Do Professional Forecasters Behave as if They Believed in the New Keynesian Phillips Curve for the Euro Area?

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Abstract

This paper finds that participants in the European Central Bank's Survey of Professional Forecasters have submitted forecasts that are consistent with a (mostly forward-looking) New Keynesian Phillips Curve for the euro area. The estimation results suggest that euro-area inflation forecasts have reacted less to unemployment forecasts after the start of the financial crisis but another cost measure (energy inflation) remains significant. This finding is consistent with a flatter Phillips Curve in the euro area. However, the reasons suggested by the International Monetary Fund for this finding, namely a better anchoring of inflation expectations and increases in structural unemployment do not seem to find support in the survey data. Instead, downward wage rigidities may be playing a prominent role.

Keywords: New Keynesian Phillips curve, inflation, unemployment, panel data, Survey of Professional Forecasters, downward wage rigidities.

JEL classification: E31, J30.

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1. Introduction

In the International Monetary Fund's World Economic Outlook it was recently suggested that inflation rates in advanced economies have become less responsive to output and unemployment during the current economic crisis (IMF, 2013).[2] With reference to the United States, Astrayuda, Ball and Mazumder (2013) labelled this phenomenon the *deflation puzzle*: with unemployment rates as high as those experienced during the Great Recession, the Phillips curve suggests that inflation should have been much lower.[3]

Unemployment rates in the euro area reached historical highs of 12% in April and May 2013 but inflation rates remained at that time relatively close to the European Central Bank's inflation objective: 1.2% in April and 1.4% in May 2013. This may be a sign of a change in the relationship between unemployment and inflation as described by the IMF. More recently, the inflation rate in the euro area has fallen to negative territory, triggering concerns of deflation, but mostly because of the mechanical impact of lower oil prices, with core inflation still above zero while unemployment remains at relatively high rates.

The New Keynesian Phillips Curve (NKPC) is probably the most influential contemporaneous theory on the determination of inflation at business-cycle frequencies.[4] According to this paradigm, forward-looking entrepreneurs set prices as mark-ups over a combination of current and future expected marginal production costs. After aggregating across firms, inflation is a function of the expected future path of aggregate *real* marginal costs:

$$\pi_t - \pi^{ss} = \beta E_t(\pi_{t+1} - \pi^{ss}) + \kappa(mc_t - mc_t^{ss}) \qquad [1]$$

where π_t is the inflation rate at time t, β is the discount factor of the entrepreneur, E_t denotes the rational-expectations operator with information up to period t, π^{ss} is the steady-state inflation rate, mc_t is the aggregate real marginal cost faced by entrepreneurs and mc_t^{ss} stands for the value of real marginal costs in

[2] The countries included in the IMF study are Canada, Switzerland, Germany, Spain, France, Italy, Japan, Netherlands, Norway, Sweden, the United Kingdom and the United States. The IMF attributed this event mainly to "the strengthening of central banks' credibility" leading to more stable inflation expectations. Acedo Montoya and Döhring (2011), in a European Commission Economic Paper, also pointed out that "the combination of stable inflation expectations, sluggish price adjustment and an only moderate impact of the output gap on inflation helps understanding the stability of core inflation despite large and persistent output gaps in the aftermath of the crisis".

[3] Astrayuda I, Ball L, Mazumder S. 2013. Inflation dynamics and the great recession: an update, presented at the seminar Inflation, Unemployment and Monetary Policy organised by the Sveriges Riksbank in February 2013.

[4] For a microfounded derivation of the NKPC, see for example Woodford (2003).

the steady state. Galí and Gertler (1999) popularised the NKPC when they published parameter estimates of a hybrid version of [1] using the labour income share as proxy for real marginal costs, which are unobservable. They found that the NKPC approximated inflation developments in the US reasonably well. The unemployment gap (i.e. the difference between the unemployment rate and the natural rate of unemployment), the output gap and the rate of capacity utilisation are common proxies for real marginal costs in many empirical specifications of [1] (see Linde, 2005, Mankiw, 2001 or Roberts, 2001, among many others).

The parameter κ completes the description of equation [1]. It is the slope of the NKPC and a function of firms' mark-ups and the severity of price rigidities in product markets. Intuitively, the higher the slope the more responsive inflation will be to developments in marginal costs. This parameter is, therefore, of crucial importance to monetary policymakers. When the Phillips curve is very steep monetary expansions, which increase the output gap and decrease the unemployment gap, would lead to more inflationary pressures than similar policies under a flatter curve.

A weaker link between inflation and unemployment, as found by the IMF, does not necessarily mean however that the NKPC is less valid. As the IMF itself but also the ECB (2012) noted, the structural unemployment rate may have increased during the crisis, which implies that unemployment may have increased by more than the unemployment gap. Or it may be that the validity of the unemployment gap as a proxy for real marginal costs has diminished in the recent past because costs have been more influenced by changes in other variables, like energy and commodity prices.[5]

This paper uses data from the ECB's Survey of Professional Forecasters (SPF) to estimate a forward-looking version of the NKPC for the euro area after the start of the financial crisis. It compares the results with estimations for the pre-crisis period. The main objective of the paper is to investigate if the forecasts submitted by professional forecasters to the ECB are consistent with the theoretical implications derived from the NKPC.

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[5] Schmidtt-Grohé and Uribe (2013) affirmed that "since the onset of the great recession in peripheral Europe, nominal hourly wages have not fallen from the high levels they had reached during the boom years in spite of widespread increases in unemployment", suggesting that maybe the unemployment rate is a worse proxy for real marginal costs than before the great recession. Matheson and Stavrev (2013) find that "the importance of import-price inflation has increased" recently for inflation developments in the US.

The ECB's SPF collects expectations of several macroeconomic variables for the euro area submitted by professional forecasters.[6] The cross-sectional variation added by the panel provides a substantial amount of additional information, more variability in the variables of interest, more degrees of freedom for statistical inference and more efficiency of the estimators than the time-series analysis alone.[7]

How could the SPF data be used to estimate the parameters in equation [1]? The ECB publishes expectations of the year-on-year inflation rate one and two years ahead submitted by SPF panellists. These are expectations of headline inflation, as core inflation is not surveyed. The ECB also publishes individual SPF expectations of some proxies for the marginal costs one year ahead.[8] This dataset allows the estimation of the parameters in equation [1] by a transformation: multiplying both sides of equation [1] by the lead operator,[9] taking rational expectations and assuming for a moment that a unit of time is one year:

$$E_{it}(\pi_{t+1} - \pi^{ss}) = \beta E_{it}(\pi_{t+2} - \pi^{ss}) + \kappa E_{it}(mc_{t+1} - mc_{t+1}^{ss}) \qquad [2]$$

$E_{it}\pi_{t+1}$ is the expected year-on-year inflation rate one year ahead $E_{it}\pi_{t+2}$ is the expected year-on-year inflation rate two years ahead and $E_{it}mc_{t+1}$ is the expected real marginal cost one year ahead. Note the subscript i next to the rational-expectations operator. It refers to panellist i in the SPF. To the extent that forecasts of inflation and marginal costs differ among SPF panellists, the cross-sectional information provided by the survey would be valuable for the estimation of the parameters in equation [2].

While a comparison between estimates of equation [2] with pre-crisis and post-crisis SPF data would not directly reveal if the NKPC has changed or not, it would provide information on whether professional forecasters submitted expectations consistent with a change in the NKPC. As the panellists of the ECB's SPF are among the most important financial institutions, research centres, business organisations and labour unions in Europe, their views are informative and, most probably, influential for the determination of macroeconomic outcomes in the euro area.[10]

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[6] Koop and Onorante (2012) estimate Phillips curves for the euro area using expectations from the ECB's SPF at the aggregated level. This paper uses individual forecasts.

