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Working Paper

Will it float? The New Keynesian Phillips curve tested on OECD panel data

Discussion Papers, No. 463

Provided in Cooperation with:
Research Department, Statistics Norway, Oslo

Suggested Citation: Bjørnstad, Roger; Nymoen, Ragnar (2006) : Will it float? The New Keynesian Phillips curve tested on OECD panel data, Discussion Papers, No. 463, Statistics Norway, Research Department, Oslo

This Version is available at:
https://hdl.handle.net/10419/192445

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Discussion Papers No. 463, July 2006
Statistics Norway, Research Department

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Will it float?
The New Keynesian Phillips curve tested on OECD panel data

Abstract: Galí, Gertler and Lòpez-Salido (2005), GGL, assert that the hybrid New Keynesian Phillips curve, NPC, is robust to different choices of estimation procedure and so some forms of specification bias. Specifically, the dominance of forward-looking behavior is robust according to GGL. We assess the NPC on a panel data set from OECD countries and find that the forward rate of inflation dominates also on the panel data set. However, when variables consistent with alternative inflation models are introduced in the models, the forward term is no longer significant. Such an outcome is predicted by the incomplete competition model of inflation, ICM, meaning that the ICM encompasses the NPC. The opposite does not apply. The non-robustness of the OECD panel data NPC is in alignment with a previous encompassing test on euro-area data, as well as tests on data from the UK and from Norway. GGL on their part do not test the robustness of the NPC features with respect to existing inflation models.

Keywords: New Keynesian Phillips Curve, forward looking price setting, panel data model, encompassing.

JEL classification: C23, C52, E12, E31

Acknowledgement: Many thanks to Pål Boug and Erling Holmøy for useful comments. The numerical results were produced by PcGive 10, Hendry and Doornik (2001) and Doornik and Hendry (2001), and RATS v. 5.00, Doan (2000).

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1 Introduction

The hybrid New Keynesian Phillips Curve, hereafter NPC, is an integral part of the standard model of monetary policy. This position is due to its stringent theoretical derivation, as laid out in Clarida, Galí, and Gertler (1999), but also the successful estimation of NPC models on time series data from different countries. In particular, the studies of Galí and Gertler (1999, henceforth GG), and Galí, Gertler and Lopez-Salido (2001, henceforth GGL) give empirical support for the NPC, in the form of correctly signed coefficients and a reasonable good data fit — using US as well as euro area data. Rudd and Whelan (2005) and Lindè (2005) criticize several aspects of the estimation and inference procedures used by GGL, but this line of critique is rebutted in a recent paper by GGL (2005), who re-assert that the NPC, in particular the dominance of forward-looking behavior, is robust to choice of estimation procedure and specification bias.

However, there are reasons to be sceptical to the NPC's status as a proven model of inflation. First, the discussion between GGL and the mentioned critics are within the realm of "statistical inference" and not of "scientific inference" to quote a distinction drawn by Koopmans and Reiersøl (1950). Statistical inference deals with inference from sample to population, hence the essential concerns are the use of the appropriate distribution theory, the

use of optimal estimation techniques and so forth. Scientific inference deals with the interpretation of the population in terms of subject matter theory, see Aldrich (1995).

Central to scientific inference is a concern for *all* the properties and implications of a chosen or maintained interpretation of the correlations (not just a chosen favourable traits), and also mindfulness of alternative hypotheses and explanations of the estimates obtained. Background knowledge is indispensable for scientific inference. In the case of the NPC an important body of background knowledge exists in the form of previous econometric inflation modelling. GGL pay only summary attention to the information content of existing models, and its potential relevance for the significance of the NPC. Thus, the encompassing principle, as laid out in Hendry (1995, Ch. 14), in particular whether the NPC model can explain the properties of earlier models, is not investigated in the series of papers by GG and GGL. As pointed out by e.g., Hendry (1988) the encompassing principle is particularly useful for testing models with rational expectations against models with subjective or 'backward looking' expectations. In line with this, recent research on euro-area data, as well as on time series from the UK and Norway, show that the hybrid NPC model in fact fails to meet the encompassing principle, see Bårdsen, Jansen, and Nymoen (2004), Bårdsen, Eitrheim, Jansen, and Nymoen (2005, Ch. 7) and Boug, Cappelen, and Swensen (2006).

Second, as pointed out by Fuhrer (2005), there is an issue of a certain internal inconsistency. The typical NPC fails to deliver the expected result that inflation persistence is 'inherited' from the persistence of the forcing variable. Instead, the derived inflation persistence, using estimated NPCs, turns out to be completely dominated by 'intrinsic' persistence (due to the accumulation of disturbances of the NPC equation). Quite contrary to the intended interpretation by GGL, Fuhrer (2005) shows that the NPC fails to explain actual inflation persistence by the persistence that inflation inherits from the forcing variable. Fuhrer summarizes that the lagged inflation rate is not a 'second order add on to the underlying optimizing behavior of price setting firms, it *is* the model'.

