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## Estimating New Keynesian Import Price Models

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Research Department

Estimating New Keynesian import price models

by

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# Estimating New Keynesian import price models\*

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## Abstract

We estimate a range of New Keynesian import price models for Norway and the UK. Contrary to standard pass-through regression analysis, this approach allows us to make a distinction between the parameters in theoretical price-setting rules and parameters in the expectations mechanisms. We find positive and significant effects of expected future import price growth for Norway. The estimates for the UK do not lend much support to the hypothesis that price-setting rules are forward-looking. For both countries, the results favour a specification that incorporates both local- and producer currency pricing, but no effect of lagged import price growth. We find mixed evidence of pricing-to-market: only for the UK do the results suggest a role for domestic prices or costs in explaining import prices.

**JEL classification:** C32, C52, F41

**Keywords:** Import prices, exchange rate pass-through, New Keynesian open economy models, GMM

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## 1 INTRODUCTION

The development of import prices is important for policy-makers in small open economies and for inflation-targeting central banks in particular. Of special interest for monetary policy is the responsiveness of import prices to changes in the nominal exchange rate; that is, the degree of exchange rate pass-through.

New Keynesian open-economy models have become popular tools for analysing exchange rate pass-through and the effects of monetary policy. These models are dynamic stochastic general equilibrium (DSGE) models that allow for imperfect competition and nominal rigidities, typically in the form of sticky prices.<sup>1</sup> A number of seminal models in the New Keynesian literature (e.g., those in Obstfeld & Rogoff (1995) and Galí & Monacelli (2005)) assume that prices are set (and sticky) in the currency of the producer (so-called producer currency pricing, PCP). They also assume that the law of one price (LOP) holds at all times for traded goods. This implies that the exchange rate pass-through to import prices is complete and immediate, in keeping with the Mundell-Fleming model.

The LOP-PCP framework is rejected by empirical studies, however: they typically find that import prices respond incompletely to changes in the exchange rate (at least in the short run); see e.g., Campa & Goldberg (2005). Hence many New Keynesian models now allow for incomplete pass-through. Following Betts & Devereux (1996, 2000), the most common approach is to assume that international product markets are segmented and that prices are set and sticky in the currency of the importing country (local currency pricing, LCP).<sup>2</sup> This implies that import prices respond only gradually to changes in the exchange rate.

The standard LCP model assumes that the exporters' mark-ups are constant in the long-run (flexible price) equilibrium. This means that the long-run pass-through is complete. The model also implies that import prices are independent of the prices and costs in the importing country (at a constant exchange rate). Recently, several authors have allowed for a non-constant mark-up in the flexible-price equilibrium by assuming that (i) the demand elasticities facing a firm depend on the firm's price relative to those of its competitors (see e.g., Bergin & Feenstra, 2001; Gust & Sheets, 2006) or (ii) the distribution of traded goods to consumers requires local goods and services (see e.g., Corsetti & Dedola, 2005). These models imply that import prices depend on domestic prices or costs in the importing country. Following Bergin & Feenstra (2001), we will refer to models with this feature as 'pricing-to-market' models. Pricing-to-market provides an additional source of incomplete pass-through besides local currency price stickiness.

A key feature of import price equations in the New Keynesian models is that they are forward-looking: current import prices depend on expected future import prices and thus (implicitly) on the expected future path of the driving variables. Despite this feature, exchange rate pass-through has usually been estimated by regressing import prices on current and lagged values of the exchange rate and other variables believed to affect import prices.<sup>3</sup> If indeed price setters are forward-looking, the coefficients in such regressions will depend on the parameters in the price-setting rules and on the parameters in the expectations mechanisms. These mechanisms will in turn depend on the regime of monetary policy. The New Keynesian models thus predict that the coefficients in conventional pass-through regressions will vary with changes in the expectations mechanisms and with changes in the monetary policy regime; that is, the regressions are susceptible to the Lucas (1976) critique.

In this paper, we estimate and evaluate a range of New Keynesian import price equations using

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<sup>1</sup>New Keynesian open-economy models are also referred to as 'new open economy macroeconomics' (NOEM) models.

<sup>2</sup>A partial list of papers that have adopted the LCP framework includes Bacchetta & van Wincoop (2000), Devereux & Engel (2003), Chari et al. (2002), and Laxton & Pesenti (2003).

<sup>3</sup>A recent contribution is Campa & Goldberg (2005). See Goldberg & Knetter (1997) for a general discussion of pass-through regressions.

generalised method of moments (GMM). Contrary to standard pass-through regression analysis, this approach allows us to make a clear distinction between the parameters in the price-setting rules and the parameters in the expectations mechanisms. GMM has been widely used to estimate individual equations in New Keynesian DSGE models, including the New Keynesian Phillips Curve (e.g., Galí & Gertler, 1999; Galí et al., 2001; Batini et al., 2005), the Euler equation for output (e.g., Fuhrer & Rudebusch, 2004) and forward-looking monetary policy rules (e.g., Clarida et al., 1998).

An alternative approach is to estimate the parameters of New Keynesian import price equations as part of fully-specified DSGE models; see for example Bergin (2006), Adolfson et al. (2007) and Choudhri et al. (2005). An advantage of the general equilibrium approach is that cross-equation restrictions implied by the DSGE model are exploited in estimation. If the model is correctly specified, this increases efficiency relative to single-equation GMM estimation. At the same time, however, imposing the cross-equation restrictions could make the estimates of the parameters in the import price equation sensitive to misspecification in other parts of the model. Single-equation analysis does not rely on a specific completing model for the driving variables.

We use the Calvo (1983) model of random price adjustment as a unifying framework for deriving New Keynesian import price equations. This eases the comparison across model specifications. The overlapping contracts model of Taylor (1980) and the linear quadratic adjustment cost framework of Rotemberg (1982) would give rise to similar price dynamics (see Roberts, 1995). Building on previous empirical studies, we only consider models that allow for incomplete pass-through. We first derive and discuss a standard (purely forward-looking) LCP model where current import price growth depends on the expected future price growth and the level of import prices relative to foreign marginal costs measured in the importing country's currency. Consumers are assumed to have constant elasticity of substitution (CES) preferences over differentiated goods; that is, the elasticities of demand for individual goods are assumed to be constant. We extend the model to allow firms that do not re-optimize prices in a given period to index their prices to past import price growth and to allow a subset of foreign exporters to engage in PCP. Finally, we consider two pricing-to-market models: a model with translog preferences and a model with distribution costs. To our knowledge, no previous studies have estimated all these versions of the New Keynesian import price equation.<sup>4</sup>

The models are estimated on data from 1980Q1 to 2003Q1 for two small open economies: the UK and Norway. The GMM estimates obtained for the UK do not lend much support to the hypothesis that the price-setting rules are forward looking: the coefficient on expected future import price growth is either statistically insignificant, economically implausible, or both. The evidence of forward-looking price-setting is stronger for Norway: the coefficient on the forward-term is positive and, in most cases, statistically significant. For both countries, the estimation results favour a specification that allows for both PCP and LCP. By contrast, we find little evidence of indexation to past import price growth.

For Norway, the estimated coefficients on foreign costs and the pricing-to-market variables are statistically insignificant and close to zero in most cases. This contrasts with the results obtained for the UK: the coefficients on the foreign cost variables are statistically significant and, moreover, the pricing-to-market models suggest a role for domestic prices or costs in explaining import prices.

The remainder of the paper is organised as follows. Section 2 provides a survey of New Keynesian import price equations. Section 3 discusses the data and the econometric methodology, and section 4 presents the GMM estimation results. Section 5 concludes the paper.

## 2 NEW KEYNESIAN IMPORT PRICE EQUATIONS

This section derives the import price equations that provide the theoretical starting point for the empirical analysis. Throughout, the world is assumed to consist of two countries: home and for-

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<sup>4</sup>Freystätter (2003) estimates forward-looking import price equations using GMM on Finnish data. The equations allow for PCP and LCP, but they do not allow for inflation indexation or pricing-to-market.

eign. We model the price-setting decisions of foreign exporters that produce differentiated goods for sale in the home country. Product markets are characterised by monopolistic competition. We also assume that international product markets are segmented; that is, we allow firms to set distinct prices for the home and foreign markets. The price that firms would choose if prices had been perfectly flexible is referred to as the ‘frictionless’ price. The associated mark-up is referred to as the ‘frictionless’ mark-up. In section 2.1 we consider models where the frictionless mark-up is constant. In section 2.2 we consider models where the frictionless mark-up is a function of domestic prices in the importing country, so-called pricing-to-market models.

## 2.1 Models with a constant frictionless mark-up

This section derives three New Keynesian import price equations with a constant frictionless mark-up: a purely forward-looking model where all exporters engage in LCP (section 2.1.1), a ‘hybrid’ LCP model where firms that do not re-optimize prices in a given period index their prices to past import price growth (section 2.1.2) and a model where a share of exporters engages in PCP and a share engages in LCP (section 2.1.3).

### 2.1.1 A purely forward-looking model with local currency pricing

Foreign firms are indexed by  $i \in [0, 1]$ . Households in the home economy derive utility from the consumption of a composite foreign good  $Y_{F,t}$ , defined as a CES aggregate of differentiated goods<sup>5</sup>

$$Y_{F,t} \equiv \left[ \int_0^1 Y_{F,t}(i)^{\frac{\varepsilon-1}{\varepsilon}} di \right]^{\frac{\varepsilon}{\varepsilon-1}}, \quad (1)$$

where  $Y_{F,t}(i)$  is the imported quantity of good  $i$  in period  $t$  and  $\varepsilon > 1$  is the constant elasticity of substitution between the individual goods. The corresponding ideal price index is

$$P_{F,t} \equiv \left[ \int_0^1 P_{F,t}(i)^{1-\varepsilon} di \right]^{\frac{1}{1-\varepsilon}}, \quad (2)$$

where  $P_{F,t}(i)$  is the import price of good  $i$ , measured in the currency of the importing country. Cost-minimisation yields a conditional demand function for an individual imported good of the form

$$Y_{F,t}(i) = \left( \frac{P_{F,t}(i)}{P_{F,t}} \right)^{-\varepsilon} Y_{F,t}. \quad (3)$$

Price setting is staggered as in Calvo (1983); the probability that a firm is allowed to re-optimize its price in any given period is  $1 - \eta$ . The expected average time between price changes is thus  $1/(1 - \eta)$ . All exporters engage in LCP; that is, they set prices in the currency of the importing country. A firm that is allowed to re-optimize its price in period  $t$  sets the price  $\tilde{P}_{F,t}(i)$  to maximise

$$E_t \sum_{\tau=t}^{\infty} \eta^{\tau-t} D_{t,\tau} \left( \frac{\tilde{P}_{F,t}(i)}{S_{\tau}} - MC_{F,\tau}(i) \right) \left( \frac{\tilde{P}_{F,t}(i)}{P_{F,\tau}} \right)^{-\varepsilon} Y_{F,\tau}, \quad (4)$$

where  $D_{t,\tau}$  is a stochastic discount factor ( $D_{t,t} = 1$ ),  $S_{\tau}$  is the nominal exchange rate and  $MC_{F,\tau}(i)$  denotes the foreign firm’s marginal costs. In the following, we assume that all firms have access to the same technology and that all factors of production can be costlessly and instantaneously reallocated across firms. These assumptions imply that the marginal cost of firms that are allowed to reset prices is the same as the average marginal cost across all firms, that is,  $MC_{F,t}(i) = MC_{F,t}$ .

<sup>5</sup>The aggregate consumption index is a CES aggregate of the composite foreign good and a composite domestic good defined as a CES index of differentiated domestic goods. An alternative approach is to model a perfectly competitive firm that combines differentiated foreign goods using a CES technology.

Moreover, since all firms solve the same optimisation problem,  $\tilde{P}_{F,t}(i) = \tilde{P}_{F,t}$  for all firms that re-optimize in period  $t$ . The first-order condition can be written as

$$0 = E_t \sum_{\tau=t}^{\infty} \eta^{\tau-t} D_{t,\tau} Y_{F,\tau} \left( \frac{\tilde{P}_{F,t}}{P_{F,\tau}} \right)^{-\varepsilon} \left[ (1-\varepsilon) \frac{1}{S_{\tau}} + \varepsilon \frac{MC_{F,\tau}}{\tilde{P}_{F,t}} \right]. \quad (5)$$

If firms change prices on average every period (i.e.,  $\eta = 0$ ), the first-order condition collapses to the standard mark-up rule

$$P_{F,t} = \frac{\varepsilon}{\varepsilon-1} S_t MC_{F,t}. \quad (6)$$

Hence the frictionless price is a constant mark-up over foreign marginal costs, measured in the currency of the importing country. In this case, therefore, a change in the exchange rate is completely passed-through to import prices in the same period.

We now consider the general case where  $\eta > 0$ . Taking a log-linear approximation of (5) around a zero inflation deterministic steady-state we obtain

$$\tilde{p}_{F,t} - p_{F,t} = (1-\beta\eta) E_t \sum_{\tau=t}^{\infty} (\beta\eta)^{\tau-t} (s_{\tau} + mc_{F,\tau} - p_{F,\tau}), \quad (7)$$

where  $\beta$  is the steady-state value of the stochastic discount factor ( $0 < \beta < 1$ ) and lower case letters denote the percentage deviation of the original variable from its deterministic steady-state value. The optimal price thus depends on a weighted average of expected future marginal costs measured in the importer's currency. It follows that changes in costs or the exchange rate have stronger short-run effects on import prices if the shocks are expected to be long-lasting than if they are expected to be reversed soon. A testable implication of this model is thus that the parameters in 'backward-looking' pass-through regressions will not be invariant to changes in the stochastic processes for the exchange rate and foreign marginal costs.

Quasi-differentiation of (7) yields

$$\tilde{p}_{F,t} - p_{F,t} = (1-\beta\eta) (s_t + mc_{F,t} - p_{F,t}) + \beta\eta E_t (\tilde{p}_{F,t+1} - p_{F,t+1}). \quad (8)$$

The aggregate price index can be written as (see Woodford, 2003, p. 178)

$$P_{F,t}^{1-\varepsilon} = (1-\eta) \tilde{P}_{F,t}^{1-\varepsilon} + \eta P_{F,t-1}^{1-\varepsilon}. \quad (9)$$

Log-linearisation around a zero inflation steady-state yields

$$0 = (1-\eta) (\tilde{p}_{F,t} - p_{F,t}) - \eta \Delta p_{F,t}, \quad (10)$$

where  $\Delta$  is the difference operator ( $\Delta x_t \equiv x_t - x_{t-1}$ ). By substituting in from (8) we obtain

$$\Delta p_{F,t} = \beta E_t \Delta p_{F,t+1} - \frac{(1-\beta\eta)(1-\eta)}{\eta} (p_{F,t} - s_t - mc_{F,t}). \quad (11)$$

The equilibrium condition relates current import price inflation to expected import price inflation in the next period and to deviations of current import prices  $p_{F,t}$  from the local currency value of foreign marginal costs  $s_t + mc_{F,t}$ . The effect of the forcing term  $(p_{F,t} - s_t - mc_{F,t})$  is decreasing in  $\eta$ : a higher degree of price stickiness implies lower pass-through of changes in the exchange rate and marginal costs in the short run. In the long-run, however, there is complete pass-through of permanent changes in the exchange rate and marginal costs.