[7] See Baltagi (1995) and Gujarati and Porter (2009).

[8] Examples are the unemployment rate and the oil-price inflation rate.

[9] The lead operator, $L^{-1}(\ )$, is defined as $L^{-1}(x_t) = x_{t+1}$ for any variable x.

[10] See http://www.ecb.europa.eu/stats/prices/indic/forecast/html/index.en.html for a partial list of institutions that contributed to the survey.

Note that this paper does not investigate whether SPF panellists *should* use the NKPC to forecast inflation rates in the euro area. The analysis here is positive, not normative. This paper uses SPF data to test the implications of the NKPC theory for the conditional expectation of the inflation rate by professional forecasters in the euro area. In this regard, the approach taken in this paper is the same as in Fendel, Lis and Rülke (FLR, 2011), who tested whether expectations by professional forecasters in G7 countries were in line with the Phillips Curve. The main differences between FLR and this paper are the dataset (forecasts from Consensus Economics in FLR and ECB's SPF forecasts here), the countries analysed (G7 countries in FLR and the euro area here), the inclusion of oil price as an explanatory variable (not included in FLR but included here, see below for details) and the time period (1989-2007 in FLR and 2002-2013 here). The last difference is especially relevant as the sample in FLR does not allow to test whether the parameters in the Phillips curve have changed during the financial crisis that started in 2007.

Alternative approaches to deal with parameter instability are, among others, time-varying structural VAR techniques (see Kirchner, Cimadomo and Hauptmeier (2010) for an application to the effects of fiscal policies in the euro area), dynamic factor models with structural breaks (e.g. Koop and Onorante, *op. cit.*) and Markov-switching structural VAR models (Sims and Zha, 2006). Notwithstanding the unquestionable attractiveness of these approaches, they require the use of certain assumptions upon which results may depend. For instance, time-varing VAR models require the specification of a law of motion for the model parameters, which may be misspecified; dynamic factor averaging techniques require an assumption on the "forgetting factor" (Raftery, Karny and Ettler, 2010) which may affect the results; and Markov-switching models require an assumption on the order of the Markov process, typically that transition probabilities only depend on the current state (Hamilton, 1989).

The paper is organised as follows. Section 2 presents the econometric model to be estimated. Section 3 discusses the data used in the estimation. Section 4 contains the estimation results and Section 5 concludes with an overview of the main results and potential directions for further research.

2. The econometric model

Equation [2] includes three variables that are unobservable: rational expectations of inflation one and two periods ahead and rational expectations of the difference between real marginal costs and their steady-state value. They need to be replaced by proxies in our econometric model.

As pointed out above, the ECB's SPF provides individual inflation forecasts one and two years ahead, which may be used as proxies for the rational expectations of inflation.[11] It also provides forecasts of other variables, unemployment and oil prices, which may serve as proxies for marginal costs. Equation [2] may then be rewritten to substitute the unobservable variables with proxies. Due to the quarterly frequency of the SPF data, the time unit is assumed to be one quarter for the reminder of the paper:

$$E_{it}^{spf}(\pi_{t+4}-\pi^{ss}) = \beta E_{it}^{spf}(\pi_{t+8}-\pi^{ss}) + \kappa E_{it}^{spf}(x_{t+4}-x_{t+4}^{ss}) + \varepsilon_{it} \qquad [3]$$

As in equation [2], expected year-on-year inflation one year (i.e. four quarters) ahead is a function of expected year-on-year inflation two years ahead and the expected real marginal cost one year ahead, with x being a vector of proxies for real marginal costs. The *spf* superscript next to the expectations operator denotes a forecast by a SPF panellist, which may or may not coincide with its rational-expectations counterpart. The error term takes the form:

$$\begin{aligned}\varepsilon_{it} &= E_{it}^{spf}(\pi_{t+4}-\pi^{ss}) - E_{it}(\pi_{t+4}-\pi^{ss}) - \beta[E_{it}^{spf}(\pi_{t+8}-\pi^{ss}) - E_{it}(\pi_{t+8}-\pi^{ss})] - \kappa[E_{it}^{spf}(x_{t+4}-x_{t+4}^{ss}) - \\ &\quad - E_{it}(mc_{t+4}-mc_{t+4}^{ss})]\end{aligned} \qquad [4]$$

The interpretation of this error term is then related to the causes by which a survey forecast may differ from its rational-expectations counterpart. A survey forecast may not coincide with the rational expectation because forecasters could exhibit a form of irrationality, choosing to ignore some pieces of relevant information that are available to them. Unfortunately for this approach, the NKPC is built under the assumption of rational expectations: agents need to be rational to derive equation [1] (Mavroeidis, Plagborg-Møller and Stock, 2014). Hence, our error term cannot be interpreted as deviations from rationality.[12]

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[11] A detailed description of the data used in the empirical exercise is deferred to Section 3.

[12] Harvey and Newbold (2003) pointed out that forecasts errors from the US-SPF have non-zero mean and are not normally distributed. Croushore (2010) argued, however, that the systematic bias disappears when real-time data is considered, and Wang and Lee (2014) recently stated that forecast rationality under asymmetric loss fails to be rejected for most of the rolling periods they analysed.

Survey forecasts, however, may contain measurement errors due to rounding and occasional mistakes made during the completion of the questionnaire. More importantly, the proxies for real marginal costs are noisy, which may lead to potentially large and persistent measurement errors. Consequently, the error term defined by equation [4] is interpreted as a combination of measurement errors.

Crucially, these measurement errors may naturally lead to the presence of unobserved individual heterogeneity in our model: different panellists may have different information on how noisy the approximations to real marginal costs are, giving rise to different inflation forecasts for the same value of the proxies. Because this differential behaviour may persist over time, unobserved individual heterogeneity may appear.

Therefore, the empirical NKPC model to be estimated is:

$$E_{it}^{spf}(\pi_{t+4}-\pi^{ss})=\beta E_{it}^{spf}(\pi_{t+8}-\pi^{ss})+\kappa E_{it}^{spf}(x_{t+4}-x_{t+4}^{ss})+\upsilon_i+\eta_{it} \qquad [5]$$

with ε_{it} = υ_i + η_{it}. The measurement error is thus decomposed into a persistent individual effect, υ_i, and a transitory shock, η_{it}. The latter is assumed to be normal *iid* and includes all omitted variables (e.g. the impact of central bank communication and media on inflation expectations).[13]

Even if different forecasters used the same NKPC model [2] with the same parameters β and κ they would not have necessarily submitted the same expectations to the ECB: different forecasters are likely to have different mapping functions between the expected marginal cost and its proxies. The panel nature of the data allows dealing with this unobserved individual heterogeneity while adding useful information from the cross section of panellists.

