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Expectations in Export Price Formation

Tests using Cointegrated VAR Models

Abstract:
The formation of export prices is an area in which the linear quadratic adjustment cost (*LQAC*) model under rational expectations may be relevant in practice. This paper evaluates the empirical performance of the *LQAC*-model using Norwegian data and a new testing procedure suggested by Johansen and Swensen (1999). We find, however, that the model can be rejected for our data set. Conversely, we show in light of Hendry (1988) that there exists a data-coherent conditional equilibrium correction (*EqCM*) model, which is not subject to the Lucas critique. Our findings do not support the claim that Norwegian exporters act on expectations based models in the formation of prices.

Keywords: Expectations, export prices, *LQAC*-model, cointegrated VAR, *EqCM*-model, exogeneity, Lucas critique.

JEL classification: C51, C52, D84, E31.

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1. Introduction

While the usefulness of cointegration analysis for studying long run relationships is hardly doubted, the modelling of short run behaviour is more of an unsettled issue. Typically, data based methods are used to develop conditional equilibrium correction models (*EqCM*), but theory based methods are also widely applied in the literature. The most prominent theory based model used to explain short run dynamics is the linear quadratic adjustment cost (*LQAC*) model under rational expectations.[1] Proponents of this approach often cite the Lucas critique as motivation for discarding conditional models and argue that *LQAC*-models may circumvent the critique if economic agents indeed are forward-looking. Estimation of such a model is however not compelling evidence in itself for or against the Lucas critique, even if the proposed *LQAC*-model is not rejected. As discussed in Ericsson and Irons (1995), one should also demonstrate the weaknesses of competing models and verify that those models suffer from the Lucas critique. Conversely, rejection of a particular *LQAC*-model does not necessarily preclude economic agents from acting on alternative expectation based models. Interestingly, Hendry (1988) argues that one may rule out *any* expectations based model, not only the *LQAC*-model, by demonstrating the constancy of a conditional *EqCM*-model together with the non-constancy of the marginal processes. It is thus surprising that Hendry's (1988) approach is rarely utilised as an alternative test in the literature studying the empirical performance of the *LQAC*-model.

The formation of export prices is an area in which the *LQAC*-model under rational expectations may find relevance in practice. Certainly, Cuthbertson (1986, 1990) and Price (1991, 1992) conclude that the empirical performance of the model using UK manufacturing export price data compares favourably with a conditional *EqCM*-model. However, not much evidence is given in these studies about the constancy or otherwise of the conditional and/or marginal models. One may then in accordance with Hendry (1988) argue that the *LQAC*-models and the conditional *EqCM*-models reported in these studies are hard to distinguish empirically. Another possible ambiguity inherent in the results in Cuthbertson (1990) concerns the use of the Johansen (1988, 1991) cointegration method. Cuthbertson (1990) cannot reject the existence of two cointegration vectors at conventional levels, but despite this finding he proceed as if only one vector exists. It is now well known that such a strategy is critical when the single cointegration vectors have not been formally identified. By way of contrast to the above mentioned studies, Svendsen (1998) using Norwegian export price data finds little evidence in favour of the LQAC-model. It is, however, most likely that the tests reported in that study are misleading since crucial conditions underlying the utilised method of testing are not met by the data. Also, the export price formation implied by the standard economic theory is most likely to apply for narrow aggregates of homogenous products, not for broad aggregates of all goods and services produced in an economy, which is the case in Svendsen (1998). Recent studies of export prices, which do not consider the *LQAC*-model, may also be subject to limitations. For instance, Smith (2000), on UK data, and Bowitz and Cappelen (2000), on Norwegian data, estimate conditional *EqCM*-models without testing properly for weak exogeneity and parameter stability. Conditional analysis is *only* valid if the conditioning variables are weakly exogenous for all parameters of interest, cf. e.g. Engle *et al.* (1983) and Hendry and Neale (1988). This paper is thus motivated to take a closer look at the empirical performance of both the *LQAC*-model and the conditional *EqCM*-model using Norwegian manufacturing export price data.

The point of departure of our analyses is the standard assumption of imperfectly competitive markets. Consequently, the optimal export price is determined as a mark-up over marginal costs. As opposed to related studies, we define marginal costs as a fully specified cost-function consistent with economic

[1] See e.g. Engsted and Haldrup (1994, 1997) and the references therein.

theory. The mark-up in turn is modelled as depending on relative prices between Norwegian and competing prices. Our tests of the *LQAC*-model are based on a new testing procedure suitable for exact rational expectations, suggested by Johansen and Swensen (1999). The idea behind this approach is to formulate the restrictions implied by the Euler equation from the *LQAC*-model within a cointegrated vector autoregressive (VAR) model. Noticeably, the only necessary assumption is that the variables are integrated of order one. No assumption of exogeneity or the number of cointegrating relations needs to be made, as is often the case in alternative testing methods, see e.g. Gregory *et al.* (1993), Gregory (1994) and Engsted and Haldrup (1994, 1997).[2] Hence, we believe that the Johansen and Swensen (1999) method is more appropriate and flexible than alternative approaches. If a well-specified cointegrated VAR can be estimated, it can be compared to a cointegrated VAR with the restrictions implied by rational expectations. We thus start out by fitting a VAR to the data, then cheque for the cointegration rank or the number of cointegrating vectors and finally estimate the cointegrated VAR with rational expectations restrictions imposed.

The results turn out to be different from those reported in Cuthbertson (1986, 1990) and Price (1991, 1992) in the sense that there is overwhelming empirical evidence against the *LQAC*-model in our case. This conclusion is not sensitive to an alternative and more traditional way of measuring marginal costs, namely that of variable unit costs. Even if there are objections to the particular form of the *LQAC*-model, as discussed in the paper, the overwhelming rejection of the model indicates that any modifications are unlikely to conform to reality. We thus follow the tests proposed by Hendry (1988) and show that there exists a data-coherent conditional *EqCM*-model of Norwegian manufacturing export prices, which is not subject to the Lucas critique. These findings rule out *any* expectation based model, not only the restrictive *LQAC*-model, as an explanation of our data.

The remainder of the paper is organised as follows: Section 2 outlines the economic background, while Section 3 describes the data used in the analyses. Section 4 presents the empirical findings and Section 5 concludes.

2. The Economic Background

We assume that Norwegian products aimed for exports are differentiated from similar goods produced in other countries. The Norwegian producers are consequently assumed to face regular downward sloping demand curves. Profit maximisation then leads to the standard formulae stating that the export price (*PA*) equals a mark-up (*MU*) times marginal costs (*MC*):

$$(1) \qquad PA = MU \cdot MC \,.$$

It is not clear how one should measure or proxy the two theoretical concepts *MU* and *MC*. The standard specification of marginal costs is some variant of variable unit costs. Such costs correspond *MC* well if firms have constant returns to scale in production, but this may be rather restrictive. The most common practice is then to include a measure of capacity utilisation as a separate variable, cf. Cuthbertson (1986), Price (1992), Smith (2000) and Bowitz and Cappelen (2000). There is a further ambiguity in that it is not clear whether marginal costs should be interpreted as marginal costs related to all production factors being variable or some being quasi fixed (e.g. capital). Most studies rely on a measure of *MC* that follows from labour and materials (including energy) being variable factors and capital being fixed. We follow that tradition.

