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The new Keynesian Phillips curve: Does it fit Norwegian data?

Abstract:

We evaluate the empirical performance of the new Keynesian Phillips curve (NKPC) for a small open economy using cointegrated vector autoregressive models, likelihood based methods and general method of moments. Our results indicate that both baseline and hybrid versions of the NKPC as well as *exact* and *inexact* formulations of the rational expectation hypothesis are most likely at odds with Norwegian data. By way of contrast, we establish a well-specified dynamic backward-looking imperfect competition model (ICM), a model which encompasses the NKPC in-sample with a major monetary policy regime shift from exchange rate targeting to inflation targeting. We also demonstrate that the ICM model forecasts well both post-sample and during the recent financial crisis. Our findings suggest that taking account of forward-looking behaviour when modelling consumer price inflation is unnecessary to arrive at a well-specified model by econometric criteria.

Keywords: The new Keynesian Phillips curve, imperfect competition model, cointegrated vector autoregressive models (CVAR), equilibrium correction models, likelihood based methods and general method of moments (GMM).

JEL classification: C51, C52, E31, F31

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Sammendrag

Vi undersøker hvor godt den nykeynesianske Phillipskurven for en liten åpen økonomi passer til norske forhold. Modellen tilsier at løpende konsumprisvekst avhenger av forventet konsumprisvekst i neste periode og avvik mellom nivået på konsumprisene og prisene på arbeidskraft og import av varer og tjenester. Våre funn indikerer at modellen ikke får støtte med norske data når vi legger standard statistiske kriterier til grunn. Vi finner isteden støtte for en konkurrerende modell som inneholder betydelige effekter av konsumprisveksten i tidligere perioder i tillegg til effekter fra avviket mellom nivået på konsumprisene og prisene på arbeidskraft og import av varer og tjenester. Den konkurrerende modellen forklarer konsumprisveksten rimelig godt i estimeringsperioden, en periode som inkluderer overgangen fra valutakursstyring til inflasjonsstyring i pengepolitikken. Samtidig er den konkurrerende modellen i stand til å prognostisere konsumprisveksten nokså godt etter estimeringsperioden og gjennom finanskrisen som var en ganske turbulent periode for norsk økonomi. Våre funn indikerer at framoverskuende forventninger i prissettingen ikke synes å være viktig som forklaringsvariabel ved modellering av den norske konsumprisveksten.

1 Introduction

Within the new Keynesian economics paradigm, based on rational expectations, optimising agents and imperfect competition in markets for goods, the new Keynesian Phillips curve (henceforth NKPC) is the workhorse of understanding inflation dynamics. The baseline NKPC explains current inflation by expected future inflation and real marginal costs as the forcing variable, whereas the hybrid version of the model also includes lagged inflation terms as a way of modelling "rule of thumb" or backward-looking price setters. Since the influential papers by Galí and Gertler (1999) and Galí *et al.* (2001), who claim strong evidence in favour of the NKPC using European and US post-war data, a great number of studies have tested the empirical validity of the model based on data of both closed and open economies, see e.g. Boug *et al.* (2010), Tillmann (2009), Juillard *et al.* (2008), Bjørnstad and Nymoen (2008), Gogley and Sbordone (2008), Fanelli (2008), Kurmann (2007), Batini *et al.* (2005), Bårdsen *et al.* (2004, 2005) and Kara and Nelson (2003) among others. The studies differ with respect to data used, sample period studied and econometric methods applied, and the supportive evidence on the NKPC is rather mixed.

The open economy new Keynesian Phillips curve (henceforth OE-NKPC) differs from its closed economy counterpart in that the exchange rate and prices on imported goods matter in some way or another for domestic inflation. In the empirical OE-NKPC literature two approaches have basically been undertaken to model the hypothesised connection between the exchange rate and domestic inflation. The first approach involves treating imported goods as final consumer goods and hence introducing import prices directly into the definition of consumer prices. In so doing, the real exchange rate becomes an explanatory variable in addition to expected future inflation and real marginal costs in the NKPC model, see Bjørnstad and Nymoen (2008), Guender and Xie (2007), Guender (2006), Galí and Monacelli (2005) and Giordani (2004) for examples. The second approach involves introducing imported goods as intermediate inputs which, together with labour, produce final consumer goods. Accordingly, real marginal costs in production becomes a function of relative prices of imported inputs, see e.g. Batini *et al.* (2005), Kara and Nelson (2003) and McCallum and Nelson (1999).[1]

The two approaches have mainly been investigated from a theoretical perspective in the literature. Among the relatively few existing empirical studies, Bjørnstad and Nymoen (2008) introduce open economy features by means of the first approach and conclude that the OE-NKPC is most likely at odds with a panel data set of twenty OECD countries (including Norway). Guender and Xie (2007) also argue, us-

[1] Smets and Wouters (2002) combine the two approaches in their theoretical NKPC model by introducing imported goods as both intermediate inputs as well as final consumer goods. Svensson (2000) discusses the channels through which the exchange rate is likely to affect consumer prices in the NKPC model for small open economies.

ing the first approach, that the OE-NKPC does not receive much empirical support based on a data set of six open economies. Kara and Nelson (2003), on the other hand, apply the second approach and claim that the OE-NKPC matches reasonably well with UK data. Batini *et al.* (2005) also derive versions of the OE-NKPC with accommodating results on UK data. Although imports are theoretically modelled as intermediate inputs in that study, the empirical models are more in line with imports being modelled as final goods as prices of *total* imports (and not prices of imported material inputs) are used among the explanatory variables. Nevertheless, the rather mixed results with respect to the empirical status of the OE-NKPC call for further research.

In this paper, we evaluate by means of the first approach the empirical performance of the OE-NKPC for Norway, a small open economy where international trade plays an important role in the exchange of goods and services. Hence, the exchange rate through import prices is likely to be relevant in the determination of domestic inflation. We derive an OE-NKPC for Norwegian inflation based on the forward-looking linear quadratic adjustment cost model of Rotemberg (1982) and the theoretical principles of the imperfect competition model (henceforth ICM) for a small open economy. Our OE-NKPC thereby relates current inflation to expected future inflation and the difference between the actual price and the price target in levels as a theory consistent forcing variable, the latter hypothesised to be based on a weigthed average of unit labour costs and prices of total imports. We contribute to the existing OE-NKPC literature by focusing on both baseline and hybrid models as well as on *exact* formulations in the sense of Hansen and Sargent (1991) and *inexact* formulations in which a stochastic term is included in the models. To this end, we employ the likelihood based testing procedures suggested by Johansen and Swensen (1999, 2004, 2008) and the commonly used GMM procedure when evaluating the *exact* and the *inexact* versions of our OE-NKPC, respectively. Rather than using an arbitrary instrument set, which is often the case in related studies, we let the instruments used within the GMM framework be dictated from fitted VAR models underlying the testing of the *exact* OE-NKPC. We are consequently able to shed some light on the importance of introducing a stochastic error term to the empirical models, an often neglected econometric issue in the literature. Unlike many related studies, we pay particular attention to time series properties, and possibly cointegrated nature, of variables involved by means of fitted VAR models and likelihood based inference.

Our empirical investigation produces several noteworthy findings. First, we establish a well-specified empirical counterpart to the theory-consistent forcing variable. Accordingly, the hypothesised link between domestic inflation and the exchange rate through import prices in our OE-NKPC is supported using Norwegian data. These findings are also in line with existing models of inflation based on Norwegian data. Second, and by way of contrast, we demonstrate that both baseline and hybrid versions as well as *exact* and *inexact* formulations of our OE-NKPC are most likely at odds with the data. Hence, the forward-looking part of our OE-NKPC is rejected, whereas the part of the model containing the deviation of the price level from its target is not. Finally, we establish a well-specified dynamic ICM model of inflation with backward-looking elements only, a model which encompasses the OE-NKPC in-sample with a major monetary policy regime shift from exchange rate targeting to inflation targeting. The regime robustness of the dynamic ICM model is inconsistent with the Lucas-critique being quantitatively important in our case. Also, we find that the dynamic ICM model forecasts well post-sample and during the financial crisis in 2008 and 2009 when the exchange rate fluctuated considerably.

The rest of the paper is organised as follows: Section 2 outlines our OE-NKPC, Section 3 describes the data, Section 4 reports findings from the cointegration analysis, Section 5 reports the various tests of the OE-NKPC and Section 6 develops a dynamic backward-looking ICM model as a competing model of inflation and conducts a forecasting exercise on that model. Section 7 concludes.

