Subjective Expectations and New Keynesian Phillips Curves in Europe
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* Views expressed are those of the individual author and do not necessarily reflect official positions of De Nederlandsche Bank.
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Abstract

This paper assesses the empirical performance of the forward-looking new Keynesian Phillips curve (NKPC) in France, Germany and Italy for the period 1991.3-2004.4. Instead of imposing rational expectations, I use direct measures of inflation expectations constructed from Consensus Economics survey data. Dependent on the real marginal costs measure, I obtain significant and plausible estimates for the quarterly discount factor and the price rigidity parameter. When analyzing the role of lagged inflation, I find that only in France lagged inflation does not have explanatory power beyond predicting expected inflation. This suggests that only in France the standard forward-looking NKPC effectively captures quarterly inflation dynamics.

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1 INTRODUCTION

Presently, a prominent framework for analyzing short-run inflation dynamics is the new Keynesian Phillips curve (NKPC). According to the NKPC, current inflation depends directly on current expectations of future inflation and the real marginal costs of production; that is, the NKPC predicts that inflation is a forward-looking phenomenon. This contrasts sharply with the traditional Phillips curve, which considers inflation to be a backward-looking phenomenon. Another difference between the old and the new Phillips curve is that the latter is explicitly derived from solid microeconomic foundations and the former is not. This gives the NKPC two advantages over the traditional Phillips curve. First, it is less exposed to the Lucas (1976) critique; an important advantage when structural breaks are present. Second, the NKPC framework allows one to obtain estimates of structural parameters like the parameter governing the degree of price rigidity; such estimates provide useful insights into the nature of inflation dynamics.

A critical issue in empirical research on the NKPC is which measure of real marginal costs is to be preferred. In the spirit of the traditional Phillips curve, early research has used real economic activity measures like employment and output. Although under certain conditions real marginal costs are proportional to the output gap, the empirical results have been rather discomforting.\(^1\) The problem, as emphasized by Fuhrer and Moore (1995), is that theory indicates that inflation should lead the output gap, yet exactly the reverse is observed in the data. Additionally these authors explain that the output-gap based NKPC predicts that credible central banks can achieve disinflations without loss of output, but this contradicts the empirical regularity that disinflations have normally been costly, even in countries with highly credible central banks like, for example, Germany (see Ball, 1994 and Clarida and Gertler, 1997).

Knowing this, Galí and Gertler (1999; henceforth GG), Sbordone (2002), and Galí, Gertler and López-Salido (2001; henceforth GGLS) employ a different marginal costs measure, namely real unit labor costs. GG and Sbordone (2002) find that the cost-based NKPC gives a good description of US inflation dynamics; GGLS repeat this result for the euro area. Conversely, recent studies on relatively open economies have shown that the real unit labor costs proxy is not always adequate, as it ignores (imported) material inputs. In these studies, estimates for the coefficient on real marginal costs are often insignificant.\(^2\)

Another critical issue is that the NKPC is frequently found unable to fully capture the large degree of persistence in inflation. Irrespective of the measure of

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1 Amongst others, Walsh (2003) discusses the conditions under which the output gap is linearly related to real marginal costs.

real marginal cost, lagged inflation appears to be an important determinant of current inflation. Largely motivated by this empirical observation, GG modify the NKPC to explicitly allow for the role of lagged inflation. The crucial assumption these authors make, is that not all expectations are formed rationally; the expectations of a fraction of economic agents are assumed to be formed according to a simple rule of thumb, which involves lagged inflation. This modified NKPC allows inflation to have both a backward- and a forward-looking component and therefore GG name it “the hybrid Phillips curve”.

An important weakness of abovementioned empirical studies is that the analysis is performed under the joint assumption of (partially) rational expectations and the NKPC; this excludes an unconditional empirical test of the economic model. To circumvent this, I use direct measures of inflation expectations which are created from Consensus Economics survey data. These subjective inflation expectations have not been previously used in this context. To the extent that the Consensus survey expectations are an accurate measure of inflation expectations, this research is able to test the NKPC unconditionally. It might be clarifying to note here that when direct measures of inflation expectations are used, the popular hybrid Phillips curve has no straightforward interpretation. The reason is that it is likely that the coefficient on lagged inflation will capture variation that should be attributed to the coefficient on expected inflation.

Similar studies have been performed by Roberts (1997, 1998) and Adam and Padula (2003; hereafter AP) for the US, and by Paloviita (2005) for the aggregate euro area and eleven EMU countries. In general, these papers find that it is quite plausible that the poor empirical fit of the NKPC is caused by the restrictive assumption of rational expectations. More specifically, AP and Paloviita (2005) find that both the output gap and real unit labor costs are adequate empirical proxies for real marginal costs once direct measures of inflation expectations are used. Notwithstanding these encouraging results, AP conclude that lagged inflation remains an important determinant of current inflation; its explanatory power goes beyond its ability to predict subjectively expected inflation.

Paloviita (2005)’s analysis on Europe is performed on a yearly basis using OECD inflation forecasts to measure inflation expectations. An advantage of the Consensus Economics survey data is that these data allow the construction of subjective inflation expectations series on a quarterly basis. This data frequency facilitates comparison with recent well-known quarterly analyses that assume rational expectations like, for example, GG and GGLS. A disadvantage associated with the Consensus Economics data is that it is somewhat limited with respect to both country and time coverage. Because of this, I focus on the three largest euro area economies (France, Germany and Italy) to obtain reasonable time coverage (1991.3-2004.4). Consequently, I investigate the empirical performance of the new Keynesian Phillips curve in France, Germany and Italy using subjective inflation
expectations. As the focus of the analysis is on the aforementioned critical issues, different real marginal costs proxies are considered and the role of lagged inflation is analyzed.

To anticipate the results of the paper, I find significant and plausible estimates for the structural parameters of the model once the appropriate measure of the real marginal costs of production is used. In the case of France and Germany the introduction of open-economy variables substantially improves the empirical fit. For Italy only the output gap leads to supportive empirical results. When analyzing the role of lagged inflation, I only find for France that lagged inflation does not have explanatory power beyond predicting expected inflation. This implicates that only in the case of France the forward-looking NKPC effectively captures quarterly inflation dynamics.

In what follows, Section 2 presents the benchmark NKPC and an open-economy extension. Section 3 discusses the data and analyzes the time series properties. Section 4 reports the estimation results and tests the role of lagged inflation. The fifth and last section concludes.

2 THE NEW KEYNESIAN PHILLIPS CURVE

This section starts with a discussion of the benchmark NKPC and explains how it can be derived when expectations are subjective. Given the fact that the individual euro area countries are relatively open economies, Sub-section 2.2 presents a possible open-economy extension of the NKPC.

2.1 Benchmark model

The microeconomic theory underlying the NKPC runs as follows. In the theoretical economy, there is a continuum of monopolistically competing firms, which is normalized to one. Every one of these firms produces a differentiated good and faces a Dixit and Stiglitz (1977) demand curve. In the absence of nominal price rigidity, the firms would not be restrained to reset their prices optimally in every time period. In such a world, current prices would depend only on current nominal marginal costs (the desired markup is assumed fixed) and inflation would not be a forward-looking phenomenon. However, when a nominal price rigidity like price adjustment costs (Rotemberg, 1982) or time-dependent price-setting (Calvo, 1983) is included in the theoretical economy, firms care about future developments in
marginal cost. Following Calvo (1983) assuming that a firm has a fixed probability, $(1-\theta)$, that it can reset its price, one can derive \(^3\)

$$\pi_t = \beta E[\pi_{t+1}] + \lambda \tilde{mc}_t,$$

(1)

where $\pi_t$ is current inflation, $E[\pi_{t+1}]$ is rationally expected inflation, $\tilde{mc}_t$ is real marginal costs (averaged across firms), and $\lambda$ denotes that the variable is in percentage deviation from its steady-state value. Equation (1) is the NKPC in ‘reduced form’ because it conceals that $\lambda$ is a function of the structural parameters:

$$\lambda = \frac{(1-\theta)(1-\beta\theta)}{\theta},$$

(2)

where $0<\beta \leq 1$ is the subjective discount rate, and $\theta$ represents the fraction of firms that keep their price constant. Substituting Definition (2) into Equation (1) yields the rational expectations NKPC in structural form

$$\pi_t = \beta E[\pi_{t+1}] + \frac{(1-\theta)(1-\beta\theta)}{\theta} \tilde{mc}_t.$$

(3)

According to Equation (3), an increase in $\beta$ implies that price setters give more weight to future expected profits; as a consequence inflation is less sensitive to current marginal costs ($\lambda$ is decreasing in $\beta$). Note additionally that a rise in $\theta$ also entails a ceteris paribus decrease in $\lambda$.\(^4\) Since increased price rigidity (higher $\theta$) means longer average price duration, $D \equiv 1/(1-\theta)$, firms place less weight on current marginal costs and more on expected future developments in marginal costs.

