Rational Expectations in Price Setting - Tests Based on Norwegian Export Prices

by

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Abstract: This paper uses imperfect competition as a basis for modelling the export price for an aggregated commodity produced by the Norwegian private mainland economy. The long run solution is analysed using a cointegration technique. The dynamics is modelled according to two different approaches; a backward looking error correction model and a forward looking model where rational expectations are assumed. The dynamic structure of the forward looking model is derived from a linear quadratic adjustment cost function under rational expectations, but the empirical results do not support this specification. We can not reject super-exogeneity to be present in the backward looking error correction model. The empirical evidence are thus not consistent with rational expectations.

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1 Introduction

It is well-know in the literature that a traditional error correction model (ECM) may encompass a number of models derived from quite different theories concerning the agents' attitudes towards changing environments and the role for expectations. In addition, the practical use of ECMs has been criticized for being too ad hoc - letting the data decide the number of lags on the forcing variables involved. The error correction representation may be due to a partial adjustment model, in which agents are not at all concerned about the future path of the forcing variables, but want to avoid too rapid changes in their own decision variables due to adjustment costs. On the other hand, the specific dynamic may arise because the agents are influenced by expectations. These expectations can be purely data-based or may have been formed according to an extrapolative scheme, as for instance the adaptive expectations model, but also according to rational expectations. The latter may be the case if the forcing variables, about which expectations are formed, follow autoregressive, or vector autoregressive, data generating processes. The error correction specification will, if it is the reduced form of a rational expectation model, be subject to the Lucas critique, Lucas (1976), which states that the reduced form equation will not exhibit invariance if the processes generating expectation variables change[1]. The ECM may consequently lead to wrong policy recommendations when implemented into a macroeconometric model, if this is the case.

However, whether expectations matter and how they are formed are, and should be, subject to empirical investigations. Results from both direct and indirect tests on the formation of expectations have been presented in the literature, but the question is not yet settled. Previous analysis on microeconomic data on how Norwegian firms form their expectations concerning demand towards the firms and prices on own products on domestic and foreign markets, conclude that expectations are formed according to an extrapolative scheme, see e.g. Svendsen (1993, 1994).

Following the exogeneity literature, see e.g. Engle, Hendry and Richard (1983), the Lucas critique is considered as an example of the absence of super exogeneity. And, as super exogeneity is a testable property of an empirical relation, see Hendry (1988) and Engle and Hendry (1993), one may in certain cases be able to refuse the Lucas critique on empirical grounds.

Two questions attract our attention in this paper. One focus is on the export price of a commodity which is an aggregate of all goods and services

[1] The Lucas critique is a special case of Haavelmo's discussion of autonomous relations, see Haavelmo (1944).

produced in the Norwegian economy excluding commodities from oil production, shipping and government services, i.e. private mainland economy. The analysis is carried out within the framework of imperfect competition and we ask whether, and to which degree, Norwegian producers behave as price setters on foreign markets.

A long-run equilibrium path for export prices is derived assuming Norwegian producers to be price setters as we assume they face a downward sloping demand curve on export markets. The empirical representation for this long-run path is derived within the cointegration framework.

Having found a long-run solution for the export price, our next focus is to model the movements along this path. Embodied in this effort is the choice between forward- and backward looking models, or as one may state it, between a theoretical or ad hoc approach toward the dynamics. A forward looking model, in which the dynamic part is represented by expected changes in future costs and competing prices, is derived from the minimization of a linear quadratic adjustment cost function under rational expectations. The solution of the minimization problem imposes a set of cross-restrictions on both long- and short-run parameters known as the backward-forward restrictions.

The approach of multiperiod quadratic loss functions and rational expectations have been widely applied in various applied economic models as price equations, see e.g. Cuthbertson (1986, 1990) and Price (1992), demand for money, see e.g. Cuthbertson (1988), Muscatelli (1989), Cuthbertson and Taylor (1992), and Engsted and Haldrup (1997), demand for labour, see e.g. Sargent (1978), Nickell (1984), Burgess (1992), Price (1994), Engsted and Haldrup (1994) and firms' inventory behaviour, see e.g. Callen, Hall end Henry (1990).

We test the restrictions in the setting of a VAR-model, following a method proposed in Haldrup and Engsted (1994) which is an extension of Campbell and Shiller's work on net present values models, see Campbell and Shiller (1987). Our results do not give much evidence in favour of the linear quadratic adjustment cost model.

The backward looking model is an ECM. No a priori cross-restrictions on the lag structure are imposed but the one that fits the data best is chosen. Neither do we make any specific assumptions concerning expectations. We do, however, confront the relation with tests for super exogeneity and we cannot reject super exogeneity in relation to structural breaks that have hit the economy during our period of estimation. This is evidence against rational expectations playing a role in the formation of export prices in Norway.

The paper is organized as follows. Our price setting model is presented in section 2, while we take a closer look at the data in section 3. Section

4 includes results from the estimation of a cointegrating vector among the three variables; the export price, unit costs and the competing price, applying Johansen's maximum likelihood method. In section 5 we present the results from testing and estimating the dynamic models, while we test for super exogeneity in section 6.

2 Price-setting under imperfect competition

The theoretical framework in this paper is that of imperfect competition, as we assume Norwegian firms act as price setters in domestic as well as in foreign markets. We also assume the possibility for price discrimination. Previous studies on Norwegian export data support the assumption of imperfect competition, see e.g. Bowitz and Cappelen (1994), Lindquist (1993) and Naug (1994). Consumers are able to distinguish between products delivered by say Norwegian producers and those delivered by others. Several arguments can justify the assumption of differentiated products in trade between countries. There may be country specific differences in quality and/or degree of processing. In empirical economics, we are mainly working with aggregates of products. The composition of these aggregates may differ across countries in such a way that these aggregated commodities can be treated as differentiated products. One may argue for price setting behaviour even when homogenous products are traded on international markets, if the domestic producers have a certain market power through their share of the world market for the actual product. This may be the case even for Norwegian producers of semi-manufactured goods.

