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The New Keynesian Phillips Curve for a Small Open Economy

Abstract:

The New Keynesian Phillips Curve (NKPC) has become the benchmark model for understanding inflation in modern monetary economics. One reason for the popularity is the microfoundation of the model, which decomposes agents' behaviour into price adjustments and deviations of the price level from its target. The empirical relevance of the NKPC is, however, a matter of debate as recent studies reveal that some supportive evidence depends crucially on the econometric methods applied. We show how to evaluate the features of the model using cointegration techniques and tests based on both single-behavioural equations and cointegrated VAR models. Our results indicate that the forward-looking part of the NKPC is most likely at odds with Norwegian data. By contrast, we establish a well-specified dynamic model interpreted as a standard backward-looking mark-up price equation. We also demonstrate that the dynamic mark-up model forecasts well post-sample and during a major change in the monetary policy regime, which certainly is strong evidence in favour of this model. Consequently, we conclude that taking account of forward-looking behaviour when modelling consumer price inflation in Norway seems unnecessary to arrive at a well-specified model by econometric criteria.

Keywords: The New Keynesian Phillips Curve, mark-up pricing, single-equation estimation methods, encompassing tests, cointegrated vector autoregressive models and equilibrium correction models.

JEL classification: C51, C52, E31, F31.

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1. Introduction

The New Keynesian Phillips Curve, hereafter NKPC, is currently the most popular theory of inflation and seems to become a cornerstone in monetary policy analysis for inflation targeting central banks. One reason for the popularity is the microfoundation of the model based on the staggered wage contracts by Taylor (1979), the quadratic price adjustment cost model of Rotemberg (1982), and the staggered price model by Calvo (1983). Roberts (1995) has shown that these models can be respecified to become similar to the expectations-augmented Phillips curve of Friedman and Phelps developed a generation ago.

In its pure form, the NKPC explains inflation by expected future inflation and excess demand or marginal costs as the forcing variable. A hybrid version of the NKPC that also includes lagged inflation terms as a way of modelling "rule of thumb" or "backward-looking" price setters is usually necessary to capture adequately the inflation persistence observed in real-world data, see for instance Fuhrer and Moore (1995), Fuhrer (1997), Galí and Gertler (1999), Galí *et al.* (2001, 2005), and Estrella and Fuhrer (2002). The incorporation of lags of inflation is, though, *ad hoc* and not strictly micro-founded. Still, Galí (2003) suggests that, while the pure NKPC is rejected on statistical grounds, it is likely to be a reasonable approximation to the inflation dynamics of both Europe and the U.S. Fuhrer (2005), on the other hand, provides evidence that by and large contradicts this view on the pure NKPC as an adequate empirical model of inflation. Essentially, and contrary to what is commonly believed, his argument is that little of the persistence in the driving variable is inherited by inflation and that inflation persistence in practice is dominated by backward-looking behaviour.

Up till now the NKPC has primarily been studied within a single behavioural equation set up to evaluate inflation processes for the U.S. economy or for aggregated Euro data. Recently, Batini *et al.* (2005) have derived an open economy version of the NKPC with accommodating results on UK data. A recent study by Bårdsen *et al.* (2004), which gives critical assessments of Galí *et al.* (2001) and Batini *et al.* (2000)¹, concludes that the empirical relevance of the NKPC is fragile. The study by Bårdsen *et al.* (2004) discusses a number of econometric issues that are likely to be important when evaluating the empirical performance of the NKPC. Among them, the role of the encompassing principle is emphasised. Existing studies of inflation may provide information that the NKPC should be evaluated against. Likewise, the importance of modelling a system that includes not only the rate of inflation, but also the forcing variable is stressed. In the same vein, Lindé (2005) argues that system estimation of the NKPC by means of full information maximum likelihood (FIML) has clear advantages over single-equation estimation

¹ Batini *et al.* (2000) is the working paper version of Batini *et al.* (2005).

techniques. We may add another econometric issue typically ignored in the existing empirical NKPC-literature. Often the NKPC is evaluated within estimated models consisting of both variables in differences and levels. Since estimation is conducted without considering time series properties and cointegration relationships between variables in levels are not tested for, we argue that some studies run the risk of operating with unbalanced models with unreliable inference as a consequence.

In this paper, we investigate open economy versions of the NKPC for Norway by means of cointegration techniques, different single-equation estimation methods, the encompassing approach as well as the system approach. Our *benchmark* NKPC is based on the quadratic price adjustment cost model of Rotemberg (1982) and the theory of mark-up pricing in the open economy case. It follows that inflation is explained by expected future inflation in addition to (the deviation from) a static long run relationship between levels of consumer prices, import prices and unit labour costs as a theory-consistent forcing variable. Using standard multivariate cointegration techniques, we establish a well-specified empirical counterpart to the theory-consistent forcing variable. Applying both general method of moments (GMM) and two stage least squares (2SLS), we show that the coefficient of the forward-looking term in the benchmark NKPC is *insignificantly* different from zero whereas the forcing variable coefficient is strongly significant. Various misspecification tests indicate that these findings are robust features. So the price adjustment part of the NKPC is rejected, while the part of the model containing the deviation of the price level from its target is not.

We then show how the benchmark model can be reparameterised so that inflation is explained by expected future inflation and contemporaneous growth in unit labour costs and import prices in addition to (the deviation from) the static long run relationship (lagged one period) as forcing variables. Inspired by the *standard* way of specifying the NKPC, we evaluate the reparameterised version of the benchmark model, but without (the deviation from) the static long run relationship as a separate explanatory variable. Under these model settings, the coefficient associated with the forward-looking term becomes highly significant and close to unity in numerical value, a common finding in several studies. However, the statistical adequacy of the reparameterised model for testing purposes is questionable as that model suffers from an omitted variable bias. Our evaluation of the standard specification of the NKPC gives similar conclusion when contemporaneous growth in unit labour costs is replaced by output gap as a forcing variable. Using the cointegration relationship from our benchmark NKPC in the context of encompassing, we estimate a well-specified standard NKPC where the cointegration term turns out to be significant, whereas the forward-looking term no longer is significant. These results clearly reject the standard NKPC as an empirical model of Norwegian inflation.

We also demonstrate that utilising the testing procedure suggested by Johansen and Swensen (1999, 2004) to evaluate the various specifications of the NKPC as a system within a cointegrated vector autoregressive (VAR) model, yields results that are not in favour of the model. Finally, we establish a well-specified dynamic model of inflation with backward-looking elements only, a model which forecasts well post-sample despite that monetary policy changed from a fixed to a floating exchange rate regime following the introduction of inflation targeting in 2001. The upshot of all these tests is that the NKPC for a small open economy like the Norwegian is rejected in favour of a model interpreted as a conventional dynamic backward-looking mark-up price equation.

The paper is organised as follows: Section 2 outlines the NKPC models considered in the empirical investigations. Section 3 reports results from the cointegration analysis. Sections 4 and 5 evaluate our benchmark NKPC and the standard specification of the NKPC by means of single-equation tests, respectively. Section 6 tests the NKPC based on cointegrated VAR models. Section 7 develops a backward-looking equilibrium correction model as a competing model to the NKPC. Section 8 concludes.

2. The 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 linear quadratic adjustment cost model of Rotemberg (1982) and the models of staggered contracts developed by Taylor (1979, 1980) and Calvo (1983). We take the quadratic adjustment cost model of Rotemberg (1982) as the starting point as this model suits well with our empirical strategy. Accordingly, the representative firm chooses a sequence of prices, P_{t+j} , in order to minimise the loss function

$$(1) \quad E_t \left[\sum_{j=0}^{\infty} \delta^j [\lambda (p_{t+j} - p_{t+j}^*)^2 + (p_{t+j} - p_{t+j-1})^2] \right],$$

where E_t denotes the conditional expectation given the information contained in the information set at time t and lower case letters indicate logs, i.e., $p_{t+j} = \log(P_{t+j})$.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. Since 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

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

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 .³ We now discuss how to specify the deviation of the price level from its target. Building on existing models of inflation for Norway reported in Bowitz and Cappelen (2001), Bårdsen and Nymoén (2001), Bårdsen *et al.* (2003), and Boug *et al.* (2006), we assume that the representative firm operates in imperfectly competitive markets facing regular downward sloping demand curves. Profit maximisation then leads to the standard formula stating that the price (P) equals a mark-up (MU) times marginal costs (MC), $P = MU \cdot MC$. In the case of 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. This implies that the term in the parenthesis in (2) is proportional to the wage share as in Galí *et al.* (2001). We follow standard practice and measure marginal costs as unit labour costs (ULC).

The mark-up is usually assumed to be a constant in the NKPC-literature by referring to one particular case in Dixit and Stiglitz (1977). Subject to a number of assumptions, notably that commodities within each sub-group or industry are close substitutes among themselves, but poor substitutes for goods in other groups or industries, so-called two-stage budgeting is valid. Moreover, if the number of goods in the industry is large (denoted by n) so that $1/n$ is small, Dixit and Stiglitz (1977) show that the individual producer price has little impact on the aggregate industry price. Hence, we may assume that the individual producer ignores the effect of his price setting on the aggregate price. In a less restrictive case, the mark-up is not constant anymore, but depends on all factors affecting demand for the particular commodity.⁴ In an open economy framework, the *a priori* assumption of all goods and services being close substitutes is clearly unreasonable. We therefore allow the mark-up to depend on

² Throughout the paper, lower case letters denote logs of the corresponding upper case variables.

³ The theoretical settings in Batini *et al.* (2005) imply that a simplified version of equation (2) in that study has the same form and content as our first order condition (2). Ignoring employment adjustment costs, equation (2) in Batini *et al.* (2005) reads

as $\Delta p_t = \phi E_{t-1} \Delta p_{t+1} + \alpha_1 E_{t-1} (\ln \mu_t^* + rmc_t)$, where $\Delta p_t = p_t - p_{t-1}$, μ_t^* is the equilibrium mark-up of prices 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 equation (2) in Batini *et al.*

(2005) and utilising that the optimal price $p_t^* = \ln P_t^* = \ln \mu_t^* + mc_t$, we have $\Delta p_t = \phi E_{t-1} \Delta p_{t+1} - \alpha_1 E_{t-1} (p_t - p_t^*)$, 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 .