Because Equation (3) is derived under the assumption of rational expectations, it is not directly obvious what the subjective expectations NKPC looks like. Intuitively one would simply replace rationally expected inflation, $E[\pi_{t+1}]$, with the average of subjective inflation expectations, $\bar{E}[\pi_{t+1}]$, yet AP have shown that this is not valid per se. According to AP, the validity of this substitution depends on whether or not expectations fulfill the law of iterated expectations. This law states that today’s expectations of next period’s expectations of certain future variables are equal to today’s expectations of these future variables.

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\(^3\) In Appendix 5.7.3 of his book, Walsh (2003) gives a step-by-step derivation of the benchmark NKPC from first principles.

\(^4\) The separate effects from changes in $\beta$ and $\theta$ on $\lambda$ can be derived from the partial first-order derivatives of $\lambda = \theta^{-1} - 1 - \beta + \beta\theta$. Given $\beta \leq 1$ and $\theta \in (0,1)$, $\partial\lambda/\partial\beta = -1 + \theta < 0$ and $\partial\lambda/\partial\theta = -\theta^{-2} < 0$. 

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variables. When subjective expectations do not violate this law, the subjective expectations NKPC can be derived from first principles. The intuition behind this can be seen from the following. Since firms face a Calvo probability $\theta$ that they cannot reset their price, $1-\theta$ firms reset their price based on their beliefs about future developments in nominal marginal costs. As a result inflation in $t+1$ depends on subjective expectations at time $t$ and $t+1$. Consecutively time $t$ forecasts of inflation are forecasts of other firms’ marginal costs forecasts at time $t$ and $t+1$. The average of time $t$ inflation forecasts, $\bar{\pi}^t[\pi_{t+1}]$, sufficiently summarizes these expectations when firms do not expect that current period’s forecasts will be adjusted in a particular direction in the next period. Hence, under the assumption that expectations fulfill the law of iterated expectations, the subjective expectations NKPC can be derived as

$$\pi_t = \beta \bar{\pi}^t[\pi_{t+1}] + \lambda \bar{rmc}_t. \quad (4)$$

Since Definition (2) defines parameter $\lambda$ in Equation (4), the subjective expectations NKPC in structural form is

$$\pi_t = \beta \bar{\pi}^t[\pi_{t+1}] + \frac{(1-\theta)(1-\beta\theta)}{\theta} \bar{rmc}_t. \quad (5)$$

From the previous discussion one cannot infer which variable or combination of variables is the appropriate measure for real marginal costs; additional assumptions are needed. Walsh (2003) explains that under certain restrictions on technology, preferences and the labor market, real marginal costs can be related to a measure of the gap between potential and actual output (that is an output gap measure). GG and GGLS show that real marginal costs is proportional to real unit labor costs, when one makes the additional assumptions that production technology is Cobb-Douglas, the labor market is competitive and there are no adjustment costs. Other possible real marginal costs measures are the open-economy measures discussed in the next sub-section.

### 2.2 Open-economy extension

The preceding sub-section described the benchmark NKPC; a framework well-suited for the analysis of inflation in closed economies. GG apply this framework to US inflation and GGLS use it for the analysis of euro area inflation. As the economies of France, Germany and Italy are relatively more open than the US and the euro area, it is appealing to analyze the effect of imports on domestic inflation.

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Approaches in the open-economy NKPC literature can be distinguished by whether they consider only final consumption goods imports, only intermediate goods imports or both types of imports. Under the assumption that all imports are final consumption goods, Kara and Nelson (2003) report that their open-economy extension gives a poor description of UK inflation. Typically more favorable results are reported by authors who consider intermediate goods imports. Galí and López-Salido (2001) and Balakrishnan and López-Salido (2002) find for Spain and the UK, respectively, a significant role for relative import prices in inflation dynamics. On the other hand, Sondergaard (2003b) shows for France, Italy and Spain that allowing for imported intermediate inputs in the NKPC does not change the empirical results much.

Leith and Malley (2003) and Rumler (2005) analyze the most general open-economy NKPC models which incorporate both intermediate and consumption goods imports. For the G7 countries Leith and Malley (2003) find that their open-economy estimates are not significantly different from their closed-economy estimates. By contrast, Rumler (2005) concludes that open economy aspects matter for the fit of the NKPC. He analyzes euro area countries and the euro area aggregate.

Here I adopt the open-economy extension proposed by Galí and López-Salido (2001) and Balakrishnan and López-Salido (2002). These authors use the following constant elasticity of substitution (CES) production function

$$
Y_t = F(N_t, M_t) = \left[ \alpha_N (Z_t N_t)^{1+1/\sigma} + \alpha_M (M_t)^{1-1/\sigma} \right]^{\sigma},
$$

(6)

where $Y_t$ is aggregate production, $Z_t$ is a labor-augmenting technology shock, $\sigma$ is the elasticity of substitution between domestic and imported inputs, $N_t$ is labor input, $M_t$ is total intermediate goods imports and $\alpha_N$ and $\alpha_M$ are the weights of these factors in the production function. From cost minimization one can derive the equilibrium condition

$$
\frac{N_t}{M_t} = \left( \frac{\alpha_N}{\alpha_M} \frac{P^M_t}{W_t} \right)^{1/\sigma},
$$

(7)

where $W_t$ is the nominal wage and $P^M_t$ is an index of imported intermediate goods prices. According to Equation (7), the ratio between the production inputs is a function of their relative prices. Balakrishnan and López-Salido (2002) show in their appendix that under these circumstances, the open-economy expression for real marginal costs is

$$
\widetilde{orc}_t = \widetilde{rulc}_t + \phi (P^M_t - w_t),
$$

(8)
where lower case variables are natural logarithms and $\hat{\cdot}$ denotes once more that the variable is in percentage deviation from its steady-state value. $\phi$ is defined as

$$\phi \equiv [1/(S\mu) - 1][\sigma - 1],$$

(9)

where $\mu$ and $S$ are the steady-state price markup over nominal marginal costs and the steady-state labor income share, respectively. Note that the labor income share is equivalent to real unit labor costs.

Equation (8) states that in an open-economy, the real marginal costs of production depend on real unit labor costs (as in the closed economy) and additionally on the relative price of domestic and foreign production inputs $(p^M_t - w_t)$. Because the price of imports is strongly related to the exchange rate, this additional term allows movements in the exchange rate to affect inflation dynamics. $\phi$ determines how relative price movements affect real marginal costs. According to Definition (9), $\phi$ can take either sign as $0 < \mu S < 1$. When $\sigma > 1$, an increase in the prices of imported materials below the increase in the nominal wage will decrease real marginal costs. Moreover, as $\sigma$ grows the importance of the relative price term also grows.

Substituting Equation (8) into Equations (4) and (5) for $\widehat{rmc}$ gives us the open-economy extension of the subjective expectations NKPC in reduced and structural form, respectively.

3 DATA ISSUES

Here I describe the data that is used to estimate the different New Keynesian Phillips Curves (NKPCs) for France, Germany and Italy. Additional information on the variable definitions and data sources can be found in the Appendix. This section is organized as follows. The first sub-section reports on the unique inflation expectation series, the second is about the empirical measures of real marginal costs and the final sub-section analyses the time series properties of the series.