Figure 1 illustrates the response of import prices to an unexpected 1% depreciation of the nominal exchange rate for different values of the price stickiness parameter  $\eta$ . The period length is assumed to be one quarter. The responses are conditional on given values of foreign marginal costs,



and the discount factor  $\beta$  is set to 0.99. The exchange rate is assumed to follow a random walk. Hence, the exchange rate shock is perceived to be permanent and the long-run exchange rate pass-through is 100%. A higher degree of price stickiness reduces the short-run pass-through and also makes it more gradual. If firms change prices on average every two quarters ( $\eta = 0.5$ ), the model predicts that the pass-through is 50% in the first quarter and near complete after one year. If the average time between price changes is four quarters ( $\eta = 0.75$ ), on the other hand, the immediate pass-through is about 25%, increasing to 75% after about four quarters. In the case where firms only adjust prices on average every eight quarters ( $\eta = 0.875$ ), the short-run pass-through is 12.5%.

Figure 2 illustrates the dependence of pass-through on the (expected) persistence of the exchange rate. The exchange rate is now assumed to follow a first-order autoregressive process:  $s_t = \tau s_{t-1} + \varepsilon_{s,t}$ , where  $\varepsilon_{s,t}$  is white noise. We plot the responses to a purely temporary shock ( $\tau = 0$ ), a temporary but highly persistent shock ( $\tau = 0.95$ ) and a permanent shock ( $\tau = 1$ ), where the latter corresponds to the random walk assumption above. The values of  $\beta$  and  $\eta$  are kept fixed at 0.99 and 0.75, respectively. The figure illustrates that the degree of pass-through is increasing in the expected persistence of the exchange rate. In the case of a purely temporary shock, the impact effect on import prices is negligible. In response to the temporary but persistent shock, exchange rate pass-through is around 20% in the first quarter, compared to 25% in the case of a permanent exchange rate shock. This illustrates that, in general, the comovement of exchange rates and import prices depends on the nature of the shock that causes the variables to move. The long-run pass-through of the exchange rate shock is complete in all cases in the sense that the long-run effects on the exchange rate and import prices are the same (i.e., zero unless the exchange rate shock is perceived to be permanent).

### 2.1.2 A hybrid model with local currency pricing

In the literature on the New Keynesian Phillips Curve it is common to consider ‘hybrid’ specifications with both forward-looking and backward-looking components. In the model derived above, firms that are not allowed to re-optimize prices in period  $t$  charge a price equal to the price charged in period  $t - 1$ . Here, following Smets & Wouters (2002), we assume instead that firms that are not allowed to re-optimize prices in period  $t$  update their prices according to the partial indexation rule

$$P_{F,t}(i) = \left( \frac{P_{F,t-1}}{P_{F,t-2}} \right)^\chi P_{F,t-1}(i), \quad (12)$$

where  $\chi \in [0, 1]$  is the indexation parameter. If  $\chi = 1$  (full indexation), this scheme collapses to the dynamic indexation scheme considered by Christiano et al. (2005). The problem facing a firm that is allowed to re-optimize in period  $t$  is now

$$\max_{\tilde{P}_{F,t}(i)} E_t \sum_{\tau=t}^{\infty} \eta^{\tau-t} D_{t,\tau} \left( \frac{\tilde{P}_{F,t}(i) \left( \frac{P_{F,\tau-1}}{P_{F,t-1}} \right)^\chi}{S_\tau} - MC_{F,\tau} \right) \left( \frac{\tilde{P}_{F,t}(i) \left( \frac{P_{F,\tau-1}}{P_{F,t-1}} \right)^\chi}{P_{F,\tau}} \right)^{-\varepsilon} Y_{F,\tau}, \quad (13)$$

and the aggregate price index is

$$P_{F,t}^{1-\varepsilon} = (1-\eta) \tilde{P}_{F,t}^{1-\varepsilon} + \eta \left( \left( \frac{P_{F,t-1}}{P_{F,t-2}} \right)^\chi P_{F,t-1} \right)^{1-\varepsilon}. \quad (14)$$

The log-linearised equilibrium condition for aggregate import price growth becomes

$$\Delta p_{F,t} = \frac{\beta}{1+\beta\chi} E_t \Delta p_{F,t+1} + \frac{\chi}{1+\beta\chi} \Delta p_{F,t-1} - \frac{(1-\eta\beta)(1-\eta)}{\eta(1+\beta\chi)} (p_{F,t} - s_t - mc_{F,t}). \quad (15)$$

Hence the lagged growth in import prices enters the equation. The weight on lagged import price growth is increasing in the degree of indexation. However, the maximum weight on lagged price growth (obtained for  $\chi = 1$ ) is  $1/(1 + \beta) \simeq 0.5$  for values of  $\beta$  close to unity.

Figure 3 shows the response of import prices to a 1% permanent exchange rate shock for different values of the indexation parameter  $\chi$  when  $\beta = 0.99$  and  $\eta = 0.75$ . Varying the degree of indexation has a relatively small effect on the short-run pass-through. The differences become more pronounced after about four quarters, however. In the medium run, the pass-through is higher for higher values of the indexation parameter; that is, pass-through is higher the larger is the weight on lagged import price growth in the import price equation. With full indexation ( $\chi = 1$ ) the import price overshoots the flexible price level and pass-through exceeds 100% in the medium run.

### 2.1.3 Models with local- and producer currency pricing

Evidence on the currency denomination of foreign trade suggests that a substantial share of imports are invoiced in the exporter's currency (see e.g., Bekx, 1998).<sup>6</sup> This motivates extending the model to allow a subset  $\phi$  of foreign firms to engage in PCP and a subset  $1 - \phi$  to engage in LCP, as in e.g., Betts & Devereux (1996, 2000), Bergin (2006) and Choudhri et al. (2005). Following Bergin (2006) and Choudhri et al. (2005), we assume that PCP firms are able to segment markets, but choose to set prices in their own currency.<sup>7</sup>

Admittedly, a limitation of our framework is that the fraction of price setters engaging in PCP is assumed to be constant and independent of the other parameters in the model. A recent literature considers the optimal choice of invoicing currency in the context of NOEM models (e.g., Devereux et al., 2004; Bacchetta & van Wincoop, 2005; Goldberg & Tille, 2005). The choice between LCP and PCP is found to depend on several factors, including the exporting firm's market share in the foreign market, the degree of substitutability between foreign and domestic goods and relative monetary stability. Thus, the parameter  $\phi$  could vary over time.

Let  $P_{F,t}^P$  and  $P_{F,t}^L$  denote the prices set by PCP firms and LCP firms respectively. The aggregate import price index can then be written

$$P_{F,t} = \left[ \phi (S_t P_{F,t}^P)^{1-\varepsilon} + (1-\phi) (P_{F,t}^L)^{1-\varepsilon} \right]^{\frac{1}{1-\varepsilon}}. \quad (16)$$

We assume that the frequency of price adjustment  $\eta$  is the same for LCP and PCP firms.<sup>8</sup> The LCP firms' price-setting problem is the same as in equation (4) above, while the optimisation problem facing a PCP firm that re-optimises in period  $t$  is

$$\max_{\tilde{P}_{F,t}^P(i)} E_t \sum_{\tau=t}^{\infty} \eta^{\tau-t} D_{t,\tau} \left( \tilde{P}_{F,t}^P(i) - MC_{F,\tau} \right) \left( \frac{S_{\tau} \tilde{P}_{F,t}^P(i)}{P_{F,\tau}} \right)^{-\varepsilon} Y_{F,\tau}. \quad (17)$$

The log-linearised equilibrium conditions for LCP firms and PCP firms are

$$\Delta p_{F,t}^L = \beta E_t \Delta p_{F,t+1}^L - \frac{(1-\beta\eta)(1-\eta)}{\eta} (p_{F,t}^L - s_t - mc_{F,t}) \quad (18)$$

$$\Delta p_{F,t}^P = \beta E_t \Delta p_{F,t+1}^P - \frac{(1-\beta\eta)(1-\eta)}{\eta} (p_{F,t}^P - mc_{F,t}) \quad (19)$$

<sup>6</sup>The share of UK imports invoiced in sterling in the years 1999 to 2002 was approximately 40%. See <http://customs.hmrc.gov.uk/>.

<sup>7</sup>Betts & Devereux (1996, 2000) assume that PCP firms are unable to segment markets internationally and hence cannot price discriminate.

<sup>8</sup>This assumption has some empirical support. Using micro data for traded goods prices at the docks for the US, Gopinath & Rigobon (2006) find that the stickiness of prices invoiced in foreign currencies in terms of foreign currency is similar to the stickiness of prices invoiced in dollars in terms of dollars.

Using the definition of the aggregate price index in (16) we obtain the following expression for aggregate import price inflation

$$\begin{aligned}\Delta p_{F,t} &= \phi(\Delta s_t + \Delta p_{F,t}^P) + (1 - \phi)\Delta p_{F,t}^L \\ &= \beta E_t \Delta p_{F,t+1} - \frac{(1 - \beta\eta)(1 - \eta)}{\eta}(p_{F,t} - s_t - mc_{F,t}) + \phi(\Delta s_t - \beta E_t \Delta s_{t+1})\end{aligned}\quad (20)$$

When some firms engage in PCP ( $\phi > 0$ ), the aggregate price equation is augmented with the current change in the exchange rate and the expected change in the exchange rate in the next period. The latter term reflects that PCP firms set prices according to expected future price growth measured in their own currency ( $\Delta p_{F,t+1}^P$ ) rather than import price growth measured in the importing country's currency ( $\Delta p_{F,t+1}^P + \Delta s_{t+1}$ ) which appears in the definition of aggregate import price growth ( $\Delta p_{F,t+1}$ ).

Figure 4 shows the response of import prices to a permanent 1% exchange rate shock for different values of the share of PCP price setters  $\phi$ . The remaining parameters take the values  $\beta = 0.99$  and  $\eta = 0.75$ . The short-run pass-through is increasing in the share of exporters that engages in PCP. If all firms engage in PCP ( $\phi = 1$ ), the pass-through is complete at all horizons (this holds regardless of the degree of price stickiness). The short-run pass-through is (as noted above) about 25% in the absence of PCP firms ( $\phi = 0$ ). When the share of PCP firms is 0.5 ( $\phi = 0.5$ ), the short-run pass-through increases to 62.5%.

The model can be extended to allow for inflation indexation in the same manner as above. We assume that PCP and LCP firms index their prices to last period's aggregate inflation rate, measured in the exporting and the importing country's currency, respectively. That is, firms that do not re-optimize in period  $t$  update their prices according to the rules

$$P_{F,t}^P(i) = \left(\frac{P_{F,t-1} S_{t-2}}{P_{F,t-2} S_{t-1}}\right)^\chi P_{F,t-1}^P(i) \text{ and } P_{F,t}^L(i) = \left(\frac{P_{F,t-1}}{P_{F,t-2}}\right)^\chi P_{F,t-1}^L(i) \quad (21)$$

The equation describing the evolution of aggregate import price inflation is now<sup>9</sup>

$$\begin{aligned}\Delta p_{F,t} &= \frac{\beta}{1 + \beta\chi} E_t \Delta p_{F,t+1} + \frac{\chi}{1 + \beta\chi} \Delta p_{F,t-1} - \frac{(1 - \eta\beta)(1 - \eta)}{\eta(1 + \beta\chi)}(p_{F,t} - s_t - mc_{F,t}) \\ &\quad + \phi \left( \Delta s_t - \frac{\beta}{1 + \beta\chi} E_t \Delta s_{t+1} - \frac{\chi}{1 + \beta\chi} \Delta s_{t-1} \right)\end{aligned}\quad (22)$$

Thus, the model with both LCP- and PCP firms and inflation indexation ascribes separate roles for both lagged import price growth and the lagged change in the exchange rate in determining import prices.

## 2.2 Pricing-to-market models

In the models considered so far, the frictionless mark-up is constant. Hence, the only mechanism generating incomplete pass-through is local currency price stickiness. In this section we consider two New Keynesian open-economy models that have the feature that the elasticity of demand perceived by the exporter is non-constant: a model with translog preferences due to Bergin & Feenstra (2001) and the distribution cost model due to Corsetti & Dedola (2005). In these models, the frictionless mark-up is a function of domestic prices in the importing country. This creates a scope for price discrimination and acts to reduce the degree of pass-through. Adopting the terminology in Bergin & Feenstra (2001), we refer to models with this property as pricing-to-market models.

<sup>9</sup>Since we are assuming that the indexation parameter is the same for PCP- and LCP firms, we would obtain the same expression for aggregate import price inflation if we instead allowed PCP- and LCP firms to index their prices to last period's aggregate PCP- and LCP inflation rate, respectively.

Several previous empirical studies have found evidence of long-run pricing-to-market, also in small open economies (see e.g., Menon, 1995; Naug & Nymoen, 1996; Herzberg et al., 2003; Kongsted, 2003).

### 2.2.1 A model with translog preferences

Following Bergin & Feenstra (2001), we assume that there are a large number  $N$  of varieties of goods available in the domestic market. Of these, goods indexed  $i = 1, \dots, N_H$  are produced by domestic firms, and goods indexed  $i = N_H + 1, \dots, N$  are produced by foreign firms. The aggregate ideal price index,  $P_t$ , implied by the translog expenditure function is<sup>10</sup>

$$\ln P_t \equiv \sum_{i=1}^N \alpha_i \ln P_t(i) + \frac{1}{2} \sum_{i=1}^N \sum_{j=1}^N \gamma_{ij} \ln P_t(i) \ln P_t(j), \quad (23)$$

where  $\gamma_{ij} = \gamma_{ji}$ . The prices of the imported goods are import prices (i.e., measured ‘at the docks’). In the special case where all goods enter the expenditure function symmetrically, the parameters become

$$\alpha_i = \frac{1}{N}, \gamma_{ii} = -\frac{\gamma}{N}, \gamma_{ij} = \frac{\gamma}{N(N-1)} \text{ for } i \neq j, \quad (24)$$

where  $\gamma > 0$ . With these restrictions, the expenditure function is homogenous of degree one:  $\sum_{i=1}^N \alpha_i = 1$  and  $\sum_{i=1}^N \gamma_{ij} = \sum_{j=1}^N \gamma_{ij} = 0$ . The demand for good  $i$  is given by

$$Y_t(i) = \psi_t(i) \frac{P_t Y_t}{P_t(i)}, \quad (25)$$

where  $Y_t$  is the aggregate demand for goods in the home country and  $\psi_t(i)$  is the expenditure share on good  $i$ , defined as

$$\psi_t(i) = \frac{\partial \ln P_t}{\partial \ln P_t(i)} = \alpha_i + \sum_{j=1}^N \gamma_{ij} \ln P_t(j). \quad (26)$$

The elasticity of demand for each good is then

$$\varepsilon_t(i) = 1 - \frac{\partial \ln \psi_t(i)}{\partial \ln P_t(i)} = 1 - \frac{\gamma_{ii}}{\psi_t(i)}, \quad \gamma_{ii} < 0. \quad (27)$$

The demand elasticity depends negatively on the price of competing products. Hence, a fall in the competitors’ prices will lead to a reduction in the desired mark-up.<sup>11</sup> If prices are flexible, the optimal price set by a foreign firm satisfies

$$P_t(i) = \frac{\varepsilon_t(i)}{\varepsilon_t(i) - 1} S_t MC_{F,t}, \quad i = N_H + 1, \dots, N, \quad (28)$$

which, using the expression for the demand elasticity in (27), can be written as

$$\frac{S_t MC_{F,t}}{P_t(i)} \left( 1 - \frac{\psi_t(i)}{\gamma_{ii}} \right) - 1 = 0. \quad (29)$$

<sup>10</sup>There is no closed form solution for the direct utility function. See Feenstra (2003) for details.