Third, Bårdsen, Jansen, and Nymoen (2004) show that the euro area NPC estimated by GGL is not robust to quite detailed changes in the GMM estimation, i.e., changes that should have negligible impact under the null that the NPC is a reasonable representation of the inflation process. Moreover, the euro-area NPC is shown to be fundamentally conditioned by certain exclusion restrictions which are invalid when tested.[1] Following Mavroeidis (2005), these results can be understood in the light of the generically weak

[1] The non-robustness due to details in the GMM estimation relates to the significance of the real marginal cost term, see also Bårdsen, Eitrheim, Jansen, and Nymoen (2005, Ch. 7). These critical results are not discussed by Galí, Gertler, and Lòpez-Salido (2005); neither is the paper by Fuhrer (2005).

identification of the NPC model of GGL.

In this paper, we assess the hybrid NPC on a panel data set from OECD countries. Our first finding reproduces the typical NPC equation, in particular regarding the dominance of forward-looking behavior. However, when the scope of the evaluation is widened to address scientific inference and to encompassing, i.e. when the properties of existing models are taken into account, the evidence in favour of the NPC model dissolves. For example, the coefficient of the forward rate is not only statistically insignificant, but is estimated to be zero. Moreover, such a result is predicted by existing dynamic econometric incomplete competition models of inflation, henceforth ICM, meaning that members of this model class encompass the NPC model, while the converse does not apply.

ICMs incorporate the theoretical ideas of monopolistic competition within the equilibrium-correction inflation model of Sargan (1980), Nymoen (1991) and Bårdsen, Eitrheim, Jansen, and Nymoen (2005, Ch. 6). Basically, the ICM framework predicts that the significant relationship between the inflation rate and the inflation rate one period ahead may be a result of incorrect variables omission. In the simplest case, the omitted variable is a linear combination of unit labour costs and the real exchange rate. Hence, the ICM's encompassing implications parallels Yule's analysis of spurious correlations in economics, the correlation between two variables (here: current and future

inflation) being related to some third variable (here: a well specified equilibrium correction term), see Aldrich (1995). In this paper, we show, more generally, that the missing variable suggested by the ICM may be included in an open economy version of the NPC model with testable restrictions on the NPC model's main parameters of interest. As we will show below, these defining restrictions are clearly rejected by our OECD panel data set.

The paper is organized as follows: In section 2, we give, as a background, GGL's view about the 'state of the NPC' as a theoretically derived model of inflation with desirable empirical properties. We also explain our own stance, namely that the lack of encompassing of existing studies is a signal that maybe the NPC is out of its depth. In section 3, we explain the framework for our encompassing oriented assessment of the NPC on OECD panel data, and section 4 presents the data set and discusses some pertinent econometric issues. The results of the econometric tests are given in section 5. Section 6 concludes.

2 The state of the NPC

The hybrid NPC is given as

$$\Delta p_t = \underset{\geq 0}{a^f} \Delta p^e_{t+1} + \underset{\geq 0}{a^b} \Delta p_{t-1} + \underset{\geq 0}{b} ws_t, \tag{1}$$

where Δp^e_{t+1} is expected inflation one period ahead, in our application the period is annual, conditional on period $t-1$ information.[2] Lower case letters indicate that the variable is measured in logs. The 'pure' NPC is specified without the lagged inflation term ($a^b = 0$). In the case of the pure NPC, Roberts (1995) has shown that several New Keynesian models with rational expectations have (1) as a common representation—including the models of staggered contracts developed by Taylor (1979, 1980)[3] and Calvo (1983), and the quadratic price adjustment cost model of Rotemberg (1982). The rationale for allowing $a^b > 0$ is that the theory applies to a (significant) portion of price adjustments in period t, but not to all. Hence, in each period, a share of the overall rate of inflation is determined by last period's rate of inflation, for example because of backward-looking expectations. The third variable in (1) is the logarithm of the wage-share, ws, which is the

[2] To be precice, $\Delta p^e_t = \mathsf{E}(\Delta p_{t+1} \mid \mathcal{I}_{t-j})$ where $\mathsf{E}(\Delta p_{t+1} \mid \mathcal{I}_{t-j})$ denotes the mathematical expectation given information available in time period $t-j$. It has become custom to assume that $j = 0$.

[3] The overlapping wage contract model of sticky prices is also attributed to Phelps (1978).

preferred operational definition of firms' marginal costs of production.[4]

The main references supporting the NPC are the articles by GG and GGL mentioned in the introduction who find that the typical NPC estimation gives the following results:

1. The two null hypotheses of $a^f = 0$ and $a^b = 0$ are firmly rejected both individually and jointly.

2. The hypothesis of $a^f + a^b = 1$ is typically not rejected at conventional levels of significance, although the estimated sum is usually a little less than one.

3. The estimated value of a^f is larger than a^b, hence forward looking behavior is dominant. a^b is usually estimated in the range of 0.2 to 0.6.

4. When real marginal costs are proxied by the wage share, the coefficient b is positive and significantly different from zero.