3.1 Subjective inflation expectations from Consensus Economics

From Section 1 we know that it is desirable to have the subjective expectation series on a quarterly basis. For European countries, the only series with this characteristic are the quarterly forecasts from Consensus Economics. Nevertheless,

6 Note that $\phi = 0$ when the production technology is Cobb-Douglas ($\sigma \to 1$). In that case the open-economy extension is indistinguishable from the benchmark NKPC with real unit labor costs measuring real marginal costs.
Consensus series have not been previously used in this context. For this a possible reason could be that the expectations series need to be modified to fit the theoretical framework. Below I explain in detail how I derived inflation expectations from the quarterly forecasts.

Since 1989, Consensus Economics polls prominent financial and economic forecasters to obtain their estimates for several countries and several macroeconomic variables. The polls are done on a monthly basis and the results are published monthly in Consensus’ publication Consensus Forecasts.

At first glance the average of growth forecasts of consumer prices is a suitable proxy for expected inflation (it is updated monthly). However, closer inspection reveals that this average of monthly growth forecasts of consumer prices cannot be used. The reason for this is that a model with microeconomic foundations lays strict requirements on empirical proxies. Specifically in the NKPC case, the inflation expectation at time $t$ should correspond to the expected percentage increase in the level of the price index between $t$ and $t+1$. The monthly inflation forecasts do not satisfy this, as these refer to expectations of growth in consumer prices in the current and the next calendar year. Because these 'events' are fixed, the forecasting horizon changes from month to month.

Fortunately, Consensus Economics also publishes a quarterly special edition of Consensus Forecasts. In this special edition, averages of consumer price growth forecasts are given for individual quarters. An example of these forecasts is given in Table 1, which shows averages of forecasts made for Germany in the fourth quarter of 2004.

Table 1  Consumer price growth forecasts given in 2004.4 for Germany

<table>
<thead>
<tr>
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<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Consumer Prices (%)</td>
<td>2.0</td>
<td>1.8</td>
<td>1.4</td>
<td>1.3</td>
<td>1.2</td>
<td>1.2</td>
<td>1.3</td>
</tr>
</tbody>
</table>

*Note: Percentages are averages of expected year-on-year changes in consumer prices.*

The growth percentages are averages of expected percentage changes in consumer prices *over the previous year*. For the current example this means that in the fourth quarter of 2004, the forecasters expect, on average, a 1.8 percentage rise in consumer prices between the first quarter of 2004 and the first quarter of 2005. Interestingly, the forecasters also give an estimate of the percentage value for the current quarter. This conveys that current actual inflation is not exactly known by them. Because interest lays in quarter-to-quarter inflation expectations, these year-to-year forecasts cannot be used directly.

Hence, suitable quarterly expected inflation series need to be constructed; this can be done in three steps. The first step is a conceptual one: interpret the
growth forecasts of consumer prices as expectations of annual inflation in consumer price index (CPI). Under this interpretation, the currently expected CPI level for the current and for the next quarter can be calculated from the ex-post CPI; this second step is illustrated graphically in Figure 1 for Germany for 2004.4.

Figure 1  Construction of expected current and next quarter CPI

<table>
<thead>
<tr>
<th>Year</th>
<th>CPI</th>
<th>E[CPI]</th>
</tr>
</thead>
<tbody>
<tr>
<td>2003.4</td>
<td>104.95</td>
<td></td>
</tr>
<tr>
<td>2004.1</td>
<td>105.25</td>
<td></td>
</tr>
<tr>
<td>2004.2</td>
<td>105.94</td>
<td></td>
</tr>
<tr>
<td>2004.3</td>
<td>106.45</td>
<td></td>
</tr>
<tr>
<td>2004.4</td>
<td>107.05</td>
<td></td>
</tr>
<tr>
<td>2005.1</td>
<td>107.15</td>
<td></td>
</tr>
</tbody>
</table>

Notes: This example refers to Germany, 2004.4. The depicted growth forecasts are averages of forecasts made in the fourth quarter of 2004 for respectively that quarter and the next quarter (the first quarter of 2005). Combining the Consensus Economics’ growth forecasts and the ex-post CPI series (see the text) yields the expected CPI, E[CPI], series.

The third and final step is now straightforward: calculate for every period the expected quarter-to-quarter inflation as the quarter-to-quarter change in that period’s expected CPI.

For the three countries, plots of expected and ex-post/actual inflation are presented in Figure 2. The time frame 1991.3-2004.4 is selected on the basis of availability of the quarterly forecasts in Consensus Forecasts. The general picture is that actual and expected CPI inflation move closely together. On closer examination, it appears that forecasters in Germany and Italy made serious forecast errors in the early nineties. This is not surprising since for both countries the early nineties were turbulent years. Germany went through its reunification process and the Italian lira dropped out of the exchange rate mechanism (ERM) in September 1992. In recent years, both actual and expected inflation fluctuate around 2% annualized CPI inflation.

7 These annualized inflation series are calculated for expositional purposes; in the analysis non-annualized inflation is used.
8 Actually, the quarterly forecasts go further back. These earlier forecasts are not used, because the same forecast is published for the first two quarters of 1991 (this data irregularity applies to all countries under analysis).
9 Before the fourth quarter of 1996, averages of consumer price growth forecasts for Germany only apply to West Germany. Potentially, this could have an effect on the inflation expectation parameter in Germany; nonetheless, I did not find significant evidence of this.
10 Interestingly, it seems that economic forecasters expected an immediate surge in Italian consumer prices, yet this did not materialize.
Figure 2 Actual and expected quarter-to-quarter inflation
1991.3-2004.4, annualized percentages

France

Germany

Italy

actual inflation
expected inflation in the previous period
3.2 Real marginal costs measures

Section 2 has discussed two main alternatives for measuring real marginal costs in the closed-economy NKPC: real unit labor costs and the real output gap. In the open-economy extension, the relative price of foreign and domestic inputs is added to real unit labor costs to account for the openness of the economies under consideration. As noted in Section 1, the assumption of rational expectations is not innocuous when it comes to determining the suitable measure for real marginal costs. By using a direct measure of expected inflation, AP and Paloviita (2005) uncover that both the conventional output gap and real unit labor costs are adequate empirical measures of real marginal costs. Since this research uses direct expectation measures as well, it is clear that the analysis should be done for different empirical proxies.

Output gap

For the three countries under consideration, output gap series are derived from Eurostat series of real GDP at basic prices. As the output gap is a measure of the gap between actual output and potential output, one needs an empirical measure of potential output. Here the widely applied Hodrick-Prescott (HP) filter is used with smoothing parameter 1,600. This filter is applied to the natural logarithm of real GDP at basic prices over the time period 1991.1-2004.4.\(^{11,12}\)

The choice of real GDP at basic prices, rather than at market prices, is not an arbitrary one. In an empirical evaluation of the NKPC model, the influences of the public sector should be minimized; the model explains the behavior of a firm in market economy and ignores the role of the government in price developments. Since series at market prices include indirect taxes, the series at basic prices are preferred.\(^{13}\)

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\(^{11}\) The filter is applied over a period starting two quarters earlier than the inflation series, to lessen the end-of-sample problems of the HP filter at the beginning of the sample. Additional improvement to the inefficiency of the HP filter could be obtained by using a forecast-based augmentation to the filter (see Mise, Kim and Newbold, 2005).\(^{12}\)

Although it would be very interesting to analyze the joint performance of theory-based output gap series \(à la\) Neiss and Nelson (2002) and subjective inflation expectations, it is beyond the scope of this paper.\(^{13}\)

Ideally we would also exclude the share of production that is attributable to the government. Unfortunately, as Sondergaard (2003b) notes, this cannot be done with Eurostat data. Something that can be done with Eurostat data, is to construct output gap series that exclude agriculture, hunting, forestry and fishing; such series would closely resemble the non-farm activity measure frequently used in the literature. Yet, since my measure of price developments is the CPI, it does not seem correct to exclude this 'farm' activity. Consequently, aggregate output gap series are used.
Real unit labor costs
If production technology is Cobb-Douglas, the labor market is competitive over the whole economy and firms face no adjustment costs, it can be shown that real marginal costs is proportional to real unit labor costs, or equivalently the labor income share.