<sup>11</sup>This is also a property of the more general preference specification proposed by Kimball (1995) that has become popular in the literature on the New Keynesian Phillips Curve (see e.g., Eichenbaum & Fisher, 2007; Woodford, 2005). See Gust & Sheets (2006) for an open-economy application.

Following Bergin & Feenstra (2001), the left-hand side of the equation can be approximated as<sup>12</sup>

$$\begin{aligned} \frac{S_\tau MC_{F,\tau}}{P_t(i)} \left(1 - \frac{\Psi_\tau(i)}{\gamma_{ii}}\right) - 1 &\approx \ln \left( \frac{S_\tau MC_{F,\tau}}{P_t(i)} \left(1 - \frac{\Psi_\tau(i)}{\gamma_{ii}}\right) \right) \\ &\approx -\frac{\Psi_\tau(i)}{\gamma_{ii}} + \ln \left( \frac{S_\tau MC_{F,\tau}}{P_t(i)} \right). \end{aligned} \quad (30)$$

This allows us to rewrite the first-order condition as

$$\ln P_t(i) = \frac{1}{2}\gamma + \frac{1}{2}(\ln S_t + \ln MC_{F,t}) + \frac{1}{2} \sum_{j \neq i}^N \frac{1}{N-1} \ln P_t(j), \quad (31)$$

where we have substituted in for the expenditure share  $\Psi_t(i)$  in (26). The optimal price puts a weight of one half on marginal costs measured in the importing country's currency and one half on the competitors' prices. Holding all other prices fixed, therefore, a 1% depreciation will increase the optimal price by 0.5%. This is an important characteristic of the translog preference specification. Imposing symmetry we can rewrite (31) as

$$p_{F,t} = \frac{N-1}{N+N_H-1} (s_t + mc_{F,t}) + \frac{N_H}{N+N_H-1} p_{H,t}, \quad (32)$$

where lower case letters represent percent deviations from the deterministic steady state, and  $p_{H,t}$  and  $p_{F,t}$  denote the common price set by the domestic and foreign firms, respectively.

Equation (32) shows that the optimal frictionless price is a function of the price of import-competing goods. Translog preferences are thus a source of strategic complementarity in price-setting (see Woodford, 2003, p. 161). Holding marginal costs and the prices of import-competing goods fixed, the exchange rate pass-through is incomplete. Intuitively, a foreign firm will take into account that, if the prices of importing-competing goods are kept constant, an increase in its price will cause demand to become more elastic. This lowers the desired mark-up and reduces the incentive to raise the price following a depreciation of the exchange rate relative to the CES case. Rather than passing it through completely to the buyer, the exporter will absorb part of an exchange rate depreciation in her mark-up. The degree of (conditional) pass-through is inversely related to the share of domestic firms in the importing country. Note, however, that the import price equation is linearly homogenous in marginal costs and the price of import-competing goods. In a fully-specified general equilibrium model that satisfies nominal neutrality, the unconditional pass-through of an exogenous shock to the nominal exchange rate (i.e., taking into account the effects of the exchange rate change on domestic prices) is complete.

In Bergin & Feenstra (2001), prices are set in two-period overlapping contracts. Here, we combine the translog preference specification with the Calvo pricing model. As before, we assume that a fraction  $1 - \eta$  of firms are allowed to re-optimize prices in a given period. Following Bergin & Feenstra (op.cit) we assume that all firms engage in LCP. The price-setting problem of a foreign firm that is allowed to reset prices in period  $t$  is

$$\max_{\tilde{P}_t(i)} E_t \sum_{\tau=t}^{\infty} \eta^{\tau-t} D_{t,\tau} \left( \frac{\tilde{P}_t(i)}{S_\tau} - MC_{F,\tau} \right) Y_\tau(i), \quad (33)$$

and the optimal price  $\tilde{P}_t(i)$  satisfies

$$E_t \sum_{\tau=t}^{\infty} \eta^{\tau-t} D_{t,\tau} \frac{P_\tau Y_\tau}{S_\tau} \left( \frac{S_\tau MC_{F,\tau}}{\tilde{P}_t(i)} \left(1 - \frac{\Psi_\tau(i)}{\gamma_{ii}}\right) - 1 \right) = 0. \quad (34)$$

<sup>12</sup>The first approximation holds if  $\frac{S_\tau MC_{F,\tau}}{P_t(i)} \left(1 - \frac{\Psi_\tau(i)}{\gamma_{ii}}\right)$  is close to unity (meaning that the price is not too different from the optimal frictionless price), and the second approximation holds if  $\Psi_t(i)$  is small (see Bergin & Feenstra, 2001).

Taking a log-linear approximation of the first-order condition around a zero inflation steady state, we obtain

$$\tilde{p}_t(i) = \frac{1}{2}(1 - \beta\eta)E_t \sum_{\tau=t}^{\infty} (\beta\eta)^{\tau-t} \left( s_{\tau} + mc_{F,\tau} + \frac{N_H}{N-1} p_{H,\tau} + \frac{N - N_H - 1}{N-1} p_{F,\tau} \right). \quad (35)$$

Imposing symmetry and assuming that the number of foreign exporters ( $N - N_H$ ) is large, the aggregate import price index can be written

$$p_{F,t} = (1 - \eta)\tilde{p}_t + \eta p_{F,t-1}, \quad (36)$$

and the equation for aggregate import price inflation becomes

$$\Delta p_{F,t} = \beta E_t \Delta p_{F,t+1} - \frac{(1 - \beta\eta)(1 - \eta)}{\eta} \frac{1}{2} \frac{N + N_H - 1}{N - 1} \left( p_{F,t} - \frac{N - 1}{N + N_H - 1} (s_t + mc_{F,t}) - \frac{N_H}{N + N_H - 1} p_{H,t} \right). \quad (37)$$

The effect of the forcing term is smaller than in the CES case. A foreign firm that contemplates raising the price of its product will take into account that, since not all foreign firms are allowed to change prices in the short run, an increase in its price will cause demand to become more elastic.<sup>13</sup> This reduces the incentive to raise the price. The model with translog preferences thus requires a smaller amount of nominal rigidities to generate slow exchange rate pass-through than the models with CES preferences. Notice that this holds also when there are no import-competing firms in the domestic market (i.e.,  $N_H = 0$ ), in which case the coefficient on the forcing term is reduced by a factor of two relative to the CES case.

Figure 5 shows the response of import prices to a permanent 1% shock to the exchange rate for different values of the share of import-competing firms ( $N_H/N$ ). The price of import-competing goods is held fixed, and the remaining parameters take the values  $\beta = 0.99$  and  $\eta = 0.75$ . The figure illustrates that, with translog preferences and a positive share of import-competing firms, the long-run pass-through is incomplete as long as domestic prices  $p_{H,t}$  are constant. If foreign and domestic firms have an equal market share ( $N_H/N = 0.5$ ), the long-run pass-through is 67%; the long-run pass-through is 53% if the market share of domestic firms is 0.9 ( $N_H/N = 0.9$ ). In the case of no import-competing firms ( $N_H/N = 0$ ), the long-run pass-through is complete. However, even in the absence of import-competing firms, short- and medium-run pass-through is still lower than in the models that assume a constant demand elasticity.

## 2.2.2 A model with distribution costs

The final model we consider is the distribution cost model in Corsetti & Dedola (2005). The key assumption is that the distribution of traded goods requires the input of local goods and services such as e.g., transportation, marketing and retail services. The distribution technology is Leontief; the distribution of one unit of imported goods to domestic households requires the input of  $\mu$  units of local goods and services. If the distribution sector is perfectly competitive, the price paid by home consumers for a unit of the imported good (the ‘retail’ price),  $\bar{P}_{F,t}$ , is

$$\bar{P}_{F,t} = P_{F,t} + \mu P_{N,t}, \quad (38)$$

where  $P_{F,t}$  denotes the import price and  $P_{N,t}$  is the price of the local goods and services that are required to distribute the good in the importing country.

Consumers are assumed to have CES preferences over differentiated goods. When setting prices, the exporter takes into account that the price paid by the consumers depends on the distri-

<sup>13</sup>The effect on the demand elasticity of the fact that domestic firms’ prices are kept fixed is captured by the forcing term.

bution costs. In the flexible price case the exporter's price-setting problem is

$$\max_{P_{F,t}(i)} \left( \frac{P_{F,t}(i)}{S_t} - MC_{F,t} \right) \left( \frac{P_{F,t}(i) + \mu P_{N,t}}{\bar{P}_{F,t}} \right)^{-\varepsilon} Y_{F,t}, \quad (39)$$

and the optimal price satisfies

$$P_{F,t} = \frac{\varepsilon}{\varepsilon - 1} S_t MC_{F,t} + \frac{\mu}{\varepsilon - 1} P_{N,t}. \quad (40)$$

This equation shows that, in the presence of distribution costs, the optimal price varies across destination markets. Moreover, the desired mark-up is a function of the exchange rate and the distribution costs in the importing country. This can be seen more clearly if we rewrite (40) as

$$P_{F,t} = \frac{\varepsilon}{\varepsilon - 1} S_t MC_{F,t} \left( 1 + \frac{\mu}{\varepsilon} \frac{P_{N,t}}{S_t MC_{F,t}} \right). \quad (41)$$

The desired mark-up is a decreasing function of the exchange rate. As in the model with translog preferences, the exporter absorbs part of an exchange rate movement in her mark-up. This lowers the degree of exchange rate pass-through to import prices. Taking a log-linear approximation around a deterministic steady-state we obtain

$$p_{F,t} = \frac{1}{1 + \zeta(mk_F - 1)} (s_t + mc_{F,t}) + \frac{\zeta(mk_F - 1)}{1 + \zeta(mk_F - 1)} p_{N,t}, \quad (42)$$

where  $\zeta$  is the steady-state share of distribution costs in the retail price of imports and  $mk_F$  is the steady-state mark-up, that is

$$\zeta = \frac{\mu P_N}{\bar{P}_F} \quad \text{and} \quad mk_F = \frac{\varepsilon}{\varepsilon - 1} \left( 1 + \frac{\mu}{\varepsilon} \frac{P_N}{S MC_F} \right). \quad (43)$$

The long-run exchange rate pass-through is decreasing in the share of distribution costs in the retail price of imports. Note that the weights on foreign marginal costs and domestic distribution prices in the expression for the optimal frictionless price in (42) sum to unity. Hence, the model satisfies long-run price homogeneity. In the benchmark calibration in Corsetti et al. (2005),  $\zeta = 0.5$  and  $mk_F \simeq 1.15$ , which implies that the long-run pass-through coefficient is equal to 0.93. Thus, the long-run pass-through is close to being complete, even when the share of distribution costs in the retail price of imports is large.

As a next step, we derive a dynamic import price equation using the Calvo assumption.<sup>14</sup> An LCP exporter who is allowed to reset price in period  $t$  faces the following optimisation problem

$$\max_{\tilde{P}_{F,t}(i)} E_t \sum_{\tau=t}^{\infty} \eta^{\tau-t} D_{t,\tau} \left( \frac{\tilde{P}_{F,t}(i)}{S_\tau} - MC_{F,\tau} \right) \left( \frac{\tilde{P}_{F,t}(i) + \mu P_{N,\tau}}{\bar{P}_{F,\tau}} \right)^{-\varepsilon} Y_{F,\tau}. \quad (44)$$

Imposing symmetry, the first-order condition is

$$E_t \sum_{\tau=t}^{\infty} \eta^{\tau-t} D_{t,\tau} Y_{F,\tau} \left( \frac{\tilde{P}_{F,t} + \mu P_{N,\tau}}{\bar{P}_{F,\tau}} \right)^{-\varepsilon} \left( \frac{1}{S_\tau} - \varepsilon \frac{\left( \frac{\tilde{P}_{F,t}}{S_\tau} - MC_{F,\tau} \right)}{\tilde{P}_{F,t} + \mu P_{N,\tau}} \right) = 0. \quad (45)$$

The aggregate price index is the same as before, and the log-linearised import price equation be-

<sup>14</sup>Choudhri et al. (2005) combine distribution costs with the Calvo assumption, but specify distribution costs in terms of labour services rather than non-traded goods. Corsetti et al. (2005) derive an import price equation similar to ours assuming quadratic adjustment costs in pricing.

comes

$$\Delta p_{F,t} = \beta E_t \Delta p_{F,t+1} - \frac{(1-\eta\beta)(1-\eta)}{\eta} \left( p_{F,t} - \frac{1}{1+\zeta(mk_F-1)} (s_t + mc_{F,t}) - \frac{\zeta(mk_F-1)}{1+\zeta(mk_F-1)} p_{N,t} \right). \quad (46)$$

The import price equation takes the same general form as above: current import price inflation depends on expected future inflation and the deviation of the current price from the frictionless price. The coefficient on the forcing term is the same as in the LCP model with CES preferences.

Figure 6 shows the response of import prices to a 1% permanent exchange rate shock for different values of  $\zeta$ . The price of domestic goods and services is held fixed, and  $\beta = 0.99$  and  $\eta = 0.75$ . As is evident from the graph, the degree of pass-through is not very sensitive to the size of the share of distribution costs in the retail price of imports. For the parameter values considered here, the pass-through is still close to 90% after twenty periods, even when the share of distribution costs is as high as 0.75.

We can extend the distribution cost model to allow for PCP price setters and inflation indexation.<sup>15</sup> The equation for aggregate import price inflation then takes the same form as equation (22) above, except that the forcing term is now  $p_{F,t} - \frac{1}{1+\zeta(mk_F-1)} (s_t + mc_{F,t}) - \frac{\zeta(mk_F-1)}{1+\zeta(mk_F-1)} p_{N,t}$ . An interesting feature of the pricing-to-market models is that when the share of PCP firms is sufficiently large, the short-run pass-through of a permanent exchange rate depreciation will exceed the long-run (conditional) pass-through. The intuition is as follows: with pricing-to-market, the frictionless mark-up falls in response to an exchange rate depreciation. Because prices are sticky in the exporter's currency, it takes time before this mark-up adjustment is fully reflected in the export price. Import prices respond instantly to the exchange rate depreciation, however. By similar reasoning, in a pure PCP model with pricing-to-market, the short-run pass-through will be *decreasing* in the frequency of price adjustment: the lower is the degree of price stickiness, the stronger is the short-run effect of an exchange rate change on export prices and hence, the weaker is the effect on import prices.

### 3 EMPIRICAL IMPLEMENTATION

We estimate the New Keynesian import price equations using data for two small open economies: the UK and Norway. In this section we discuss the data used in the empirical analysis (section 3.1), the econometric model specification (section 3.2), the GMM estimator (section 3.3) and the implications of the New Keynesian import price model for the cointegration properties of the variables (section 3.4).

#### 3.1 Data

The data are seasonally adjusted, quarterly series covering the period 1980Q1–2003Q1 (see appendix A for details on the variable definitions and sources). The import price series is an index of import prices of manufactures and the exchange rate is a broad trade-weighted nominal exchange rate.

