# The New Keynesian Phillips curve tested on OECD panel data*

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**Abstract**

In their work, Galí, Gertler and Lopez-Salido, GGL, assert that the hybrid New Keynesian Phillips curve (NPC) with dominance of forward-looking behavior and real marginal costs is robust to choices of estimation procedure, details of variables definitions and choice of data samples. In an estimation on panel data from OECD countries we replicate the typical empirical NPC from country studies. However, we also test an alternative economic interpretation of the empirical correlations. Specifically we find that the expected rate of future inflation and real marginal costs serve as replacements for equilibrium correction terms that are implied by the general imperfect competition model of wage and price setting. Hence, as an explanatory model of OECD inflation, the NPC is encompassed by an existing theory.

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

**JEL classification:** *C23, C52, E12, E31*

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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 microeconomic foundation, as laid out in Clarida et al. (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 López-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 Linde (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, model evaluation entails consideration of *all* the properties and implications of a chosen or maintained interpretation of the correlations (not just chosen favorable 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, shows that the hybrid NPC model in fact fails to meet the encompassing principle, see Bårdsen et al. (2004), Bårdsen et al. (2005, Ch. 7) and Boug et al. (2006).

Second, as pointed out by Fuhrer (2006), 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 (2006) 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 et al. (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 identification of the NPC model of GGL.

In this paper, we assess the hybrid NPC on a panel data set from OECD countries assuming. The pooled estimator is biased if homogeneity is falsely imposed, but it is more efficient and it has no small sample bias as would be the case if the model was estimated for each country separately. However, the main motivation for considering a pooled estimator stems from the observation that the microfoundations of the NPC model abstract from the institutional and historical idiosyncrasies of individual countries, and that this may explain why the NPC might be inferior to models that are specified to explain exactly those features. The NPC should be expected to perform better as a model of "representative" price dynamics for a panel of countries. As a confirmation of this hypothesis, our first finding in this paper show that the typical NPC equation stands its ground very well on the OECD data set, in particular the dominance of forward-looking behavior in price setting. This result indicates that if the pooled estimator is biased, the bias is small.

Our second main finding is that when the evidence represented by existing models are taken into account, the evidence in favour of the NPC model dissolves, i.e., the NPC fails on the encompassing test. For example, the coefficient of the forward rate is not only statistically insignificant, but is estimated to be very close to zero. Moreover, such a result is predicted by existing dynamic econometric imperfect 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 et al. (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 omission of variables. 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).[2] Conversely, we show below that the equilibrium correction variables suggested by the ICM can be rationalized under the hypothesis that the NPC holds. It is then straightforward to test the null hypothesis that the NPC restrictions hold using likelihood ratio tests.

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 the NPC, although brilliantly derived as a theoretical model, is found lacking in its econometric properties. 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

[1] The non-robustness due to details in the GMM estimation relates to the significance of the real marginal cost term, see also Bårdsen et al. (2005, Ch. 7).

[2] See Aldrich (1995) for an overwiev of Yule's work on spurious correlations.

section 5. Section 6 concludes.

## 2 The empirical status 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 $\Delta p^e_{t+1}$ is expected inflation one period ahead, conditional on the information available in period $t-1$.[3] Lower case letters indicate that the variable is measured in logs, and the third variable in (1) is the logarithm of the wage-share, $ws$, which is the preferred operational definition of firms' marginal costs of production.[4].

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)[5] 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 main references supporting the NPC are the articles by GG and GGL mentioned above 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 at conventional levels of significance.

Critics of the NPC have challenged the robustness of all four typical traits, but with different emphasis and from different perspectives. The inference procedures

[3]To be precise, $\Delta p^e_t = \mathsf{E}(\Delta p_{t+1} \mid \mathcal{I}_{t-j})$ 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 $j = 0$.

[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 Gali et al. (2005). However, also for the wage-share definition, the results are non-robust to minor changes in estimation methodology, see Bårdsen et al. (2004).

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

and estimation techniques used by GG and GGL have been criticized by Rudd and Whelan (2005) and others, but GGL (2005) show that their results remain robust. However, the statistical adequacy of the NPC for US and euro area data is also brought into doubt by the results in Fanelli (2008) and Juselius (2007, Ch III) based on the vector autoregressive regression model.

