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Abstract: The paper analyses the price on domestic market for an aggregate commodity produced by Norwegian private mainland economy. The long-run solution is modelled assuming imperfect competition. The elasticities with respect to unit labour costs and competing prices vary with an indicator for competitive strength in domestic market. I consider two models for the dynamic part of the equation. Model A is a conditional ECM in current and lagged variables. Model B is derived from a multiperiod quadratic loss function which introduces rational expectations to the model. The backward-forward restrictions are not rejected. The estimated elasticities for both models are in line with the previous empirical results for the Norwegian economy. Model A is preferred to Model B, partly on the basis of informal encompassing results.

Keywords: Domestic prices, Imperfect competition, Time-varying elasticities, Multiperiod loss function, Rational expectations, Error correction models

JEL classification: C22, D43, D84,

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"A dynamic model obtained directly from optimization exercises may have "desirable theoretical features" (...) but if it does not fully capture the properties of the data or perform better than competing models it is ultimately destined to fall into the graveyard of empirical models (p. 372)." Muscatelli, V. A. (1989): A Comparison of the Rational Expectations and General-to-specific Approaches to Modelling the Demand for M1, Oxford Bulletin of Economics and Statistics 51, 353-375.

1. Introduction

In this paper I confront two different models which aim to explain the domestic price on an aggregated commodity produced by the Norwegian private mainland economy. Both models rely on the theory of imperfect competition in modelling the long-run equilibrium solutions, but they differ in how I model their dynamic parts, leading to two different regression models. The first model is in the tradition of the general-to-specific approach (Davidson et al. (1978)) and is a conditional error-correction model (ECM) with the dynamic part represented by current and lagged variables. The dynamic specification is somewhat *ad hoc* as it does not rely on a clearly specified theoretical framework. For one thing, the ECM may encompass a set of models derived from quite different assumptions concerning expectational hypotheses. If expectations actually are formed according to the rational expectations hypothesis, the ECM will not exhibit invariance if the processes generating expectations variables change. The model is thus subject to the Lucas critique. As an alternative, in the second model, I derive the dynamic part from a strict theoretical framework, namely the multiperiod quadratic loss function (Sargent (1978)). The resulting equation is a linear rational expectations model, with theoretically derived overidentifying restrictions on the parameters. These restrictions are tested for empirically. Other applied works combining a multiperiod quadratic loss function and rational expectations, are Callen et al. (1990), Cuthbertson (1986, 1988, 1990), Cuthbertson and Taylor (1992), Muscatelli (1989), Nickell (1984) and Price (1992, 1994). Of these, only Price (1992) studies prices set in domestic market (UK) and he finds support for the hypothesis that prices are set by agents who form rational expectations.

The two models do not differ much in many respects as empirical results are concerned, such as estimated long-run parameters and their significance, standard errors of regression during the estimation period and post-sample forecast abilities. The score of the conditional ECM (Model A) is somewhat better than for the forward looking model (Model B). I also find that the lead coefficients, when freely estimated, are insignificant. A joint zero restriction on these coefficients is not rejected. I conclude that the empirical evidence, presented in this paper, does not give much support to the combination of the REH and the framework of a multiperiod quadratic loss function, in explaining domestic prices. It is worth noting that, independent of whether rational expectations are assumed or not, our long-run elasticities are in line with other resuls on the price determination in Norway (see Aukrust (1977) and Bowitz and Cappelen (1994)).

The theoretical model which describes the long-run solution is given in section 2, while the two alternative dynamic specifications of the model are given in section 3. Section 4 deals with the time series properties of the data and the possibilities for one or more cointegrating vectors among the variables in the suggested long-run solution. The two next sections, section 5 and section 6 contain the main empirical results for the two models. I discuss the findings and draw some conclusions in the final section 7.

2. The long-run equilibrium solution

I study the price on a commodity which is the aggregate of goods and services produced by the Norwegian private mainland economy. The long-run solution for domestic prices is derived within the framework of imperfect competition. The commodity may be sold at domestic or foreign markets and I allow for price discrimination. In the domestic market, domestic producers are faced with the competition from foreign producers of *almost* the same commodity. I regard the commodity delivered from domestic producers and the imported commodity as two separate commodities, and assume that customers are able to distinguish between them. Different arguments may be put forward to support the assumption of heterogeneous products. First of all, the composition of the two aggregates may not be the same. This argument is of particular relevance at my level of aggregation. Second, products from different countries may differ with regard to quality and degree of processing.

Even if our results indicate a certain degree of price setting behaviour, some firms producing for the domestic market may still act as price takers. Our commodity is an aggregate including a broad range of products as for instance machinery, services and consumer goods. It should be obvious that the structures of these markets are quite unlike.

Domestic producers face a downward sloping demand curve on the domestic market and may act as price setters. The demand of our commodity is a function of the total level of demand, Y, and the price ratio PK/BH, where PK is the price on the imported commodity and BH is the price on the domestic produced commodity. Regarding the structure of production, I assume constant returns to scale for variable factors and let PV denote variable unit costs. From the first order condition for profit maximization, we may now derive the following expression for the optimal price set by domestic producers.

(1) $$BH_t = g(PK_t/BH_t, Y_t)PV_t$$

$g(BH_t/PK_t, Y_t)$ is the mark-up and depends on the structure of demand. In appendix 1, I assume that the consumers' choice between the commodity delivered from domestic versus foreign producers depend on the price ratio, PK/BH, and are independent of the level of demand. I allow the elasticities to vary with the importshare, m_t, across time, in the reduced form equation for BH which is derived from (1). The given importshare may serve as an indicator of domestic producers' competitive strength in own markets relative to the strength of foreign producers.

Variable unit costs are a function of unit labour costs (PW) and costs on intermediate products per unit. If we leave out intermediate deliveries *inside* private mainland economy, most intermediate goods in the production is imported. The price on imported, competing products, PK, is therefore used as a proxy for the costs on intermediate products. Norwegian economy is near self-sufficient with electricity and changes in the electricity price is consequently not captured by PK. I therefore include the price on electricity (PE) as a cost of production. A change in the use of electricity per unit of production over the estimation period, may lead to instability in the estimated parameters as long as the chosen proxy differs from electricity costs per unit.

I arrive at the following long-run solution for domestic prices, denoted bh*, where bh=log(BH), pk=log(PK), pw=log(PW) and pe=log(PE). The model is derived step-by-step in appendix 1. Here, the model is expanded with a white noise error term, u_t.

(2) $$bh_t^* = \beta_0 + \beta_1 pw_t + \beta_2 pk_t + \beta_3 pe_t + \beta_4 z_t + u_t, \qquad \beta_2 = 1 - \beta_1 - \beta_3$$

Equation (2) describes the long run equilibrium path for domestic prices as a function of competing prices, unit labour costs, the electricity price and the importshare. z_t is defined as $[\log(PW_t/PK_t)m_t]$. If competition increases with m_t, I expect $\beta_4 < 0$, implying that increased competition between foreign and domestic producers, increases the importance of competing prices in domestic producers' price setting rule.

The long-run elasticities of domestic prices with regard to competing prices and unit labour costs are time dependent through the variable m_t, while the elasticity with regard to electricity prices is assumed constant over time.

(3) $$El_{PW}BH_t = \beta_1 + \beta_4 m_t, \quad El_{PE}BH = \beta_3, \quad El_{PK}BH_t = \beta_2 - \beta_4 m_t = 1 - \beta_1 - \beta_3 - \beta_4 m_t$$

The theoretical framework, derived in appendix 1, imposes the restriction of homogeneity of degree one on the long-run elasticities, as the sum of the three elasticities is restricted to equal one. This is satisfied through the linear restriction on the parameters in equation (2); $\beta_2 = 1 - \beta_1 - \beta_3$.

3. Dynamic specification

Equation (2) describes a long-run equilibrium solution for domestic prices. The actual price, at a given point of time, may very well be off this path. Faced with changes in competing prices and unit costs, producers may not find it desirable or feasible to adjust to the new equilibrium price immediately. In economic literature one may find different kinds of arguments (see for instance Andersen (1994)), as "menu costs" and "staggered contracts", which impose a loss from changing the price from one period to another. This loss is weighted against the loss of being off the long-run equilibrium path by the quadratic loss function. The loss may be related to one single period, but as today's decisions will influence the discrepancies from the equilibrium path for tomorrow, the loss function is often formulated in a multiperiod framework as in equation (4).

$$(4) \quad Q_{t-1} = E_{t-1}\left\langle \sum_{s=0}^{\infty} \delta^s \left(\mu\left(bh_{t+s} - bh^*_{t+s}\right)^2 + \left(bh_{t+s} - bh_{t+s-1}\right)^2\right) \middle| \Omega_{t-1} \right\rangle$$

The discounted loss for all future periods seen from period t-1, Q_{t-1}, is a function of the expected long-run equilibrium path and the expected path for actual prices. E_{t-1} is the expectation operator, while Ω_{t-1} is the information set including at least all free, available information at time t-1. δ is the discount rate and μ is the relative weight attached to the discrepancy from the long-run equilibrium path. The loss function is derived under the assumption of one representative agent. If, in our aggregate model, the structural parameters, μ and δ, vary across the agents, our estimates of these may be unstable.

The price, bh_t, is set in order to minimize the quadratic loss function. The solution to this problem is derived by use of the forward convolution method (Sargent (1987)). The resulting equation is a linear rational expectations model. I have replaced the expected long-run equilibrium price, bh*, with the assumed model for this price (see equation (2)), and I rearrange in order to get a first order difference of domestic prices as the dependent variable. The symbol Δ, denotes a differentiated variable; $\Delta bh_t = bh_t - bh_{t-1}$, etc, while the superscript "e" denotes expectations: $\Delta bi^e_{t+s} = E_{t-1}(\Delta bi_{t+s} \mid \Omega_{t-1})$, etc.

$$(5) \quad \begin{aligned} \Delta bh_t &= -(1-\lambda)[bh_{t-1} - \beta_1 pw_{t-1} - \beta_2 pk_{t-1} - \beta_3 pe_{t-1} - \beta_4 z_{t-1}] \\ &+ (1-\lambda)\sum_{s=0}^{\infty}(\lambda\delta)^s[\beta_1 \Delta pw^e_{t+s} + \beta_2 \Delta pk^e_{t+s} + \beta_3 \Delta pe^e_{t+s} + \beta_4 \Delta z^e_{t+s}] \\ &+ (1-\lambda)(1-\lambda\delta)\sum_{s=0}^{\infty}(\lambda\delta)^s u_{t+s} \end{aligned}$$

λ is the stable root in the difference equation calculated from the first order condition to the minimization problem. When δ and λ are estimated, we can derive an estimate of μ from $\mu=(1-\lambda)(1-\delta\lambda)/\lambda$. The parameter μ decreases with λ, and if μ is unstable, so is λ. A value on λ close to one indicates that μ is close to zero for a given discount rate. In this case, agents are more concerned with period-to-period changes in the price, than with being off the long-run equilibrium path. The latter discrepancy is given relatively high weight if λ goes toward zero (and μ goes toward infinity).

The expectations variables, Δpk^e_{t+s}, Δpw^e_{t+s}, Δpe^e_{t+s} and Δz^e_{t+s} are unobservable and one needs additional assumptions to make the above relationship operational. Here, I assume that expectations are formed according to the hypothesis of rational expectations. According to this hypothesis, agents' expectations equal the mathematical expectations conditional upon a given set of information plus a white noise error term, called the prediction error, $\omega_{i,t+s}$ (i=pw, pk, pe, z). Additionally, one may assume a certain data generating process in lagged and current variables for the variables for which expectations are formed, and, as the next step, replace the unobservables with the assumed process *and* the prediction error. This is not done in this paper, in that I make use of the unbiasedness property which is derived from the hypothesis of rational expectations. This property states that the difference between the realized value and the expectations held by agents equals the prediction error[1]. One may consequently use the realizations as proxies for the expectations variables.

In equation (5), there is a total of four groups of expectational variables. In order to simplify, I assume that agents' expect the competitive strength between domestic and foreign producers to remain stable at the latest observed level ($m_{t+s} = m_{t-1}$, s=0,1,2...), and accordingly $\Delta z^e_{t+s} = m_{t-1} (\Delta pk^e_{t+s} - \Delta pw^e_{t+s})$. I then arrive at the following model which describes changes in domestic prices as a function of observable variables, solely, in addition to the error term, v_t. The model is non-linear in the parameters β_1, β_2, β_3, β_4, λ and δ.

$$\Delta bh_t = -(1-\lambda)[bh_{t-1} - \beta_1 pw_{t-1} - \beta_2 pk_{t-1} - \beta_3 pe_{t-1} - \beta_4 z_{t-1}]$$

$$(6) \qquad + (1-\lambda)\sum_{s=0}^{\infty}(\lambda\delta)^s[(\beta_1+\beta_4 m_{t-1})\Delta pw_{t+s} + (\beta_2 - \beta_4 m_{t-1})\Delta pk_{t+s} + \beta_3 \Delta pe_{t+s}] + v_t$$

$$v_t = (1-\lambda)\sum_{s=0}^{\infty}(\lambda\delta)^s[(1-\lambda\delta)u_{t+s} + (\beta_1+\beta_4 m_{t-1})\omega_{1,t+s} + (\beta_2-\beta_4 m_{t-1})\omega_{2,t+s} + \beta_3\omega_{3,t+s}]$$

The error term in equation (6) is a function of the error term from the long-run equilibrium model and the prediction errors. All errors are represented by lead structures, so autocorrelation is likely to be present in the error process described by v_t. Heteroscedasticity may arise through the inclusion of m_{t-1} in the error term. In both cases, one cannot use the standard formulae for standard errors of the estimated parameters. Consistent standard errors may be derived from the Newey and West (1987) heteroscedasticity and autocorrelation consistent variance-covariance matrix.

From the unbiasedness property (see footnote 1), one knows that the realizations, Δpk_{t+s}, Δpw_{t+s} and Δpe_{t+s}, are correlated with their respective prediction errors, $\omega_{i,t+s}$. Δpk_{t+s}, Δpw_{t+s} and Δpe_{t+s} are thus not weakly exogenous to the parameters in equation (6). I apply an errors-in-variables method, namely the 2SLS (two-stage least squares) and the non-linear 2SLS (NL-2SLS). The latter is used when I estimate subject to the over-identifying restrictions on parameters in equation (6). The NL-2SLS estimation method applied on equation (6) is not fully efficient, as I do not make use of the derived cross-restrictions between the parameters in the model for Δbh_t and in the model for the error process, v_t.

The over-identifying restrictions on parameters in equation (6) are known as the backward-forward restrictions and may be tested empirically. The lead coefficients decline geometrically and are related to the level part of the equation through the parameter λ. (1-λ) is a parallel to the error correction coefficients in traditional ECMs and gives the speed of adjustment towards the long run equilibrium path. The speed of adjustment increases when λ approaches zero. This is consistent with a high value on the parameter μ.

The dynamic part of equation (6) is derived from a set of theoretically based assumptions. I will confront the resulting forward looking regression with the results from a traditional ECM where the dynamic part only includes lagged and current variables. The dynamic part of the relation is chosen

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[1] The unbiasedness property: $\Delta pw^e_{t+s} = \Delta pw_{t+s} + \omega_{1,t+s}$, $\Delta pk^e_{t+s} = \Delta pk_{t+s} + \omega_{2,t+s}$ and $\Delta pe^e_{t+s} = \Delta pe_{t+s} + \omega_{3,t+s}$

according to what fits the data best as long as certain basic requirements are met, such as restrictions from economic theory, and the regression passes a set of misspecification tests (the "general-to-specific" approach).

The outset for the estimation of an ECM in current and lagged variables, is the following model.

$$(7) \qquad \Delta bh_t = \phi_0 + \phi_1 bh_{t-1} + \phi_2 pw_{t-1} + \phi_3 pk_{t-1} + \phi_4 pe_{t-1} + \phi_5 z_{t-1} + \sum_{i=1}^{i=I} \gamma_{1i} \Delta bh_{t-i} + \sum_{j=0}^{j=J} \gamma_{2j} \Delta pk_{t-j} + \sum_{k=0}^{k=K} \gamma_{3k} \Delta pw_{t-k} + \sum_{l=0}^{l=L} \gamma_{4l} \Delta pe_{t-l} + e_t$$

e_t is a white noise error term. ϕ_1 is the error correction coefficient. Estimates of the long-run elasticities in equation (2) are derived from the estimates of the ϕ-parameters according to the following formulaes:

$$(8) \qquad \beta_1 = -\phi_2 / \phi_1, \quad \beta_2 = -\phi_3 / \phi_1, \quad \beta_3 = -\phi_4 / \phi_1, \quad \beta_4 = -\phi_5 / \phi_1,$$

The model in equation (7) encompasses several theoretically based models. Among these are models derived from the assumption of rational expectations as well as extrapolative expectations, and models derived from different kinds of quadratic loss functions with loss related to current and/or future periods (Nickell (1985)). Equation (7) encompasses a rational expectations model derived from a somewhat more general quadratic loss function than (4), if we assume an autoregressive data generating processes for the variables for which expectations are formed. In equation (6), I made no assumptions concerning the data generating processes for pw, pk and pe.