Critics of the NPC have challenged the robustness of all four typical traits, but with different emphasis and from different perspectives. The inference procedures and estimation techniques used by GG and GGL have been criticized by Rudd and Whelan (2005) and others but GGL (2005) show that

[4]Other close-at-hand measures are the output-gap or the rate of unemployment. However it is the wage-share which most often yields the expected sign on the estimated coefficient of marginal costs, see Galí, Gertler, and Lòpez-Salido (2005). However, also for the wage-share definition, the results are non-robust to minor changes in estimation methodology, see Bårdsen, Jansen, and Nymoen (2004).

their initial results remain robust. However, the empirical validity of the NPC remains damaged in the light of a vector autoregressive regression model on euro area data, see Fanelli (2006).

Bårdsen, Jansen, and Nymoen (2004) and Bårdsen, Eitrheim, Jansen, and Nymoen (2005, Ch 7) have assessed the NPC from another perspective, namely that of encompassing. For several countries, models already exists which (claim to) explain inflation, and it is generally advisable to test a new model, the NPC in this case, against such models. Bårdsen, Jansen, and Nymoen (2004) concentrate on the dynamic incomplete competition model (ICM) of wage and price setting mentioned in the introduction, and find that the NPC model fails to account for the properties of these existing models. Conversely, the dynamic ICM models seems to be able to account for many NPC properties.[5]

For example, based on the ICMs for UK and Norway presented in Bårdsen, Fisher, and Nymoen (1998), it can be hypothesized that the wage-share variable in GGL's euro area NPC is a mis-representation of the true underlying equilibrium correction variable, and therefore that the estimation results for b is probably not as robust as GGL will have us to believe. Using GGLs

[5] Our focus is the encompassing capability of the NPC vis-a-vis, the European tradition of equilibirum correction based inflation modelling. Equally interesting is the testing of the NPC against the North American Phillips-curves, see Gordon (1997) which pre-dates the US data NPC of Galí and Gertler (1999) by several decades, yet GGL omit that information from the assessment of their new model.

data set Bårdsen, Jansen, and Nymoen (2004) show that the significance of the wage share is fragile and depends on the exact implementation of the estimation method used, thus refuting that 4. above is robust on euro-area data.

Bårdsen, Jansen, and Nymoen (2004) also show that the NPC model, and the ICM, can be written as a price adjustment model in equilibrium correction form, see Sargan (1980) and Nymoen (1991). However, compared to the dynamic ICM, the NPC is a highly restrictive equilibrium correction model. On the one hand this means that the NPC can potentially parsimoniously encompass the ICM, but on the other hand it is also possible that the ICM class of models can successfully explain the seemingly robust features of the NPC. The test results, on euro data, UK data and Norwegian data, show that features 1-3 can be explained in the light of the ICM. The crux of the argument is the mis-representation of the equilibrium correction part of the model. When that part of the model is re-specified, with equilibrium correction terms consistent with the wage curve and the long-run price setting equation which are typical of the ICM framework, the hypothesis $a^f = 0$ can no longer be rejected, and $a^f + a^b$ is estimated to be less than one. Both findings are best understood on the premise that, with the (tentatively) correct equilibrium correction terms in place, the model is no longer the differenced data (random walk) model of prices which the NPC model effectively is, see

Fuhrer (2005). Finally, since the significance of a^f is non-robust, it cannot be taken for granted that property 3. holds. On the contrary, a^b seems to be larger than a^f for the investigated data sets. In the case of Norway this is confirmed by the results in Boug, Cappelen, and Swensen (2006).

3 The framework

In this paper, we make use of data from 20 OECD countries, so the closed economy NPC in (1) is a limitation. Recently, Batini, Jackson, and Nickell (2005) have derived an open economy NPC from theoretical principles, showing that the main theoretical content of the NPC generalizes, but that consistent estimation of the parameters a^f, a^b and b requires that the model is augmented by variables which explain inflation in the open economy case. Hence, the open economy NPC (OE-NPC) is

$$\Delta p_t = \underset{\geq 0}{a^f} \Delta p^e_{t+1} + \underset{\geq 0}{a^b} \Delta p_{t-1} + \underset{\geq 0}{b} ws_t + c\, x_t, \tag{2}$$

where x_t, in most cases a vector, contains the open-economy variables, and c denotes the corresponding coefficient vector. The change in the real import price, $\Delta(pi_t - p_t)$ in our notation, is the single most important open economy

augmentation of the NPC. The results in Batini, Jackson, and Nickell (2005) are, broadly speaking, in line with GG and GGL properties 1-4 above, but as noted above, those properties are not robust when tested against the existing UK model in Bårdsen, Fisher, and Nymoen (1998).