The standard assumption of profit maximization under imperfect competition leads to the well-known result that marginal revenue equals marginal costs in all markets. This condition implicitly defines the export price, PA, as a function of the price on competing products (PX), the level of total demand in both markets, the price on variable production factors (PV), the produced quantity, and the stock of capital. Assumptions on the form of the demand and production functions may give us a simplified price function. The first simplifying step is to assume constant return to scale for variable factors. The optimal price, PA, can thus be expressed as a function of costs per unit and a mark-up depending on the structure of demand. A further simplification can be made if we assume the demand function to be derived from a CES (Constant Elasticity of Substitution) utility function. The consumers' choice between the good delivered from Norwegian or foreign producers will thus depend solely on the price ratio between the two of them.

In equation (1) the relation is made linear in the parameters through a logarithmic transformation and the equation describes the price set by Norwegian producers as a function of the price on competing products and variable unit costs. We denote the export price derived from this equation pa^*, to indicate that the equation defines the long-run equilibrium path, or say a long-run target.

$$pa_t^* = \beta_0 + \beta_1 pv_t + \beta_2 px_t + \upsilon_t, \qquad \beta_1 + \beta_2 = 1 \tag{1}$$

where $pa = log(PA)$, $px = log(PX)$ and $pv = log(PV)$. The equation is augmented with a white noise error term, υ_t.

We note that the elasticity of export prices with respect to unit costs is given by β_1 and the elasticity with respect to competing prices by β_2. It follows from our assumptions that the price PA should be homogenous of degree one (static homogeneity) in the competing price (PX) and unit costs (PV) which is implied by the restriction on the sum of the two price elasticities, β_1 and β_2.

It may be shown that $\beta_1 = 1/\sigma$ and $\beta_2 = (\sigma - 1)/\sigma$, where σ is the elasticity of substitution between the product delivered from Norwegian producers and the product delivered from foreign producers. The price, PA, can be described solely as a function of unit costs ($\beta_2 \to 0$) if the elasticity of substitution (σ) approaches one. If σ approaches ∞, the products are identical (homogenous) and we are back to the theory of competitive markets ($\beta_1 \to 0$). Norwegian producers will have no explicit market power in this situation, and consequently PA will approach PX.

3 The data

Estimations are carried out on quarterly seasonal unadjusted data. Variable definitions and data sources are provided in appendix A. We study the export price of a commodity which is an aggregate of all goods and services produced in the Norwegian economy excluding commodities from oil production, shipping and government services, i.e. private mainland economy. The level of aggregation may cover sectors with different strategic position in their respective markets and with different production structure. If the composition of the aggregate has changed during the estimation period, the differences between the sectors may lead to instability in our estimated parameters.

A weighted average of Norwegian import prices serves as a proxy for competing prices (PX). The weight given each price is the weight the respective product has in the Norwegian export basket. PX is measured in Norwegian currency (NOK). Other proxies were considered; among other things a weighted average of import prices in the main foreign markets for Norwegian producers, a proxy that has been used in previous research, see Svendsen (1995). The results were, however, not satisfactory. One reason may be that it is based on a different composition of goods than what found in aggregated export from the Norwegian mainland economy. The weights we use in constructing PX in this paper does reflect this particular composition.

Costs (PV) are represented by labour costs per unit, inclusive of net sector taxes for the sector we are studying. Because this sector includes most of private production activity for the mainland economy, intermediate deliveries from other sectors are mainly imported. These costs are proxied by the index of competing prices (PX). As the composition of Norwegian imports of intermediate goods will differ from the composition of our export, the full effect of costs related to intermediate deliveries will not be captured by our equation.

Figure 1 displays the two ratios PA/PV and PA/PX for the period 1970(1) to 1992(4). Norwegian export prices (PA) have increased less than unit costs (PV) but more than competing prices (PX) over the entire period. The differences are not dramatic. The movements in the ratios throughout the period give a more complex picture. We observe that Norwegian export prices took part in the international price increase that succeeded OPEC I. Exports of energy intensive goods make up a substantial part of our aggregate. Norwegian producers were less affected by increased petroleum prices due to cheap hydroelectric power in Norway. However they were affected as a large part of their intermediates is imported goods. The most dramatic change in the ratio between Norwegian export prices and competing prices, takes place in the 1980s with the Norwegian export prices increasing relative to competing prices from 1983 to 1989. During the same period, we have a slight downward movement in the relation between export prices and unit costs. The data indicate that both variable costs and the price on competing product are important in explaining the Norwegian export price.

4 Time series properties and cointegrating vectors

An econometric model is balanced if it is made up of stationary variables, i.e. I(0)-variables, and/or cointegrating vectors of non-stationary variables. We cannot reject the hypothesis that the variables pa, pv, px are I(1)-variables,

Table 1: Test for cointegration. 1971(2)-1992(4). VAR(3).

	$\mathbf{L}_{\max}$	$\mathbf{L}_{trace}$
r$\leq$ 2	2.22	2.22
r$\leq$ 1	5.36	7.58
r$\leq$ 0	25.16*	32.74*

Notes: $L_{\max}$ and L_{trace} are the maximum eigenvalue and trace test for the number of cointegrating vectors. * indicates significance at a 5% level.

but their first order differentiated are not[2], according to the results from (Augmented) Dickey-Fueller tests. The long-run part will form a stationary element if the level variables cointegrate. The next step is thus to evaluate the cointegrating properties.