4. Data and Time-Series Properties

The estimations are carried out on quarterly seasonally unadjusted data from 1972:1 to 1992:1. Seven observations (1992:2-1993:4) are saved for post-sample forecasts. Appendix 2 gives further information on data sources and variable definitions. The analysed price, BH, is the price index on domestic deliveries from Norwegian private mainland economy. I define PW as unit labour costs inclusive of net sector taxes for private mainland economy. In the model, the competing price, PK, represents both the price on competing products from foreign producers and the price on imported intermediate goods. I proxy this variable by the price index (in Norwegian currency) on total imports[2]. As electricity price, PE, I apply the price index on electricity delivered in domestic market. The variable m_t is defined according to equation (9), where I_t is defined as total imports excluding petroleum and shipping, and Q_t is the value added for private mainland economy, both measured in real terms. The indicator increases as the level of imports relative to domestic production increases. An eventual increase may be a result of both an increased importshare of intermediate goods and increased import penetration on domestic markets for final products.

(9) $$m_t = I_t/(I_t+Q_t)$$

Figure 1 (bh_t, pw_t and pe_t), 2 (bh_t and pk_t) and 3 (m_t and z_t) graph the time-series involved in my analysis. All series have a positive trend, but while unit costs, competing prices and the import share (m_t) increase less than domestic prices over the entire period, the electricity price increases more. For all variables, the rate of increase are higher in the first half period than in the second. Competing prices increase more than domestic prices during the first period, but less during the second while unit costs have a lower rate of change than domestic prices in the first period, and about the same in the second. The rate of change in the importshare is higher in the first period than in the second. The marked movements in m_t during the 1970s are due to import movements. The fall in m_t from 1978:2 to 1978:3 follows a eight percent devaluation of the Norwegian currency (NOK) in the beginning of the same year. The seasonal patterns are more pronounced in pw_t, pe_t and m_t than in bh_t and pk_t and they are not stable. During estimation I take account of the seasonal patterns by dummies which may change at certain points (described in section 5).

[2] Since PK also is a proxy for costs of production, I use the price index on total imports (i.e. including petroleum and shipping) in order to capture costs related to the use of petroleum. In equation (9), defining the importshare, I use total imports excluding petroleum and shipping. If I had measured the importshare without excluding the petroleum sector, the measure would have been highly influenced by the growth in the Norwegian petroleum sector during the estimation period.

Figure 1: Domestic prices (bh), unit labour costs (pw) and price on electricity (pe). 72:1-92:1. Source: Statistics Norway.

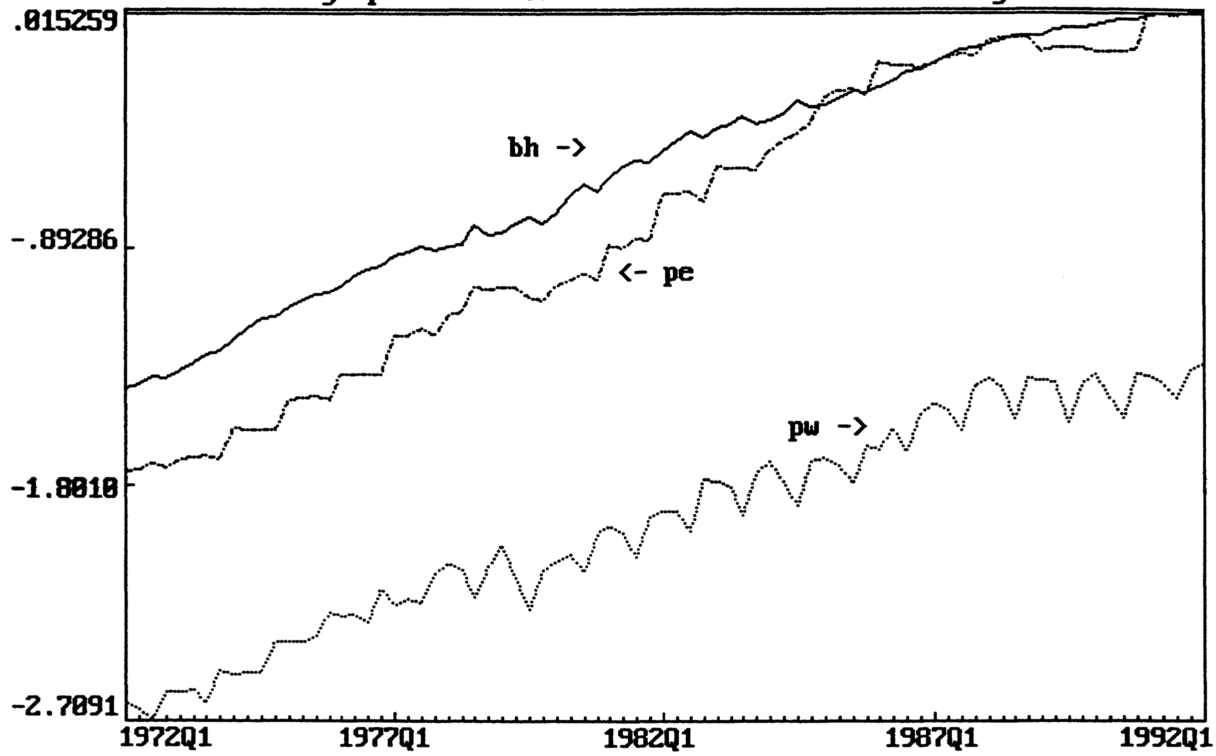


Figure 2: Domestic prices (bh) and competing prices (pk). 72:1-92:1. Source: Statistics Norway.

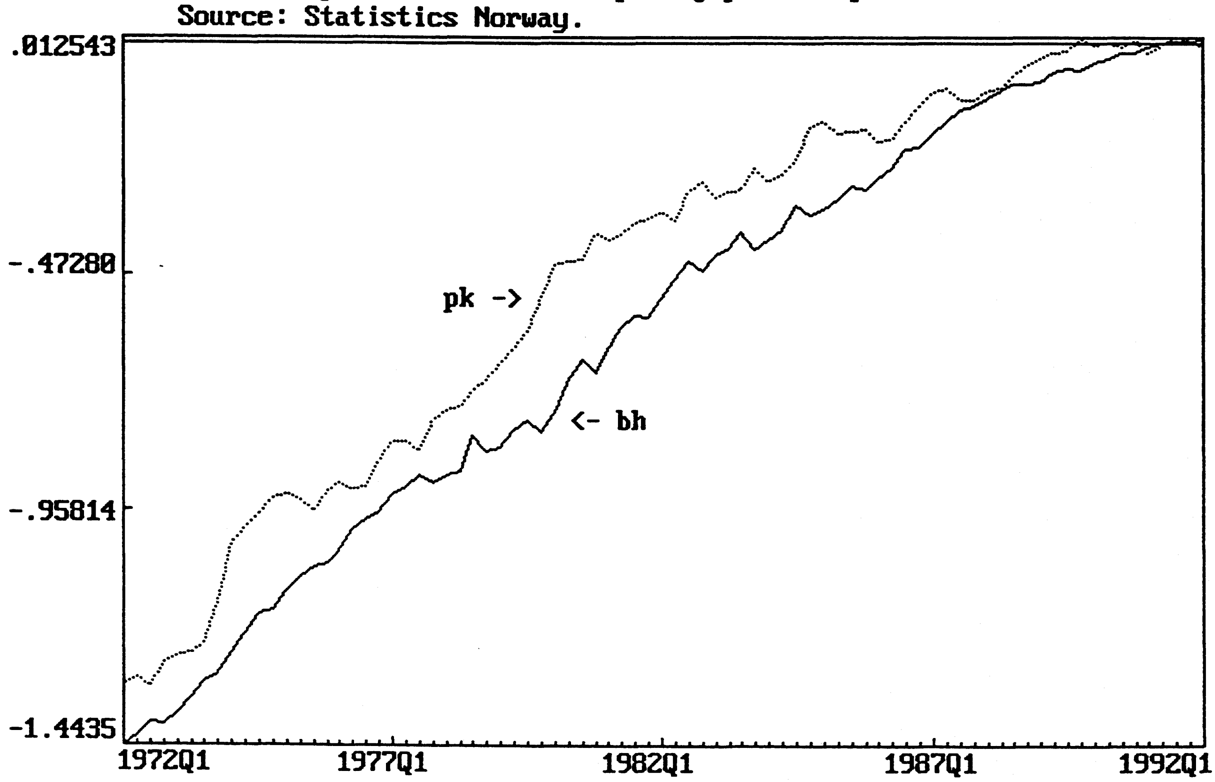


Figure 3: Indicator for competitive strength (m) and log(PK/BH)m (z). 72:1-92:1. Source: Statistics Norway.

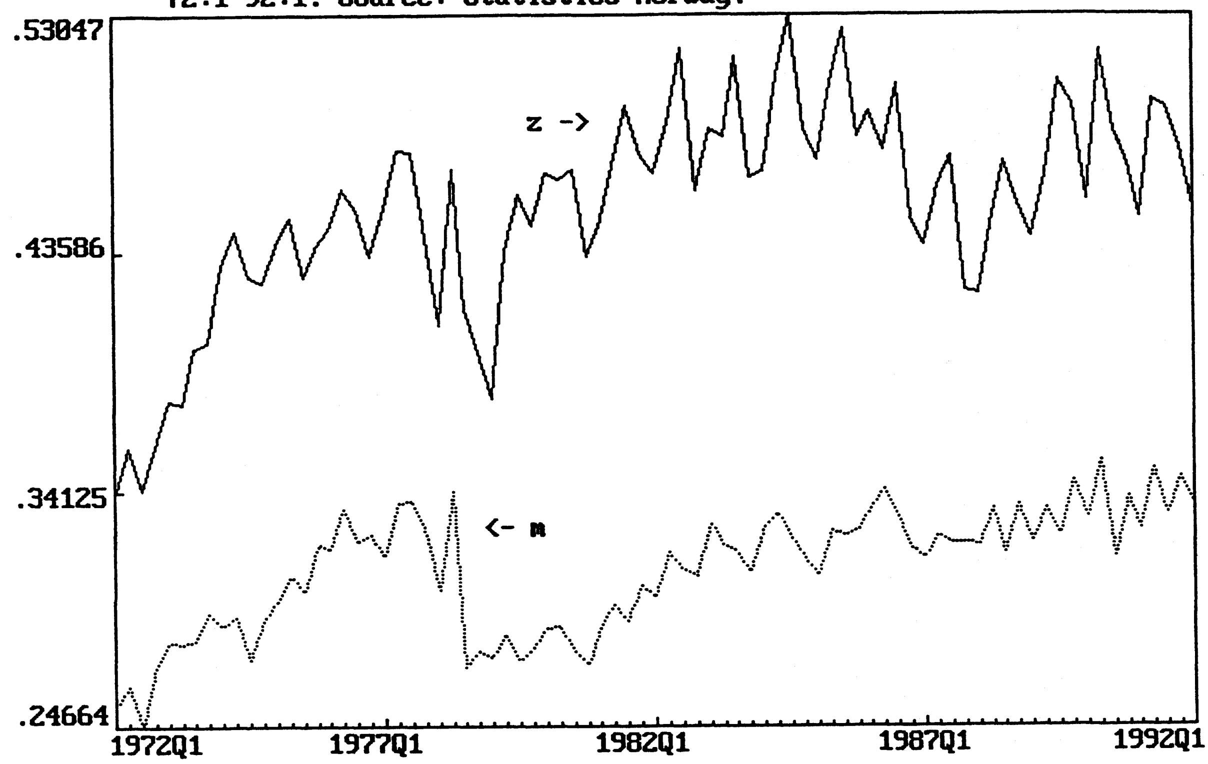
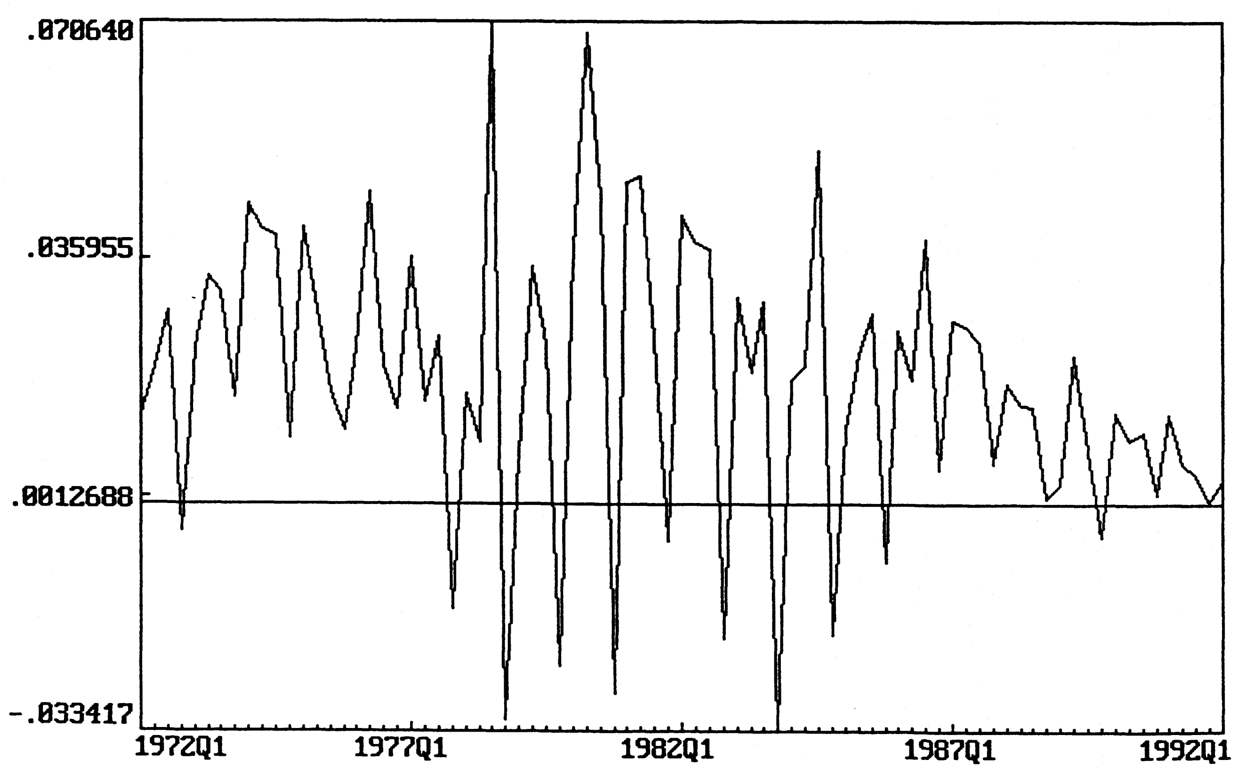


Figure 4: bh-bh(-1). 72:1-92:1. Source: Statistics Norway.



The time-series properties[3] of the applied series are analysed using Augmented Dickey-Fueller (ADF) tests. I report the resulting statistics in table I.

Table I: Augmented Dickey-Fueller tests. 72:1-92:1

Variable	Test[1)]	"T-value"[2)]	Variable	Test [1)]	"T-value"[2)]
bh	ADF(4)	-2.51	Δbh	ADF(3)	-1.92
pw	ADF(7)	-1.82	Δpw	ADF(6)	-3.84
pe	ADF(4)	-2.46	Δpe	ADF(3)	-3.58
pk	ADF(6)	-1.39	Δpk	ADF(5)	-4.26
z	ADF(4)	-2.38	Δz	ADF(6)	-4.80

[1] The order of the tests are indicated in the parentheses, and are chosen according to the highest significant lag in the ADF regression.
[2] Critical value at a 5% level is -2.90 (computed using the response surface estimates in MacKinnon (1990)).

We cannot reject the hypothesis that all the variables with the exception of the dependent variable, bh_t, are I(1), according to the ADF-tests. The rates of change in these variables, Δpw_t, Δpe_t, Δpk_t and Δz_t are thus stationary. As bh_t is concerned, the results indicate that the inflation rate, Δbh_t, is non-stationary and that bh_t is I(2) over the analysed period[4]. Figure 4 shows that changes in domestic prices are characterized by positive values in the first part of the period, high volatility in the mid-period and then followed by a period with decreasing values. There have been several instances of price regulations during the estimation period, and specially between 1977 and 1980. The Norwegian currency has been devaluated several times, with the largest ones in 1978 (8%) and 1986 (12%). The observed shift in the inflation rate after 1987 is partly due to a shift in policy, with more concern about inflation relative to unemployment, combined with low activity in domestic and international economy. Figure 4 also indicates that structural changes in the seasonal pattern have occurred twice during the period, the first at the beginning of 1978 and the second around 1985. Over a longer period of time, I assume domestic prices will appear as an I(1)-variable. I find that if I exclude the three last years from the sample, the test statistics are very close to not reject the hypothesis that domestic prices are I(1). I have chosen to treat it as an I(1)-variable during my analysis. Possible problems with the estimated regression may however be a result of the time-series properties of my dependent variable.