[2] Following Hendry and Neale (1988), weak exogeneity of forcing variables is precluded if economic agents act on expectations based models of those variables. It is thus puzzling to us that some tests of the *LQAC*-model in the literature are based on the assumption of exogeneity.

However, our measure of marginal costs is based on a fully specified cost function in line with neoclassical producer behaviour. We may outline the construction of *MC* as follows: First, we specify a Cobb-Douglas production function relating gross output to the inputs of capital, labour, materials and energy. Next, minimising costs with respect to variable inputs allows us to derive a *dual* cost function for variable factor costs as a function of gross output, the capital stock, the input prices of labour, materials and energy and a time trend reflecting technical change. Marginal costs are then found by differentiating this function with respect to gross output. Thus, we specify *MC* as

(2) $$MC = f(Y, K, W, PU, PM, t, \lambda),$$

where Y is gross output, K is the capital stock, W is the wage rate, PU is a price index of energy inputs, PM is a price index of materials, t is a time trend and λ is a vector containing the parameters of $f(\cdot)$. Finally, we attain a (data based) marginal cost variable by utilising estimates of the different parameters of λ.[3] The rationale of using the cost function (2) is that our MC is more theoretically consistent than the standard approach. A capacity utilisation measure such as the modified Wharton index includes gross output, the capital stock and a trend. Additionally, utilisation measures are usually regarded as stationary variables, which in our view unnecessary complicates the cointegration analysis in the present case, cf. Rahbek and Mosconi (1999). And, as noticed in the introduction, the Johansen and Swensen (1999) method requires all variables to be integrated of order one, i.e., I(1).

The mark-up in (1) is in general a function of all the variables that enter the demand function facing exporters. If we denote the competing price that Norwegian exporters face for *PF* and the income or demand shift variable for *I*, we have in general:

(3) $$MU = m(PA/PF, I).$$

In our chosen specification, we have excluded the income element from the mark-up factor. We have not included separate demand effects in the analyses for two reasons. First, any demand effects are likely to be captured by competing prices and the marginal cost measure. Second, most previous studies pay little attention to demand effects and we wish to compare our results with those reported in existing studies. From (1) and (3) we may now obtain:

(4) $$PA = g(MC, PF).$$

Equation (4) is homogenous of degree one in *MC* and *PF*. Although our price equation is derived from a theory of monopolistic competition, it also encompasses the main alternative, namely that of "the law of one price" or perfect competition for homogenous goods. In the latter case the export price is equal to the price of the competitors, so that $PA = PF$. On the other hand, perfect competition also implies $PA = MC$, but now the causality runs the other way in that world market prices determine domestic marginal costs.

We may interpret (4) as a *static* equilibrium describing an exporter decision rule in the *long run*. Hence, it will serve as the starting point for the cointegration analysis in Section 4. It is however likely that short run deviations from the long run target in response to shocks are rational in an environment with adjustment costs. Accordingly, an economic model that takes this possibility into account seems a natural next step in modelling the export price. In a set-up involving rational expectations, the forward-looking *LQAC*-model

[3] Applying Shepard's lemma to the *dual* cost function, we may show that the optimal levels of variable factors of production under cost-minimisation are functions of the same parameters of λ as in (2). Thus, the estimates of λ in this paper are deduced from data-coherent conditional *EqCM*-models of variable factors of production. The estimates of λ are not reported, but are provided upon request.

generally attempts to reconcile costs incurred by not hitting the target on the one hand and costs resulting from changing the variables on the other hand. A possible formulation for the situation considered in this paper is that the export prices are chosen in order to minimise the intertemporal quadratic loss function

$$(5) \qquad E_t\left[\sum_{j=0}^{\infty} \beta^j [\theta(pa_{t+j} - pa^*_{t+j})^2 + (pa_{t+j} - pa_{t+j-1})^2]\right],$$

where E_t denotes the conditional expectation given the information contained in the information set at time t, pa^*_t represents the *long run* target and lower case letters indicate logs, i.e., $pa_{t+j} = \log(PA_{t+j})$. The parameter β represents the discount rate and θ the relative cost parameter of the two terms of the loss function. As can be seen from (5), exporters determine a sequence of export prices so as to minimise the expected present discounted value of all future squared deviations from the long run target. However, since changes in the export price will be penalised as well, immediate adjustment towards the long run target will be non-optimal unless θ is large. Thus, the introduction of the second term in (5) makes the decision rule *dynamic* rather than *static*, as suggested by (4).

As is well known, see e.g. Sargent (1978), Nickell (1985) and Engsted and Haldrup (1994, 1997), the first order condition for this minimisation problem is an Euler equation of the form

$$(6) \qquad \Delta pa_t = \beta E_t[\Delta pa_{t+1}] - \theta(pa_t - pa^*_t).$$

When the long run target is a function of the form

$$(7) \qquad pa^*_t = \gamma_1 pf_t + \gamma_2 mc_t,$$

and if we let $X_t = (pa_t, pf_t, mc_t)'$, the Euler equation reduces to

$$(8) \qquad (\beta, 0, 0)E_t[X_{t+1}] - (1 + \beta + \theta, -\theta\gamma_1, -\theta\gamma_2)X_t + (1,0,0)X_{t-1} = 0.$$

We notice that (7) is a logarithmic transformation of (4), such that γ_1 and γ_2 give the long run elasticity of the export price with respect to the competing price and the marginal costs, respectively. The homogeneity restriction predicted by our theory implies that $\gamma_1 + \gamma_2 = 1$. In the case of perfect competition for homogenous goods, the homogeneity restriction reduces to $\gamma_1 = 1$.

Another possible formulation is that the long run target has the form

$$(9) \qquad pa^*_t = \gamma_1 E_t[pf_{t+1}] + \gamma_2 E_t[mc_{t+1}].$$

This seems as a sensible choice when the aim is to take estimated future values of the target into account. Proceeding as earlier, the Euler equation may now be written

$$(10) \qquad (\beta, \theta\gamma_1, \theta\gamma_2)E_t[X_{t+1}] - (1 + \beta + \theta, 0, 0)X_t + (1,0,0)X_{t-1} = 0.$$

Regardless of the form of the target, we observe that (8) and (10) for fixed values of β and θ contain *linear* restrictions involving the conditional expected value of the observations *one-step-ahead* and the present and lagged observed values. In other words, the restrictions satisfying the Euler equations entail restrictions on the long run relationships as well as some restrictions on the short run parameters. When the information set is specified the conditional expectation $E_t[X_{t+1}]$ can be worked out and the restrictions stated explicitly. This way of formulating the restrictions is the foundation for the testing procedure advocated by Johansen and Swensen (1999). As explained in that paper, the idea is to start out with a well specified VAR and test, using a likelihood ratio test, the restrictions deduced from the Euler equation of the *LQAC*-model as restrictions on the coefficients of the VAR. That is, one has to work out the maximum likelihood estimator of the cointegrated VAR, both with and without the rational expectation restrictions imposed, in order to construct the likelihood ratio test. The procedure is further discussed below in relation to our particular VAR.