⁴ Cf. equation (32) in Dixit and Stiglitz (1977).

relative prices in a way that also accommodate the view that for small open economies, producers can be price takers on world markets. Specifically, we let $MU = m_0 (P / PI)^{m_1}$, where PI denotes the competing price that producers face and $m_0 > 0$ and $m_1 \leq 0$ reflect conditions on the demand side of the product markets. With $m_1 < 0$ an increase in the competing price allows the producer to increase her mark-up over marginal costs. By replacing marginal costs with unit labour costs and taking logs of the involved variables, we can then solve out for the producer's price target to obtain

$$(3) \quad p_t = \gamma_0 + \gamma_1 pi_t + \gamma_2 ulc_t + \varepsilon_t,$$

where $\gamma_0 = m_0^{1/(1-m_1)}$, $\gamma_1 = -m_1 / (1 - m_1)$, $\gamma_2 = 1 / (1 - m_1)$ and ε_t is a stochastic error term. We notice that (3) is homogeneous of degree one in competing prices and unit labour costs since $\gamma_1 + \gamma_2 = 1$. The comprised parameter γ_0 is a constant containing demand characteristics, whereas γ_1 and γ_2 are interpreted as the partial effect on p with respect to pi and ulc , respectively. Although the price equation (3) is derived from a theory of imperfect competition, it also contains the main alternative as a special case, namely that of the law of one price or perfect competition for homogenous goods. In the latter case m_1 approaches infinity, such that the price is equal to the price of the competitors, i.e., $P = PI$. 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.⁵ Equation (3) is a *static* model of the price target so that the right hand side of (3) is equivalent to p_t^* in (2). Inserting (3) in (2) gives our benchmark open economy NKPC

$$(4) \quad \Delta p_t = \delta E_t \Delta p_{t+1} - \lambda [p_t - \gamma_0 - \gamma_1 pi_t - (1 - \gamma_1) ulc_t] + \varepsilon_t.$$

Of course, we may extend (4) by lags of inflation in order to obtain hybrid versions with both forward-looking and backward-looking elements included. 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 (4) to be a balanced equation. We observe that the expression in the brackets of (4) is a theory-consistent driving variable that may form a cointegration relationship with testable restrictions, which is an empirical question. Existence of cointegration means that the constant γ_0 has a specific interpretation,

⁵ In the 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 *et al.* (2001).

namely the mean of the estimated cointegration relationship.⁶ The first step in our modelling strategy when moving from theory to empirics is to investigate whether the empirical counterpart of (4) can be expressed as a cointegration relationship among the level variables. Since the answer to this question is yes, as demonstrated in the next section, we continue our modelling strategy by estimating and testing (4) by means of both single-equation and system methods.

The evaluation of (4) will serve as a point of reference for further analyses. In particular, we want to compare (4) with alternative open economy specifications of the NKPC, which are inspired by the standard modelling strategy in the NKPC-literature. Typically, the following hybrid version of the NKPC is estimated as a single behavioural equation, see e.g. Galí and Gertler (1999), Galí *et al.* (2001, 2005) and Rudd and Whelan (2005):

$$(5) \quad \Delta p_t = \delta_1 E_t \Delta p_{t+1} + \delta_2 \Delta p_{t-1} + \omega y_t,$$

where y_t is excess demand or marginal costs, usually approximated by the output gap in the economy, unit labour costs (in logs), or the wage share (in logs). Some authors refer to (5) with $\delta_1 \approx 1$ and $\delta_2 = 0$ as the NKPC model. We may augment (5) in an *ad hoc* manner by including open economy features such as import prices, and test the significance of the forward term of inflation within such an extended NKPC. In either case, estimating (5) differs from our estimation strategy in several aspects. First, time series properties and potential cointegration relationships among levels variables are not accounted for. Second, empirical evaluation of (5) may be associated with possible incompleteness, as the exogeneity status of the forcing variable is not discussed. As emphasised by Bårdsen *et al.* (2004), a two-equation system consisting of (5) and an equation for y_t , is needed in order to allow for the possibility of feed-back from inflation on the forcing variable. Third, using standard output gap measures based on the Hodrick- Prescott procedure, as in e.g. Batini *et al.* (2005) and Paloviita (2006), are not theory-consistent driving variables, see Neiss and Nelson (2005) for a discussion. Inspired by

⁶ When abstracting from adjustment costs of employment, the open economy NKPC in equation (9) in Batini *et al.* (2005) is consistent with and has the same form and interpretation as our equation (4). To see this, we may substitute equations (6) and

(7) in Batini *et al.* (2005) into $\Delta p_t = \phi E_{t-1} \Delta p_{t+1} + \alpha_1 E_{t-1} (\ln \mu_t^* + rmc_t)$ from footnote 3 and collect terms in the parenthesis to obtain $\Delta p_t = \phi E_{t-1} \Delta p_{t+1} - \alpha_1 E_{t-1} [p_t - \kappa - \vartheta_1 (z_{p,t} + s_{L,t}) - \vartheta_2 (y_t - y_t^*) - \vartheta_3 p_t^w - \vartheta_4 p_{m,t}]$, where $\kappa = (\mu_0 - \ln \alpha) / (\mu_2 + \mu_3)$, $\vartheta_1 = 1 / (\mu_2 + \mu_3)$, $\vartheta_2 = \mu_1 / (\mu_2 + \mu_3)$, $\vartheta_3 = \mu_2 / (\mu_2 + \mu_3)$, $\vartheta_4 = \mu_3 / (\mu_2 + \mu_3)$ 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, respectively. Accordingly, the expression in the brackets of the reparameterised version of equation (9) in Batini *et al.* (2005) may form a cointegration relationship with testable restrictions analogous to the possible cointegration relationship in our equation (4). Since 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.

Bårdsen *et al.* (2004), we compare estimates of (4) and (5) by means of the encompassing principle in order to answer the question of which, if either, encompasses the other model.

3. Cointegration analysis⁷

The econometric evaluation of Norwegian inflation is conducted using quarterly, seasonally unadjusted data that spans the period 1983Q1–2001Q1. As in the study by Bårdsen *et al.* (2005), we measure quarterly inflation by the official consumer price index (*CPI*) rather than by the GDP deflator used in Galí *et al.* (2001) and Batini *et al.* (2005) among others. 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 as unrelated to the micro price setting behaviour as any concept within the national accounts. Moreover, there is a sizeable influence on the CPI from open economy features such as import prices or the exchange rate, see Boug *et al.* (2005). Thus, we argue that the consumer price index is a more relevant price series for evaluating the NKPC for Norway than the GDP deflator.

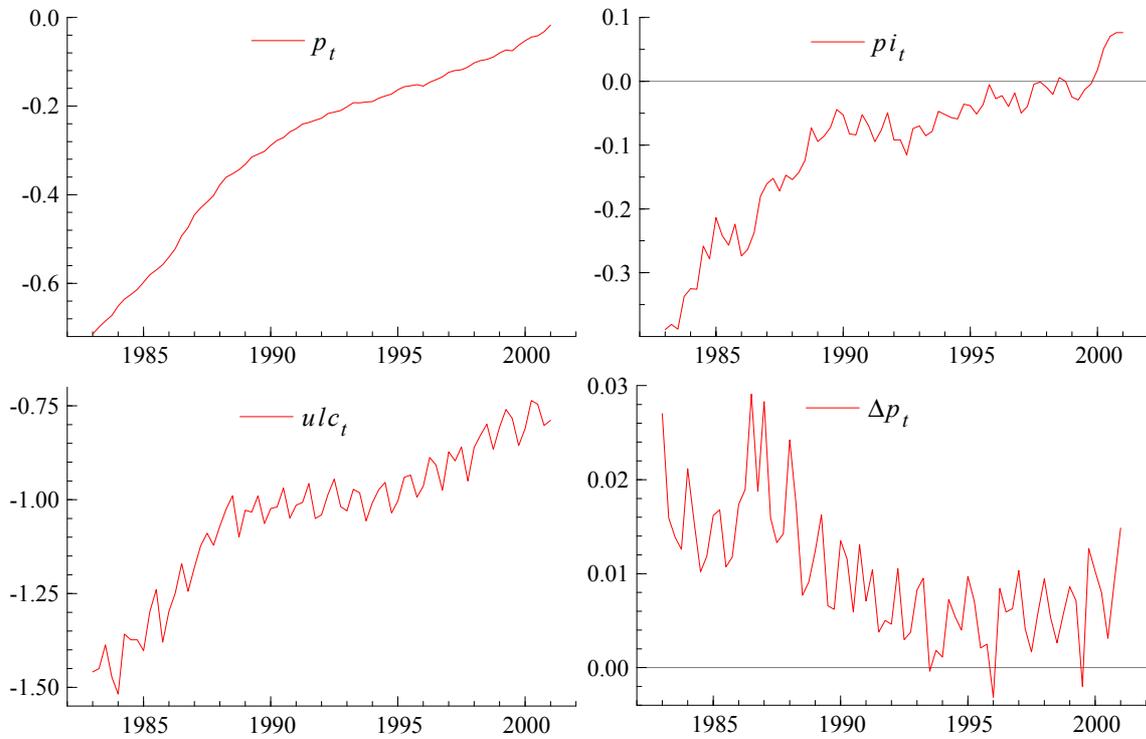
As pointed out by Rogoff (2003) among others, the 1970s was a period of high inflation rates compared to the last two decades for most OECD countries. While the Norwegian inflation rate on average was 8 per cent per year in the 1970s, the inflation rate dropped to around 2 per cent on a yearly basis in the 1990s. It may well be that there were different inflation mechanisms at work during periods of high inflation compared to low inflation periods. Moreover, the 1970s and the early 1980s were characterised by massive governmental price controls.⁸ Finally, as noted earlier, the Norwegian central bank followed a policy of exchange rate targeting in various forms over the last decades, cf. Bowitz and Cappelen (2001), until late March 2001 when monetary policy changed to inflation targeting, which could have caused the price formation to *shift* in accordance with the Lucas critique. These are the reasons for our choice of sample period. We extend the sample, however, by nineteen quarters to conduct *out-of-sample* forecasting over the period 2001Q2–2005Q4 and to shed light on the Lucas critique and its potential quantitative importance.

⁷ The empirical estimation and testing of the models are conducted using PcGive 10.3, see Hendry and Doornik (2001) and Doornik and Hendry (2001). Throughout the paper, square brackets [...] and parenthesis (..) contain *p*-values and standard errors, respectively.

⁸ Mavroeidis (2004) concludes that the estimates of a NKPC model are less reliable when the sample covers periods where inflation has been under effective policy control.

We employ the deflator for total imports (PI) as a proxy for the competing price, whereas total labour costs relative to value added in the private mainland economy serves as a proxy for unit labour costs (ULC), see the Appendix for details. Figure 1 displays the log of the consumer price index (p_t), the log of the import price deflator (pi_t) and the log of unit labour costs (ulc_t), together with the inflation rates (Δp_t) over the sample period. The two price series as well as the series for unit labour costs exhibit a clear upward trend, but with no apparent mean reverting property, suggesting p_t , pi_t and ulc_t to be I(1). Therefore, a reduced rank VAR is a candidate as an empirical model. However, it is possible that the time series are I(2) rather than I(1) over the sample period. In the cointegration analysis below we investigate both alternatives.