Bårdsen et al. (2004) and Bårdsen et al. (2005, Ch. 7) have assessed the NPC from another perspective, namely that of encompassing. For several countries, inflation models existed already, and it is generally advisable to test a new model, the NPC in this case, against such models. Bårdsen et al. (2004) concentrate on the dynamic imperfect 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 this existing model. Conversely, the dynamic ICM model seems to be able to account for many NPC properties.[6]

For example, based on the ICMs for UK and Norway presented in Bårdsen et al. (1998), it can be hypothesized that the wage-share variable in GGL's euro-area NPC is a misrepresentation 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 GGL's data set Bårdsen et al. (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 result number 4. above is robust on euro-area data.

Bårdsen et al. (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 misrepresentation 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. In the case of Norway, this is confirmed by the results in Boug et al. (2006).

There is no suggestion in the theory about how we should choose the time period $t$ in equation (1), as month, quarter or year. The applications and tests on country data just cited use quarterly data, whereas with panel data the annual period is the only practical choice. This give rises to the questions of temporal aggregation consequences and of comparison of results based on the two periodicities. However, Galí and Gertler (1999) noted that prior to their work on quarterly data the only successful estimation of NPCs had used annual data, indicating that if anything,

[6]Our focus is the encompassing capability of the NPC vis-a-vis, the European tradition of equilibrium 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 exclude that information from the assessment of their new model.

the annual frequency favours the NPC. Consistent with this view we show below that the typical features 1-4 above are replicated on our annual data set, indicating robustness with respect to temporal aggregation.

## 3 An encompassing perspective

In this paper, we make use of data from 20 OECD countries, so the closed economy NPC in (1) is a limitation. Batini et al. (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 et al. (2005) are, broadly speaking, in line with GG's and GGL's properties 1.-4., but as noted, those properties are not robust when tested against the existing UK model in Bårdsen et al. (1998).

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

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

where $ulc$ denotes unit labour costs (in logs) and $pd$ is the log of the price level on domestic goods and services. We let $(1-\gamma)$ denote a constant import share, and define the aggregate price by the definition:

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

If we solve for $pd$, insert in (3) and re-write, we obtain 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

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

or

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

where we have conveniently defined $\alpha^f$, $\alpha^b$, $\beta$ and $\psi$ as new coefficients. 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 $\Delta p^e_{t+1}$. Second, the econometric ECM is restricted since the coefficients of $\Delta ulc_t$, $\Delta pi_t$ 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 $\gamma$.

(6) assumes a constant import share which may be too stylized for many purposes. However, how this restriction affects the comparison between rival models of inflation is not clear *a priori*, and in section 5 we take care to show that it does not tilt the evidence against the NPC in the encompassing test.

As mentioned above, an alternative model for price formation is the imperfect competition model, ICM, where prices are set as a mark-up over unit labour cost and where the mark up depends on relative prices:

$$(7)\qquad pd = m_0 - m_1\,(pd - pi) + ulc,$$

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

$$(8)\qquad p = \mu_0 + \mu_1 ulc + (1-\mu_1)\,pi,$$

where $\mu_1 = \frac{\gamma}{1+m_1}$ and $\mu_0 = m_0\ \mu_1$. Due to for example shocks, incomplete information and adjustment costs, prices are rarely at this optimal level. Therefore it has become standard to present the ICM in equilibrium correction form, where (8) is used to define the equilibrium correction variable, and where variables that are likely to be important in the short run are also included in the model. For simplicity, we assume that the dynamic part of the NPC is valid, and therefore include the same variables also in the equilibrium correction model, which then becomes:

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

A comparison of the two models, the OE-NPC in (6) and equation (9), reveals that the OE-NPC is a special case of the (9). The standard ICM, without the lead in inflation, is also a special case of (9). Therefore we will refer to (9) as the minimum nesting model or the encompassing model, cf Hendry (1995, ch 14.6). The OE-NPC in (6) implies the following restrictions on the coefficients in (9): $H^a_0$: $\beta_3 = \beta_1 + \beta_2$ and $H^b_0$: $\beta_4 = -\beta_2$. Hence, the rejection of $H^a_0$ and/or $H^b_0$ are not consistent with OE-NPC parsimoniously encompassing model (9). The same applies if $H^c_0$: $\alpha^f = 0$ cannot be rejected statistically based on estimation of (9): this test-outcome is inconsistent with the main assumption of the NPC, namely that a significant proportion of price setters are forward looking in the rational expectations sense. Finally $H^d_0$: $\alpha^b = 0$ can also be tested using (9).