We have used tests developed in Johansen (1988) to test the number of cointegrating relations among the I(1)-variables included in the vector (pa,pv,px). The results[3] are reported in table 1. The model is augmented with centered seasonal dummies, a dummy for a structural change in the seasonal pattern at the end of 1977, lagged values of changes in the unemployment rate and dummies for outliers. These variables enter the VAR unrestricted and most of them are needed for the equation for pv in the VAR to pass the normality-test. Critical values are calculated according to Osterwald-Lenum (1992). We reject the hypothesis of zero cointegrating vectors against the alternative of one vector, while we cannot reject that the number of cointegrating vectors are equal to or less than one against the alternative of two vectors. We conclude that there is one cointegrating vector.

The cointegrating vector corresponding to the highest eigenvalue, calculated by use of the Johansen maximum likelihood procedure, is (1, -0.55, -0.49) when normalized on pa, and with the next two components referring to px and pv respectively. Table 2 reports the estimated cointegrating vectors and p-values for the test of long-run homogeneity, weak exogeneity and the combined hypothesis. Each one of the hypotheses are not rejected, but the combined one is.

The estimated cointegrating vectors indicates that unit costs and competing prices both have an impact on export prices.

[2] The following "t-values" were obtained by the DF or ADF tests: pa (-1.94), pv (-2.29), px (-2.12), Δpa (-6.86), Δpv (-9.28) and Δpx (-9.70).

[3] All empirical results are derieved using the PcGive 9.0 and PcFiml 9.0, see Doornik and Hendry (1997a, 1997b).

Table 2: Estimates of the cointegration vectors and tests of restrictions.

Imposed restriction(s)	pa_t	px_t	pv_t	$\boldsymbol{\alpha}'$	χ^2**-test of restriction(s)**^{a}
Unrestricted	1	$\underset{(.09)}{-.55}$	$\underset{(.08)}{-.49}$	$(\underset{(.07)}{-.27}, \underset{(.08)}{-.14}, \underset{(.11)}{.19})$	
Long-run homogeneity	1	$\underset{(.11)}{-.40}$	$-.60$	$(\underset{(.06)}{-.20}, \underset{(.06)}{-.12}, \underset{(.09)}{.12})$	.0876
Weak exogeneity	1	$\underset{(.09)}{-.65}$	$\underset{(.08)}{-.39}$	$(\underset{(.06)}{-.32}, *, *)$	.0563
Long-run homogeneity and weak exogeneity	1	$\underset{(.11)}{-.54}$	$-.46$	$(\underset{(.06)}{-.24}, *, *)$	.0302*

Notes: α' is the vector of standarized loadings. Standard errors are provided in parantheses. [a]Reported figures are p-values.* indicates significance at a 5% level.

5 Dynamic models of price setting behaviour

Equation (1) is often looked upon as a long-run equilibrium path, or a moving target that firms try to catch up with. The short-run movements, or the dynamics, around this target may be modelled in several different ways. We will apply two different strategies proposed in the literature, a backward looking and a forward looking model. The two dynamic models have an error correction term, representing the long-run equilibrium path, in common, but they differ with respect to important features. The difference between the two models lies in our assumptions concerning expectations and the cross-restrictions imposed on the parameters in the forward looking model.

5.1 A backward looking error correction model

The results in this section is based on the estimation of an error correction model (ECM) for export prices for the Norwegian private mainland economy. A general representation of the model is given in equation (2).

$$\Delta pa_t = \gamma_0 + \sum_{r=1}^{R} \gamma_{1r} \Delta pa_{t-r} + \sum_{l=0}^{L} \gamma_{2l} \Delta pv_{t-l} + \sum_{m=0}^{M} \gamma_{3m} \Delta px_{t-m}$$
$$-\tau_0 \left(pa - \beta_1 pv - \beta_2 px\right)_{t-1} + \varpi_t \tag{2}$$

We recognize the long-run solution within parentheses, while the speed of adjustment towards this path is given by the error correction coefficient,

τ_0. γ_{1r}, γ_{2l} and γ_{3m} are parameters in the short-run dynamics, describing the movements around the long-run path. ϖ_t is assumed to be a white noise error term.

We do, somewhat imprecisely, refer to this model as a backward looking ECM but note that its dynamics rely on looser theoretical ground than the dynamics in the forward looking model which we present in the next section. The lagged terms in the dynamics can be defended by the presence of adjustments costs, adaptive expectations or other sorts of extrapolative expectations, but may also be consistent with rational expectations if the expectations terms are generated from an autoregressive (or vector autoregressive) process, see Nickell (1985). Consequently, no specified assumptions are made with regard to these matters as different sets of assumptions may lead to the error correction representation in equation (2).