3. The Data

The empirical analyses are conducted using quarterly, seasonally unadjusted data that spans the period 1978(1)–1998(4).[4] In line with related studies, we study export prices of manufactures. Specifically, we model the export price (*PA*) of machinery and metal products excluding ships, oil platforms, chemicals, pulp and paper and basic consumption goods, i.e., food, clothes, beverages and tobacco. The chosen aggregate accounts for approximately 25 per cent of total Norwegian manufacturing exports. Our relatively narrow aggregate is motivated so as to reduce the risk of instability in estimated parameters caused by broad aggregation of heterogeneous products. We employ the Norwegian import price (*PF*) of machinery and metal products as a proxy for the competing price. Referring to Naug and Nymoen (1996), one may argue that *PF* is a poor proxy for competing prices. However, preliminary investigations showed in sharp contrast to Naug and Nymoen (1996) that changes in foreign prices and the exchange rate are quickly and completely passed through to the import price of machinery and metal products. One explanation for the different findings may be that the mentioned study models a much broader aggregate than what *PF* represents. Another argument for using our proxy of competing prices is that *PF* is product specific, while the available foreign price and the exchange rate are not. The data for *PA* and *PF* are indices with 1996 as base year and are expressed in the Norwegian currency (NOK).[5] As explained in Section 2, a data based marginal cost variable (*MC*) serves as a proxy for production costs. We refer to the Appendix for further details on the data definitions and sources. All variables are in logarithms and denoted with lower case letters in the succeeding analyses. To account for seasonality in the data, we include three centred seasonal dummies, labelled CS_{1t}, CS_{2t} and CS_{3t}, in all regressions.

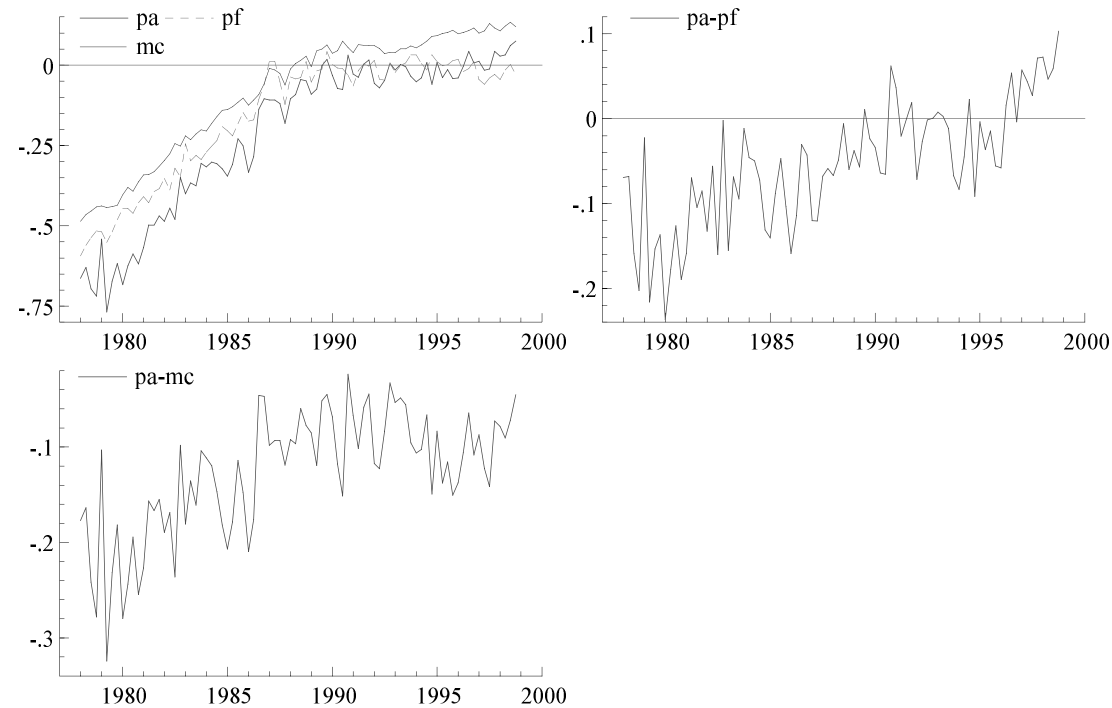
We now examine the data with reference to time series properties. Figure 1 displays the export price, the competing price and the marginal costs, together with the two ratios $(pa–pf)_t$ and $(pa–mc)_t$ over the sample period. We observe that the two price series as well as the marginal costs exhibit a clear upward trend, suggesting pa_t, pf_t and mc_t to be I(1) or possibly I(2). The former hypothesis of I(1)-variables is supported by various augmented Dickey-Fuller tests.[6] Consequently, it seems safe to carry out the Johansen and Swensen (1999) methodology on our data set. We also notice that Norwegian export prices have increased relative to both competing prices and marginal costs over the entire period. The movements in the ratios over the business cycle offer a somewhat more complex picture. For instance, pa_t increased relative to pf_t from 1980 to 1984, decreased considerably in 1985 and then picked up steadily from 1986 to 1990. Overall, our data set seems to suggest that both competing prices and marginal costs are important candidates for explaining the Norwegian export price of machinery and metal products. It is not obvious, however, from Figure 1 whether the data series have the property of forming a cointegrating vector(s) that is consistent with our economic model.

[4] The data is taken from the Quarterly National Accounts (QNA), published by Statistics Norway unless otherwise stated. Due to a revision of the QNA, we were prevented from extending the series by data before 1978.

[5] It may be a problem that the underlying model in Section 2 explains the prices in levels, while the available data are expressed as indices.

[6] To simplify matters, we follow previous studies in ignoring any possible presence of seasonal integration in our data series [cf. Hylleberg *et al*. (1990)]. As such, it is worthwile to mention that the conventional ADF-tests used here only give rough guidance to the properties of integration. The results of the ADF-tests are not reported, but are provided upon request.

Figure 1. The log of export prices (*pa*), competing prices (*pf*) and marginal costs (*mc*)



4. Empirical Results

In this section we consider the empirical content of the *LQAC*-model under rational expectations. Our strategy in testing the model is as follows: First, we utilise the Johansen (1988, 1991) method to establish the empirical counterpart (or lack thereof) of the long run target (7). The results from the cointegration analysis are then used to test the *LQAC*-model within the Johansen and Swensen (1999) framework. To shed further light on the empirical role of expectations based models, we also develop a conditional *EqCM*-model, and examine whether this model suffers from the Lucas critique. For this purpose, we apply the testing procedures suggested by Hendry (1988).