The tests of the significance of the forward and lagged inflation terms, $H^c_0$: $\alpha^f = 0$ and $H^d_0$: $\alpha^b = 0$, are basically panel data versions 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 restrictions for the leads and lags of $ulc$, have so far not been considered systematically. However, 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 $\beta_4$ and $H_0^b$, since $\beta_4 = -\beta_2$ no longer follows logically from the NPC. This is because $\beta_4$ is a composite parameter also when the NPC is the valid model. For that reason, our focus in section 5 will be on hypotheses $H_0^a$, $H_0^c$, $H_0^d$, and not on $H_0^b$.

The ICM implies fewer testable restrictions on the nesting model in (9). The ICM does however require that $\beta_1 > 0$ and $0 > \beta_2 > -\beta_1$. Subject to this, the standard ICM, without a lead-term parsimoniously encompasses model (9) if $H_0^c$: $\alpha^f = 0$ is not rejected statistically.

Nesting of the two rivaling models in an encompassing model is convenient for formal testing of parsimonious encompassing. More generally, the standard ICM (i.e., without the lead in inflation) has the potential of encompassing the NPC, since the ICM can explain the properties of the NPC when estimated in the conventional form given by equation (1), or (2). This is because the ICM, if correct, implies that for example $(ulc_{i,t-1} - p_{i,t-1})$ is positively correlated with both $\Delta p_t$ and $\Delta p_{t+1}^e$, and so omitting $(ulc_{i,t-1} - p_{i,t-1})$ from the inflation equation, means that the estimated value of $\alpha^f$ will be biased upwards. The same is true if we consider $(ulc_{i,t-1} - p_{i,t-1})$, or a linear combination of the two. Inclusion of $(ulc_{i,t-1} - p_{i,t-1})$ and $(ulc_{i,t-1} - pi_{i,t-1})$ in the list of instruments used when estimating the NPC will not solve this problem with bias, although it will reduce the problem of weak instruments. Statistical significance of Sargan's (1964) test for the validity of overidentifying instruments, dubbed $\chi^2_{\mathsf{ival}}$ in table 3 below, is a sign that bias due to omitted variables may be an issue, since that test is interpretable as test of whether the structural equation (i.e., the NPC) parsimoniously encompass the implied unrestricted reduced form, where all the instruments are included. Hence a logically test outcome, if the ICM model encompasses the NPC, would for example be that estimation of (2) gives the expected signs and magnitudes of $a^f$ and $a^b$ at the same time as $\chi^2_{\mathsf{ival}}$ is significant, and the estimated value of $a^f$ is reduced when $(ulc_{i,t-1} - p_{i,t-1})$ and $(ulc_{i,t-1} - pi_{i,t-1})$ are moved from the list of instruments and instead are included in the inflation equation. Bårdsen et al. (2005, Ch 7) report such a constellation of evidence for three different data sets: the euro area, the UK and Norway. Below we investigate whether a similar constellation can be found for the panel of OECD data.

Finally, the misplacement of explanatory variables in the list of instruments will most likely lead to autocorrelated residuals in the NPC equation. Hence, the ICM if it is a congruent model, predicts the misspecification of the NPC, see Bontemps and Mizon (2001) This aspect of encompassing may be difficult to detect though, since a first order moving average process in the NPC residuals are consistent with that model, due to replacement of the expected lead with the actual lead.

## 4 Data and econometric issues

As already mentioned, 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 one lead and one lag we loose the observations from 1960 and 2004.

The main data in the analysis are retrieved from OECD's Main Economic Indicator (MEI) 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 price index is 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.