3. The data

As the aim is to estimate equation [5] for the euro area as a whole, all variables described in this section are euro-area aggregates. The source of the expected variables in [5] is the ECB's SPF, which is conducted since 1999 Q1. It surveys expectations of inflation, GDP growth, unemployment, policy rates, compensation per employee, oil prices and exchange rates for several forecast horizons. 100 forecasters have participated

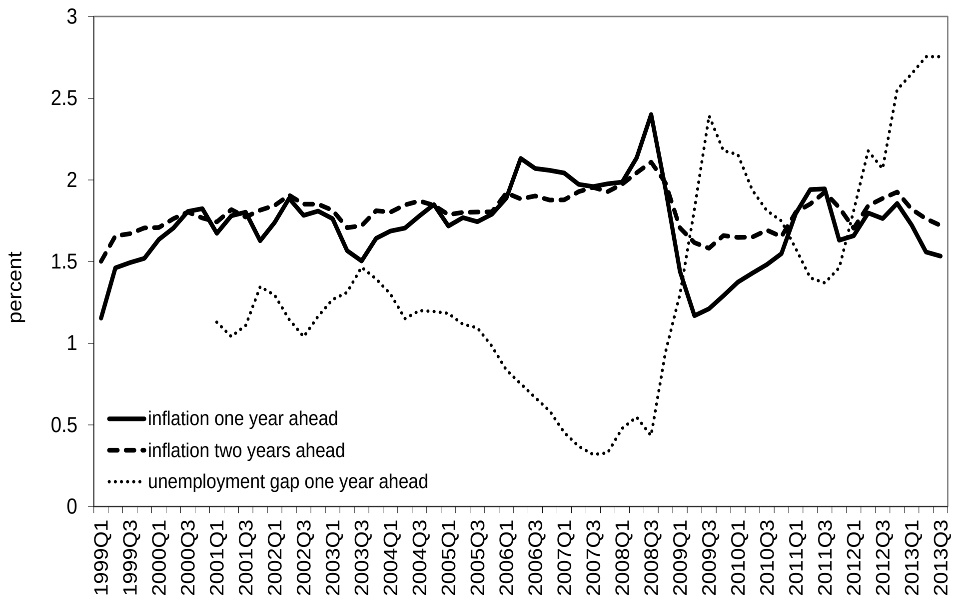
[13] Lamla and Maag (2012) developed a learning model where inflation forecasts are individual-specific by allowing the forecasters to interpret the same media reports differently. However, their empirical results suggest that professional forecasters' disagreement does not depend on media coverage.

at least once in the survey, although the average participation rate is around 60 forecasters per round. The panel is unbalanced, as many forecasters have discontinued their participation in the survey over time and have been replaced by new panellists.[14]

The survey is conducted quarterly, in January, April, July and October, and the questionnaires are sent out to the participants immediately after Eurostat publishes the final estimate of the inflation rate in the euro area for the previous month, typically on the 16th day of the month. The forecasters have around one week to return the questionnaire. At the time of completing the questionnaire, let's say the 2013 Q3 questionnaire, which was filled in in July, the participants knew the inflation rate with a one-month lag (June), the GDP growth rate with a two-quarter lag (2013 Q1) and the unemployment rate with a two-month lag (May).

Focusing on inflation expectations, there are six different inflation forecasts available from the SPF, differing in the forecast horizon. In this paper we use the one-year and two-year ahead inflation forecasts as $E_{it}^{spf}(\pi_{t+4})$ and $E_{it}^{spf}(\pi_{t+8})$ in equation [5] respectively. These forecasts refer to year-on-year inflation rates, as quarter-on-quarter inflation forecasts are not surveyed in the ECB's SPF.[15] The average inflation expectations across participants for these two forecast horizons in each survey round since 1999 Q1 are shown on Figure 1.

Figure 1: Average expectations from the ECB's SPF

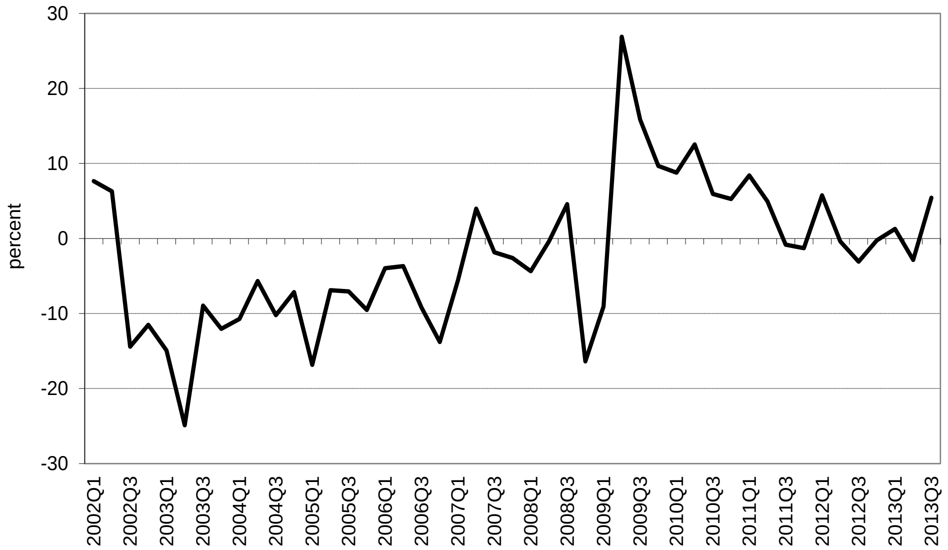


[14] Visit http://www.ecb.europa.eu/stats/prices/indic/forecast/html/index.en.html for a full description of the survey.

[15] In the 2013 Q3 example, these forecasts refer to the year-on-year inflation rate in June 2014 and June 2015 respectively.

The proxies for the expected real marginal cost include the forecasts of the unemployment gap and oil-price inflation. The expected unemployment gap is constructed as the expected unemployment rate one year ahead minus the expected unemployment rate five calendar years ahead, which is assumed to be a proxy for the natural rate of unemployment.[16] These individual forecasts are obtained from the SPF.[17] The time series for the average unemployment gap across forecasters is also shown on Figure 1. As predicted by the NKPC, there seems to be a negative relationship between the expected unemployment gap and the *difference* between the expected inflation rates one and two years ahead.

Figure 2: Average year-on-year expected increase in oil prices from the ECB's SPF



A complementary proxy used for the expected marginal cost is the year-on-year expected increase in oil prices four quarters ahead, due to energy being an important input in the production function (see Berndt and Wood, 1975 and 1979, Griffin and Gregory, 1976, and Kemfert and Welsch, 2000, among many others). Moreover, volatile energy prices have had a significant impact on HICP developments in the last decade. The SPF surveys the expected price of oil (Brent, in dollars) since 2002 Q1.[18] We could have included the expected increase in oil prices in euros because the SPF also surveys the dollar/euro exchange rate. We

[16] Many SPF forecasters report their expectations of the natural rate of unemployment when they are asked about the expected unemployment rate five calendar years ahead (53% according to ECB, 2014).

[17] In the 2013 Q3 example, these forecasts refer to the unemployment rate in May 2014 and the average unemployment rate in 2018 respectively. Expectations of the unemployment rate five calendar years ahead are published quarterly since 2001 Q1.

[18] In the 2013 Q3 example, this forecast refers to the expected average price of oil in 2014 Q2.

decided not to because of reverse-causality concerns, with high expected inflation triggering a monetary policy response that may affect the external value of the euro.[19] The time series of the average expected increase in oil prices four quarters ahead across SPF participants is shown on Figure 2.

As indicated in the previous section, there are measurement errors in our econometric model and instruments are needed to obtain consistent estimates of its parameters. Lags of the regressors will be used when appropriate, i.e. when a test of the over-identifying restrictions suggests that the instruments are valid. Lags of some macroeconomic variables, namely the unemployment gap, labour costs, the inflation rate and the increase in the price of oil are also included as instruments. In particular:

- the *unemployment-gap instrument* in any given quarter is defined as the difference between the unemployment rate in the middle month of the quarter, seasonally adjusted, and the average unemployment forecast five calendar years ahead from the SPF conducted in that quarter. The average unemployment forecast is used instead of the individual forecast to mitigate measurement errors in this variable as much as possible. Forecasters know the value of this instrument with a one-quarter lag.[20]
- the *labour-costs instrument* in any given quarter is the year-on-year percentage change in the quarterly labour-cost index for the euro area (total labour costs, business economy, seasonally adjusted) published by Eurostat. Due to its publication lag, forecasters know the value of this instrument with a two-quarter lag.[21] Note that the SPF also collects expectations of compensation per employee but are not used in this paper because they are forecasts for the next calendar year, not one year ahead.

