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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. As a explanatory model of OECD inflation, the NPC is encompassed by an existing theory.

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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 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, scientific inference requires consideration of *all* the properties and implications of a chosen or maintained interpretation of the correlations (not just 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 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 homogeneous slope coefficients across countries. 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. In line with this way of reasoning, our first finding in this paper is 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.

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 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 equilibirum correction variables suggested by the ICM can be rationalized under the hypothesis that the NPC holds. It is then straigth-forward 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 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

[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.

tests are given in 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 Δp^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. 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)[4] 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 preferred operational definition of firms' marginal costs of production.[5]

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

[5] 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).

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, models already exist which (claim to) explain inflation, 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 GGLs 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. 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 (2006). 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 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

[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 omit that information from the assessment of their new model.

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, 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 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 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. above, but as noted above, 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

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

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

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

or

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

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 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 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 are believed to be important in the shorter run make up the short run part. For simplicity 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:

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

Hence, a comparison of 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, namely H_0^a: $\beta_3 = \beta_1 + \beta_2$ and H_0^b: $\beta_4 = -\beta$. Hence, the rejection of H_0^a and/or H_0^b are inconsistent with the OE-NPC. The same applies if H_0^c: $\alpha^f = 0$ cannot be rejected statistically based on estimation of (9): this test-outcome is inconsistent the main assumption of the NPC, namely that a significant proportion price setters are forward looking in the rational expectations sense. Finally H_0^d: $\alpha^b = 0$ can also be tested using (9).[7]

[7] 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.

The tests of the significance of the forward and lagged inflation terms, H_0^c: $\alpha^f = 0$ and H_0^d: $\alpha^b = 0$, 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.

The ICM interpretation implies fewer testable restriction on (9). Under the given dynamic specification, the ICM does however require that $\beta_1 > 0$ and $0 > \beta_2 > -\beta_1$. 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

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.

There is a separate 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.[8] 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

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

The variables are the same as in the previous sections. We have added an extra subscript i for each country, country-specific fixed effects, θ_i, 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.

[8] 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 presence of Δp^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 Δp^e_{t+1} 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 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 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 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 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 part of model (10) seem to be stationary.

We also test for cointegration between the variables that make up the equi-

Table 1: Panel unit root tests.

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)

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.

librium 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[9] 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 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 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.

[9]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.

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.

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.[10]

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.

[10] 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

	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)
$ws_{i,t}$	-0.011 (0.01)		
$(ulc_{i,t} - p_{i,t})$		-0.005 (0.008)	
$(ulc_{i,t-1} - p_{i,t-1})$			0.053 (0.014)
$(ulc_{i,t-1} - pi_{i,t-1})$			-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 (0.002)	0.005 (0.002)
$\Delta VAT_{i,t}$	0.003 (0.0005)	0.003 (0.0004)	0.003 (0.0004)
# observ	567	567	567
$\hat{\sigma} \cdot 100$	1.29	1.29	1.00
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, heteroscedasticity consistent standard errors are in parentheses, (..). $\hat{\sigma}$ denotes the estimated residual standard error. χ^2_{ival} denotes Sargan's (Sargan, 1964) specification test which is χ^2 distributed under the null of valid instruments (degrees of freedom are 10, 10 and 8 respectively). $N_{\mathsf{AR\text{-}1}}$ and $N_{\mathsf{AR\text{-}2}}$ have a standard normal distribution under the null of no 1. and 2. order autoregressive errors, respectively.

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 (BRR) and its lags, and an index of employment protection (EP) 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 models M1 and M1'.

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 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. In the outset, it can not 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 its not consistent with the large and singificant estimate of the coefficient of the forward term (which should be equally affected if aggregation was the issue). However, as mentioned above, Bårdsen et al. (2004) have documented that the wage-share coefficient is non-robust, even on the euro-area data used by GGL.

That the $\mathsf{M1}$ results 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.

As shown in the previous sections, significance of the forward-term in $\mathsf{M1}$ should carry over to $\mathsf{M2}$ 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 $\mathsf{M2}$. The coefficient is in fact estimated to be very close to zero. The dominance of the forward term in $\mathsf{M1}$ 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 $(ulc_{i,t-1} - p_{i,t-1})$, $(ulc_{i,t-1} - pi_{i,t-1})$, $\Delta ulc_{i,t}$ and $\Delta pi_{i,t}$ it is also evident that both H_0^a: $\beta_3 = \beta_1 + \beta_2$ and H_0^b: $\beta_4 = -\beta_2$ will be rejected at any reasonable level of significance: the estimated coefficient of $(ulc_{i,t-1} - p_{i,t-1})$ is 0.32, which is 10 times the size predicted by the NPC, and β_4 is more than 4 times bigger than $-\beta_2$.[11]

The diagnostic tests at the bottom of the table also convey bad news for the NPC: In $\mathsf{M1}$, the Sargan test statistic χ^2_{ival} is significant, and there is indication of quite significant residual autocorrelation (also of 2. order). For $\mathsf{M2}$ there are no signs of mis-specification. In sum, the results for M2 provide quite convincing evidence that inflation equilibrium corrects with respect to an open economy long-run price equation. Hence, our interpretation of the cointegration tests in Table 2 is supported and strengthened by the results for the dynamic econometric model M2. This substantive 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 M2 is just as significant and the estimated coefficient of the forward term is 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. Overall, small changes in Table 4 compared with Table 3 indicate that the problems of contemporaneous correlations and panel heteroscedasticity are fairly small.

[11] The 't-statistic' is 46.8 and 8.4 when testing H_0^a and H_0^b, respectively.

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

	M1	M1'	M2
$\Delta p_{i,t+1}$	0.59 (0.08)	0.59 (0.08)	0.03 (0.10)
$\Delta p_{i,t-1}$	0.44 (0.05)	0.43 (0.05)	0.38 (0.04)
$ws_{i,t}$	-0.007 (0.02)		
$(ulc_{i,t} - p_{i,t})$		0.007 (0.012)	
$(ulc_{i,t-1} - p_{i,t-1})$			0.051 (0.014)
$(ulc_{i,t-1} - pi_{i,t-1})$			-0.019 (0.005)
$\Delta ulc_{i,t}$			0.30 (0.05)
$\Delta pi_{i,t}$			0.10 (0.015)
$\Delta(pi_{i,t} - p_{i,t})$	0.05 (0.02)	0.05 (0.02)	
$\Delta po_{i,t}$	0.004 (0.005)	0.004 (0.005)	0.005 (0.003)
$\Delta VAT_{i,t}$	0.004 (0.001)	0.004 (0.001)	0.003 (0.0008)
# observ	567	567	567
$\hat{\sigma} \cdot 100$	1.30	1.30	0.98
$N_{\text{AR-1}}$	$-9.08[0.000]$	$-9.03[0.000]$	$-1.13[0.26]$
$N_{\text{AR-2}}$	$-6.66[0.000]$	$-6.61[0.000]$	$-0.46[0.64]$

Notes: Square brackets, [..], contain p-values, heteroscedasticity consistent standard errors. $N_{\text{AR-1}}$ and $N_{\text{AR-2}}$ have a standard normal distribution under the null of no 1. and 2. order autoregressive errors, respectively.

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, the main contribution of our analysis is that we are able to show that this result is to be expected also when the NPC is a seriously flawed model, and that the typical NPC is encompassed by an existing framework for inflation modeling known as the imperfect competition model (ICM) of wage and price setting. Specifically, our analyses 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 also very close to zero.

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.[12] This should help ensuring consistency in the dataset.

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.

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

[12] 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

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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