Amongst others, Sondergaard (2003b) and Coenen and Levin (2004) note that in measuring the labor income share, it is important to include labor compensation for self-employed workers. In the absence of data, the standard approach is to assume that self-employed workers earn the average of compensations received by employees. I adopt this approach and define real unit labor costs as the product of two ratios. The first ratio is total compensation to employees divided by nominal GDP at basic prices (the employee share of income); the second ratio is total employment divided by the number of employees (the correction factor for self-employed labor). Once more, GDP at basic prices is used to minimize influences from the public sector (see Batini, Jackson and Nickell, 2000).

To be comparable with the output gap, I construct real unit labor costs series for the period 1991.1-2004.4. The natural logarithm of these series, \( \text{rulc} \), is graphed for the three countries in Figure 3. These series cannot be used directly since theory links inflation and real marginal costs in deviation from steady state \( \text{rmc} \). A frequently used proxy of real unit labor costs’ steady-state is the sample average. In Figure 3, \( \text{rulc} \) sample averages are represented by horizontal, dotted lines. Using these sample averages \( \text{rmc} \) is defined as \( \text{rulc} \) minus its sample average.

In addition to this measure, I define \( \text{rulc} \) in deviation from its linear trend as a measure of \( \text{rmc} \). Figure 3 shows clearly that there are downward trends in all three countries in the period 1991-2004.\(^\text{14}\) In the empirical analysis I give an interpretation of these downward trends.

Open-economy measures
In Sub-section 2.2 I have explained that the open-economy measure of real marginal costs depends on real unit labor costs and the relative price of labor and imported intermediate goods \( (p^M_t - w) \). I measure the intermediate import prices index \( p^M_t \) by the natural logarithm of the Hamburgisches Welt-Wirtschafts-Archiv (HWWA) index of world market prices of raw materials (in national currency). The empirical measure for \( w \) is the natural logarithm of total compensation received per employee (in national currency). The natural proxy for the steady-state

\(^{14}\) An interesting additional measure could result from fitting a linear trend to the \( \text{rulc} \) series before 1999 and a constant thereafter. Nevertheless, the empirical results using such a measure were dissatisfying.
Figure 3  Natural logarithm of real unit labor costs
1991.1-2004.4

France

Germany

Italy

- linear trend in rulc
- rulc
- sample average rulc
of the relative price term \((p_t^M - w_t)\) is its sample average. Above I have described two possible empirical measures of real unit labor costs in deviation from steady-state. Here I use real unit labor costs in deviation from trend.\(^{15}\)

What remains to be addressed is the empirical implementation of \(\phi\). According to Definition (9), \(\phi\) depends on the steady-state markup \(\mu\), the steady-state labor income share \(S\) and the elasticity of substitution \(\sigma\). Because I use real unit labor cost in deviation from trend, the obvious proxy for \(S\) is the linear trend in real unit labor costs. Given the assumptions about the economy, the steady-state markup is equal to the fixed desired markup \(\mu = \varepsilon / (\varepsilon - 1)\), where \(\varepsilon\) is the elasticity of demand. A frequently used value for \(\varepsilon\) is 1.1 which amounts to a steady-state markup of \(\mu = 1.1\) (see GGLS, Leith and Malley, 2003 and Rumler, 2005). Since the value of \(\sigma\) is unknown and its estimation lies outside the scope of this paper, I follow the literature in varying the elasticity of substitution around the Cobb-Douglas benchmark value \(\sigma \rightarrow 1\).

Figure 4 plots the open-economy measure for the three economies for different values of \(\sigma\), namely \(\sigma = 0.9\), \(\sigma \rightarrow 1\) and \(\sigma = 1.1\). These values have also been used by Balakrishnan and López-Salido (2002) for the UK.\(^{16}\) The figure shows that the measures move considerably apart between (the end of) 1997 and 2001. This period of turbulence coincides with the Asian crisis of 1997 and its aftermath. An inspection of the HWWA indices and the compensation series reveals that in this turbulent period, movements in raw material prices dominate the relative price term \((p_t^M - w_t)\). This observation combined with the intuitive appeal of a positive relation between intermediate (import) prices and marginal costs gives an indication that open-economy measures with \(\sigma > 1\) will outperform the measures with \(\sigma < 1\). This seems in line with recent evidence on the elasticity of substitution between capital and labor.\(^{17}\) Berthold, Fehn and Thode (2002) find for France and Germany for the period 1970-1995 that the substitution elasticity between capital and labor is significantly greater than one.

3.2 Time series properties

Since the econometric consequences of nonstationarity are quite severe, it is good practice to analyze the univariate stochastic processes that best characterize the data. Remarkably, influential papers like GG and GGLS pass over stationarity

\(^{15}\) I have also experimented with real unit labor costs in deviation from its sample average. When using that proxy the empirical results are far less favorable.

\(^{16}\) Because these values are arbitrary, I follow Balakrishnan and López-Salido (2002) in examining a range of \(\sigma\)s in the empirical analysis.

\(^{17}\) I did not find any previous estimations of a CES production function using imported intermediate goods and labor as production factors.
Figure 4  Open-economy real marginal costs measures
1991.1-2004.4, percentage deviation from steady state

France

-3  -2  -1  0  1  2  3

91 92 93 94 95 96 97 98 99 00 01 02 03 04

sigma = 0.9
sigma => 1
sigma = 1.1

Germany

-3  -2  -1  0  1  2  3

91 92 93 94 95 96 97 98 99 00 01 02 03 04

sigma = 0.9
sigma => 1
sigma = 1.1

Italy

-6  -4  -2  0  2  4  6

91 92 93 94 95 96 97 98 99 00 01 02 03 04

sigma = 0.9
sigma => 1
sigma = 1.1
analysis. A possible explanation may be that in the current context, the risk of being accused of “spurious regression” is low; the NKPC is a macroeconomic relationship based on solid microeconomic foundations. Yet, the use of nonstationary time series has the additional adverse effect that the standard properties of econometric estimators no longer hold.\textsuperscript{18}

Firstly the stationarity of the data is tested by means of augmented Dickey-Fuller (ADF) tests. The null hypothesis of the ADF test is that the tested variable has a unit root. When the null hypothesis can be rejected at the five percent significance level, I conclude that the tested variable appears to be (trend) stationary. The number of lags to include in the testing regression is determined by following the general-to-specific approach advocated by Ng and Perron (1995). This approach amounts to including a sufficient number of lagged differences (I use a maximum lag length of 12) into the testing equation and deleting them one by one until the null hypothesis that the last lagged difference is insignificant can be rejected at the five percent significance level. The number of remaining lags is the number of lags that enters the actual ADF test.

The ADF test results are as follows. For the output gap series, the null is soundly rejected in all countries; this indicates that all output gap series are stationary. Furthermore I find that \textit{rule} in deviation from trend appears to be stationary for France and Germany; this cannot be concluded for Italy. The other closed-economy \textit{rmc} proxy, \textit{rule} in deviation from sample average, shows signs of nonstationary in all countries under consideration. Regarding the open-economy proxies, only for Italy the null hypothesis of a unit root cannot be rejected. Finally, the ADF tests for the inflation series indicate that in all countries both actual and expected inflation are nonstationary.\textsuperscript{19}

Since the ADF test has severe power problems in the presence of a structural break, alternative tests have been developed that allow for the presence of a structural break. The modifications can be distinguished by whether the structural break is regarded as exogenous or endogenous. Zivot and Andrews (1992) argue that an endogenous break test is more appropriate since it avoids the problem of data-mining associated with an exogenous break test. Table 2 shows the results of the Zivot and Andrews (1992) test for actual and expected inflation.\textsuperscript{20}

\textsuperscript{18} Of course, as Engle and Granger (1987) have pointed out, a linear combination of nonstationary time series may be stationary. When such a cointegrating relationship exists, parameter estimators are said to be super consistent (see for example Davidson and MacKinnon, 1993).