Figure 1. Consumer prices (p_t), import prices (pi_t), unit labour costs (ulc_t) and inflation (Δp_t)



We adapt the multivariate cointegration method suggested by Johansen (1995) to find an empirical counterpart of (4). The starting point of the I(1) cointegration analysis and the tests that follow is an equilibrium correction representation of a VAR model of order k

$$(6) \quad \Delta x_t = \sum_{i=1}^{k-1} \theta_i \Delta x_{t-i} + \pi x_{t-1} + \phi_0 D_t + \phi_1 t + \varepsilon_t,$$

where $\varepsilon_t \sim \text{NIID}(0, \Sigma)$ with zero expectation and covariance matrix Σ , x_t is a $(p \times 1)$ vector of modelled variables, D_t is a vector of deterministic components (intercepts, seasonal dummies and impulse dummies) and t is a linear deterministic trend *restricted* to lie in the cointegrating space, thereby restricting the system to at most a linear deterministic trend in levels of economic variables involved. The deterministic components in D_t are kept *unrestricted* in (6). 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 π . The null hypothesis of r cointegrating vectors may be formulated as $H_0 : \pi = \alpha\beta'$, where α and β are $p \times r$ matrices, $\beta' x_t$ comprises r cointegrating $I(0)$ linear combinations and α contains the adjustment coefficients. In the following analyses, x_t contains three potential $I(1)$ variables, namely the consumer price index (p_t), the import price deflator (pi_t) and unit labour costs (ulc_t). Initially, we estimated a 5th order VAR based on this information set. A battery of lag reduction tests suggests a lag order of three. Diagnostic tests for the preferred VAR are reported in Table 1. The residual misspecification tests show that the third order VAR produces residuals with statistically acceptable properties.⁹

Table 1. Residual misspecification tests of the third order VAR

Equation	Statistics						
	AR_{1-5} $F(5, 51)$	$ARCH_{1-4}$ $F(4, 48)$	$NORM$ $\chi^2(2)$	HET $F(20, 35)$	AR^V_{1-5} $F(45, 116)$	$NORM^V$ $\chi^2(6)$	HET^V $F(120, 180)$
Δp_t	1.04 [0.41]	0.48 [0.75]	5.02 [0.08]	0.60 [0.89]			
Δpi_t	0.78 [0.57]	1.20 [0.32]	2.01 [0.37]	1.19 [0.31]			
Δulc_t	0.24 [0.94]	0.58 [0.68]	0.25 [0.88]	0.74 [0.76]			
VAR					1.30 [0.13]	7.05 [0.32]	0.72 [0.97]

Notes: AR_{1-5} is Harvey's (1981) test for until 5th order residual autocorrelation; $ARCH_{1-4}$ is the Engle (1982) test for until 4th order autoregressive conditional heteroskedasticity in the residuals; $NORM$ is the normality test described in Doornik and Hansen (1994) and HET is a test for residual heteroskedasticity due to White (1980). Similar tests for the entire VAR are denoted by V [see Hendry and Doornik (2001)]. $F(\cdot)$ and $\chi^2(\cdot)$ represent the null distributions of F and χ^2 , with degrees of freedom shown in parenthesis.

For our VAR to be considered as a valid starting point of the cointegration analysis, it should also contain constant parameters. Recursively estimated one step residuals and sequences of break-point Chow tests show that the system is reasonably constant over the sample.¹⁰ The next step is thus to investigate the cointegration properties between the selected variables by means of our preferred

⁹ We notice that the preferred VAR includes two impulse dummies that account for outliers in the Δp_t -equation in 1986Q3 and 1996Q1, see Bårdsen *et al.* (2002a) for motivation. The normality tests (which are a joint test for skewness and excess kurtosis) are rejected for the Δp_t -equation and the system as a whole without these dummies, essentially due to excess kurtosis. Even though the properties of cointegration estimators are more sensitive to deviations from normality due to skewness than to excess kurtosis [see Juselius and MacDonald (2000)], we include the two impulse dummies in order to be on the safe side with respect to statistical inference. Additionally, the preferred VAR includes an impulse dummy to mop up an outlier in the Δulc_t -equation in 1984Q1. The Δulc_t -equation suffered from severe residual autoregressive heteroskedasticity without this impulse dummy. Noticeably, neither the stability properties of the preferred VAR nor the cointegration analyses below are significantly affected by the three mentioned impulse dummies.

system. Table 2 contains results from applying the method suggested by Johansen (1995) to determine the rank of the third order VAR, both with and without the linear trend.

Table 2. Johansen's cointegration tests

Information set: $[p, pi, ulc, D_t, t]$, eigenvalues: 0.269, 0.196, 0.076				
Information set: $[p, pi, ulc, D_t]$, eigenvalues: 0.256, 0.153, 0.018				
Hypothesis	Third order VAR <i>with</i> a linear trend		Third order VAR <i>without</i> a linear trend	
	λ_{trace}	λ^a_{trace}	λ_{trace}	λ^a_{trace}
$r = 0$	44.59 [0.032]*	39.09 [0.114]	35.09 [0.010]*	30.77 [0.038]*
$r \leq 1$	21.69 [0.153]	19.02 [0.285]	13.51 [0.097]	11.84 [0.166]
$r \leq 2$	5.76 [0.501]	5.05 [0.596]	1.36 [0.243]	1.19 [0.274]

Unrestricted cointegrating vector (without a linear trend):

$$\hat{p} = \hat{\gamma}_0 + 0.278pi + 0.683ulc$$

(0.246) (0.135)

Weak exogeneity tests:	p	pi	ulc
$\chi^2(1)$	9.189 [0.002]**	1.715 [0.190]	1.806 [0.179]

Notes: Doornik and Hendry (2001, p. 175) point out that the sequence of trace tests leads to a consistent test procedure, but no such result is available for the maximum eigenvalue test. Hence, current practice is to only consider the former. The λ_{trace} and λ^a_{trace} statistics are the trace statistics without and with degrees-of-freedom-adjustments. r denotes the cointegration rank. The p -values, which are reported in PcGive, are based on the approximations to the asymptotic distributions derived by Doornik (1998). It should be noted that the inclusion of impulse dummies in the VAR affects the asymptotic distribution of the reduced rank test statistics and therefore the critical values are only indicative. However, the bias induced by such deterministic components is supposed to be minor as only three impulse dummies are included in our case. The asterisk * and ** denote rejection of the null hypothesis at the 5 per cent and 1 per cent significance levels, respectively. The weak exogeneity tests, which are asymptotically distributed as $\chi^2(1)$ under the null [see Johansen (1995)], are calculated under the assumption that $r = 1$.

The trace statistics without small sample adjustments (λ_{trace}) in the case of a linear trend restricted to lie in the cointegrating space rejects the null of no cointegration at the 5 per cent significance level, but the null of at most one cointegrating vector is not rejected. Based on the trace statistics with small sample adjustments (λ^a_{trace}), the null of no cointegration is a borderline case at the 10 per cent level. In the case of no linear trend, both trace statistics suggest that there is one cointegrating vector between p , pi and ulc at the 5 per cent level. Assuming the rank to be unity, a likelihood ratio test of model reduction [see Doornik and Hendry (2001, p. 51)] from a cointegrated VAR with the linear trend to a cointegrated VAR without the linear trend, yields $\chi^2(1) = 1.314$ with a p -value of 0.252. So the linear trend is *insignificant* in the reduced rank VAR. Also, an I(2) analysis 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

¹⁰ These tests are available on <http://people.ssb.no/bou>.

zero.¹¹ We therefore proceed the I(1) analysis under the assumption of one cointegrating vector (i.e., $r=1$) without a linear trend restricted to lie in the cointegrating space. The 2 log likelihood value of the reduced rank VAR, to be used in Section 6, is estimated to 1992.84.

The estimate of the *unrestricted* cointegrating vector (normalised on p) reported in Table 2 is interpretable as an equation for consumer prices in Norway as the estimated coefficients for import prices and unit labour costs are economically reasonable with expected signs. Besides, the results of the weak exogeneity tests suggest that the cointegrating vector enters the Δp_t -equation only. We also notice that the sum of the estimated coefficients inherent in the vector is not far from unity, as predicted by theory. To complete the cointegration analysis, we thus tested for, and could not reject, the existence of long run homogeneity between p , pi and ulc . Imposing the homogeneity restriction and weak exogeneity of ulc gives $\chi^2(2)=2.221$ (with a p -value of 0.329) and the following *restricted* estimate of the cointegrating vector (normalised on p):¹²

$$(7) \quad \hat{p} = \hat{\gamma}_0 + 0.416pi + 0.584ulc$$

(0.097)

The recursively estimated parameter of ulc in (7) is fairly constant and a sequence of $\chi^2(2)$ test statistics confirms the validity of the homogeneity restriction and the weak exogeneity of ulc for any sample ending between 1990 and 2001 (the data until 1990Q1 were used for initialisation). To sum up, we interpret (7) as a long-run consumer price equation that corresponds well with the theory of mark-up pricing and the fact that for a small open economy like the Norwegian, foreign prices are expected to matter somewhat, i.e., producers have some market power in the product markets.¹³ In particular, equation (7) implies an estimate of $m_1 = -0.7$, its economic interpretation being that when prices of foreign competitors increase by one per cent, the mark-up increases by 0.7 per cent. Intuitively, this is a reasonable result since higher foreign prices allow domestic firms to increase their mark-up in a situation with some market power.

¹¹ Results from the I(2) analysis are available on <http://people.ssb.no/bou>.

¹² Instead, imposing the homogeneity restriction and weak exogeneity of pi is almost rejected by the data. On the other hand, imposing the homogeneity restriction and weak exogeneity of pi and ulc jointly are clearly rejected.

¹³ The estimation results in (7) are in line with previous findings on Norwegian data. Already, Aukrust (1977, p. 123) points 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. Interestingly, Bårdsen, Fisher and Nymoen (1998) estimate a corresponding long run open economy price mark-up equation for the UK with somewhat different information set and a different sample period.