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 ($\Delta po_t$) and the change in the indirect tax rate ($\Delta VAT_{i,t}$) as well as the change in the real import price $\Delta(pi_{i,t} - p_{i,t})$. The resulting equation is denoted NPC1 in the next section.[7] The oil price is denominated in US dollars and $\Delta po_t$ 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

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

The variables are the same as in the previous sections. We have added an extra subscript $i$ for each country, country-specific fixed effects, $\theta_i$, and a stochastic error term $\varepsilon_{i,t}$. This model is denoted Minimum nesting model 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 $\Delta p^e_{t+1}$ in the model causes two econometric problems. The first is a relatively minor one, and arises because estimation proceeds by substitution of $\Delta p^e_{t+1}$ by the observable $\Delta p_{t+1}$, 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 with valid instruments, and we can do the same on our panel data set. Second, and more fundamentally, models with forward-looking rational expectation terms are not easily identified, see Pesaran (1987) and Mavroeidis (2004). In brief, rational expectations force a situation where valid instruments may also be weak instruments. As a practical solution, we include the 2. order lag of variables like inflation in the instrument list, which contributes to identification if the marginal model of e.g., $ulc_t$ does not depend on $\Delta p_{t-1}$. Other available variables may also be used as instruments. For example, since $\Delta ulc_t$ 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

[7] Of course, since we normalize on $\Delta p_t$, it is nominal import price growth that appears on the right-hand-side of the estimated equation.

Table 1: Panel unit root tests.

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

Note: The Levin-Lin-Chu test assumes common unit root processes (see 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) assume individual root processes.
P-values are given in parentheses.

and the unemployment benefit replacement ratio can be used as instruments. The full set of instruments is given in connection with the econometric results in the 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 was 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, Ch. 4) 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 have the same magnitude as in other studies, the bias is believed to be small.

The principle of balanced equations requires that the variables are either stationary or cointegrated. Macroeconomic time series are typically non-stationary, and we therefore have to investigate the order of integration of the main variables

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

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

Note: Tests 1-4 are based on the within panel estimator (see Hsiao, 1986). Tests 5-7 use the between dimension, see Pedroni (1999). The test are performed using Pedroni's RATS code (Pedroni, 2006). P-values are given in parentheses.

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 $\Delta p$. Hence, the unit root analysis indicate that the growth rates included in the dynamic 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 and with a rank equal to 1. 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[8] and presented in the same order as in Pedroni (1999).

While macro panels typically exhibit cross-sectional dependence, the panel unit root tests and the Pedroni panel data cointegration tests all assume cross-country independence. As shown by Banerjee, Marcellino and Osbat (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 thus to some extent correct for this form of cross-sectional dependence in the panel. Therefore we considered three cases regarding the cointegration space; one without time dummies and deterministic trends, one where time dummies were included, but not deterministic trends, and one where heterogeneous

[8]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.

deterministic trends and time dummies were included. The tests for cointegration are conducted in a static regression setting. The trends and dummies are included in order to correct for any potential contemporaneous correlation in the residuals not necessarily present in a preferred dynamic model.

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 favour 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.

The GMM estimator assumes spherical errors. Consequently, we should test for homoscedasticity and error independence in the panel regressions. Therefore, in the next section we present two estimators for each model. First, in 3, ordinary GMM estimators are presented, and then Table 4 introduces GMM estimators with Cross-Section SUR (PCSE) corrections of standard errors and covariances where we use the same instruments as in Table 3. The Cross-Section SUR estimator is robust to both panel heteroscedasticity and contemporaneous correlations in the errors.[9]

## 5 Econometric results

Table 3 reports the estimation results for the econometric OECD inflation models. NPC1 is the empirical equivalent to equation (2) above and represents the model that has been estimated on several data sets with results that are summarized in section 2. In NPC1, real marginal costs are measured in accordance with equation (3) above, i.e., by the wage share of gross value added, $ws_{i,t}$. NPC2 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. But as seen from the table, this distinction has virtually no effect on the estimation results. NPC3 is the empirical equivalent to equation (6) above, and is a reparameterisation of equation (2) with the additional assumption of a constant share of import prices in the consumer price index. The Minimum nesting model is the encompassing model equivalent to equation (9) above, i.e. the estimated equilibrium correction model, which under the assumption of a constant import share 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 $ws_{i,t-1}$, second lag of the unemployment rate ($UR_{t-2}$), the gross replacement rate ($BRR$) and its lag, and an index of employment protection ($EP$) and its lag. $(ulc_{i,t-1} - pi_{i,t-1})$ and $(ulc_{i,t-1} - p_{i,t-1})$ are additional instruments in the two models NPC1 and NPC2.