We start out with a general specification of the short-term dynamics in accordance to the general-to-specific approach advocated by Davidson, Hendry, Srba and Yeo (1978). The short-run dynamics in our preferred equation, are chosen according to what fits the data best as far as the estimated parameters satisfy fundamental restrictions proposed by economic theory and the estimated regression passes different tests for misspecification. The export price equation is required to satisfy the following restrictions; positive long-run parameters ($\beta_1, \beta_2 > 0$) and no rejection of the homogeneity restriction ($\beta_1 + \beta_2 = 1$). Equation (3) is our preferred backward looking ECM[4].

$$\begin{aligned}\Delta pa_t \quad = \quad & \underset{(.04)}{.15} + \underset{(.06)}{.35} \cdot\Delta px_t + \underset{(.07)}{.30} \cdot\Delta px_{t-1} + \underset{(.07)}{.27} \cdot\Delta px_{t-2} + \underset{(.07)}{.13} \cdot\Delta px_{t-4} \\ & + \underset{(.05)}{.20} \cdot\Delta pv_t - \underset{(.05)}{.15}\cdot \Delta pv_{t-2} - \underset{(.06)}{.21} \cdot pa_{t-1} + \underset{-}{.11} \cdot px_{t-1} + \underset{(.03)}{.10} \cdot pv_{t-1} \\ & + \underset{(.01)}{.03} \cdot CSeason_t \cdot p77 + \underset{(.01)}{.05} \cdot CSeason_{t-1} + \underset{(.01)}{.04} \cdot CSeason_{t-2}\end{aligned} \tag{3}$$

$$Static\ long-run\ solution \ : \ pa = .15 + .52 \cdot px + .48 \cdot pv$$

$$Method: OLS, N = 87(71(2) - 92(4)), R^2 = 0.60, \sigma = .0181, RSS = .0246,$$

$$DW = 1.95, ARCH_4 : F(4, 67) = 0.95, HETERO : F(20, 54) = 0.88,$$

$$AR_{1-5} : F(5, 70) = 0.73, NORM : \chi^2(2) = 3.96, RESET : F(1, 74) = 0.13$$

[4]Lagged endogenous variables (one to four lags) have been included in preliminary estimations, but were far from being significant, and these results are not reported

Standard errors are provided in parentheses and significance probabilities in brackets. The equation passes all diagnostic tests for functional form misspecification.[5] and it seems to track the movements in Δpa_t reasonable well according to figure 2. Recursive graphics are shown in figures 3 and 4(a,b) and we note that the recursive coefficients are reasonable stable and that the equation passes the break-point Chow test at a good margin.

Equation (3) is estimated subject to the restriction of static homogeneity ($\beta_1 + \beta_2 = 1$). The restriction is not rejected by the data according to a Wald test with a p-value of 0.192. The elasticity of competing prices (β_2) decreases from 0.62 to 0.52, while the elasticity of (β_1) unit costs increases from 0.41 to 0.48 when the homogeneity restriction is imposed. We cannot reject the hypothesis of dynamic homogeneity[6] (at a probability significance of 0.537). If imposed, the long-run equilibrium solution will depend solely on the long-run level of pv and px, and be unaffected by their long-run growth rate.

The speed of adjustment towards the long-run equilibrium path, given by the error correction coefficient, is slow and the coefficient stays relatively stable between -.30 and -.20 in all regressions undertaken in our search for a parsimoniously one. At all levels it is reduced when long-run homogeneity is imposed. We found the same pattern in the cointegration analysis.

The long-run equilibrium part of the equation and the short-run dynamics are estimated simultaneously by use of OLS. The OLS estimates of the long-run parameters will equal the estimates derived by Johansen's maximum likelihood method if weak exogeneity is imposed in the latter, and the same order of the VAR is assumed. The OLS estimates are however more efficient. The results from section 4 indicate that we may have some problem with the weak exogeneity together with long-run homogeneity, but the long-run effects of unit costs and competing prices (β_1 and β_2) in equation (3) are very close to the unrestricted cointegrating vector and the vector estimated subject to long-run homogeneity and weak exogeneity.

While the three variables, pa, pv and px, were treated as endogenous in the VAR-model in section 4, we condition on the contemporaneous variables Δpv_t and Δpx_t in equation (2) which is a conditional econometric model. The single-equation approach behind the estimation of a conditional econometric model, leads to valid inference only if we cannot reject the hypothesis that the

[5] AR_{1-5} is the Lagrange Multiplier test for residual serial correlation, see Harvey (1990). $ARCH_4$ is a test for heteroscedasticity, see Engle (1982). $NORM$ is a test for the normality of the residuals, see Doornik and Hansen (1994). Next follows White's test for heteroscedasticity, see White (1980) and the $RESET$-test for functional form misspecification, see Ramsey (1969).

[6] Dynamic homogeneity is defined as $\sum_{r=1}^{R} \gamma_{1r} + \sum_{l=0}^{L} \gamma_{2l} + \sum_{m=0}^{M} \gamma_{3m} = 1$

conditioning variables are weakly exogenous for the parameters of interest, see Engle et al. (1983). If weak exogeneity is rejected, useful information is lost and the estimated coefficients may not be independent of changes in the process generating the conditioning variables. We cannot reject that both Δpv_t and Δpx_t are weakly exogenous for the parameters of interest in equation (3) according to the Wu-Hausman test[7], Wu (1973). The p-value of the Lagrange Multiplier test for omitted variables is 0.244 (its observed F-statistics equals the Wu-Hausman statistics). Following the proposals in Urbain (1992) we have also tested whether the marginal processes for Δpv_t and Δpx_t display error correcting behaviour by testing a zero restriction on the error correction term from equation (3) in the two marginal processes. The p-values of the Wald-tests are .055 (Δpx_t), .298 (Δpv_t) and .117 (Δpx_t and Δpv_t). We interpret these results as evidence for Δpv_t and Δpx_t being weakly exogenous for the long-run parameters.

