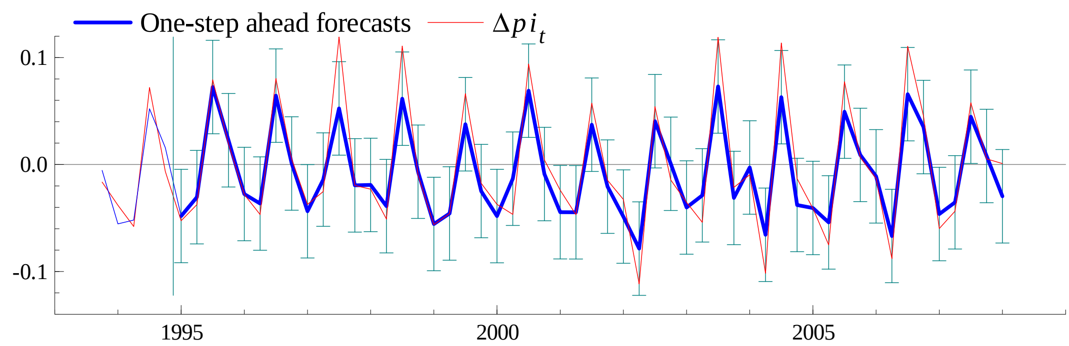
Our estimate of the speed of adjustment of import prices following a shift in the exchange rate accords with Menon (1996), who finds that pass-through is complete within two quarters for most products in the Australian context. However, Naug and Nymoen (1996) find relatively slow speed of adjustment of import prices, which may partly be viewed in light of a sample period where monetary policy was that of a fixed exchange rate regime. Small exchange rate fluctuations during that period (cf. Figure 3, panel *b*) may thus have been understood as transitory by foreign firms, in which case it may have been rational to respond relatively slow, if at all.

Turning to stability properties of the specific model, we first notice that the model shows no sign of misspecification as reported below (18). This model property is further confirmed by recursive break point Chow statistics and recursively estimated coefficients, which provide evidence of reasonable constancy from the early 1990s. We now ask whether the model is able to predict import prices of clothing out-of-sample to shed some more light on its robustness with respect to the trade and monetary policy regime changes during the period $1995-1998$ and in late March 2001, respectively. If pass-through indeed has changed in the wake of these policy shifts, we should expect instabilities in the estimated model as indicated by poor out-of-sample forecasting ability. To this end, we shall use simple one-step ahead forecasts by reestimating (18) based on observations until $1994Q4$, and leaving 53 quarters ($1995Q1-2008Q1$) for out-of-sample forecasts.[24] Figure 4 depicts actual values of Δpi_t together with its one-step ahead forecasts and 95 per cent confidence intervals to each forecast in the forecasting period (shown by the vertical error bars of $\pm$2SE).

We observe that the forecasts only misses significantly the observed values of

[24] We acknowledge that the reestimated model will be based on the same model design that was used in the estimation with the full sample period ending in $2008Q1$.