To derive testable implications of the NPC on our country data set we make use of the identity

$$ws_t = ulc_t - pd_t, \tag{3}$$

where ulc denotes unit labour costs (in logs) and pd is the log of the price level on domestic goods and services. Let $(1-\gamma)$ be the share of imports, then the aggregate price level is defined as

$$p_t = \gamma\, pd_t + (1-\gamma)\, pi_t. \tag{4}$$

If we solve this for pd, insert in (3) and re-write, we get the following equation for the wage-share:

$$ws_t = -\frac{1}{\gamma}\left[p_{t-1} - \gamma\, ulc_{t-1} - (1-\gamma)\, pi_{t-1}\right] + \Delta ulc_t - \frac{1}{\gamma}\Delta p_t + \frac{1-\gamma}{\gamma}\Delta pi_t. \tag{5}$$

We can then re-write the open economy NPC as

$$\begin{aligned}\Delta p_t &= \frac{a^f}{\left(1+\frac{b}{\gamma}\right)}\Delta p^e_{t+1} + \frac{a^b}{\left(1+\frac{b}{\gamma}\right)}\Delta p_{t-1} - \frac{b}{(\gamma+b)}\left[p_{t-1} - \gamma\ ulc_{t-1} - (1-\gamma)\, pi_{t-1}\right] \\ &\quad + \frac{\gamma\ b}{(\gamma+b)}\Delta ulc_t + \frac{b\,(1-\gamma)}{(\gamma+b)}\Delta pi_t + \frac{\gamma\ c}{(\gamma+b)}\, x_t,\end{aligned}$$

or

$$\begin{aligned}\Delta p_t &= \alpha^f \Delta p^e_{t+1} + \alpha^b \Delta p_{t-1} + \beta(ulc_{t-1} - p_{t-1}) - \beta\,(1-\gamma)\,(ulc_{t-1} - pi_{t-1}) \\ &\quad + \beta\ \gamma\ \Delta ulc_t + \beta\,(1-\gamma)\,\Delta pi_t + \psi x_t,\end{aligned} \tag{6}$$

where we have defined α^f, α^b, β and ψ as new coefficients for simplification. This equation brings out that the NPC has an interpretation as an equilibrium correction model (ECM), of the price level, see Sargan (1980) and Nymoen (1991), but with two important remarks. First, the usual ECM for inflation is extended by the inclusion of the forward-looking term Δp^e_{t+1}. Second, the econometric ECM is restricted since the coefficients of Δulc_t, Δpi_t and the ECM terms, $(ulc_{t-1} - p_{t-1})$ and $(ulc_{t-1} - pi_{t-1})$, are restricted to be functions of b and γ.

As mentioned above, an alternative model for price formation is the incomplete competition model, ICM, where prices are set as a mark-up over

unit labour cost and where the mark up depends on relative prices:

$$pd = m_0 - m_1 (pd - pi) + ulc \tag{7}$$

where $0 \leq m_1 \leq 1$. By using (4) we get

$$p = \mu_0 + \mu_1 ulc + (1 - \mu_1) pi, \tag{8}$$

where $\mu_1 = \frac{\gamma}{1+m_1}$ and $\mu_0 = m_0 \ \mu_1$. Due to for example incomplete information or adjustment costs, prices are rarely – if ever – at this optimal level. Therefore it has become popular to present the ICM in equilibrium correction form, where (8) is the long run part and where variables that is believed to be important in the shorter run make up the short run part. For comparison let us say that the dynamic part of the NPC is the true one, and therefore include the same variables also in the ICM. Then the ICM would look like this:

$$\begin{aligned} (9)\ \Delta p_t &= \alpha^f \Delta p^e_{t+1} + \alpha^b \Delta p_{t-1} + \beta_1 (ulc_{t-1} - p_{t-1}) + \beta_2 (ulc_{t-1} - pi_{t-1}) \\ &\quad + \beta_3 \Delta ulc_t + \beta_4 \Delta pi_t + \psi x_t. \end{aligned}$$

Hence, a comparison of the the two rivaling models, the OE-NPC in (6) and

the ICM in (9), reveals that the only difference between the two is that while the OE-NPC implies restrictions on the coefficients, the ICM is much less restrictive, i.e. under the given dynamic specification only that $\beta_1 > 0$ and $0 > \beta_2 > -\beta_1$. Empirical tests of the coefficient-restrictions implied by the OE-NPC may settle the issue. Consider the following two hypothesis: H_0^a: $\beta_3 = \beta_1 + \beta_2$ and H_0^b: $\beta_4 = -\beta_2$. The rejection of H_0^a and/or H_0^b would therefore appear to be telling evidence against the OE-NPC.

As noted above, OE-NPC models are usually specified with the rate of change in the real import price as one of the elements in x_t. Equation (9) is consistent with that interpretation, the only caveat applies to β_4 and H_0^b, since $\beta_4 = -\beta_2$ no longer follows logically from the NPC. This is because β_4 is a composite parameter also when the NPC is the valid model.

There are additional properties of the open economy NPC that can be tested on our panel data set of OECD inflation data. For example, we can test the significance of the forward and lagged inflation terms, by testing the null-hypothesis of H_0^c: $\alpha^f = 0$ and H_0^d: $\alpha^b = 0$. This is basically the panel data version of the usual econometric assessment of the NPC on country (or area) data referred to above, GG and GGL in particular. The two former hypotheses H_0^a and H_0^b, which capture the implied NPC restriction of the leads and lags of ulc, have so far not been considered systematically.