2 An open economy NKPC model

As explained by Roberts (1995), there are several routes from a theoretical set up of firm's pricing behaviour that lead to the NKPC model, including the forward-looking linear quadratic adjustment cost model of Rotemberg (1982) and the models of staggered contracts developed by Taylor (1979, 1980) and Calvo (1983). We assume that the representative firm, based on Rotemberg (1982), chooses a sequence of prices (P_{t+j}) to minimise the loss function

$$E_t\left[\sum_{j=1}^{\infty}\delta^j\left[\lambda(p_{t+j}-p^*_{t+j})^2+(p_{t+j}-p_{t+j-1})^2\right]\right], \tag{1}$$

where E_t denotes the conditional expectation given the information contained in the information set at time t and lower case letters indicate natural logarithms of a variable, i.e., $p_{t+j}=\ln(P_{t+j})$.[2] The variable p^*_t is the price target or the static equilibrium price, whereas δ represents the discount rate and λ the relative cost parameter of the two terms of the loss function. Hence, firms determine a sequence of prices so as to minimise the expected present discounted value of the sum of all future squared deviations from the target and squared changes in the price itself. Because changes in the price will be penalised, immediate adjustment towards the target will be non-optimal unless λ is large. The first order condition of this minimisation problem gives the Euler equation

$$\Delta p_t=\delta E_t\Delta p_{t+1}-\lambda(p_t-p^*_t), \tag{2}$$

[2] Throughout the paper, lower case letters denote natural logarithms of the corresponding upper case variables.

where the first difference of p_t, $\Delta p_t = p_t - p_{t-1}$, defines current inflation, while $E_t \Delta p_{t+1}$ is expected inflation one period ahead conditional upon information available at time t.

We now show how to introduce, from theoretical principles, open economy features into (2). Building on existing ICM models of inflation for Norway, see e.g. Bårdsen *et al.* (2005, ch. 8.7), we assume that the representative firm operates in imperfectly competitive markets facing regular downward sloping demand curves. Profit maximisation then implies that prices are set as a mark-up over marginal costs. Assuming a value added Cobb-Douglas production function in labour and capital with capital as a quasi-fixed factor, unit labour costs are proportional to marginal costs. We follow Bjørnstad and Nymoen (2008) and Batini *et al.* (2005) among others and let the mark-up depend on relative prices such that

$$pd_t = m_0 - m_1(pd_t - pi_t) + ulc_t, \tag{3}$$

where pd_t denotes the price level on domestic consumer goods and services, pi_t is the price level on imported consumer goods and services, ulc_t denotes unit labour costs and $0 \leq m_1 \leq 1$ and reflects conditions on the demand side of the product markets. It follows from (3) that an increase in the competing price allows the domestic producer to increase her mark-up over unit labour costs. We then let $(1 - \kappa)$ denote the constant import share and define the aggregate consumer price level by the commonly used identity of the form [see e.g. Svensson (2000), Galí and Monacelli (2005) and Bjørnstad and Nymoen (2008)]

$$p_t \equiv \kappa pd_t + (1 - \kappa)pi_t. \tag{4}$$

By using (4), we can solve for the producer's price target to obtain

$$p_t = \gamma_0 + \gamma_1 ulc_t + \gamma_2 pi_t, \tag{5}$$

where $\gamma_0 = m_0\gamma_1$, $\gamma_1 = \kappa/(1 + m_1)$ and $\gamma_2 = (1 - \gamma_1)$. We notice that (5) is homogeneous of degree one in competing prices and unit labour costs. Although (5) is derived from the ICM model, it also contains the law of one price or perfect competition for homogenous goods as a special case. In the latter case m_1 approaches infinity, such that the domestic price is equal to the cometing price, i.e., $pd_t = pi_t$. Intuitively, this is reasonable because the closer substitutes the products are the smaller is the market power of each producer and accordingly also the mark-up. Our specification of the mark-up allows for a general model to be tested, with a constant mark-up as a special case.[3] Equation (5) is a static model of the price

[3] In the closed economy NKPC literature it is common to assume that producers face isoelastic demand curves so that the mark-up is a constant, see e.g. Galí and Gertler (1999) and Galí *et al.* (2001).

target so that the right hand side of (5) is equivalent to p_t^* in (2). Inserting (5) in (2) yields the baseline OE-NKPC model in our context

$$\Delta p_t = \delta E_t \Delta p_{t+1} - \lambda eqcm_t, \tag{6}$$

where $eqcm_t = p_t - \gamma_0 - \gamma_1 ulc_t - \gamma_2 pi_t$. Before looking at the hybrid version of (6), we point out that the dynamic backward-looking ICM model, without any forward-lokking term, is a special case of the OE-NKPC. To see this, we may reparameterise (6) such that

$$\Delta p_t = \eta_1 E_t \Delta p_{t+1} + \eta_2 \Delta ulc_t + \eta_3 \Delta pi_t - \eta_4 eqcm_{t-1}, \tag{7}$$

where $\eta_1 = \delta/(1+\lambda)$, $\eta_2 = \lambda\gamma_1/(1+\lambda)$, $\eta_3 = \lambda\gamma_2/(1+\lambda)$ and $\eta_4 = \lambda/(1+\lambda)$. The dynamic backward-looking ICM model is therefore nested by and is a special case of the OE-NKPC when $\eta_1 = 0$. If the hypothesis $\eta_1 = 0$ cannot be rejected the OE-NKPC is said to be parsimoniously encompassed by the dynamic backward-looking ICM model in the terminology of Hendry (1995).[4] Such a test outcome is further inconsistent with the main assumption of the NKPC that a significant proportion of price setters are forward looking in accordance with the rational expectation hypothesis.

Inspired by the influencial studies by Galí and Gertler (1999) and Galí, Gertler and López-Salido (2001), we specify the following hybrid version of (6) with both forward-looking and backward-looking elements included:

$$\Delta p_t = \gamma_f E_t \Delta p_{t+1} + \gamma_b \Delta p_{t-1} - \lambda_h eqcm_t. \tag{8}$$

We see that (8) reduces to (6) when $\gamma_b = 0$. Accordingly, if the baseline OE-NKPC is valid, then $\gamma_f = \delta$, $\gamma_b = 0$ and $\lambda_h = \lambda$. Generally, the parameter spaces $0 \leq \delta \leq 1$, $0 \leq \lambda$ and $0 \leq \gamma_f, \gamma_b \leq 1$, $0 \leq \lambda_h$ are required to provide an admissible economic interpretation of an estimated baseline and hybrid OE-NKPC, respectively. If we regard inflation as a stationary process, the deviation of the price level from its target value must also be a stationary process in order for (6) and (8) to be balanced equations. We notice that the expression $eqcm_t$ in (6) and (8) is a theory-consistent driving variable that may form a cointegration relationship with testable restrictions.

The open economy NKPC in Batini *et al.* (2005) is consistent with and has the same form and interpretation as (6), see Appendix 1 for details. However, as opposed to Batini *et al.* (2005), we pay particular attention to the time series properties, and possibly cointegrated nature, of variables involved. We test the empirical relevance of (6), (7) and (8) based on Norwegian data, cointegration techniques, likelihood based methods and GMM.

[4] Generally speaking, parsimonious encompassing requires a small model to explain the results of a larger model within which it is nested, cf. Hendry (1995, p. 511).

3 Data

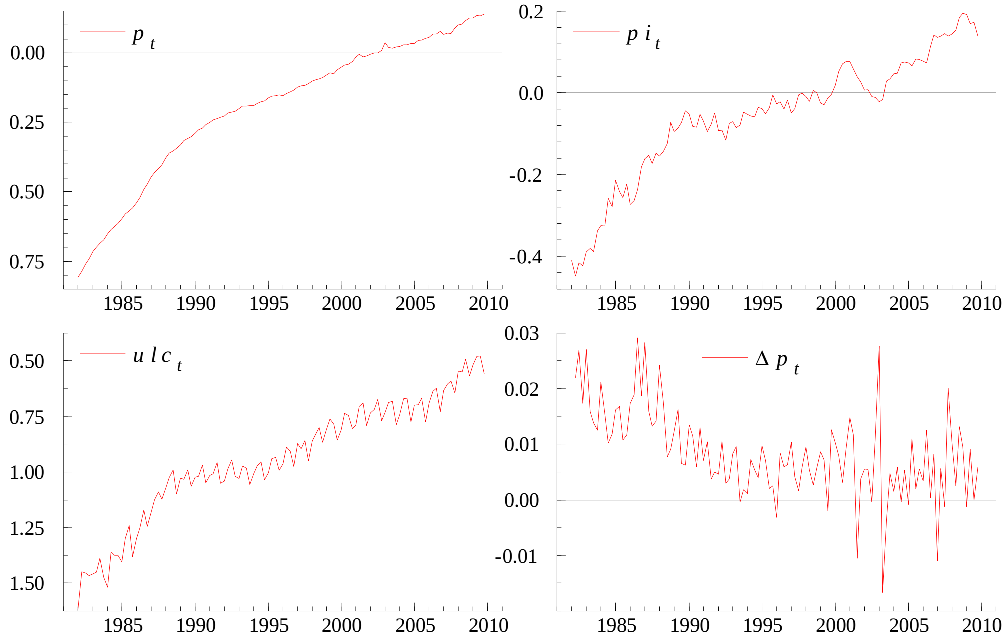
The empirical analysis is based on quarterly, seasonally unadjusted data that spans the period $1982Q1 - 2009Q4$, of which data from the period $1982Q1 - 2005Q4$ and $2006Q1 - 2009Q4$ are used for estimation and *out-of-sample* forecasting, respectively. Mavroeidis (2004) concludes that the estimates of the NKPC are less reliable when the sample covers periods in which inflation has been under effective policy control. The starting point of our estimation period is thus motivated by the fact that the 1970s and the early 1980s were characterised by massive governmental price controls. If the expectational term in the OE-NKPC relationship is the most important influences on the correlation between exchange rate movements and inflation, then we would expect the relationship to depend closely on monetary policy regime in force. We explored this hypothesis by ending the estimation period in $2001Q1$ rather than in $2005Q4$ as monetary policy changed fundamentally from exchange rate targeting to inflation targeting in late March 2001.[5] It turned out, however, that the estimates of both the OE-NKPC and the dynamic backward-looking ICM model are virtually unchanged irrespective of which of the two ending points is used. These findings suggest that the Lucas critique lacks force in our empirical case. We extend the estimation sample by sixteen quarters for *out-of-sample* forecasting to shed light on any change in the link between exchange rates movements and domestic inflation following the financial crisis in 2008 and 2009.