To implement the price-setting rules empirically, we need a measure of the marginal costs of foreign firms. Following Batini et al. (2005), we assume that the marginal cost of producing value added output depends on unit labour costs and the price of raw materials input. Specifically, we assume that the log-linearised equation for foreign marginal costs is given by

$$mc_{F,t} = (1 - \delta)ulc_{F,t} + \delta p_{COM,t}, \quad (47)$$

<sup>15</sup>The model in Choudhri et al. (2005) has distribution costs and a combination of LCP- and PCP price setters. Laxton & Pesenti (2003) combine distribution costs and a specification of adjustment costs in pricing that implies a linearised equation for aggregate import price inflation that is observationally equivalent to a Calvo model with full indexation.



where  $ulc_{F,t}$  denotes foreign unit labour costs,  $p_{COM,t}$  denotes the price of raw materials and  $\delta$  is a parameter to be estimated. The dataseries for foreign unit labour costs are constructed using data for domestic unit labour costs and trade-weighted relative unit labour costs in manufacturing industries. As a proxy for the price of raw materials we use an index of the world price of metals constructed by the IMF. The commodity price index is converted into foreign currency using the trade-weighted nominal exchange rate.

To estimate the pricing-to-market models we need a measure of domestic prices; in the distribution cost model import prices depend on the price of local goods and services, and the model with translog preferences predicts that import prices depend on the prices of import competing products. We use domestic unit labour costs as a proxy for distribution costs in the importing country. As a proxy for the price of import-competing goods we use a producer price index for manufactures sold in the domestic market.

### 3.2 Econometric model specification

The import price equations derived above can be obtained as restricted versions of the following general specification:

$$\begin{aligned} \Delta p_{F,t} = & \alpha_1 E_t \Delta p_{F,t+1} + \alpha_2 \Delta p_{F,t-1} + \alpha_3 (\Delta s_t - \alpha_1 E_t \Delta s_{t+1} - \alpha_2 \Delta s_{t-1}) \\ & + \alpha_4 (s_t + ulc_{F,t} - p_{F,t}) + \alpha_5 (s_t + p_{COM,t} - p_{F,t}) + \alpha_6 (p_{H,t} - p_{F,t}) + u_t, \end{aligned} \quad (48)$$

where the error term  $u_t$  is assumed to be a mean zero, serially uncorrelated process and can be interpreted as arising from e.g., exogenous variations in the desired mark-up (see e.g., Adolfson et al., 2007).  $\Upsilon = \{\alpha_1, \alpha_2, \alpha_3, \alpha_4, \alpha_5, \alpha_6\}$  depend on the underlying structural parameters:

$$\begin{aligned} \alpha_1 &= \frac{\beta}{1 + \beta\chi} \\ \alpha_2 &= \frac{\chi}{1 + \beta\chi} \\ \alpha_3 &= \phi \\ \alpha_4 &= \frac{1}{1 + \beta\chi} \frac{(1 - \beta\eta)(1 - \eta)}{\eta} \xi \rho (1 - \delta) \\ \alpha_5 &= \frac{1}{1 + \beta\chi} \frac{(1 - \beta\eta)(1 - \eta)}{\eta} \xi \rho \delta \\ \alpha_6 &= \frac{1}{1 + \beta\chi} \frac{(1 - \beta\eta)(1 - \eta)}{\eta} \xi (1 - \rho) \end{aligned}$$

When the frictionless mark-up is constant  $\xi = \rho = 1$ . In the distribution cost model,  $p_{H,t}$  is the price of local goods and services used in the distribution of imported goods,  $\xi = 1$  and  $\rho = 1/(1 + \zeta(mk_F - 1))$ , where the latter is inversely related to the distribution cost parameter  $\mu$ . In the model with translog preferences,  $p_{H,t}$  represents the price of import-competing goods,  $\xi = \frac{1}{2} \frac{N + N_H - 1}{N - 1}$  and  $\rho = \frac{N - 1}{N + N_H - 1}$ . The latter is inversely related to the share of domestic firms in the domestic market.

To increase the generality of the results, we do not impose all the restrictions implied by the Calvo model with indexation. For example, an equation like (48) could be derived from a Calvo model without indexation if a share of the firms that are allowed to change their prices did not set their prices optimally, but applied a rule of thumb based on the recent pricing behaviour of their competitors (see Galí & Gertler, 1999). The exact interpretation of the coefficients in the import price equation would be different, however. Equation (48) could also be derived from a model with quadratic costs of price adjustment as in Rotemberg (1982). The presence of lagged import price growth in the equation could then be motivated by quadratic costs of adjusting the level of import price growth relative to the previous period's import price growth (see e.g., Price,

1992; Ireland, 2001). Again, although the form of the linearised import price equation would be the same, the interpretation of the coefficients would be different. Finally, the model is potentially consistent with alternative models of pricing-to-market that imply that the optimal frictionless price can be written as a linear combination of the exporters' marginal costs and domestic prices in the importing country.

### 3.3 The GMM estimator

We first estimate the models using GMM. Limited information methods such as GMM do not rely on a specific completing model for the driving variables. This is an advantage given the lack of a satisfactory structural model for the exchange rate and the challenge involved in recovering a stable reduced form model for the exchange rate over a period that covers several monetary policy regimes.<sup>16</sup>

Let  $\varepsilon_{t+1} \equiv \Delta p_{F,t+1} - E_t \Delta p_{F,t+1}$  and  $v_{t+1} \equiv \Delta s_{t+1} - E_t \Delta s_{t+1}$ , and let  $\mathcal{F}_t$  denote the exporting firm's information set at time  $t$ . Then, according to the rational expectations hypothesis,  $E_t[\varepsilon_{t+1}|\mathcal{F}_t] = E_t[v_{t+1}|\mathcal{F}_t] = 0$ . Replacing the expected values with the actual realisations of the variables and adding a constant term we obtain the following estimating equation

$$\begin{aligned} \Delta p_{F,t} = & \alpha_0 + \alpha_1 \Delta p_{F,t+1} + \alpha_2 \Delta p_{F,t-1} + \alpha_3 (\Delta s_t - \alpha_1 \Delta s_{t+1} - \alpha_2 \Delta s_{t-1}) \\ & + \alpha_4 (s_t + ulc_{F,t} - p_{F,t}) + \alpha_5 (s_t + p_{COM,t} - p_{F,t}) + \alpha_6 (p_{H,t} - p_{F,t}) + \omega_t, \end{aligned} \quad (49)$$

where  $\omega_t \equiv u_t - \alpha_1 (\varepsilon_{t+1} - \alpha_3 v_{t+1})$  is a linear combination of the rational expectations forecast errors and the 'structural' disturbance term  $u_t$ . By construction, the disturbance  $\omega_t$  is correlated with the regressors, which implies that ordinary least squares will not yield consistent estimates of the parameters in the model.

Let  $z_t \in \mathcal{F}_t$  denote a  $q \times 1$  vector of variables satisfying  $E_t[u_t z_t] = 0$ . Then, it follows from the definition of the disturbance term  $\omega_t$  that  $z_t$  is a vector of valid instruments, that is

$$E_t[\omega_t z_t] = 0, \quad t = 1, \dots, T \quad (50)$$

These moment conditions provide the basis for the GMM estimation. The GMM estimator is given by

$$\hat{\Upsilon} = \arg \min_{\Upsilon} \left( \frac{1}{T} \sum_{t=1}^T \omega_t(\Upsilon) z_t \right)' W_T \left( \frac{1}{T} \sum_{t=1}^T \omega_t(\Upsilon) z_t \right), \quad (51)$$

where  $W_T$  is a positive semi-definite weighting matrix. Under certain regularity conditions (see Hall, 2005, p. 50), the GMM estimator is consistent and asymptotically normal. The asymptotic variance of  $\hat{\Upsilon}$  is minimised by setting the weighting matrix  $W_T$  equal to a consistent estimate of the inverse of the long-run covariance matrix of the sample moments:

$$\begin{aligned} S &= \lim_{T \rightarrow \infty} \text{Var} \left[ \frac{1}{\sqrt{T}} \sum_{t=1}^T \omega_t(\Upsilon) z_t \right] \\ &= \Gamma_0 + \sum_{i=1}^{\infty} (\Gamma_i + \Gamma_i'), \end{aligned}$$

where  $\Gamma_i$  is the  $i^{\text{th}}$  autocovariance matrix of the sample moments.

The composite disturbance term  $\omega_t$  can be shown to have a first-order moving-average representation (see Pesaran, 1987, p. 191).<sup>17</sup> This compels us to use a heteroskedasticity and autocorrelation

<sup>16</sup>This advantage may come at a cost of lower efficiency as not all cross-restrictions implied by the rational expectations hypothesis are imposed during estimation. Moreover, as emphasised by Pesaran (1987) and Mavroidis (2004), the strength of identification in GMM depends on properties of the processes governing the driving variables.

<sup>17</sup>Thus, first-order residual autocorrelation is not in itself a valid cause for rejecting the New Keynesian import price

consistent (HAC) estimate of  $S$ ,

$$\widehat{S}_T = \widehat{\Gamma}_0 + \sum_{i=1}^l \mathbf{v}_i \left( \widehat{\Gamma}_i + \widehat{\Gamma}_i' \right), \quad (52)$$

where  $\widehat{\Gamma}_i$  are the sample autocovariances, and  $l$  denotes the bound on how many autocovariances are used to form the estimate. To ensure that  $\widehat{S}_T$  is positive semi-definite in finite samples, the autocovariances are weighted using the kernel  $\mathbf{v}_i$ . Below we use the Bartlett kernel as proposed by Newey & West (1987)

$$\mathbf{v}_i = \begin{cases} 1 - \frac{i}{b+1} & \text{for } \frac{i}{b+1} \leq 1 \\ 0 & \text{for } \frac{i}{b+1} > 1 \end{cases} \quad (53)$$

where  $b$  is the bandwidth parameter chosen by the investigator. Since the behaviour of the HAC estimator of the covariance matrix can be highly sensitive to the choice of bandwidth parameter, den Haan & Levin (1996) recommend using more than one approach when estimating the covariance matrix. Below we therefore consider three different choices of bandwidth: a fixed bandwidth equal to 1, a fixed bandwidth equal to 3, and the bandwidth selected by the data-based method proposed by Newey & West (1994). The latter depends on the autocovariances of the moment conditions and typically results in a quite large bandwidth for our data.

A test of the over-identifying restrictions can in principle be based on the  $J$ -test statistic of Hansen (1982)

$$J = \left( \frac{1}{\sqrt{T}} \sum_{t=1}^T \boldsymbol{\omega}_t \left( \widehat{\Upsilon} \right) z_t \right)' \widehat{S}_T^{-1} \left( \frac{1}{\sqrt{T}} \sum_{t=1}^T \boldsymbol{\omega}_t \left( \widehat{\Upsilon} \right) z_t \right) \xrightarrow{d} \chi^2(q-r), \quad (54)$$

where  $r$  is the number of parameters to be estimated and  $q-r$  denotes the number of over-identifying restrictions. However, Monte Carlo evidence in Mavroidis (2005) suggests that the finite-sample power of the  $J$ -test to detect misspecification in forward-looking inflation equations is low, particularly when the number of instruments is large or the HAC estimate of the covariance matrix allows for a very general correction for autocorrelation.

Finally, a fundamental condition for consistency of the GMM estimator is that the population moment condition in (50) is satisfied at only one value in the parameter space (see e.g., Hall, 2005, p. 51). If this condition is satisfied, we say that the parameter vector  $\Upsilon$  is identified. However, the literature on weak identification in GMM estimation (see Stock et al. (2002) for a survey) has demonstrated that generic identification is not sufficient to ensure reliable inference in finite samples. If the parameters are weakly identified; that is, if the instruments are only weakly correlated with the endogenous variables, conventional point estimates and confidence intervals based on the asymptotic normal approximation will be misleading, even in large samples. In the models with both PCP and LCP, we need instruments for the rate of exchange rate depreciation in period  $t+1$ . The fact that it has proven difficult to beat the random walk forecast of exchange rates suggests that weak identification might be of particular concern when estimating these models.

### 3.4 The cointegration implications of the New Keynesian import price models

The asymptotic properties of the GMM estimator are derived under the assumption that the variables in the model are stationary. GMM estimation of equation (49) implicitly assumes that  $s_t + ulc_{F,t} - p_{F,t}$ ,  $s_t + p_{COM,t} - p_{F,t}$  and  $p_{H,t} - p_{F,t}$  are stationary or cointegrated.

Cointegration is a testable implication of the theoretical model. Focusing on the case where the variables in the model are at most integrated of order one,  $I(1)$ , the New Keynesian import price models imply that import prices are cointegrated with the optimal frictionless price. The models

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equation. However, autocorrelation could also be due to model misspecification caused by e.g. omitted variables.

with a constant frictionless mark-up imply that import prices should be cointegrated with foreign marginal costs measured in domestic currency, that is

$$p_{F,t} - mc_{F,t} - s_t \sim I(0). \quad (55)$$

In this case, the long-run exchange rate pass-through, measured as the long-run elasticity of import prices with respect to the exchange rate, keeping marginal costs fixed, is complete. With our measure of marginal costs, the models with a constant frictionless mark-up imply that

$$p_{F,t} - s_t - (1 - \delta)ulc_{F,t} - \delta p_{COM,t} \sim I(0), \quad (56)$$

or, equivalently, that  $s_t + ulc_{F,t} - p_{F,t}$  and  $s_t + p_{COM,t} - p_{F,t}$  are cointegrated with cointegration parameter  $\delta/(1 - \delta)$ . We note that this would hold also if  $s_t + ulc_{F,t} - p_{F,t}$  and  $s_t + p_{COM,t} - p_{F,t}$  themselves were stationary, in which case there would be two cointegrating relations among the variables.

The pricing-to-market models predict that import prices should be cointegrated with foreign marginal costs and the price of domestic goods, that is

$$p_{F,t} - \rho (mc_{F,t} + s_t) - (1 - \rho)p_{H,t} \sim I(0). \quad (57)$$

A version of (57) has served as the theoretical starting point of many empirical studies of exchange rate pass-through.<sup>18</sup> It is common to interpret a significant coefficient on domestic prices (i.e.,  $\rho < 1$ ) in the cointegrating regression as evidence of long-run pricing-to-market. Notice, however, that (57) would hold if relative prices and costs themselves were stationary, that is, if

$$p_{F,t} - mc_{F,t} - s_t \sim I(0) \text{ and } p_{H,t} - p_{F,t} \sim I(0). \quad (58)$$

In this case, the theory predicts that there should be two (or three) cointegrating vectors relating the variables. Hence, a finding that  $p_{F,t} - s_t - mc_{F,t} \sim I(0)$  is consistent with the pricing-to-market hypothesis.

To investigate the cointegration properties of the data, we first inspect the variables graphically and then conduct formal unit root and cointegration tests. Figure 7 plots foreign unit labour costs, commodity prices and domestic prices/costs relative to the import price index,  $s_t + ulc_{F,t} - p_{F,t}$ ,  $s_t + p_{COM,t} - p_{F,t}$  and  $p_{H,t} - p_{F,t}$ , for the UK and Norway. It is evident from the graphs that import prices increased less than foreign unit labour costs over the sample period. One possible explanation is that the price of raw materials increased less than unit labour costs. The graphs show that this is indeed the case. It cannot be the full explanation, however; there is no apparent downward trend in the ratio of commodity prices to import prices. The plots thus indicate that an implication of the models with a constant frictionless mark-up, namely that import prices is a constant mark-up on foreign marginal costs in the long run, does not hold in the data.

According to the pricing-to-market models, a fall in import prices relative to foreign unit labour costs could be explained by a fall in domestic unit labour costs or domestic producer prices relative to import prices. However, for both countries, the domestic cost- and price indices increased more than import prices over the sample period. Thus, the long-run implications from the theoretical import price equations in section 2 are seemingly rejected by the data.