\textsuperscript{19} All ADF test results are available upon request.

\textsuperscript{20} The Zivot and Andrews test has also been used to verify the (non)stationarity of the real marginal costs proxies. Since the results are in line with the ADF test results, these are omitted.
Table 2  Zivot and Andrews unit root test

<table>
<thead>
<tr>
<th>Country</th>
<th>Variable</th>
<th>Test statistic ($^{5%} ZA_t$)</th>
<th>Break date</th>
</tr>
</thead>
<tbody>
<tr>
<td>France</td>
<td>$\pi_t$</td>
<td>-9.37</td>
<td>1997.1</td>
</tr>
<tr>
<td></td>
<td>$\tilde{F}[\pi_{t+1}]$</td>
<td>-6.69</td>
<td>1999.3</td>
</tr>
<tr>
<td>Germany</td>
<td>$\pi_t$</td>
<td>-7.76</td>
<td>1994.2</td>
</tr>
<tr>
<td></td>
<td>$\tilde{F}[\pi_{t+1}]$</td>
<td>-6.43</td>
<td>1994.4</td>
</tr>
<tr>
<td>Italy</td>
<td>$\pi_t$</td>
<td>-5.52</td>
<td>1996.2</td>
</tr>
<tr>
<td></td>
<td>$\tilde{F}[\pi_{t+1}]$</td>
<td>-5.89</td>
<td>1999.4</td>
</tr>
</tbody>
</table>

Notes: Zivot and Andrews (1992) unit root test results for actual inflation ($\pi_t$) and expected inflation ($\tilde{F}[\pi_{t+1}]$); sample: 1991.3 – 2004.4. The unit root null hypothesis is tested against the alternative of stationarity with a structural break in trend and intercept at an unknown point in time. $^{5\%} ZA_t$ is the appropriate asymptotic five percent critical value of the Zivot and Andrews test statistic. The tests are executed with GAUSS code provided by Junsoo Lee (http://www.cba.ua.edu/~jlee/gauss/).

The results suggest that when we allow for a one-time endogenous break, all inflation series are stationary over the period 1991.3-2004.4. Although it is not easy to give straight explanations of the specific break dates in Table 2, it is intuitively clear that all breaks most probably relate to the transition towards the Economic and Monetary Union (EMU). Concluding, it seems that rule in deviation from its sample average is nonstationary for all countries under consideration. Moreover, for Italy there is also evidence of a unit root in rule in deviation from its linear trend and in the open-economy measures. For comparison reasons I do not exclude these series from the analysis.

4 EMPIRICAL ANALYSIS

In this section I report estimates of the subjective expectations NKPC for France, Germany and Italy using quarterly data. Subsequently to the discussion of the estimation results, the role of lagged inflation is analyzed. Let us start however with some estimation issues.

4.1 Estimation issues

Because direct measures of inflation expectations are used, it is in principle possible to use ordinary least squares (OLS) to estimate the NKPC parameters. In fact, AP apply OLS in their paper on subjective expectations and the NKPC in the
US. However, OLS might be inconsistent since the explanatory variables are most likely measured with error. To assess the (in)consistency of OLS, Hausman (1978) tests have been used. The test results indicate that the null hypothesis of OLS consistency should be rejected at conventional significance levels for all countries considered. For this reason I prefer using an instrumental variables (IV) estimator.

As there are many IV estimators, a specific one needs to be chosen; here the two-stage least squares (2SLS) estimator is used instead of the more widely applied efficient generalized method of moments (GMM) estimator. Although the efficient GMM estimator is asymptotically never worse than the 2SLS estimator, it can have poor small sample properties as it makes use of fourth moment estimates (see Hayashi, 2000).

Under the assumption that OLS is inconsistent due to measurement errors, no obvious instruments arise. Evidently, instruments need to be both relevant and valid. Instruments are relevant when they are highly correlated with the explanatory variables and valid when they are orthogonal to the disturbance term. Obtaining relevant instruments is not the hard part. Given the considerable persistence in macroeconomic variables, lagged variables are generally highly correlated with current variables and thus highly relevant. However obtaining valid instruments is far more troublesome. The best one can do is assuming that the measurement errors are completely random and testing for the joint validity of the moment conditions. Like AP, this paper uses the Sargan test or overidentifying restrictions test for this purpose.\textsuperscript{21} The Sargan test statistic is easily calculated by taking \( N \) (the number of observations) times the uncentered \( R^2 \) of an auxiliary regression of the 2SLS residuals upon the full set of instruments (see Verbeek, 2004). The null hypothesis of the Sargan test is that the instruments are jointly valid.

Until now an unaddressed issue is how estimates for structural form parameter \( \theta \) can be derived. Obviously 2SLS cannot be applied to the NKPC in structural form, since the NKPC is nonlinear in the structural parameters. Fortunately the value of \( \theta \) can be determined from the point estimates of the reduced form parameters \( \beta \) and \( \lambda \).\textsuperscript{22} These point estimates are straightforwardly obtained by applying 2SLS to the reduced form NKPC.

\textsuperscript{21} An overidentifying restrictions test is only possible when there are more instruments available than there are parameters to be estimated.
\textsuperscript{22} Given point estimates for \( \beta \) and \( \lambda \), Definition (2) implies that \( \theta \) can take on two values: \( \theta = (2\beta)^{\dagger}(1+\lambda+\beta \pm \sqrt{4\beta+(1+\lambda+\beta)^2}) \). Only the implied value of \( \theta \) that is consistent with it being a probability is reported. The corresponding standard error can be derived from estimating the NKPC in structural form using nonlinear 2SLS.
4.2 Benchmark estimation results

This sub-section gives the benchmark estimation results for the three countries considered. The instrument sets used for France and Germany include lagged expected inflation, the lagged real marginal costs proxy, two lags of the short-term interest rate and a constant. The instrument set for Italy contains lagged expected inflation, the lagged short-term interest rate, two lags of the real marginal costs proxy, an ERM-crisis dummy and a constant. Relevance of these instruments is tested by regressing each explanatory variable on the related instrument set and testing the null hypothesis that all slope parameters are jointly zero. Because all null hypotheses are rejected at a significance level of one percent, I conclude that the relatively small instrument sets are capable of explaining the explanatory variables.24

Table 3 presents the estimates of Equation (4), when using the output gap (column two), real unit labor costs in deviation from sample average (column four) and real unit labor costs in deviation from trend (column six), respectively, as a proxy for real marginal costs. Below the parameter point estimates, Newey-West standard errors (robust to serial correlation up to three lags) are given in parentheses.25 Also reported are Sargan test statistics and corresponding p-values. The test results indicate that all instrument sets used are valid at the five percent significance level.

As Table 3 shows, the results differ greatly across both countries and real marginal costs proxies. On the positive side, all estimates for parameter $\beta$ are close to one, as is implied by the theory. Although for Italy the $\beta$ values are estimated slightly above one, estimating the model under the restriction $\beta = 1$ does not lead to significantly different results. With respect to parameter $\lambda$, the results are less encouraging. Since the coefficient on real marginal costs ($\lambda$) should be both positive and significant, almost all benchmark results are contradictory to the theory. The problem is that the positive point estimates of $\lambda$ are measured with relatively large standard errors.26 A noteworthy exception is the $\lambda$ estimate for Italy when using the output gap: the estimate is positive and significant at the five percent level.