4. Single-equation tests of the benchmark NKPC

The standard practice in empirical evaluation of the NKPC is to test the significance of the forward term of inflation by means of different single-equation estimation techniques, namely that of 2SLS and GMM. Model (4) cannot be estimated directly due to the fact that $E_t \Delta p_{t+1}$ involves unobservable expectations of inflation one period ahead, and consequently is a latent variable. Hence, we follow common practice and replace it by the realised value, Δp_{t+1} , in order to derive the estimating equation:

$$(8) \quad \Delta p_t = \delta \Delta p_{t+1} - \lambda eqcm_t + e_t,$$

where $eqcm_t = p_t - \gamma_0 - \gamma_1 p_t - (1 - \gamma_1) ulc_t$, $e_t = \varepsilon_t - \delta \eta_{t+1}$ is the gross error, ε_t is the stochastic error term from (4), $\eta_{t+1} = \Delta p_{t+1} - E_t \Delta p_{t+1}$ is the expectation error in predicting future inflation and δ is the coefficient attached to the forward-looking term. We notice that estimating (8) by means of the "errors in variables" method induces first order moving average errors by construction since ε_t and η_t are correlated, see e.g. Bårdsen *et al.* (2002b) for details. Estimated serial correlation thus corroborates forward-looking behaviour, but it may also be a sign of model misspecification, as discussed in Nymoen (2003) and Bårdsen *et al.* (2004). Nevertheless, the potential need to correcting for serially correlated errors motivates the use of GMM rather than 2SLS as a single-equation estimation method. Under rational expectations and the assumption that the error term ε_t is white noise, it follows that

$$(9) \quad E_{t-1} \{ (\Delta p_t - \delta \Delta p_{t+1} + \lambda eqcm_t) z_{t-1} \} = 0,$$

where z_{t-1} is a vector of instruments dated $t-1$ and earlier. The orthogonality conditions given by (9) provide the basis for estimation with GMM. Since the instrument set includes only lagged variables, we implicitly treat $eqcm_t$ as endogenous.¹⁴ A potential shortcoming of our approach, as pointed out by e.g. Galí *et al.* (2005), is that estimation of (8) 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.¹⁵

¹⁴ This could be justified by the fact that $p_t = \Delta p_t + p_{t-1}$, and thus includes the left hand side variable in (8).

¹⁵ See also Rudd and Whelan (2005) who propose the same bias with GMM if estimates of the closed (or reduced) form of the NKPC (obtained by solving out for expected future inflation) are significantly different from those obtained from estimating the structural form, cf. equation (8) in our case. However, Galí *et al.* (2005) show that when one estimates the closed form equation in a way that incorporates the restrictions of the structural form, then the parameter estimates are almost identical to those obtained by estimating the structural form. Lindé (2005) argues by means of a Monte Carlo experiment and simulated data from a small macroeconomic model including a NKPC, an aggregate demand equation and an interest policy rule that GMM may produce the opposite small sample bias, namely support for a backward-looking Phillips curve, although the true Phillips curve is highly forward-looking. In essence his argument is that the downward bias in the coefficient of the

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 and too general corrections for serial correlation seriously weaken the power of the J test, thus obscuring specification problems and distorting GMM based inference. We address these issues below by using different heteroskedasticity autocorrelation consistent (HAC) estimators (allowing only few periods of serial correlation) and relatively few instruments that may also play a role as additional explanatory variables in (9).

Throughout the evaluation of (8), we use lagged inflation $(\Delta p_{t-1}, \Delta p_{t-2})$, lagged import price growth $(\Delta pi_{t-1}, \Delta pi_{t-2})$, lagged growth in unit labour costs $(\Delta ulc_{t-1}, \Delta ulc_{t-2})$, lagged equilibrium correction term $(eqcm_{t-1})$, the constant term and dummies for seasonal effects $(sdum_t)$ as potential instruments.¹⁶ 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 as 40 instruments. Equation (10) presents GMM results of (8) for the period 1983Q1–2001Q1 when iterating over both coefficients and weighting matrix, with fixed bandwidth based on Newey and West (1987).¹⁷

$$(10) \quad \Delta \hat{p}_t = -0.062 \Delta p_{t+1} - 0.092 eqcm_t + 0.045 + sdum_t$$

$$(0.131) \quad (0.011) \quad (0.005)$$

$$\text{GMM, } T = 73 \text{ (1983Q1–2001Q1)}$$

$$\chi^2_J(5) = 5.406 [0.368], \hat{\sigma} = 0.0046$$

where $\hat{\sigma}$ denotes the estimated residual standard error and χ^2_J is the J -statistics of the validity of the overidentifying instruments. An intercept is freely estimated in (10) 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. Noteworthy, the fact

forward-looking term arises when inflation is not intrinsically persistent, but inherits persistence in the model via inertia in the aggregate demand (i.e., the forcing variable in the NKPC) and the policy rule. In contrast, Lindé (2005) claims on the basis of his model settings that FIML is superior to GMM for the purpose of obtaining reliable estimates of the NKPC parameters. Galí *et al.* (2005) find this claim doubtful and point out that Lindé's (2005) Monte Carlo exercise is heavily distorted in favour of FIML. Nonetheless, inspired by Lindé's (2005) findings, we estimate our NKPC using a system approach in Section 6 as a comparable exercise to the single-equation estimates herein.

¹⁶ The dummies $D86Q3$ and $D96Q1$ used in the VAR in Section 3 to account for special events in the economy are not included in the set of instruments to simplify the model and facilitate GMM estimation in Eviews5.

that the estimated constant comprises both the mean of the cointegration relationship, elements of short run dynamics as well as being influenced by the scaling of the variables (see the Appendix) makes the level of the mark-up as such non identifiable. We observe that the equilibrium correction term is highly significant, an aspect of (10) which supports our benchmark NKPC. However, the numerical and statistical insignificance of the forward-looking term contradict the theoretical model. So deleting the forward-looking term, (10) reduces to a simple partial adjustment model in the CPI level. Our findings stand in sharp contrast to several existing studies, 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).

We now judge the robustness of these results with respect to different choices made about the GMM estimation strategy. First, we investigate any distortions to the GMM estimates caused by the use of different HAC estimators. When instead using the data dependent bandwidth selection method proposed by Andrews (1991), we get

$$(11) \quad \Delta \hat{p}_t = -0.044 \Delta p_{t+1} - 0.090 eqcm_t + 0.044 + sdum_t$$

$$(0.134) \quad (0.011) \quad (0.005)$$

$$\text{GMM, } T = 73 \text{ (1983Q1–2001Q1)}$$

$$\chi^2_J(5) = 5.229 [0.389], \hat{\sigma} = 0.0046$$

Compared to (10), the estimated coefficients in (11) are virtually unchanged and the J -statistics does not seem to be much affected by the different HAC estimators. Also, we do not obtain marked differences in the GMM estimates of (8) when using the variable Newey-West (1994) method.

Next, we investigate any sensitivity with regards to the set of instruments. Typically in applied work, a hybrid version of the NKPC, with both forward-looking and backward-looking elements, is needed in order to explain inflation dynamics adequately. By including the first lag of inflation from the list of instruments as an additional regressor, we obtain (with fixed bandwidth based on Newey and West (1987))

¹⁷ The Newey-West fixed bandwidth is based solely on the number of observations in the sample, which in our case is given by $\text{int}\left[4(73/100)^{2/9}\right] = 3$.

$$(12) \quad \Delta \hat{p}_t = -0.038 \Delta p_{t+1} + 0.088 \Delta p_{t-1} - 0.077 eqcm_t + 0.038 + sdum_t$$

(0.142) (0.143) (0.024) (0.011)

$$\text{GMM, } T = 73 \text{ (1983Q1–2001Q1)}$$

$$\chi^2_J(4) = 5.326 [0.255], \hat{\sigma} = 0.0045$$

As seen, the first lag of inflation is far from being significant and the results with respect to the forward-looking term, the equilibrium correction term and the J -statistics respond little to this respecification of the model. We reach the same conclusion when conducting similar exercises with all the other variables included in 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 the case of our benchmark model.

Finally, we investigate robustness with respect to single-equation estimation methods. Since the coefficient of the forward-looking term δ is found to be insignificantly different from zero in all regressions above, which is consistent with no serial correlation of first order (at least in theory) in the data, we may apply a simpler estimation method than GMM. We employ 2SLS (with no serial correlation correction), which produces¹⁸

$$(13) \quad \Delta \hat{p}_t = -0.074 \Delta p_{t+1} - 0.099 eqcm_t + 0.047 + dummies_t$$

(0.189) (0.019) (0.008)

$$2SLS, T = 73 \text{ (1983Q1–2001Q1), } \hat{\sigma}_{IV} = 0.0038$$

$$\chi^2_{\text{Sargan}}(5) = 4.777 [0.444]$$

$$AR_{1-1}: F(1, 65) = 1.580 [0.213]$$

$$AR_{1-2}: F(2, 64) = 0.799 [0.454]$$

$$AR_{1-5}: F(5, 61) = 1.006 [0.422]$$

$$ARCH_{1-4}: F(4, 58) = 0.257 [0.905]$$

$$NORM: \chi^2(2) = 1.981 [0.371]$$

$$HET: F(8, 57) = 1.728 [0.112]$$

¹⁸ The 2SLS estimation is conducted using PcGive 10.3, cf. Hendry and Doornik (2001) and Doornik and Hendry (2001). We use the same instrument set as in the GMM analyses and continue to treat $eqcm_t$ as endogenous. The battery of test statistics reported below (13) is as follows: AR_{1-1} , AR_{1-2} and AR_{1-5} are the Harvey (1981) test for first order, until second order and until 5th order residual autocorrelation, respectively, $ARCH_{1-4}$ is the Engle (1982) test for until 4th order autoregressive conditional heteroskedasticity in the residuals, $NORM$ is the normality test described in Doornik and Hansen (1994), HET is the White (1980) test for residual heteroskedasticity due to cross products of the regressors and χ^2_{Sargan} is the Sargan (1964) specification test of the validity of the instruments. The regressor labelled $dummies_t$ includes the impulse dummies $D86Q3$ and $D96Q1$ from the VAR in Section 3 in addition to the seasonal dummies. We should emphasise that the impulse dummies are only needed to render normality distributed residuals and that the coefficient estimates in (13) are not influenced by these variables.

We observe that the 2SLS coefficient estimates are practically identical to the GMM estimates reported above. The p -value of the Sargan test is evidence against misspecification with respect to the predictive power of the set of instruments. Adding each of the variables in the instrument set individually to the right hand side of (13) make them, as before, insignificant. Contemporaneous effects from growth in import prices and unit labour costs are also insignificant as regressors (treated endogenously and exogenously) in (13). Besides, the diagnostic tests indicate that (13) is well specified econometrically, as no serial correlation (neither first nor higher order) and heteroskedasticity are revealed in the residuals. No detectable residual autocorrelation is by itself a finding that contradicts the theory that agents are acting in accordance with the NKPC, which otherwise could have been the case if the residuals of (13) had exhibited first order serial correlation.