[19] Regression results with the dollar/euro exchange rate included as an additional regressor confirmed this concern. These results are available from the author upon request.

[20] In 2013 Q3, this instrument is the difference between the unemployment rate in August 2013 and the average five-calendar-years-ahead forecast of the unemployment rate in the 2013 Q3 SPF round. It became part of the information set available to SPF panellists in 2013 Q4.

[21] In 2013 Q3, this instrument is the year-on-year percentage change in the labour cost index in 2013 Q3. It was in the information set available to SPF panellists in 2014 Q1.

- the *inflation-rate instrument* in any given quarter is defined as the year-on-year inflation rate in the last month of the quarter published by Eurostat. Forecasters know the value of this variable with a one-quarter lag.[22]
- the *oil-price-inflation instrument* in any given quarter is defined as the year-on-year percentage change in the average price of oil (Brent, in dollars) over the quarter. It was obtained from the ECB´s Statistical Data Warehouse as the price in euros multiplied by the dollar/euro exchange rate. Forecasters know the value of this instrument with a one-quarter lag.[23]

4. Estimation results

Results with time-series data

We first present estimations obtained with aggregated time-series data (ignoring the existence of the SPF panel with individual data) of the parameters of the NKPC augmented with a lagged-inflation term. This additional term has traditionally been included in estimations of the NKPC to improve its fit (Fuhrer and Moore, 1995, Sbordone, 2006), reflecting some form of adaptive expectations. As the consequences of omitting a relevant regressor (potential inconsistency) are much worse than those of including an irrelevant regressor (potential loss in efficiency), lagged inflation is included in the empirical specification. The resulting specification is the so-called hybrid NKPC:

$$E_t^{spf}(\pi_{t+4} - \pi^{ss}) = \gamma(\pi_t - \pi^{ss}) + \beta E_t^{spf}(\pi_{t+8} - \pi^{ss}) + \kappa E_t^{spf}(x_{t+4} - x_{t+4}^{ss}) + \xi_t \qquad [6]$$

where π_t is the latest available inflation rate before each SPF round is conducted and belongs to the forecasters' information set. Intuitively, equation [6] states that the year-on-year inflation forecast one year (four quarters) ahead is a function of a constant (which combines all time-invariant terms in [6]), the year-on-year inflation forecast two years (eight quarters) ahead, the expected unemployment gap one year ahead, the expected increase in the price of oil one year ahead and the latest realised year-on-year inflation rate published.

[22] In 2013 Q3, this instrument is the year-on-year inflation rate in September 2013. It became part of the information set available to SPF panellists in 2013 Q4.

[23] In 2013 Q3, this instrument is the year-on-year increase in the price of oil (in dollars) in September 2013. It was in the information set available to SPF panellists in 2013 Q4.

The Generalised Method of Moments (GMM) estimator is used due to the potential endogeneity of the regressors, as it replaces endogenous regressors with instruments in the orthogonality conditions. The choice of instruments must take into account that the error term in our model is likely to be autocorrelated because i) the model is misspecified since proxies for the real marginal cost are included; and ii) the dependent variable is a year-on-year expectation which is measured quarterly. The instrument list includes the first lag of the expected inflation rate two years ahead, the second lag of the expected unemployment gap and the first lag of the expected increase in the price of oil. The instrument list also includes two lags of the unemployment-gap instrument, the second lag of the labour-costs instrument, one lag of the inflation-rate instrument and one lag of the oil-price-inflation instrument.[24]

Table 1: Estimated parameters of equation [6]

$$E_t^{spf}(\pi_{t+4}-\pi^{ss})=\gamma(\pi_t-\pi^{ss})+\beta E_t^{spf}(\pi_{t+8}-\pi^{ss})+\kappa_1 E_t^{spf}(u_{t+4}-u_{t+4}^{ss})+\kappa_2 E_t^{spf}(oil_{t+4})+\xi_t$$

	constant	γ	Β	κ1	κ2
OLS. Sample: 2002 Q1 – 2013 Q3					
Point estimates	-1.094***	0.043*	1.565***	-0.072***	0.000
(*HAC standard errors*)	(*0.298*)	(*0.026*)	(*0.174*)	(*0.025*)	(*0.002*)
GMM. Sample: 2002 Q1 – 2013 Q3					
Point estimates	-0.231	0.127***	0.965***	-0.047***	0.002
(*HAC standard errors*)	(*0.478*)	(*0.043*)	(*0.301*)	(*0.016*)	(*0.002*)

Note: *** denotes significance at the 1% level. ** denotes significance at the 5% level. * denotes significance at the 10% level.

Table 1 shows the estimation results for the sample period 2002 Q1 – 2013 Q3 because oil-price forecasts are not available before 2002. Newey-West HAC standard errors are computed because, as mentioned above, the errors are likely to be autocorrelated. While OLS estimates of the coefficients are still consistent under autocorrelation of the residuals, the estimates of the standard errors are biased and inconsistent. The

[24] The Sargan test does not reject the over-identifying restrictions of the model (p value: 0.499). The null hypothesis of no correlation between each instrument and the residuals in the sample was not rejected: the p-values of the tests were 0.209 for the first lag of the expected inflation rate two years ahead, 0.204 for the second lag of expected unemployment gap, 0.675 for the first lag of the expected increase in the price of oil, 0.591 and 0.910 for the two lags of the unemployment-gap instrument, 0.219 for the second lag of the labour-costs instrument, 0.986 for the lag of the inflation-rate instrument and 0.643 for the lag of the oil-price-inflation instrument. The first lag of the expected unemployment gap one year ahead was not included in the list because the null hypothesis of no correlation with the residuals was rejected at the 10% significance level.

latter may be corrected either by Feasible Generalised Least Squares (which under some conditions is efficient) or by estimating HAC standard errors (which produces consistent estimates of the standard errors). OLS estimation results are also shown for comparison.

The estimation results suggest that past inflation is statistically significant but the forward-looking part of the NKPC dominates.[25] The coefficient of the expected unemployment gap has a negative sign, as expected, and is statistically significant.[26] The coefficient of the expected oil-price inflation rate has the expected positive sign but is not significant at the 10% level. These results, however, should be taken with caution because the sample size is very small (T=46).[27]

It could be argued that these estimates are heavily influenced by the events occurred during the first half of the sample, until the start of the financial crisis in 2007, and that the NKPC does not hold thereafter. To verify this claim, GMM regressions across sub-samples could be run but the number of observations would be extremely low and the results too unreliable. It is at this point when the attention is turned to estimates of the NKPC parameters using the panel of individual expectations from the ECB's SPF.

The ECB's SPF individual forecasts may help identifying the parameters of the NKPC to the extent that there is enough variation across forecasts within each cross section (i.e. within each survey round). To verify that not all forecasters submitted the same forecasts to the ECB, Figure 3 shows boxplots of every cross section in the panel for the four forecasts included in the analysis: inflation expectations one and two years ahead, the expected unemployment gap one year ahead and the expected oil-price inflation rate one year

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[25] The null hypothesis $\gamma+\beta=1$ is not rejected (p-value: 0.726).