Though equation (6) is seen to encompass two different strands of the

literature, the NPC and the ECM approach to inflation modelling, it remains very restrictive since it assumes perfect competition. The alternative econometric model, the incomplete competition model, ICM, since it assumes incomplete competition in both price and wage setting, only requires that $\beta_1 > 0$ and $0 > \beta_2 > -\beta_1$ for being logically consistent with the idea that in a stable long-run situation, the price level is a mark-up on unit labour costs, and that the mark-up depends on competitiveness, see Nymoen (1991) and Bårdsen, Eitrheim, Jansen, and Nymoen (2005, Ch. 6). However, notice that the ICM does not imply H^c: $\alpha^f = 0$. Hence a structural ICM for inflation with elements of forward-looking behavior is a constructive alternative to both the NPC and the ICM with (only) backward-looking expectations.

4 Data and econometric issues

We use a data set for annual wages and prices for 20 OECD countries, for the time period 1960-2004. For some of the countries the time period is shorter, so the panel is unbalanced. Because of leads and lags we loose the observations from 1960 and 2004.

The main data in the analysis is retrieved from the MEI OECD database. The definitions and data sources are given in appendix A, but we note that while almost all previous papers use data for the manufacturing sector we

use the OECD unit labour cost index that covers the whole economy. The import prices are constructed by taking the ratio of the value and the volume of imported goods and services. Furthermore, we use the consumer price index as a measure for the endogenous variable.

There are seperate open economy price adjustment equation for each country in the panel. As a benchmark model we first estimate the NPC model (2) with the following variables in the x vector: the rate of change in the oil price (Δpo_t) and the change in the indirect tax rate (ΔVAT_t) as well as the change in the real import price $\Delta(pi_t - p_t)$. The resulting equation is denoted M1 in the next section.[6] The oil price is denominated in US dollars and Δpo_t therefore captures cost shocks that are common to the countries in the panel.

However, as we have seen above, the relationship between the NPC and the dynamic ICM model is brought out by the open economy inflation equation (9), which we repeat here as

$$
\begin{aligned}
(10) \qquad \Delta p_{i,t} &= \alpha^f \Delta p^e_{i,t+1} + \alpha^b \Delta p_{i,t-1} + \beta_1 (ulc_{i,t-1} - p_{i,t-1}) \\
&\quad + \beta_2 (ulc_{i,t-1} - pi_{i,t-1}) + \beta_3 \Delta ulc_{i,t} + \beta_4 \Delta pi_{i,t} \\
&\quad + \psi_1 \Delta po_{i,t} + \psi_2 \Delta VAT_{i,t} + \varepsilon_{i,t}.
\end{aligned}
$$

[6] Of course, since we normalize on Δp_t, it is nominal import price growth that appears on the right-hand-side of the estimated equation.

The variables are the same as in the previous sections, but we have added an extra subscript i for each country and a stochastic error term $\varepsilon_{i,t}$. This model is denoted M2 in the next section. As we have seen above, the validity of the NPC hinges not only on the significance of the forward term (rejection of H_0^c: $\alpha^f = 0$), but also on H_0^a: $\beta_3 = \beta_1 + \beta_2$ not being rejected.

The presence of the Δp_{t+1}^e in the model causes two econometric problems. The first is a relatively minor one, and arises because estimation proceeds by substitution of Δp_{t+1}^e by the observable Δp_{t+1}, which induces a moving average disturbance term in the estimated model, even if the original equation has white noise errors, see Blake (1991). Usually this problem is tackled by the use of GMM estimation, and we can do the same on our panel data set. Second, and more fundamentally, models with forward-looking rational expectations term are not easily identified, see Pesaran (1987) and Mavroeidis (2004). In brief, rational expectations forces a situation where valid instruments may also be weak instruments. As a practical solution, we include the 2. order lags of variables like inflation in the instrument list, which helps identification if the marginal model of e.g., ulc_t does not depend on Δp_{t-1}. Other available variables may also be used as instruments. For example, since Δulc_t is on the right hand side, we can use lags of rates of unemployment as instruments since we do not expect the rate of unemployment to affect inflation through other channels than unit labour costs. The same line

of reasoning motivates that variables measuring employment protection and the unemployment benefit replacement ratio can be used as instruments. The full set of instruments is given in connection with the results section below.

Nickell (1981) shows that OLS estimation may be inconsistent when applied to models that include fixed effects and a lagged dependent variable. The bias is of the order $1/T$, where T is the time dimension of the panel. In our case the time dimension varies from 21 to 37, therefore it is likely that the 'Nickell bias' will be very small. Moreover, this is largely confirmed by Judsen and Owen (1999) who show that OLS estimation of dynamic fixed effects models perform well for $T = 30$, i.e. with a T dimension similar to ours. Even when $T = 20$, the fixed effects estimator were almost as good as the alternatives (GMM and Anderson-Hsiao).