As a further test of misspecification dictated by existing results from several decades of empirical modelling of inflation, we enlarge the information set with two different measures of capacity utilisation, namely the output gap and the unemployment rate.¹⁹ Interestingly, none of these variables are significant in (13) as separate regressors with either its first or second lag in addition to using them as instruments. Moreover, adding these variables one by one as contemporaneous variables treating them either as endogenous or exogenous does not alter the insignificance status of these variables in (13).²⁰ Hence, measures of capacity utilisation do not seem to be important variables when testing forward-looking behaviour in the context of our NKPC. At last, motivated by the findings in Bårdsen and Nymoen (2001), we add contemporaneous growth in electricity prices as an additional explanatory variable. Apart from the fact that electricity price growth is only significant when treated as exogenous, the results in (13) is still unaffected by this model respecification. We also notice that lags of electricity price growth are far from being significant as separate regressors.

5. Single-equation tests of the standard NKPC

So far we have examined the empirical counterpart of (4). Here, we want to relax our modelling strategy and evaluate different versions of (5), but augmented with open economy features, as alternative models of inflation in the spirit of the standard NKPC. We may show the main difference

¹⁹ The output gap variable is defined as the percentage deviation of gross domestic product from its underlying trend estimated by the Hodrick-Prescott filter, see the Appendix for details. The interpretation of the respecified model with such capacity measures may be a Phillips curve derived from price setting with the equilibrium mark-up influenced by the state of the business cycle in addition to relative prices already captured through the eqm_t term, cf. e.g. Batini *et al.* (2005) for details.

²⁰ These results are available on <http://people.ssb.no/bou>.

between the benchmark model and the standard model by noting that (4) can be *reparameterised* as a forward-looking equilibrium correction model that reads as (ignoring the constant and the error term for the present purpose)

$$(14) \quad \Delta p_t = \eta_1 E_t \Delta p_{t+1} + \eta_2 \Delta p_i_t + \eta_3 \Delta ulc_t - \eta_4 [p - \gamma_1 p_i - (1 - \gamma_1) ulc]_{t-1},$$

where $\eta_1 = \delta / (1 + \lambda)$, $\eta_2 = \lambda \gamma_1 / (1 + \lambda)$, $\eta_3 = \lambda (1 - \gamma_1) / (1 + \lambda)$ and $\eta_4 = \lambda / (1 + \lambda)$. By letting Δulc_t and Δp_i_t be marginal costs and open economy features, respectively, equation (5) in the present situation differs from (14), or equivalently (4), in the former's exclusion of the deviation of the price level from its target as a separate forcing variable. Estimating (14) without the term in brackets by means of GMM with fixed bandwidth based on Newey and West (1987), we obtain the following open economy NKPC in line with the standard specification:

$$(15) \quad \Delta \hat{p}_t = 0.969 \Delta p_{t+1} + 0.086 \Delta p_i_t + 0.077 \Delta ulc_t + 0.0014 + sdum_t$$

(0.076) (0.049) (0.015) (0.0011)

$$\text{GMM, } T = 73 \text{ (1983Q1–2001Q1)}$$

$$\chi^2(4) = 4.703 [0.453], \hat{\sigma} = 0.0054$$

The same set of instruments used for identification of (8) is also applied here and both Δp_i_t and Δulc_t are treated as endogenous variables in the estimation. We recognise that the coefficient associated with the forward-looking term now becomes highly significant and close to unity in numerical value, a common finding in several studies. However, the statistical adequacy of (15) for testing purposes is questionable as it suffers from an omitted variable bias. One simple way of obtaining a statistically adequate model in the present situation is to include the deviation of the price level from its target (lagged one period) from the list of instruments as an explanatory variable in (15), cf. Mavroeidis (2005). As we have seen, such a model is nothing else, but a reparameterisation of (4) which was rejected by the data in the previous section.

Another model motivated by the standard specification of the NKPC replaces the contemporaneous growth in unit labour costs in (15) with the output gap (GAP_t) as a proxy for excess demand or marginal costs. In this case, we obtain by means of 2SLS without serial correlation correction

$$(16) \quad \Delta \hat{p}_t = 1.043 \Delta p_{t+1} - 0.00036 GAP_t + 0.064 \Delta pi_t - 0.0044 + dummies_t$$

$$(0.269) \quad (0.00037) \quad (0.097) \quad (0.0022)$$

$$2SLS, T = 73 (1983Q1-2001Q1), \hat{\sigma}_{IV} = 0.0051$$

$$\chi^2_{Sargan} (3) = 3.597 [0.308]$$

$$AR_{1-1}: F(1, 64) = 7.425 [0.008]$$

$$AR_{1-2}: F(2, 63) = 4.620 [0.013]$$

$$AR_{1-5}: F(5, 60) = 3.745 [0.005]$$

$$ARCH_{1-4}: F(4, 57) = 0.700 [0.595]$$

$$NORM: \chi^2(2) = 3.850 [0.146]$$

$$HET: F(10, 54) = 4.718 [0.000]$$

where lagged inflation ($\Delta p_{t-1}, \Delta p_{t-2}$), lagged import price growth ($\Delta pi_{t-1}, \Delta pi_{t-2}$), lagged output gap (GAP_{t-1}, GAP_{t-2}) and the constant term, dummies for seasonal effects and special events in the economy (labelled *dummies*) are applied as instruments. As before, both GAP_t and Δpi_t are treated as endogenous variables in the estimation. Several interesting aspects are seen from the results. First, we recognise again that the coefficient attached to the forward term is close to unity and now looks highly significant. However, both the output gap and the import price growth are statistically insignificant, an aspect of (16) which contradicts the NKPC. In line with what is commonly reported, we do not estimate sizeable coefficients on the driving variables, see e.g. Fuhrer (2005). Noteworthy, we obtain similar parameter estimates as in (16) using GMM. Using GMM the forward term vanishes when (16) is enlarged, and significantly so, with the first lag of inflation as a separate regressor, whereas the estimates by and large are unaffected by such a hybrid version using 2SLS.²¹ The diagnostic tests of (16) indicate presence of severe residual autocorrelation, which may, as noted earlier, be a sign of econometric misspecification due to omitted variables rather than the “errors in variables” method. Accordingly, equation (16) provides evidence both in favour of and against the alternative NKPC (5) as an empirical model of Norwegian inflation dynamics. The uncertainty about whether the detected autocorrelation really is a symptom of misspecification or in fact is induced by the replacement of $E_t \Delta p_{t+1}$ with Δp_{t+1} in the estimation, calls for further evaluations of (16) so as to make reliable judgment of its empirical success.

To this end, we take advantage of the above evidence of long run price determination in order to test its implications for the encompassing property of (16). Drawing closely on Bårdsen *et al.* (2004), the encompassing procedure in our case may be outlined as follows: Assume that there exists a set of

²¹ Bårdsen *et al.* (2005, p. 144) achieve similar results as in (16), albeit with wage share proxying marginal costs, using GMM and Norwegian data over the sample period 1972Q4 – 2001Q1. Also, Bårdsen *et al.* (2005, p. 184) find the forward-term to be insignificant when adding the first lag of inflation into their model and using GMM.

instruments $z = [z1 \ z2]$, where the subset $z1$ is sufficient for identification of (16) and the subset $z2$ is defined by empirical findings in Section 3. Then, using $z1$ as identifying instruments, estimate the model $\Delta p_t = \delta E_t \Delta p_{t+1} + \omega_1 GAP_t + \omega_2 \Delta p_i_t + \omega_3 z2_t$ under the assumption of rational expectations about future inflation. Under the null hypothesis that the NKPC is the correct model, then $\omega_3 = 0$ is implied. Hence, non-rejection of $\omega_3 = 0$ confirms the NKPC as an empirical model of inflation. On the other hand, non-rejection of $\delta = 0$ and rejection of $\omega_3 = 0$, the encompassing implication of the NKPC is rejected. We let the equilibrium correction term, $eqcm_t = p_t - 0.416 p_i_t - 0.584 ulc_t$, be the additional instrument ($z2_t$) in our case. The additional instrument represents information about price determination that was not used in the estimation of (16). The encompassing principle now states that, if rational expectations indeed are correct, then the extension of the information set should not influence the significance of Δp_{t+1} in (16). The estimation results (using 2SLS without autocorrelation correction) are presented in equation (17).²²

$$(17) \quad \Delta \hat{p}_t = 0.021 \Delta p_{t+1} + 0.00036 GAP_t + 0.096 \Delta p_i_t - 0.067 eqcm_{t-1} + 0.029 + dummies_t$$

(0.443)
(0.00035)
(0.066)
(0.034)
(0.016)

$$2SLS, \quad T = 73 \text{ (1983:1–2001:1)}, \quad \hat{\sigma}_{IV} = 0.0039$$

$$\chi^2_{\text{Sargan}}(3) = 3.941 \text{ [0.268]}$$

$$AR_{1-1}: F(1, 63) = 0.208 \text{ [0.650]}$$

$$AR_{1-2}: F(2, 62) = 0.106 \text{ [0.899]}$$

$$AR_{1-5}: F(5, 59) = 0.413 \text{ [0.838]}$$

$$ARCH_{1-4}: F(4, 56) = 0.107 \text{ [0.980]}$$

$$NORM: \chi^2(2) = 0.619 \text{ [0.734]}$$

$$HET: F(12, 51) = 1.289 \text{ [0.254]}$$

Compared to (16), we notice that the estimated coefficient of the forward term in (17) reduces from 1.04 to 0.02, and thus becomes insignificant. By contrast, the equilibrium correction term, which should be of no importance if (16) is the true model, is significant at the 5 per cent level. Moreover, there is no longer sign of autocorrelation in the residuals. These findings taken together support the hypothesis that (16) in fact is misspecified, its underlying cause being omitted variables. Thus, the outcomes of the encompassing test provide clear evidence against (16) as an alternative open economy NKPC for Norwegian inflation dynamics.²³

²² The same set of instruments used for identification of (16) is also applied here. The results in (17) are virtually unchanged when applying GMM.

²³ Noticeably, the rejection of the encompassing properties of (16) is regardless of whether $eqcm_t$ (treated as endogenous) or $eqcm_{t-1}$ is included as a separate regressor in the model.

In light of the findings in this section, we may claim that applying our modelling strategy to (4) appears to be a good strategy when evaluating open economy NKPC for Norway. We conclude from the range of estimation results reported so far that the numerical and statistical insignificance of the forward-looking element is a robust feature and that open economy versions of the NKPC do not seem to fit Norwegian data well. This conclusion is based on the premise that the testing of the forward term is conducted within statistically adequate models. But, as emphasised by Bårdsen *et al.* (2004), Lindé (2005) and Mavroeidis (2005) among others, system estimation of NKPC may have clear advantages over single-equation alternatives. Although no particular sign of model misspecification associated with our GMM and 2SLS procedures are revealed, we now turn to test various specifications of the NKPC using (i) the Johansen and Swensen (1999, 2004) cointegrated VAR approach and (ii) the reduced rank VAR in accordance with the results in Section (3) to further check the robustness of the single-equation results.