In line with Bårdsen *et al.* (2005), we measure quarterly inflation by the official consumer price index rather than by the GDP deflator normally used in the NKPC literature. The actual prices that agents in the economy set are on gross output and not on value added. Deflators based on value added are typically residuals in the national accounts, in particular those following the principle of double-deflating. Hence, the GDP deflator is less related to the micro price setting behaviour than other concept within the national account. Thus, we argue that the consumer price index is a more relevant price series for evaluating the OE-NKPC for Norway than the GDP deflator. We employ the deflator for total imports as a proxy for the price level on imported consumer goods and services, whereas total labour costs relative to value added in the private mainland economy serves as a proxy for unit labour costs, see Appendix 2 for details. Figure 1 displays the log of the consumer price index (p_t), the log of the import prices (pi_t) and the log of unit labour costs (ulc_t), together with the inflation rates (Δp_t) over the sample period.

We notice that the consumer price inflation shows rather large changes in the quarters $1986Q3$, $2001Q3$, $2003Q1$ and $2003Q2$ during the estimation period. These changes are most likely associated with the 12 per cent devaluation of the Norwegian currency in May 1986, the drop in the VAT rate on food from 24 per cent to 12 per cent in July 2001 and the substantial increase and decrease of the electricity prices during the first and second quarter of 2003, respectively. That

[5] See Boug *et al.* (2006) for details.

Figure 1: Time series for p_t, pi_t, ulc_t and Δp_t



inflation increased considerably in the third and not in the second quarter of 1986 is due to delayed pass-through from exchange rate changes to import prices and consumer prices. Fluctuations in electricity prices are to a large extent related to natural causes (e.g. temperature) and not much to immediate economic phenomena as electricity is mainly based on hydropower. We control for the mentioned episodes in the empirical analysis by impulse dummies labelled $D86Q3$, $D01Q3$, $D03Q1$ and $D03Q2$. The consumer price inflation also shows some huge fluctuations during the forecasting period, especially in the quarter $2007Q1$ and in the years 2008 and 2009 which are likely to be attributed to the considerable fall in electricity prices and the large movements in the exchange rate during the recent financial crisis, respectively. We further notice that the time series exhibit a clear upward trend, but with no apparent mean reverting property, suggesting p_t, pi_t and ulc_t are nonstationary $I(1)$ series. Therefore, a reduced rank VAR is a candidate as an empirical model. However, the time series may exhibit a quadratic trend such that the time series are $I(2)$ rather than $I(1)$ over the sample period. We investigate both alternatives in the cointegration analysis below.

4 Cointegration analysis

We adapt the cointegration rank test suggested by Johansen (1995, p. 167) to find an empirical counterpart of (5). The point of departure of the $I(1)$ analysis and the tests that follow is an equilibrium correction representation of a p-dimensional VAR of order k written as

$$(9) \qquad \Delta X_t = \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \Pi X_{t-1} + \Phi_0 D_t + \Phi_1 + \Phi_2 t + \varepsilon_t, t = k+1, \ldots, T,$$

where $X_t = (p_t, ulc_t, pi_t)'$, D_t includes seasonal dummies labelled SD_{it} $(i = 1, 2, 3)$ and the impulse dummies $D86Q3$, $D01Q3$, $D03Q1$ and $D03Q2$ as described above, t is a linear deterministic trend and $\varepsilon_{k+1}, \ldots, \varepsilon_T$ are independent Gaussian variables with expectation zero and (unrestricted) covariance matrix Ω. The initial observations of $X_1, \ldots, X_k$ are kept fixed. We follow common practice and restrict the linear trend to lie in the cointegrating space, whereas the deterministic components D_t and Φ_1 are kept unrestricted in (9). If X_t is $I(1)$, presence of cointegration implies $0 < r < p$, where r denotes the rank or the number of cointegrating vectors of the impact matrix Π. The null hypothesis of r cointegrating vectors may be formulated as H_0: $\Pi = \alpha\beta'$, where α and β are p x r matrices, $\beta' X_t$ comprises r cointegrating $I(0)$ linear combinations and α contains the adjustment coefficients.

We find that $k = 3$ is the appropriate choice of lag length to arrive at a model with no serious misspecification in the residuals.[6] Table 1 reports the findings from applying the cointegration rank test to the data based on the VAR of order three.

Table 1: Tests for cointegration rank

r	λ_i	λ_{trace}	λ^a_{trace}
$r = 0$	0.262	47.22 [0.016]*	42.65 [0.052]
$r \leq 1$	0.136	18.95 [0.290]	17.12 [0.414]
$r \leq 2$	0.056	5.36 [0.555]	4.84 [0.626]

Notes: r denotes the cointegration rank and λ_i are the eigenvalues from the reduced rank regression, see Johansen (1995). The λ_{trace} and λ^a_{trace} are the trace statistics without and with degress of freedom adjustments, respectively. The p-values in square brackets, which are reported in OxMetrics, are based on the approximations to the asymptotic distributions derived by Doornik (1998). It should be noted that inclusion of impulse dummies in the VAR affects the asymtotic distribution of the reduced rank test statistics. Thus, the critical values are only indicative. The asterisk * denotes rejection of the null hypothesis at the 5 per cent significance level.

We observe that the rank should be set to unity at the 5 per cent significance level (albeit the λ^a_{trace} statistics is a borderline case), indicating existence of one

[6] We notice that the preferred VAR includes two additional impulse dummies to mop up relatively large residuals in the quarters 1984Q1 and 1996Q1 in the ulc_t-equation and the p_t-equation, respectively. These dummies are labelled $D84Q1$ and $D96Q1$. The ulc_t-equation suffers from severe residual autoregressive heteroskedasticity without $D84Q1$, while the p_t-equation has clear non-normal residuals (skewness and excess kurtosis) without $D96Q1$. We emphasise that the cointegration analysis below are not significantly affected by any of the impulse dummies $D84Q1$, $D86Q3$, $D96Q1$, $D01Q3$, $D03Q1$ and $D03Q2$ included in the preferred VAR.

cointegration relationship between consumer prices, unit labour costs and import prices. Testing the $I(2)$ hypothesis by means of Johansen (1995), which combines testing the rank of Π as before and the potential additional reduced rank restriction on the long run matrix of the model in first differences, proposes that the number of $I(2)$ relations is zero in our case. It may be that including a quadratic deterministic trend would yield an even more satisfactory model fit. From an economic perspective, however, such a trend in levels of the variables is not a sensible long run property.

The null hypothesis that the linear trend can be eliminated from the VAR, assuming the rank to be unity, is not rejected by a likelihood ratio test. The p-value is 0.388 based on a χ^2 approximation with one degree of freedom. The corresponding maximised value of the 2 log likelihood, to be used in Section 5.1, is 2536.72. Imposing a further restriction of homogeneity between p_t, ulc_t and pi_t entails a reduction in the value of the 2 log likelihood of 0.00097, which corresponds to a p-value of 0.912 using the same χ^2 approximation. The issue of joint weak exogeneity of ulc_t and pi_t is more debatable. Here the -2 log likelihood ratio value is 7.909. The p-value based on approximating the null distribution with a χ^2 distribution with two degrees of freedom is 0.02. However, investigating weak exogeneity more closely, using both parametric and non-parametric bootstrap methods, reveals that the asymptotic approximations are not accurate in our case. A bootstrap of the likelihood ratio test, using the estimated values of the VAR coefficients not imposing weak exogeneity and resampling the residuals, yields a p-value of 0.515. The outcome of a non-parametric bootstrap is similar. Hence, we conclude that the cointegration vector enters the Δp_t-equation only.