The two regression models are balanced if the variables included in the models are all I(1) and the level terms cointegrate. A balanced equation will lead to valid inference. Whether the level terms cointegrate, and eventually how many cointegrating vectors there are among them, are tested by use of the maximum likelihood procedure developed by Johansen (1988). He proposes two different tests of the number of cointegrating vectors, the max eigenvalue test and the trace test. The results of the tests[5] are given in table II. According to the first of these tests, we have to reject that there are at least one cointegrating vector while the second test does not reject that there are at least two vectors that cointegrate. The conclusion is thus not obvious. I do, however, report the estimated vector corresponding to the highest eigenvalue in table III. The estimated values (column (a)) are quite reasonable, but two of the coefficients are not significant. The restriction of static homogeneity is not rejected according to a likelihood ratio (LR) test at a significance probability of 0.73. The coefficients estimated subject to this restriction (column (b)), are all close to the constrained coefficients and the estimated elasticities for 92:1[6] are $El_{PW}BH=0.40$, $El_{PK}BH=0.41$ and $El_{PE}BH=0.20$.

[3] All empirical results are derived by the software package Microfit 3.21(Pesaran and Pesaran (1991)).
[4] The ADF(2) test statistic for $\Delta\Delta bh_t$ is -21.42 which clearly exceeds numerically the critical value of -2.90.
[5] Critical values are calculated according to Osterwald-Lenum (1992).
[6] m_t equals 0.336 in 92:1.

Table II: Johansen maximum likelihood procedure. Cointegration LR test.
N=81 (72:1 to 92:1). VAR(5)-model for bh, pw, pk, pe, z. Additional I(0) variables included in the VAR: centered seasonal dummies. Eigenvalues: .29, .24, .16, .11, .06

Max eigenvalue test				Eigenvalue trace test			
H_0	H_1	Statistic	95% Critical Value	H_0	H_1	Statistic	95% Critical Value
$r = 0$	$r = 1$	27.76	33.32	$r = 0$	$r > 1$	77.56	70.60
$r \leq 1$	$r = 2$	21.86	27.14	$r \leq 1$	$r > 2$	49.80	48.28
$r \leq 2$	$r = 3$	14.06	21.07	$r \leq 2$	$r > 3$	27.93	31.53
$r \leq 3$	$r = 4$	9.01	14.90	$r \leq 3$	$r > 4$	13.87	17.95
$r \leq 4$	$r = 5$	4.86	8.18	$r \leq 4$	$r > 5$	4.86	8.18

Table III: Estimated cointegrated vector using the Johansen procedure, normalized on domestic prices (bh), corresponding to the eigenvalue 0.29

Variable	(a)	(b)
pw..	.49	.43[1)]
pk..	.36*	.36[1)]
pe..	.17*	.20[1)]
z..	-.13	-.10

* (**): significant at a 5% (1%) significance level.
[1] Estimated subject to the restriction of static homogeneity. Significance probabilities are not available in Microfit 3.21 when the vector is estimated subject to a restriction.

5. An ECM for domestic prices

The cointegrating vector shown in the last section, is derived from a VAR-analysis, that treats all variables as endogenous. In the proceeding, I leave this multi-variate framework and base the analysis on a single-equation approach. I do not want the analysis to rely too heavily on the cointegration analysis, both because the doubt about the dependent variable being I(1) and the inconclusive results concerning the number of cointegrating vectors. In addition I found that two of the estimated parameters in the vector related to the highest eigenvalue, are insignificant at a 5% level.

The outset for estimating the backward looking ECM is the regression model in equation (7). I began with up to five lags on the differences. The model is expanded with centered seasonal dummies (d1, d2, d3) and dummies for structural changes in the seasonal pattern by the end of 1977 (dk77) and 1985 (dk85). There have been several occasions of politically set restrictions on prices and wages during the estimation period. Therefore, I have included two more dummies, one that captures the effects on prices in periods with regulations and a second that captures the effects of catching up with the trend after regulations stop (see Bowitz and Cappelen (1994)). The two dummies were, however, never significant.

Equation (7) is a conditional econometric model, in that I condition on the current variables, Δpw_t, Δpk_t and Δpe_t. The conditioning variables should be weakly exogenous for the parameters of interest to get valid inference from our regression (Engle et al. (1983)). This is not the case if the variables Δbh_t, Δpw_t, Δpk_t and Δpe_t are determined simultaneously in a multi-variate system. If so, valuable information is lost when one relies on a single-equation approach. The electricity prices are partly given by long-term contracts[7] and meteorological factors, while I assume that the price index on total import is decided abroad[8]. Consequently, I assume that Δpk_t and Δpe_t are weakly exogenous in our regression model. Theory suggests that domestic prices and labour unit costs are decided simultaneously, so one cannot treat Δpw_t as weakly exogenous in the regression for Δbh_t. Ordinary least squares (OLS) estimators are thus inconsistent. Consistent estimators are derived by use of an errors-in-variables method, 2SLS with instruments for Δpw_t[9]. The coefficient for Δpw_t never became significant during estimations and the variable is excluded from my preferred equation, which is estimated by OLS. The only conditioning variable in the preferred equation is Δpe_t.

My preferred ECM, is reported as equation (b) in table IV. The regression is estimated subject to the restriction of static homogeneity ($\beta_2 = 1 - \beta_1 - \beta_3$). I test the restriction in relation to the unconstrained regression, reported as equation (a). The restriction is not rejected according to the observed Wald-statistics ($\chi^2_{WALD}(1)$). Not surprisingly, as the unconstrained elasticities sum up to 0.99, I find that coefficients and test-statistics are insensitive to the cross-restriction on the parameters.

All coefficients except the coefficient of Δbh_{t-3}, which has a t-value of 1.86, are significant. The inclusion of this variable improves the Cusum and Cusumq plots. The three elasticities are significant at a 1% level. A variable deletion test of the zero-restrictions on the excluded variables (with I=L=4 and J=K=5), does not lead to rejection at a significant level of 0.92 according to a Wald-test.

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7 The long-term contracts are mainly given to the electricity-intensive industry.

8 Some recent empirical results on the determination of Norwegian import prices may question this assumption (Naug and Nymoen (1993)).

9 Instruments for Δpw_t in addition to the weakly exogenous variables in the regression: UR_{t-1}, ΔUR_{t-3}, $\Delta trtn_t$, $\Delta ytsx_{t-2}$ (see appendix 2 for definition of variables). These are selected, applying variable deletion tests, from a larger set of variables.

Table IV: ECM for domestic prices. Dependent variable: Δbh_t. OLS. Estimation period: 72:1 to 92:1. Forecasts: 92:2 to 93:4

	Equation (a)			Equation (b)[4]		
Regressor	Coeff.	(s.e.)	(N-W s.e)[5]	Coeff.	(s.e.)	(N-W s.e.)
Constant	.15**	(.04)[2]	(.04)	.16**	(.03)	(.03)
bh_{t-1}	-.30**	(.05)	(.05)	-.30**	(.05)	(.05)
pk_{t-1}	.10**	(.02)	(.02)	.11**	(.02)[4]	(.02)
pw_{t-1}	.16**	(.03)	(.03)	.16**	(.03)[4]	(.03)
pe_{t-1}	.03*	(.02)	(.02)	.03**	(.01)[4]	(.01)
z_{t-1}	.16**	(.05)	(.04)	.15**	(.05)	(.04)
Δbh_{t-3}	.17	(.09)	(.11)	.15	(.08)	(.10)
Δpe_t	.08*	(.03)	(.02)	.08*	(.03)	(.02)
Δpe_{t-3}	-.11**	(.04)	(.04)	-.11**	(.03)	(.03)
d1	-.01**	(.003)	(.002)	-.01**	(.003)	(.002)
dk77*d3	-.02**	(.004)	(.003)	-.02**	(.004)	(.003)
dk85*d3	.02**	(.004)	(.004)	.02**	(.004)	(.004)
$El_{PW}BH$[1]	.35**	(.08)	(.08)	.37**	(.04)[4]	(.05)
$El_{PK}BH$[1]	.53**	(.05)	(.04)	.53**	(.05)[4]	(.04)
$El_{PE}BH$	.11**	(.05)	(.05)	.10**	(.03)[4]	(.02)
$\hat{\beta}_4$	-.53**	(.16)	(.06)	-.51**	(.14)	(.09)
R^2		.82			.82	
100 SER[7]		942			936	
DW		1.74			1.73	
$\chi^2_{SC}(4)$[2]		3.70			3.61	
$\chi^2_{RESET-IV}(1)$		.65			.57	
$\chi^2_N(2)$		.22			.21	
$\chi^2_{HET}(1)$		3.84*			3.78	
$\chi^2_{ARCH}(4)$		6.32			5.85	
$\chi^2_{CHOW}(7)$		2.60			2.65	
t_{ECM}		-6.28**			-6.40**	
Wu-Hausman[6]		.68			.81	
$\chi^2_{WALD}(1)$[3]		.15				

* (**): Significant at a 5% (1%) level.
1 Calculated for t=1992:1.
2 Critical values 5% (1%) level: $\chi^2(1)$=3.84 (6.63), $\chi^2(2)$=5.99 (9.21), $\chi^2(4)$=9.49 (13.28), $\chi^2(7)$=14.07 (18.48).
3 Test of the restriction; $El_{PW}BH+El_{PK}BH+El_{PE}BH = 1$ (static homogeneity).
4 Estimated subject to the restriction of static homogeneity.
5 Newey-West standard errors.
6 Significance probabilities: (a) 0.413 and (b) 0.373. The statistic follows a $F(v_1,1)$-distribution with $v_1 = 68$ in (a) and $v_1=69$ in (b).
7 SER: Standard error of regression

I report a set of statistics testing for misspecification. The RESET-test (Ramsey (1969)), $\chi^2_{RESET}(1)$, is a test for functional form misspecification, while the Jarque-Bera's test statistic, $\chi^2_N(2)$, is a test for the normality of regression residuals (Jarque and Bera (1980)). $\chi^2_{SC}(4)$ is Godfrey's test statistic for residual serial correlation (Breusch and Godfrey (1981), App.B). I report two different tests for heteroscedasticity. $\chi^2_{HET}(1)$ is based on the regression of squared fitted values on squared residuals (Koenker (1981)), while $\chi^2_{ARCH}(1)$ is the autoregressive-conditional heteroscedasticity test statistic (Engle (1982)). Equation (b) passes all our tests for misspecification, but we note that the observed value on $\chi^2_{HET}(1)$ is quite high. The null of no homoscedasticity is rejected at a 5% level in equation (a), which is not estimated subject to the homogeneity restriction. Both equations pass the ARCH-test for heteroscedasticity. The standard errors of the coefficients, calculated according to the standard OLS-formulae, are no longer valid if heteroscedasticity is present. Consistent standard errors (N-W s.e.) based on the Newey-West heteroscedasticity and autocorrelation consistent estimator of the variance-covariance matrix do not alter our conclusions concerning the significance of the long-run elasticities or the validity of the homogeneity restriction. The adjusted t-value for the coefficient of Δbh_{t-3} decreases, while the one for the coefficient of Δpe_t increases.

The adequacy of predictions is tested by use of Chow's second test (Chow (1960), and the resulting statistics is reported as $\chi^2_{CHOW}(7)$ in the table. The assumption of Δpe_t being weakly exogenous in the regression for Δbh_t is tested by use of a Wu-Hausman test (Wu (1973)). The reported statistics, do not lead to a rejection of the hypothesis of weak exogeneity.

The results in section 4 were inconclusive on whether our level variables cointegrate or not. Kremers et al. (1992) propose a test for cointegration among the level variables, based on the t-ratio of the error correction coefficient. This statistic is reported as t_{ECM} in the table. The observed values are well above the critical values which are calculated using the surface estimates in MacKinnon (1991), so we cannot reject that the level variables in the ECM cointegrate according to this test. The validity of the test does however depend on whether the cointegrating vector only appears in the ECM for Δbh_t and not also in the ECMs for the other variables which I include in the vector.

In my preferred equation, I find that changes in domestic prices mostly are a result of previous discrepancies from the long-run path, which is reduced by 30% in each period, and to a minor extent a function of short term movements in the explanatory variables. The only variables that I include in the dynamic part are changes in the electricity price (Δpe_t and Δpe_{t-3}) and lagged values of the endogenous variable. The sum of the coefficients of Δpe_t and Δpe_{t-3} equals -.03 and is not different from zero at a significance level of 0.53. The long-run level of BH_t depends on the long-run levels of PW_t, PK_t, PE_t and m_t if the model complies with dynamic homogeneity[10]. This restriction is clearly rejected by a Wald-test ($\chi^2_{WALD}(1)=123.63$), and the long-run growth rate in PE_t will affect the long-run level of BH_t. Dynamic inhomogeneity in macroeconomic models may lead to counter-intuitive results. A permanent increase in the rate of inflation will lower the mark-up rate and give higher real-wages. Even if this result are consistent with observed data during the estimation period, they may be quite wrong for subsequent periods. A new policy-rule that leads to a permanent shift in the rate of inflation, may alter the structural parameters in such a way that real-wages remain unchanged. Nickell (1988) warns against the use of models that exhibit dynamic inhomogeneity in analysis that entail large shifts in the inflation. Instead, he proposes a more explicit modelling of expectations variables to account for that these variables depend on policy-rules.

The elasticities with respect to unit labour costs ($El_{PW}BH$) and to competitive prices ($EL_{PK}BH$) vary over time. The estimate of β_4 is negative. Consequently, $El_{PK}BH$ ($El_{PW}BH$) increases (decreases) with the import share, m_t, which measure the degree of competition from abroad. The time paths of the two elasticities (figure 5) illustrate the effect of a more open economy. The competitive price elasticity increases while the unit labour cost elasticity decreases during the estimation period. $El_{PW}BH$ varies over the interval $\langle 0.37, 0.42\rangle$ and $El_{PK}BH$ varies over the interval $\langle 0.48, 0.54\rangle$. Their average values are 0.39 for $El_{PW}BH$ and 0.52 for $El_{PK}BH$.

Figure 6 and figure 7 show the results of the Cusum and Cusumq tests of structural stability (Brown et al. (1975)). The Cusumq plot suggests a sudden instability to occur in the regression coefficients in 1987. This instability coincides with a reduced rate of increase in domestic prices. Recursive plots of estimated coefficients are found in figure a-k, appendix 4. Most coefficients remain stable during the last part of the estimation period, from about 1987, but for several of them, a minor shift takes place in 1986/1987. I have run the regression over different intervals (t=72:1,...,T with T=80:1,...,88:4), to get

[10] Dynamic homogeneity is equivalent with the following restriction on the short-run parameters in (7):

$$(1/\phi_1)\left(\sum_{i=1}^{i=I}\gamma_{1i} + \sum_{j=0}^{j=J}\gamma_{2j} + \sum_{k=0}^{k=K}\gamma_{3k} + \sum_{l=0}^{l=L}\gamma_{4l} - 1\right) = 0$$

a time series for the predictive failure test, $\chi^2_{CHOW}(p)$ (p equals number of post sample periods). The null hypothesis of predictive failure is clearly rejected for all choice of T.

Figure 8 and figure 9 graph actual values of Δbh_t and bh_t with their respective values from static simulations of the regression. The last seven periods in figure 9 are out of sample forecasts. According to figure 8, the regression does not always track the rate of change. It does, however, track a lower inflation rate after 1987, quite well.

Figure 5: Time-varying elasticities. Backward looking ECM of domestic prices.

