Empir Econ (2013) 45:757–788
DOI 10.1007/s00181-012-0629-6

Trade liberalisation and exchange rate pass-through: the case of textiles and wearing apparels

Andreas Benedictow · Pål Boug

Received: 8 March 2011 / Accepted: 26 June 2012 / Published online: 16 September 2012

Abstract 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 the presence of non-tariff barriers to trade, variables reflecting such a link is rarely included in empirical models. In this paper, we estimate a pricing-to-market model for Norwegian import prices on textiles and wearing apparels, controlling for non-tariff barriers to trade and shift in imports from high- to low-cost countries. We apply the cointegrated VAR approach and develop measures of foreign prices based on superlative price indices (including the Törnqvist and Fischer price indices) and a data calibration method necessary to approximate relative price levels across countries. Our measures of foreign prices thereby account for inflationary differences *and* varying import shares and price level differences (known as the China effect) among trading partners. We show that these measures of foreign prices, unlike standard measures used in the pricing-to-market literature, are likely to produce unbiased estimates of pass-through. Once the China effect is controlled for, we find little evidence that pass-through has changed alongside trade liberalisation.

Keywords Trade liberalisation · The China effect · Import prices · pricing-to-market · Exchange rate pass-through · Vector autoregressive models

JEL Classification C22 · C32 · C43 · E31

A. Benedictow · P. Boug (✉)
Research Department, Statistics Norway, P.O.B. 8131 Dep., 0033 Oslo, Norway
e-mail: pal.boug@ssb.no

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 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 (2005), Atkeson and Burstein (2008), Bugamelli and Tedeschi (2008), Thomas and Marquez (2009) and Gust et al. (2010). Also, empirical studies of small open economies show 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 e.g. Menon (1995b), Menon (1996), Naug and Nymoen (1996), Alexius (1997), Kenny and McGettigan (1998) and Doyle (2004).

However, previous studies usually ignore the Bhagwati hypothesis that the presence of non-tariff barriers to trade may affect pass-through, see Bhagwati (1991). The hypothesis 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.

In this paper, we estimate a model for Norwegian import prices on textiles and wearing apparels (henceforth clothing) that controls for the shift in imports from high- 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. 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 coincides well with 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 conventional 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 trade liberalisation, which led to the massive increase in 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 been dubbed the *China effect* and is likely to be important when quantifying pass-through in regression models. 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 three different measures of foreign prices to be used in the estimation of pass-through. Index number theory advocates the use of the so-called superlative index number formulas when the aim of the aggregation problem is to account for flexible substitution effects between commodities caused by relative price level changes, see e.g. Diewert (1976, 1978). Our first two measures of foreign prices are thus based on the Törnqvist and Fischer price indices, which both belong to the class of superlative price indices. The appealing aggregation properties are, however, somewhat counterbalanced by the fact that available data on foreign prices on clothing are price *indices* and not price *levels*, which makes the superlative price indices (like any other index number formulas) not directly ready for numerical calculations in our context. If the available set of price indices is plugged directly into the superlative price indices, only inflationary impulses implied by price changes and substitution between goods with different price changes are accounted for in the final price aggregate. We, therefore, suggest a data calibration method based on purchasing power parities to account for not only *inflationary differences* as is typical in the pricing-to-market literature, but also *varying import shares* and *differences in price levels*—that is the China effect—among trading partners when constructing the superlative price index measures of foreign prices.[1] Our third measure of foreign prices is based on the often used geometric mean price index with constant import shares as weights, a measure which fails to take account of the China effect. By comparing the estimates of pass-through that come out of modelling the import price of clothing with the 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.

We find that the China effect on traded goods prices is substantial in the clothing industry. Our calculations suggest that the shift in imports from high- to low-cost countries on average has reduced the international price impulses on imports of clothing by around 2 percentage points per year since the early 1990s. Controlling for these effects by means of the superlative price index measures of foreign prices, we estimate import price models consistent with the pricing-to-market hypothesis. Specifically, the pass-through and pricing-to-market elasticities are significantly estimated to 0.44 and 0.56, respectively, irrespective of using the Törnqvist or the Fischer price index measure of foreign prices in the regression model. In contrast, we find that the use of the geometric mean price index measure of foreign prices with constant weights biases the estimates due to international price impulses being substantially overestimated. We also establish that the estimated dynamic model is reasonably stable in sample

[1] To our knowledge, no previous studies have estimated pricing-to-market models with a superlative price index measure of foreign prices. Generally, there are few academic papers which examine the impact of increased imports from China and other low-cost countries on traded goods prices and overall inflation in developed countries. Thomas and Marquez (2009) estimate a pricing-to-market model for aggregated US import prices with a geometric Paasche price index (with varying weights) measure of foreign prices, but do not decompose the price measure into its different inflationary and price level components as we do in the present paper. Nickell (2005) computes the China effect on traded goods prices based on the geometric Paasche price index, but does not estimate a model when analysing the impact of a changing trade pattern on import prices and overall consumer price inflation in the UK. Wheeler (2008) also calculates the China effect on traded goods prices following the operational route in Nickell (2005). In addition, Wheeler (2008) estimates panel regressions of UK inflation by goods category on the level and growth of the import share from China as the main determinants.

and exhibits no serious forecasting failures around the dates of the shifts in trade policy. That no serious structural breaks are detected may reflect that likely pass-through effects of changes in trade policy are controlled for through the superlative price index measures of foreign prices. Consequently, once the effect of shifts in imports towards low-cost countries is controlled for, we find little evidence that the properties of the import price equation have changed alongside trade liberalisation.

Based on our findings, we claim that the choice of aggregation formula for foreign prices matters for the quantification of an import price model, an issue typically ignored in the pricing-to-market literature. We emphasise that the empirical example of clothing is *not* a special case as trade liberalisation has led to increased exports of several product categories (not just clothing) from China and other low-cost countries to Norway and other high-cost countries over the last two decades or so. There exist a large parallel literature on trade that explicitly discusses and demonstrates the importance of choosing the relevant price index to incorporate new products (or countries entering or leaving the market for a particular good) into an aggregate of international prices, see e.g. Feenstra (1994) and Broda and Weinstein (2006). Gaulier et al. (2008) present extensive empirical evidence that the choice of aggregation method matters for the calculation of international prices and show that the Törnqvist and Fischer price indices provide similar results in practice.

The rest of the paper is organised as follows: Sect. 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 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 Sect. 5 presents the estimated dynamic model for import prices on clothing. Section 6 concludes.