4.1 Cointegration Analysis

The cointegration analysis commenced from a 5th order VAR in $X_t = (pa_t, pf_t, mc_t)'$.[7] This model was augmented with an unrestricted intercept and the three centred seasonal dummies. Preliminary investigations suggested that a 4th order VAR is adequate to render residuals with statistically acceptable properties. Further simplifications yielded a mc_t-equation that suffered from residual heteroskedasticity and autocorrelation. Residual misspesification tests for the 4th order VAR are reported in Table 1. It should be noticed that the preferred VAR includes two additional unrestricted

[7] We follow existing studies and simplify matters by assuming that the export market can be distinguished from the home market. This assumption may pose caveats if the export price and the home price are set simultaneously by Norwegian manufacturers. We investigated this possibility on our data set expanded with data on the home price of machinery and metal products. Using Johansen's procedure, we found some evidence of two significant cointegrating vectors among the four variables. A *priori*, one of the vectors should most naturally be interpreted as a long run export price equation, whereas the other vector should be interpreted as a long run home price schedule. The results from testing economically meaningful identifying restrictions on the two vectors were, however, not very encouraging. We thus decided to leave this issue on the agenda for future research.

impulse dummies that account for outliers in the pa_t-equation in 1979(1) and 1979(2). This may be unfortunate since the information set was chosen to explain pa_t. However, the conclusions from the cointegration analysis are not significantly affected by these variables.

Table 1. Residual misspesification tests[1]

	Statistics						
	$AR_{1\text{-}5}$	$ARCH_{1\text{-}4}$	$NORM$	HET_1	$AR^V_{1\text{-}5}$	$NORM^V$	HET^V_1
Equation	$F(5, 58)$	$F(4, 55)$	$\chi^2(2)$	$F(24, 38)$	$F(45, 137)$	$\chi^2(6)$	$F(144,200)$
pa	0.314	0.771	1.224	0.754			
pf	1.024	0.140	2.246	0.624			
mc	0.893	1.596	3.491	0.660			
VAR					0.836	1.689	0.549

[1] $AR_{1\text{-}5}$ is Harvey's (1981) test for 5th order residual autocorrelation; $ARCH_{1\text{-}4}$ is the Engle (1982) test for 4th order autoregressive conditional heteroskedasticity in the residuals; *NORM* is the normality test described in Doornik and Hansen (1994) and HET_1 is a test for residual heteroskedasticity due to White (1980). Similar tests for the entire VAR is denoted by ^{V} [see Doornik and Hendry (1997)]. $F(\cdot)$ and $\chi^2(\cdot)$ represent the null distributions of F and χ^2, with degrees of freedom shown in parenthesis.

For our VAR to be considered as a valid starting point of the cointegration analysis, it should also contain reasonably constant parameters. Recursive estimates (with $\pm$ 2 standard errors) and sequences of break-point Chow tests (scaled by their 5 per cent critical values) are shown in Figure 2. We conclude that the system is constant over the sample. The next step is thus to investigate the cointegration properties between the selected variables.

Figure 2. Recursive test statistics for the preferred VAR

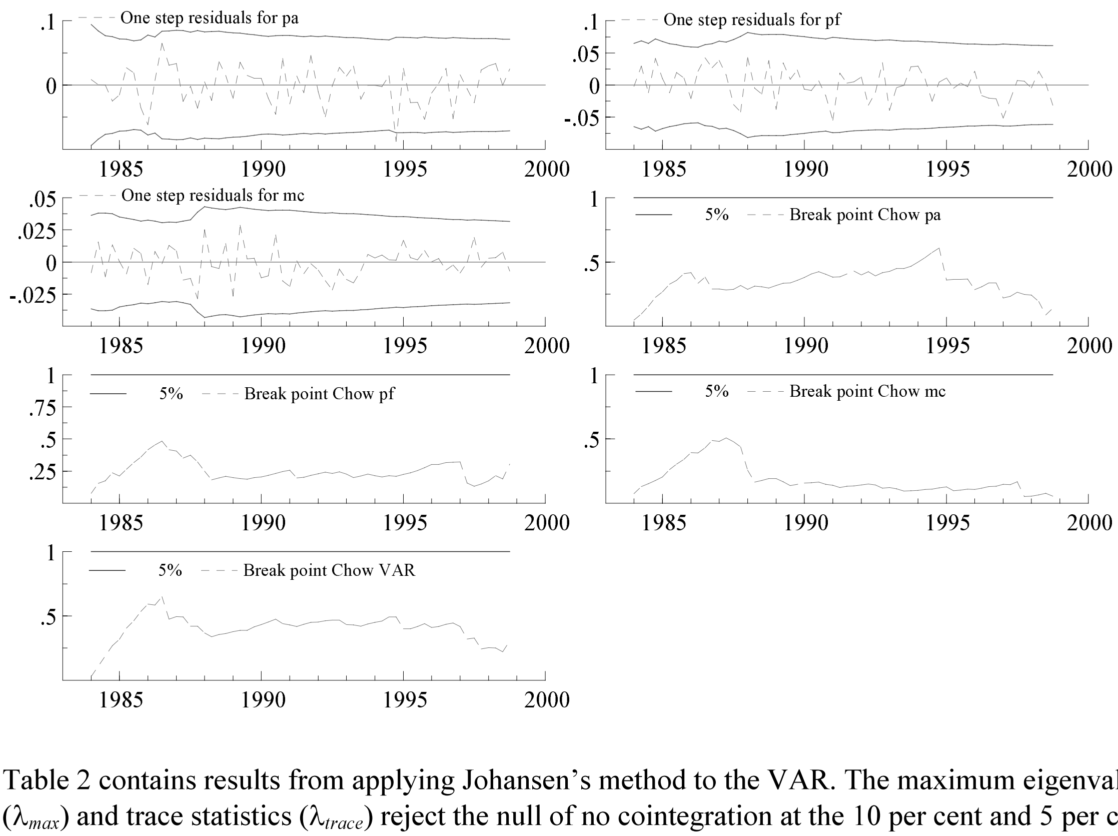


Table 2 contains results from applying Johansen's method to the VAR. The maximum eigenvalue (λ_{max}) and trace statistics (λ_{trace}) reject the null of no cointegration at the 10 per cent and 5 per cent levels, respectively, but the null of at most one cointegration vector is not rejected by any of the statistics. Recursive estimation of the eigenvalues (not reported) supports our conclusion that there is a single cointegrating vector between *pa*, *pf* and *mc*. Accordingly, the tests of the *LQAC*-model will be performed under the assumption that the cointegration rank equals unity. The estimate of the *unrestricted* cointegrating vector (normalised on *pa*) is interpretable as a long run export price equation. As seen, the estimates of the long run export price elasticities (i.e., γ_1 and γ_2) are economically reasonable and statistically significant. Besides, the results of the weak exogeneity tests [see Johansen and Juselius (1990)] imply that the cointegrating vector enters the pa_t-equation only, albeit the mc_t-equation may be a borderline case in this respect. An explanation of the latter result is that production costs increase due to, say, increased wages in the labour market. The wage increase in turn may lead to a depreciation of the NOK (due to a loss in the competitiveness) and a subsequent increase in the export price. Another channel from which the marginal costs may not be weakly exogenous is through the competing price. That is, the competing price affects the price index for material inputs that enters our marginal cost function, cf. equation (2).