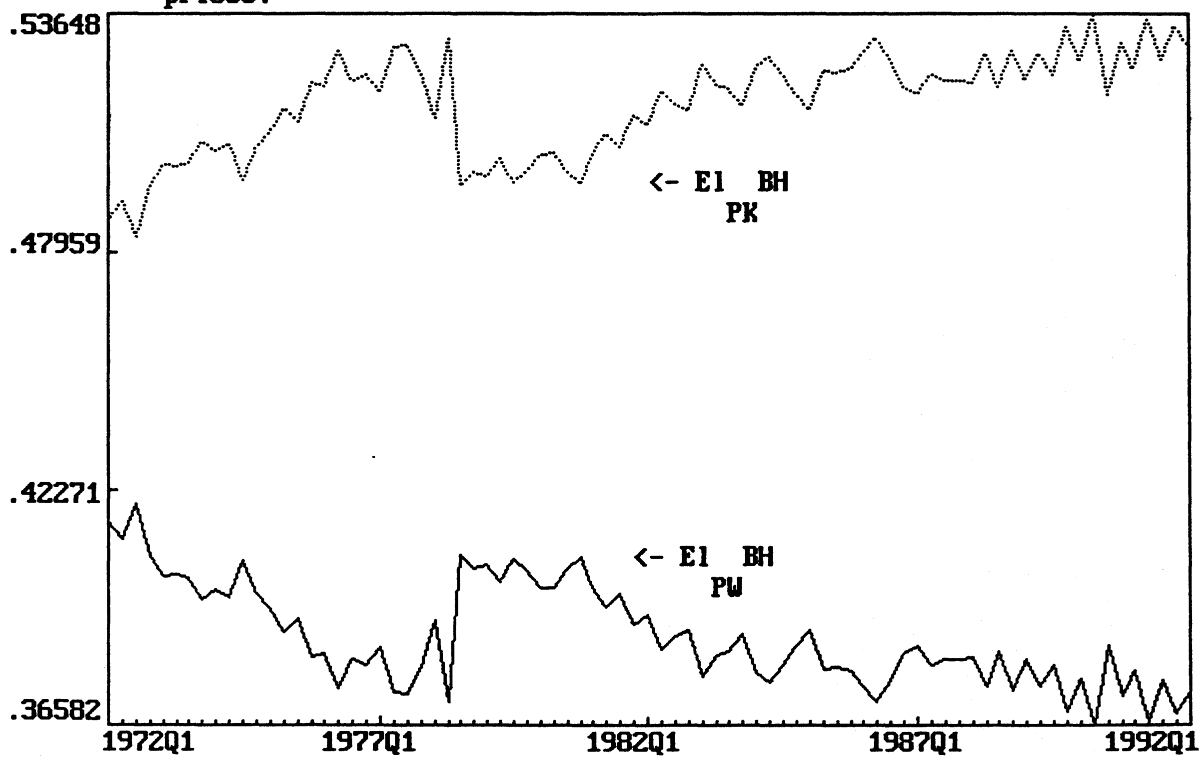
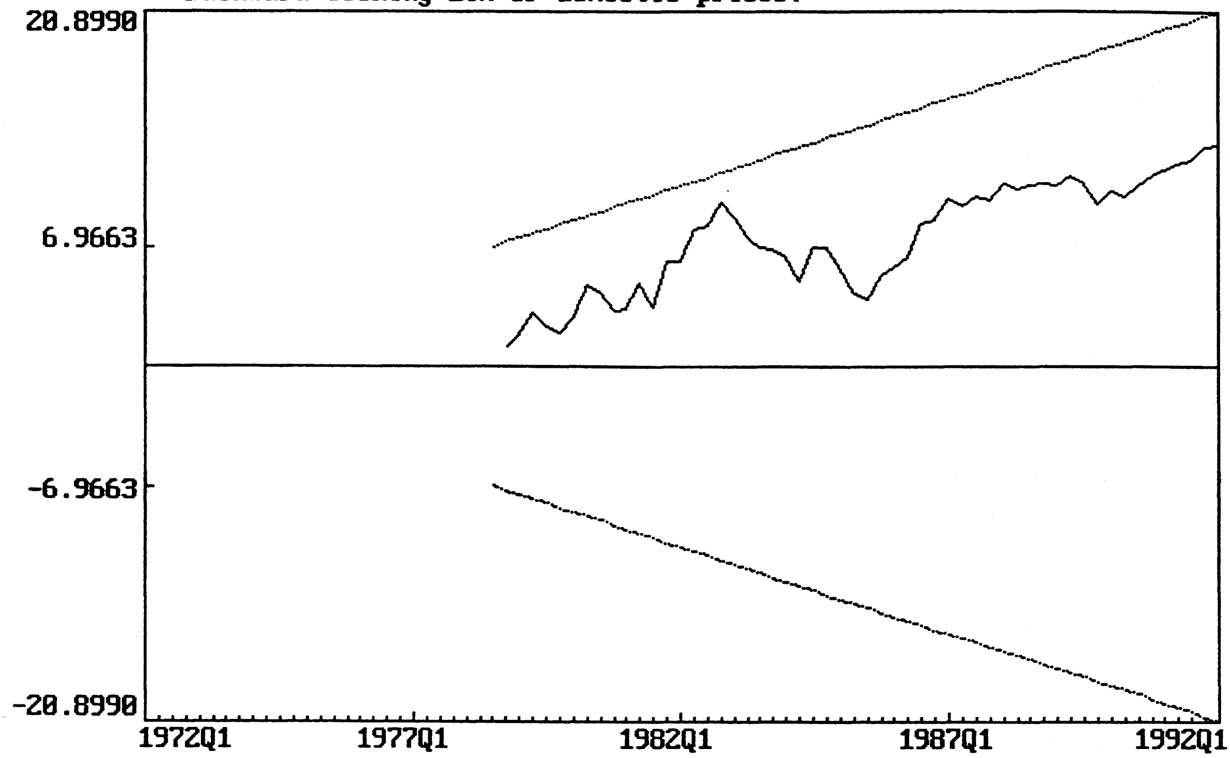
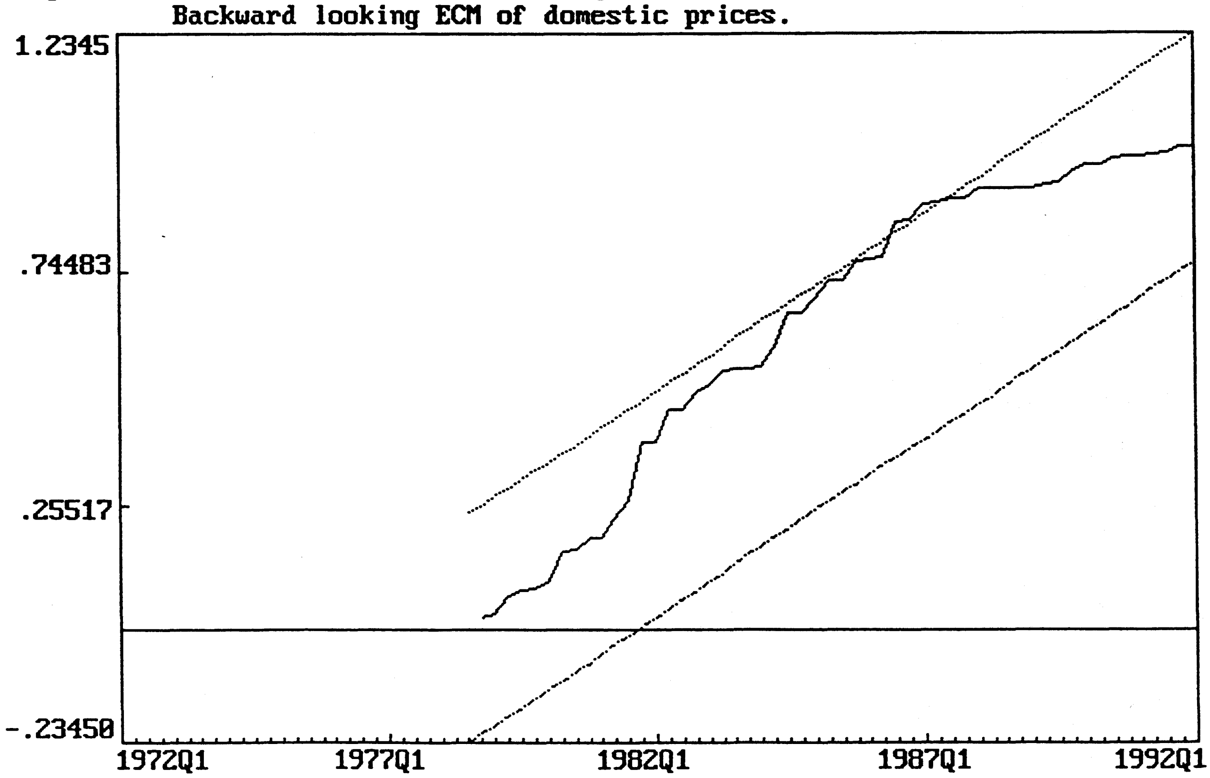


Figure 6: Plot of cumulative sum of recursive residuals (CUSUM).
Backward looking ECM of domestic prices.



The straight lines represent critical bounds at 5% significance level

Figure 7: Plot of cumulative sum of squares of recursive residuals (CUSUMQ).
Backward looking ECM of domestic prices.



The straight lines represent critical bounds at 5% significance level

Figure 8: Plot of actual and fitted values of bh(t)-bh(t-1). Backward looking ECM of domestic prices.

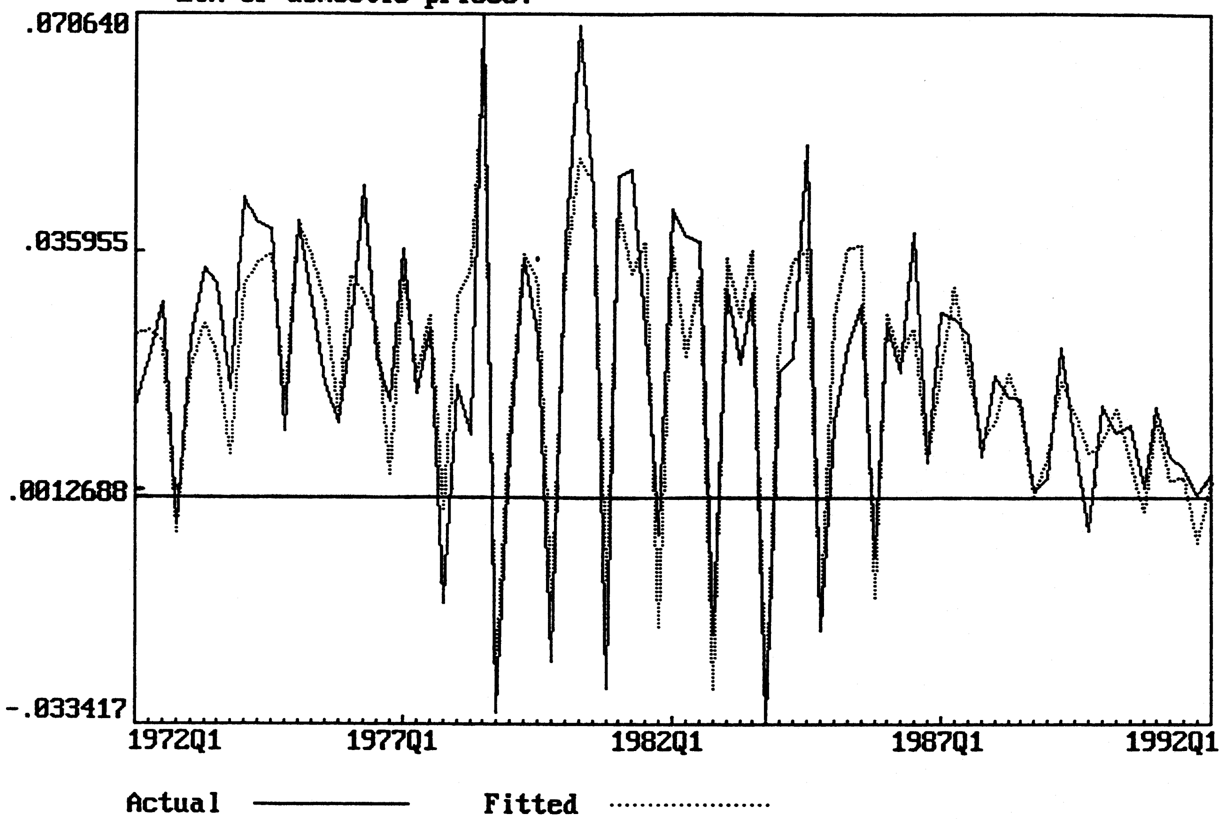
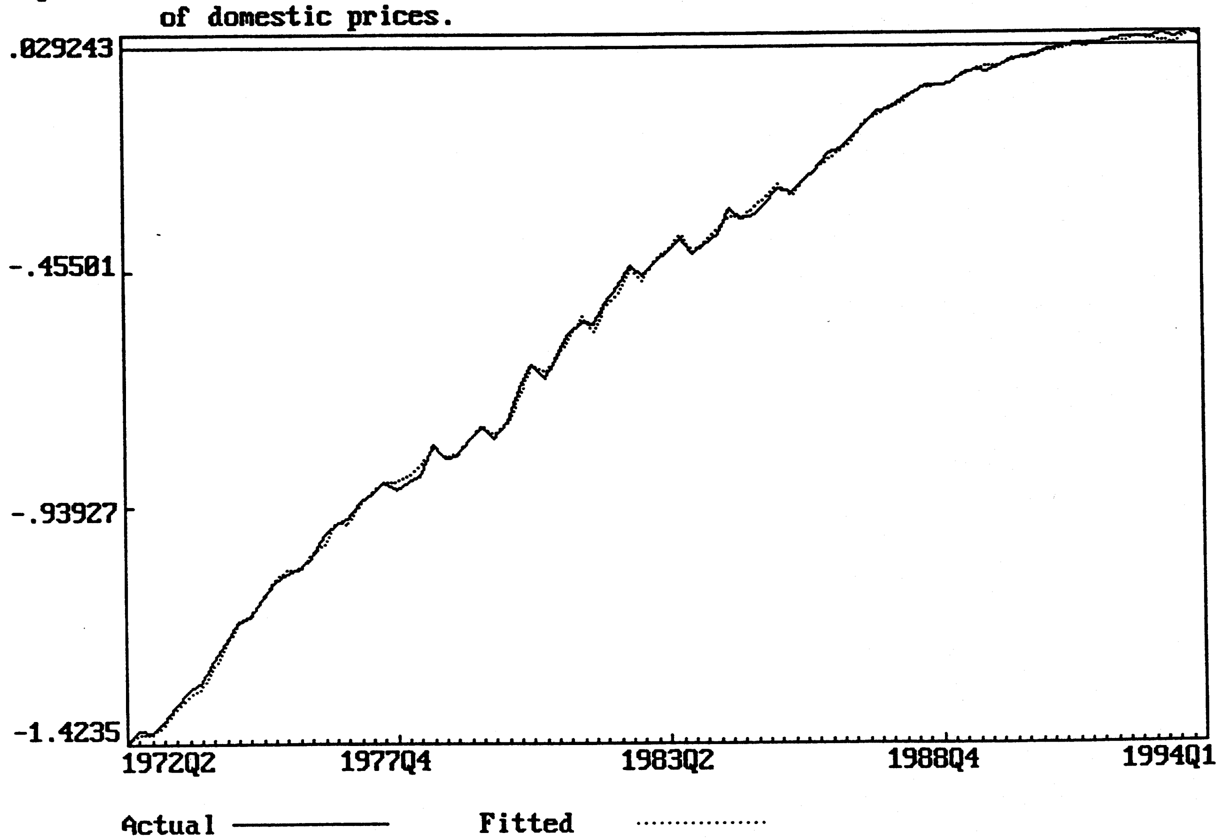


Figure 9: Plot of actual and fitted values of bh(t); Backward looking ECM of domestic prices.



6. A forward looking equation for domestic prices

Table V presents the empirical results for the non-linear forward looking model in equation (6), with the number of leads strictly restricted to equal one[11]. The regression is augmented with seasonal dummies and dummies for price and wage regulations during estimation, as in the backward looking ECM. Insignificant dummies are subsequently excluded from the regression. Different requirements have to be met by the estimated coefficients. First, from the price setting theory, we get restrictions on the sign of the long-run elasticities in addition to the restriction of static homogeneity. Equation (a) is estimated without imposing the homogeneity restriction. The restriction may be imposed according to the result of a $\chi^2_{WALD}(1)$-test and the coefficients in equation (b) are estimated subject to this restriction. I concentrate on equation (b).

The three elasticities are all significant when the restriction is imposed. So are the other coefficients, with the exception of $\hat{\delta}$. The estimated value of the discount factor is within the interval $\langle 0,1 \rangle$, which is required to give a meaningful interpretation, but the coefficient is not at all precisely estimated and the interval of $\pm$ 2 s.e. does cover both zero and one. A value on λ in the interval $\langle 0,1 \rangle$, ensures stability of the model. The interval of $\pm$ 2 s.e. around $\hat{\lambda}$ lies within this interval. The estimated value on λ indicates a somewhat slower reduction of the discrepancy from the long-run path in the previous period, than indicated by the results for the backward looking model.

The backward-forward restrictions follow from the framework of a multiperiod quadratic loss function and are imposed prior to estimation. The unconstrained equation, which is linear in the parameters, is estimated and the results are reported in appendix 3. The resulting Wald-statistics from testing the over-identifying (backward-forward) restrictions are repeated in the bottom part of table V. For equation (a) the table also includes the result of testing the homogeneity restriction and the backward-forward restrictions simultaneously. In neither case, are the backward-forward restrictions rejected by the data. For the full evaluation of the reported results, it is of importance to note that when I increase the length of the leads, the restrictions are no longer valid.

The reported tests for misspecification are, in addition to those already commented, Sargan's statistic for a general test of misspecification of the model and the validity of the instruments (Sargan (1964)). Chow's test for the adequacy of predictions and the autoregressive-conditional heteroscedasticity test are not available for NL-2SLS-procedure. The reported test-statistics indicate the possibility for heteroscedastic residuals, and so I report the Newey-West standard errors.