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). Markets for clothing are typically characterised by imperfect competition between firms producing differentiated products. Furthermore, these markets are segmented due to trade barriers, transportation costs and imperfect information. Profit maximisation under these 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 the presence of (and removal of) non-tariff barriers to trade build on Naug and Nymoen (1996) and Menon (1996).

2.1 Pricing-to-market

Consider a representative foreign firm producing a differentiated product of clothing exported to n 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

$X_i = X_i(PX_i \cdot ER_i, PQ_i, DP_i)$, where PX_i is the firm's export price measured in the exporter's currency, ER_i is the bilateral exchange rate with respect to 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

$$\prod(PX_1, \ldots, PX_n) = \sum_{i=1}^{n} PX_i \cdot X_i\,(PX_i \cdot ER_i, PQ_i, DP_i) - C\left[\sum_{i=1}^{n} X_i\,(PX_i \cdot ER_i, PQ_i, DP_i)\,, W\right], \tag{1}$$

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

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

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 $\eta_i = \eta_i((PX_i \cdot ER_i), PQ_i, DP_i)$ is the elasticity of demand in market i. As every export price reflects conditions in each particular 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.

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

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[2]

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

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

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

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 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

Before the Uruguay Round 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. 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 10–15 years. Substantial reduction over time in tariff rates on imports of clothing has likewise pulled in the same direction.[3]

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. The

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

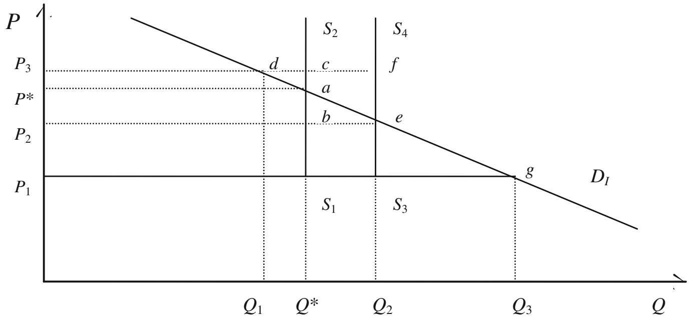

Fig. 1 Pass-through with presence and removal of quota restrictions

effects of quantity restrictions on pass-through do not depend on particular market structures. We, therefore, extend Menon (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 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 shift 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 so 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 horizontal 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 Fig. 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 the 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

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 three measures of foreign prices based on (i) the Törnqvist price index (pf_t^T) with varying import shares as weights, (ii) the Fischer price index (pf_t^F) with varying import shares as weights and (iii) 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:

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

where $\psi = \gamma_1/(1 + \gamma_1)$ and $\delta = \gamma_2/(1 + \gamma_1)$. Because the unemployment rate enters (5) without a logarithmic transformation the markup is specified as $\lambda_t = K(VC/PI)_t^{\gamma_1} \exp^{(\gamma_2 UR_t)}$. A testable implication of LOP from (5) is that $(pi - pf^i - er)_t$ is stationary or forms a long run cointegration relationship when the variables involved all are non-stationary. 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.

Finally, 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 u_t is thus correlated with the export price measure and (5) only forms a cointegration relationship when the measurement errors are stationary. We show in Sect. 4 that a well-specified VAR model and a significant cointegrating vector exist using our data set. Thus, judged by statistical criteria, measurement errors seem to be a minor problem. The disturbance term in (5) also 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 would be overestimated in (5) to the extent that u_t is correlated with UR_t. Again, our estimated import price model is well specified, indicating that u_t and UR_t are not much correlated.

3.1 The measure of foreign prices

As numerous index number formulas with different aggregation properties exist, see e.g. Balk (2008) for a survey, we are faced with the problem of which to choose to account properly for the China effect in the measure of foreign prices. Two commonly used index number formulas in the literature on trade are the Laspeyres and Paasche price indices, which both measure the evolution of prices between a base and a comparison period for a given basket of goods. These price indices, however, produce biased measures of price evolutions for two main reasons. Firstly, as the assigned weights are from a single period only, they fail to capture any substitution effect among different goods from one period to another. Whereas the Laspeyres price index tends to overestimate price growth because it uses the weights from the base period, the Paasche price index tends to underestimate price growth by assigning larger weights to products with increased quantity following a relative price decrease. Secondly, both the Laspeyres and Paasche price indices fail to capture disappearance of old or appearance of new products between the base period and the comparison period. Ignoring new products generally leads to overestimation of price growth, see e.g. Feenstra (1994) and Broda and Weinstein (2006).

Both sources of bias discussed above may be dealt with by chaining the Laspeyres and Paasche price indices. Such chained indices account for changes in the basket of goods, including new products, through variations in the weights assigned to each product entering the basket. However, the chained indices still suffer from a measurement bias as they only use information from one of the two periods in each elementary index. Feenstra (1997) show empirically that the Laspeyres and Paasche price indices represent the upper and lower bounds of the real price development.

The superlative Törnqvist and Fischer price indices, on the other hand, use information from both the base and the comparison period by composing the Laspeyres and Paasche price indices, see e.g. Diewert (1976, 1978). Whereas the Törnqvist price index is defined as the geometric mean of the geometric Laspeyres and Paasche price indices, the Fischer price index is defined as the geometric mean of the arithmetic Laspeyres and Paasche price indices, see e.g. Balk (2008). The fact that all information at hand is utilised motivates us to apply the Törnqvist and Fischer price indices as the underlying index number formulas for the measure of foreign prices. The Törnqvist price index (PF^T) in our context equals

$$PF_t^T \equiv \left(\prod_{j=1}^{n} PF_{j,t}^{s_{j,t-1}} \prod_{j=1}^{n} PF_{j,t}^{s_{j,t}} \right)^{1/2} = \prod_{j=1}^{n} PF_{j,t}^{\bar{s}_{j,t}}, \tag{6}$$

where the expressions $\prod_{j=1}^{n} PF_{j,t}^{s_{j,t-1}}$ and $\prod_{j=1}^{n} PF_{j,t}^{s_{j,t}}$ are the geometric Laspeyres and Paasche price indices, respectively, $\bar{s}_{j,t} \equiv \frac{s_{j,t}+s_{j,t-1}}{2}$ for $j = 1, \ldots, n$, $s_{j,t-1}$ and $s_{j,t}$ are the value shares of imports from trading partner j in the base period $t-1$ and the comparison period t, respectively, $0 \leq s_{j,h} < 1$ and $\sum_{j=1}^{n} s_{j,h} = 1$ for $h = t-1, t$. We observe that PF_t^T is a weighted geometric average of the foreign price indices ($PF_{j,t}$), the weights being the arithmetic means of the value import shares of the base and comparison period. Aggregating the foreign price indices by means of (6) directly will only capture inflationary impulses because a price index by construction measures the percentage change in a price relative to a base period. We, therefore, suggest a calibration method based on purchasing power parities to approximate relative price levels to accommodate both inflationary impulses and price level differences across countries in the measure of foreign prices.