6. Cointegrated VAR tests

The basic idea behind the procedure suggested by Johansen and Swensen (1999, 2004) is to start with a well specified VAR and test, using a likelihood ratio test, the implications of the NKPC on the coefficients of the VAR. Therefore, one has to work out the maximum likelihood estimator of the coefficients, both with and without the expectation restrictions imposed, in order to construct a likelihood test. It is essential that the expectation restrictions are *exact*. Although (4) contains errors (ε_t), it implies an *exact* rational expectation formulation. To see this, we date (4) at time $t+1$ and take the conditional expectation given the information in $x_t = (p_t, pi_t, ulc_t)$ at time t to get the *exact* restrictions $E_t \Delta p_{t+1} = \delta E_t \Delta p_{t+2} - \lambda [E_t p_{t+1} - \gamma_0 - \gamma_1 E_t pi_{t+1} - (1 - \gamma_1) E_t ulc_{t+1}]$. However, restrictions of this form are not immediately amenable to the test procedure in Johansen and Swensen (1999, 2004), which only works for *one-step-ahead* expectations. We therefore treat a simplification of (4) where the error term (ε_t) is excluded. The equation (4) may then be expressed as (ignoring the constant until further notice)

$$(18) \quad c_1' E_t [x_{t+1}] + c_0' x_t + c_{-1}' x_{t-1} = 0,$$

where $c_1 = (\delta, 0, 0)'$, $c_0 = -(1 + \delta + \lambda, -\lambda\gamma_1, -\lambda\gamma_2)'$ and $c_{-1} = (1, 0, 0)'$. We remark that (18) for fixed values of δ , λ and $\gamma = \gamma_1 = 1 - \gamma_2$ contains restrictions involving the conditional expected value of the observations *one-step-ahead* and the present and lagged observed values. Having specified the

information set, the conditional expectation $E_t[x_{t+1}]$ can be worked out and the restrictions stated explicitly. These restrictions entail constraints on the long run relationship as well as on the short run parameters on the VAR model, see Boug *et al.* (2006) for details. The approach in Johansen and Swensen (1999, 2004) requires that these restrictions contain no unknown parameters. In (4) the parameter δ , λ and γ are unknown, but for fixed values we can find the maximum value of the likelihood $L(\delta, \lambda, \gamma)$. The maximum likelihood estimator of δ , λ and γ can then in principle be found by employing an appropriate numerical optimizer.

In the present case there is an additional problem. The reduced rank VAR specified in Section 3 contains seasonal dummies and three impulse dummies. The restrictions deduced by the procedure we just described also involve constraints on the deterministic part of the model. Since we focus on the non-deterministic part, we simply drop some restrictions and test for the remaining. The question that then naturally arises is what the substantial interpretation of the hypothesis is. As explained in Boug *et al.* (2006), dropping restrictions on the deterministic part means that (18) can be formulated as

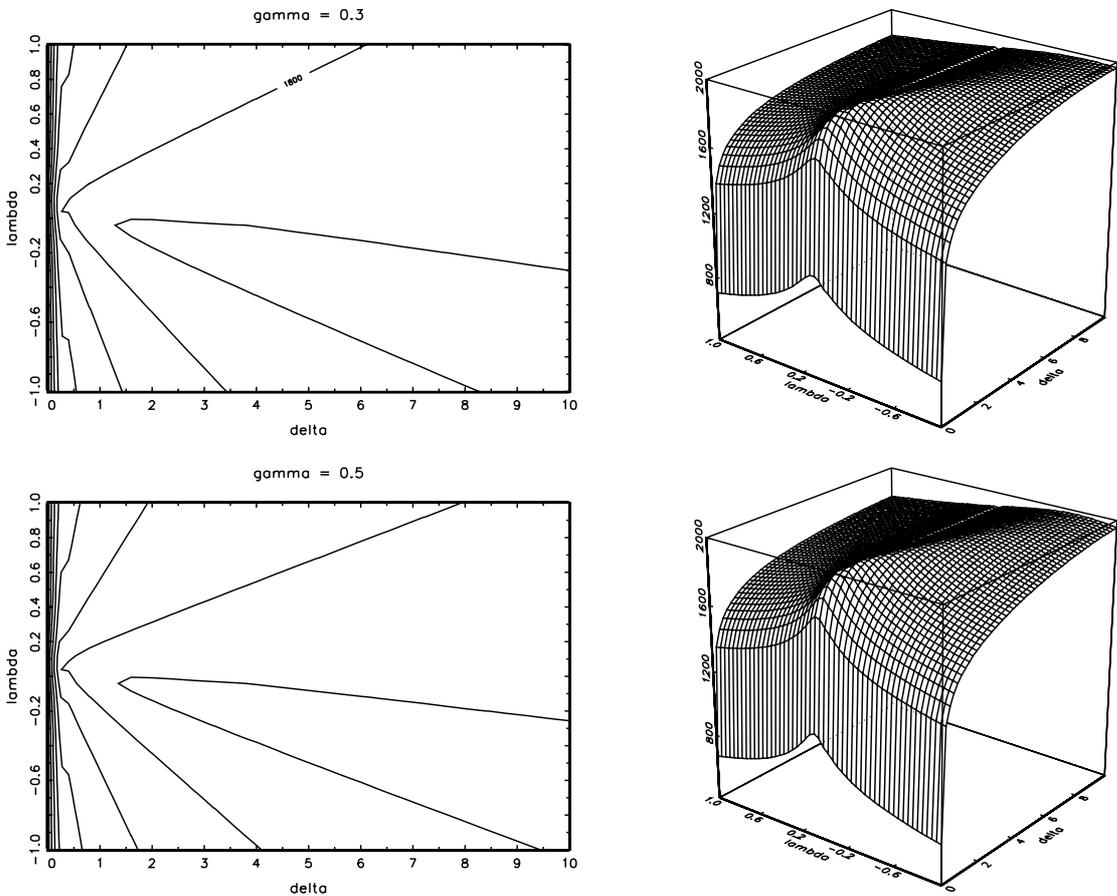
$$(19) \quad c_1' E_t[x_{t+1} - \phi D_{t+1}] = -c_0' x_t - c_{-1}' x_{t-1},$$

where D_t is the vector with deterministic variables (seasonal dummies, impulse dummies and the constant term). A possible interpretation is that an unbiased forecast of the non-deterministic part of the future observation can be constructed from the present and previous observations. In Figures 2 and 3 values of the $2 \log L(\delta, \lambda, \gamma)$ are plotted for $0 < \delta$ and $-1 < \lambda < 1$ and for fixed values of γ and γ_0 . Notably, the latter two parameters are determined so that the corresponding likelihood surfaces are close being as large as possible. Before inspecting the values closer it is worth pointing out that setting $\lambda = 0$ means that the cointegration vector must equal zero, which implies that the rank is equal to zero. This hypothesis was rejected in Section 3. It will therefore be rejected even more strongly when we, as here, impose additional restrictions.

In Figure 2 the deterministic part of the NKPC model is specified as containing both the dummies and the constant term. The upper level curve corresponds to 1950, and the difference between the values of the two level curves is 150. When $\gamma = 0.3$, the maximum value is 1978.68 and corresponds to $\lambda = -0.20$ and $\delta = 3.4$. Similarly, when $\gamma = 0.5$ the maximum value is 1980.85 and corresponds to $\lambda = -0.20$ and $\delta = 3.8$. The likelihood ratio tests can now be carried out by comparing these values to the 2 log likelihood value of the reduced rank VAR model with rank equal to unity, i.e., the value 1992.84 (from Section 3). With δ , λ and γ estimated and six degrees of freedom the $-2 \log$ likelihood ratio test

statistics are 11.99 and 14.16 corresponding to p -values between 0.025 and 0.05. Hence, the restrictions involved in (19) are rejected at conventional significance levels. Also, we notice that the estimated values of δ and λ are far outside their parameter regions with reasonable economic interpretations, i.e., $0 < \delta < 1$ and $\lambda > 0$. Likelihood ratio tests with δ and λ equal to values within these regions indicate, as we can see from Figure 2, that (19) is overwhelmingly rejected. Specifically, we end up with the boundary values $\hat{\delta} = 1$ and $\hat{\lambda} = 0$, their economic interpretation in terms of the loss function (1) being that all the weight is put on penalising price changes.

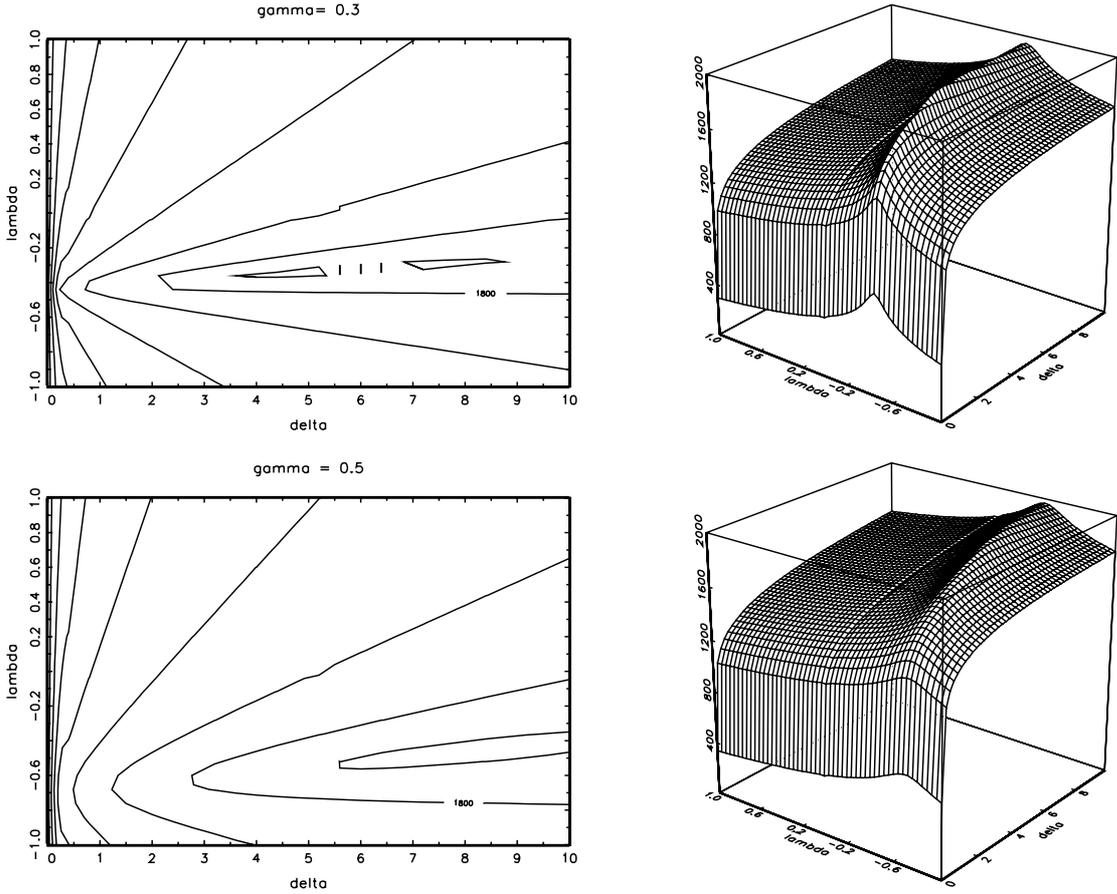
Figure 2. $2 \log L(\delta, \lambda, \gamma)$. D_t contains dummies and the constant term



In Figure 3 the deterministic part of the NKPC model is specified as containing the dummies only and the intercept takes the value of -0.2 . Thus, the hypothesis may be expressed as (19), where the deterministic variables (D_t) now only consist of dummies and contains no constant term. The general impression from Figure 2 still remains true. When $\gamma = 0.3$, the maximum value is 1972.74 and corresponds to $\lambda = -0.42$ and $\delta = 5.8$. We remark that the ridge of the surface declines as δ increases.