The estimated long-run parameters are about the same values as those estimated for the backward looking ECM. The time paths for the two time-varying elasticities are also quite similar. We may derive an estimate on μ, the weight attached to the deviation from the long-run equilibrium path relative to the loss of changing the price from one period to another[12]. The derived estimate, which is a function of λ and δ, is about 0.2. In the loss function, more weight is thus given to future period-to-period changes than to future deviations from the long-run path. This is a quite common result in empirical work on quadratic intertemporal loss functions. Due to the high standard error for $\hat{\delta}$, however, the estimate on μ is not very precise.

[11] The backward-forward restrictions are rejected if higher order leads are included in the model, mainly due to negative coefficients in the unconstrained regression.

[12] $\lambda\mu = (1-\lambda)(1-\lambda\delta)$

Tabell V: A forward looking ECM for domestic prices estimated subject to the backward-forward restrictions. Dependent variable: Δbh_t. NL- 2SLS[13]. 72:1 to 92:1

Parameters	Equation (a) Estimate	(N-W s.e.)	Equation (b)[1)] Estimate	(N-W s.e.)
Constant	.16*	(.08)[1)]	.12**	(.03)
$\hat{\lambda}$	.74**	(.05)	.76**	(.04)
$\hat{\beta}_1$	.59**	(.12)	.53**	(.07)[1)]
$\hat{\beta}_2$	.38**	(.05)	.37**	(.06)[1)]
$\hat{\beta}_3$	.06	(.07)	.10**	(.03)[1)]
$\hat{\beta}_4$	-.48*	(.23)	-.56**	(.16)
$\hat{\delta}$	.68*	(.34)	.58	(.33)
d2	.01**	(.004)	.01**	(.002)
dk77*d3	-.02**	(.003)	-.02**	(.003)
dk85*3	.02**	(.004)	.02**	(.003)
$El_{PW}BH$[2)]	.43*	(.17)	.34**	(.07)[1)]
$El_{PK}BH$[2)]	.54**	(.07)	.56**	(.07)[1)]
$El_{PE}BH$	.06	(.07)	.10**	(.03)[1)]
R^2	.76		.76	
100 SER	1065		1054	
DW	2.02		1.98	
$\chi^2_{SM}(p)$[3)]	32.18	(p=29)	33.35	(p=30)
$\chi^2_{SC}(4)$	4.79		5.12	
$\chi^2_{RESET}(1)$	.03		.13	
$\chi^2_{N}(2)$	1.45		1.33	
$\chi^2_{HET}(1)$	5.54*		5.69*	
$\chi^2_{WALD}(1)$[4)]	.54		-	
$\chi^2_{WALD}(6)$[5)]	5.97		3.88	
$\chi^2_{WALD}(7)$[6)]	6.46		-	

*(**): Significant at a 5% (1%) level.
[1] Estimated subject to the restriction of static homogeneity.
[2] Calculated for t=1992:1.
[3] Critical values 5% (1%) level: $\chi^2(1)=3.84$ (6.63), $\chi^2(2)=5.99$, $\chi^2(4)=9.49$, $\chi^2(6)=12.59$, $\chi^2(7)=14.07$, $\chi^2(29)=42.46$, $\chi^2(30)=43.77$.
[4] Test of static homogeneity; $El_{PW}BH+El_{PK}BH+El_{PE}BH=1$.
[5] Test of backward-forward restrictions.
[6] Test of backward-forward restrictions and the restriction of static homogeneity.

I conclude, from the plots of actual and fitted values of Δbh_t (figure 10) and bh_t (figure 11), that the forward looking model fits the actual series quite well. As for the backward looking model, figure 10 reveals several occasions of underprediction of the rate of change in the first half of the estimation period, but we can also see from both figures that the model tracks the levelling out towards the end of the estimation period and in the post sample period (figure 11), very well.

In table VI I report the implicit lead coefficients derived from equation (b). The coefficients depend on the indicator of competitive strength. The coefficients are calculated for t=92:1. Four of the six coefficients, including both coefficients for expected competing prices, are significant. This is an

[13] Additional instruments: Δpe_{t-1} Δpe_{t-2} Δpe_{t-3} Δpv_{t-2} Δpv_{t-3} $\Delta ytsx_{t-1}$ $\Delta ytsx_{t-2}$ Δpk_{t-1} Δpk_{t-2} Δpk_{t-3} Δpk_{t-4} $trtn_{t-1}$ $\Delta trtn_{t-1}$ $\Delta trtn_{t-2}$ UR_{t-1} ΔUR_{t-3} v_{t-1} Δv_{t-1} Δv_{t-2} Δv_{t-3} d1 d2 d3 dk88 dk85 d74q1 pstopin pstopout.

improvement, compared with the unconstrained results in appendix 3, and is partly gained from the increased degrees of freedom due to the over-identifying restrictions.

Table VI: Implicit coefficients in the lead structure. Estimated subject to static homogeneity and backward-forward restrictions. t=92:1

Variable	Δpw_t	Δpw_{t+1}	Δpk_t	Δpk_{t+1}	Δpe_t	Δpe_{t+1}
Coefficient	.083**	.037	.136**	.060*	.024*	.011
N-W s.e.	.021	.027	.027	.029	.011	.008

* (**): Significant at a 5% (1%) level.

Figure 10: Plot of actual and fitted values of bh(t)-bh(t-1). Forward looking ECM of domestic prices.

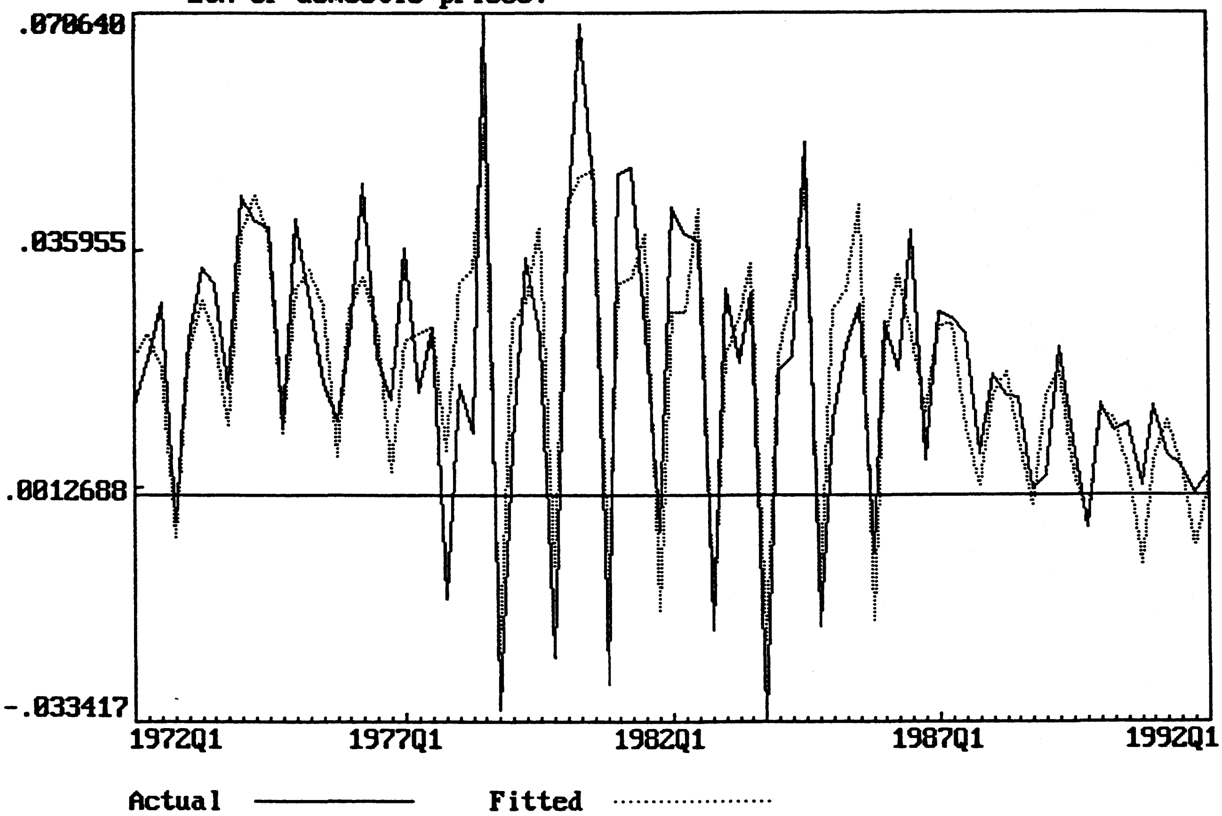
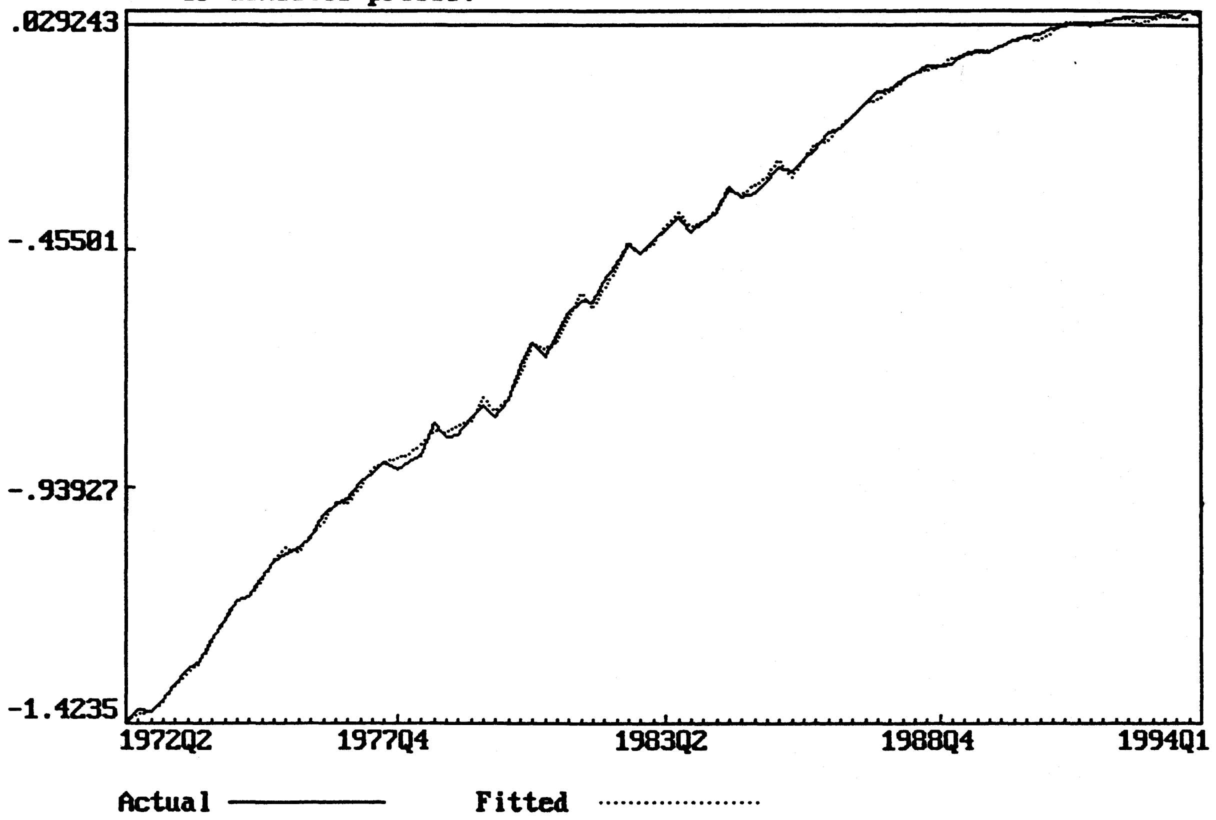


Figure 11: Plot of actual and fitted values of bh(t). Forward looking ECM of domestic prices.



7. Discussion and conclusions

The long-run part of the two models, Model A (equation (b) in table IV) and Model B (equation (b) in table V) are derived from the same theoretical framework, assuming imperfect competition and time-varying elasticities. The two models do however differ in their dynamic specification, i.e. in the way the movements towards the long-run equilibrium is modelled. The dynamic part of Model B is estimated subject to the hypothesis of rational expectations. The expectations are introduced in the model through the multiperiod quadratic loss function. As Model A is concerned, no particular assumptions are made neither regarding whether expectations are part of the decision problem, nor how they eventually are formed. The dynamic specification is chosen according to the general-to-specific strategy. We just note that different sets of assumptions may result in the general model in equation (7).

The estimates of the long-run parameters are quite similar for the two models. The estimated elasticities support the hypothesis that costs of production are determinants for the commodity price at domestic market. At the end of the estimation period, the elasticities of unit labour costs and the electricity price, add to 0.47 in Model A and 0.44 in Model B. And, in addition, we remember that in the two models competing prices both count for the effect of imported goods being a factor of production, and for the competition from foreign producers. I am not able to distinguish between these two effects at my level of aggregation. The reported results in this paper are in line with other results on the price determination in Norway, and are regarded as somewhat of stylized facts (see Aukrust (1977) and Bowitz and Cappelen (1994)).

The main difference between the two models lies in their dynamic parts and consequently in the way they react to shocks or permanent shifts in the exogenous variables. While only lagged levels of PW and PK and lagged levels and changes in PE are included in Model A, Model B is expanded with current and leaded changes in PW, PK and PE. A permanent shift in PK or PW which starts in period t results in a shift in BH already in period t-1 in Model B, but first in period t+1 in Model A. Since shifts in exogenous prices are foreseen in Model B *and* current and future deviations from the long-run path represent a loss, the movements towards the new long-run equilibrium path starts before the permanent shift occurs. But as period-to-period changes also represent a loss, BH moves, just as in Model A, asymptotically towards the new equilibrium. 90 % of the gap between actual and equilibrium price is closed after 5-7 periods in both models. It takes somewhat longer time before the entire gap disappear in Model B than it takes in Model A. This is due to different estimates of the error-correction coefficient (represented by $-(1-\lambda)$ in the forward looking model), and not to the different approaches. The effects in Model B of a permanent shift in PE follows the same pattern as with shifts in PW and PK. In Model A, BH is adjusted in the same period as the shift first appears, but follows then a cyclical pattern relative to the reference scenario before the gap asymptotically is closed.

Two main differences appear when the two models are subject to a one-period shock in PW and PK. First, in Model A the effect on BH occurs one period after the shock, at the same time as the agents realize they have been off the equilibrium path in the previous period. In Model B, the shock is predicted in advance and prices are partly adjusted one period before the shock occurs. Second, the reaction to the shock, measured as the maximum one-period percentage deviation from the reference path, is more pronounced in Model A. In Model B, the agents know they are faced with a shock and not a permanent shift, while they in Model A just act according to what they observe have happened in the previous period. A shock in PE gives an immediately reaction in BH in Model A, followed by a period with cyclical deviations from the reference path. 90% of the maximum deviation from the reference path have disappeared after 8-9 periods independent of model or shocked exogenous variable.

Figure 12: 10% increase in PW, PK and PE from t and out simulation period.

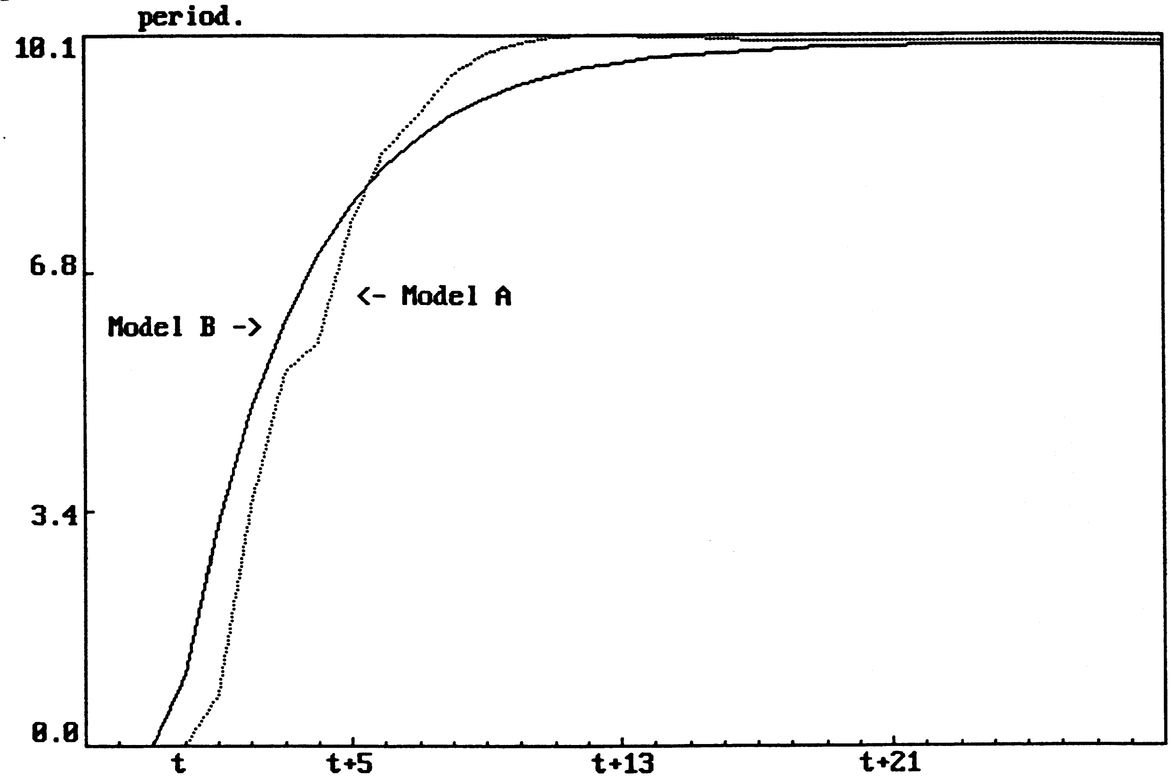


Figure 13: Shock. 10% increase in PW, PK and PE in t.