The first step of our data calibration method involves constructing calibration coefficients for each trading partner, labelled λ_j, by the formula

$$\lambda_j = \frac{\frac{GDP_j^{NOM}}{GDP_j^{PPP}}}{\frac{GDP_{numeraire}^{NOM}}{GDP_{numeraire}^{PPP}}}, \tag{7}$$

where GDP_j^{NOM} and GDP_j^{PPP} are nominal GDP and purchasing power parity adjusted volume of GDP for trading partner j, respectively, and $GDP_{numeraire}^{NOM}$ and $GDP_{numeraire}^{PPP}$ are corresponding GDP figures for the numeraire country. We point out that the calibration coefficients are unitless and easy to interpret for our purposes. For instance, a λ_j equal to 0.5 would imply that the overall price level in country j is 50 % of that in the numeraire country.

The second step of our data calibration method involves multiplying the calibration coefficients from (7) with the corresponding price indices from (6). Formally, we rewrite (6) as

$$PF_t^T = \prod_{j=1}^{n} PF_{j,t}^{*\bar{s}_{j,t}}, \tag{8}$$

where $PF_{j,t}^* = \lambda_j PF_{j,t}$. The calibrated price indices $PF_{j,t}^*$ in (8)—which are to be interpreted as relative price levels—equal the relative price levels calculated from (7) in the base period, in which the price indices $PF_{j,t}$ are set equal to unity. In all other periods, the calibrated price indices develop according to the development of the levels of respective original price indices. Formula (8) is an aggregate of foreign prices that accounts for the *total* price effects of the shift in imports towards low-cost-countries.

Analogous to (8), the Fischer price index (PF^F) in our context equals

$$PF_t^F = \left(\frac{\sum_{j=1}^{n} s_{j,t-1} PF_{j,t}^*}{\sum_{j=1}^{n} s_{j,t} \frac{1}{PF_{j,t}^*}} \right)^{1/2}, \tag{9}$$

where the expressions $\sum_{j=1}^{n} s_{j,t-1} PF_{j,t}^*$ and $\frac{1}{\sum_{j=1}^{n} s_{j,t} \frac{1}{PF_{j,t}^*}}$ are the Laspeyres and Paasche price indices (calibrated with λ_j for our purposes), respectively. Aggregating the foreign price indices by means of (9) will, just like (8), produce a final index number that measures the *total* price effects of the shift in imports towards low-cost-countries. We show below that (8) and (9) generate similar foreign price aggregates in our case. Hence, we shall in the following decomposition of the *total* price effects into inflation and price level effects concentrate on the Törnqvist price index as the underlying index number formula. To simplify matters in the exposition without loss of generality, we only consider two trading partners ($j = 1, 2$). First, taking natural logarithms of (8) and differencing once, we obtain

$$\begin{aligned}\Delta pf_t^T &= pf_t^T - pf_{t-1}^T \\ &= \bar{s}_{1,t} pf_{1,t}^* + \bar{s}_{2,t} pf_{2,t}^* \\ &\quad - \bar{s}_{1,t-1} pf_{1,t-1}^* - \bar{s}_{2,t-1} pf_{2,t-1}^*,\end{aligned} \tag{10}$$

where Δ indicates the first difference operator. Then, adding and subtracting $\bar{s}_{1,t} pf_{1,t-1}^*$ and $\bar{s}_{2,t} pf_{2,t-1}^*$ to the right hand side of (10), making use of the adding up condition of the value shares of imports and collecting terms, we get an expression for the percentage change in the aggregate foreign price in period t that reads as

$$\begin{aligned}\Delta pf_t^T &= \bar{s}_{1,t} \Delta pf_{1,t}^* + \bar{s}_{2,t} \Delta pf_{2,t}^* \\ &\quad + \Delta \bar{s}_{1,t} \left(pf_{1,t-1}^* - pf_{2,t-1}^* \right).\end{aligned} \tag{11}$$

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 first two terms on the right hand side of (11) show that increasing inflation on clothing from each of the trading partners contribute to increasing inflationary impulses faced by Norwegian importers. The larger the price increase and the larger the import share, the larger is the inflationary impulse (measured in foreign currency) in Δpf_t^T. The last term on the right hand side of (11) constitutes the total effect of the price level differences, that is 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 deflationary impulse in Δpf_t^T. We notice that the China effect is zero only in the special cases when the import shares are constant ($\Delta \overline{s}_{1,t} = 0$) and when the composition of trade changes between countries with identical price levels ($pf^*_{1,t-1} - pf^*_{2,t-1} = 0$). Although the bilateral distribution of the China effect can be sensitive to the choice of numeraire country, the size of the *aggregated* China effect calculated from (11) is not. The level of aggregate foreign prices may now be calculated by setting PF_t^T equal to unity in the base period and letting the level of the price index from then on be determined consecutively by the measured growth rates from (11).

The standard practise in related studies is to weight together some proxy for foreign prices by means of a geometric 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 (PF^G) in our context, analogous to (8), equals

$$PF_t^G = \prod_{j=1}^{n} PF_{j,t}^{*\overline{s}_j}, \tag{12}$$

where the exponent $\overline{s}_j$ now is the constant value share of imports from trading partner j, $0 \leq \overline{s}_j < 1$ and $\sum_{j=1}^{n} \overline{s}_j = 1$. Following Naug and Nymoen (1996), we set $\overline{s}_j$ equal to the average of each import share over the sample period. Again, differencing once the natural logarithms of (12), adding and subtracting $\overline{s}_1 pf^*_{1,t-1}$ and $\overline{s}_2 pf^*_{2,t-1}$, making use of the adding up condition of the value shares of imports and collecting terms, we get

$$\Delta pf_t^G = \overline{s}_1 \Delta pf^*_{1,t} + (1 - \overline{s}_1)\, \Delta pf^*_{2,t}. \tag{13}$$

We see that (13) only accounts for inflationary differences among the trading partners. Accordingly, international price impulses are overestimated to the extent that China effects are present. On this background, we expect that the estimate of the degree of pass-through will reflect an omitted variable bias when PF_t^G is used instead of the superlative price indices 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 (11). 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. A linear trend is just a 'measure of ignorance' and at best it represents an omitted variable in the regression model. We, therefore, 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 superlative price indices.