---

23 For Italy I use a dummy variable to capture the exchange-rate turbulence of the Italian lira in the period 1992.3-1993.1. Because the dummy appears in the regression model, it is also included in the instrument set.
24 F-statistics range from 7.98 (for expected inflation in France) to 269.35 (for real unit labor costs in deviation from its sample average in Italy). Since the lowest F-statistic is very much larger than the corresponding one percent critical value (3.74), all null hypotheses are rejected at the one percent significance level.
25 Especially for Germany and Italy Ljung-Box autocorrelation tests indicate that there is significant serial correlation in the residuals. Hence, normal standard errors are inappropriate. Following the suggestion of Newey and West, the number of lags considered has been determined from the number of observations.
26 In view of the relatively small sample periods this is not surprising.
Table 3  Results benchmark NKPC  
1991.4-2004.4 (53 observations) 
\[ \pi_t = \beta \hat{F}(\pi_{t-1}) + \lambda \hat{m}_{t-1}, \lambda = (1-\theta)/(1-\beta) / \theta \] 

<table>
<thead>
<tr>
<th></th>
<th>Output gap</th>
<th>Real unit labor costs</th>
<th>Real unit labor costs</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>unrestricted</td>
<td>( \beta = 1 )</td>
<td>unrestricted</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>deviation from average</td>
</tr>
<tr>
<td>France</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( \beta )</td>
<td>0.901**</td>
<td>1.000</td>
<td>0.918**</td>
</tr>
<tr>
<td></td>
<td>(0.090)</td>
<td></td>
<td>(0.091)</td>
</tr>
<tr>
<td>( \lambda )</td>
<td>0.065</td>
<td>0.065</td>
<td>-0.057</td>
</tr>
<tr>
<td></td>
<td>(0.052)</td>
<td>(0.056)</td>
<td>(0.0536)</td>
</tr>
<tr>
<td>( \theta )</td>
<td>0.807**</td>
<td>0.776**</td>
<td>0.733**</td>
</tr>
<tr>
<td></td>
<td>(0.092)</td>
<td>(0.084)</td>
<td>(0.081)</td>
</tr>
<tr>
<td>Sargan</td>
<td>7.055</td>
<td>8.046</td>
<td>3.766</td>
</tr>
<tr>
<td>(p-value)</td>
<td>(0.070)</td>
<td>(0.090)</td>
<td>(0.288)</td>
</tr>
<tr>
<td>Germany</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( \beta )</td>
<td>0.969**</td>
<td>1.000</td>
<td>0.974**</td>
</tr>
<tr>
<td></td>
<td>(0.109)</td>
<td></td>
<td>(0.122)</td>
</tr>
<tr>
<td>( \lambda )</td>
<td>-0.044</td>
<td>-0.046</td>
<td>0.003</td>
</tr>
<tr>
<td></td>
<td>(0.111)</td>
<td>(0.114)</td>
<td>(0.046)</td>
</tr>
<tr>
<td>( \theta )</td>
<td>0.962**</td>
<td></td>
<td>0.788**</td>
</tr>
<tr>
<td></td>
<td>(0.413)</td>
<td></td>
<td>(0.120)</td>
</tr>
<tr>
<td>Sargan</td>
<td>1.936</td>
<td>1.968</td>
<td>0.859</td>
</tr>
<tr>
<td>(p-value)</td>
<td>(0.586)</td>
<td>(0.742)</td>
<td>(0.835)</td>
</tr>
<tr>
<td>Italy</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( \beta )</td>
<td>1.013**</td>
<td>1.000</td>
<td>1.053**</td>
</tr>
<tr>
<td></td>
<td>(0.058)</td>
<td></td>
<td>(0.069)</td>
</tr>
<tr>
<td>( \lambda )</td>
<td>0.203**</td>
<td>0.202**</td>
<td>-0.027**</td>
</tr>
<tr>
<td></td>
<td>(0.066)</td>
<td>(0.066)</td>
<td>(0.012)</td>
</tr>
<tr>
<td>( \theta )</td>
<td>0.637**</td>
<td>0.640**</td>
<td>-0.029</td>
</tr>
<tr>
<td></td>
<td>(0.042)</td>
<td>(0.046)</td>
<td>(0.023)</td>
</tr>
<tr>
<td>Sargan</td>
<td>7.303</td>
<td>7.408</td>
<td>1.414</td>
</tr>
<tr>
<td>(p-value)</td>
<td>(0.063)</td>
<td>(0.116)</td>
<td>(0.702)</td>
</tr>
</tbody>
</table>

Notes: Newey-West standard errors (robust to serial correlation up to three lags) are in parentheses. ** and * denote significance at the five and ten percent level, respectively. Row Sargan corresponds to the Sargan test of the overidentifying restrictions (p-values are in parentheses). Dummy estimates for Italy are not shown.
An important result from the benchmark estimations is that the widely advocated measure of real marginal costs, real unit labor costs in deviation from sample average, is empirically rather unsuccessful. A possible explanation is that the assumption of a constant desired markup of prices over marginal costs is too restrictive. Indeed, as Coenen and Levin (2004) explain, downward trends in real unit labor costs can be interpreted as increasing desired markup levels. This can be seen from the definition of the actual markup. Because the actual markup is defined as the ratio of price over nominal marginal costs, it is equally defined as the inverse of real marginal costs. For this reason, downward trends in real marginal costs represent, by definition, increasing markup levels. Of course, this need not reflect that the desired markup levels have increased over the time period considered, but it is a distinct possibility. Comparing the results in columns four and six in Table 3, we observe that for France and Germany the empirical results improve when we allow for (deterministically) increasing desired markups over time. Nevertheless, the \( \lambda \) estimates remain insignificant at the five percent significance level. This strongly suggests that the real unit labor costs proxy is inappropriate in the current context.

When technically possible, implied values of the price stickiness parameter \( \theta \) have been calculated from the point estimates of \( \beta \) and \( \lambda \). Unfortunately, only the \( \theta \) estimate for Italy is reliable because \( \lambda \) must be estimated significantly to identify \( \theta \). This is perhaps best explained with an example. Let us consider the insignificant, positive \( \lambda \) estimate for Germany when using real unit labor costs in deviation from trend. Under the restriction \( \beta = 1 \), the point estimate is 0.061 and the corresponding 95% confidence interval ranges from -0.111 to 0.233. Since \( \lambda \) needs to be nonnegative to identify \( \theta \) (see footnote 22), parameter \( \theta \) is not identified for a large part of the 95% confidence interval. Hence, due to the large uncertainty surrounding the positive point estimate of \( \lambda \), the corresponding \( \theta \) estimate for Germany is deemed highly unreliable. Other papers that address this issue are Guay and Pelgrin (2004) and Rumler (2005).

The \( \theta \) estimate for Italy suggests that the proportion of firms that keep prices constant any given quarter is about 64%. Since \( \theta \) estimates imply average price duration values according to the definition \( D = 1/(1-\theta) \), the average price duration in Italy is estimated around 2.8 quarters. This seems low when compared with evidence reported in studies on micro consumer price data where the average price duration ranges from four to five quarters in the euro area (see Dhyne et al., 2005). Nonetheless, Benigno and López-Salido (2002) also find that the degree of price stickiness in Italy is relatively low compared to other euro area countries.\(^{27}\)

\(^{27}\) The analysis of Benigno and López-Salido (2002) is quite different from this analysis; most importantly they assume rational expectations and use a longer, different sample (1970-1998) for all countries considered.
4.3 Open-economy results

The previous sub-section has shown that the NKPC does not fit the data when the real marginal costs of production are approximated by real unit labor costs. To investigate whether this is due to assuming away open-economy factors, I analyze the open-economy extension proposed by Galí and López-Salido (2001) and Balakrishnan and López-Salido (2002).

Table 4 shows the estimates of Equation (4) when \( rmc_t \) is replaced with the real marginal costs measure specified in Equation (8). The results are given, assuming a relatively low elasticity in column two, the benchmark elasticity in column four and a relatively high elasticity in column six. As before, Newey-West standard errors (robust to serial correlation up to three lags) are in parentheses. The instrument sets used are as defined in the previous sub-section. According to the Sargan test results all instrument sets are valid at the five percent significance level. Relevance of the instrument sets is also confirmed; the lowest F-statistic is 8.01 and the corresponding five percent critical value is 3.74.

A clear result emerges from Table 4, namely allowing for a relatively high degree of substitution between labor and imported inputs substantially improves the fit of the NKPCs for France and Germany. Under the assumption that \( \sigma = 1.1 \), the coefficients on the open-economy measures of real marginal costs are positive and significant at the five percent level. Interestingly, this is also reported for the UK by Balakrishnan and López-Salido (2002). For Italy the open-economy results are less encouraging; all \( \lambda_s \) are estimated with the wrong sign.