For $\delta= 10.0$ the value at the ridge is 1944.45. When $\gamma= 0.5$ the maximum value is 1979.34 and corresponds to the values $\lambda= -0.48$ and $\delta= 7.4$. Accordingly, the likelihood ratio test statistics are somewhat larger than when no restrictions on the constant term are imposed. The degrees of freedom of the statistics are only increased by one, so the picture is similar to the previous case.

Figure 3. 2 $\log L(\delta, \lambda, \gamma)$. D_t contains dummies only ($\gamma_0 = -0.2$)



Finally, we remark that the standard formulation (5) of the NKPC also corresponds to a situation where the cointegration vector equals zero. As explained earlier this means that it is not necessary to test the hypothesis, since it is already tested in determining the reduced rank as unity. We conclude from the wide range of cointegrated VAR tests of the NKPC that the single-equation findings reported above are not sensitive to the choice of estimation method. Regardless of estimation method applied, the NKPC is most likely at odds with the Norwegian data.

7. A dynamic backward-looking model

An alternative to the NKPC as a model of inflation is the standard backward-looking equilibrium mark-up price model, see e.g. Bårdsen *et al.* (1998), Bowitz and Cappelen (2001), Bårdsen *et al.* (2003), and Boug *et al.* (2006). To illustrate the main difference between the two models, consider the following *backward-looking* equilibrium correction representation of the reduced rank VAR established above (ignoring deterministic components and the error term for ease of exposition):

$$(20) \quad \Delta p_t = \sum_{i=1}^2 \psi_{1i} \Delta p_{t-i} + \sum_{i=0}^2 \psi_{2i} \Delta p_{t-i} + \sum_{i=0}^2 \psi_{3i} \Delta ulc_{t-i} - \eta [p - \gamma_1 pi - (1 - \gamma_1) ulc]_{t-1},$$

where the expression in brackets represents deviation from (7) as an equilibrium correction mechanism (*eqcm*). Abstracting from lagged difference terms, $\psi_{11} = \psi_{12} = \psi_{21} = \psi_{22} = \psi_{31} = \psi_{32} = 0$, equation (20) differs from the *forward-looking* equilibrium correction model in (14) in the former's exclusion of forward-looking expectations. It is an empirical question whether our data set is able to discriminate between the two rival models. As the NKPC is rejected by the data, we next try to develop a well-specified dynamic backward-looking model of inflation relying on a general-to-specific modelling strategy applied on (20).

Simplifications from the general to specific model is performed using PcGets1, see Hendry and Krolzig (2001). In accordance with the evidence above, the equilibrium correction term $eqcm = p - 0.416pi - 0.584ulc$ is included in the general model, lagged one period. Rather than imposing the impulse dummies from the VAR analysis a priori, we now let PcGets1 choose any such dummies based on an outlier detection procedure.²⁴ Briefly speaking, PcGets1 first test 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. Since PcGets1 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. PcGets1 picks the following specific model in our case:

²⁴ We specify the detection size of outliers (measured in standard deviation) to be 2.3 in PcGets1. Using the default value of 2.56 in PcGets1 does not provide normally distributed residuals in the final model.

$$\begin{aligned}
(21) \quad \Delta \hat{p}_t &= 0.183 \Delta p_{t-1} + 0.205 \Delta p_{t-2} + 0.024 \Delta ulc_t - 0.045 eqcm_{t-1} \\
&\quad (0.089) \quad (0.090) \quad (0.008) \quad (0.013) \\
&0.023 + 0.018 D86Q3 - 0.012 D96Q1 + sdum_t \\
&\quad (0.006) \quad (0.003) \quad (0.003)
\end{aligned}$$

$$\begin{aligned}
&OLS, T = 73 (1983Q1-2001Q1), \hat{\sigma} = 0.00321 \\
&AR_{1-5}: F(5, 59) = 1.765 [0.134] \\
&ARCH_{1-4}: F(4, 56) = 0.517 [0.723] \\
&NORM: \chi^2(2) = 1.583 [0.453] \\
&HET: F(12, 51) = 1.451 [0.174]
\end{aligned}$$

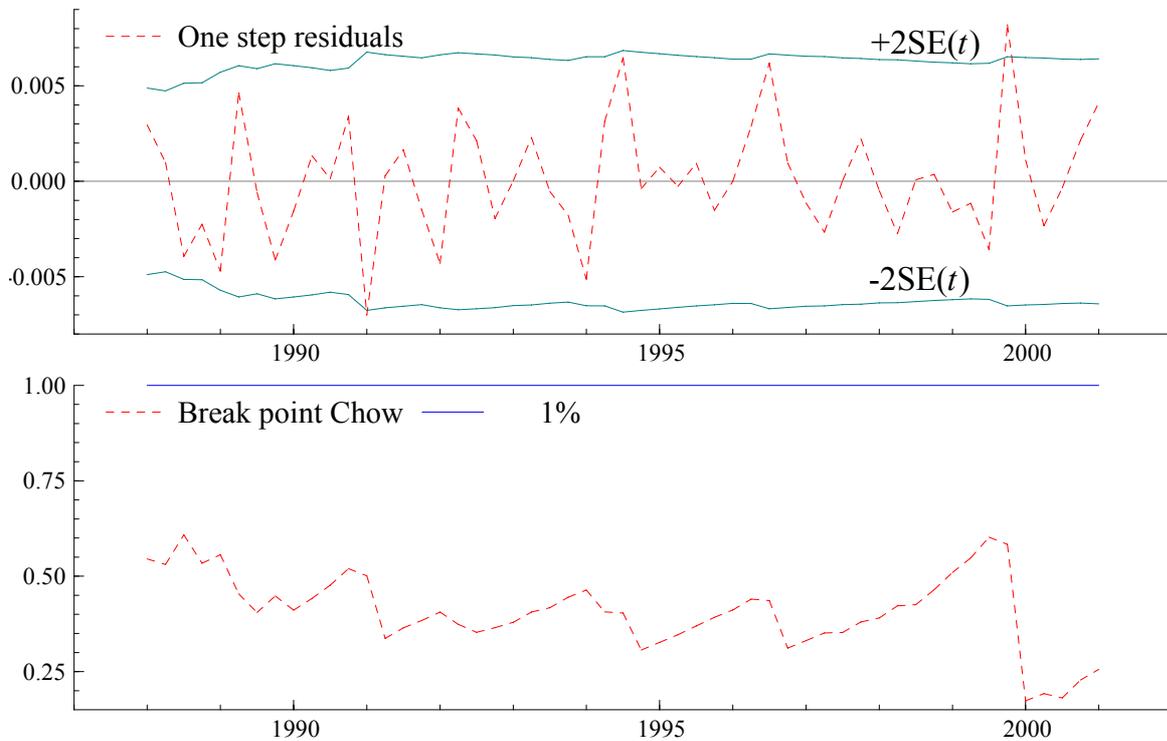
We notice that (21) is derived from a single equation analysis and not from a system. One may apply single equation analysis with the long run relationship(s) estimated and deduced from a VAR model in cases where the conditioning variables are error correcting, but are weakly exogenous for the short run parameters, see Urbain (1992) and Boswijk and Urbain (1997). The contemporaneous variable Δulc_t may be weakly exogenous for the short run parameters in (21) since the predicted counterpart from the VAR (in Section 3) invoked as a separate regressor yields a p -value of 0.813 (cf. Wu-Hausman test). The parameters in (21) are thus consistently estimated by OLS.

Several features about Norwegian inflation dynamics stand out from (21). First, we observe that the diagnostics tests reveal no symptom of misspecification and the economic variables entering the model are all highly significant. For instance, the $eqcm_{t-1}$ appears with a t -value of -3.351 , hence adding force to the results obtained from the cointegration analysis. The NKPC is variance encompassed by the backward-looking equilibrium correction model as the estimated residual standard error is reduced from 0.46 per cent, cf. equation (10), to 0.32 per cent. Interestingly, the dummies in 1986 and 1996 correspond with the chosen dummies in Section 3. We also see that the estimated short run effect on inflation from changes in unit labour costs is far less than its long run counterparts. Taken together, the minor short run effects and the small magnitude of the estimated loading coefficient (0.045), imply very slow consumer price adjustment in the face of shocks in unit labour costs.

Finally, empirical evidence of constancy of (21) may be judged from recursive test statistics shown in Figure 4, see Doornik and Hendry (2001). Neither the one-step residuals with ± 2 estimated equation standard errors (denoted $\pm 2 SE(t)$ in the figure) nor the sequence of break point Chow tests at the 1 per cent significance level indicate non-constancy. Similarly, aside from a minor fluctuation in the early

1990s in some of the estimates, all recursive estimates vary little, especially relative to their estimated uncertainty.