[26] Mazumder (2011) questioned the fundamental validity of the NKPC empirical model on the grounds that the most commonly used proxy for real marginal costs, the labour income share, yielded positive estimates of the slope, κ, because it was countercyclical before the crisis. In our dataset, his critique remains valid: the correlation of the unemployment-gap instrument defined in Section 3, a countercyclical variable, with the variable π_{t+1}- π_{t+2}, is 0.14 during the sample 2001 Q1–2007 Q3. This notwithstanding, for the short sample after the crisis started, 2007 Q4-2013 Q3, this correlation makes more sense: the correlation of π_{t+1}- π_{t+2} with the unemployment-gap instrument is -0.40. More importantly, when we use aggregate SPF inflation expectations instead of actual inflation, the correlations of the expected unemployment gap with the variable $E_{it}^{spf}(\pi_{t+1})$ - $E_{it}^{spf}(\pi_{t+2})$ are -0.67, -0.77 and -0.61 for the 2001 Q1–2013 Q3, 2001 Q1–2007 Q3 and 2007 Q4–2013 Q3 samples respectively.

[27] Similar results were obtained by Brissimis and Maginas (2008) with aggregated survey data for the US. They claim that past inflation may be found to be statistically significant when final inflation figures are used by the econometrician instead of real-time data. In our sample period, however, the revisions to euro-area inflation figures are very small.

ahead. Indeed, there is significant variation across forecasts within every cross-section, with possibly a few exceptions only. [28] Therefore, the panel of forecasts may add valuable information for the estimation of the NKPC parameters.[29]

Results with panel data

The aim in this subsection is to obtain estimates of the parameters in equation [5] with panel data for two different sub-samples: the pre-crisis period, from 2002 Q1 to 2007 Q3, and the crisis period, from 2007 Q4 to 2013 Q3.[30] As before, we expand the econometric model to explore the statistical relevance of past inflation rates:

$$E_{it}^{spf}(\pi_{t+4}-\pi^{ss})=\gamma(\pi_t-\pi^{ss})+\beta E_{it}^{spf}(\pi_{t+8}-\pi^{ss})+\kappa E_{it}^{spf}(x_{t+4}-x_{t+1}^{ss})+\upsilon_i+\eta_{it} \qquad [7]$$

The properties of the unobservable individual component, υ_i, are crucial to the estimation strategy. If the contemporaneous correlations between the regressors and the individual components and between the regressors and the disturbance were both zero, we could estimate equation [7] on pooled data, with all SPF forecasts in each sub-sample treated as if they belonged to a single cross section.

This assumption, however, seems too strong given the interpretation of the error term [4] as a measurement error: let's assume that SPF panellists believe that the NKPC is the right model. In order to forecast inflation, they would like to know the expected real marginal costs faced by firms. Unfortunately, they do not observe this variable but two proxies (the expected unemployment gap and the expected increase in the price of oil). Every survey round, each panellist should then make an unobserved adjustment to these proxies

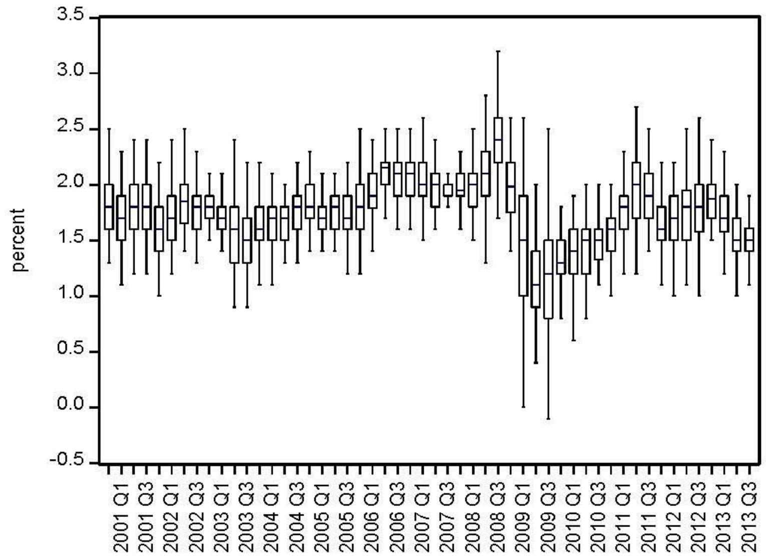
[28] Each box on Figure 3 represents the inter-quartile range (IQR) for each cross section. The line inside each box denotes the median observation from each cross section. The lines above each box represent the range of observations in each cross section between the first quartile and the first quartile minus 1.5 times the IQR (sometimes known as the upper *whiskers*). The lines below each box represent the range of observations between the third quartile and the third quartile plus 1.5 times the IQR (the lower *whiskers*). Note that a few forecasts fall outside the box and the whiskers.

[29] The best examples of possibly too low variation across forecasts within a cross section are inflation expectations one year ahead in 2007 Q3, and inflation expectations two years ahead in 2007 Q3 and Q4.

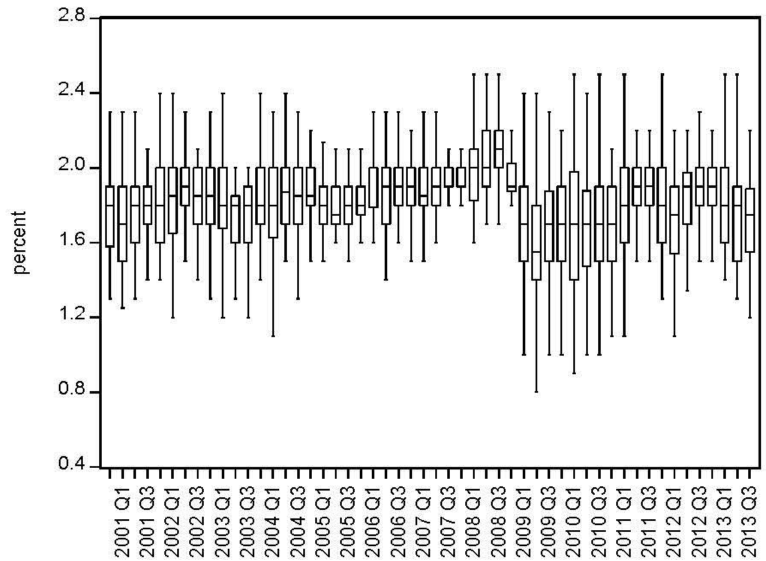
[30] This partition is motivated on the fact that the negative effects from the US housing crisis started to spread out to the world financial markets in August 2007 (see New York Times. 2011. The bank run we knew so little about. 2 April, available at http://www.nytimes.com/2011 /04/03/business/03gret.html?_r=0). As the 2007 Q3 SPF takes place in July, the first survey in "crisis" mode was 2007 Q4. Moreover, many measures of macroeconomic uncertainty computed with data from the ECB's SPF start to pick up in the second half of 2007 (see López-Pérez V. 2014. Measures of macroeconomic uncertainty for the ECB's survey of professional forecasters. Available at http://hdl.handle.net/10317/4248).