One possible interpretation of the decline in import prices relative to foreign costs is that it captures a decline in tariffs and transportation costs over the sample. It may also reflect a shift in imports from high-cost to low-cost countries spurred by trade liberalisation. This fall in the price level is not picked up in our measure of unit labour costs, which, since it is a weighted average of unit labour costs indices with a common base-year value, will only pick up differences

<sup>18</sup>See e.g., Naug & Nymoen (1996) and Herzberg et al. (2003) for analyses of Norwegian and UK import prices respectively.

in cost inflation (see Høegh-Omdal & Wilhelmsen (2002), Røstøen (2004) and Nickell (2005) for a discussion of this point). These factors could also help explain why import prices have fallen relative to domestic prices and costs.

The effects of trade-liberalisation are not captured by the theoretical models considered in this paper. In the empirical analysis we approximate these effects by means of a linear trend. Specifically, we detrend the variables  $p_{H,t} - p_{F,t}$ ,  $s_t + ulc_{F,t} - p_{F,t}$  and  $s_t + p_{COM,t} - p_{F,t}$  prior to the GMM estimation by regressing each variable on a constant and a deterministic trend. The detrended variables are plotted in figure 8. The visual impression from the graphs is that  $s_t + ulc_{F,t} - p_{F,t}$ ,  $s_t + p_{COM,t} - p_{F,t}$  and  $p_{H,t} - p_{F,t}$  could be trend-stationary.

Table 1 reports the results of two different unit root tests on the detrended series: the augmented Dickey-Fuller (ADF) test (Dickey & Fuller, 1979; Said & Dickey, 1984) and the Kwiatkowski-Phillips-Schmidt-Shin (KPSS) test (Kwiatkowski et al., 1992).<sup>19</sup> The null hypothesis in the ADF test is that the variable has a unit root, while the null hypothesis in the KPSS test is that the variable is stationary. The KPSS test does not reject the null hypothesis of stationarity for any of the detrended series. Moreover, the ADF test rejects the unit root hypothesis at the 5% level for all series except the series for Norwegian unit labour costs relative to import prices. It is well-known that it is difficult to distinguish empirically between non-stationary processes and highly persistent yet stationary processes. However, on the basis of the unit root tests and the visual impression of the series, the assumption that the detrended series are stationary does not seem unreasonable.

#### 4 GMM ESTIMATION RESULTS

This section presents the GMM estimates of the New Keynesian import price equations for the UK and Norway. We first report estimation results for the models with a constant frictionless mark-up: a purely forward-looking LCP model, a hybrid LCP model and a model that allows for both PCP and LCP. Then, we report estimates of the two pricing-to-market models.

For the models with a constant frictionless mark-up the GMM estimation is based on the following sets of instruments:

$$z_{1,t} = \left\{ \begin{array}{l} \sum_{i=0}^2 \Delta s_{t-i}, \sum_{i=0}^2 \Delta ulc_{F,t-i}, \sum_{i=0}^2 \Delta p_{COM,t-i}, \sum_{i=1}^2 \Delta p_{F,t-i}, \\ s_{t-1} + ulc_{F,t-1} - p_{F,t-1}, s_{t-1} + p_{COM,t-1} - p_{F,t-1} \end{array} \right\}$$

$$z_{2,t} = \left\{ \begin{array}{l} \sum_{i=1}^2 \Delta s_{t-i}, \sum_{i=1}^2 \Delta ulc_{F,t-i}, \sum_{i=1}^2 \Delta p_{COM,t-i}, \sum_{i=1}^2 \Delta p_{F,t-i} \\ s_{t-1} + ulc_{F,t-1} - p_{F,t-1}, s_{t-1} + p_{COM,t-1} - p_{F,t-1} \end{array} \right\}$$

The set  $z_{1,t}$  contains current values and two lags of the first difference of the driving variables  $\Delta s_t$ ,  $\Delta ulc_{F,t}$  and  $\Delta p_{COM,t}$ , two lags of import price growth  $\Delta p_{F,t}$  and lagged values of real unit labour costs  $s_t + ulc_{F,t} - p_{F,t}$  and real commodity prices  $s_t + p_{COM,t} - p_{F,t}$ . For the current values of the driving variables to be valid instruments it must be the case that (i) the variables can be observed by the exporter before she sets prices in period  $t$ , and (ii) the variables are exogenous in the sense that  $E_t [u_t \Delta s_t] = E_t [u_t \Delta ulc_{F,t}] = E_t [u_t \Delta p_{COM,t}] = 0$ . These conditions are strict. First, because of time lags in gathering and processing information, the exporters may base their pricing decisions on expectations dated at time  $t - 1$  rather than at time  $t$ . Second, measurement errors in (our proxy of) marginal costs could make it correlated with the error term in the import price equation. For these reasons, we also report results for the instrument set  $z_{2,t}$ , which only contains variables dated  $t - 1$  or earlier.<sup>20</sup>

As mentioned above, in the models with both PCP and LCP we need instruments for the rate

<sup>19</sup>The results are obtained using EViews version 5.

<sup>20</sup>Using only lagged instruments is common in the literature on the New Keynesian Phillips Curve (see e.g., Galí & Gertler, 1999; Galí et al., 2001).

of exchange rate depreciation in period  $t + 1$ . For these models, we also considered an extended instrument set which included current and lagged values of the short-term interest rate differential between the importing country and its trading partners, as well as estimates of the output gap in the importing country and the output gap of total OECD.<sup>21</sup> However, the main conclusions in this section were not affected by this extension of the instrument set.

Tables 2 and 3 report the GMM estimates of the parameters in the purely forward-looking LCP model with a constant frictionless mark-up.<sup>22</sup> For the UK, the coefficients on the levels terms are positive and, in most cases, statistically significant. The  $J$ -test does not reject the validity of the over-identifying restrictions. The coefficient on the forward-term is negative or close to zero in all cases, however. We also note that the estimates are highly sensitive to the choice of instrument set and the choice of bandwidth parameter in the HAC estimate of the covariance matrix. This could be a symptom of weak identification (see Nason & Smith, 2005).<sup>23</sup> An estimate of  $\delta$ , the weight on commodity prices in marginal costs, can be computed from the ratio of the coefficients on the level terms (see section 3.2). The estimate of  $\delta$  varies from 0.41 to 0.44 when estimation is based on the instrument set  $z_{1,t}$  and from 0.34 to 0.37 when the instrument set is  $z_{2,t}$ .

The evidence of forward-looking price-setting is stronger for Norway: the coefficient on the forward-term is positive and, in most cases, statistically significant. The coefficient is also not significantly different from one. Moreover, the over-identifying restrictions are not rejected using the  $J$ -test. The coefficients on the level terms are statistically insignificant, however. This holds irrespective of the choice of instrument set or the choice of bandwidth parameter. When estimation is based on the instrument set  $z_{1,t}$ , the implicit estimate of  $\delta$  varies from 0.21 to 0.28 depending on the choice of bandwidth parameter. However, when the instrument set is  $z_{2,t}$ , the estimate of  $\delta$  varies from 0.03 when the bandwidth is based on the data-based method to 0.51 when the bandwidth is one.

The results for the ‘hybrid’ model with local currency pricing are reported in tables 4 and 5. For the UK, the coefficient on lagged import price growth is small and imprecisely estimated. The other coefficient estimates are similar to what was obtained in the purely forward-looking LCP model. In particular, the coefficient on the forward-term is still negative. For Norway, the coefficient on lagged import price growth is somewhat larger than for the UK. The estimates are far from being statistically significant, however. The coefficient on future import price growth is slightly smaller compared with the purely forward-looking model. Overall, the estimation results lend little support to the LCP model with indexation as a model of UK or Norwegian import prices of manufactures. This is consistent with the findings reported by Smets & Wouters (2002), who do not find evidence of strong indexation in euro-area import prices.

Next, we turn to the model that allows a subset of exporters to engage in PCP. The results are reported in tables 6 and 7. The key result emerging from these tables is the following: the coefficient on the exchange rate term is positive and both numerically and statistically significant. This holds for both datasets and across instrument sets and bandwidth parameters. The estimated share of PCP firms depends strongly on the choices of instrument set and bandwidth parameter, however. For the UK the estimated share of PCP firms ranges from 0.42 to 0.74; for Norway it ranges from 0.31 to 0.96. The remaining parameters are also affected by the inclusion of the exchange rate term. For the UK, the coefficient on the forward-term is positive (but still statistically

<sup>21</sup>The interest rate and output gap series were taken from OECD’s Economic Outlook database. The following countries were included in the measure of the trading partners’ interest rate: Australia, Canada, Japan, the euro area, Sweden, the US, Switzerland and in the case of Norway; the UK. The weights are based on the trade-weights used to construct the effective exchange rate (fixed 1995 weights).

<sup>22</sup>The results are obtained using the simultaneous-updating GMM estimator in EViews 5.

<sup>23</sup>In the purely forward-looking models, the coefficient on the forward-term corresponds to the subjective discount factor  $\beta$ . It has been noted by several authors that it is difficult to obtain accurate estimates of this parameter in single-equation rational expectations models (see e.g., the discussion in Gregory et al. (1993)). Some authors therefore fix the value of  $\beta$  prior to estimation.

insignificant) when estimation is based on the instrument set  $z_{2,t}$ .<sup>24</sup> For Norway, the coefficient on the forward-term is somewhat smaller than in the pure LCP model. The coefficients on the level terms are still small and statistically insignificant, however.

The results illustrate that the  $J$ -test has low power to detect misspecification: in the pure LCP model, the  $J$ -test did not reject the validity of the over-identifying restrictions when the instrument set contained current values of the exchange rate. The results do not offer support for the exact specification of the LCP-PCP model, but nevertheless constitute strong evidence against the pure LCP model.<sup>25</sup> This finding is consistent with the results reported by Choudhri et al. (2005), who estimate open-economy DSGE models on data for non-US G7 countries using an impulse response matching approach and find that the best-fitting model incorporates a combination of PCP and LCP.

Tables 8 to 11 report estimation results for the pricing-to-market models for the two different measures of domestic prices and costs. The estimation is based on the following instrument sets:

$$z_{3,t} = \left\{ z_{1,t}, \sum_{i=0}^2 \Delta p_{H,t-i}, p_{H,t-1} - p_{F,t-1} \right\}$$

$$z_{4,t} = \left\{ z_{2,t}, \sum_{i=1}^2 \Delta p_{H,t-i}, p_{H,t-1} - p_{F,t-1} \right\}$$

For the UK, the coefficient on domestic producer prices is positive and statistically significant, except in the case where the instrument set is  $z_{3,t}$  and the bandwidth in the HAC estimate of the covariance matrix is set equal to one. Ignoring the latter case, the estimate of the coefficient on domestic prices in the implied expression for the optimal frictionless price  $(1 - \rho)$  lies in the range 0.33-0.40. The estimate of long-run (conditional) exchange rate pass-through thus lies in the range 0.60-0.67. The implied estimate of  $\delta$  varies from 0.30 to 0.34. The coefficient on domestic unit labour costs is positive in all of the distribution cost models. Five of the six estimates are significant at the 10% level. The implied estimate of  $1 - \rho$  now varies from 0.15 to 0.22, and the estimate of  $\delta$  lies in the range 0.33-0.41. Thus, the long-run (conditional) pass-through is somewhat larger in this model: the estimates lie in the range 0.78-0.85.

The estimates of the purely forward-looking LCP model with pricing-to-market do seem to suggest a role for domestic prices and costs in explaining UK import prices. This conclusion is robust to extending the model to allow for indexation to past import price growth. The estimated coefficient on lagged import price inflation is now negative. The evidence of pricing-to-market is somewhat weaker if we extend the model to allow for both PCP and LCP, however. In this case, the coefficient on domestic producer prices or domestic unit labour costs is statistically insignificant and in some cases, negative. As in the models without pricing-to-market variables, the estimated share of PCP firms is positive and statistically significant. Moreover, the coefficient on the forward term is negative in most cases.

For Norway, the coefficient on domestic prices and costs is statistically insignificant. The coefficient on domestic producer prices is in many cases negative. The coefficient on domestic unit labour costs is positive, however. Again, the evidence of forward-looking price-setting is stronger for Norway than for the UK: the coefficient on the forward-term is positive and, in most cases, statistically significant. Moreover, the estimate of the discount factor  $\beta$  is economically plausible. For example, when the estimation of the distribution cost model is based on the instrument set  $z_{3,t}$  and the bandwidth is selected using the Newey-West method, the estimate of the discount factor is 0.99 and is statistically significant at the 1% level. In this case, the implied estimate of long-run

<sup>24</sup>When we use the Newey-West method to select the bandwidth in the HAC estimate of the weighting matrix, the coefficient on the forward-term is statistically significant at the 5% significance level if we use a one-sided test. Given our prior about the sign of effect of expected future import price growth, it could be argued that the one-sided test is more relevant than a two-sided test.

<sup>25</sup>These conclusions are robust to extending the LCP-PCP model to allow for inflation indexation.

(conditional) exchange rate pass-through ( $\rho$ ) is 0.58 and the estimate of  $\delta$  is 0.14. If we interpret the effect of domestic unit labour costs as the effect of distribution costs (see section 2.2.2), this estimate of long-run pass-through seems unreasonably low. More plausibly, the domestic unit labour cost variable is acting as a proxy for the price of import-competing products.

Extending the pricing-to-market models to allow for indexation to lagged import price growth, we reach similar conclusions as above: the coefficient on lagged import price growth is fairly small and statistically insignificant. The coefficient on the forward-term is still positive and statistically significant in most cases, whereas the coefficients on the levels terms, including the coefficients on the pricing-to-market variables, are statistically insignificant. Finally, extending the model to allow for both PCP and LCP has the effect of lowering the coefficient on the forward-term in most cases, although the coefficient remains statistically significant. The coefficient on domestic prices and costs is still statistically insignificant. As in the LCP-PCP model without pricing-to-market variables, the coefficient on the exchange rate term is positive and statistically significant.

## 5 CONCLUDING REMARKS

A key issue in the empirical literature on the New Keynesian Phillips Curve has been to determine whether price setters are ‘forward-looking’, in the sense that expected future prices matter for the determination of current prices. By contrast, most of the empirical work on import prices has taken the form of reduced-form pass-through regressions, with no attempt to distinguish between expectational dynamics and dynamics arising from other sources. This paper makes a first attempt to fill this gap by estimating New Keynesian import price equations derived from the Calvo model of staggered price setting.

Taken at face value, the GMM estimates obtained for the UK do not lend much support to the hypothesis that the price-setting rules are forward looking: the coefficient on expected future import price growth is either statistically insignificant, economically implausible, or both. The evidence of forward-looking price-setting is stronger for Norway: the coefficient on the forward-term is positive and, in most cases, statistically significant. For both countries, the estimation results favour a specification that allows for both PCP and LCP. By contrast, there seems to be little evidence of indexation to past import price growth.

For Norway, the estimated coefficients on foreign costs and the pricing-to-market variables are statistically insignificant and close to zero in most cases. This contrasts with the results obtained for the UK: the coefficients on the foreign cost variables are statistically significant and, moreover, the pricing-to-market models suggest a role for domestic prices or costs in explaining import prices.

The fact that the estimation results for the UK and Norway are so different is somewhat puzzling. The differences in the results could be related to differences in the country- or commodity composition of imports. The estimation of the pricing-to-market models requires proxies for variables which are inherently hard to measure: the price of local goods and services used in distribution and the price of import-competing goods. The commodity composition of manufacturing imports and the domestic production of manufactures is likely to be different. In particular, Norway imports manufactured goods (e.g., motor vehicles) for which there do not exist domestic substitutes. Such measurement problems could be part of the explanation why we do not obtain a significant effect of domestic prices for Norway.