We obtain the following restricted cointegrating vector (normalised on p_t) when the restrictions of homogeneity between p_t, ulc_t and pi_t, weak exogeneity of both ulc_t and pi_t and no linear trend in β are imposed (standard errors in parenthesis):

$$p_t = const. + \underset{(0.084)}{0.604} ulc_t + 0.396 pi_t. \tag{10}$$

To sum up, we interpret (10) as a long-run consumer price equation that corresponds well with the theory of mark-up prising and that for a small open economy like the Norwegian, open economy features such as import prices are expected to matter somewhat. The estimates in (10) are in line with previous findings on Norwegian data, see e.g. Bårdsen *et al.* (2005, p. 182). More than three decades ago, Aukrust (1977, p. 123) pointed out that the total direct effect on consumer prices to be expected, under Norwegian conditions, from a proportionate increase of all import prices can be put at 0.33 per cent. Hence, (10) is also in line with the Scandinavian model of inflation, cf. Lindbeck (1979).

5 Tests of the OE-NKPC model

An important econometric issue when testing the OE-NKPC concerns whether the model is specified in its *exact* or *inexact* form by introducing a stochastic error term u_t. Generally, absence of an unobserved disturbance term ($u_t = 0$) may be a restrictive and nontrivial assumption as there are several justifications for why such a term could be included in the model, see e.g. Sbordone (2005). To shed light on the importance of the disturbance term in our empirical case, we evaluate both versions of the model in this paper. However, as demonstrated by Boug *et al.* (2010), the *exact* version of the NKPC is algebraically less involved and produces much simpler rational expectations (RE) restrictions on a bivariate VAR than what follows from the *inexact* version under the assumption of u_t being a sequence of innovations, i.e., $E_t(u_{t+1}) = 0$. Hence, the numerical treatment of the *exact* model using likelihood based methods is also much simpler than the *inexact* model. When a trivariate VAR is the underlying model, as is the case in the present study, the numerical treatment of the *inexact* NKPC is even more complicated to handle within likelihood based methods. As a consequence, we employ the testing procedure suggested by Johansen and Swensen (1999, 2004, 2008) and the commonly used GMM procedure when evaluating the *exact* and the *inexact* versions of the OE-NKPC, respectively.

5.1 The e*xact* OE-NKPC

The basic idea behind the procedures suggested by Johansen and Swensen (1999, 2004, 2008) is to start with a well-specified VAR model and test, using a likelihood ratio test, the implications of the OE-NKPC for the parameters of the VAR. To construct a likelihood ratio test, we need to work out the maximum likelihood estimator of the parameters, both with and without the *exact* RE restrictions imposed on the model. Generally, the way the likelihood ratio test is constructed depends on whether the VAR includes restricted or unrestricted deterministic terms or a homogeneity restriction between the variables involved.

We recall from Section 4 that a well-specified reduced rank VAR (henceforth CVAR) with unrestricted deterministic terms (the constants, the seasonals and the impulse dummies) and no restricted deterministic trend passed a test for homogeneity between p_t, ulc_t and pi_t. Consequently, we will test the *exact* OE-NKPC below by means of the procedures developed in Johansen and Swensen (2008) with some necessary modifications to make them relevant in our empirical context. First, taking the conditional expectation of $c'\Delta X_{t+1}$ and using the empirical counterpart of (9), we get

$$c'E_t[\Delta X_{t+1}] = c'\Pi X_t + c'\Gamma_1 \Delta X_t + c'\Gamma_2 \Delta X_{t-1} + c'\Phi_0 D_{t+1} + c'\Phi_1, \tag{11}$$

where $c' = (1, 0, 0)$. Then, we make use of the fact that the *exact* form of the baseline

OE-NKPC in (6) can be expressed as

$$E_t[\Delta p_{t+1}] = (\lambda/\delta)(p_t - \gamma_1 ulc_t - \gamma_2 pi_t) + (1/\delta)\Delta p_t - \lambda\gamma_0/\delta. \tag{12}$$

Defining $c = d_1 = (1,0,0)'$, $d_2 = (0,0,0)'$ and $d = (1, -\gamma_1, -\gamma_2)'$, and letting $\tau = (\lambda/\delta)$, $\tau_1 = (1/\delta)$ and $\mu = -\lambda\gamma_0/\delta$, we observe that (12) has exactly the same form of the RE restrictions as covered by equation (2) in Johansen and Swensen (2008), i.e.,

$$c' E_t[\Delta X_{t+1}] = \tau d' X_t + \tau_1 d_1' \Delta X_t + \mu. \tag{13}$$

The *exact* RE model in (13) expressed as restrictions on the coefficients of (11) will imply restrictions also on the deterministic parts of the model. Because we are not focusing on the properties captured by the seasonals and the impulse dummies from the fitted VAR, we simply drop the RE restrictions on these deterministic components. As explained in Boug *et al.* (2006) this amounts to formulating the *exact* RE restrictions as

$$c' E_t[\Delta X_{t+1} - \Phi_0 D_{t+1}] = \tau d' X_t + \tau_1 d_1' \Delta X + \mu. \tag{14}$$

Equating (14) and (11) rewritten as $c' E_t[\Delta X_{t+1} - \Phi_0 D_{t+1}]$ implies that the following RE restrictions must be satisfied in the *exact* baseline OE-NKPC case:

$$c'\Pi = \tau d', \quad c'\Gamma_1 = \tau_1 d_1', \quad c'\Gamma_2 = 0 \quad \text{and} \quad c'\Phi_1 = \mu. \tag{15}$$

To find the profile or concentrated likelihood, the structural parameters γ_1 and γ_2, and hence d, are considered as known, while τ and τ_1 are allowed to vary. Accordingly, following the procedures in Johansen and Swensen (2008), we may compute for each value of $\gamma = \gamma_1 = 1 - \gamma_2$ the value of the likelihood when the homogeneity restriction and the restrictions in (15) are satisfied concurrently. Moreover, the restrictions in (15) imply that the marginal part of the model takes the form

$$\Delta p_t = \tau[p_{t-1} - \gamma ulc_{t-1} - (1-\gamma)pi_{t-1}] + \tau_1 \Delta p_{t-1} + \phi' D_t + \epsilon_{1t}, \tag{16}$$

where $\phi = (\phi_1, \ldots, \phi_{10})$ are the coefficients on the seasonals, the impulse dummies and the constant from the VAR. We notice that there are five restrictions involved as the coefficients of Δulc_{t-1}, Δpi_{t-1}, Δp_{t-2}, Δulc_{t-2} and Δpi_{t-2} are constrained to zero. The conditional part of the model involves an unrestricted regression of $\Delta(ulc_t, pi_t)'$ on Δp_t, $[p_{t-1} - \gamma ulc_{t-1} - (1-\gamma)pi_{t-1}]$, Δp_{t-1}, Δulc_{t-1}, Δpi_{t-1}, Δp_{t-2}, Δulc_{t-2}, Δpi_{t-2}, constant, the seasonals and the impulse dummies. For fixed values of γ the numerical value of the likelihood can be evaluated, and hence the maximum likelihood estimate of γ be determined by a numerical optimisation routine. Once the maximum likelihood estimate of γ is known, the estimates of τ and τ_1 can be found by ordinary least squares (OLS) from the marginal part of the model.

The structural parameters δ and λ then follow from the definitions $\tau = (\lambda/\delta)$ and $\tau_1 = (1/\delta)$. As we pointed out in Section 2 the two formulations (6) and (7) are reparameterisations of each other. Due to the invarians property of the maximum likelihood estimators under reparameterisations the results just reported also apply for the alternative formulation in (7).

All the procedures described above are also applicable to the *exact* form of the hybrid OE-NKPC in (8). Specifically, the hybrid OE-NKPC can be formulated on the form of equation (2) in Johansen and Swensen (2008) by letting c, d and d_1 be as earlier and defining $d_2 = (1, 0, 0)'$, $\tau = \lambda_h/\gamma_f$, $\tau_1 = 1/\gamma_f$, $\tau_2 = -\gamma_b/\gamma_f$ and $\mu = -\lambda_h\gamma_0/\gamma_f$. Using these definitions of the parameters to modify the equations (13), (14) and (15) we may again compute for each value of γ the value of the likelihood when both the homogeneity restriction and the restrictions implied by the *exact* RE hypothesis are satisfied. Because the models are nested we shall use the top down procedure by means of likelihood ratio tests when testing the models against each other. Table 2 summarises the outcome of the likelihood ratio tests of the *exact* OE-NKPC.

Table 2: Likelihood ratio tests of the exact OE-NKPC. Nested models

Model	$2\log L$	$-2\log LR$	df.	p-value
CVAR without homogeneity restriction	2536.72[1]			
CVAR with homogeneity restriction	2536.71[1]	0.01	1	0.92
Exact hybrid OE-NKPC	2535.15[2]	1.56	4	0.82
Exact baseline OE-NKPC	2529.64[2]	5.60	1	0.02

[1] Maximal values of the likelihood *without* the RE restrictions imposed.
[2] Maximal values of the likelihood *with* the RE restrictions imposed.