3.2 Data[4]

We use quarterly, seasonally unadjusted time series covering the period 1986Q1–2008Q1. The import price (pi_t) is an implicit deflator for imports of clothing with Norwegian substitutes measured in Norwegian currency. The products comprising the deflator are priced *cif* at the Norwegian border. Hence, prices include costs of insurance and freight, but exclude tariffs. The deflator is a chained geometric mean price index calculated by weighting together each unit price, which is 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 continuously updated in accordance with the development in the country composition of clothing imports, the deflator reflects the shifts in imports from high- to low-cost countries over time.

To construct the superlative price indices (pf_t^T and pf_t^F), we need data on import shares, export prices and price levels for each one of the main trading partners. The foreign trade statistics provide time series of import shares by country. The main exporters of clothing to Norway 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 % of Norwegian imports of clothing as an average over the sample period.[5] Because the euro area is treated as one country, we abstract from any import substitution from high- to low-cost countries within this area.

It proved difficult to find long and consistent proxies for export prices for 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 $1986Q1$.

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 for the level of productivity in each country. Because such data are not available, we use purchasing power parity-adjusted GDP figures provided by IMF. Table 1 shows the

[4] See the Appendix for details about data definitions and sources.

[5] 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 % in 2008, but is left out of the analysis due to lack of relevant price data.

Table 1 Average relative price levels (λ_j). 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

average calculated international relative price levels (λ_j) over the period 1991–2008 based on (7).[6]

As the euro area is chosen as numeraire country, λ_{eu} equals unity. Our calculations show that the overall price level in China is 41 % of that in the euro area. The corresponding figure for Turkey is 56 %. Hence, both China and Turkey stand out as low-cost countries in our study. We recognise that the relative price levels in Table 1 are good proxies only to the extent that relative price levels on clothing are similar to relative GDP deflators across countries, an assumption that needs 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 the case, the calculated superlative price index measures of foreign prices, based on the figures in Table 1 will overestimate the true negative price level impulses to the Norwegian economy. The superlative price indices 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 superlative price indices, and thereby the sensitivity of the estimates of pass-through and pricing-to-market, when the relative price levels for China and Turkey in Table 1 are increased and decreased by 50 %, other things equal. Nevertheless, the relative price levels in Table 1 are used as benchmarks to calibrate the respective export price indices in (8) and (9).

Figure 2 displays the country-specific export prices ($pf_{j,t}$), measured in foreign currency and normalised to unity in 1986Q1. We observe 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 the 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 to reduced export possibilities for most high-cost countries in the successive

[6] 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.

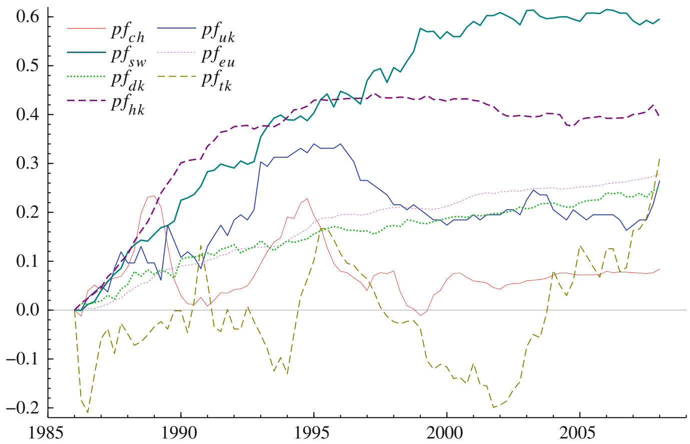

Fig. 2 Time series for foreign prices ($pf_{j,t}$). Source: See the Appendix

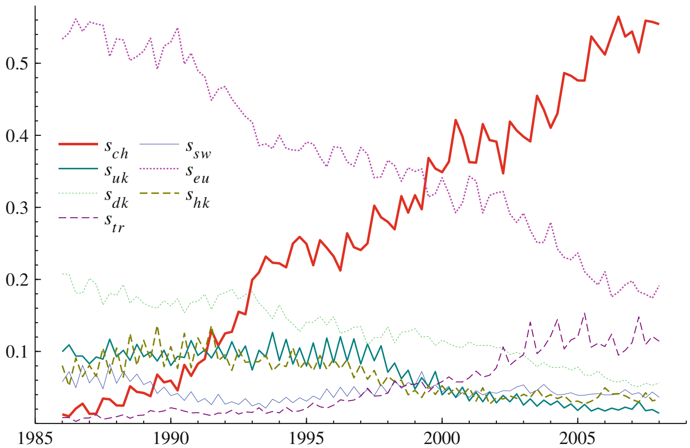

Fig. 3 Time series for import shares ($s_{j,t}$). Source: See the Appendix

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

Figure 3 displays the country-specific import shares ($s_{j,t}$), which sum to unity in each period. We see that the import share from China increased from a few per cent in 1986 to around 55 % in 2008. The import share from the euro area fell likewise from around 55 % in 1986 to around 20 % in 2008. After a substantial increase in the import share from the mid 1990s, Turkey supplied more than 10 % of Norwegian imports of clothing in 2008. Whereas, the import share from Sweden was relatively stable around 5 % 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

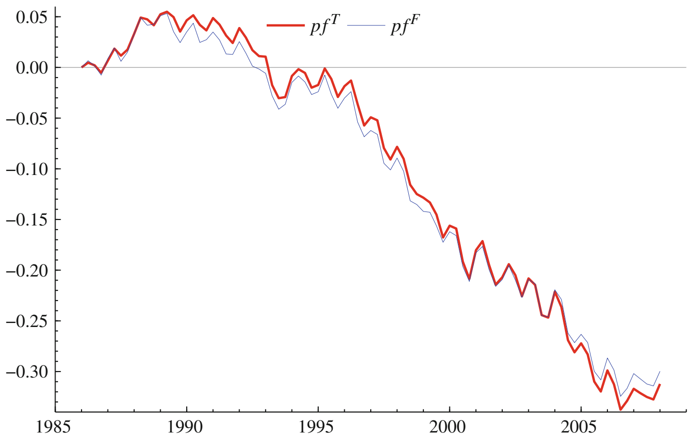

Fig. 4 Time series for foreign prices (pf_t^T and pf_t^F)