The failure of not assuming rational expectations, when agents' actions partly do rely on this kind of expectations can lead to biased estimates on the included parameters. Let us look at an example: The costs of production (PV) increase, for instance due to increased wages in the domestic labour market. The isolated reaction for the individual firm under imperfect competition is to raise its selling price in all markets together with facing reduced revenues. For the entire economy the wage increase may lead to a reduction in the trade surplus with a subsequent depreciation/devaluation of the domestic currency. The depreciation/devaluation gives room for an increase in the export price measured in domestic currency (PA).

Firms, that behave according to the rational expectations hypothesis and correctly foresee the reduced value of the currency, will begin to raise their export prices immediately after costs have increased, both due to increased costs and to the expected change in the exchange rate. The econometrician, who have only included current and lagged values of unit costs and competing prices, will overestimate the effect of costs in the export price relation, and thus the firms' market power. In order to get unbiased estimates, one has to include leading variables as well.

5.2 Empirical tests of the LQAC-model

The forward looking model we focus one, is the linear quadratic adjustment cost (LQAC) model. The model is derived from a multiperiod quadratic loss

[7] This is a test for independence between the residual and the conditioning variables. Reduced form equations are estimated for Δpv_t and Δpx_t. Next, we test the significance of the residuals from these two equations in our preferred equation for Δpa_t estimated by use of ordinary least squares (OLS).

function (Q), which imposes quite strict restrictions on the dynamics.

$$Q = E_t \left[\sum_{s=0}^{\infty} \delta^s \left(\mu \left(pa_{t+s} - pa^*_{t+s} \right)^2 + \left(pa_{t+s} - pa_{t+s-1} \right)^2 \right) |\Omega_t \right] \quad (4)$$

According to the loss function, firms may suffer a loss both from being off the long-run equilibrium path, pa^*_{t+s}, and from changing the actual price, for instance due to loss of reputation towards customers. The actual price, pa_{t+s}, is a result of minimizing the total loss. μ represents the weight given to the discrepancy from the long-run path relative to the costs related to changing prices. δ is the discount factor, defined over the range $\langle 0, 1 \rangle$. We assume rational expectations, so that agents' beliefs concerning future prices and costs can be expressed as the mathematical expectation of the variables conditional upon information available at time t, Ω_t. E_t is the expectations operator.

It turns out that we may parameterize the solution to the optimization problem[8] as an error correction model with forward looking expectations. This is a relevant representation of the solution when dealing with I(1)-variables that cointegrate.

$$\triangle pa_t = (1-\lambda)\beta_0 + (1-\lambda) \sum_{s=0}^{\infty} (\lambda\delta)^s \left[\beta_1 \triangle pv^e_{t+s} + \beta_2 \triangle px^e_{t+s} \right]$$

$$- (1-\lambda) \left[pa - \beta_1 pv - \beta_2 px \right]_{t-1} + (1-\lambda)(1-\lambda\delta) \upsilon_t \quad (5)$$

We have replaced pa^*_t by our assumed relation for the long-run equilibrium path (equation (1)) and sat $pv^e_{t+s} = E_t \left(pv_{t+s} |\Omega_t \right)$ and $px^e_{t+s} = E_t \left(px_{t+s} |\Omega_t \right)$. λ is the stable root in the difference equation calculated from the first order condition to the above minimization problem[9] and $(1-\lambda)$ appear as a parallel to the adjustment coefficient in the backward looking ECM. A value on λ close to one corresponds to firms being relatively more concerned about the loss from changing prices than from being off the long-run path. The adjustment process is slow when this is the case and the influence from expected changes in the forcing variables is increased. $\upsilon_t = 0$ correspond to the case of a non-stochastic long-run price target and the model in (5) becomes the exact linear rational expectations model.

[8] The solution to the optimisation problem is shown, among others, in Sargent (1978) and Nickell (1985).

[9] One may show that $\lambda\mu = (1-\lambda)(1-\lambda\delta)$.

Geometrically declining weights, that are related to the backward looking parameter (λ), are imposed on the lead structure in equation (5). Different strategies, involving both single-equation methods and VAR-modelling, have been proposed in the literature in order to obtain estimates on the parameters and to test the restrictions that are known as the backward-forward restrictions.

In this paper we follow the method proposed in Engsted and Haldrup (1994) and test the LQAC-model in the setting of a VAR-model. The method, which is an extension of Campbell and Shiller's work on net present value models, see Campbell and Shiller (1987), is based on prior estimates of the parameters. We will briefly resume the method.

Manipulating equation (5) and using the lag operator, L, takes us to

$$(1-\lambda L)\left[pa_t-\beta_1 pv_t-\beta_2 px_t\right]+\lambda\left[\beta_1 \triangle pv_t+\beta_2 \triangle px_t\right]-(1-\lambda)\beta_0=$$

$$(1-\lambda)\sum_{s=1}^{\infty}(\lambda\delta)^s\left[\beta_1 \triangle pv^e_{t+s}+\beta_2 \triangle px^e_{t+s}\right]+(1-\lambda)(1-\lambda\delta)\upsilon_t \tag{6}$$

We define a variable S_t, called "the spread", defined as the left hand side of equation (6), with the coefficients replaced by their estimates, $\hat{\lambda}$, $\hat{\beta}_1$ and $\hat{\beta}_2$, retrieved somewhere else. If the LQAC-model under rational expectations is true this variable will be an optimal predictor of the present value of the future changes in the forcing variables. Engsted and Haldrup then show that this implication can be tested within the following VAR.

$$\begin{pmatrix}\Delta pv_t\\ \Delta px_t\\ S_t\end{pmatrix}=\sum_{j=1}^{J}C_j\begin{pmatrix}\Delta pv_{t-j}\\ \Delta px_{t-j}\\ S_{t-j}\end{pmatrix}+e_t \tag{7}$$

Testable cross-restrictions on the parameters in the case of an exact linear rational expectation model, i.e. $\upsilon_t=0$, are derived in Engsted and Haldrup (1994).