We observe that the likelihood ratio tests indicate that the *exact* hybrid OE-NKPC is not rejected, whereas the baseline model is. Hence, allowing for lagged inflation to enter the OE-NKPC does improve the performance of the model compared to the baseline model. These impressions are also evident from the plots of the concentrated likelihood functions displayed in Figure 2.

The curve corresponding to the CVAR with the homogeneity restriction as the only restriction imposed has its maximum at $\hat{\gamma} = 0.359$, which is not much different from the maximum likelihood estimate of $\hat{\gamma} = 0.379$ in the *exact* hybrid OE-NKPC case. However, the corresponding maximum likelihood estimates of $\gamma_f = 5.162$ and $\gamma_b = -0.747$ are both outside the intervall $[0, 1]$, which do not make sense economically. To investigate these findings further we compute estimates of the structural parameters γ_f, γ_b and λ_h for some reasonable values of γ in addition to the maximum likelihood estimate of $\hat{\gamma} = 0.379$. The computed estimates are displayed in Table 3.

We see that the estimates for γ_f, γ_b and λ_h in all cases are outside the region of having admissable economic interpretations. Additional evidence of estimates with economically meaningless interpretation is provided by Figur 3, which plots

Figure 2: Concentrated likelihood functions as functions of γ. CVAR with homogeneity restriction (solid line), exact hybrid OE-NKPC (short dashed line) and exact baseline OE-NKPC (long dashed line)

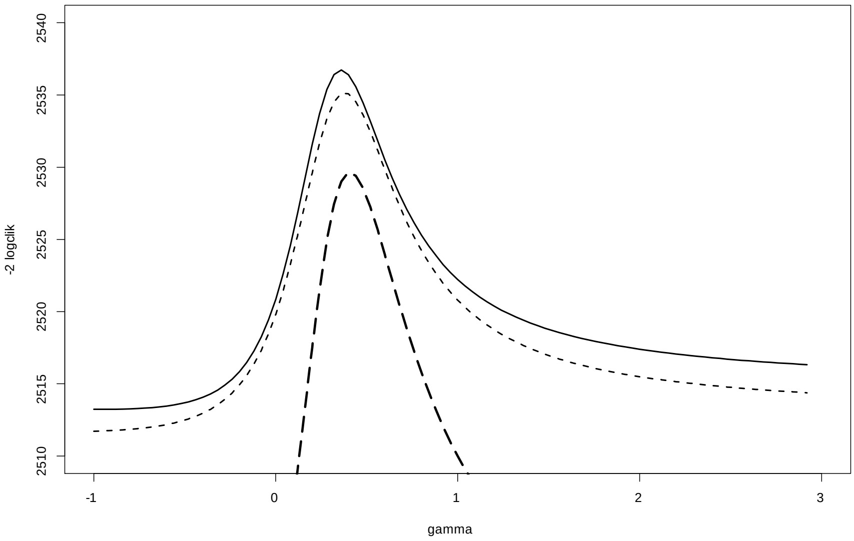


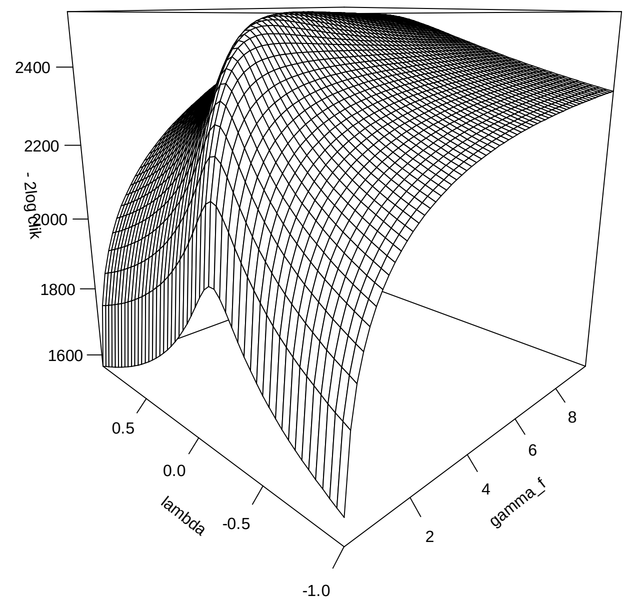
Table 3: Some parameter estimates of the exact hybrid OE-NKPC

γ	γ_f	γ_b	λ_h
0.20	3.567	-0.735	-0.143
0.38	5.162	-0.747	-0.267
0.60	5.844	-0.866	-0.230
0.90	4.830	-0.916	-0.105

Notes: The estimates of γ_f, γ_b and λ_h are computed for reasonable values of γ.

the concentrated likelihood surface $2 \log cL(\gamma_f, \gamma_b, \gamma, \lambda_h)$ as a function of γ_f and λ_h for $\gamma = 0.379$ and $\gamma_b = -0.747$. Allowing only economically meaningful parameter values ($0 \leq \gamma_f, \gamma_b, \gamma \leq 1$ and $\lambda_h \geq 0$) yields a maximal value of $2 \log L$ equal to 2436.40, corresponding to $\hat{\gamma_f} = 1.0, \hat{\gamma_b} = 0.0005, \hat{\lambda}_h = 3.93E - 6$ and $\hat{\gamma} = 0.23$, which are on the border of the permissible region. The likelihood ratio test for the null hypothesis that these estimates belong to the permissable region has a non-standard asymptotic distribution, which is a convex combination of χ^2 distributions with different degrees of freedom, see Boug *et al.* (2010) for details. In this case, the critical values are smaller than the critical values computed from a standard $\chi^2(4)$ distribution. Because the difference of the maximal values of $2 \log L$ is so large ($2535.15 - 2436.40 = 98.75$) and therefore exceeds all relevant critical values using a standard $\chi^2(4)$ distribution, the likelihood ratio test also rejects the null hypothesis that $0 \leq \gamma_f, \gamma_b, \gamma \leq 1$ and $\lambda_h \geq 0$.

Figure 3: Surface plot of concentrated likelihood function as a function of γ_f and λ_h for $\gamma = 0.379$ and $\gamma_b = -0.747$. The exact hybrid OE-NKPC



We conclude from the likelihood ratio tests reported in this section that although the number of cointegrating relations and the homogeneity restriction suggested by the theory are supported by the Norwegian data, the dynamic structure implied by the *exact* OE-NKPC model is not. Accordingly, the *exact* RE hypothesis embedded in the theoretical model seems too simple to be in accordance with the data.

5.2 The *inexact* OE-NKPC

We now turn to the *inexact* form of the OE-NKPC and its implications for the RE restrictions on the trivariate cointegrated VAR. Due to presence of an unobserved disturbance term, the restrictions in (14) now take the form

$$c' E_t[\Delta X_{t+1} - \Phi_0 D_{t+1}] = \tau d' X_t + \tau_1 d_1' \Delta X_t + \mu + u_t, \tag{17}$$

where u_t is assumed, like in Boug *et al.* (2010), to be a sequence of innovations, i.e., $E_t(u_{t+1}) = 0$. To see how the form in (17) has important implications for the RE restrictions that differ from those implied by (14), we make use of methods similar to those used by Boug *et al.* (2010) for bivariate VARs.[7] First, the fitted reduced rank VAR may be written on level form as

$$X_t = A_1 X_{t-1} + A_2 X_{t-2} + A_3 X_{t-3} + \Phi_0 D_t + \Phi_1 + \epsilon_t, \tag{18}$$

and the restrictions in (17) correspondingly as

$$c' E_t[X_{t+1} - \Phi_0 D_{t+1}] = c_0' X_t + c_{-1}' X_{t-1} + u_t, \tag{19}$$

for c as earlier, $c_0 = c + \tau d + \tau_1 d_1$ and $c_{-1} = -\tau_1 d_1$. Then, rewriting (19) at time $t+1$, using the law of iterated expectations and inserting one-step ahead forecasts from the VAR, the following restrictions on the coefficients of the VAR must be satisfied in the *inexact* baseline OE-NKPC case when $E_t(u_{t+1}) = 0$:

$$\begin{aligned} c'(A_1^2 + A_2) + c_0' A_1 + c_{-1}' &= 0 \\ c'(A_1 A_2 + A_3) + c_0' A_2 &= 0 \\ c'(A_1 A_3) + c_0' A_3 &= 0 \end{aligned} \tag{20}$$

$$c'[A_1(\Phi_0 D_{t+1} + \Phi_1) + \Phi_1] + c_0'(\Phi_0 D_{t+1} + \Phi_1) = 0. \tag{21}$$

The model (18) with reduced rank equal to unity contains $18 + 3 + 2 = 23$ autoregressive parameters in addition to the coefficients of the deterministic terms. There are not more than 9 restrictions on the coefficients of the VAR in (20), so using the reversed engineering approach of Kurmann (2007), expressing the likelihood in terms of the parameters of the inflation equation and the structural parameters δ, γ and λ, we end up with at least 17 freely varying parameters in addition to those from (21). However, the maximum likelihood estimates are computationally troublesome to obtain due to the rather complicated nature of the restrictions in (20), and is beyond the scope of the present paper. To simplify matters, we therefore follow a number of related studies and adopt the GMM approach to evaluate *inexact* versions of the OE-NKPC.