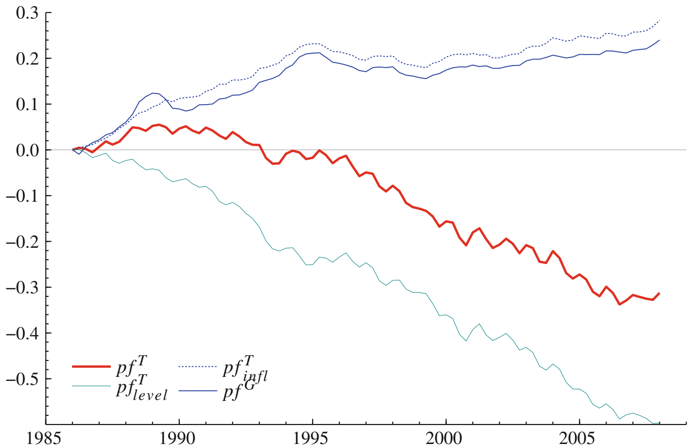

Fig. 5 Time series for foreign prices (pf_t^T , $pf_{t,infl}^T$, $pf_{t,level}^T$ and pf_t^G)

during the same period. Overall, the shift in imports towards low-cost countries at the expense 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.

Figure 4 displays the computed Törnqvist (pf_t^T) and Fischer (pf_t^F) price aggregates (measured in foreign currency) based on (8) and (9), respectively. Practically speaking, we see that the two superlative price indices generate identical aggregates with a substantial fall in international export prices on clothing during the sample period. Our calculations indicate that the price aggregates are roughly 30 % lower in 2008 compared to 1986, which implies on average a yearly decrease of around 1.3 percentage points.

Figure 5 displays the Törnqvist price index measure of foreign prices (pf_t^T) and its two components, the inflation effects ($pf_{t,\inf l}^T$) and the price level effects ($pf_{t,level}^T$) based on (11), together with the geometric mean price index measure of foreign prices with constant weights (pf_t^G) based on (12). According to our calculations, the shift in imports from high- to low-cost countries—the China effect – 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 effects were moderate, reflecting little substitution of imports towards low-cost countries due to strict trade regulations. The international price impulses were, however, pulled higher and dominated by inflationary effects 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^T from 1995 onwards. Even though the last quota restriction was lifted in 1998, the price level effects continued to pull down pf_t^T during the last decade, which indicates that trade liberalisation may have had long lasting effects on international export prices on clothing. We also observe that the development in pf_t^G parallels the development in $pf_{t,\inf l}^T$. More importantly though, is the substantial differences in pf_t^T and pf_t^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 somewhat less than 30 % through the sample period rather than a price fall of the same magnitude. We believe that pf_t^T provides a better measure of the true international price development faced by Norwegian importers of clothing given the significant change in the import pattern over time.

Figure 6 displays the Törnqvist price aggregate (pf_t^T) based on (8) together with the Törnqvist price aggregate based on a 50 % increase ($pf_{t,high}^T$) and decrease ($pf_{t,low}^T$) in the relative price levels for China and Turkey in Table 1. We notice that the development in pf_t^T is rather sensitive to different assumptions made about the relative price levels for China and Turkey. Whereas a 50 % increase in the relative price levels for China and Turkey makes an international price fall of only 5 % ($pf_{t,high}^T$) from

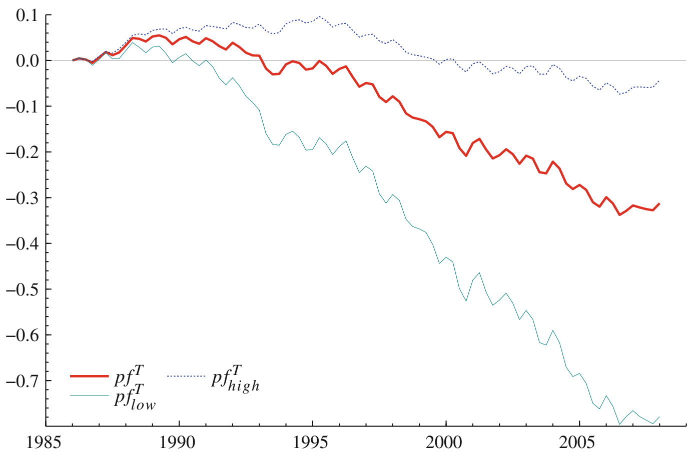

Fig. 6 Time series for foreign prices (pf_t^T, $pf_{t,high}^T$ and $pf_{t,low}^T$)

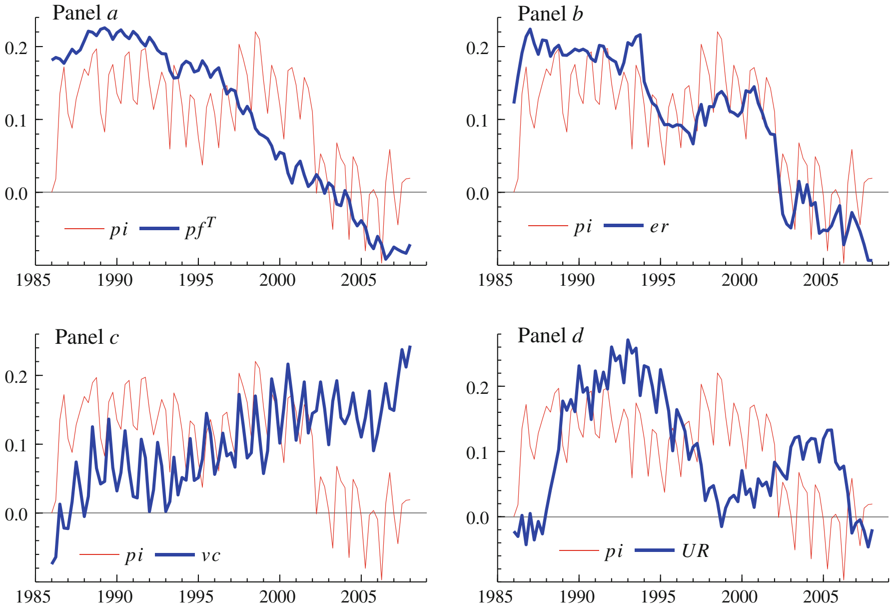

Fig. 7 Time series for pi_t, pf_t^T, er_t, vc_t and UR_t. Sources: See the Appendix

1986 to 2008, a 50 % decrease in the same relative price levels produces a price fall of as much as 55 % ($pf_{t,low}^T$) in the same period. We return to this issue below and analyse whether the sensitivity in pf_t^T produces a serious sensitivity in the estimates of pass-through and pricing-to-market.