Table 2. Johansen's cointegration tests[1]

Information set: (*pa*, *pf*, *mc*)
Eigenvalues: 0.220, 0.128, 0.040

	Statistics					
Hypothesis	λ_{max}	λ^a_{max}	95%	λ_{trace}	λ^a_{trace}	95%
$r=0$	19.90	16.92	21.00	34.15*	29.03	29.7
$r\leq1$	11.00	9.35	14.10	14.25	12.11	15.4
$r\leq2$	3.25	2.76	3.80	3.25	2.76	3.8

Estimate of the *unrestricted* cointegrating vector:[2]

$$pa = \hat{\alpha}_0 + \underset{(0.252)}{0.525} pf + \underset{(0.256)}{0.585} mc$$

Weak exogeneity tests:[3]	*pa*	*pf*	*mc*
$\chi^2(1)$	7.409**	2.396	3.269
p-value	(0.007)	(0.122)	(0.071)

[1] *r* denotes the cointegration rank, i.e., the number of cointegrating vectors. The λ_{max} and λ_{trace} statistics are the maximum eigenvalue and trace statistics, whereas λ^a_{max} and λ^a_{trace} are the corresponding statistics with a degrees-of-freedom-adjustment, cf. Reimers (1992). The 95 per cent quantiles are taken from Table 1 in Osterwald-Lenum (1992). Asterisk * and ** denote rejection of the null hypothesis at the 5 per cent and 1 per cent significance levels.[2] The figures in parentheses are standard errors.[3] The tests, which are asymptotically distributed as $\chi^2(1)$ under the null, are calculated under the assumption that r=1.

We also notice that the sum of the unrestricted estimates of the export price elasticities (i.e., $\gamma_1 + \gamma_2$) is not far from unity, as predicted by theory. To complete the cointegration analysis, we thus tested for, and could not reject, the existence of long run homogeneity between *pa*, *pf* and *mc*. The restrictions of long run homogeneity and weak exogeneity of *pf* and *mc*, both individually and jointly, were however rejected by the data. Imposing the homogeneity restriction only gives $\chi^2(1)$=0.487 (with a *p*-value of 0.485) and the following *restricted* estimate of the cointegrating vector (normalised on *pa*):

$$pa = \hat{\alpha}_0 + \underset{(0.381)}{0.645} pf + 0.355 mc, \tag{11}$$

where the standard error is in parenthesis. Perhaps a worrying feature of the cointegrating vector in (11) is the relatively small *t*-value of 1.693 on the coefficient of *pf*. This value may, however, be considered statistically significant because, for a one-sided test, its *p*-value is 0.047. A sequence of $\chi^2(1)$ test statistics (not shown) also confirms the validity of the homogeneity restriction for any sample size between 1985 and 1998. Similarly, the recursively estimated parameter of *pf* in (11) is reasonably stable. In previous studies of Norwegian export prices, the homogeneity restriction is also found to hold, see Svendsen (1998) and Bowitz and Cappelen (2000). However, the cost and competing price elasticity in these studies differ somewhat from those reported in (11) as they are both estimated to 0.5. Interestingly, Cuthbertson (1990) also find equal weights of the elasticities when homogeneity is imposed in a study of the UK manufacturing export price. It should be stressed from the outset that a direct comparison between our estimates and those in the mentioned studies is impossible due to use of different aggregation levels and different treatments of the marginal cost variable.

Finally, the adjustment coefficient associated with (11) is estimated to −0.277. Economically, the estimate means that around 28 per cent of disequilibrium in (11) are corrected within one quarter. The short run dynamics in *pa* is not however, fully described by this coefficient. Given its importance for policy, we shall return to this issue shortly.

4.2 Tests of the *LQAC*-model

Having established the number of the cointegration vectors in our data set, we proceed to test the empirical performance of the *LQAC*-model. A few words concerning our particular VAR is however needed before we report results from the tests. We recall that a VAR of the form

(12) $$X_t = A_1 X_{t-1} + ... + A_4 X_{t-4} + \Phi D_t + \varepsilon_t, t = 1979(1),...,1998(4),$$

with reduced rank equal to one was shown to have satisfactory statistical properties. Here D_t represents the constant, the centred seasonals and the impulse dummies, as previously explained. The restrictions satisfied by the Euler equation and deduced from the *LQAC*-model described in Section 2 may be written

(13) $$c_1' E_t[X_{t+1}] + c_0' X_t + c_{-1}' X_{t-1} = 0.$$

Using the Johansen and Swensen (1999) procedure, these conditions may be tested as restrictions on the coefficients of the reduced rank VAR. However, the model contains dummy variables so a slightly different formulation makes more sense. We first notice that

(14) $$E_t[X_{t+1}] = \Phi D_{t+1} + A_1 X_t + ... + A_4 X_{t-3}.$$

It seems reasonable to consider restrictions involving the part of the forecasts of future values that do not involve the deterministic part, and test restrictions of the form

(15) $$c_1' E_t[X_{t+1} - \Phi D_{t+1}] = -c_0' X_t - c_{-1}' X_{t-1}.$$

Inserting the previous expression for $E_t[X_{t+1}]$, (15) becomes

(16) $$c_1' A_1 X_t + ... + c_1' A_4 X_{t-3} = -c_0' X_t - c_{-1}' X_{t-1},$$

such that there are no restrictions on the coefficients of the deterministic part of the VAR. We can then use the same procedure as described in Johansen and Swensen (1999) with the exception that the deterministic terms are included without restrictions when the marginal part of the model is estimated. Since the method is based upon likelihood ratio tests we have to find the maximum likelihood estimators of the coefficients of the VAR under the imposed restrictions. We now provide some details of how that is done in the present situation.

Consider the first formulation of the Euler equation in Section 2. Then $c_1' = (\beta,0,0)$, $c_0' = -(1+\beta+\theta,-\theta\gamma_1,-\theta\gamma_2)$ and $c_{-1}' = (1,0,0)$. Under these restrictions the cointegration vector, $d = -(c_1 + c_0 + c_{-1})$, equals $\theta(1,-\gamma_1,-\gamma_2)$. The long run homogeneity restriction is therefore $1-\gamma_1-\gamma_2 = 0$. Imposing this in addition to the restrictions from the Euler equation yields $c_0' = -[1+\beta+\theta,-\theta\gamma_1,-\theta(1-\gamma_1)]$. Thus, the c'-vectors depend on the unknown parameters β, θ and $\gamma=\gamma_1$. Proceeding as in Johansen and Swensen (1999) we can then evaluate the likelihood of the reduced rank VAR with the restrictions from the Euler equation assuming β, θ and γ to be known. We thus obtain a value of the likelihood as a function of β, θ and γ over a grid of possible values.