By using a linear trend to capture the effects of trade liberalisation, we have implicitly assumed that these effects have been constant over the sample period. This is a strict assumption. The plots of the Norwegian dataseries indicate that the downward trend in import prices relative to domestic prices and costs became more pronounced in the last part of the sample. A more flexible approach would be to allow for an unobservable stochastic trend in the model.



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## A VARIABLE DEFINITIONS AND SOURCES

- $P_F$ : Import price of manufactured goods (local currency).<sup>26</sup> Seasonally adjusted (SA). Source: OECD International Trade and Competitiveness Indicators (ITCI)
- $S$ : Nominal effective exchange rate. Sources: Bank of England (Broad effective exchange rate index [XUQABK82]) and Norges Bank (Trade-weighted exchange rate [TWI]).
- $ULC_F$ : Unit labour costs in foreign manufacturing (foreign currency). Trade-weighted. SA. Source: OECD Economic Outlook. Constructed as  $\frac{ULCM}{EXCHEB \times ULCMDR}$  where  $ULCM$  is unit labour costs in domestic manufacturing industries [Q.GBR.ULCM/Q.NOR.ULCM],  $EXCHEB$  is the nominal effective exchange rate [Q.GBR.EXCHEB/Q.NOR.EXCHEB] and  $ULCMDR$  is relative unit labour costs [Q.GBR.ULCMDR/Q.NOR.ULCMDR].
- $P_{COM}$ : Index of world metal prices. The original series is measured in US dollars. We convert the index to the currency of the trading partners using the official effective exchange rate  $S$ . Source: IMF International Financial Statistics [00176AYDZF...].
- $P_H$ : Producer price index (home sales). Seasonally adjusted using the X12-ARIMA routines implemented in EViews 5. Sources: OECD Main Economic Indicators [GBR.PPIAMP01.IXOB.Q] and Statistics Norway [Commodity price index for the industrial sectors (VPPI). Total industry. Domestic market.] Or; Unit labour costs for the total economy. Source: OECD Economic Outlook [Q.GBR.ULC/Q.NOR.ULC].

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<sup>26</sup>All variables are converted to a common baseyear 1995=1.

Table 1: Univariate unit root tests. Detrended series. 1980Q1-2003Q2<sup>a</sup>

	UK		Norway	
	ADF	KPSS	ADF	KPSS
$s_t + ulc_{F,t} - p_{F,t}$	-2.32*	0.16	-2.57*	0.13
$s_t + p_{COM,t} - p_{F,t}$	-3.99**	0.06	-3.74**	0.11
$p_{H,t} - p_{F,t}$ <sup>b</sup>	-2.36*	0.18	-2.07*	0.23
$p_{H,t} - p_{F,t}$ <sup>c</sup>	-2.58*	0.15	-1.25	0.21

a The numbers in the table are the ADF  $t$ -statistics and the KPSS  $LM$ -statistics. Single asterisks (\*) and double asterisks (\*\*) denote statistical significance at the 5% level and the 1% level, respectively. The ADF test equation does not include any deterministic terms. The selection of lag-order for the ADF test is based on the Akaike Information Criterion (AIC) with the maximum number of lagged differenced terms set to four. The critical values for the ADF test are taken from MacKinnon (1996). The KPSS test equation includes a constant. The KPSS tests were run using a Bartlett kernel and the bandwidth is selected using the Newey & West (1994) method. The 1% and the 5% asymptotic critical values are 0.739 and 0.463, respectively (see table 1 in Kwiatkowski et al., 1992).

b  $p_{H,t}$  denotes producer price index for manufactures.

c  $p_{H,t}$  denotes domestic unit labour costs for the total economy.

Table 2: GMM estimates of a purely forward-looking model with local currency pricing. UK data 1980Q4–2003Q1.

$$\Delta p_{F,t} = \alpha_0 + \alpha_1 \Delta p_{F,t+1} + \alpha_4 (s_t + ulc_{F,t} - p_{F,t}) + \alpha_5 (s_t + p_{COM,t} - p_{F,t}) + \omega_t$$

Instruments	Bandwidth	$\alpha_0$	$\alpha_1$	$\alpha_4$	$\alpha_5$	$J - stat$
$z_{1,t}$	1	0.00 (0.002)	0.08 (0.212)	0.08* (0.039)	0.06** (0.018)	$\chi^2(10) = 13.06$ [0.22]
$z_{1,t}$	3	0.00 (0.002)	-0.09 (0.196)	0.13** (0.039)	0.09** (0.020)	$\chi^2(10) = 10.61$ [0.39]
$z_{1,t}$	Newey-West	-0.00 (0.002)	-0.02 (0.138)	0.13** (0.035)	0.10** (0.019)	$\chi^2(10) = 8.46$ [0.58]
$z_{2,t}$	1	0.00 (0.003)	-0.45 (0.299)	0.14** (0.051)	0.08** (0.024)	$\chi^2(7) = 7.95$ [0.34]
$z_{2,t}$	3	0.00 (0.003)	-0.52 (0.285)	0.19** (0.048)	0.10** (0.024)	$\chi^2(7) = 6.98$ [0.43]
$z_{2,t}$	Newey-West	0.00 (0.003)	-0.50 (0.277)	0.19** (0.042)	0.11** (0.021)	$\chi^2(7) = 5.89$ [0.55]

Note: Numbers in parenthesis denote standard errors, numbers in square brackets denote  $p$ -values. Asterisks \*\* and \* denote statistical significance at the 1% and 5% level, respectively

Table 3: GMM estimates of a purely forward-looking model with local currency pricing. Norwegian data 1980Q4–2002Q4.

$$\Delta p_{F,t} = \alpha_0 + \alpha_1 \Delta p_{F,t+1} + \alpha_4 (s_t + ulc_{F,t} - p_{F,t}) + \alpha_5 (s_t + p_{COM,t} - p_{F,t}) + \omega_t$$

Instruments	Bandwidth	$\alpha_0$	$\alpha_1$	$\alpha_4$	$\alpha_5$	$J - stat$
$z_{1,t}$	1	0.00 (0.002)	0.71** (0.268)	0.09 (0.087)	0.03 (0.023)	$\chi^2(10) = 8.16$ [0.61]
$z_{1,t}$	3	0.00 (0.002)	0.82** (0.216)	0.03 (0.059)	0.01 (0.018)	$\chi^2(10) = 8.06$ [0.62]
$z_{1,t}$	Newey-West	0.00 (0.001)	0.85** (0.176)	0.04 (0.048)	0.01 (0.015)	$\chi^2(10) = 6.91$ [0.73]
$z_{2,t}$	1	0.00 (0.002)	0.67 (0.419)	0.09 (0.089)	0.03 (0.027)	$\chi^2(7) = 7.31$ [0.40]
$z_{2,t}$	3	0.00 (0.002)	0.85* (0.361)	0.01 (0.065)	0.01 (0.023)	$\chi^2(7) = 6.57$ [0.48]
$z_{2,t}$	Newey-West	0.00 (0.001)	0.88** (0.300)	0.02 (0.046)	0.00 (0.016)	$\chi^2(7) = 5.27$ [0.63]

Note: Numbers in parenthesis denote standard errors, numbers in square brackets denote  $p$ -values. Asterisks \*\* and \* denote statistical significance at the 1% and 5% level, respectively

Table 4: GMM estimates of a hybrid model with local currency pricing. UK data 1980Q4–2003Q1.

$$\Delta p_{F,t} = \alpha_0 + \alpha_1 \Delta p_{F,t+1} + \alpha_2 \Delta p_{F,t-1} + \alpha_4 (s_t + ulc_{F,t} - p_{F,t}) + \alpha_5 (s_t + p_{COM,t} - p_{F,t}) + \omega_t$$

Instruments	Bandwidth	$\alpha_0$	$\alpha_1$	$\alpha_4$	$\alpha_5$	$\alpha_2$	$J - stat$
$z_{1,t}$	1	0.00 (0.002)	0.08 (0.212)	0.08 (0.045)	0.06** (0.019)	0.02 (0.111)	$\chi^2(9) = 13.01$ [0.16]
$z_{1,t}$	3	0.00 (0.002)	-0.09 (0.209)	0.14** (0.049)	0.10** (0.022)	-0.03 (0.106)	$\chi^2(9) = 10.62$ [0.30]
$z_{1,t}$	Newey-West	-0.00 (0.002)	-0.05 (0.140)	0.13** (0.041)	0.10** (0.020)	0.03 (0.087)	$\chi^2(9) = 8.38$ [0.50]
$z_{2,t}$	1	0.00 (0.003)	-0.43 (0.325)	0.13* (0.065)	0.07** (0.026)	0.02 (0.113)	$\chi^2(6) = 8.01$ [0.24]
$z_{2,t}$	3	0.00 (0.003)	-0.54 (0.318)	0.20** (0.064)	0.10** (0.027)	-0.04 (0.116)	$\chi^2(6) = 6.86$ [0.33]
$z_{2,t}$	Newey-West	0.00 (0.003)	-0.52 (0.298)	0.22** (0.059)	0.12** (0.025)	-0.07 (0.095)	$\chi^2(6) = 5.82$ [0.44]

Note: Numbers in parenthesis denote standard errors, numbers in square brackets denote  $p$ -values. Asterisks \*\* and \* denote statistical significance at the 1% and 5% level, respectively

Table 5: GMM estimates of a hybrid model with local currency pricing. Norwegian data 1980Q4–2002Q4.

$$\Delta p_{F,t} = \alpha_0 + \alpha_1 \Delta p_{F,t+1} + \alpha_2 \Delta p_{F,t-1} + \alpha_4 (s_t + ulc_{F,t} - p_{F,t}) + \alpha_5 (s_t + p_{COM,t} - p_{F,t}) + \omega_t$$

Instruments	Bandwidth	$\alpha_0$	$\alpha_1$	$\alpha_4$	$\alpha_5$	$\alpha_2$	$J - stat$
$z_{1,t}$	1	0.00 (0.002)	0.64* (0.288)	0.12 (0.090)	0.04 (0.023)	0.13 (0.110)	$\chi^2(9) = 7.39$ [0.60]
$z_{1,t}$	3	0.00 (0.002)	0.69** (0.238)	0.07 (0.067)	0.02 (0.019)	0.12 (0.099)	$\chi^2(9) = 7.86$ [0.55]
$z_{1,t}$	Newey-West	0.00 (0.001)	0.74** (0.179)	0.05 (0.048)	0.01 (0.015)	0.04 (0.084)	$\chi^2(9) = 6.66$ [0.67]
$z_{2,t}$	1	0.00 (0.002)	0.61 (0.427)	0.12 (0.093)	0.04 (0.027)	0.13 (0.111)	$\chi^2(6) = 6.36$ [0.38]
$z_{2,t}$	3	0.00 (0.002)	0.77* (0.370)	0.05 (0.073)	0.02 (0.023)	0.10 (0.105)	$\chi^2(6) = 6.17$ [0.40]
$z_{2,t}$	Newey-West	0.00 (0.001)	0.84** (0.307)	0.03 (0.051)	0.00 (0.018)	0.04 (0.095)	$\chi^2(6) = 5.24$ [0.51]

Note: Numbers in parenthesis denote standard errors, numbers in square brackets denote  $p$ -values. Asterisks \*\* and \* denote statistical significance at the 1% and 5% level, respectively

Table 6: GMM estimates of a model with producer- and local currency pricing. UK data 1980Q4–2003Q1

$$\Delta p_{F,t} = \alpha_0 + \alpha_1 \Delta p_{F,t+1} + \alpha_3 (\Delta s_t - \alpha_1 \Delta s_{t+1}) + \alpha_4 (s_t + ulc_{F,t} - p_{F,t}) + \alpha_5 (s_t + p_{COM,t} - p_{F,t}) + \omega_t$$

Instruments	Bandwidth	$\alpha_0$	$\alpha_1$	$\alpha_4$	$\alpha_5$	$\alpha_3$	$J - stat$
$z_{1,t}$	1	0.00 (0.002)	-0.13 (0.207)	0.13** (0.033)	0.07** (0.014)	0.42** (0.056)	$\chi^2(9) = 9.86$ [0.36]
$z_{1,t}$	3	0.00 (0.002)	-0.02 (0.158)	0.13** (0.033)	0.08** (0.014)	0.44** (0.052)	$\chi^2(9) = 9.26$ [0.41]
$z_{1,t}$	Newey-West	0.00 (0.002)	0.01 (0.132)	0.15** (0.029)	0.09** (0.013)	0.42** (0.041)	$\chi^2(9) = 7.12$ [0.62]
$z_{2,t}$	1	0.00 (0.002)	0.14 (0.191)	0.08* (0.036)	0.05** (0.018)	0.74** (0.141)	$\chi^2(6) = 7.28$ [0.30]
$z_{2,t}$	3	0.00 (0.002)	0.20 (0.177)	0.09* (0.036)	0.06** (0.017)	0.70** (0.109)	$\chi^2(6) = 6.86$ [0.33]
$z_{2,t}$	Newey-West	0.00 (0.001)	0.23 (0.148)	0.09** (0.031)	0.07** (0.016)	0.65** (0.084)	$\chi^2(6) = 5.79$ [0.45]

Note: Numbers in parenthesis denote standard errors, numbers in square brackets denote  $p$ -values. Asterisks \*\* and \* denote statistical significance at the 1% and 5% level, respectively

Table 7: GMM estimates of a model with producer- and local currency pricing. Norwegian data 1980Q4–2002Q4.

$$\Delta p_{F,t} = \alpha_0 + \alpha_1 \Delta p_{F,t+1} + \alpha_3 (\Delta s_t - \alpha_1 \Delta s_{t+1}) + \alpha_4 (s_t + ulc_{F,t} - p_{F,t}) + \alpha_5 (s_t + p_{COM,t} - p_{F,t}) + \omega_t$$

Instruments	Bandwidth	$\alpha_0$	$\alpha_1$	$\alpha_4$	$\alpha_5$	$\alpha_3$	$J - stat$
$z_{1,t}$	1	0.00 (0.002)	0.43 (0.293)	0.09 (0.081)	0.03 (0.026)	0.46** (0.162)	$\chi^2(9) = 7.97$ [0.54]
$z_{1,t}$	3	0.00 (0.002)	0.53* (0.243)	0.06 (0.061)	0.02 (0.023)	0.45** (0.138)	$\chi^2(9) = 6.83$ [0.65]
$z_{1,t}$	Newey-West	0.00 (0.001)	0.56* (0.239)	0.06 (0.058)	0.02 (0.021)	0.44** (0.131)	$\chi^2(9) = 6.46$ [0.69]
$z_{2,t}$	1	0.00 (0.002)	0.72 (0.453)	0.07 (0.120)	0.03 (0.043)	0.96* (0.450)	$\chi^2(6) = 5.26$ [0.51]
$z_{2,t}$	3	0.00 (0.002)	0.69 (0.380)	0.08 (0.085)	0.04 (0.032)	0.71 (0.370)	$\chi^2(6) = 5.22$ [0.52]
$z_{2,t}$	Newey-West	0.00 (0.002)	0.74 (0.374)	0.07 (0.079)	0.03 (0.031)	0.65 (0.352)	$\chi^2(6) = 4.93$ [0.55]

Note: Numbers in parenthesis denote standard errors, numbers in square brackets denote  $p$ -values. Asterisks \*\* and \* denote statistical significance at the 1% and 5% level, respectively