The implied \( \theta \) values indicate that in France about 70% and in Germany about 75% of the firms keep prices unchanged any given quarter. These values entail that the average price duration in France is about 3.4 quarters and in Germany it is about 4.0 quarters.

As the selected substitution elasticity values are somewhat arbitrary, I investigate how different values of \( \sigma \) affect the slope (\( \lambda \)) estimates; the results are depicted in Figure 5. The plots for France and Germany tell a similar story. When the elasticity of substitution increases beyond the earlier assumed value 1.1, the slope estimate becomes smaller yet highly significant. According to the theory, a lower coefficient on current marginal costs means that firms care relatively more about future cost developments. As the estimates for the subjective discount factor \( \beta \) are largely unaffected, Definition (2) implies that the price rigidity parameter \( \theta \) is estimated higher when a larger elasticity of substitution is assumed. This is an intuitive result. When it is relatively easy to substitute away from an expensive production factor, the need to adjust prices in response to cost developments is mitigated.

From the previous sub-section we know that the higher values of \( \theta \) also imply longer average price duration values. In the case of France the implied average price duration ranges from about 3.4 quarters (\( \sigma = 1.1 \)) to about 6.0
Table 4  Open-economy results  
1991.4-2004.4 (53 observations) 

\[ \pi_t = \beta \bar{F}(\pi_{t-1}) + \lambda \omega_{t-1}, \quad \lambda = (1 - \theta)(1 - \beta \theta)/\theta \]

<table>
<thead>
<tr>
<th></th>
<th>( \sigma = 0.9 )</th>
<th>( \sigma \to 1 )</th>
<th>( \sigma = 1.1 )</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>unrestricted ( \beta = 1 )</td>
<td>unrestricted ( \beta = 1 )</td>
<td>unrestricted ( \beta = 1 )</td>
</tr>
<tr>
<td>France</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( \beta )</td>
<td>0.906** (0.090)</td>
<td>0.916** (0.086)</td>
<td>0.934** (0.067)</td>
</tr>
<tr>
<td>( \lambda )</td>
<td>-0.000 (0.056)</td>
<td>0.120 (0.073)</td>
<td>0.146** (0.042)</td>
</tr>
<tr>
<td>( \theta )</td>
<td>1.000 (0.597)</td>
<td>0.733** (0.081)</td>
<td>0.702** (0.045)</td>
</tr>
<tr>
<td>Sargan</td>
<td>6.534 (0.088)</td>
<td>7.040 (0.071)</td>
<td>5.008</td>
</tr>
<tr>
<td></td>
<td>7.389 (0.117)</td>
<td>7.683 (0.104)</td>
<td>5.570</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>(0.234)</td>
</tr>
<tr>
<td>Germany</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( \beta )</td>
<td>0.972** (0.120)</td>
<td>0.937** (0.108)</td>
<td>0.934** (0.103)</td>
</tr>
<tr>
<td>( \lambda )</td>
<td>0.010 (0.087)</td>
<td>0.071 (0.078)</td>
<td>0.098** (0.042)</td>
</tr>
<tr>
<td>( \theta )</td>
<td>0.915** (0.377)</td>
<td>0.788** (0.120)</td>
<td>0.752** (0.067)</td>
</tr>
<tr>
<td>Sargan</td>
<td>0.877 (0.831)</td>
<td>0.754 (0.860)</td>
<td>1.309</td>
</tr>
<tr>
<td></td>
<td>0.920 (0.922)</td>
<td>1.074 (0.898)</td>
<td>1.668</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>(0.797)</td>
</tr>
<tr>
<td>Italy</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( \beta )</td>
<td>1.014** (0.070)</td>
<td>1.016** (0.072)</td>
<td>1.014** (0.070)</td>
</tr>
<tr>
<td>( \lambda )</td>
<td>-0.041 (0.026)</td>
<td>-0.029 (0.023)</td>
<td>-0.016</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>-0.016</td>
</tr>
<tr>
<td>( \theta )</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Sargan</td>
<td>5.050 (0.168)</td>
<td>5.657 (0.130)</td>
<td>6.508</td>
</tr>
<tr>
<td></td>
<td>5.148 (0.272)</td>
<td>5.785 (0.216)</td>
<td>6.672</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>(0.154)</td>
</tr>
<tr>
<td>Notes: Newey-West standard errors (robust to serial correlation up to three lags) are in parentheses. ** and * denote significance at the five and ten percent level, respectively. Row Sargan corresponds to the Sargan test of the overidentifying restrictions (p-values are in parentheses). Dummy estimates for Italy are not shown.</td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>
Figure 5  Influence of elasticity of substitution on slope ($\lambda$) estimates and corresponding 95% confidence intervals

France

Germany

Italy

Elasticity of substitution
quarters ($\sigma = 1.4$); in the case of Germany it ranges from about 4.0 quarters ($\sigma = 1.1$) to about 4.8 quarters ($\sigma = 1.4$). These implied price durations are in line with the four to five quarters reported in studies on micro consumer price data (see previous sub-section).

For Italy the experiment confirms the poor empirical fit reported in Table 4. Irrespective of the value of the elasticity of substitution, the slope estimates are at odds with the theory. Nonetheless, Figure 5 indicates that the slope estimate increases when open-economy factors are given a greater weight in marginal costs developments.

4.4 The role of lagged inflation

A robust result in the literature is the NKPC’s inability to fully capture the typically large degree of persistence in inflation. This implies that the standard forward-looking NKPC is rejected by the data. Motivated by this finding, theorists have developed modifications of the NKPC that allow for a significant role of lagged inflation in the inflation process. A popular modification is GG’s hybrid Phillips curve. Nevertheless, in the current context the hybrid Phillips curve has no straightforward interpretation. The reason is that the coefficient on lagged inflation is likely to capture variation that should be attributed to (the coefficient on) subjectively expected inflation.

Therefore the role of lagged inflation is evaluated in a different manner. I follow AP in testing the explanatory power of lagged inflation beyond its ability to predict subjectively expected inflation. This can be implemented by testing the significance of that part of the variation in lagged inflation which is orthogonal to expected inflation. The orthogonalized part of lagged inflation can be derived from a regression of lagged inflation on expected inflation, since the calculated residuals represent (by definition) that part of the variation in lagged inflation that is orthogonal to expected inflation. Admittedly, this approach biases results towards rejecting a role for lagged inflation or equivalently, towards not rejecting the null hypothesis that the coefficient on orthogonalized lagged inflation (henceforth $orth(\pi_{t-1})$) equals zero. For this reason I prefer relatively high significance levels (10% and 20%) to lessen the probability of making type II errors.

Table 5 presents the result of this experiment when using the output gap (column two), real unit labor costs in deviation from trend (column three) and the open-economy measure with $\sigma = 1.1$ (column four), respectively, as a proxy for real marginal costs. The instrument sets used are as defined in the previous sub-section with the generated variable $orth(\pi_{t-1})$ superadded. All instrument sets are

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28 The experiment has been performed for all real marginal costs measures. Since the results do not depend on the measure used, only a selection of the results is shown in Table 5.
Table 5  The role of lagged inflation  
1991.4-2004.4 (53 observations)

\[
\hat{\pi}_t = \hat{\beta} F_{[\hat{\pi}_{t+1}]} + \lambda (\sigma) \hat{r}_{mc_{t}} + \gamma \text{ orth} (\hat{\pi}_{t-1})
\]