The GMM approach in our context requires identifying relevant instruments and does not necessitate strong assumptions on the underlying model, as is the case with the VAR and the likelihood based procedure used above. First, we define in line with Galí and Gertler (1999) among others the RE forecast error as $v_{t+1} \equiv \Delta p_{t+1} - E_t \Delta p_{t+1}$. Then, $E_t[v_{t+1}] = 0$ according to the RE hypothesis. Replacing $E_t \Delta p_{t+1}$ in (6) by its realised value Δp_{t+1} we obtain the following modified equation in the case of the *inexact* baseline OE-NKPC:

$$\Delta p_t = \delta \Delta p_{t+1} - \lambda eqcm_t + \xi_t, \tag{22}$$

where $\xi_t = u_t - \delta v_{t+1}$ is a linear relationship between the stochastic error term u_t and the forecast error v_{t+1} in predicting future inflation and the homogeneity restriction

[7] See also Kurmann (2007) and Fanelli (2008).

$\gamma_2 = 1 - \gamma_1$ is satisfied in $eqcm_t$. We notice that estimating (22) by means of the errors in variables method induces first order moving average errors by construction, see e.g. Bårdsen *et al.* (2005, p. 291) for details. Estimated serial correlation thus corroborates forward-looking behaviour in the RE sense, but it may also be a sign of model misspecification, as discussed in Bårdsen *et al.* (2004). Nevertheless, the possibility of serially correlated errors motivates the use of GMM. Under the RE hypothesis in (6), we also have that

$$E_t \{(\Delta p_t - \delta \Delta p_{t+1} + \lambda eqcm_t) z_{t-1}\} = 0, \tag{23}$$

where z_{t-1} is a vector of instruments dated at time $t-1$ and earlier. The orthogonality conditions in (23) provide the basis for the GMM estimation of the *inexact* baseline OE-NKPC in our context Because the instrument set includes only lagged variables, we implicitly treat $eqcm_t$ as an endogenous variable.[8] We use the corresponding GMM set up to estimate *inexact* versions of (7) and (8).

A potential shortcoming of our approach, as pointed out by e.g. Galí *et al.* (2005), is that GMM estimates may be biased in favour of finding a significant role for expected future inflation, even if that role is truly absent or negligible, if the instrument set includes variables that directly cause inflation, but are omitted as regressors in the model specification. Similarly, Mavroeidis (2005) argues that NKPC models are likely to suffer from underidentification, and that identification in empirical applications is achieved by confining important explanatory variables to the set of instruments, with misspecification as a result. In principle, misspecification can be tested using Hansen's (1982) J test of overidentifying restrictions. However, Mavroeidis (2005) shows that using too many instruments seriously weaken the power of the J test, thus obscuring specification problems and distorting GMM based inference. We address these issues below by using relatively few instruments that may also play a role as additional explanatory variables.

Throughout the evaluation of (22) and the corresponding equations for (7) and (8), we use the following set of instruments similar to the predetermined variables from the reduced rank VAR established above:

$$z_{t-1} = \left\{ \begin{array}{c} \Delta p_{t-1}, \Delta p_{t-2}, \Delta pi_{t-1}, \Delta pi_{t-2}, \Delta ulc_{t-1}, \\ \Delta ulc_{t-2}, eqcm_{t-1}, \theta, SD_{it} \end{array} \right\}, \tag{24}$$

where θ denotes a constant term.[9] The number of instruments used in our analyses is small compared to e.g. Batini *et al.* (2005), who base their study on as much

[8] This could be justified by the fact that $p_t = \Delta p_t + p_{t-1}$, and thus includes the left hand side variable as a right hand side variable in (22). It is common in the NKPC literature to use only lagged instruments, see e.g. Galí and Gertler (1999) and Galí *et al.* (2001). One obvious reason is that some current information may not be available at the date when price setters form their expectations.

[9] The impulse dummies $D84Q1$, $D86Q3$, $D96Q1$, $D01Q3$, $D03Q1$ and $D03Q2$ used in the VAR to account for outliers and special events in the economy are not included in the set of instruments to facilitate GMM estimation in EViews6.

as 40 instruments. Table 4 reports GMM estimates of the *inexact* counterparts to (6), (7) and (8) for the sample period $1982Q4 - 2005Q3$ when iterating over both coefficients and weighting matrix, with fixed bandwidth based on Newey and West (1987).[10]

Table 4: GMM estimates of the inexact OE-NKPC

	Model (6)	Model (7)	Model (8)
Δp_{t+1}	-0.639 (0.390)	0.147 (0.220)	-0.636 (0.385)
Δp_{t-1}			0.031 (0.115)
Δpi_t		-0.027 (0.075)	
Δulc_t		0.043 (0.017)	
$eqcm_t$	-0.164 (0.043)		-0.160 (0.044)
$eqcm_{t-1}$		-0.077 (0.024)	
constant	0.084 (0.022)	0.039 (0.011)	0.082 (0.022)
SD_{2t}	-0.0077 (0.0036)		-0.0076 (0.0036)
SD_{3t}	-0.0137 (0.0031)	-0.0070 (0.0010)	-0.0136 (0.0031)
Observations	92	92	92
$\hat{\sigma}$	0.0071	0.0057	0.0071
$\chi^2_J(\cdot)$	2.887 [0.823]	3.004 [0.699]	2.894 [0.716]

Notes: Sample period is $1982Q4 - 2005Q3$, $\hat{\sigma}$ denotes the estimated residual standard error, $\chi^2_J(\cdot)$ is the J statistics of the validity of the instruments with degrees of freedom $(\cdot)$ being 6, 5 and 5 for models (6), (7) and (8), respectively, parenthesis (..) contain standard errors and square brackets contain p-values.

An intercept is freely estimated in all three models in line with standard practice, which is reasonable as we do not correct for the mean in the inflation series prior to estimation. Also, there is no reason to believe that the long run mean of inflation should be zero. The fact that the estimated constant comprises the mean of the cointegration relationship, elements of short run dynamics as well as being influenced by the scaling of the variables (see Appendix 2) makes the level of the mark-up as such non identifiable.

We observe that the equilibrium correction term is highly significant in all three models, an aspect of the data which supports the *inexact* versions of (6), (7) and (8). However, the statistical *in*significance of the forward-looking term contradicts the theoretical OE-NKPC in all three cases. The GMM results with respect

[10]The Newey-West fixed bandwidth is based on the number of observations in the sample, which in our case is given by $\text{int}[4(92/100)^{2/9}] = 3$. None of the impulse dummies from the VAR are significant at the 5 per cent significance level when added individually to the models reported in Table 4.

to the forward-looking term, the equilibrium correction term, $\hat{\sigma}$ and the J statistics are hardly affected when comparing models (6) and (8), the latter estimated by including the first lag of inflation from the list of instruments as an additional regressor in the model. We notice that Δp_{t-1} is far from being significant in model (8), which is also the case for all the other variables from the instrument set. The argument of Mavroeidis (2005) that identification of the NKPC is achieved by confining important explanatory variables to the set of instrument does not seem to be relevant in our empirical case.

We conclude from the range of GMM estimates in this section that neither the *inexact* baseline nor the *inexact* hybrid OE-NKPC fit Norwegian data well. Accordingly, we claim that introducing a disturbance term to our model in the way it is interpreted here is not important when evaluating the empirical performance of the OE-NKPC. Our findings stand in sharp contrast to several existing studies using GMM, which present evidence that the NKPC is a good approximation of inflation dynamics in the US and Europe, cf. Galí and Gertler (1999), Galí *et al.* (2001) and Batini *et al.* (2005) to mention some few examples.