We may also use an orthogonality test; multiply equation (6) by $\left(1-(\lambda\delta)^{-1}L\right)$ and do some manipulations:

$$\lambda\delta S_t - S_{t-1} + (1-\lambda)\lambda\delta\left(\beta_1 \Delta pv_t + \beta_2 \Delta px_t\right) =$$

$$(1-\lambda)\sum_{j=0}^{\infty}(\lambda\delta)^{j+1}\Delta E_t\left(\beta_1 \Delta pv_{t+j} + \beta_2 \Delta px_{t+j}\right) + (1-\lambda)(1-\lambda\delta)\left(\lambda\delta\upsilon_t - \upsilon_{t-1}\right) \tag{8}$$

As the right-hand side of the equation is pure innovations under the null of rational expectations in addition to a MA(1)-error, we may consequently test whether it is orthogonal to the agents information set, Ω_{t-1} if $\upsilon_t = 0$, and Ω_{t-2} if $\upsilon_t \neq 0$.

Estimates of the parameters[10], β_1, β_2, and λ are needed to construct the variable S_t. We also need an estimate of δ in order to conduct a likelihood-ratio (LR) test of the cross-restrictions. $\hat{\beta}_1$ and $\hat{\beta}_2$ are taken from the cointegrating vector with long-run homogeneity imposed, i.e. $(\hat{\beta}_1, \hat{\beta}_2) = (0.6, 0.4)$ (see table 2).

Following Engsted and Haldrup (1994) we may use the error correction coefficient from an ECM for Δpa_t as an estimate of $(1-\lambda)$ if Δpv_t and Δpx_t are weakly exogenous with respect to the long-run parameters for Δpa_t and Δpa_t does not Granger-cause Δpv_t and Δpx_t. These conditions are not met by our data. What we do, is to calculate S_t for three different values[11] of λ, in order to test the cross-restrictions and to conduct the orthogonality test on a broader range of possibilities. If we reject the cross-restrictions for all different values of λ, we accept this as evidence against the combined framework of rational expectations and LQAC-models.

Three different estimates of S_t were derived with $\lambda = (0.7, 0.8, 0.9)$. For all three cases, a VAR(4)-model was the preferred one while estimating equation (7). The LR-test of the cross-restrictions were conducted for four different values of the discount factor, $\delta = (0.7, 0.8, 0.9, 0.99)$. The LR-statistics, which follow a $\chi^2(12)$distribution, are presented in table 3. As indicated by (**), none of our results are significant at a 0.01 significance level. In fact, for all our combinations of (λ, δ) the probability values for the restrictions to be valid, are $p = 0.000$.

The next question is whether the orthogonality proposition holds. The term on the right-hand side of equation (8) has been computed for our chosen

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[10] The estimate on $\hat{\beta}_0$ does not influence the test results. We have, however, set $\hat{\beta}_0 = \frac{1}{T}\sum_{t=1}^{T} pa_t - \hat{\beta}_1 pv_t - \hat{\beta}_2 px_t = 0.58$.

[11] The values form an interval that covers different estimates on the adjustment coefficient indicated by the cointegration analysis, the ECM and attempts (neither successful, nor reported) to estimate the Euler-equation.

Table 3: Likelihood Ratio test of the cross-restrictions.

$\boldsymbol{\lambda} \backslash \delta$	**.99**	**.90**	**.80**	**.70**
.9	79.1**	88.9**	102.4**	119.5**
.8	85.1**	95.9**	110.6**	129.0**
.7	92.9**	105.0**	121.4**	141.5**

Table 4: The ortogonality test with information dated t-2,t-3. p-values.

$\boldsymbol{\lambda} \backslash \delta$	**.99**	**.90**	**.80**	**.70**
.9	.040*	.032*	.023*	.016*
.8	.035*	.026*	.017*	.010**
.7	.025*	.016*	.009**	.004**

values on (λ, δ), and then regressed on two different information sets[12], the first dated t-1,...,t-3 and the second t-2, t-3. The first one is relevant for our reference model which is the exact rational expectation model (i.e. $\upsilon_t = 0$). The second takes into account that υ_t may be a stochastic white noise error-term (i.e. $\upsilon_t \neq 0$) for instance because the econometricians have a smaller information set than the firms.

Testing on the information set that includes t-1, we find the hypothesis to be rejected with a p-value of 0.000 for all combinations of (λ, δ) considered in this paper. If, however, $\upsilon_t \neq 0$, these results may be a result of correlation between the MA(1)-error in equation (8) and information dated t-1. We report p values for the $F(r_1, r_2)$-statistic used to test the hypothesis of a joint zero-restriction of the entire information-set in table 4 when information dated t-1 is excluded from the set. According to the results, we can still reject the orthogonality property to be valid for all combinations of (δ, λ), but at higher p-values. We conclude that there are information that, if more efficiently used, could have reduced the need for changing the expectations of the future path of the forcing variables, also when we control for the possibility that υ_t is a white noise error-term.

Campbell and Shiller (1987) propose an informal way to evaluate the net present value models. The idea is to evaluate whether an eventual rejection of the model is due to fundamental deviations from the model or solely to economically unimportant factors as a white noise error-term. From equation

[12] The information sets include $\Delta pa_{t-j}, \Delta px_{t-j}, \Delta pv_{t-j}$,(j=(1),2,3). Obviously, these variables are rather strong candidates to be included in a forward-looking firm's informationset. In addition, seasonal dummies are included.