Figure 7 displays the time series for the import price of clothing (pi_t) together with the Törnqvist price index measure of foreign prices (pf_t^T) in panel *a*, the exchange rate (er_t) in panel *b*, the domestic variable unit costs (vc_t) in panel *c*, and the unemployment rate (UR_t) in panel *d*. The exchange rate series is a chained geometric mean index the construction of which parallels that of pf_t^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^T, er_t, vc_t and UR_t are adjusted in Fig. 7 to match that of pi_t, which is normalised to unity in 1986Q1.

It is evident that pi_t, pf_t^T and er_t all exhibit a clear downward trend throughout the sample period, whereas vc_t shows some upward trend. At the same time, import prices of clothing relative to foreign prices measured in Norwegian currency $(pi - pf^T - er)_t$ increased from 1986 to 2008, which may be explained by the fact that variable unit costs relative to import prices $(vc - pi)_t$ 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_t as $(pi - pf^T - er)_t$ increased somewhat more than $(vc - pi)_t$. As indicated by panel *d*, the development in pi_t may also partly be explained by the development in the domestic demand pressure (UR_t). Specifically, the apparent fall in pi_t during the first half of the 1990s and during the years between 1999 and 2006 coincides well with increased UR_t in the same periods. Likewise, the increase in pi_t during the second half of the 1990s matches rather closely with decreased UR_t. 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 to non-stationary time series properties. The unemployment rate, on the other hand, is stationary by construction. However, we follow Bjørnstad and Nymoen (1999) in the subsequent analysis and treat UR_t as if it is non-stationary within the sample period.

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 start with an unrestricted p-dimensional VAR of order k having the form

$$X_t = \sum_{i=1}^{k} \Pi_i X_{t-i} + \mu + \varpi t + \varepsilon_t, \quad t = k+1, \ldots, T, \tag{14}$$

where X_t is a $(p \times 1)$ vector of modelled variables at time t, μ represents constants and seasonals, ϖ is a $(p \times 1)$ coefficient vector of a linear deterministic trend t, $\Pi_1, \ldots, \Pi_k$ are $(p \times p)$ coefficient matrices of lagged level variables and $\varepsilon_{k+1}, \ldots, \varepsilon_T$ are independent Gaussian variables with expectation zero and (unrestricted) $(p \times p)$ covariance matrix Ω. The initial observations of $X_1, \ldots, X_k$ are kept fixed. The question now is how (14) can be re-parameterised 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 $\Pi = -(I - \Pi_1 - \cdots - \Pi_k)$.

The way the CVAR is formulated in our context depends on the exogeneity status or otherwise of the unemployment rate series. Firstly, we consider the case when the unemployment rate series is endogenous in the system, hence (14) is a five-dimensional VAR in $X_t = (pi_t, pf_t^i, er_t, vc_t, UR_t)'$, $i = T, F, G$. Once $X_t \sim I(1)$, 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 $5 \times r$ 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

$$\Delta X_t = \Gamma_1 \Delta X_{t-1} + \alpha\beta' X_{t-1} + \mu + \delta t + \varepsilon_t, \tag{15}$$

where $\beta' X_{t-1}$ is an $r \times 1$ 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 a (5×5) coefficient matrix of the lagged differentiated variables. Next, we 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 four-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 (15) as (again assuming $k = 2$)

$$\Delta X_{1,t} = A_1 \Delta X_{2,t} + \Gamma_{1,1} \Delta X_{t-1} + \alpha_1 \beta' X_{t-1} + \mu_1 + \delta_1 t + \varepsilon_{1,t}, \tag{16}$$

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 (16) as a non-modelled variable. Because the number of relevant variables to be included in (14), 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 (16), we first estimate the full system in (15) and test formally rather than assume the weak exogeneity status of the unemployment rate in that system. 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.

Strictly speaking, the cointegration rank does not need to be determined from the partial system once it has been determined from the full system. Nevertheless, we re-determine the cointegration rank from (16) for the sake of comparison with the rank determination from (15). However, as noted by Harbo et al. (1998), the asymptotic 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 cointegration rank, 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 constants 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. Not surprisingly, we find that the two calculated superlative price indices produce similar cointegration analyses and near identical corresponding dynamic import price models. Accordingly, we shall below concentrate on the empirical findings based on the Törnqvist price index measure of foreign prices.[7]

4.1 Cointegration analysis based on PF_t^T

Irrespective of specifying a full five-dimensional VAR in $X_t = (pi_t, pf_t^T, er_t, vc_t, UR_t)'$ or a partial four-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

[7] Results from the cointegration analysis and the modelling of the dynamic import price equation based on the Fischer price index measure of foreign prices are available from the authors upon request.

Table 2 Tests for cointegration rank based on PF_t^T

Full CVAR system				Partial CVAR system			
r	λ_i	λ_{trace}	p-val$_{Ox}$	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$_{Ox}$ are the significance probabilities from OxMetrics and $5\,\%_{Harbo}$ are the critical values (5 % significance level) from Table 2 in Harbo et al. (1998)

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.[8] 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 measure of foreign prices assuming $k = 3$.

We notice that the null hypothesis of no cointegration can be rejected at the 5 % 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. 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 the estimates of which 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 parsimonious, 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 % 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 excluded

[8] 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. As noted by Franses and Lucas (1998), standard cointegration tests are sensitive to atypical events such as outliers and structural breaks.

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

from β. The linear trend, however, is clearly *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 (16) assuming $r = 1$. Table 3 summarises results from these tests.

Firstly, 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 Sect. 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)

$$pi_t = const. + 0.444 pf_t^T + \underset{(0.016)}{0.444} er_t + 0.556 vc_t - \underset{(0.003)}{0.020} UR_t, \tag{17}$$

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 (17), 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.[9] The pass-through and pricing-to-market elasticities are significantly estimated to 0.44 and 0.56, respectively. Also, the estimated import price equation includes significant negative effects of the unemployment rate such that decreases in domestic demand pressure (proxied by increases in the unemployment rate) cause prices of imports to fall somewhat.