Table 8: GMM estimates of a pricing-to-market model with local currency pricing. Pricing-to-market variable: Domestic producer prices. UK data 1980Q4–2003Q1.

$$\Delta p_{F,t} = \alpha_0 + \alpha_1 \Delta p_{F,t+1} + \alpha_4 (s_t + ulc_{F,t} - p_{F,t}) + \alpha_5 (s_t + p_{COM,t} - p_{F,t}) + \alpha_6 (p_{H,t} - p_{F,t}) + \omega_t$$

Instruments	Bandwidth	$\alpha_0$	$\alpha_1$	$\alpha_4$	$\alpha_5$	$\alpha_6$	$J - stat$
$z_{3,t}$	1	0.00 (0.002)	0.46* (0.181)	0.07 (0.062)	0.06* (0.024)	-0.00 (0.066)	$\chi^2(13) = 16.57 [0.22]$
$z_{3,t}$	3	-0.00 (0.002)	-0.57** (0.197)	0.31** (0.060)	0.16** (0.030)	0.23* (0.096)	$\chi^2(13) = 11.43 [0.58]$
$z_{3,t}$	Newey-West	-0.00 (0.002)	-0.40* (0.184)	0.42** (0.062)	0.19** (0.029)	0.30** (0.103)	$\chi^2(13) = 9.02 [0.77]$
$z_{4,t}$	1	0.01 (0.003)	-0.87 (0.347)	0.39** (0.102)	0.17** (0.046)	0.37** (0.132)	$\chi^2(9) = 7.31 [0.61]$
$z_{4,t}$	3	0.00 (0.003)	-0.83** (0.261)	0.40** (0.082)	0.18** (0.043)	0.34** (0.114)	$\chi^2(9) = 7.04 [0.63]$
$z_{4,t}$	Newey-West	0.00 (0.003)	-0.70** (0.211)	0.41** (0.077)	0.19** (0.040)	0.34** (0.110)	$\chi^2(9) = 5.90 [0.75]$

Note: Numbers in parenthesis denote standard errors, numbers in square brackets denote  $p$ -values. Asterisks \*\* and \* denote statistical significance at the 1% and 5% level, respectively

Table 9: GMM estimates of a pricing-to-market model with local currency pricing. Pricing-to-market variable: Domestic producer prices. Norwegian data 1980Q4–2002Q4.

$$\Delta p_{F,t} = \alpha_0 + \alpha_1 \Delta p_{F,t+1} + \alpha_4 (s_t + ulc_{F,t} - p_{F,t}) + \alpha_5 (s_t + p_{COM,t} - p_{F,t}) + \alpha_6 (p_{H,t} - p_{F,t}) + \omega_t$$

Instruments	Bandwidth	$\alpha_0$	$\alpha_1$	$\alpha_4$	$\alpha_5$	$\alpha_6$	$J - stat$
$z_{3,t}$	1	0.00 (0.002)	0.81** (0.236)	0.05 (0.078)	0.02 (0.021)	-0.01 (0.051)	$\chi^2(13) = 8.55 [0.81]$
$z_{3,t}$	3	0.00 (0.002)	0.83** (0.193)	0.02 (0.055)	0.01 (0.017)	-0.02 (0.036)	$\chi^2(13) = 8.67 [0.80]$
$z_{3,t}$	Newey-West	0.00 (0.001)	0.84** (0.150)	0.03 (0.046)	0.01 (0.014)	0.01 (0.032)	$\chi^2(13) = 7.87 [0.85]$
$z_{4,t}$	1	0.00 (0.002)	0.91* (0.361)	0.03 (0.084)	0.01 (0.025)	-0.02 (0.068)	$\chi^2(9) = 6.97 [0.64]$
$z_{4,t}$	3	0.00 (0.002)	0.91** (0.300)	-0.01 (0.062)	-0.00 (0.020)	-0.01 (0.046)	$\chi^2(9) = 6.61 [0.68]$
$z_{4,t}$	Newey-West	0.00 (0.001)	0.97** (0.237)	0.01 (0.050)	-0.01 (0.016)	0.02 (0.038)	$\chi^2(9) = 5.34 [0.80]$

Note: Numbers in parenthesis denote standard errors, numbers in square brackets denote  $p$ -values. Asterisks \*\* and \* denote statistical significance at the 1% and 5% level, respectively

Table 10: GMM estimates of a pricing-to-market model with local currency pricing. Pricing-to-market variable: Domestic unit labour costs. UK data 1980Q4–2003Q1.

$$\Delta p_{F,t} = \alpha_0 + \alpha_1 \Delta p_{F,t+1} + \alpha_4 (s_t + ulc_{F,t} - p_{F,t}) + \alpha_5 (s_t + p_{COM,t} - p_{F,t}) + \alpha_6 (p_{H,t} - p_{F,t}) + \omega_t$$

Instruments	Bandwidth	$\alpha_0$	$\alpha_1$	$\alpha_4$	$\alpha_5$	$\alpha_6$	$J - stat$
$z_{3,t}$	1	0.00 (0.002)	-0.07 (0.205)	0.13** (0.045)	0.09** (0.020)	0.04 (0.033)	$\chi^2(13) = 14.94$ [0.31]
$z_{3,t}$	3	-0.00 (0.003)	-0.02 (0.208)	0.30** (0.038)	0.18** (0.024)	0.09 (0.047)	$\chi^2(13) = 12.95$ [0.45]
$z_{3,t}$	Newey-West	-0.00 (0.003)	0.01 (0.209)	0.35** (0.037)	0.20** (0.026)	0.10 (0.051)	$\chi^2(13) = 8.13$ [0.84]
$z_{4,t}$	1	0.00 (0.003)	-0.41 (0.302)	0.29** (0.060)	0.14** (0.029)	0.12* (0.048)	$\chi^2(9) = 8.17$ [0.52]
$z_{4,t}$	3	0.00 (0.003)	-0.16 (0.301)	0.36** (0.052)	0.18** (0.029)	0.14* (0.057)	$\chi^2(9) = 7.39$ [0.60]
$z_{4,t}$	Newey-West	-0.00 (0.003)	0.04 (0.275)	0.38** (0.045)	0.20** (0.031)	0.14* (0.060)	$\chi^2(9) = 6.33$ [0.71]

Note: Numbers in parenthesis denote standard errors, numbers in square brackets denote  $p$ -values. Asterisks \*\* and \* denote statistical significance at the 1% and 5% level, respectively

Table 11: GMM estimates of a pricing-to-market model with local currency pricing. Pricing-to-market variable: Domestic unit labour costs. Norwegian data 1980Q4–2002Q4.

$$\Delta p_{F,t} = \alpha_0 + \alpha_1 \Delta p_{F,t+1} + \alpha_4 (s_t + ulc_{F,t} - p_{F,t}) + \alpha_5 (s_t + p_{COM,t} - p_{F,t}) + \alpha_6 (p_{H,t} - p_{F,t}) + \omega_t$$

Instruments	Bandwidth	$\alpha_0$	$\alpha_1$	$\alpha_4$	$\alpha_5$	$\alpha_6$	$J - stat$
$z_{3,t}$	1	0.00 (0.002)	0.84** (0.226)	0.08 (0.095)	0.02 (0.023)	0.03 (0.036)	$\chi^2(13) = 8.59$ [0.80]
$z_{3,t}$	3	0.00 (0.002)	0.92** (0.185)	0.03 (0.068)	0.01 (0.018)	0.02 (0.030)	$\chi^2(13) = 8.16$ [0.83]
$z_{3,t}$	Newey-West	-0.00 (0.001)	0.99** (0.153)	0.04 (0.060)	0.01 (0.015)	0.03 (0.024)	$\chi^2(13) = 7.01$ [0.90]
$z_{4,t}$	1	0.00 (0.002)	0.75 (0.406)	0.12 (0.103)	0.04 (0.029)	0.03 (0.042)	$\chi^2(9) = 8.28$ [0.51]
$z_{4,t}$	3	0.00 (0.002)	0.92** (0.344)	0.03 (0.074)	0.01 (0.025)	0.03 (0.034)	$\chi^2(9) = 7.49$ [0.59]
$z_{4,t}$	Newey-West	-0.00 (0.001)	0.95** (0.347)	0.05 (0.063)	0.00 (0.019)	0.04 (0.031)	$\chi^2(9) = 5.93$ [0.75]

Note: Numbers in parenthesis denote standard errors, numbers in square brackets denote  $p$ -values. Asterisks \*\* and \* denote statistical significance at the 1% and 5% level, respectively

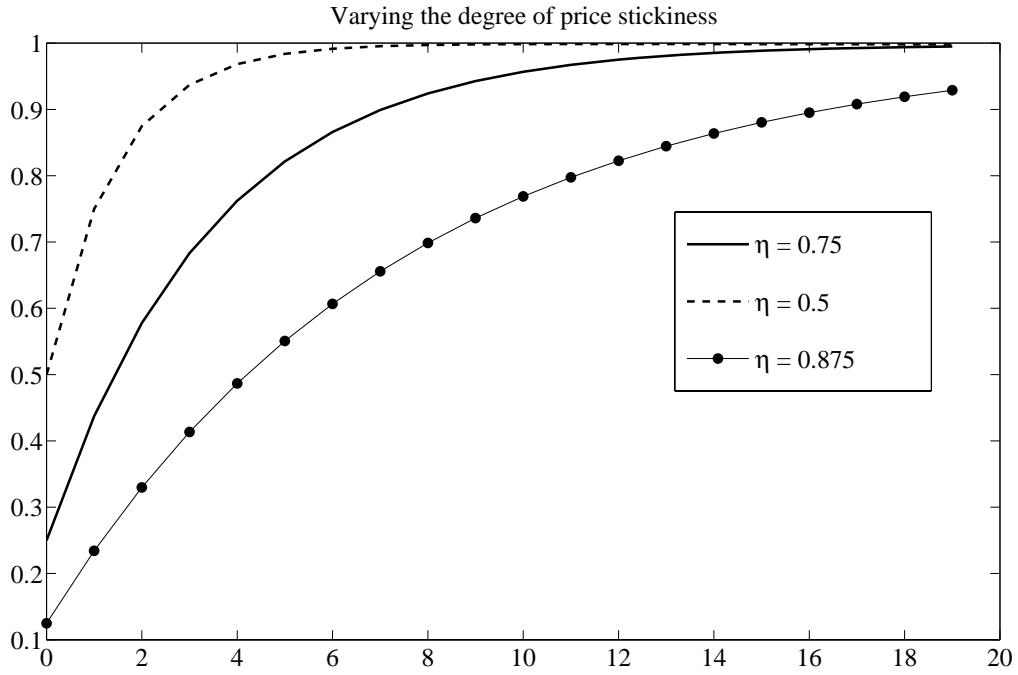


Figure 1: The response of import prices to 1% permanent exchange rate shock for different degrees of price stickiness  $\eta$  in the purely forward-looking LCP model.  $\beta = 0.99$ .

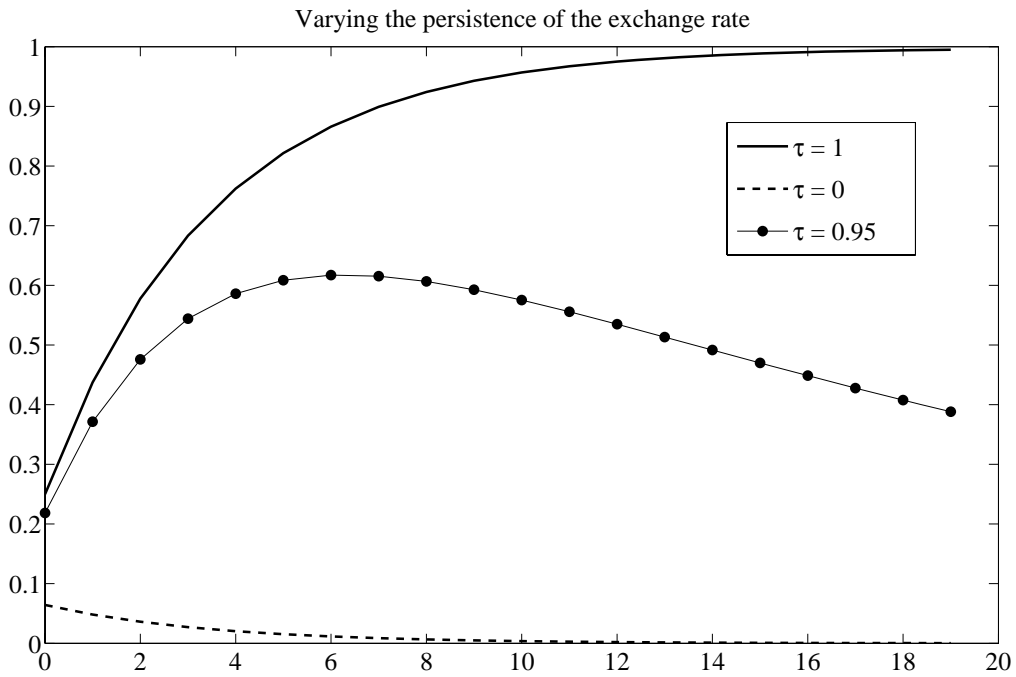


Figure 2: The response of import prices to 1% exchange rate shock for different degrees of persistence  $\tau$  in the exchange rate in the purely forward-looking LCP model.  $\beta = 0.99, \eta = 0.75$ .

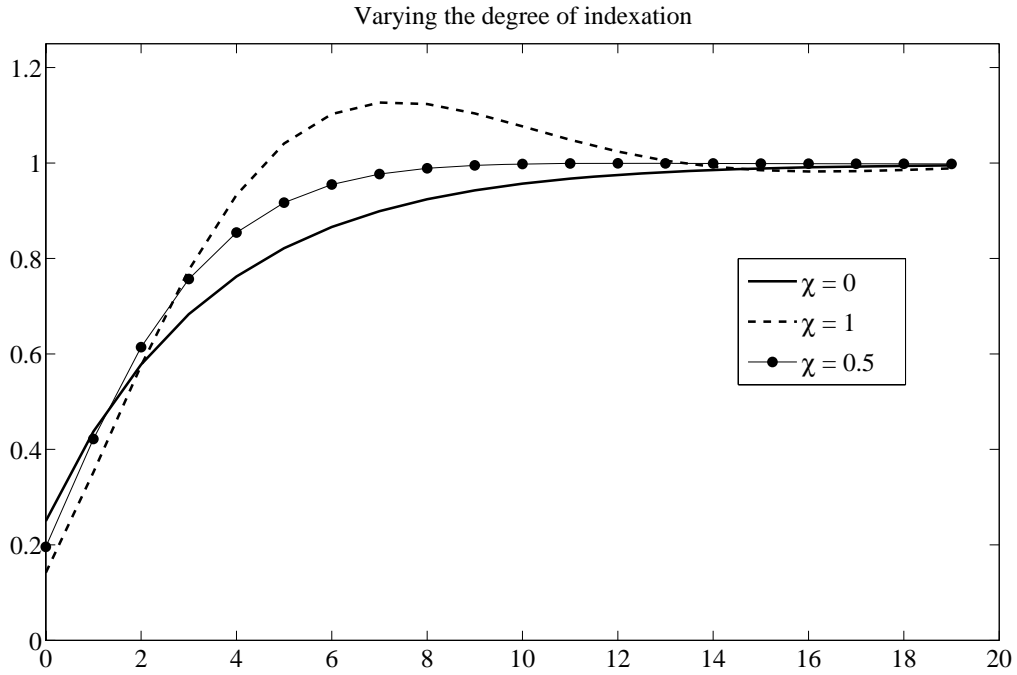


Figure 3: The response of import prices to 1% permanent exchange rate shock for different values of the indexation parameter in the hybrid LCP model.  $\beta = 0.99, \eta = 0.75$ .