(6) we note that S_t equals the present value of future changes in the forcing variables and an error-term. The part of the right-hand side which represent the present value can be solved for the future forcing variables conditional on the information set Ω_t, resulting in an unconditional VAR-forecast[13] of the present value of future changes in the forcing variables. We call this variable, $S_t^* = E(S_t \,|\Omega_t)$. The discrepancy between S_t and S_t^* under the null of rational expectations should be equal to the error term ,$(1-\lambda)(1-\lambda\delta)\upsilon_t$. A graphic presentation of S_t and S_t^* (see figure 5) calculated for $(\lambda = 0.9, \delta = 0.99)$, indicate a substantial difference between the two of them, leading to the informal conclusion that the rejection of the cross-restrictions are due to fundamental deviations from the model.

6 Test for super exogeneity

We have not gained sufficient support for the combined framework of rational expectations and the LQAC-model. The refusal may be due to the expectations or the dynamic structure, or both of them, being wrongly specified. Our preferred ECM is specified solely in backward (or contemporaneous) terms. It may nevertheless be the reduced form of an rational expectation model, if the forcing variables may be described in terms of autoregressive processes. If this in fact is the situation, our preferred model will be subject to the Lucas critique, see Lucas (1976), as a change in the processes governing the forcing variables will lead to a structural change in our model describing the movements in Norwegian export prices.

The relevance of the Lucas critique, is a special case of the failure of a model to meet the requirements for super exogeneity, see Hendry (1988), requirements which should be met if the model is to be used in policy analysis. In our model, Δpv_t and Δpx_t are super exogenous for our parameters of interest if (i) they are weakly exogenous and (ii) changes in their marginal processes do not introduce changes in the conditional model describing Δpa_t, i.e. equation (3). If super exogeneity is supported by the data, the Lucas critique looses its relevance and so does the proposal that the model is a reduced rational expectation model.

The evidence in favour of Δpx_t being weakly exogenous are not unambiguous (as we saw in section 4), but if we accept it we may continue to test the second part of the requirements. In appendix B, marginal models for Δpv_t and Δpx_t are reported (see equations (9) and (11)), while recursive graphics

[13] The VAR-forecast is based on our estimates of λ and δ, and the coefficient matrixe from estimating the VAR-model in equation (7). For further details, see Engsted and Haldrup (1994).

Table 5: Test for superexogenity

Marginal model for:	**Variables**	**Results**[a]
Δpv_t	$p922, p7879 \cdot u_{t-1}$	$F(2,73) = 0.15\,[0.86]$
Δpx_t	$p7475, p811, p911$	$F(3,72) = 0.35\,[0.79]$

Notes: [a]We test H_0 :Reported variables are wrongly omitted in equation (3). p-values given in brackets.

are reported in figures 6 and 7. For Δpx_t, both Chow tests , the recursive residual sum of squares and the 1-step residuals clearly indicate instability around 1990. The model for Δpv_t is only just about to pass the Break-point Chow test in 1991-92, while the three other plots indicate another period of instability around 1978-79. Both marginal models are expanded with variables (including dummies) picking up the instabilities (see equations (10) and (12)) and according to figures 8 and 9, the instabilities are clearly reduced.

The marginal model for Δpv_t becomes more stable when adding a dummy capturing the wage regulation in 1978-79 ($p7879 \cdot u_{t-1}$) and a second dummy for an outlier ($p922$). The marginal model for Δpx_t is augmented with two dummies that may be due to OPEC I and II ($p7475$ and $p811$) and an outlier in the early 1990ties ($p911$).

The next step is to test whether these variables add any explanatory power to our conditional model. The results from these tests are reported in table 5.

None of the two variable-sets are significant in our conditional model. Neither are the variables individually significant when added to equation 3, according to theirs t-values. Of special interest is the dummy ($p7879$) picking up the effects of the wage regulation in 1978 and 1979 and the two dummies reflecting the oil-price shocks ($p7475$ and $p811$). The whole effects of these occurrences are fully represented by the presence of Δpv_t and Δpx_t in our equation, which leads us to the conclusion that the agents are concerned about the actual and previous changes in costs and competing prices, or if they behave according to expectations, these are formed according to some sort of extrapolative expectation models (for instance adaptive expectations).

7 Conclusions

We have studied the export price for an aggregated commodity produced by the Norwegian private mainland economy. A long-run solution is derived from cointegration analysis lending support to the assumption of imperfect competition. Two different strategies have been tried out in modelling the short-run dynamics. The first one, an ECM in which no explicit assumptions are made with regard to expectations, seems to fit the data reasonable well. The other one, a LQAC-model under rational expectations, implies cross-restrictions on the parameters that our data do not support. The restrictions are tested within the framework of a VAR-model. We cannot reject that our ECM possess super exogeneity in relation to structural breaks that have hit the economy during our period of estimation. These findings provide no warranty for our model to be invariant if faced with future interventions. But, they are evidence against rational expectations playing a role in the formation of export prices in Norway.

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APPENDIX

A Variable definitions and data sources

PA: Export price index of the commodity produced by private mainland economy

PX: Price index in competing markets

PV: Variable unit costs inclusive of net sector taxes for private mainland economy

U: The unemployment rate.