Referring to the notation in Johansen and Swensen (1999), $b' = c_1' = (\beta,0,0)$ in our case, so we may set the 3 x 2 matrix a equal to the matrix (e_2 e_3), where e_2 and e_3 are the 3 x 1 unit vectors having one as the second and third element, respectively, and else consisting of zeros. Since the rank of the matrix c_1', q, and the reduced rank of the VAR model, r, both are equal to unity in our case, we first determine the conditional model by regressing $a'\Delta X_t = (\Delta pi_t, \Delta mc_t)'$ on $b'\Delta X_t - c_{-1}'\Delta X_{t-1}, d'X_{t-1}, \Delta X_{t-1},\ldots,\Delta X_{t-3}$, the constant, the seasonals and the impulse dummies. Notice that $b'\Delta X_t - c_{-1}'\Delta X_{t-1} = (\beta\Delta pa_t - \Delta pa_{t-1},0,0)$ and $d'X_{t-1} = \theta[pa_{t-1} - \gamma pf_{t-1} - (1-\gamma)mc_{t-1}]$. To find the contribution from the marginal model we then, in accordance with the modifications above, regress $b'\Delta X_t - d'X_{t-1} - c_{-1}'\Delta X_{t-1} = \beta\Delta pa_t - \Delta pa_{t-1} - \theta pa_{t-1} + \theta\gamma pf_{t-1} + \theta(1-\gamma)mc_{t-1}$ on the constant, the seasonals and the impulse dummies. Let the residuals from the first regression be R_{1t} and let the mean sum of the matrices $R_{1t}R_{1t}'$ be S_{11}. Compute the product of the determinant of S_{11} and the mean sum of squares from the last regression. Finally, the value of the maximised likelihood for fixed values of (β, θ, γ) is this product divided by $b'b$ det($a'a$) = β^2.

After some experimentation to obtain the largest values of the likelihood we used $\beta = 0.90, 0.91,\ldots,0.99$, $\theta = 0.05, 0.10,\ldots,0.25$ and $\gamma=0.2,0.4,\ldots,0.8$ as a grid. The log likelihood then varied between 877.0 and 886.6. Compared to the value 931.5, which is the log likelihood of the cointegrated VAR without the rational expectation restrictions imposed, this yields a –2 log likelihood ratio statistic between 89.8 and 109.0. Regardless of whether the parameter values β, θ and γ are considered as fixed, corresponding to 12 degrees of freedom, or estimated, corresponding to 9 degrees of freedom, there is thus overwhelming empirical evidence against the *LQAC*-model in our case. There is not much difference using the other possible formulation of the target value, as described in Section 2. In this case the log likelihood varied between 877.7 and 886.8.

One may speculate over the reasons for the poor performance of the *LQAC*-model in the present situation. Initially, we tested two different specifications of the *LQAC*-model. In the first specification the loss function (5) was defined for *pa* with *mc* as the cost variable, whereas in the second specification it was defined with average unit variable production costs as the cost variable. However, we have only reported results based on *mc* as the two specifications produced similar outcomes. The form of the loss function (5) is also of course not the only one possible. In particular, losses are assumed to have symmetric effects in the formulation we have used. This may be questionable. Over- and undershooting of a target value may have different effects in general, and may also be the case for adjustment costs. Additionally, the loss function does not reflect that the data are quarterly. A loss function taking this into account may be more reasonable, see e.g. Todd (1990) for some work in this direction. Despite these objections the fact remains that the *LQAC*-model is strongly at odds with our data set, so any modifications must be substantial if the model is to have satisfactory empirical properties.

It may also be worth pointing out that we have considered what is called the *exact* linear rational expectation model. Another formulation is to assume that the target has the form $pa_t^* = \gamma_1 pf_t + \gamma_2 mc_t + e_t$, where e_t is a white noise error. Such a formulation emerges when the economic agents have knowledge, which is not accessible to the econometrician. For details of how such models can be estimated, see Dolado *et al.* (1991) and Gregory *et al.* (1993).

4.3 The Conditional *EqCM*-model and the Lucas Critique

A potential shortcoming of formal tests of rational expectation models, as touched upon above, is that they may not have much economic content. As Engsted and Haldrup (1994) point out: "If the test does not reject the model it could be due to low power against important alternatives, and if it does reject, we do not know whether it is caused by fundamental deviations from the model, the information set used in the analysis, or simply by economically unimportant factors". In a similar vein, the rejection of the *LQAC*-model does not preclude Norwegian exporters from acting on alternative expectations based models in the formation of prices. Thus, we are motivated to test such a hypothesis by means of Hendry's (1988) approach. Testing the constancy or otherwise of a conditional *EqCM*-model of export prices and the marginal processes of the conditioning variables seems promising, given the major shocks (e.g. exchange rate and oil price shocks) that hit the Norwegian economy during the eighties and nineties. We adopt the *general-to-specific* modelling approach advocated by Davidson *et al.* (1978) in the search for a data-coherent conditional *EqCM*-model. Applying the same information set as in the cointegration analysis and using deviations from (11) as an equilibrium correction mechanism (*EqCM*), we proceed from the following model:[8]

$$(17)\qquad \begin{aligned}\Delta pa_t = \beta_0 + \sum_{i=1}^{4}\beta_{1i}\Delta pa_{t-i} + \sum_{i=0}^{4}\beta_{2i}\Delta pf_{t-i} + \sum_{i=0}^{4}\beta_{3i}\Delta mc_{t-i}\\ + \delta[pa - 0.645pf - 0.355mc]_{t-1} + e_t,\end{aligned}$$

where Δ is the difference operator, the expression in $[\cdot]$ is the *EqCM* and e_t is the error term, assumed to be white noise. As it stands, (17) is likely to represent an overparameterisation. We thus present a parsimonious export price model (18), which is a valid simplification of (17). It may be worth noting that (18) is derived from a single equation analysis rather than a system one. Following Boswijk and Urbain (1997), one may apply single equation analysis with the long run relationship(s) estimated and deduced from a VAR model in cases where the conditioning variables are error correcting but weakly exogenous for the short run parameters. We recall that *mc* may not be weakly exogenous for the cointegration parameters. Preliminary estimations and subsequent simplifications of (17) were conducted with and without instruments for Δmc_t. This variable was, however, far from being significant in any of the regressions considered, and (18) thus omits contemporaneous effects from mc_t. Noticeably, no impact effect from Δmc_t amounts to the restriction that the covariance between the residuals of the pa_t-equation and the mc_t-equation in the VAR is zero.

$$(18)\qquad \begin{aligned}\Delta pa_t = const. &- \underset{(0.087)}{0.360}\Delta_2 pa_{t-1} - \underset{(0.090)}{0.292}\Delta pa_{t-3} - \underset{(0.081)}{0.130}\Delta pa_{t-4} + \underset{(0.120)}{0.233}\Delta pf_t + \underset{(0.262)}{0.731}\Delta mc_{t-1}\\ &- \underset{(0.070)}{0.214}(pa - 0.65pf - 0.35mc)_{t-1} - \underset{(0.009)}{0.025}CS_{1t} + \underset{(0.038)}{0.195}D79(2)\end{aligned}$$

Method: *OLS* T=79[1979(2)–1998(4)] R^2=0.601 σ=3.57% DW=2.10
$AR_{1\text{-}5}$:$F(5, 65)$=1.145 $ARCH_{1\text{-}4}$:$F(4, 62)$=0.843 $NORM$:$\chi^2(2)$=1.054
HET_1:$F(14, 55)$=1.663 HET_2:$F(35, 34)$=1.301 $RESET$:$F(1, 69)$=2.303.