The pooled panel data regression is valid only under the assumption that the slope coefficients are homogeneous across countries. As shown by Pesaran and Smith (1995), if homogeneous coefficients are falsely imposed, the pooled estimator is inconsistent even if T approaches infinity. However, as pointed out by Baltagi (1995) the pooled model can yield more efficient estimates at the expense of bias, and one must therefore balance the two concerns. We have nevertheless assumed homogeneous coefficients, and since the estimated coefficients are in the same magnitude as in other studies, the bias is believed to be small.

Table 1: Panel unit root tests, 1960–2004. P-values in parenthesis.

Null: Unit root, levels Individual effects and linear trends	p	ulc	pi
Levin-Lin-Chu, t-stat	1.75 (0.96)	1.99 (0.98)	3.86 (1.00)
Im-Pesaran-Shin, W-stat.	4.22 (1.00)	6.06 (1.00)	6.94 (1.00)
ADF – Fisher, χ^2- stat.	15.1 (1.00)	13.0 (1.00)	8.84 (1.00)
PP – Fisher, χ^2- stat.	1.07 (1.00)	17.9 (1.00)	4.23 (1.00)
Null: Unit root, differences Individual effects and linear trends	Δp	Δulc	Δpi
Levin-Lin-Chu, t-stat	-3.49 (0.00)	-7.09 (0.00)	-14.1 (0.00)
Im-Pesaran-Shin, W-stat.	-2.82 (0.00)	-5.64 (0.00)	-10.6 (0.00)
ADF – Fisher, χ^2- stat.	63.1 (0.01)	96.4 (0.00)	182.0 (0.00)
PP – Fisher, χ^2- stat.	41.3 (0.41)	89.6 (0.00)	308.7 (0.00)

The principle of balanced equations requires that the variables are either stationary or cointegrated. However, macroeconomic time series are typically non-stationary, and we therefore have to investigate the order of integration of the main variables in our study. Unit- root tests have in general low power, and in order to improve power we have performed four different panel unit root tests; The Levin-Lin-Chu test (Levin et al., 2002), the Im-Pesaran-Shin test (Im et al., 2003), the Fisher-ADF test and the Fisher-PP test (Maddala and Wu, 1999, and Choi, 2001). The results are reported in Table 1. The null hypothesis of a unit-root is not rejected for any of the variables. However, the null of I(2) is clearly rejected, except in the PP-test for Δp. Hence, the unit root analysis indicate that the growth rates included in the dynamic

Table 2: Pedroni (1999) panel cointegration tests. Heterogenous intercepts included. P-values in parenthesis

Null of no cointegration							
Test number	1	2	3	4	5	6	7
No time dummies, no trend							
Test statistics	−1.0	2.0	1.7	1.7	2.9	2.1	1.4
	(0.32)	(0.05)	(0.09)	(0.09)	(0.00)	(0.04)	(0.16)
With time dummies, no trend							
Test statistics	1.7	−0.1	−0.3	−0.8	1.4	0.2	−0.9
	(0.09)	(0.92)	(0.76)	(0.42)	(0.16)	(0.84)	(0.37)
With time dummies and heterogenous deterministic trends							
Test statistics	1.3	0.4	−0.5	−2.1	1.8	−0.6	−3.0
	(0.19)	(0.69)	(0.62)	(0.04)	(0.07)	(0.55)	(0.76)

part of model (10) seem to be stationary.

We also test for cointegration between the variables that make up the equilibrium part of the ICM inflation equation. Pedroni (1999) suggests a suite of 7 tests designed to test the null hypothesis of no cointegration in dynamic panels with multiple regressors. The first four tests are based on the within panel estimator (see Hsiao, 1986), and are listed as tests 1–4 in Table 2. The last three tests are labelled Group Mean Panel Tests by Pedroni, and are calculated by pooling along the between dimension. The test statistics are calculated using RATS[7].

While macro panels typically exhibit cross-sectional dependence, the panel unit root tests and the Pedroni panel data cointegration test all assume cross-

[7]RATS v. 5.00, Doan (2000). Many thanks to professor Peter Pedroni for providing us with the RATS codes used to calculate the relevant test statistics.

country independence. As shown by Banerjee et al. (2004, 2005) using Monte Carlo simulations, falsely assuming cross-sectional independence causes severe size distortions. The inclusion of common time dummies could capture some of the common shocks and as thus correct for this form of cross-sectional dependence in the panel. Therefore we considered three cases regarding the cointegrating space; one without time dummies and deterministic trends, one where time dummies were included, but not deterministic trends, and one where heterogeneous deterministic trends and time dummies were included.

The Pedroni-tests in Table 2 show that the null of no cointegration is only rejected in some of the tests, hence the formal evidence in favor of cointegration is weak. However, since the estimated coefficients in our models – both in the OE-NPC and the ICM – resembles quite well the findings in single-country analysis and the cointegration tests have low power, we continue our modelling strategy assuming that the long run variables are in fact cointegrated. After all, our most important benchmark is the existing literature cited previously.