<table>
<thead>
<tr>
<th>Output gap</th>
<th>Real unit labor costs deviation from trend</th>
<th>Open-economy measure (( \sigma = 1.1 ))</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td><strong>France</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( \hat{\beta} )</td>
<td>0.892#</td>
<td>0.908#</td>
</tr>
<tr>
<td>(0.084)</td>
<td>(0.079)</td>
<td>(0.064)</td>
</tr>
<tr>
<td>( \hat{\lambda} )</td>
<td>0.054</td>
<td>0.100#</td>
</tr>
<tr>
<td>(0.053)</td>
<td>(0.070)</td>
<td>(0.046)</td>
</tr>
<tr>
<td>( \hat{\gamma} )</td>
<td>0.153</td>
<td>0.129</td>
</tr>
<tr>
<td>(0.201)</td>
<td>(0.182)</td>
<td>(0.189)</td>
</tr>
<tr>
<td>Sargan (p-value)</td>
<td>6.407</td>
<td>6.558</td>
</tr>
<tr>
<td>(0.093)</td>
<td>(0.087)</td>
<td>(0.173)</td>
</tr>
<tr>
<td><strong>Germany</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( \hat{\beta} )</td>
<td>0.842#</td>
<td>0.871#</td>
</tr>
<tr>
<td>(0.081)</td>
<td>(0.081)</td>
<td>(0.078)</td>
</tr>
<tr>
<td>( \hat{\lambda} )</td>
<td>-0.016</td>
<td>0.008</td>
</tr>
<tr>
<td>(0.078)</td>
<td>(0.064)</td>
<td>(0.044)</td>
</tr>
<tr>
<td>( \hat{\gamma} )</td>
<td>0.435#</td>
<td>0.427#</td>
</tr>
<tr>
<td>(0.093)</td>
<td>(0.095)</td>
<td>(0.096)</td>
</tr>
<tr>
<td>Sargan (p-value)</td>
<td>7.335</td>
<td>1.117</td>
</tr>
<tr>
<td>(0.062)</td>
<td>(0.773)</td>
<td>(0.504)</td>
</tr>
<tr>
<td><strong>Italy</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( \hat{\beta} )</td>
<td>0.969#</td>
<td>0.966#</td>
</tr>
<tr>
<td>(0.036)</td>
<td>(0.038)</td>
<td>(0.039)</td>
</tr>
<tr>
<td>( \hat{\lambda} )</td>
<td>0.096#</td>
<td>-0.005</td>
</tr>
<tr>
<td>(0.046)</td>
<td>(0.013)</td>
<td>(0.012)</td>
</tr>
<tr>
<td>( \hat{\gamma} )</td>
<td>0.559#</td>
<td>0.630#</td>
</tr>
<tr>
<td>(0.104)</td>
<td>(0.100)</td>
<td>(0.099)</td>
</tr>
<tr>
<td>Sargan (p-value)</td>
<td>7.005</td>
<td>1.622</td>
</tr>
<tr>
<td>(0.072)</td>
<td>(0.654)</td>
<td>(0.564)</td>
</tr>
</tbody>
</table>

Notes: Newey-West standard errors (robust to serial correlation up to three lags) are in parentheses. \# and \# denote significance at the ten and twenty percent level, respectively. Row Sargan corresponds to the Sargan test of the overidentifying restrictions (p-values are in parentheses). Dummy estimates for Italy are not shown.
found to be relevant since the lowest F-statistic (7.31) is substantially larger than the corresponding one percent critical value (3.22). Furthermore, the instrument sets appear also valid as the Sargan test never rejects the null hypothesis of instrument validity at the five percent significance level.

In Table 5 the estimates of the coefficient on orthogonalized lagged inflation, $\gamma$, are primarily important. For Germany and Italy all $\gamma$ estimates are significantly different from zero at conventional significance levels, irrespective of the real marginal costs proxy used. From this I conclude that in Germany and Italy the role of lagged inflation goes beyond predicting subjectively expected inflation. By contrast, for France the estimates of $\gamma$ tell a different story. Irrespective of the marginal costs measure used, all estimates are insignificant at the twenty percent significance level. When comparing the $\gamma$ estimates for France, we observe that using an open-economy measure yields the smallest estimated coefficient on orthogonalized lagged inflation. This finding sustains my preference for real marginal costs measures that explicitly take open-economy factors into account.

5 CONCLUSIONS

This paper has examined the empirical performance of the standard forward-looking new Keynesian Phillips curve in France, Germany and Italy. Instead of imposing rational expectations, direct measures of inflation expectations have been used. This approach has the advantage that it enables an unconditional test of the new Keynesian Phillips curve.

Two key critical issues which arise in the literature have been addressed. The first is which empirical measure of the real marginal costs of production is appropriate in the sense that the parameter estimates are as implied by the theory. For Italy I have found evidence in support of the output gap; for France and Germany however, the output gap did not work well empirically. For these countries I have only obtained plausible estimates when taking into account open-economy factors affecting real marginal costs and subsequently the inflation process. This is an intuitive finding considering the openness of these economies compared to the United States and the euro area.

The second issue addressed in this paper is the role of lagged inflation in the inflation process. For Germany and Italy I have shown that lagged inflation is a significant determinant of current inflation; its role in the inflation process goes beyond predicting subjectively expected inflation. In contrast, for France I have found that lagged inflation does not have explanatory power beyond predicting expected inflation. This suggests that the standard forward-looking new Keynesian Phillips curve effectively captures quarterly inflation dynamics in France.
Since subjective inflation expectations are likely to depend on lagged inflation, it is possible that the subjective expectations new Keynesian Phillips curve is empirically similar to the hybrid Phillips curve of Galí and Gertler (1999). However, for policy recommendations only the new Keynesian Phillips curve can be used. The reason is that the Lucas (1976) critique applies to the hybrid Phillips curve because it imposes that a fixed fraction of the price-setters is backward-looking. The subjective expectations new Keynesian Phillips curve does not impose a role for lagged inflation in the inflation process yet allows for it via the subjective expectations. Nonetheless, my results for Germany and Italy strongly indicate that the explanatory power of lagged inflation goes beyond predicting subjective inflation expectations. This finding cannot be reconciled with the forward-looking new Keynesian Phillips curve and poses new questions for future research.
Appendix

Data Sources

- **Actual quarter-to-quarter inflation** is calculated as the first difference of the natural logarithm of seasonally adjusted CPI. The original quarterly CPI series cover the period 1990.3-2004.4 and are from the Bank for International Settlements (BIS). These CPI series are seasonally adjusted with the Census X12-method.

- **Expected quarter-to-quarter inflation** is constructed for the period 1991.3-2004.4 from seasonally adjusted CPI (see the first bullet) series and inflation forecast series. The forecast series are from Consensus Economics Inc. and are averages of forecasts made by surveyed forecasters. Sub-section 3.1 gives an elaboration of the construction of the expected inflation series.

- **Quarterly output gap series** are derived from Hodrick-Prescott filtering (lambda 1600) the natural logarithm of seasonally adjusted real GDP at basic prices over the period 1991.1-2004.4. The unadjusted quarterly real GDP series are from Eurostat (http://europa.eu.int/comm/eurostat/); these are seasonally adjusted with the X12-method from the US Census bureau.

- **Quarterly real unit labor costs series** are constructed for the period 1991.1-2004.4 by combining the following series: compensation of employees, nominal GDP at basic prices, total employment in number of persons and the number of employees. Compensation of employees consists of wages and salaries, and of employers’ social contributions. These time series are all from Eurostat (http://europa.eu.int/comm/eurostat/). Before putting to use, the real unit labor costs series are seasonally adjusted with the Census X12-method. See sub-section 3.2 for further information.

- **Quarterly series of relative prices of domestic and foreign inputs** are derived for the period 1991.1-2004.4 from the following series: compensation of employees (see previous bullet), the total number of employees (see previous bullet), the Hamburgisches Welt-Wirtschafts-Archiv (HWWA) index of world market prices of raw materials (in euros) and the national exchange rate vis-à-vis the euro which is from Eurostat (http://europa.eu.int/comm/eurostat/). The exchange rate is used to obtain the (relative) prices in national currency. The relative price series are seasonally adjusted (X12) before calculated in deviation from steady-state.

- **Short-term interest rates**, used as instruments in the analysis, are national three-month money market interest rates. These series are obtained from the BIS for the period 1991.1-2004.4.
References


Sondergaard, L. (2003a), Inflation Dynamics in the Traded Sectors of France, Italy and Spain, mimeo, Department of Economics, Georgetown University.


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