Figure 3: Cross-sectional variation of ECB's SPF forecasts

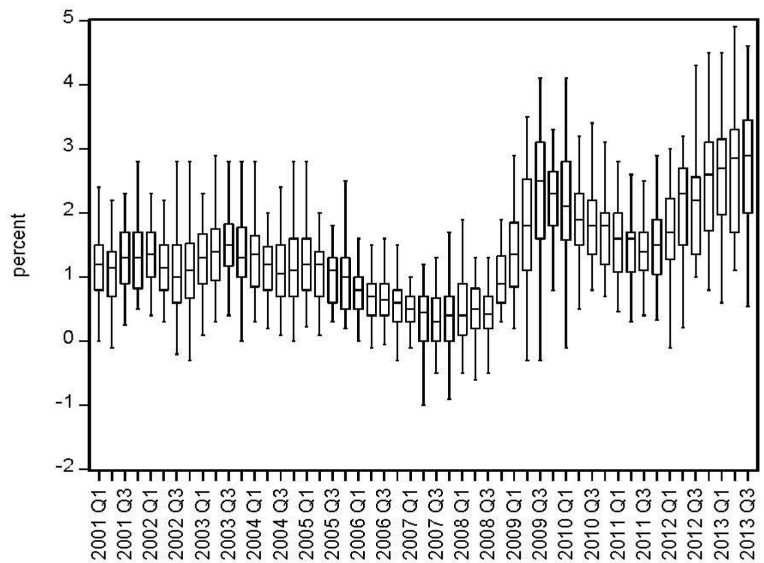
a. Expected inflation one year ahead



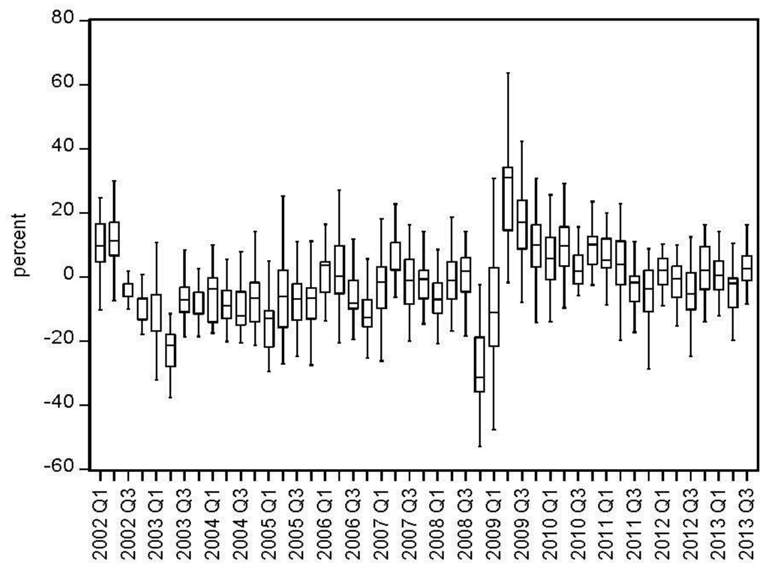
b. Expected inflation two years ahead



c. Expected unemployment gap one year ahead



d. Expected oil-price inflation one year ahead



to obtain her best guess of the real marginal cost. In the likely event that these unobserved adjustments systematically differ among forecasters, individual unobservable effects may be part of the error term.[31]

The question is then whether these individual effects are correlated with the regressors in equation [7]. In principle they could, because forecasters that believe in a larger gap between expected real marginal costs one year ahead and its proxies may forecast higher inflation rates one year ahead. And if real marginal costs are persistent they may *also* forecast higher inflation rates two years ahead. In this scenario of correlation between the individual effects and the regressors, we would need to rely on a "fixed-effects" estimator.

Information on whether the correlation between the individual component and the regressors is quantitatively important may be obtained: Arellano (2003) points out that, in the presence of measurement errors and unobserved heterogeneity, the OLS estimator exhibits two biases. The first is the usual measurement error bias, which increases in absolute value with the variance of the measurement error. The second bias comes from the covariance between the individual component and the regressors.

In the context of model [7], if we found evidence that the unobserved individual heterogeneity co-moves significantly with the regressors, the case for fixed effects would become stronger. Following Arellano (2003) again, equation [7] is estimated in levels by OLS, where the estimated parameters will be affected by the two biases described above. Then the model is re-estimated in deviations from individual averages by OLS, where the estimated parameters will include the measurement-error bias only. Finally, the model may be re-estimated in first differences by OLS: the estimated parameters are likely to be even more biased in the presence of measurement error under the plausible assumption that the persistence in real marginal costs is high compared to the persistence in the measurement errors (Wooldridge, 2011).

Table 2 shows the OLS estimates of the three different specifications of equation [7] described in the previous paragraph. The top panel contains the results for the first sub-sample (2002 Q1 – 2007 Q3) and the bottom panel those for the second sub-sample (2007 Q4 – 2013 Q3). In our model, the measurement-error

[31] Different information sets across forecasters may explain the discrepancy.

bias should be negative for γ, β, and the coefficient of expected oil-price inflation, κ_2, This bias should be positive for the coefficient accompanying expected unemployment, κ_1.[32]

Table 2: Estimated parameters of equation [7]

$$E_{it}^{spf}(\pi_{t+4}-\pi^{ss})=\gamma(\pi_t-\pi^{ss})+\beta E_{it}^{spf}(\pi_{t+8}-\pi^{ss})+\kappa_1 E_t^{spf}(u_{t+4}-u_{t+4}^{ss})+\kappa_2 E_t^{spf}(oil_{t+4})+\upsilon_i+\eta_{it}$$

	constant	γ	β	κ_1	κ_2
Sub-sample 2002 Q1 – 2007 Q3					
OLS levels	1.066*** (0.164)	0.018 (0.030)	0.472*** (0.064)	-0.129*** (0.028)	0.005*** (0.001)
OLS deviations	1.248*** (0.205)	0.009 (0.028)	0.393*** (0.090)	-0.153*** (0.020)	0.004*** (0.001)
OLS first differences	-	0.064** (0.027)	0.074 (0.089)	-0.064* (0.038)	0.005*** (0.001)
Sub-sample 2007 Q4 – 2013 Q3					
OLS levels	0.170 (0.137)	0.239*** (0.028)	0.566*** (0.064)	-0.000 (0.022)	0.006*** (0.002)
OLS deviations	0.228 (0.143)	0.225*** (0.029)	0.558*** (0.058)	-0.011 (0.022)	0.006*** (0.002)
OLS first differences	-	0.155*** (0.027)	0.427*** (0.091)	0.004 (0.031)	0.004** (0.002)

Note: HAC standard errors in parentheses. *** denotes significance at the 1% level. ** denotes significance at the 5% level. * denotes significance at the 10% level.

There seems to be strong indications of measurement-error bias: when we compare the point estimates on the “deviations” rows (which include the measurement-error bias) with those on the “first-differences” rows (which exacerbate the measurement-error bias), the difference is relatively large. This finding supports the case for the use of instruments in our estimation.

The second bias, the unobserved-heterogeneity bias, should affect neither γ nor κ_2: neither lagged inflation nor the expected price of oil ought to co-move much with the average gap estimated by each forecaster between the unobservable real marginal cost and its proxies. This bias, however, could push β upwards because inflation forecasts two years ahead may be positively correlated with expected real marginal costs one year ahead if marginal costs are persistent. Therefore, inflation forecasts two years ahead are likely to capture part of the effect from the systematic measurement error on the dependent variable. For analogous

[32] This measurement-error bias is also called “attenuation” bias (Wooldridge, 2011), as the OLS estimates of the parameters of the regressors measured with error are biased towards zero.

reasons, κ_1 may be biased downwards, as the correlation between the unemployment gap and real marginal costs is expected to be negative.

The estimation results do not find strong evidence of correlation between the unobservable individual effect and the regressors. When the point estimates on the "levels" rows (which include the measurement-error bias and the heterogeneity bias) are compared with those on the "deviations" rows (which include the measurement-error bias only), the differences are minor. The heterogeneity biases of β and κ_1 are small, especially in the second sub-sample, and always statistically insignificant.[33] These findings do not strongly support the case for the "fixed effects" estimator.[34]

When the individual unobserved heterogeneity is not correlated with the regressors, the "random effects" estimator is appropriate. In this case, the individual effects are considered random variables extracted from a distribution which is independent from the distribution of the regressors. Table 3 shows the values of the GMM "random effects" estimators of the parameters of equation [7] for the two sub-samples and the full sample. Period SUR estimated robust standard errors are obtained to account for the potential autocorrelation of the residuals within a cross-sectional unit (e.g. due to the overlapping forecast horizons).