5 Econometric results

Table 3 reports the estimation results for the econometric OECD inflation models. As explained above, M1 represents the model that has been estimated on several data sets with results that are summarized in section 2. In M1, real marginal costs are measured in accordance with equation (3) above, i.e., by the wage share of gross value added. M1' instead uses unit labour costs deflated by the consumer price index, which may be a better measure than $ws_{i,t}$, since the change in the consumer price index is the left hand side variable. M2 is the estimated equilibrium correction model (10), which encompasses both the NPC and the ICM interpretation.

The models are estimated using GMM, where $\Delta p_{i,t+1}$, $\Delta ulc_{i,t}$ and $\Delta(pi_{i,t} - p_{i,t})$ are treated as endogenous explanatory variables. The following variables are used as instruments in all models: $\Delta p_{i,t-2}$, $\Delta pi_{i,t-1}$, $\Delta po_{i,t-1}$, $\Delta ulc_{i,t-1}$ and $ws_{i,t-1}$, the gross replacement rate and its lags, and an index of employment protection and its lags. $(ulc_{i,t-1} - pi_{i,t-1})$ and $(ulc_{i,t-1} - p_{i,t-1})$ are additional instruments in the two M1 models.

As can be seen, the results for M1 and M1' are well aligned with GGL's typical hybrid NPC model. In fact, the first three typical features listed in section 2 are clearly recognizable in the column with results for M1. Both the lagged and leading inflation terms have significant coefficients; the sum of

Table 3: GMM estimation results for an OECD panel data set. Heteroscedasticity consistent standard errors in parenthesis.

	M1	M1'	M2
$\Delta p_{i,t+1}$	0.56 (0.03)	0.57 (0.03)	−0.01 (0.12)
$\Delta p_{i,t-1}$	0.47 (0.03)	0.46 (0.02)	0.38 (0.03)
$ws_{i,t}$	−0.010 (0.01)		
$(ulc_{i,t} - p_{i,t})$		−0.005 (0.008)	
$(ulc_{i,t-1} - p_{i,t-1})$			0.053 (0.0014)
$(ulc_{i,t-1} - pi_{i,t-1})$			−0.020 (0.006)
$\Delta ulc_{i,t}$			0.32 (0.06)
$\Delta pi_{i,t}$			0.11 (0.014)
$\Delta(pi_{i,t} - p_{i,t})$	0.05 (0.01)	0.05 (0.01)	
$\Delta po_{i,t}$	0.005 (0.002)	0.005 (0.02)	0.005 (0.02)
$\Delta VAT_{i,t}$	0.003 (0.0005)	0.003 (0.0004)	0.003 (0.0004)
# observ	567	567	567
$\hat{\sigma} \cdot 100$	1.29	1.29	1.0
χ^2_{ival}	41.49[0.000]	41.96[0.000]	10.96[0.204]
$N_{\mathsf{AR\text{-}1}}$	−3.07[0.002]	−3.02[0.002]	−0.26[0.81]
$N_{\mathsf{AR\text{-}2}}$	−2.34[0.019]	−2.35[0.019]	−0.30[0.76]

Notes: Square brackets, [..], contain p-values, standard errors are in parenthesis, (..). $\hat{\sigma}$ denotes the estimated residual standard deviation. χ^2_{ival} denotes Sargan's (Sargan, 1964) specification test which is χ^2 distributed under the null of valid instruments. $N_{\mathsf{AR\text{-}1}}$ and $N_{\mathsf{AR\text{-}2}}$ have a standard normal distribution under the null of no 1. and 2. order autoregressive residuals.

the coefficients cannot be statistically distinguished from unity, and forward-looking behavior dominates. The only anomaly is the insignificance of the wage-share coefficients, which contradicts the typical NPC feature 4. However, as mentioned above, Bårdsen, Jansen, and Nymoen (2004) have documented that the wage-share coefficient is non-robust, even on the euro-area data used by GGL. That the M1 results are corroborating the typical finding on US and euro-area data, as well as on other country data sets may be taken as an indication that the problem with between country correlation is not too large. Usually, time dummies are included to correct for one type of cross sectional dependence. However, handling this potential problem by means of time dummies is unsatisfactory in this model since the model includes a lead as well as a lag of the left-hand side variable, with over-fitting as a result.

As shown in the previous sections, significance of the forward-term in M1 should carry over to M2 if the NPC is the right theoretical framework. However, we observe the opposite, namely that the hypothesis H_0^c: $\alpha^f = 0$ is not rejected in M2. The coefficient is in fact estimated to zero. The dominance of the forward term in M1 is thus due to $\Delta p_{i,t+1}$ being correlated with $(ulc_{i,t-1} - p_{i,t-1})$ and $(ulc_{i,t-1} - pi_{i,t-1})$; there is no genuine correlation between the predictable part of $\Delta p_{i,t+1}$ and $\Delta p_{i,t}$. By considering the coefficients (and standard errors) of $(ulc_{i,t-1} - p_{i,t-1})$, $(ulc_{i,t-1} - pi_{i,t-1})$ and $\Delta ulc_{i,t}$ it is also evident that H_0^a: $\beta_3 = \beta_1 + \beta_2$ will be rejected at any level

of significance: the estimated coefficient is 0.32, which is 10 times the size predicted by the NPC.[8]