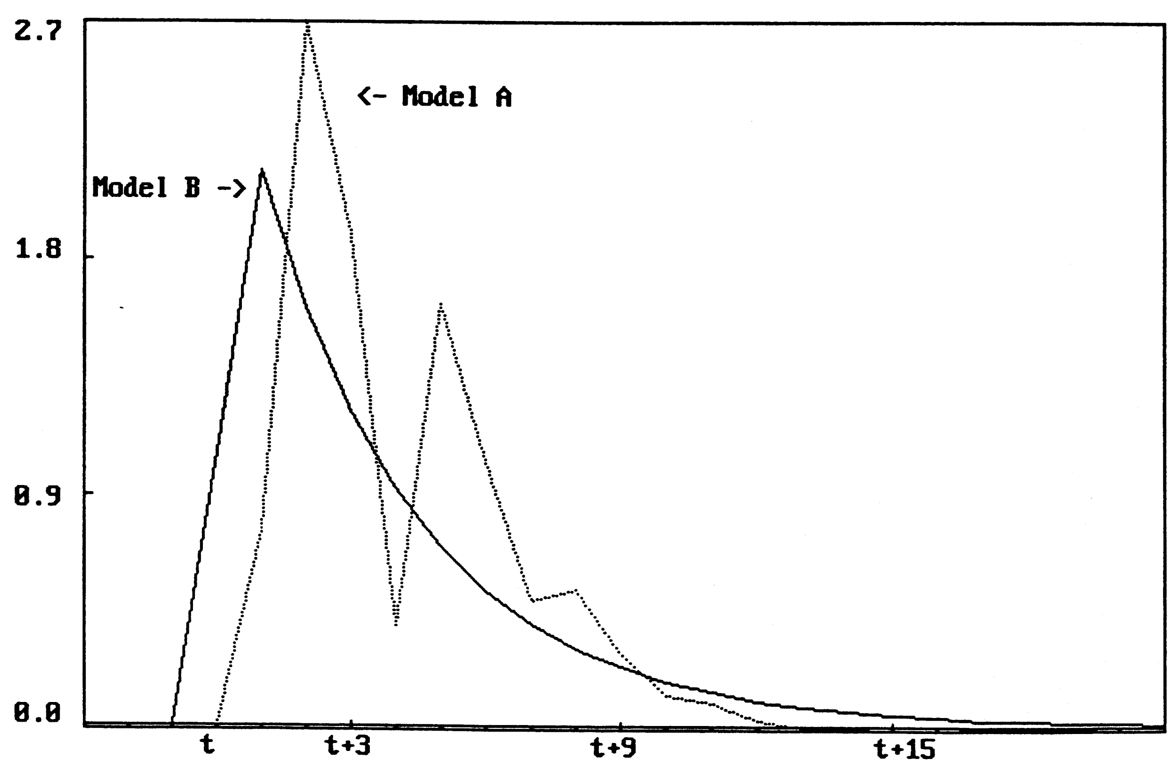


Figure 12 shows the percentage deviations in BH from the reference price, when the two models are subject to a permanent 10% shift in PW, PK and PE, while figure 13 shows the similar deviations following a 10% increase in PW, PK and PE in one single period.

I have used the two models to derive dynamic forecasts for Δbh_t for the sub sample period, 92:2-93:4. Table VII contains summary statistics, measuring the forecast adequacy. It shows that the mean prediction error is somewhat lower for model B than for Model A. This is due to one period of underprediction of the rate of change, while model A overpredicts the rate of change in the whole sub sample period. Model A performs better according to the three other statistics. In addition, Model A shows a higher degree of accuracy during the estimation period with a standard error of regression of 0.0094 (Table IV) compared with 0.0105 for Model B (Table V). I conclude that the results are in favour of Model A both according to forecast abilities and accuracy during estimation period, but the differences between the two models are small.

Table VII: Measures of forecast adequacy. Dynamic forecasts. Estimation period: 72:1-92:1. Forecast period: 92:2-93:4

	Model A	Model B
Mean prediction error	.00610	.00595
Sum of squares of prediction errors	.00005	.00007
Root mean sum of squares of prediction errors	.00728	.00838
Mean sum of absolute prediction errors	.00610	.00626

[1] Model A: Equation (b), table IV. Model B: Equation (b), table V.

Encompassing tests are not easy available when the first submodel is a linear OLS regression and the other is a non-linear 2SLS regression. Some evidence may however be drawn from encompassing tests of Model A vs the linear 2SLS regression in appendix 3 (Model C). Model C is the unconstrained version of Model B, i.e. not estimated subject to the backward-forward restrictions. I have estimated a model encompassing Model A and Model C, using 2SLS. Variable deletion tests imply that Model A encompasses Model C (significance probability 0.839) but that Model C does not encompass Model A (significance probability 0.010). If one expands Model B with the variables Model A and Model B do not have in common (i.e. Δpe_{t-3}, Δbh_{t-3} and d1), a joint zero restriction on these variables is rejected at a significance level of 0.015. The results mentioned above are, even not formally correct executed encompassing tests, evidence in favour of Model A.

The dynamic part of Model B is, as already mentioned, mainly derived from theoretical assumptions. The backward-forward restrictions on the estimated coefficients are not rejected by data. But, a look at the freely estimated coefficients in appendix 3, reveal that the coefficients in the lead structure are not significant different from zero at a 5% level. Joint zero restrictions on the coefficients for Δpw_t, Δpw_{t+1}, Δpe_t, Δpe_{t+1}, Δpk_t and Δpk_{t+1} are neither rejected with a significance probability of 0.686. The relative low accuracy of the estimated lead coefficients in the unrestricted version of Model B, makes it more easy to impose the over-identifying restrictions. On the other hand, the same restrictions result in significant estimates on some of the lead coefficients (see table VI).

The empirical results presented in this paper support the assumptions of imperfect competition in the commodity market and that of time-varying elasticities. The results on the long-run elasticities are in line with the prevailing views on the impacts of wages and foreign prices on domestic prices for the Norwegian economy. Most previous results are derived from disaggregated analyses. In a disaggregated study of domestic prices in the Norwegian economy, Bowitz and Cappelen (1994) find that for most commodities the long-run solution depends solely on unit costs. The price on intermediate goods from both domestic and foreign producers is included in their unit cost variable. The long -run elasticities in Bowitz and Cappelen, derived from simulations of the price-block within a macroeconometric model (KVARTS), are close to the ones presented in this paper. Bowitz and

Cappelen base their study on ECMs in current and lagged variables and does not involve expectations. The elasticities are not allowed to vary across time.

To conclude, the long-run elasticities presented in my paper are quite insensitive to my assumptions concerning how expectations are formed and the dynamics arrived at. The combination of the assumed long-run solution, the multiperiod quadratic loss function and rational expectations is not rejected, but the evidence in favour of this particular model is not strong. This is not the same as saying that expectations do not matter or that they are not formed according to the hypothesis of rational expectations. The results from the unconstrained version of model B, model C, do however not support that expectations, if they matter, follow this particular hypothesis.

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Appendix 1: A Price Model with Time-varying Elasticities

(A1) $BH_t = g(PK_t / BH_t, Y_t) PV_t$

The price, BH, is a function of variable unit costs, PV, and a mark-up, g(BH/PK, Y), where PK is the competing price and Y is the level of demand.

If we assume weak separability in demand for domestic and imported goods, (A1) defines BH as a function of PV and PK.

(A2) $BH_t = h(PV_t, PK_t)$

The h-function is determined from both the supply- and demand-side. We have chosen a general specification and allow the elasticities to vary with the importshare, m_t.

(A3) $BH_t = e^{\alpha_0} PV_t^{\alpha_t} PK_t^{1-\alpha_t}$ *where* $\alpha_t = \alpha_1 + \alpha_4 m_t$

The importshare, m_t, may be interpreted as an indicator for economic changes that affect the elasticities. Such changes may for instance be movements towards a more open economy that reduces the market power of domestic producers. If we instead of a general specification assume that the demand function is derived from a CES (Constant Elasticity of Substitution) utility function, α_4 equals zero and α_1 is a function of the elasticity of substitution between domestic and imported goods.

Variable unit costs is described as a function of unit labour costs, PW, the price of intermediate products, PH, and the electricity price, PE. If the short-run cost-function is Cobb-Douglas we have

(A4) $PV_t = PH_t^{1-\kappa_1-\kappa_2} PW_t^{\kappa_1} PE_t^{\kappa_2}$

We combine (A3) and (A4), make the relation linear in the parameters through a logarithmic transformation, and rearrange.

$$\text{(A5)} \quad \begin{aligned} bh_t &= \alpha_0 + (1-\alpha_1-\alpha_4 m_t) pk_t + (1-\kappa_1-\kappa_2)(\alpha_1+\alpha_4 m_t) ph_t \\ &\quad + \kappa_1(\alpha_1+\alpha_4 m_t) pw_t + \kappa_2(\alpha_1+\alpha_4 m_t) pe_t \end{aligned}$$

where bh=log(BH), pk=log(PK), ph=log(PH), pw=log(PW) and pe=log(PE).

We continue with assuming that the price on intermediate products may be proxyed by the competing price, PH=PK.

$$\text{(A6)} \quad \begin{aligned} bh_t &= \alpha_0 + \kappa_1\alpha_1 pw_t + (1-(\kappa_1+\kappa_2)\alpha_1) pk_t + \kappa_2\alpha_1 pe_t \\ &\quad + \left[\kappa_1 pw_t - (\kappa_1+\kappa_2) pk_t + \kappa_2 pe_t\right]\alpha_4 m_t \end{aligned}$$

We define

$$\text{(A7)} \quad \begin{aligned} \beta_0 &= \alpha_0, \quad \beta_1 = \kappa_1\alpha_1, \quad \beta_2 = 1-(\kappa_1+\kappa_2)\,\alpha_1, \\ \beta_3 &= \kappa_2\alpha_1, \quad \beta_4 = \kappa_1\,\alpha_4 \end{aligned}$$

In order to arrive at an expression that is more easy to handle, we assume $\kappa_2\alpha_4 m_t = 0$. The assumption is justified for small values on κ_2, i.e. if the eleasticity with respect to electricity is low in the short-run cost function. The long-run solution for domestic prices thus describes bh_t as a function of pk_t, pw_t, pe_t and $z_t = \log(PW_t/PK_t)m_t$.

$$\text{(A8)} \qquad bh_t = \beta_0 + \beta_1 pw_t + \beta_2 pk_t + \beta_3 pe_t + \beta_4 z_t$$

Through the definitions of β_1, β_2 and β_3, $\beta_2 = 1-\beta_1-\beta_3$. Accordingly, (A8) is homogenous of degree one in PW, PK and PE.

Appendix 2: Data and definitions of variables

BH_t	Domestic price index of the commodity produced by Private mainland economy
PW_t	Variable unit costs inclusive of net sector taxes for Private mainland economy
PK_t	Price index of total Norwegian imports
PE_t	Domestic price index of electricity
I_t	Norwegian import excluding petroleum and shipping, real terms
Q_t	Value added for private mainland economy, real terms
UR_t	Unemployment rate, according to Labour Force Sample Survey (LFSS)
V_t	Exchange rate expressed as Norwegian currency per unit of foreign currency
$TRTN_t$	Average tax rate of households, according to National Account
$YTSX_t$	Net sector taxes for Private mainland economy
ds_t (s=1,2,3)	Centered seasonal dummies; ds_t = 1 if quarter = s, ds_t = -1 if quarter = 4, 0 otherwise
$dk77_t$	$dk77_t$=1 if t ≤ 77:4, 0 otherwise
$dk85_t$	$dk85_t$=1 if t ≤ 85:4, 0 otherwise
$d74q1_t$	$d74q1_t$=1 if t=74:1, 0 otherwise
$pstopin_t$	Dummy for price and wage regulations. Pstopin ∈ [0, 1]. Pstopin equals zero in periods with no regulations. Source: Bowitz and Cappelen (1994).
$pstopout_t$	Dummy for lagged effects of price and wage regulations. Pstopout ∈ [0, 1]. Pstopout equals zero in periods with no lagged effects. Source: Bowitz and Cappelen (1994).

Data are taken from the Quarterly National Account, published by Statistics Norway. All price indices equal 1 in 1991. Prices are given in Norwegian currency. Lower case denotes the natural logarithm of the variable while the symbol Δ denotes a differentiated variable.

Appendix 3: Unconstrained regressions of the forward looking model

Table 3.1: Forward looking ECM for Δbh_t. 2SLS[7]. 72:1 to 92:1

Regressor	Equation (a) Coeff.	N-W s.e.	Equation (b)[1] Coeff.	N-W s.e.
Constant	.20*	(.08)	.13**	(.04)
bh_{t-1}	-.28**	(.07)	-.26**	(.07)
pw_{t-1}	.18**	(.06)	.14**	(.04)[1]
pk_{t-1}	.11**	(.02)	.10**	(.02)[1]
pe_{t-1}	.01	(.02)	.02	(.01)[1]
z_{t-1}	.10*	(.05)	.14*	(.06)
Δpw_t	.08	(.06)	.04	(.04)
Δpw_{t+1}	.06	(.04)	.03	(.03)
Δpk_t	.09	(.08)	.07	(.08)
Δpk_{t+1}	.06	(.13)	.07	(.13)
Δpe_t	.02	(.05)	.04	(.04)
Δpe_{t+1}	.06	(.07)	.03	(.08)
d2	.01*	(.005)	.01*	(.004)
dk77*d3	-.02**	(.004)	-.02**	(.004)
dk85*d3	.02**	(.006)	.02**	(.007)
$El_{PW}BH$[2]	.51**	(.13)	.35**	(.07)[1]
$El_{PK}BH$[2]	.51**	(.07)	.56**	(.08)[1]
$El_{PE}BH$[2]	.03	(.07)	.09*	(.04)[1]
R^2	.78		.79	
100 SER	1054		1031	
DW	1.93		1.82	
$\chi^2_{SM}(p)$[3]	25.37	(p=23)	28.04	(p=23)
$\chi^2_{SC}(4)$	3.74		4.98	
$\chi^2_{RESET}(1)$	.03		.16	
$\chi^2_{N}(2)$	2.83		.87	
$\chi^2_{HET}(1)$	7.97**		7.94**	
$\chi^2_{WALD}(6)$[4]	5.97		3.88	
$\chi^2_{WALD}(1)$[5]	1.49		-	
$\chi^2_{WALD}(7)$[6]	6.42			

* (**): Significant at a 5% (1%) level.

1 Estimated subject to the restriction of static homogeneity.

2 Calculated for t=92:1.

3 Critical values 5% (1%) level: $\chi^2(1)$=3.84 (6.63), $\chi^2(2)$=5.99, $\chi^2(4)$=9.49, $\chi^2(6)$=12.59, $\chi^2(7)$=14.07, $\chi^2(23)$=35.17.

4 Test of backward-forward restrictions.

5 Test of: $El_{PW}BH+El_{PK}BH+El_{PE}BH = 1$ (static homogeneity).

6 Test of backward-forward restrictions and static homogeneity.

7 Additional instruments: Δpe_{t-1} Δpe_{t-2} Δp_{t-3} Δpw_{t-2} Δpw_{t-3} $\Delta ytsx_{t-1}$ $\Delta ytsx_{t-2}$ Δpk_{t-1} Δpk_{t-2} Δpk_{t-3} Δpk_{t-4} $trtn_{t-1}$ $\Delta trtn_{t-1}$ $\Delta trtn_{t-2}$ UR_{t-1} ΔUR_{t-3} v_{t-1} Δv_{t-1} Δv_{t-2} Δv_{t-3} d1 d2 d3 dk77 dk85 d74q1 pstopin pstopout.

Appendix 4: Recursive coefficients (figure a-k), backward looking ECM

Figure a: Coef. of Constant and its 2 S.E. bands based on recursive OLS
Backward looking ECM for domestic prices.