As can be seen, the results for NPC1, NPC2 correspond closely to the typical hybrid NPC model. In fact, the first three typical features listed in section 2 are clearly recognizable. Both the lagged and leading inflation terms have significant

[9] For further details on this estimation procedure see the EViews User's Guide, Quantity Micro Software (2005, Ch. 29).

Table 3: GMM estimation results for an OECD panel data set

| | NPC1<br>Eq. (2) | NPC2<br>Eq. (2) | NPC3<br>Eq. (6) | Minimum nesting model<br>Eq. (9) |
|---|---|---|---|---|
| $\Delta p_{i,t+1}$ | 0.56<br>(0.03) | 0.57<br>(0.03) | 0.50<br>(0.04) | 0.003<br>(0.127) |
| $\Delta p_{i,t-1}$ | 0.47<br>(0.03) | 0.46<br>(0.02) | 0.47<br>(0.03) | 0.43<br>(0.04) |
| $ws_{i,t}$ | −0.011<br>(0.01) | | | |
| $(ulc_{i,t} - p_{i,t})$ | | −0.005<br>(0.008) | | |
| $(ulc_{i,t-1} - p_{i,t-1})$ | | | 0.012<br>(0.010) | 0.058<br>(0.016) |
| $(ulc_{i,t-1} - pi_{i,t-1})$ | | | −0.008<br>(0.003) | −0.022<br>(0.007) |
| $\Delta ulc_{i,t}$ | | | 0.004<br>(−−) | 0.35<br>(0.07) |
| $\Delta(pi_{i,t} - p_{i,t})$ | 0.05<br>(0.01) | 0.05<br>(0.01) | 0.06<br>(0.02) | 0.12<br>(0.02) |
| $\Delta po_t$ | 0.005<br>(0.002) | 0.005<br>(0.002) | 0.007<br>(0.002) | 0.006<br>(0.002) |
| $\Delta VAT_{i,t}$ | 0.003<br>(0.0005) | 0.003<br>(0.0004) | 0.004<br>(0.0004) | 0.003<br>(0.0004) |
| # observ | 567 | 567 | 567 | 567 |
| $\hat{\sigma} \cdot 100$ | 1.29 | 1.29 | 1.26 | 1.11 |
| $\chi^2_{\mathsf{ival}}$ | 41.5[0.000] | 42.0[0.000] | 42.2[0.000] | 11.1[0.195] |
| $N_{\mathsf{AR\text{-}1}}$ | −3.07[0.002] | −3.02[0.002] | −2.95[0.003] | −0.48[0.63] |
| $N_{\mathsf{AR\text{-}2}}$ | −2.34[0.019] | −2.35[0.019] | −2.36[0.018] | −0.31[0.76] |

Notes: Square brackets, [..], contain p-values, heteroscedasticity consistent standard errors are in parentheses, (..). $\hat{\sigma}$ denotes the estimated residual standard error. $\chi^2_{\mathsf{ival}}$ denotes the Sargan Sargan (1964) specification test which is $\chi^2$ distributed under the null of valid instruments (degrees of freedom are 10, 10, 9 and 8 respectively). $N_{\mathsf{AR\text{-}1}}$ and $N_{\mathsf{AR\text{-}2}}$ have a standard normal distribution under the null of no 1st and 2nd order autoregressive errors, respectively.

coefficients; the sum of the coefficients cannot be statistically distinguished from unity, and the lead term gets the highest estimated coefficient. The only anomaly is the insignificance of the wage-share coefficients, which contradicts the typical NPC feature 4. In the outset, it cannot be ruled out that the aggregation bias analyzed by Imbs et al. (2005) is responsible for the insignificance of the wage share coefficient, but this view is not consistent with the large and significant estimate of the coefficient of the forward term (which should be equally affected if aggregation was the issue). Moreover, as mentioned above, Bårdsen et al. (2004) have documented that the wage-share coefficient is non-robust, even for estimation on the euro-area data used by GGL.

That the panel data results for the NPC are corroborating the typical finding on US and euro-area data, as well as on data of other countries 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.