6 A competing model of inflation

That the forward-looking term in the OE-NKPC is statistically *in*significant motivates us to evaluate empirically the dynamic backward-looking ICM model as a competing model of inflation. We recall from Section 2 that the dynamic backward-looking ICM model is nested by and is a special case of the OE-NKPC when the forward-looking term is excluded from the model. Hence, to the extent that our data set is able to discriminate between the two competing models of inflation, we should find a well-specified dynamic backward-looking ICM model as judged by econometric criteria. To establish such a model in this section, we shall rely on a general-to-specific modelling strategy using the autometrics procedure available in OxMetrics, see Doornik and Hendry (2009). The point of departure is the general conditional model

$$
\begin{aligned}
(25)\quad \Delta p_t \; = \; & \theta + \sum_{i=1}^{2} \varphi_{1,i} \Delta p_{t-i} + \sum_{i=0}^{2} \varphi_{2,i} \Delta ulc_{t-i} + \sum_{i=0}^{2} \varphi_{3,i} \Delta pi_{t-i} + \eta eqcm_{t-1} \\
& + \theta_1 SD_{1t} + \theta_2 SD_{2t} + \theta_3 SD_{3t} + \phi_1 D84Q1 + \phi_2 D86Q3 \\
& + \phi_3 D96Q1 + \phi_4 D01Q3 + \phi_5 D03Q1 + \phi_6 D03Q2 + e_t,
\end{aligned}
$$

which is justified by the cointegration analysis and the reduced rank VAR established above, both in terms of the number of lags and the weak exogeneity test results of pi_t and ulc_t, see Boswijk and Urbain (1997). The error term e_t in (25) is assumed to be white noise. Briefly speaking, autometrics first tests the general model for misspecification to ensure data coherence. If data coherence is satisfied, then the general model is simplified by excluding statistically insignificant variables. Because autometrics controls for any invalid reduction by means of diagnostic tests, the specific model choice will not loose any significant information about the relationship from the available data set. As a result, the specific model parsimoniously encompasses the general model and is not dominated by any other model. Autometrics picks the following specific model in our case together with diagnostic tests[11] and standard errors in parenthesis:

$$
\begin{aligned}
(26)\ \Delta p_t \quad = \quad & \underset{(0.068)}{0.184}\Delta p_{t-1} + \underset{(0.065)}{0.144}\Delta p_{t-2} - \underset{(0.012)}{0.053} eqcm_{t-1} + \underset{(0.005)}{0.025} + \underset{(0.0011)}{0.0056} SD_{1t} \\
& + \underset{(0.0012)}{0.0034} SD_{2t} - \underset{(0.0011)}{0.0039} SD_{3t} + \underset{(0.003)}{0.019} D86Q3 - \underset{(0.003)}{0.011} D96Q1 \\
& - \underset{(0.003)}{0.015} D01Q3 + \underset{(0.004)}{0.020} D03Q1 - \underset{(0.004)}{0.026} D03Q2 \\
& \\
& OLS, T = 93\ (1982Q4 - 2005Q4), \hat{\sigma} = 0.0033 \\
& AR_{1-5}\text{: } F(5, 76) = 1.86\ [0.11],\ ARCH_{1-4}\text{: } F(4, 85) = 1.89\ [0.12], \\
& NORM\text{: } \chi^2(2) = 1.49\ [0.47],\ HET\text{: } F(9, 78) = 1.75\ [0.09].
\end{aligned}
$$

Several features about Norwegian inflation dynamics stand out from (26). First, we observe that the diagnostics tests reveal no symptom of misspecification in the model and $\hat{\sigma}$ is reduced considerably from the corresponding estimates for the OE-NKPC. Second, the economic variables entering the model are all highly significant. Consumer price inflation in Norway seems to be rather persistent as represented by the significant autoregressive coefficients of Δp_{t-1} and Δp_{t-2}.[12] The $eqcm_{t-1}$ appears with a t-value of -4.35, hence adding force to the results obtained from the cointegration analysis. Third, the sign of the impulse dummies corresponds well with the expected effects of the associated economic events described above. Fourth, we see that no significant contemporaneous short run effects on inflation from import prices and unit labour costs are inherent in (26). No contemporaneous short run effects and the small magnitude of the estimated loading coefficient (0.053) together imply very slow consumer price adjustment in the face of shocks in import prices and unit labour costs.

Empirical evidence of constancy of (26) may be judged from recursive test statistics, see Doornik and Hendry (2009). Neither one-step residuals with 2 estimated equation standard errors nor a sequence of break point Chow tests at the 1 per cent significance level indicate non-constancy. All recursive estimates vary little, especially relative to their estimated uncertainty. That no significant structural

[11] AR_{1-5} is a test for until 5th order residual autocorrelation; $ARCH_{1-4}$ is a test for until 4th order autoregressive conditional heteroskedasticity in the residuals; $NORM$ is a joint test for residual normality (no skewness and excess kurtosis) and HET is a test for residual heteroskedasticity, see Doornik and Hendry (2009). The numbers in square brackets are p-values.

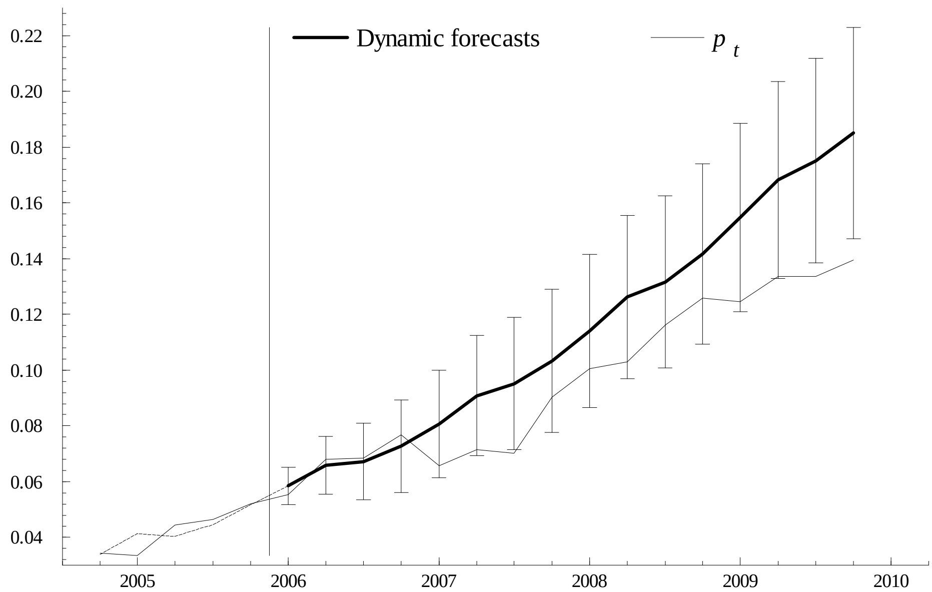
[12] Batini (2006) also finds this type of inflation persistence to be substantial in the harmonised index of consumer prices for the whole euro area and in the consumer price index for Italy, France and Germany.

breaks are detected around the date of the shift in monetary policy regime from exchange rate targeting to inflation targeting (late March 2001) points to (26) *not* being subject to the Lucas critique.

Having identified a well-specified dynamic ICM model in-sample, we study the out-of-sample forecasting performance to shed light on its robustness with respect to relatively large movements in the exchange rate during the recent financial crisis, which took off after the bankruptcy of Lehman Brothers September 15th 2008. Taylor (2000) argues that the extent to which a firm matches exchange rate movements by changing its own price depends on how persistent the movements are expected to be. For a retail firm that adds services to its imports of final goods, a depreciation of the exchange rate will raise the costs of the imports evaluated in domestic currency. If the depreciation is viewed as temporary, the retail firm will according to Taylor (2000) pass through less of the depreciation to its own price.

If the price setting behaviour indeed changed significantly following the financial crisis, we should expect instabilities in the estimated Δp_t-equation as, for example, indicated by poor out-of-sample forecasting ability. To assess the forecasting performance of (26), we employ sixteen quarters ($2006Q1-2009Q4$) of out-of-sample observations, including the period of the financial crisis. Figure 4 depicts actual values of p_t together with dynamic forecasts, adding bands of 95 per cent confidence intervals to each forecast in the forecasting period.

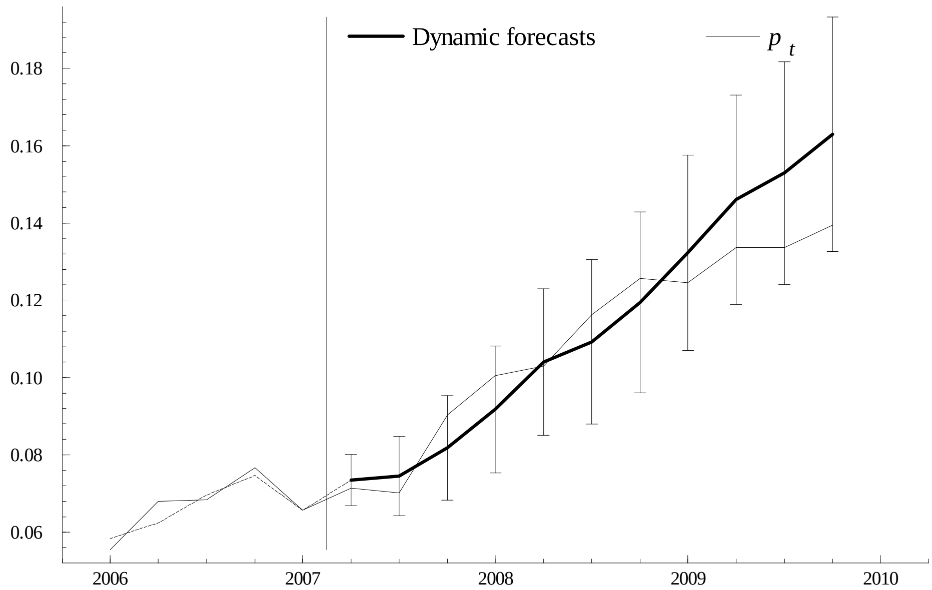
Figure 4: Actual values and dynamic forecasts of p_t



A majority of the actual values of p_t stay within their corresponding confidence intervals over the forecasting period. The actual value of p_t is close to be outside the confidence interval in the first quarter of 2007. The point in time in which the first sign of instability occurs does *not*, however, coincide with the period of the financial crisis. Rather, the actual value of p_t in $2007Q1$ and the values thereafter are likely to be influenced by the huge and transitory fall in electricity prices during the first quarter of 2007. Consequently, the dynamic forecasts overpredict the actual values of p_t thereafter.