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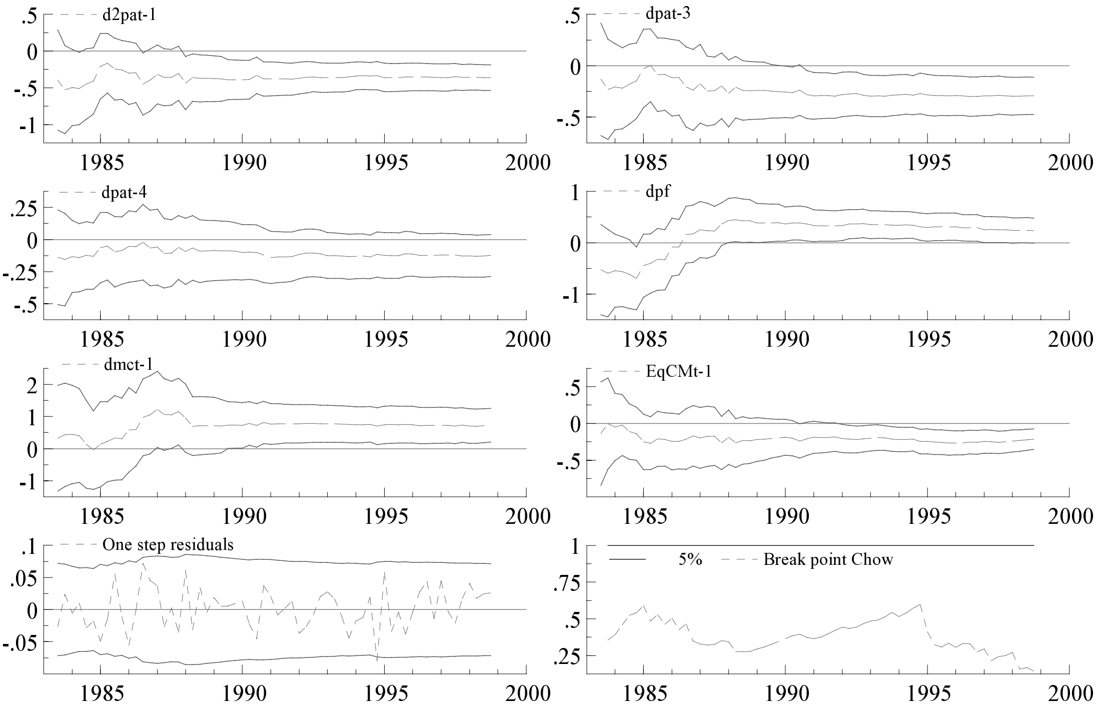
[8] The unrestricted impulse dummies and the seasonals from the VAR are suppressed for ease of exposition.

Below the conditional *EqCM*-model we report several test statistics. Estimated standard errors are in parentheses, and T, R^2, σ and DW are the number of observations, the squared multiple correlation coefficient, the residual standard error and the Durbin-Watson statistic, respectively. In addition to $AR_{1\text{-}5}$, $ARCH_{1\text{-}4}$, *NORM* and HET_1 defined in Table 1, we report HET_2 and *RESET*. The former tests whether the squared residuals depend on the levels, squares and cross products of the regressors [cf. White (1980)], while the latter tests for functional form misspesification [cf. Ramsey (1969)]. None of the diagnostics are significant at the 1 per cent level. The economic variables entering (18) are overall highly significant. For instance, the *EqCM* appears in the model with a *t*-value of −3.06, hence adding force to the results obtained from the cointegration analysis. Besides, the adjustment coefficient of −0.214 is virtually unchanged from the adjustment coefficient associated with (11). Regarding competing prices, the estimated impact elasticity of 0.23 is considerably smaller than its long run counterpart. Accordingly, Norwegian firms seem to smooth the export prices with respect to changes in the competing prices. The apparent slow adjustment of *pa* to shifts in *pf* may reflect the fact that the Norwegian exchange rate was fixed within certain bands for the most part of the sample period. If it is costly to change the export price, it may be wise to adjust slowly to changes in the competing price that are caused by temporary and *small* fluctuations in the exchange rate. One puzzle with this interpretation is that the sample period contains some *major* discrete changes in the exchange rate[9] that may have been viewed as permanent, in which case exporters may have adjusted the export price more fastly. If the short run pass-through depends on the particular type of fluctuations in the exchange rate, then (18) may be subject to the Lucas critique in the sense of being unstable owing to policy interventions, cf. Favero and Hendry (1992). However, empirical evidence of constancy of our conditional *EqCM*-model may be judged from Figure 3, which plots recursive test statistics and recursive estimates for the coefficients on key variables entering (18).

Figure 3. Recursive estimates and test statistics of (18)



[9] For instance, the Norwegian currency was devaluated by 12 per cent in May 1986. After the European currency turmoil in December 1992, the NOK was depreciated by 4 per cent. Then, in January 1997, the NOK was appreciated to its highest level against European currencies since December 1992.

The recursive estimates vary little, especially relative to their estimated uncertainty. Similarly, neither the one-step residuals (with ± 2 standard errors) nor the sequence of break point Chow tests (scaled by their 5 per cent critical values) reject constancy. The conclusion of (18) being autonomous to all conceivable interventions over the sample may be questionable. One may reject invariance due to the effects on the export price of the dummy *D*79(2), that could capture effects of OPEC II [cf. Favero and Hendry (1992)]. That is, Norwegian exporters of machinery and metal products may have adjusted *pa* in accordance with *anticipated* changes in *pf*, following the oil price shock in the late 1970s. To the extent that such expectations are present, the Lucas critique applies to (18). It is, however, interesting that (18) remained constant in spite of the reversed oil price shock in the mid-1980s.