CPI: The Consumer price index

$CSeason$: Centered seasonal equal 0.75 in first quarter, -0.25 else

$p77$: Dummy equal 1 for t≤1977(4), 0 else (dummy for change in seasonal pattern from 1978)

$p922$: Dummy equal 1 for t=1992(2), 0 else

$p7879$: Dummy for wage regulation law, equal 1 for t=78(3), 78(4), 79(2)-80(1), 0 else

$p7475$: Dummy equal 1 for t=1974(4), -1 for t=1975(4), 0 else

$p811$: Dummy equal 1 for t=1981(1), 0 else

$p911$: Dummy equal 1 for t=1991(1), 0 else.

$pa = ln(PA)$, $px = ln(PX)$, $pv = ln(PV)$, $u = ln(U)$, $cpi = ln(CPI)$

Data are taken from the Quarterly National Account, published by Statistics Norway. The proxy for competing prices (PX) is calculated as a weighted average of disaggregated Norwegian import prices (in NOK). The weight given each importprice equals the weight the respective product has in the Norwegian export basket.

All prices are given in Norwegian currency (NOK).

B Marginal models

Variable definitions are given in Appendix A.

B.1 Marginal model for Δpv_t:

$$\begin{aligned}\Delta pv_t \;=\; & -\underset{(.12)}{.43} - \underset{(.09)}{.25}\,\Delta pv_{t-1} - \underset{(.09)}{.18}\,\Delta pv_{t-2} - \underset{(.09)}{.36}\,pv_{t-1} + \underset{(.08)}{.34}\,cpi_{t-1} - \underset{(.01)}{.06}\,u_{t-1} \\ & - \underset{(.02)}{.06}\cdot CSeason_t - \underset{(.02)}{.07}\cdot CSeason_{t-1} - \underset{(.01)}{.19}\cdot CSeason_{t-2} + \underset{(.01)}{.07}\,CSeason_{t-2}\cdot p77\end{aligned} \tag{9}$$

$Method: OLS, N = 87(71(2) - 92(4)), R^2 = 0.92, \sigma = .0270, RSS = .0561,$

$DW = 1.98, AR_{1-5}: F(5,72) = 0.26, ARCH_4: F(4,69) = 3.29^*,$

$NORM: \chi^2(2) = 1.21, HETERO: F(15,61) = 0.83,$

$FORM: (47,29) = 0.85, RESET: F(1,77) = 13.40^{**}$

B.2 Stabilized marginal model for Δpv_t:

$$\begin{aligned}\Delta pv_t \;=\; & \underset{(.11)}{-.27} - \underset{(.09)}{.40}\,\Delta pv_{t-1} \underset{(.08)}{-.28}\,\Delta pv_{t-2} - \underset{(.07)}{.25}\,pv_{t-1} + \underset{(.07)}{.24}\,cpi_{t-1} - \underset{(.01)}{.05}\,u_{t-1} \\ & - \underset{(.02)}{.05}\cdot CSeason_t - \underset{(.02)}{.06}\cdot CSeason_{t-1} - \underset{(.01)}{.19}\cdot CSeason_{t-2} \\ & + \underset{(.01)}{.07}\,CSeason_{t-2}\cdot p77 - \underset{(.02)}{.06}\,p922 - \underset{(.02)}{.08}\,p7879\cdot u_{t-1}\end{aligned} \tag{10}$$

$Method: OLS, N = 87(71(2) - 92(4)), R^2 = 0.94, \sigma = .0228, RSS = .0391,$

$DW = 1.87, AR_{1-5}: F(5,70) = .14, ARCH_4: F(4,67) = .91,$

$NORM: \chi^2(2) = 2.12, HETERO: F(18,56) = .48,$

$RESET: F(1,74) = 4.61^*$

B.3 Marginal model for Δpx_t:

$$\begin{aligned}\Delta px_t \quad = \quad & \underset{(.004)}{.01} + \underset{(.11)}{.42}\ \Delta pa_{t-1} - \underset{(.13)}{.20}\ \Delta pa_{t-4} + \underset{(.11)}{.05}\ \Delta px_{t-4} \\ & + \underset{(.01)}{.02}\ CSeason_{t-2} - \underset{(.02)}{.04}\ CSeason_{t-2} \cdot p77\end{aligned} \tag{11}$$

$Method : OLS, N = 87(71(2) - 92(4)), R^2 = 0.21, \sigma = .0286, RSS = .0664,$

$DW = 2.38, AR_{1-5} : F(5, 76) = 1.10, ARCH_4 : F(4, 73) = 0.94,$

$NORM : \chi^2(2) = 24.66^{**}, HETERO : F(9, 71) = 1.79,$

$FORM : F(18, 62) = 1.28, RESET : F(1, 80) = 1.07$

B.4 Stabilized marginal model for Δpx_t:

$$\begin{aligned}\Delta px_t \quad = \quad & \underset{(.003)}{.01} + \underset{(.09)}{.65}\ \Delta pa_{t-1} - \underset{(.10)}{.21}\ \Delta pa_{t-4} + \underset{(.08)}{.08}\ \Delta px_{t-4} + \underset{(.01)}{.02}\ CSeason_{t-2} \\ & - \underset{(.01)}{.04}\ CSeason_{t-2} \cdot p77 + \underset{(.02)}{.09}\ p7475 + \underset{(.02)}{.06}\ p811 - \underset{(.02)}{.11}\ p911\end{aligned} \tag{12}$$

$Method : OLS, N = 87(71(2) - 92(4)), R^2 = 0.55, \sigma = .0220, RSS = .0377,$

$DW = 2.21, AR_{1-5} : F(5, 73) = 1.25, ARCH_4 : F(4, 70) = 1.30,$

$NORM : \chi^2(2) = 7.28^{*}, HETERO : F(13, 64) = 0.98,$

$FORM : F(22, 55) = 1.37, RESET : F(1, 77) = 0.17$