To take a closer look at this hypothesis for the forecasting failure of (26), we reestimate the model over the period $1982Q4 - 2007Q1$ with an impulse dummy in $2007Q1$ as a separate regressor controlling for the substantial fall in electricity prices during that quarter. Hence, eleven observations are now available for forecasting. The reestimated equation is virtually unchanged from (26) with respect to both parameter estimates and diagnostics. Figure 5 plots actual values of p_t together with dynamic forecasts when the reestimated model is used for forecasting.

Figure 5: Actual values and dynamic forecasts of p_t



We observe that the forecasting failure of (26) is eliminated. Thus, the general impression of the out-of-sample forecasting ability of (26) is reasonably good despite relatively large exchange rate movements in the wake of the financial crisis. We may argue in light of Taylor (2000) that the exchange rate movements during the financial crisis were perceived as transitory rather than permant shocks such that firms found it rational not to alter their pricing behaviour.

To sum up, the economic properties inherent in (26) seem consistent with the actual inflation persistence in Norway. Our data set is able to discriminate between the OE-NKPC and the dynamic backward-looking ICM model, the former being rejected in favour of the latter. Of course, relying on Bårdsen *et al.* (2005, p. 183), other and more elaborate dynamic backward-looking ICM models in which electricity prices and unemployment are allowed to play a role may exist. However, the purpose here has been to emphasise that discrimination between the two rival models is possible through testable restrictions using the same information set throughout.

7 Conclusions

In this paper, we have evaluated the empirical performance of an open economy version of the NKPC (labelled OE-NKPC) as a model of Norwegian inflation. Our starting point was the forward-looking linear quadratic adjustment cost model of Rotemberg (1982) and the theoretical principles of the incomplete competition model (ICM) for a small open economy. We showed that our OE-NKPC relates current inflation to expected future inflation and the difference between the actual price and the price target in levels, a difference being a theory consistent forcing variable determined by a weigthed average of unit labour costs and prices of total imports. The OE-NKPC thus includes variables both in levels and differences that demand econometric care with respect to both time series properties and cointegrated nature of these variables in the model. Such econometric issues have typically been ignored in related studies on open economies data.

We first established by means of reduced rank regressions a cointegrating vector between the price level and the target level in line with the ICM model and existing evidence on Norwegian data. By way of contrast, we then found using cointegrated VAR models and likelihood based testing procedures that the *exact* OE-NKPC, both in terms of the baseline and the hybrid version, does not receive much backing from the data. We obtained similar findings when various *inexact* OE-NKPC models were evaluated within the GMM framework. Accordingly, we claim that introducing a disturbance term to the model in the way it is interpreted here is not important when evaluating the empirical performance of the OE-NKPC as a model of Norwegian inflation. Finally, we established a well-specified dynamic ICM model, which in addition to the theory consistent forcing variable includes backward-looking terms only. We found that the dynamic ICM model encompasses the OE-NKPC, is reasonably stable in-sample with a major monetary policy regime shift from exchange rate targeting to inflation targeting and forecasts well post-sample and during the recent financial crisis. All these findings are strong evidence in favour of the dynamic ICM model and the Lucas critique does not seem to be important in our empirical context. We conclude that including forward-looking behaviour when modelling consumer price inflation in Norway seems unnecessary to arrive at a well-specified model by econometric criteria.

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Appendix 1

The theoretical settings in Batini *et al.* (2005) imply that a simplified version of equation (2) in that study has the same structure as our first order condition (2). Ignoring employment adjustment costs and a stochastic error term, equation (2) in Batini *et al.* (2005) reads

$$\Delta p_t = \phi E_{t-1}\Delta p_{t+1} + \alpha_1 E_{t-1}(\ln \mu_t^* + rmc_t), \tag{27}$$

where $\Delta p_t = p_t - p_{t-1}$, μ_t^* is the equilibrium mark-up on nominal marginal costs (MC_t) and $rmc_t = \ln(MC_t/P_t) = mc_t - p_t$. Substituting $rmc_t = mc_t - p_t$ into (27) and utilising that the optimal price $p_t^* = \ln P_t^* = \ln \mu_t^* + mc_t$, we have that

$$\Delta p_t = \phi E_{t-1}\Delta p_{t+1} - \alpha_1 E_{t-1}(p_t - p_t^*), \tag{28}$$

which is identical to our equation (2) except that expectations are formed on the basis of information available at the end of period $t-1$ rather than at time t. When still abstracting from adjustment costs of employment and a stochastic error term, the operational OE-NKPC in Batini *et al.* (2005) is consistent with and has the same form and interpretation as our equation (6). To see this, we substitute the following expressions for rmc_t and $\ln \mu_t^*$ in Batini *et al.* (2005)

$$\begin{aligned} rmc_t &= -\ln \alpha + s_{L,t} + \mu_3(p_{m,t} - p_t) \\ \ln \mu_t^* &= \mu_0 + z_{p,t} + \mu_1(y_t - y_t^*) + \mu_2(p_t^w - p_t) \end{aligned} \tag{29}$$

into (27) and collect terms to obtain

$$\Delta p_t = \phi E_{t-1}\Delta p_{t+1} + \alpha_1 E_{t-1}[p_t - k - \upsilon_1(z_{p,t} + s_{L,t}) - \upsilon_2(y_t - y_t^*) - \upsilon_3 p_t^w - \upsilon_4 p_{m,t}], \tag{30}$$

where $k = (\mu_0 - \ln \alpha)/(\mu_2 + \mu_3)$, $\upsilon_1 = 1/(\mu_2 + \mu_3)$, $\upsilon_2 = \mu_1/(\mu_2 + \mu_3)$, $\upsilon_3 = \mu_2/(\mu_2 + \mu_3)$, $\upsilon_4 = \mu_3/(\mu_2 + \mu_3)$ and $z_{p,t}$, $s_{L,t}$, $(y_t - y_t^*)$, p_t^w and $p_{m,t}$ denote product market competition, labour share, state of the business cycle, world price of domestic GDP (in domestic currency) and price of total imports (in domestic currency), respectively.

Although imports are theoretically modelled as intermediate inputs in Batini *et al.* (2005), the operational equation in (30) with $p_{m,t}$ measuring the price of *total* imports is more in line with our approach when introducing open economy features to the NKPC. Nevertheless, the expression in the brackets of (30) may form a cointegration relationship with testable restrictions analogous to the hypothesis of cointegration relationship in our equation (6). Because estimation is conducted without considering time series properties and cointegration relationships between variables in levels are not tested for, Batini *et al.* (2005) run the risk of operating with unbalanced models with unreliable inference as a consequence.

Appendix 2

P: The official consumer price index (2002 = 1). Source: Statistics Norway.

PI: Implicit deflator of total imports (2002 = 1). Source: Statistics Norway, the Quarterly National Accounts.

ULC: Unit labour costs defined as YWP/QP, where YWP and QP are total labour costs and value added in the private mainland economy, respectively. Source: Statistics Norway, the Quarterly National Accounts.

$D84Q1$: Impulse dummy used to account for a large residual in the ulc_t-equation of the VAR. Equals unity in the first quarter of 1984, zero otherwise.

$D86Q3$: Impulse dummy used to control for the 12 per cent devaluation of the Norwegian currency in May 1986. Equals unity in the third quarter of 1986, zero otherwise.

$D96Q1$: Impulse dummy used to account for a large residual in the p_t-equation of the VAR. Equals unity in the first quarter of 1996, zero otherwise.

$D01Q3$: Impulse dummy used to control for the drop in the VAT rate on food from 24 per cent to 12 per cent in July 2001. Equals unity in the third quarter of 2001, zero otherwise.

$D03Q1$: Impulse dummy used to control for the large increase in electricity prices during the first quarter of 2003. Equals unity in the first quarter of 2003, zero otherwise.

$D03Q2$: Impulse dummy used to control for the large decrease in electricity prices during the second quarter of 2003. Equals unity in the second quarter of 2003, zero otherwise.

$D07Q1$: Impulse dummy used to control for the large decrease in electricity prices during the first quarter of 2007. Equals unity in the first quarter of 2007, zero otherwise.