Figure 4: Actual values and one-step ahead forecasts of Δpi_t



Δpi_t once, namely in the third quarter of 1997. The point in time of the forecasting failure does coincide with the time period in which a majority of the quota restrictions on trade had already been abolished. However, the particular forecasting failure may very well be explained by the Asian financial crises rather than the shift in trade policy itself. As seen from Figure 3 (panel b), the import prices on clothing increased alongside the strong appreciation pressure of the Norwegian currency following the crises. In any case, the fact that 15 out of 16 forecasts during the period $1995 - 1998$ are inside the confidence intervals (albeit $1998Q3$ is a borderline case) points to pass-through being constant throughout the trade liberalisation period. Also, a Chow test statistic of parameter constancy between the sample and the forecasting periods, see Hendry and Doornik (2001, p. 241), is far from being significant, as indicated by $F[53,25] = 0.53$ and its p-value of 0.98. Moreover, the reestimated model is close to the one in (18) with respect to parameter estimates and diagnostics. We therefore conclude that the out-of-sample forecasting ability of the estimated import price model is satisfactory despite *major* regime shifts in both trade and monetary policies during the forecasting period. That no serious forecasting failures are detected during the second half of the 1990s may reflect that possible effects on pass-through of changes in the trade policies are controlled for through pf^T, effects which may otherwise be reflected in unstable estimates of the model. Because the model exhibits no forecasting failures around the date of $2001Q1$ suggests that pass-through has remained unchanged despite the introduction of inflation targeting. These findings may be explained by the fact that foreign firms also experienced relatively high exchange rate volatility during the 1990s, cf. Figure 3 (panel b). After leaving the fixed exhange rate system in 1992 in favour of a managed floating regime, the exchange rate behaved more like free float following several episodes of speculative attacks against the Norwegian currency. It is therefore not surprising if foreign firms beliefs about the permanent nature of (large) exchange rate fluctuations did already change, if at all, during the period of the managed floating regime. We have established, however, that the estimated import price model is stable also throughout the 1990s, which contradicts such a hypothesis.[25]

6 Conclusions

Economic theory predicts that presence of non-tariff barriers to trade is potentially important when quantifying the degree of pass-through to traded goods prices. In this paper, we have applied the cointegrated VAR approach and estimated a pricing-to-market model for Norwegian import prices of clothing over the period $1986-2008$, controlling explicitly for potential pass-through effects of the gradual removal of non-tariff barriers to trade and the switch in imports from high-cost to low-cost countries. The novelty of the paper, we believe, is that the measure of foreign prices as one of the explanatory variables in the model is based on the total differentiation of the Törnqvist price index. As such, we have allowed not only for inflationary differences as is common in previous studies, but also varying import shares and differences in price levels (known as the China effect) among trading partners when constructing the measure of foreign prices.

One important finding in this paper is that the China effect on traded goods prices is substantial in the clothing industry. Our calculations suggest that the shift in imports from high-cost to low-cost countries since the early 1990s on average has reduced the international price impulses on clothing imports by around 2 percentage points per year. With the Törnqvist price index based measure of foreign prices we established an import price model for clothing consistent with the pricing-to-market hypothesis. Specifically, we found the pass-through and pricing-to-market elasticities to be 0.45 and 0.55, respectively. We also found that these estimates are reasonably stable, which contradicts the implications of the Bhagwati hypothesis that gradual removal of non-tariff barriers to trade has pushed pass-through upwards, other things equal. That is, once the China effect is controlled for through our measure of foreign prices, we find little evidence that the long run slopes of the import price model have changed significantly alongside trade liberalisation. By way of contrast, we found that relying on the alternative measure of foreign prices based on the often used geometric mean price index with constant import shares as weights may overestimate international price impulses and thereby produce biased estimates of pass-through and pricing-to-market. These findings thus point to the potential problem of omitted variable bias in our empirical case if the China effect is *not* explicitly controlled for in the regression model. Of course, we may approximate the China effect through a linear trend in the model together with the geometric mean based measure of foreign prices. However, we showed that such a model is still likely

[25] We also controlled for any instabilities in the estimated model by means of the outlier detection procedure available in PcGets. It turned out that no significant marginal outliers were detected by this procedure during the periods of trade liberalisation and shift in monetary policy.

to produce some biasedness in the estimates of pass-through and pricing-to-market, albeit not critically. Because the China effect exhibits some non-linearities during the sample period, we believe that the Törnqvist based measure of foreign prices is a more flexible and reliable approach than the linear trend, which implicitly assumes that the China effect has been constant throughout the sample period. We further established that the dynamic part of the estimated import price model is reasonably stable in-sample. Finally, a forecasting exercise on the estimated dynamic model does not lend much support to the hypothesis that pass-through has changed in the wake of the trade policy shifts during the second half of the 1990s and the monetary policy regime shift in 2001.