Among the four model, it is the minimum nesting model which is most likely to be affected by weak instruments problems, since it has two instruments less than NPC1 and NPC2, i.e., $(ulc_{i,t-1} - p_{i,t-1})$ and $(ulc_{i,t-1} - pi_{i,t-1})$. Hence we first test the strength of the other 13 instruments used in the minimum encompassing model. The result is $F = 57$, which is much higher than the rule-of-thumb of 10, see Stock et al. (2002). As for NPC1 and NPC2, the 'first stage regression' for those model includes the two other significant predictors. In line with this, the test statistic for instrument irrelevance is $F = 76$ for NPC1 and NPC2. Incidentally, Bårdsen et al. (2004) report $F = 71$ for their re-estimation of the euro area NPC.

Since the typical features of NPC is also present in NPC3, which is an empirical valid reparameterisation of the New Keynesian Phillips curve on this panel data set, we can start to investigate the encompassing capabilities of the NPC by considering the coefficients (and standard errors) of $(ulc_{i,t-1} - p_{i,t-1})$, $(ulc_{i,t-1} - pi_{i,t-1})$, $\Delta ulc_{i,t}$ and $\Delta pi_{i,t}$ in NPC3 and in the Minimum nesting model. It is evident that $H_0^a$: $\beta_3 = \beta_1 + \beta_2$ will be rejected at any reasonable level of significance. For example, the estimated coefficient of $(ulc_{i,t-1} - p_{i,t-1})$ in the general model is 0.35, which is 10 times the size predicted by the NPC. The 't-value' is 4.5 and the 't-probability' is practically speaking *zero*. Hence on this test, the NPC does not parsimoniously encompass the minimum nesting model. Conversely, the hypothesis $H_0^c$: $\alpha^f = 0$ is not rejected, as can be seen directly from the estimation results for minimum nesting model, implying that the standard ICM, without a lead in inflation, does parsimoniously encompass the unrestricted (nesting) model.

The diagnostic tests at the bottom are also of interest for those who want to test the NPC as a valid econometric model. First, Sargan's test statistic $\chi^2_{\mathsf{ival}}$, which has an interpretation as an encompassing test, is significant for NPC1. As discussed in section 3, this outcome can be predicted on the basis of ICM being the valid model. Second, there is quite significant residual autocorrelation, while for the Minimum nesting model there are no signs of mis-specification. In sum, the results for the Minimum nesting model provide quite convincing evidence that inflation equilibrium corrects with respect to an open economy long-run price equation. Hence, our

Table 4: GMM estimation results for an OECD panel data set with Cross-Section SUR (PCSE) corrections of standard errors and covariances.

| | NPC1 | NPC2 | NPC3 | Minimum nesting model |
|---|---|---|---|---|
| $\Delta p_{i,t+1}$ | 0.59<br>(0.08) | 0.59<br>(0.08) | 0.50<br>(0.08) | 0.04<br>(0.11) |
| $\Delta p_{i,t-1}$ | 0.44<br>(0.05) | 0.43<br>(0.05) | 0.44<br>(0.05) | 0.43<br>(0.04) |
| $ws_{i,t}$ | −0.007<br>(0.02) | | | |
| $(ulc_{i,t} - p_{i,t})$ | | 0.007<br>(0.012) | | |
| $(ulc_{i,t-1} - p_{i,t-1})$ | | | 0.021<br>(0.013) | 0.055<br>(0.016) |
| $(ulc_{i,t-1} - pi_{i,t-1})$ | | | −0.013<br>(0.006) | −0.021<br>(0.006) |
| $\Delta ulc_{i,t}$ | | | 0.08<br>(−−) | 0.33<br>(0.06) |
| $\Delta(pi_{i,t} - p_{i,t})$ | 0.05<br>(0.02) | 0.05<br>(0.02) | 0.06<br>(0.02) | 0.11<br>(0.02) |
| $\Delta po_{i,t}$ | 0.004<br>(0.005) | 0.004<br>(0.005) | 0.006<br>(0.005) | 0.006<br>(0.003) |
| $\Delta VAT_{i,t}$ | 0.004<br>(0.001) | 0.004<br>(0.001) | 0.004<br>(0.001) | 0.003<br>(0.001) |
| # observ | 567 | 567 | 567 | 567 |
| $\hat{\sigma} \cdot 100$ | 1.31 | 1.30 | 1.25 | 1.08 |
| $N_{\text{AR-1}}$ | −9.08[0.000] | −9.03[0.000] | −8.06[0.000] | −1.46[0.15] |
| $N_{\text{AR-2}}$ | −6.66[0.000] | −6.61[0.000] | −5.98[0.000] | −0.52[0.60] |

Notes: See notes to Table 3.

interpretation of the cointegration tests in Table 2 is supported by the results for the dynamic econometric Minimum nesting model. This substantial conclusion is also robust to changes in the estimation methodology.