9 Although significantly estimated, the adjustment coefficient for pf_t^T is only 40 % of that for pi_t.

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 the 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. 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 % for some of the quota-protected textiles and wearing apparels studied. Menon (1996) partly views this finding in light of the Bhagwati hypothesis as 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 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, it is likely that pass-through has not changed dramatically since the mid 1990s. Recursive estimates of the pass-through coefficient in (17) are reasonably stable in the years after 1995. Also, recursively estimated $\chi^2_{(5)}$ indicate that the restrictions in (17) are supported by the data throughout the second half of the sample period.

We complete the cointegration analysis based on pf_t^T by examining potential sensitivity in the estimate of pass-through based on different assumptions made about the relative price levels for China and Turkey. As already revealed from Fig. 6, the calculated development in pf_t^T is somewhat sensitive to a 50 % increase ($pf_{t,high}^T$) and decrease ($pf_{t,low}^T$) in the relative price levels for China and Turkey in Table 1. We obtain the following estimated cointegrating vectors with $pf_{t,high}^T$ and $pf_{t,low}^T$ replacing pf^T, all other modelling issues equal:

$$pi_t = const. + 0.604 pf_{t,high}^T + \underset{(0.019)}{0.604} er_t + 0.396 vc_t - \underset{(0.003)}{0.019} UR_t, \tag{18}$$

$$pi_t = const. + 0.306 pf_{t,low}^T + \underset{(0.015)}{0.306} er_t + 0.694 vc_t - \underset{(0.005)}{0.023} UR_t. \tag{19}$$

Similar to (17), 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 (18) and (19), 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 relative price levels for China and Turkey. The estimate of pass-through increases and decreases by 33 % when $pf_{t,high}^T$ and $pf_{t,low}^T$ replace pf_t^T, which we consider as a rather moderate sensitivity in the estimate given the rather substantial magnitude of the shift in the relative price levels. We shed some further light on the sensitivity in the estimates of pass-through and pricing-to-market due to the potential problem of omitted variable bias in the subsequent cointegration analysis based on pf_t^G rather than pf_t^T.

Table 4 Tests for cointegration rank based on PF_t^G

Full CVAR system				Partial CVAR system			
r	λ_i	λ_{trace}	$p\text{-val}_{Ox}$	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, $p\text{-val}_{Ox}$ are the significance probabilities from OxMetrics and $5\,\%_{Harbo}$ are the critical values (5 % significance level) from Table 2 in Harbo et al. (1998)

4.2 Cointegration analysis based on PF_t^G

As with the Törnqvist price index 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. Again, 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 when the measure of foreign prices are based on the geometric mean price index with constant weights.

Again, the rank should be set to unity in the case of the full CVAR system at the 5 % 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 conduct 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 that a specific variable does not enter the cointegrating relation is rejected for all the variables pi_t, pf_t^G, er_t, vc_t and UR_t, a finding in line with the analysis above using the Törnqvist price index 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. Note that H_2 is now 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 restrictions corresponding to 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)

$$pi_t = const. + 0.601 pf_t^G + \underset{(0.101)}{0.601} er_t + 0.399 vc_t - \underset{(0.005)}{0.020} UR_t - \underset{(0.00047)}{0.00207} t, \quad (20)$$

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

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 estimate obtained with the Törnqvist price index measure of foreign prices. More important though, when comparing (17) and (20), 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 (20) and *not* in (17).

One possible interpretation of these findings is that the effects of shifts in imports from high- 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 (20), effects which are *not* explicitly picked up by pf_t^G alone. As seen from Fig. 5, 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 (20) 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 0.44 from (17) and the 2 percentage points yearly decrease in the price level term $pf_{t,level}^T$. However, the fact that $pf_{t,level}^T$ exhibits some apparent cycles, especially around 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. For this reason, we suspect the estimates of pass-through and pricing-to-market in (20) to be somewhat biased compared to those in (17). 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 (20), results which are unlikely given the facts about the clothing industry outlined in Sect. 2. Although not accepted by the data[10], these findings also point to the likely problem of an omitted

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

variable bias if shifts in imports towards low-cost countries and trade liberalisation effects are *not* explicitly controlled for in the import price equation for clothing.

We have seen that using the Törnqvist price index 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 the geometric mean price index with constant weights together with a linear trend (which at best represents the China effect) in the regression model to quantify pass-through consistently.

5 A dynamic import price model

The degree of pass-through may, just like trade policy, be linked to 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. We shall here test the hypothesis that pass-through has changed alongside trade liberalisation and particular exchange rate fluctuations by examining stability properties of an estimated dynamic equilibrium correction model (henceforth $EqCM$). 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

$$\Delta pi_t = const. + \sum_{i=1}^{2} \varphi_{1,i} \Delta pi_{t-i} + \sum_{i=0}^{2} \varphi_{2,i} \Delta (pf^T + er)_{t-i} + \sum_{i=0}^{2} \varphi_{3,i} \Delta vc_{t-i}$$
$$+ \sum_{i=0}^{2} \varphi_{4,i} \Delta UR_{t-i} + \eta EqCM_{t-1} + \eta_1 S1_t + \eta_2 S2_t + \eta_3 S3_t + e_t. \quad (21)$$

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 $\Delta(pf^T + er)_t$ is denominated in Norwegian currency, a short run restriction imposed from the outset in line with the corresponding 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 (21) with two lags, whereas the $EqCM$ term [defined in accordance with (17)] is lagged one period. The error term e_t is assumed to be white noise. Simplifications from the general to the specific model is performed using the autometrics option in OxMetrics 6, see Doornik and Hendry (2009). Autometrics picks the following *specific* model in our case together with diagnostic tests[11] and the estimated standard errors below the point estimates (sample period: 1986Q4–2008Q1):

[11] AR_{1-5} is Harvey (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 (2008), *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.

$$
\begin{aligned}
\Delta pi_t = & \underset{(0.055)}{-0.336}\Delta pi_{t-1} + \underset{(0.088)}{0.479}\Delta(pf^T + er)_t - \underset{(0.006)}{0.021}\Delta UR_t - \underset{(0.006)}{0.020}\Delta UR_{t-1} \\
& \underset{(0.067)}{-0.385}[pi_{t-1} - 0.44(pf^T + er)_{t-1} - 0.56vc_{t-3} + 0.020UR_{t-2}] \\
& \underset{(0.024)}{+0.188} - \underset{(0.009)}{0.065}S1_t - \underset{(0.007)}{0.109}S2_t
\end{aligned}
\tag{22}
$$

Diagnostic tests:

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

The equilibrium correction term enters significantly in (22). The 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. The *EqCM* 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.[12] Accordingly, import prices respond quickly and with some overshooting to shocks in foreign prices (denominated in foreign currency) and exchange rates. However, the specific model also contains 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 (22). 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. Fig. 7, 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 (22) 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

[12] Both Δpf_t^T and Δer_t enter *in*significantly as separate explanatory variables in (22), 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 (22). Hence, Δpf_t^T and Δer_t may be regarded as weakly exogenous for the short run parameters in the specific model. Thus, the parameters are consistently estimated by OLS, see Urbain (1992).