We acknowledge though that the empirical analysis is based on an operational measure of foreign prices which is encumbered with some uncertainty. As emphasised in the text, data on comparable price levels on clothing are not available. We have therefore relied on the purchasing power parity adjusted GDP deflator in each country as a basis for calibrating each export price index of clothing when calculating the China effect over time. The available data are also somewhat inadequate when it comes to China in particular, whose export prices of clothing are proxied by consumer prices in the first part of the sample period. That said, robustness analysis revealed that these limitations in the data do not cause much sensitivity in the estimates of pass-through and pricing-to-market. Also, the finding that the estimated dynamic model is well specified throughout the sample period points further to measurment errors being a minor problem in the present context. However, an issue not addressed in this paper is the potential role for expectational dynamics arising from foreign firms being forward-looking in their price setting behaviour. If foreign firms indeed are forward-looking, the coefficients in the regression models considered herein will depend not only on the parameters in the price setting rule, but also on the parameters in the expectations mechanism. Estimating a New Keynesian import price model for clothing by means of likelihood based methods in the spirit of Boug *et al.* (2006, 2010) is left for future work.

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Appendix

PI: Chained geometric Paasche price index for imports of clothing (cif) measured in the Norwegian currency. $1986Q1 = 1$. Source: Statistics Norway, the Quarterly National Accounts (QNA).

PF^T: Törnqvist price index based measure of export prices of clothing measured in foreign currency. $1986Q1 = 1$. The index is based on import shares, export price indices and GDP deflators of the main Norwegian trading partners of clothing (China, the Euro area, UK, Sweden, Denmark, Hong Kong and Turkey), cf. equation (7) in the text.

PF^G: Geometric mean price index based measure of export prices of clothing measured in foreign currency, cf. equation (9) in the text. $1986Q1 = 1$.

PF_{ch}: China: Producer price index of clothing (from $1997Q1$) and consumer price index all products (from $1996Q1$) measured in the Chinese currency. Source: Reuters EcoWin.

PF_{eu}: The Euro area: Producer price index of clothing measured in EURO. Source: Reuters EcoWin.

PF_{uk}: United Kingdom: Export price index of clothing measured in the UK currency. Source: National statistics online, http://www.statistics.gov.uk/statbase/.

PF_{sw}: Sweden: Export price index of clothing measured in the Swedish currency. Source: National statistics online, http://www.ssd.scb.se/databaser/.

PF_{dk}: Denmark: Industrial output price index of clothing measured in the Danish currency. Source: Reuters EcoWin.

PF_{hk}: Hong Kong: Producer price index of clothing (from $1990Q1$) and consumer price index all products (from $1986Q1$) measured in the Hong Kong currency. Source: Reuters EcoWin.

PF_{tr}: Turkey: Export price index of clothing (from $2004Q1$), export price index of manufactures (from $1995Q1$) and import price index total (from $1986Q1$) measured in the Turkish currency. Source: Reuters EcoWin.

S_j: Value import shares of clothing from country j (China, the Euro area, UK, Sweden, Denmark, Hong Kong and Turkey). Source: Statistics Norway, the Foreign Trade Statistics.

GDP deflators are calculated from nominal GDP and PPP adjusted real GDP. Source: IMF, the World Economic Outlook Database, http://www.imf.org/external/pubs/ft/weo/2009/01/weodata/index.aspx.

ER: Chained geometric Laspeyres index for the exchange rate basket based on S_j and the bilateral exchange rates between Norway and China, the Euro area, UK, Sweden, Denmark, Hong Kong and Turkey. $1986Q1 = 1$. Source: Statistics Norway and Norges Bank.

VC: Domestic variable unit costs of clothing defined as the sum of costs of variable factor inputs relative to total production of clothing. $1986Q1 = 1$. Source: Statistics Norway, QNA.

UR: Unemployment rate defined as the number of unemployment as a percentage of the labour force. Source: Statistics Norway, the Labour Force Survey.