The results are consistent with the findings in the literature for the pre-crisis period, with a negative effect on expected inflation from the expected unemployment gap and a positive effect from expected oil-price inflation. Moreover, the backward-looking component of inflation turns out to be insignificant, suggesting that SPF participants provided inflation forecasts that are consistent with a purely forward-looking Phillips curve for this sub-sample.

For the period after the start of the financial crisis, the estimations of the parameters of the "random effects" model vary somehow with respect to the "pre-crisis" period. Although the forward-looking part of the Phillips curve is still prominent, the backward-looking part is now statistically significant.[35] More importantly, the expected unemployment gap is no longer significant. This finding supports the claim by the

[33] In the model estimated by OLS in orthogonal deviations for the first sub-sample (row 2 on Table 3), the F-test of the null hypothesis "Ho: β=0.472 and κ_1=-0.129" has a p-value of 0.402.

[34] The Hausman test did not reject the null hypothesis "Ho: the difference between the random-effects and the fixed-effects estimates is small" for the full simple and the two sub-samples. Under the null, the random-effects estimator is preferred to the fixed-effects estimator because it is more efficient.

[35] The null hypothesis Ho: $\gamma + \beta = 1$ cannot be rejected (p-value of the F statistic: 0.669).

IMF about the Phillips curve being now flatter than before the crisis.[36] It is also consistent with the results by Fendel, Lis and Rülke (2011) who estimated the Phillips curve to be flatter during economic downturns in G7 countries.

Table 3: Estimated parameters of equation [7]:

$$E_{it}^{spf}(\pi_{t+4}-\pi^{ss})=\gamma(\pi_t-\pi^{ss})+\beta E_{it}^{spf}(\pi_{t+8}-\pi^{ss})+\kappa_1 E_t^{spf}(u_{t+4}-u_{t+4}^{ss})+\kappa_2 E_t^{spf}(oil_{t+4})+\upsilon_i+\eta_{it}$$

	constant	γ	β	κ1	κ2
Sub-sample 2002 Q1 – 2007 Q3					
GMM[37] "Random effects"	0.013 (0.307)	0.105 (0.092)	0.961*** (0.139)	-0.109*** (0.034)	0.011*** (0.003)
Sub-sample 2007 Q4 – 2013 Q3					
GMM[38] "Random effects"	-0.163 (0.199)	0.233*** (0.036)	0.720*** (0.128)	0.026 (0.026)	0.010*** (0.003)
Full sample 2002 Q1 – 2013 Q3					
GMM[39] "Random effects"	0.028 (0.163)	0.177*** (0.024)	0.796*** (0.093)	-0.056*** (0.016)	0.009*** (0.002)

Note: Period SUR robust standard errors in parentheses. *** denotes significance at the 1% level. ** denotes significance at the 5% level. * denotes significance at the 10% level.

Is this bad news for the NKPC? Not necessarily because the second proxy for expected real marginal costs, the expected increase in the price of oil, is statistically significant, has the correct sign and the hypothesis

[36] An econometrician that does not take into account the structural break in the relationship between inflation and unemployment would still find a negative slope of the Phillips curve, although a bit flatter than in the first sub-sample.

[37] Number of observations: 447. The list of instruments includes the first lag of the expected inflation rate two years ahead, the second lag of the inflation instrument, the labour-costs instrument and the oil-price inflation instrument, and the first and second lags of the expected unemployment gap, the expected increase in the price of oil and the unemployment-gap instrument. The Sargan test does not reject the over-identifying restrictions of the model (p value: 0.581).

[38] Number of observations: 415. The list of instruments includes the first lag of the expected inflation rate two years ahead and the expected increase in the price of oil, the second lag of the labour-costs instrument, and the first and second lags of the expected unemployment gap, the unemployment-gap instrument, the inflation-rate instrument and the oil-price-inflation instrument. The Sargan test does not reject the over-identifying restrictions of the model (p value: 0.328).

[39] Number of observations: 978. This number is higher than the sum of the numbers of observations used in each sub-sample mainly because the estimation for the second sub-sample does not use lagged observations from the first sub-sample as instruments, which results in the loss of some observations. The results are the same if lagged instruments from the first sub-sample are allowed in the estimation for the second sub-sample. The list of instruments includes the first lag of the expected inflation rate two years ahead, the expected increase in the price of oil and the inflation-rate instrument, the first and second lags of the expected unemployment gap, the unemployment-gap instrument and the oil-price inflation instrument, and the second lag of the labour-costs instrument. The Sargan test does not reject the over-identifying restrictions of the model (p value: 0.493).

that its magnitude remains unchanged is not rejected at conventional significance levels.[40] That is, during the financial crisis, the SPF panellists still provided forecasts that are consistent with a “hybrid” but strongly forward-looking NKPC. Their forecast, however, could be interpreted as if they considered oil-price inflation rates a better proxy than the unemployment gap for changes in real marginal costs. This result is consistent with the findings by Paradiso and Rao (2012) for the United States and Australia. They estimated a different version of the Phillips curve, without the forward-looking component, and found that the coefficient of the oil price in the Phillips curve has been increasing over the last few decades. They concluded that “the oil price is likely to play a significant role in future inflation rates”.

One of the key assumptions of the GMM random-effects estimator is that the instruments are strictly exogenous. This assumption may seem too strong in this context, where measurement errors are present. If the measurement errors are correlated with the regressors, as in the classical error-in-variables model, the use of lagged values of the regressors would make the random-effects estimator inconsistent. It could be argued that the results of the Sargan tests of the overidentifying restrictions suggest that the instruments are valid but, in order to dissipate any potential concern in this regard, equation [7] is also estimated in orthogonal forward deviations.

The orthogonal-forward-deviations model subtracts from each variable in equation [7] the mean of all its future values. Therefore, this estimator does not require strict exogeneity for consistency, just sequential exogeneity of the regressors, which is a weaker assumption. Under sequential exogeneity, lagged values of the regressors are valid instruments. As a robustness check of the results shown in Table 3, equation [7] is estimated by GMM in orthogonal forward deviations. Table 4 shows the estimation results.

The results are qualitatively the same as before: (i) the expected unemployment gap is significant in the first sub-sample but not in the second; (ii) the expected increase in the price of oil is significant in both periods; and (iii) the NKPC becomes partially backward-looking in the second sub-sample. Therefore, the results shown in the paper are robust to the sequential-exogeneity assumption.