In Table 4 the equations are estimated with GMM with Cross-Section SUR (PCSE) corrections of standard errors and covariances using the same instruments as before, but now with a Cross-Section SUR instrument weighting matrix. This estimation technique corrects for both panel heteroscedasticity and contemporaneous correlations in the errors. The estimated coefficients change very little, in particular the estimated equilibrium correction coefficient in the Minimum nesting model is just as significant and the estimated coefficient of the forward term is again not significantly different from zero. Interestingly, the estimated coefficient of changes in the oil price is not longer significant. This confirms our interpretation above, namely that this term in the models in Table 3 corrects for some common shocks in the panel. The small changes in Table 4 compared with Table 3 indicate that the problems of contemporaneous correlations and panel heteroscedasticity are fairly small.[10]

[10]Interestingly the mulity-country study of Dees et al. (2008) first presents results which support the dominance of the forward-term in the standard NPC (and the dynamic homogeneity), when heterogeneity is accounted for, but then show that the dominance of the forward-term vanishes once variables representing foreign inflation is included in the NPC. These findings seem to be consistent our results.

# 6 Conclusion

GGL claim that the NPC represents a significant advance in inflation modelling which finally substantiates the dominance of forward-looking behavior in price adjustment. In this paper we have lifted the empirical testing of the NPC model from the calm waters of US and euro area data to the vast data ocean represented by a panel data set from 20 OECD countries. We are able replicate all typical features of estimated NPC model — thus the New Keynesian Phillips curve appears to hold its ground.

However, one main result of our analysis is that researcher's will be able to empirical confirm the NPC also when it is not an encompassing model. Hence an encompassing approach is a necessary ingredient in the empirical assessment of the NPC. We also find that the typical NPC is encompassed by an existing model known as the imperfect competition model (ICM) of wage and price setting. Specifically, our analysis show conclusively that the expected rate of future inflation and the wage-share serve as replacements for ICM specific equilibrium correction terms. Adding these terms to the NPC model critically affect the estimated coefficient of the forward term, not only is the coefficient insignificant, the point estimate is practically zero. Hence, the search goes on to find a congruent encompassing model which has lead in inflation as a statistically and economically significant explanatory variable.

# A Data definitions and sources

The data consist 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 Organization for Economic Cooperation and Development (OECD) Economic Outlook and Main Economic Indicators (MEI) Databases.[11] This should help ensuring consistency in the data set.

*Description of the variables*

$P$ : *Consumer prices.* $P$ is a consumer price index, 2000=100, retrieved from the Main Economic Indicator (MEI) OECD database.

$PI$ : *Price of imports.* The ratio of import value and import volume, both in domestic currency, is used as a proxy for the price of imports. The series are retrieved from the MEI OECD database.

$PO$ : *Price of oil.* The world dated price of Brent crude oil measured in USD per barrel is retrieved from the MEI OECD database.

$UR$ : *Rate of unemployment.* The OECD standardized unemployment rates give the number of unemployed persons as a percentage of the civilian labour force. The series are retrieved from the MEI OECD database.

$ULC$ : *Unit Labour Costs.* $ULC$ is an index of unit labour costs (2000=100) and is retrieved from the MEI OECD.

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

Table 5: 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 |

*VAT* : *Indirect tax rate.* This is standard VAT rates in per cent 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' respective national bureaus of statistics. VAT rates for the United States are missing and are therefore assumed to be constant in the analysis.

*EP: Employment protection.* The data comprise an index of the degree of employment protection, and are provided by Dr. Luca Nunziata, Nuffield College, University of Oxford, UK, see Nunziata (2005). The series are extended with the 1995 value for the years 1996–2004.

*BBR: Benefit Replacement Ratio.* The data comprise an index of unemployment benefits in per cent of the average wage level, and are provided by Dr. Luca Nunziata, Nuffield College, University of Oxford, UK, see Nunziata (2005). The series are extended with the 1995 value for the years 1996–2004.

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