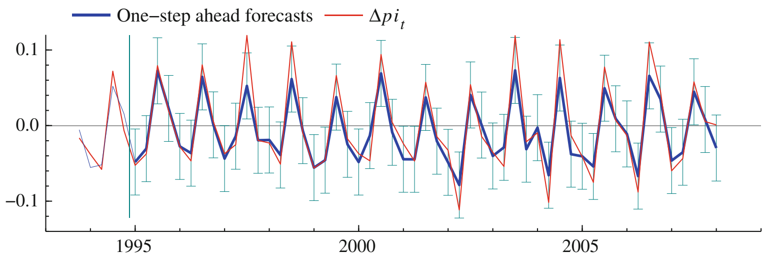

Fig. 8 Actual values and one-step ahead forecasts of Δpi_t

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. Fig. 7, panel *b*) may thus have been viewed as transitory by foreign firms, in which case it may have been rational to respond slower, if at all.

Turning to parameter stability properties of the specific model, we first notice that the model shows no sign of misspecification as reported below (22). 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 trade policy changes and exchange rate fluctuations during the sample period. If pass-through has changed, 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 (22) based on observations until $1994Q4$, and leaving 53 quarters (1995Q1–2008Q1) for out-of-sample forecasts. Figure 8 depicts actual values of Δpi_t together with its one-step ahead forecasts and 95 % confidence intervals to each forecast in the forecasting period (shown by the vertical error bars of $\pm$2SE).

We observe that the forecasts only miss significantly the observed values of 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 well be explained by the Asian financial crises rather than the shift in trade policy itself. As seen from Fig. 7 (panel *b*), the import prices on clothing increased following a sharp depreciation of the Norwegian currency in 1997. Nevertheless, 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 fairly constant throughout the trade liberalisation period. Also, a Chow test statistic of parameter constancy between the sample and the forecasting periods 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 (22) 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 shifts in trade policy 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 trade policy are controlled for through pf_t^T, effects which may otherwise be reflected in unstable estimates of the model.

After the introduction of inflation targeting in 2001Q1, we should expect higher pass-through if foreign exporters believed the exchange rate changes to be more permanent in nature than before. The fact that the model exhibits no forecasting failures suggests that pass-through has remained unchanged also around the date of the shift in monetary policy. These findings may be explained by the fact that foreign firms also experienced relatively high exchange rate volatility during the 1990s, cf. Fig. 7 (panel *b*). After leaving the fixed exchange 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 perceptions of the permanent nature of (large) exchange rate fluctuations changed, 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.[13]

6 Conclusions

Economic theory predicts that the 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 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- to low-cost countries. The novelty of the paper, we believe, is that the measure of foreign prices is based on superlative price indices (including the Törnqvist and Fischer price indices) and a data calibration method necessary to approximate relative price levels across countries. As such, we allowed not only for inflationary differences as is common in previous pricing-to-market 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.

We found that the China effect on traded goods prices is substantial in the clothing industry. Our calculations suggest that the shift in imports from high- 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 superlative price index measures of foreign prices, we established import price models for clothing consistent with the pricing-to-market hypothesis. Specifically, we found the pass-through and pricing-to-market elasticities to be 0.44 and 0.56, respectively, irrespective of using the Törnqvist or the Fischer price index measure of foreign prices in the regression model. 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

[13] We also controlled for any instabilities in the estimated model by means of the outlier detection procedure available in OxMetrics 6. It turned out that no significant outliers were detected by this procedure during the periods of trade liberalisation and shift in monetary policy.

trade has pushed pass-through upwards, other things equal. That is, once the China effect is controlled for through the measure of foreign prices, we found little evidence that the long run properties of the import price model have changed significantly alongside trade liberalisation. By way of contrast, we found that the often used geometric mean price index with constant weights overestimates international price impulses and thereby produces 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. We may approximate the China effect through a linear trend in the model together with the geometric mean price index with constant weights. However, we showed that such a model is still likely to produce some biasedness in the estimates of pass-through and pricing-to-market. Because the China effect exhibits some apparent cycles, we argue that the superlative price index measures of foreign prices are superior to the geometric price index with constant weights combined with the linear trend, which implicitly assumes that the China effect has been constant throughout the sample period. We further established that the dynamic 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.

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.

Acknowledgments We are grateful to the coordinating editor Robert M. Kunst, an anonymous referee, Roger Bjørnstad, Thomas von Brasch, Ådne Cappelen, Erling Holmøy, John Muellbauer, Ragnar Nymoen and Terje Skjerpen for helpful comments and suggestions. Estimation results are obtained using OxMetrics 6, see Doornik and Hendry (2009). The usual disclaimer applies.

Appendix

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

PF^T: Törnqvist price index measure of export prices of clothing, measured in foreign currency. $1986Q1 = 1$, cf. Eq. (8) in the text.

PF^F: Fischer price index measure of export prices of clothing, measured in foreign currency. $1986Q1 = 1$, cf. Eq. (9) in the text.

PF^G: Geometric mean price index (with constant weights) measure of export prices of clothing, measured in foreign currency. $1986Q1 = 1$, cf. Eq. (12) in the text.

PF_{ch}: China: Producer price index of clothing (from $1997Q1$) and consumer price index all products (from $1986Q1$), measured in 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 UK currency. *Source*: National statistics online, http://www.statistics.gov.uk/statbase/.

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

PF_{dk}: Denmark: Industrial output price index of clothing, measured in 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 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 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.

λ_j: Calculated relative price levels for country j based on nominal GDP and PPP adjusted real GDP, cf. Eq. (7) in the text. *Source*: IMF, the World Economic Outlook Database, http://www.imf.org/external/pubs/ft/weo/2009/01/weodata/.

ER: Chained geometric mean 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. 1986$Q1 = 1$. *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. 1986$Q1 = 1$. *Source*: Statistics Norway, QNA.

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

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