Figure 4. Recursive test statistics of (20)

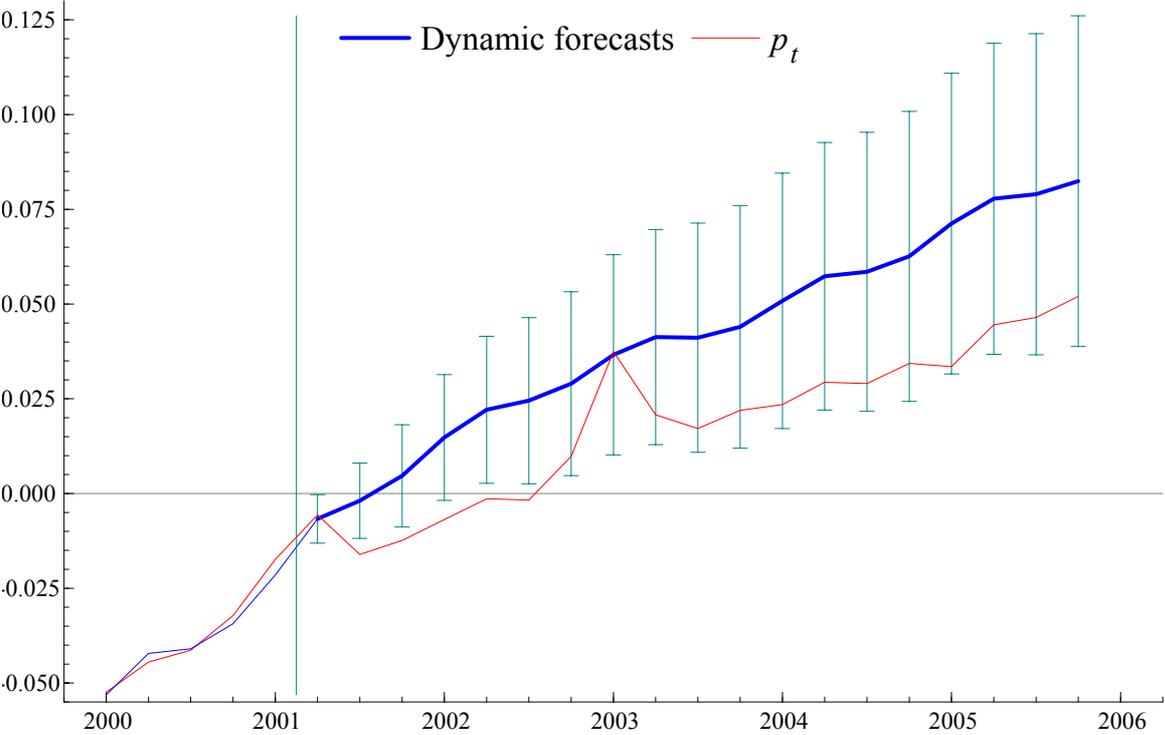


Having identified a well-specified backward-looking mark-up model *in-sample*, we study the *out-of-sample* forecasting performance to shed light on its robustness with respect to the regime shift in monetary policy in late March 2001. If the price setting behaviour has changed significantly following the introduction of inflation targeting, 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 (21), we employ nineteen quarters (2001Q2–2005Q4) of *out-of-sample* observations, including the period after the formal change in the monetary policy regime. Figure 5 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. A majority of the actual values of p_t stay within their corresponding confidence intervals over the forecasting period, with some exceptions occurring in the beginning of the forecasting period. The point in time in which the instability occurs *does not*, however, coincide with the time of the formal change in the monetary policy regime. Rather,

the actual value of p_t in 2001Q3, which is the first value staying outside the confidence intervals, and the values thereafter until 2002Q4 are influenced by the huge (and permanent) drop in the VAT rate of food from 24 to 12 per cent in July 2001.²⁵ Consequently, the dynamic forecasts overpredict the actual values of p_t during the entire forecasting period, except at 2003Q1 when the electricity prices increased substantially on a transitory basis. To take a closer look at these arguments for the forecasting failure of (21), we reestimate the model over the period 1983Q1–2001Q3 with an impulse dummy in 2001Q3 as a separate regressor to mop up the change in the VAT rate. Hence, seventeen observations are now available for forecasting. The reestimated equation is virtually unchanged from (21) with respect to both parameter estimates and diagnostics.

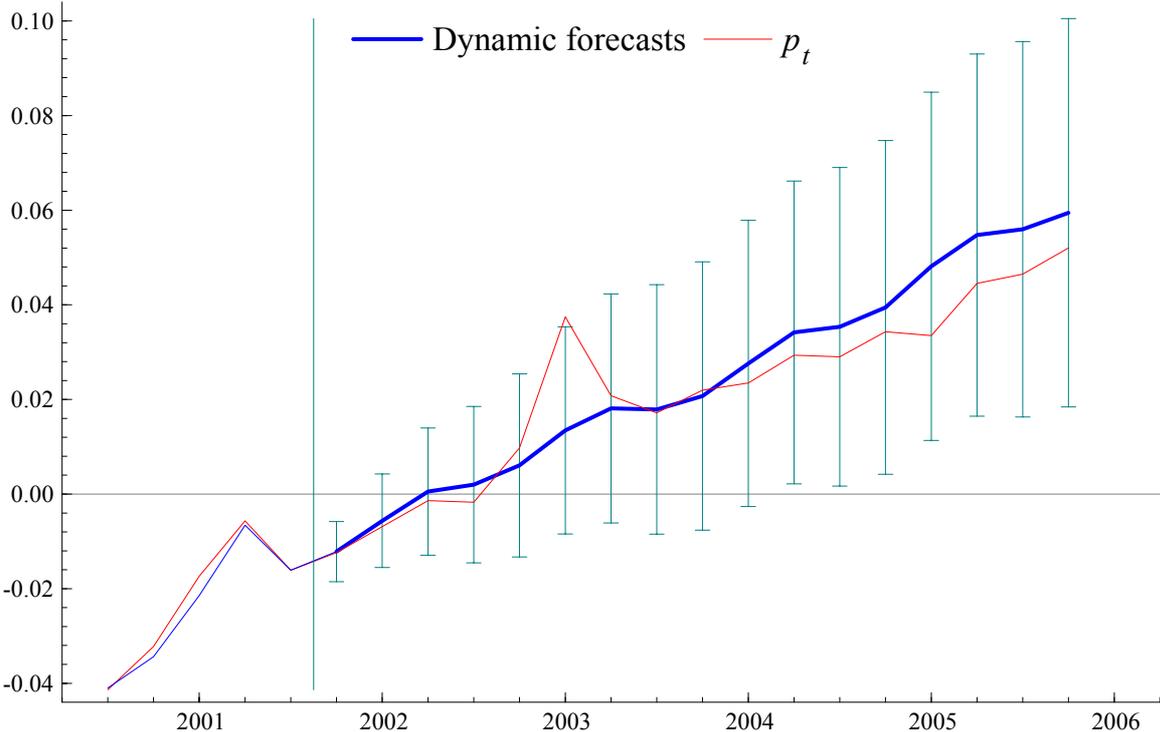
Figure 5. Actual values of p_t and dynamic forecasts with 95 per cent bands



²⁵ The items in the CPI affected by this VAT-change made up nearly 12 per cent of total CPI. Assuming full pass-through of the tax change to consumer prices implies a decline in the CPI of 1.2 percentage points, which is close to what we observe in Figure 5.

Figure 6 plots actual values of p_t together with dynamic forecasts when the reestimated model is used for forecasting. We observe that the forecasting failure of (21) now is eliminated, albeit the value of p_t in 2003Q1 is a borderline case. Thus, the general impression of the *out-of-sample* forecasting ability of (21) is reasonably good despite a *major* regime change in monetary policy. The regime robustness is inconsistent with the Lucas-critique being quantitatively important in our case.

Figure 6. Actual values of p_t and dynamic forecasts with 95 per cent bands



To sum up, the economic properties inherent in (21) seem consistent with the actual inflation persistence in Norway. We have seen that our data set is able to discriminate between the NKPC and the dynamic backward-looking mark-up model, the former being rejected in favour of the latter. Of course, following Bårdsen and Nymoen (2001), other and more elaborate dynamic mark-up models in which electricity prices and unemployment are allowed to play a role may exist.²⁶ 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.

²⁶ By enlarging the information set to include electricity prices and the unemployment rate, PcGets1 produces a more complex model than (21) with these variables influencing the short run dynamics of inflation in a significant way. Also, the forecasting ability of this model is even better than (21). These results are available on <http://people.ssb.no/bou>.

8. Conclusions

We have evaluated the empirical performance of open economy versions of the NKPC as models of Norwegian inflation by means of different estimation techniques, encompassing tests and tests based on cointegrated VAR models. Our starting point was the quadratic adjustment cost model of Rotemberg (1982) that relates the current inflation rate to expected future inflation and the difference between the actual price and the price target in levels. Thus, our NKPC model includes variables both in levels and differences that demands some care when applying econometric methods. We first established a cointegrating relationship between the price level and the target determined by explanatory factors dictated by economic theory for a small open economy. Based on the cointegrating relationship, we then tested the NKPC using single equation methods and various misspecification tests. The conclusion from this battery of tests is clear; there is no role for forward-looking behaviour in the model. The autocorrelation structure of the errors implied by the forward-looking model is also rejected. Thus, the NKPC reduces to a partial adjustment model of the consumer price level!

We have also estimated standard versions of the NKPC model found in the literature, but augmented with import prices as open economy features. When estimating these alternative NKPC as single behavioural equation and using standard test of significance, we found some supportive results for an open economy version of the NKPC. However, we demonstrated that these models are statistically inadequate for testing purposes as the forward term vanishes in respecified models that include relevant information from economic theory. We also found that the various specifications of the NKPC are most likely at odds with the data when evaluated within cointegrated VAR models. Finally, as opposed to the NKPC, we established a well-specified dynamic model of Norwegian inflation with backward-looking terms only. The dynamic model is fairly stable *in-sample* and forecasts well *post-sample* during a period of a major change in monetary policy, which is strong evidence in favour of this model. Thus, it is possible to improve on the simple partial adjustment model that resulted from testing the NKPC model.

We conclude that including forward-looking behaviour when modelling consumer price inflation in Norway seems unnecessary in order to arrive at a well-specified model by both econometric and economic theory selection criteria. Our findings are at odds with several empirical studies of the NKPC on both American and European data. In light of our results, it would be useful to re-evaluate the conclusions of existing studies using cointegration techniques and tests based on both single-behavioural models and cointegrated VAR models.

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Data definitions and sources

<i>P</i>	The official consumer price index (2002=1). Source: Statistics Norway.
<i>PI</i>	Implicit deflator of total imports (2002=1). Source: Statistics Norway.
<i>ULC</i>	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.
<i>GAP</i>	Output gap defined as the percentage deviation of seasonally adjusted real gross domestic product from its underlying trend estimated by the Hodrick-Prescott filter (with the smoothing parameter set at 40 000). Source: Statistics Norway.
<i>D84Q1</i>	Dummy variable used to account for an outlier and instability in the equation for Δulc_t in the VAR. It equals unity in the first quarter of 1984, zero otherwise.
<i>sdum_t</i>	A comprised dummy that includes seasonal dummies.
<i>dummies_t</i>	A comprised dummy that includes seasonal dummies and dummies for special events in the economy, see Bårdsen <i>et al.</i> (2002a) for details.

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