The diagnostic tests at the bottom of the table also convey bad news for the NPC: In M1, the Sargan test χ^2_{ival} is significant, and there is indication of quite significant residual autocorrelation (also of 2. order). For M2 there are no signs of mis-specification. Moreover, M2 is easy to interpret as a simple price equation consistent with a different supply shocks (demand shocks might be said to be under-represented in this model), but also to last periods deviation between the price level and a hypothetical long-run price equation which functions as an attractor. The t-statistic of the $(ulc_{i,t-1} - p_{i,t-1})$ terms indicate significance, and the implied estimate for the weight on unit labour cost in the long-run price equation is 0.64 which is of reasonable magnitude, although one would of course expect that a better estimate would allow for heterogeneity between countries. Thus, the results for M2 indicate that the variables that enter the long run part of the model are cointegrated even though the formal panel cointegration tests in Section 4 were inconclusive on this point.

[8] The 't-statistic' is 46.8.

6 Conclusions

GGL claim that the NPC represents a significant advance in inflation modelling which finally substantiates the dominance of forward-looking behavior in price adjustment. We have argued that the scientific inference method used by GGL and others is doubtful since it leaves out any systematic assessment of their findings in the light of existing models and of alternative hypotheses. In line with Bårdsen, Jansen, and Nymoen (2004) we show that the model class made up of dynamic incomplete competition models (ICMs) can explain both why the forward-term dominates in GGLs findings, but also why that dominance may be more apparent than a genuine feature of price dynamics.

The estimation results in this paper give little support to the main theoretical ideas of the NPC, namely the hypothesized significant roles of the forward looking term and the wage share as proxy of marginal costs. Our analysis suggests that the expected inflation rate and the wage share may be acting as replacements for equilibrium correction terms that are better approximations of actual price setting behavior, consistent with the ICM. Furthermore, we show that the econometric model of inflation would improve markedly by adding the lagged real unit labour costs and the ratio between unit labour costs and import prices as separate explanatory vari-

ables. These improvements critically affect the estimated coefficient of the forward term, not only is the coefficient insignificant, it is also estimated to zero.

A Data definitions and sources

The data consists of annual time series from as early as 1960 for some countries and up to 2004 for all the 20 OECD countries given in the table below. Some of the variables do not exist for the whole period, and similarly some countries' variables are not available. Consequently, we use an unbalanced panel data set.

Most of the data used in this paper is retrieved from or constructed by using the Organisation for Economic Co-operation and Development (OECD) Economic Outlook and Main Economic Indicators (MEI) Databases.[9] This should help ensuring consistency in the dataset.

Description of the variables

P : *Consumer price index.* The P variable is constructed by using a Purchasing Power Parity index (PPP) and multiplying it with the consumer price index for USA in order to get comparable consumer prices between the OECD countries in the sample. The PPP variable is in its simplest form,

[9] By using Xvision Fame 8.0.2, a programme licensed by SunGard Data Management Solutions.

Table 4: Listing of countries in the data set.

Name of country	**Number in database**
Australia	1
Austria	2
Belgium	3
Canada	4
Denmark	5
Finland	6
France	7
Germany	8
Ireland	9
Italy	10
Japan	11
Netherlands	12
New Zealand	13
Norway	14
Portugal	15
Spain	16
Sweden	17
Switzerland	18
UK	19
USA	20

id est consumer price index in local currency divided by consumer price in USD. The calculation gives us:

$$P_i = PPP_i \cdot CPI_{US_INDEX} = \frac{CPI_i}{CPI_{US}} \frac{CPI_{US}}{CPI_{US_2000}} = \frac{CPI_i}{CPI_{US_2000}}$$

The denominator (CPI in US for year 2000) is simply a constant and just adds to the constant in the regression.

PI : Price of imports. The ratio of import value and import volume is

used as a proxy for the price of imports.

PO : *Price of oil.* The world dated price of Brent crude oil measured in USD.

UR : *Rate of unemployment.* The OECD standardised unemployment rates give the number of unemployed persons as a percentage of the civilian labour force.

ULC : *Unit Labour Costs.* ULC is an index of unit labour costs (2000=100) provided by the OECD.

VAT : *Indirect tax rate.* This is standard VAT rates in percent for the different OECD countries. VAT rates for the EU is retrieved from DOC/1635/2005 - EN. VAT rates for Japan, New Zealand, Norway, Canada and Australia is obtained from the countries' repective national beureaus of statistics.

EP: Employment protection. The data comprises an index of the degree of employment protection provided by Dr. Luca Nunziata, Nuffield College, University of Oxford, UK. See Nunziata (2005).

BBR: Benefit Replacement Ratio. The data comprises an index of unemployment benefits in percent of the average wage level. Provided by Dr. Luca Nunziata, Nuffield College, University of Oxford, UK. See Nunziata (2005).

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