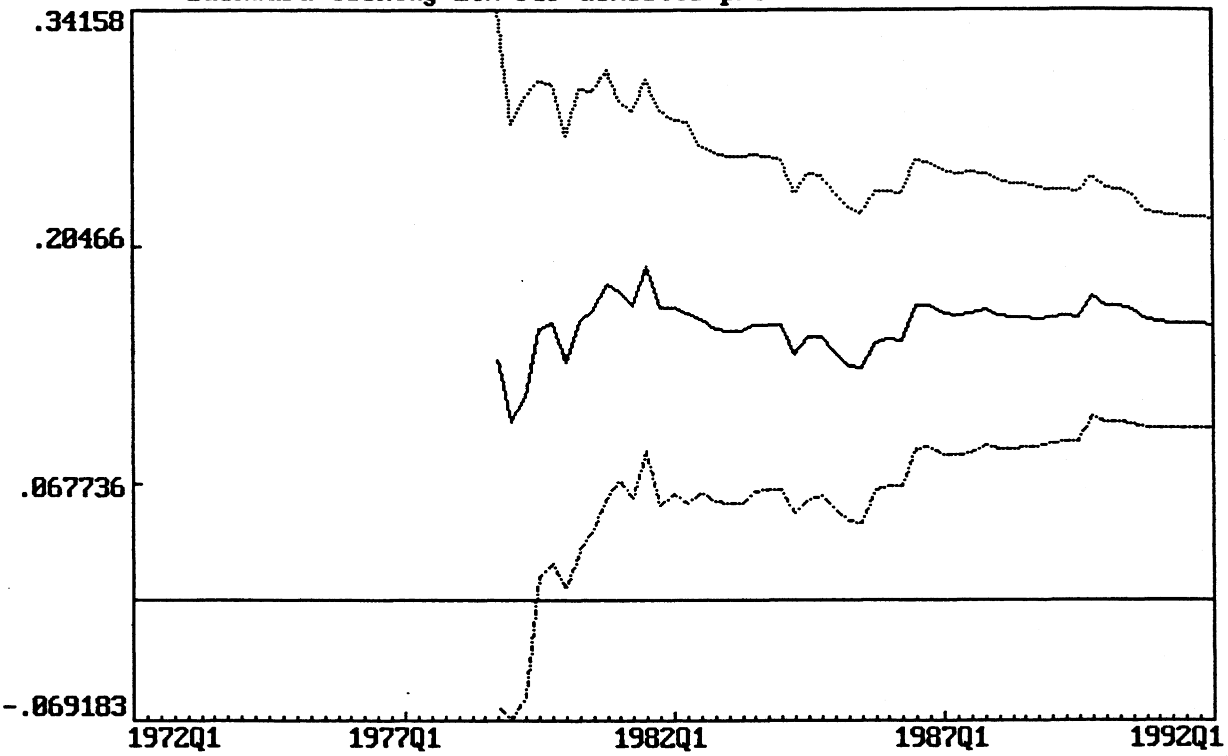


Figure b: Coef. of (bh-pk)(-1) and its 2 S.E. bands based on recursive OLS
Backward looking ECM of domestic prices.

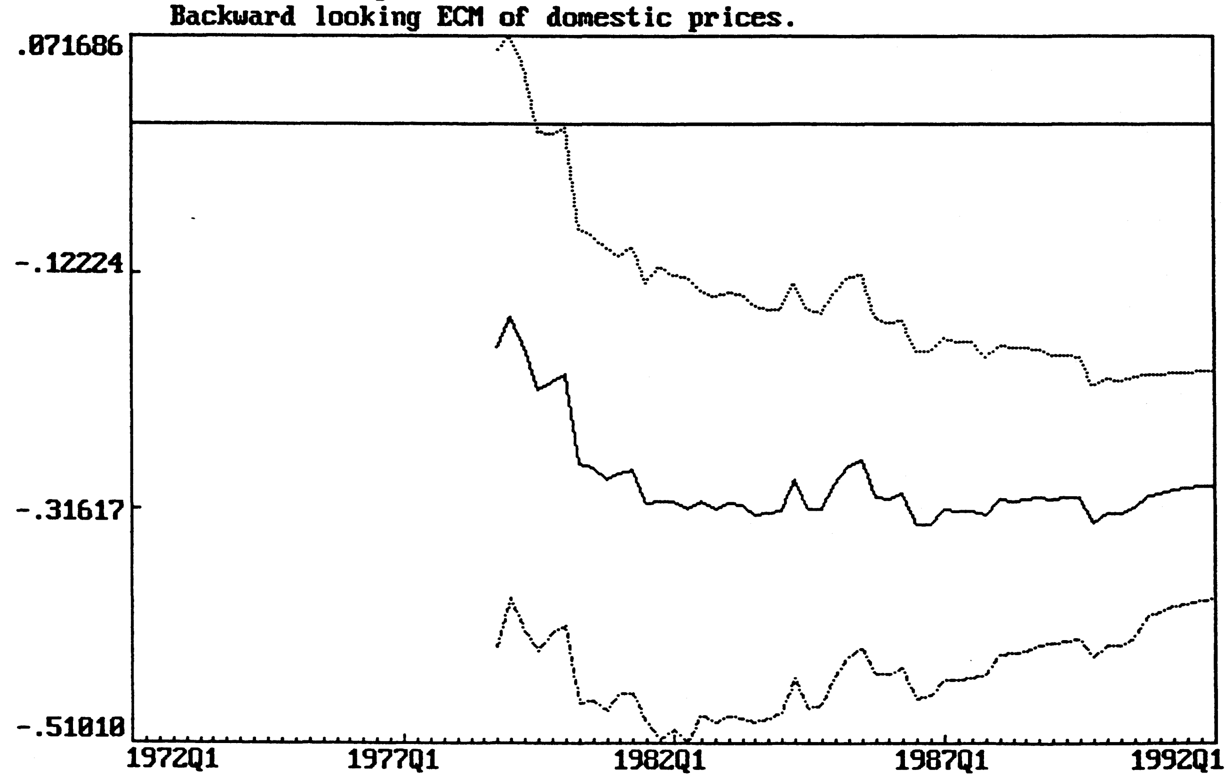


Figure c: Coef. of (pw-pk)(-1) and its 2 S.E. bands based on recursive OLS
Backward looking ECM of domestic prices.

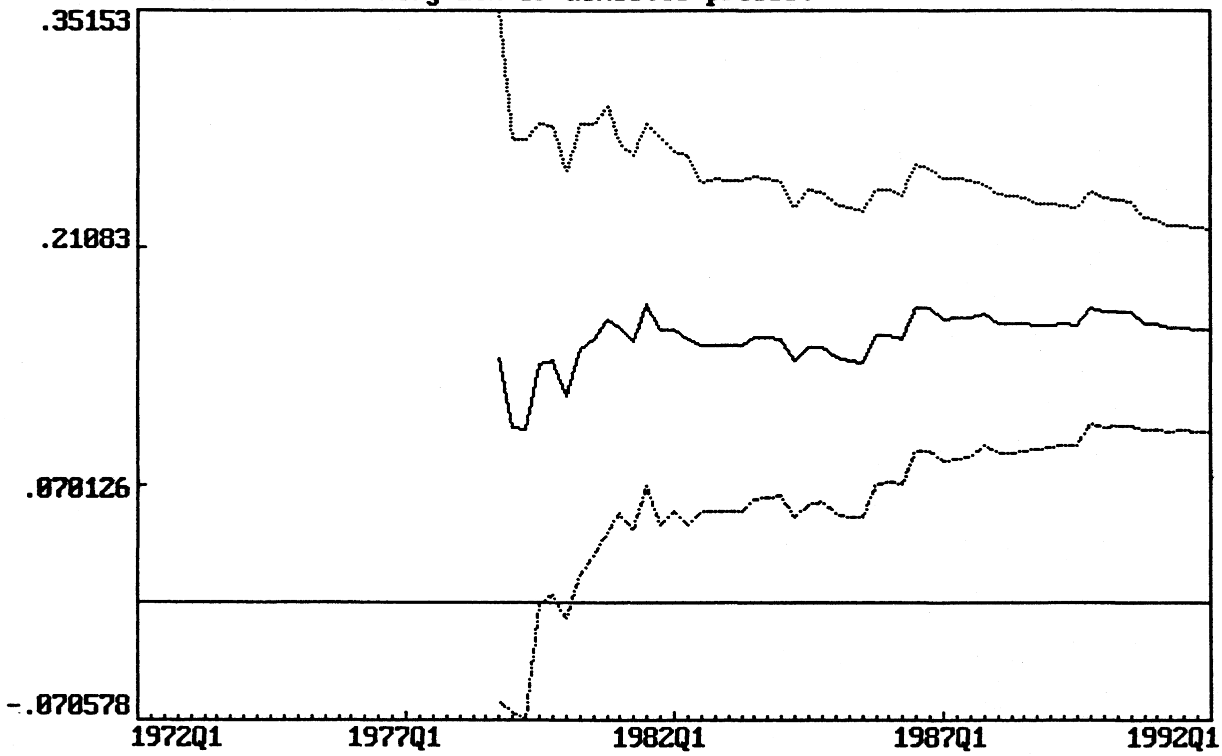


Figure d: Coef. of (pe-pk)(-1) and its 2 S.E. bands based on recursive OLS
Backward looking ECM of domestic prices.

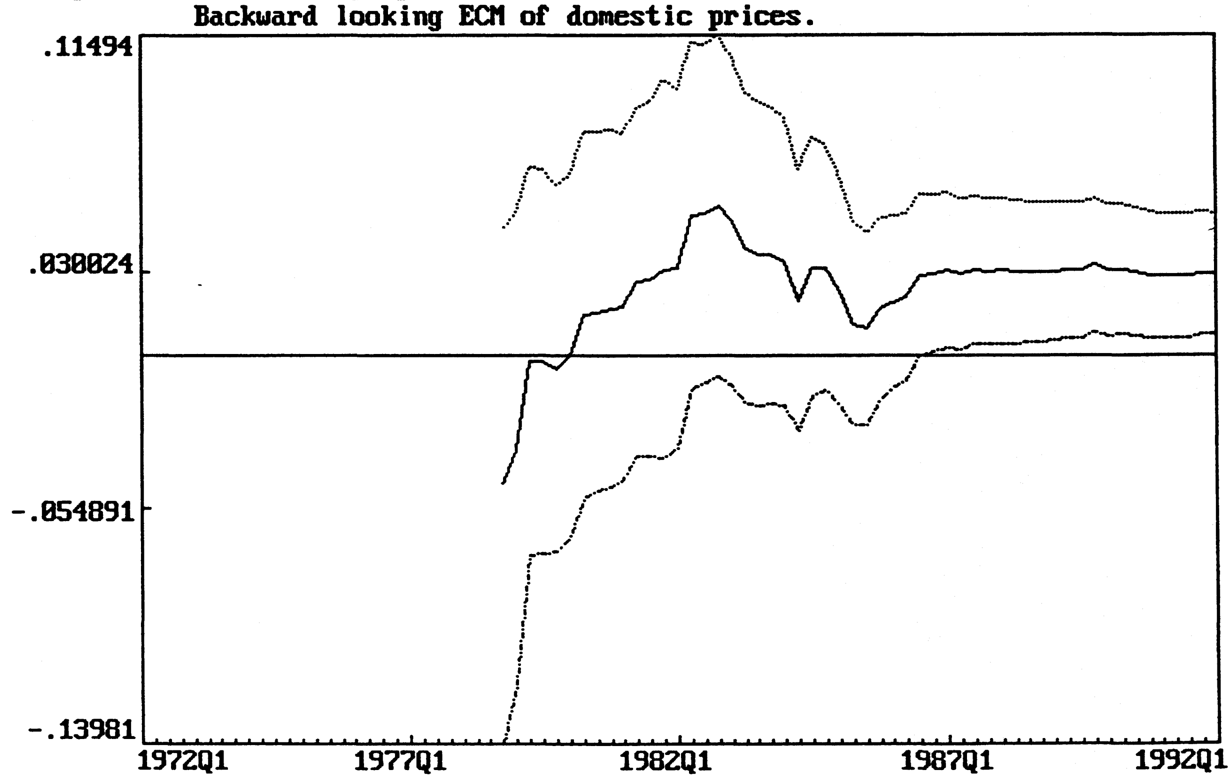


Figure e: Coef. of z(-1) and its 2 S.E. bands based on recursive OLS.
Backward looking ECM of domestic prices.

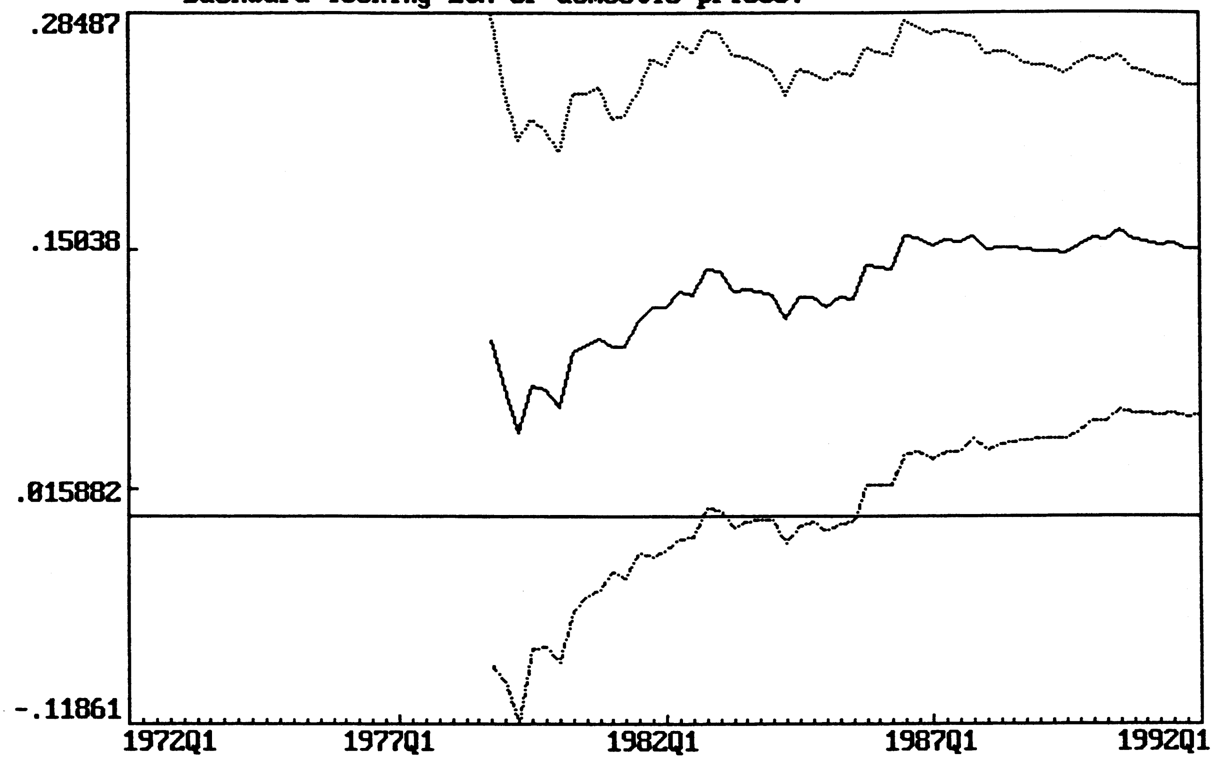


Figure f: Coef. of dbh(-3) and its 2 S.E. bands based on recursive OLS.
Backward looking ECM of domestic prices.

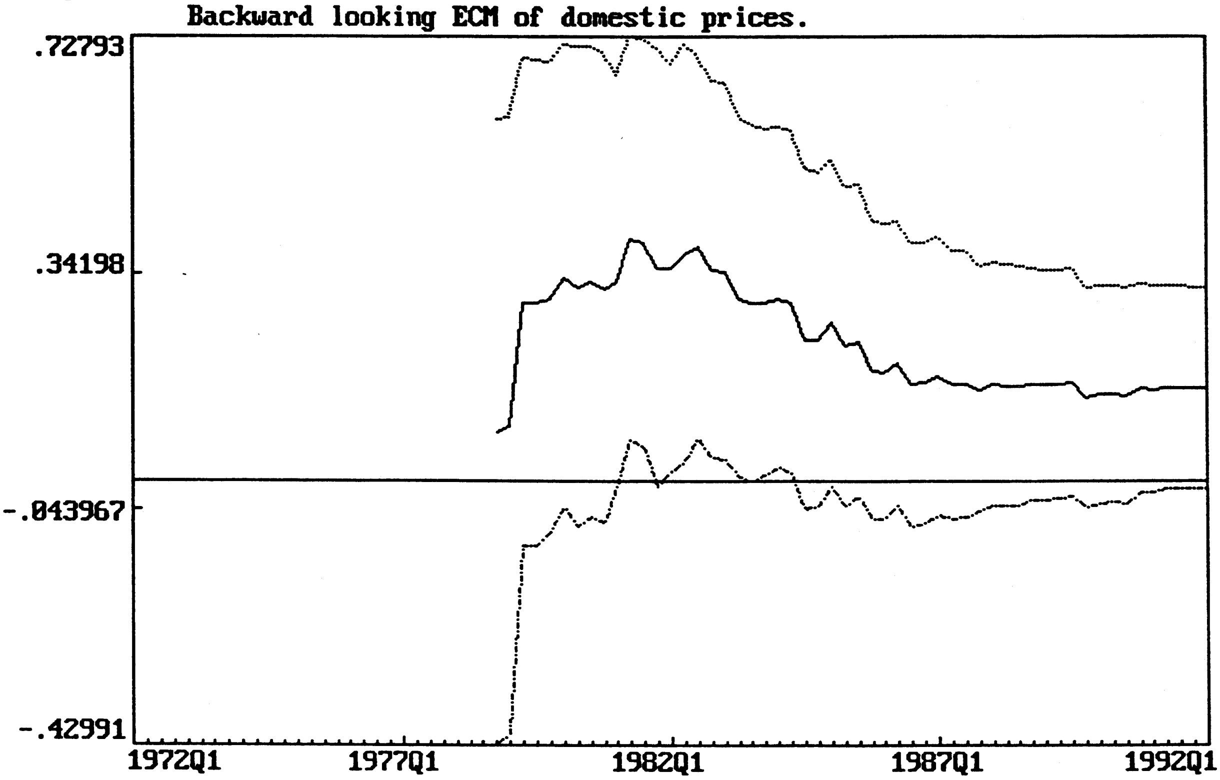


Figure g: Coef. of dpe and its 2 S.E. bands based on recursive OLS.
Backward looking ECM of domestic prices.