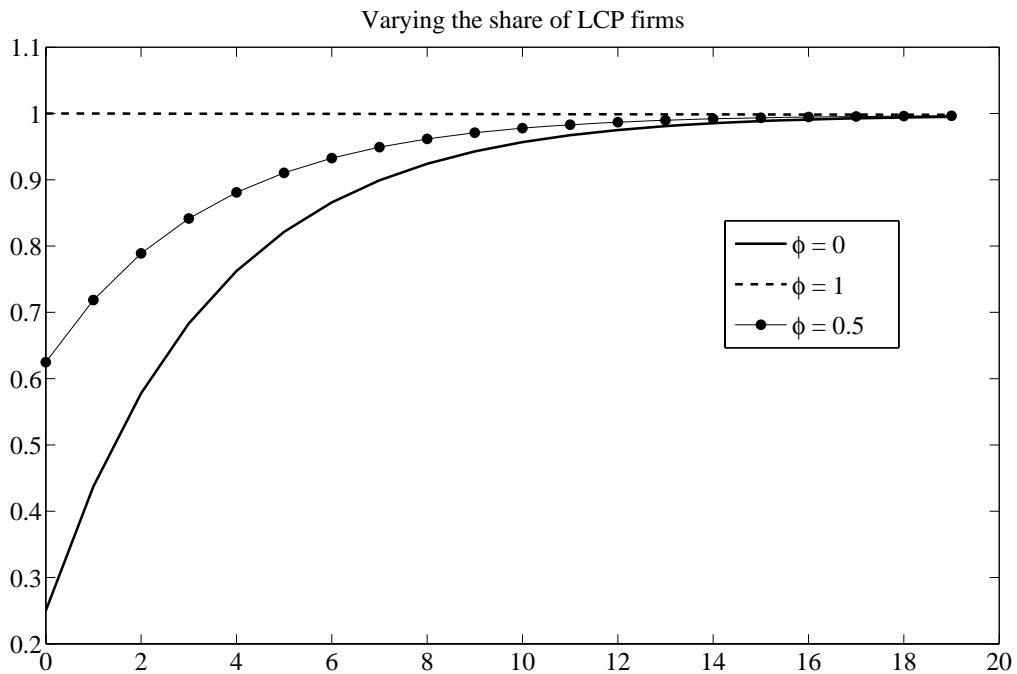


Figure 4: The response of import prices to 1% permanent exchange rate shock for different values of the share of LCP firms  $\phi$  in the LCP-PCP model.  $\beta = 0.99, \eta = 0.75$ .

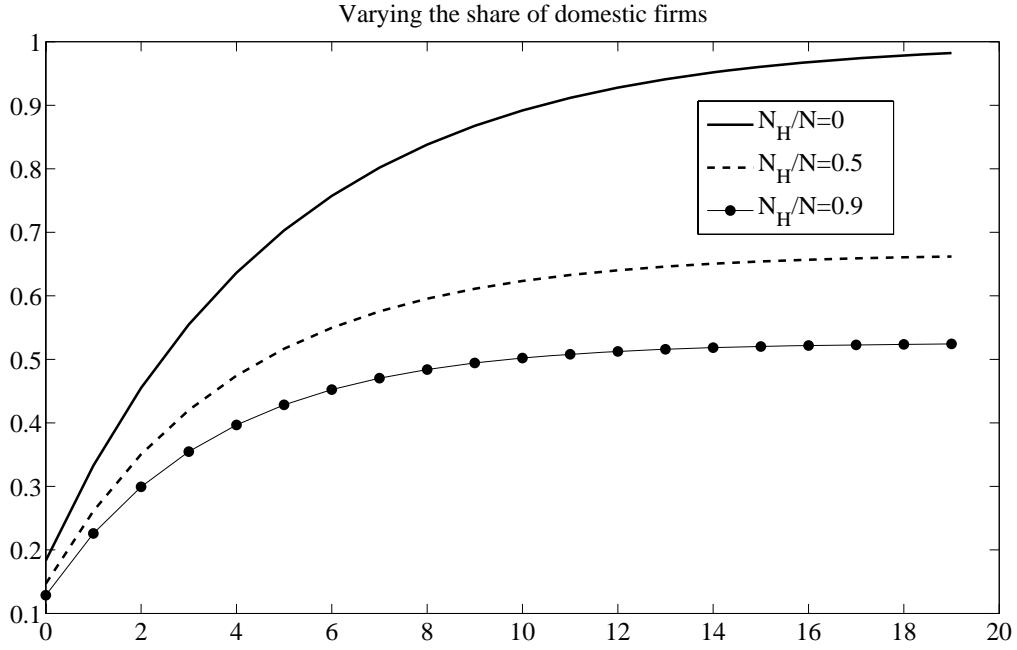


Figure 5: The response of import prices to 1% permanent exchange rate shock for different shares of import-competing firms  $N_H/N$  in the model with translog preferences.  $\beta = 0.99, \eta = 0.75$ .

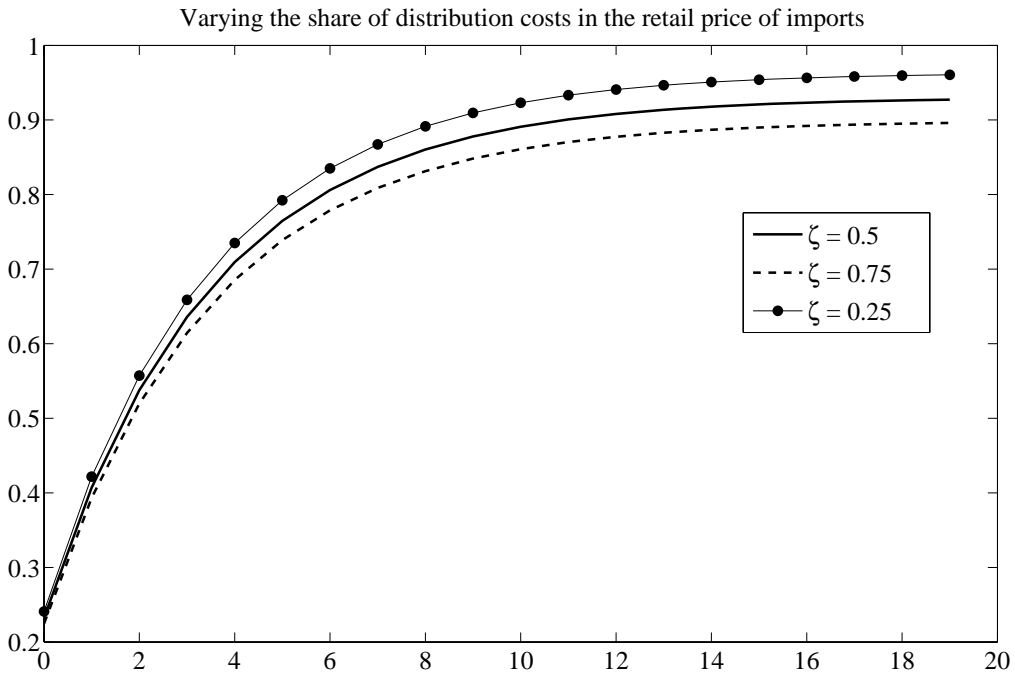


Figure 6: The response of import prices to 1 % permanent exchange rate shock for different values of the steady-state share of distribution costs in the retail price of imports  $\zeta$  in the distribution cost model.  $\beta = 0.99, \eta = 0.75$ .

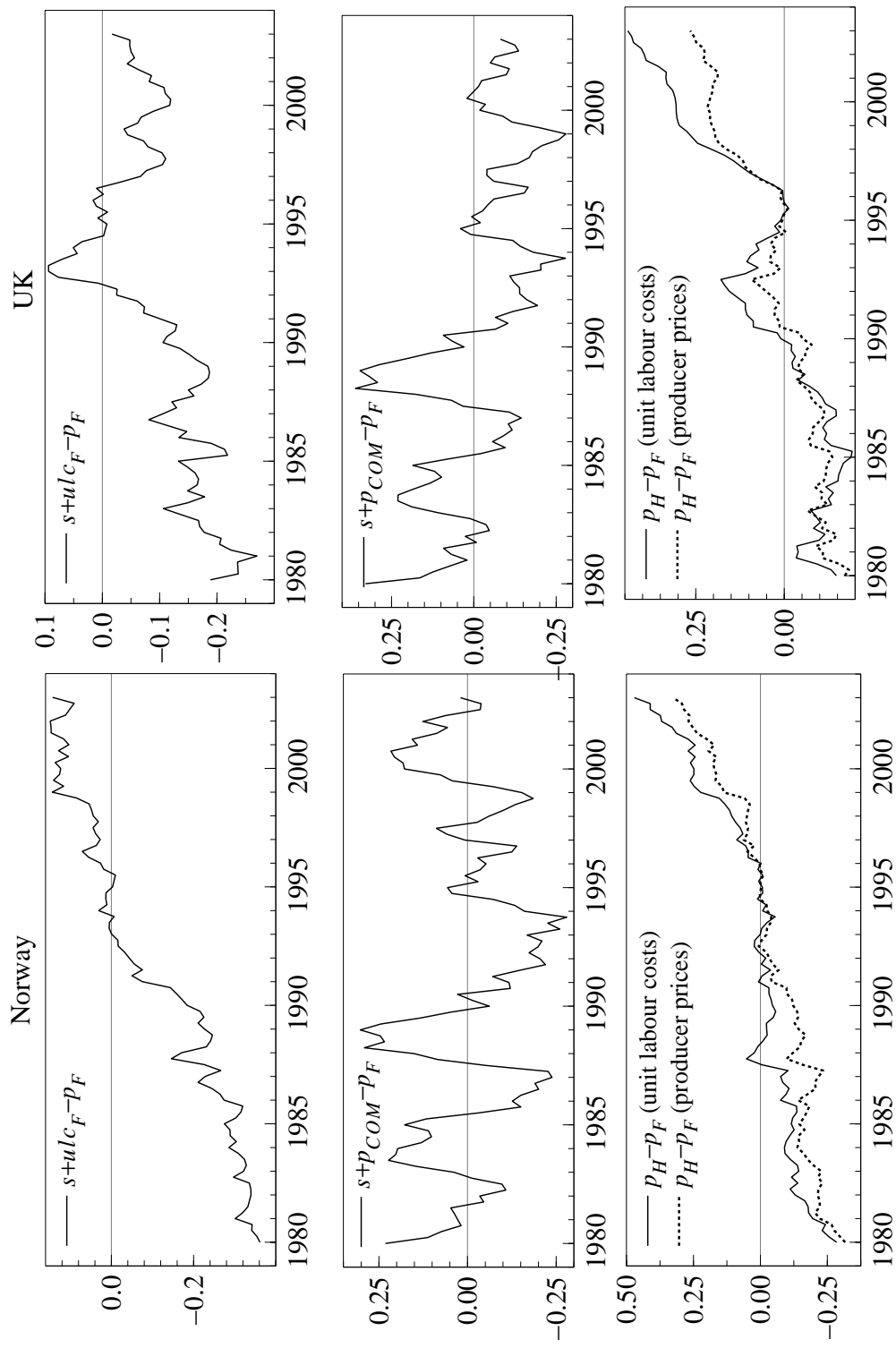


Figure 7: Data series 1980Q1–2003Q1

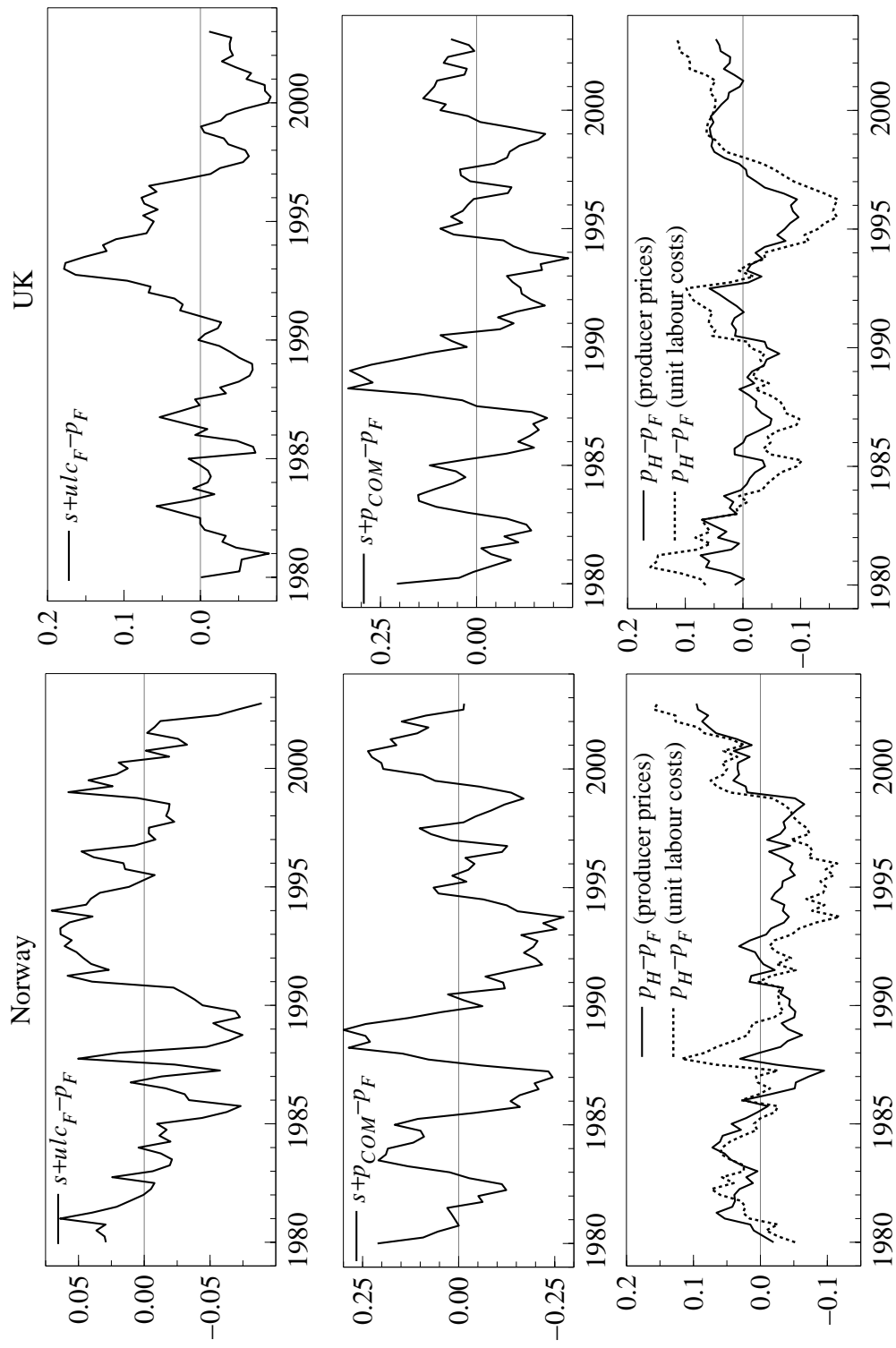


Figure 8: Detrended data series 1980Q1–2003Q1

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