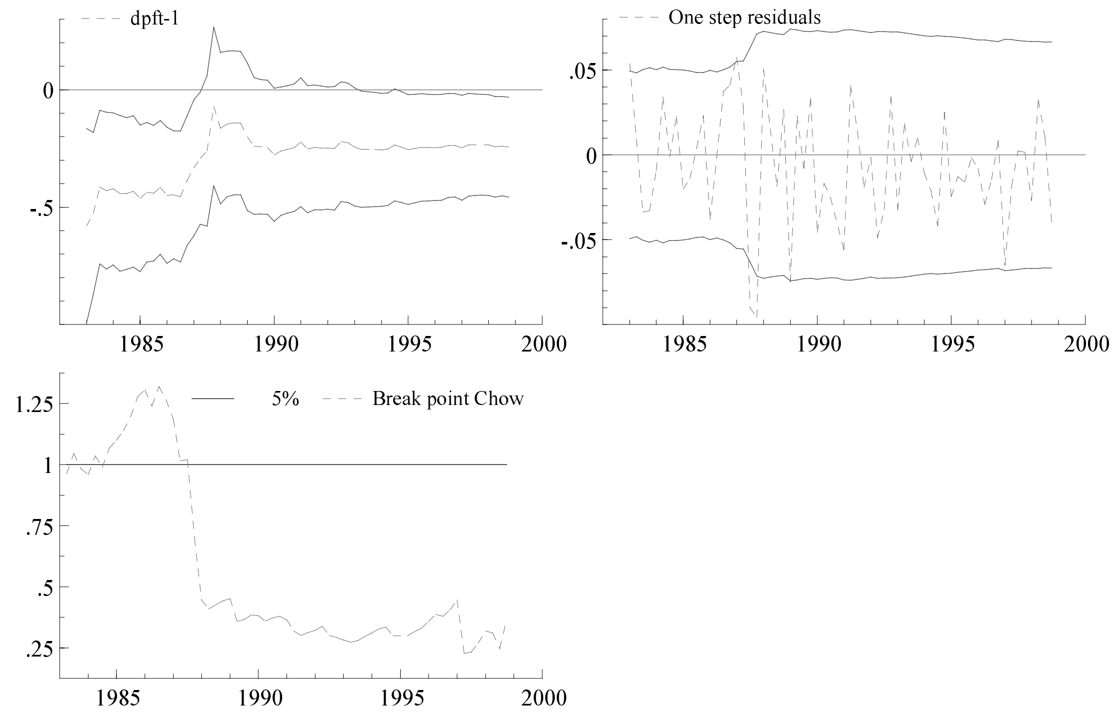
We now focus on the concepts of Hendry (1988), and turn to showing the corresponding empirical evidence of nonconstancy of the marginal process for Δpf_t. Noticeably, no marginal model for marginal costs is considered because that variable enters (18) only at a lag. Starting with a fifth-order autoregressive process for Δpf_t and simplifying, we obtain the following specification (with standard errors in parentheses):

(19) $$\underset{}{\Delta pf_t} = const. - \underset{(0.106)}{0.244}\Delta pf_{t-1} - \underset{(0.009)}{0.021}CS_{2t}$$

Method: *OLS* T=75[1978(3)–1998(4)] R^2=0.118 σ=3.33% *DW*=2.05
$AR_{1\text{-}5}$:$F(5, 74)$=0.654 $ARCH_{1\text{-}4}$:$F(4, 71)$=2.145 *NORM*:$\chi^2(2)$=2.058
HET_1:$F(3, 75)$=2.989 HET_2:$F(4, 74)$=2.255 *RESET*:$F(1, 78)$=0.057.

Figure 4 displays recursive test statistics and recursive estimates for (19), with constancy rejected. For instance, the sequentially estimated equation standard errors increase by about 40 per cent during the sample. That is, a marked worsening of fit, as also detected by the significant break point Chow statistics.

Figure 4. Recursive estimates and test statistics of (19)



Following Hendry (1988), the nonconstancy of the marginal model for Δpf_t together with the constancy of the conditional model for Δpa_t has two important implications in the present situation. First, the Lucas critique cannot apply to (18). This implication holds even if the Δpf_t-equation ignores *some* information relevant to the process generating competing prices. The combination of empirical evidence is thus inconsistent with the hypothesis that Norwegian exporters act on expectations based models. Second, our findings imply that we cannot reject that Δpf_t is weakly exogenous for the short run parameters in (18), parameters of which thereby are consistently estimated by *OLS*.

From an economic perspective, it may be puzzling that Norwegian exporters do not seem to form expectations about future values of competing prices. However, the conditional *EqCM*-model (18) may nevertheless have a forward-looking interpretation, one in which the forward-looking aspects arise from *data-based* predictors rather than model-based predictors of the competing price, cf. Hendry and Ericsson (1991). Such behaviour may be rational because of high costs of information collection and processing. Even so, we do not find significant evidence that (18) in the face of Hendry (1988) is subject to the Lucas critique.[10]

5. Conclusions

The applied literature on export price formation has typically relied on cointegration methods when analysing the determinants of export prices in the long run. On the other hand, some authors utilise data based methods and others apply theory based methods, mainly the *LQAC*-model under rational expectations, when it comes to the modelling of short run behaviour of export prices. One may argue that the latter approach is a possible way to circumvent the Lucas critique if the export price is indeed determined by forward-looking behaviour. However, one may point to the fact that the former approach by means of the *general-to-specific* method has proved successful in some cases, and that the Lucas critique may be overstated.[11] In this paper we have taken a closer look at the empirical performance of both an *LQAC*-model and a conditional *EqCM*-model using data for Norwegian export prices of machinery and metal products.

Our point of departure when analysing export prices has been the standard approach of imperfectly competitive markets. The optimal price is accordingly given as a mark-up over marginal costs. Unlike related studies, we have defined marginal costs as a fully specified cost-function in line with theory. We then derived from cointegration analyses a single long run relationship for the export price with the competing price and marginal costs as its long run determinants. Next, we applied a new, and we believe better, method of testing the *LQAC*-model's ability to explain short run dynamics in the export price of machinery and metal products. The results were however, not very promising as far as the empirical success of the *LQAC*-model is concerned. In fact, our conditional *EqCM*-model is found to be reasonably stable in spite of substantial shocks that hit the Norwegian economy during the eighties and nineties. These findings rule out any expectation based model, not only the restrictive *LQAC*-model, as an explanation of our data.

Our findings may be of some interest from a policy perspective. The knowledge that export prices respond to changes in competing prices and domestic costs is important for modelling and forecasting the effects of changes in the exchange rate. If the true model is the conditional *EqCM*-model, then

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[10] Favero and Hendry (1992) show that Hendry's (1988) approach may lack power for even large instabilities in the marginal processes. Thus, we realise that our conclusions from the recursive tests are subjective. That said, Hendry and Ericsson (1991, p. 855-859) report recursive test statistics and figures that are comparable to Figure 3 and 4, and conclude that the Lucas critique does not apply to the conditional money demand model for UK.

[11] See Ericsson and Irons (1995) and the references therein.

currency policies that increase the competing price will depress inflationary effects on export prices more than what would be predicted by the *LQAC*-model under rational expectations.

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Data Definitions and Sources

PA	Export price index of machinery and metal products expressed in Norwegian currency (1996=1). Source: Statistics Norway.
PF	Import price index of machinery and metal products expressed in Norwegian currency (1996=1). Source: Statistics Norway.
MC	Databased proxy for marginal costs.
CS_{it}	Centred seasonal dummy for quarter i, equals 0.75 in quarter i, −0.25 otherwise, i=1,2,3.
*D*79(1)	Dummy variable used to account for an outlier in the equation for *PA* in the VAR. Equals 1 in 1979(1), zero otherwise.
*D*79(2)	Dummy variable used to account for an outlier in the equation for *PA* in the VAR. Equals −1 in 1979(2), zero otherwise.