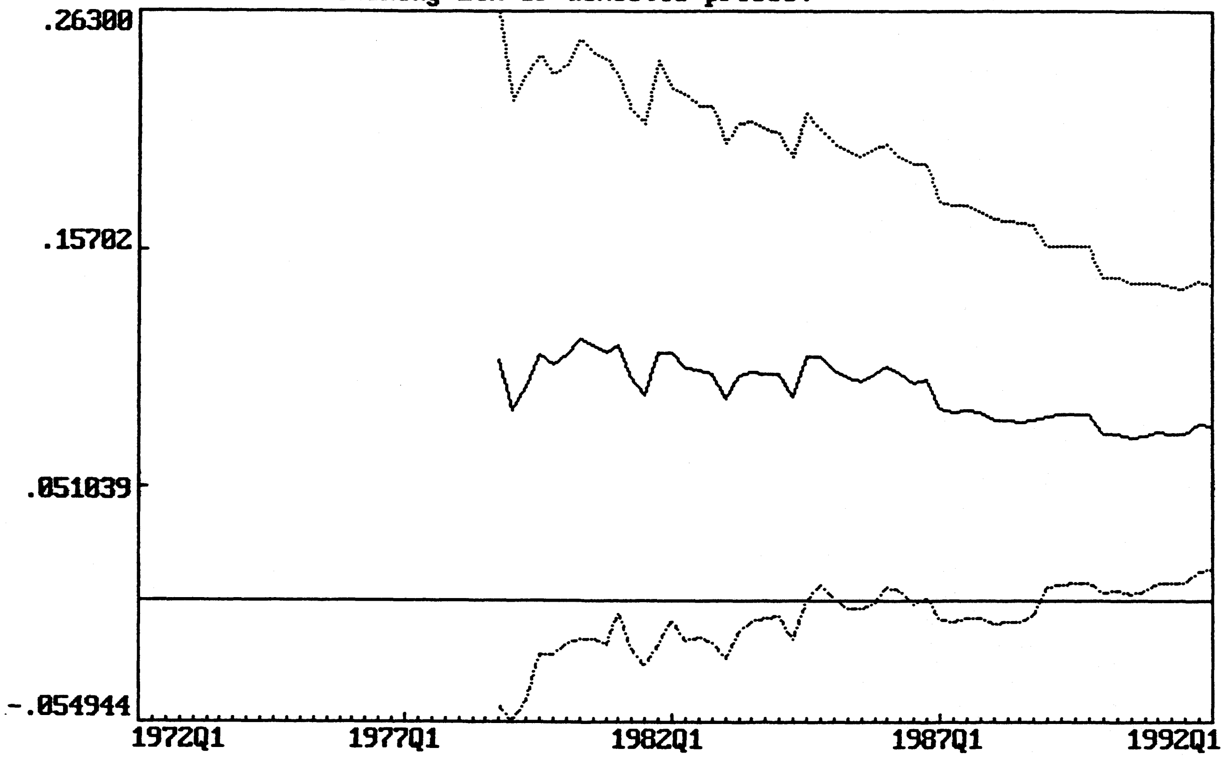


Figure h: Coef. of dpe(-3) and its 2 S.E. bands based on recursive OLS.
Backward looking ECM of domestic prices.

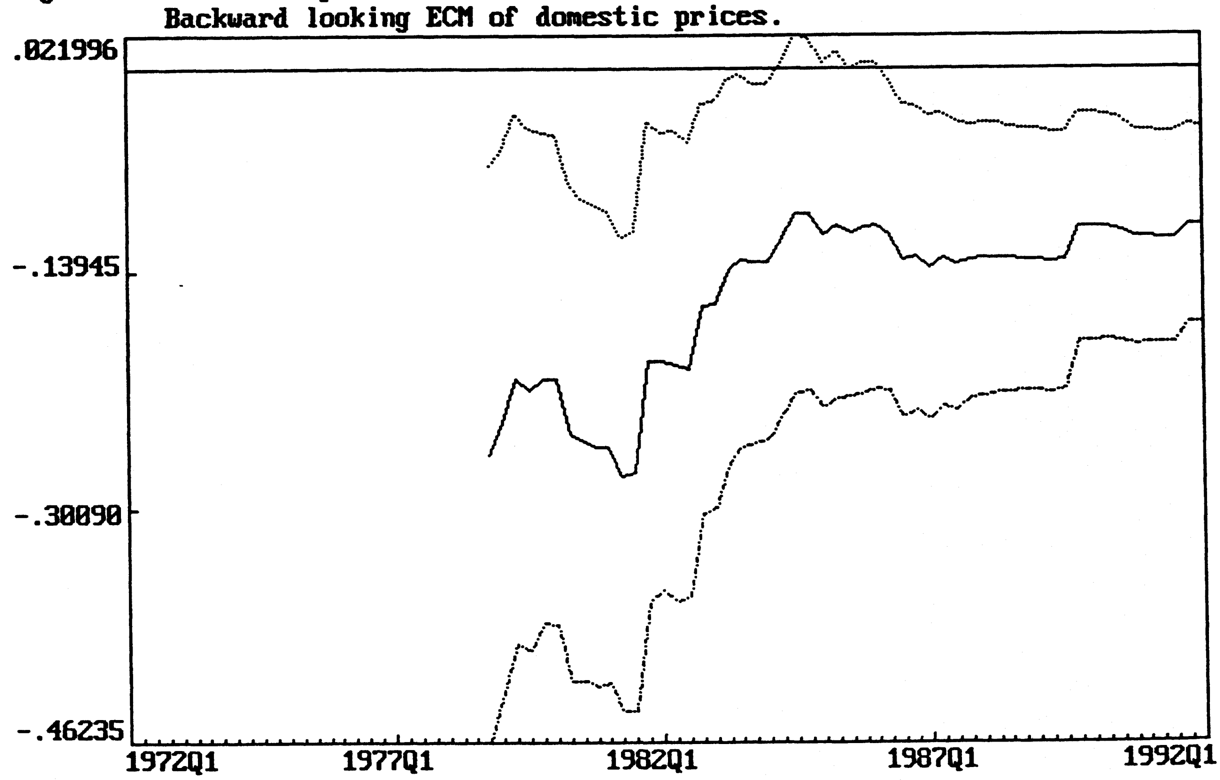


Figure i: Coef. of d1 and its 2 S.E. bands based on recursive OLS.
Backward looking ECM of domestic prices.

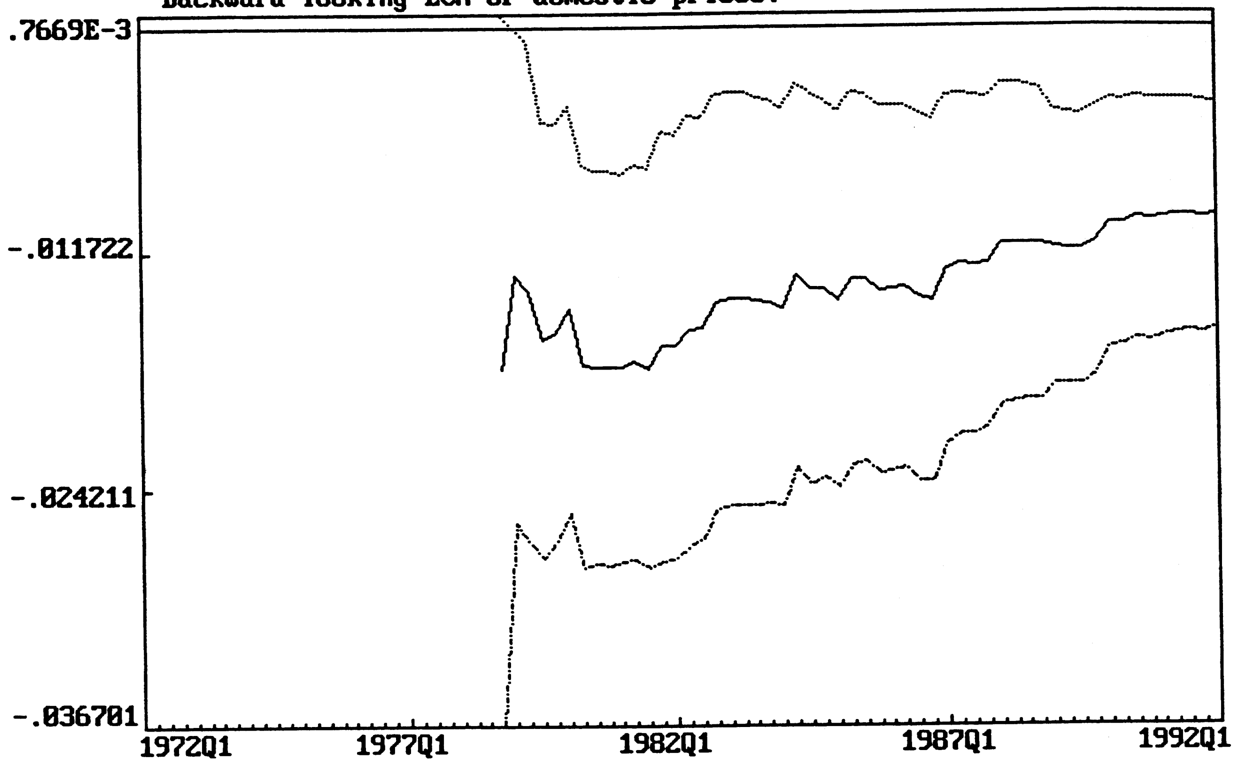


Figure j: Coef. of dk77*d3 and its 2 S.E. bands based on recursive OLS.
Backward looking ECM of domestic prices.

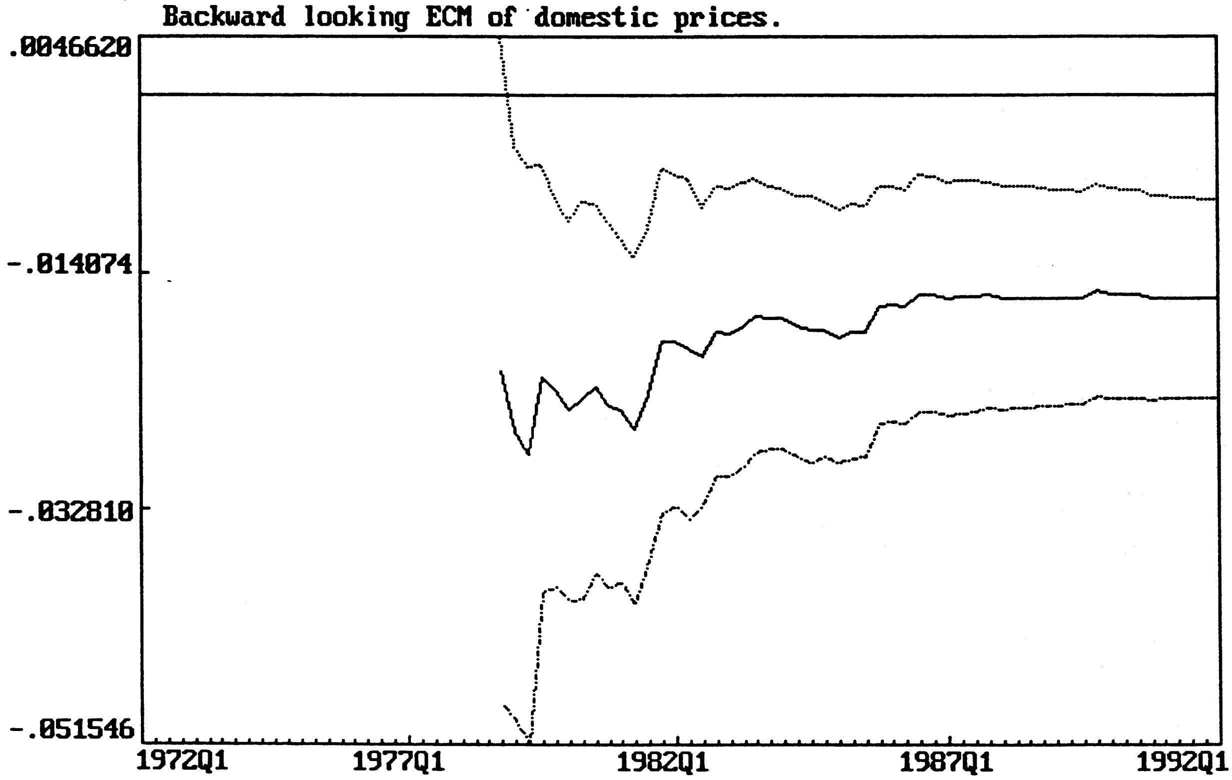
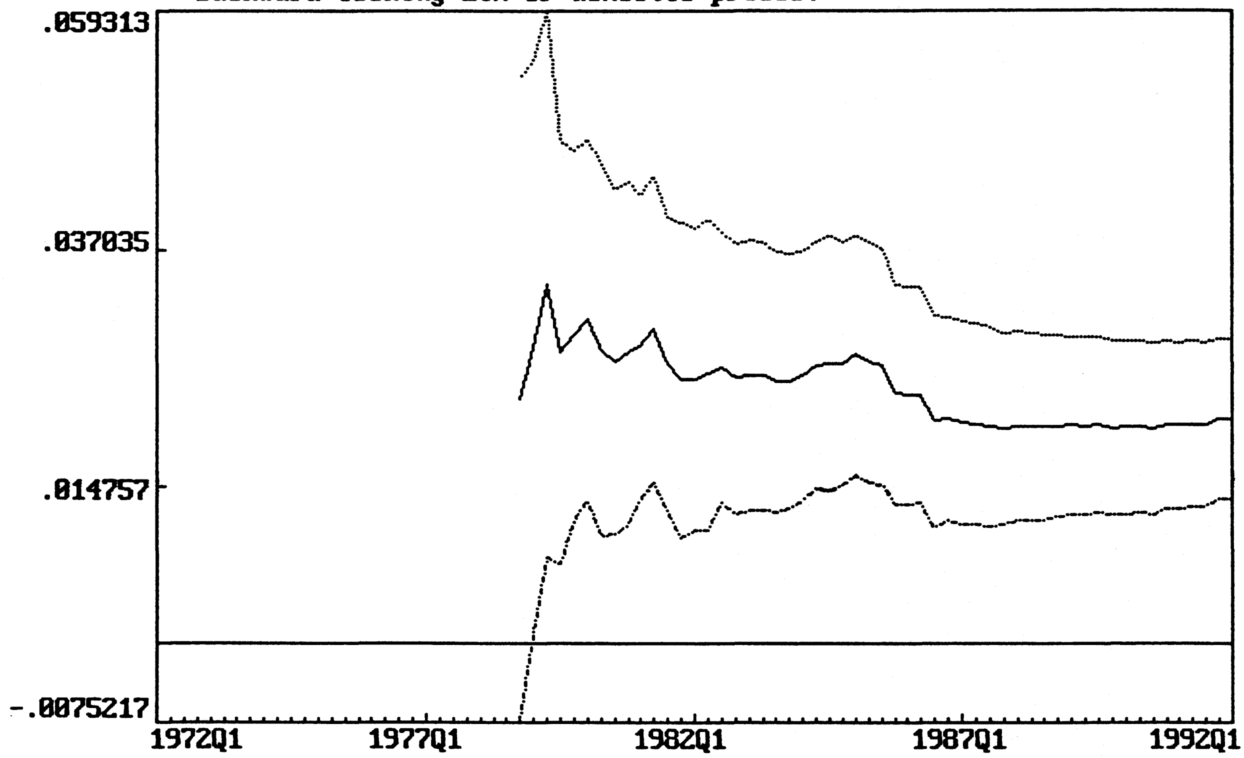


Figure k: Coef. of dk85*d3 and its 2 S.E. bands based on recursive OLS.
Backward looking ECM of domestic prices.



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