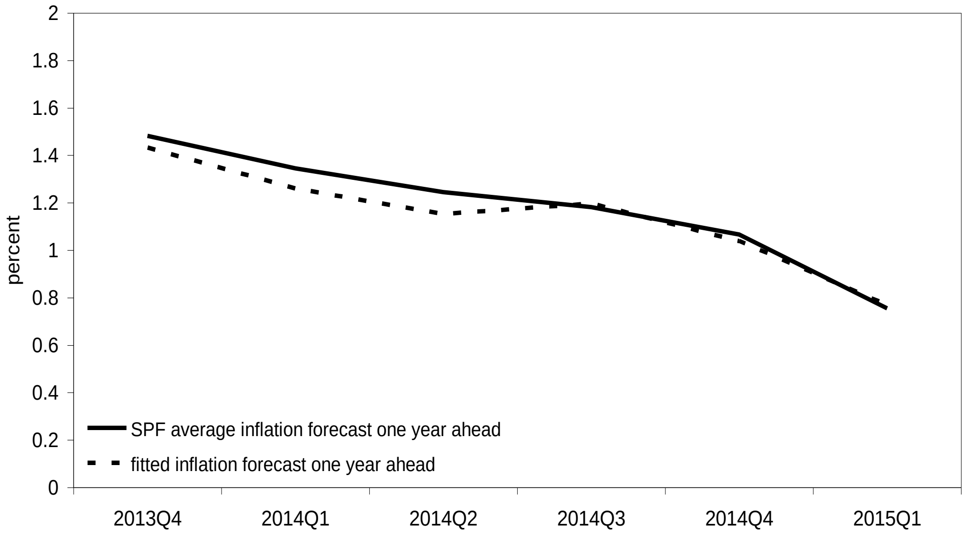
[40] The null hypothesis Ho: $\kappa_2 = 0.008$ (the magnitude of the short-run effect of oil prices on inflation has not changed), has a p-value of 0.762. The null hypothesis Ho: $\kappa_2 /(1-\gamma)= 0.008$ (the magnitude of the long-run effect of oil prices on inflation has not changed), has a p-value of 0.622.

Table 4: Estimated parameters of equation [7] in orthogonal deviations:

	constant	γ	β	κ_1	κ_2
Sub-sample 2002 Q1 – 2007 Q3					
GMM[41] "Orthogonal deviations"	-	0.097 (0.063)	1.128*** (0.175)	-0.177*** (0.036)	0.005** (0.002)
Sub-sample 2007 Q4 – 2013 Q3					
GMM[42] "Orthogonal deviations"	-	0.236*** (0.047)	0.811*** (0.212)	0.019 (0.032)	0.011*** (0.003)

Note: Period SUR robust standard errors in parentheses. *** denotes significance at the 1% level. ** denotes significance at the 5% level. * denotes significance at the 10% level.

Figure 4: Out-of-sample comparison between fitted and SPF forecasts



As the sample used in the estimation ends in 2013 Q3, an out-of-sample comparison between the SPF average inflation forecasts one year ahead since 2013 Q4 and the inflation forecast one year ahead predicted by the estimated NKPC may help ascertain the goodness of fit of model. With that aim, the random-effects model estimated with post 2007 Q4 data is fed with the SPF average inflation forecasts two years ahead, the

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[41] Number of observations: 508. The list of instruments includes the first lag of the expected inflation rate two years ahead, the expected unemployment gap, the expected increase in the price of oil and the oil-price inflation instrument, the second lag of the inflation instrument and the labour-costs instrument, and the first and second lags of the unemployment-gap instrument. The Sargan test does not reject the over-identifying restrictions of the model (p value: 0.785).

[42] Number of observations: 387. The list of instruments includes the first lag of the expected inflation rate two years ahead and the expected increase in the price of oil, the second lag of the labour-costs instrument, and the first and second lags of the expected unemployment gap, the oil-price inflation instrument, the unemployment-gap instrument. The Sargan test does not reject the over-identifying restrictions of the model (p value: 0.394).

average oil-price inflation forecasts one year ahead and the lagged inflation rate for the period 2013 Q4 to 2015 Q1.[43] Figure 4 displays the comparison between the SPF forecasts and the model forecasts.

The distance between the SPF average inflation forecast and the forecast predicted by the model is small. The average distance is 0.04 percentage points and the maximum distance is 0.09 percentage points. These numbers are not economically significant and suggest that the model does a reasonable job explaining the conditional expectation of SPF inflation forecasts one year ahead. In fact, when equation [7] is estimated with data from 2007 Q4 to 2015 Q1, the results confirm that the relationship stays reasonably stable (see Table 5).

Table 5: Estimated parameters of equation [7] with data up to 2015 Q1

$$E_{it}^{spf}(\pi_{t+4}-\pi^{ss})=\gamma(\pi_t-\pi^{ss})+\beta E_{it}^{spf}(\pi_{t+8}-\pi^{ss})+\kappa_1 E_t^{spf}(u_{t+4}-u_{t+4}^{ss})+\kappa_2 E_t^{spf}(oil_{t+4})+\upsilon_i+\eta_{it}$$

	constant	γ	β	κ1	κ2
Sub-sample 2007 Q4 – 2015 Q1					
GMM[44] "Random effects"	-0.175 (0.158)	0.202*** (0.034)	0.767*** (0.098)	0.016 (0.022)	0.006*** (0.002)

Note: See the notes of Table 3.

As an additional robustness check, and with the aim of minimising the effects of the unbalanced nature of the panel on the results, the same GMM estimations whose results appear in Table 3 are conducted excluding the panellists that did not provided inflation, unemployment and oil-price forecasts in less than half of all survey rounds. The results are reported in Table 6 and are almost identical to the results obtained with all the forecasters, with the unemployment gap losing relevance as a proxy of marginal costs after the start of the crisis.[45]

[43] The value of statistically insignificant estimated parameters is set to zero.

[44] Number of observations: 526. The list of instruments for this estimation includes the first lag of the expected inflation rate two years ahead and the expected increase in the price of oil, the first and second lags of the expected unemployment gap, the second lag of the inflation instrument and the oil-price inflation instrument, the second and third lags of the unemployment-gap instrument and the third lag of the labour-costs instrument. The Sargan test does not reject the over-identifying restrictions of the model (p value: 0.422).

[45] The p-values associated to the F-statistic used to test the null hypothesis "the parameter estimates reported in Table 6 are different from those reported in Table 3" are 0.986 and 0.962 for the first and second sub-samples respectively.

Table 6: Estimated parameters of equation [7] with data from forecasters with at least 50% participation

$$E_{it}^{spf}(\pi_{t+4}-\pi^{ss})=\gamma(\pi_t-\pi^{ss})+\beta E_{it}^{spf}(\pi_{t+8}-\pi^{ss})+\kappa_1 E_t^{spf}(u_{t+4}-u_{t+4}^{ss})+\kappa_2 E_t^{spf}(oil_{t+4})+\upsilon_i+\eta_{it}$$

	constant	γ	β	κ_1	κ_2
Sub-sample 2002 Q1 – 2007 Q3					
GMM[46] "Random effects"	0.047 (0.323)	0.144 (0.098)	0.906*** (0.150)	-0.107*** (0.038)	0.012*** (0.003)
Sub-sample 2007 Q4 – 2013 Q3					
GMM[47] "Random effects"	-0.048 (0.208)	0.218*** (0.036)	0.689*** (0.134)	0.016 (0.026)	0.010*** (0.003)

Note: See the notes of Table 3.

Why may the unemployment gap become a worse proxy for real marginal costs during the crisis? The answer is probably related to the impact of downward wage rigidities on inflation developments in many euro-area countries: when unemployment is relatively low, wages increase more than inflation pushing real marginal costs upwards. Firms have then incentives to increase prices. This behaviour gives rise to a negative relationship between unemployment and inflation in "good times" (first sub-sample).

In "bad times", when the unemployment rate is well above the natural rate of unemployment, wages should decelerate its pace of increase and even fall in nominal terms. Downward nominal and real wage rigidities, however, make declines in nominal and real wages less likely. In this scenario, real marginal costs do not decrease as much as they would in the absence of wage rigidities and the negative relationship between unemployment and inflation could disappear (second sub-sample).

Is there evidence in the ECB's SPF dataset of the existence of downward wage rigidities in the euro area? Figure 5 reproduces the average expected unemployment gap one year ahead shown on Figure 1 together with the average expected rate of increase in nominal and real compensation per employee for the next calendar year from the SPF.[48] From 2005 to 2008, the shrinking expected unemployment gap coincided with

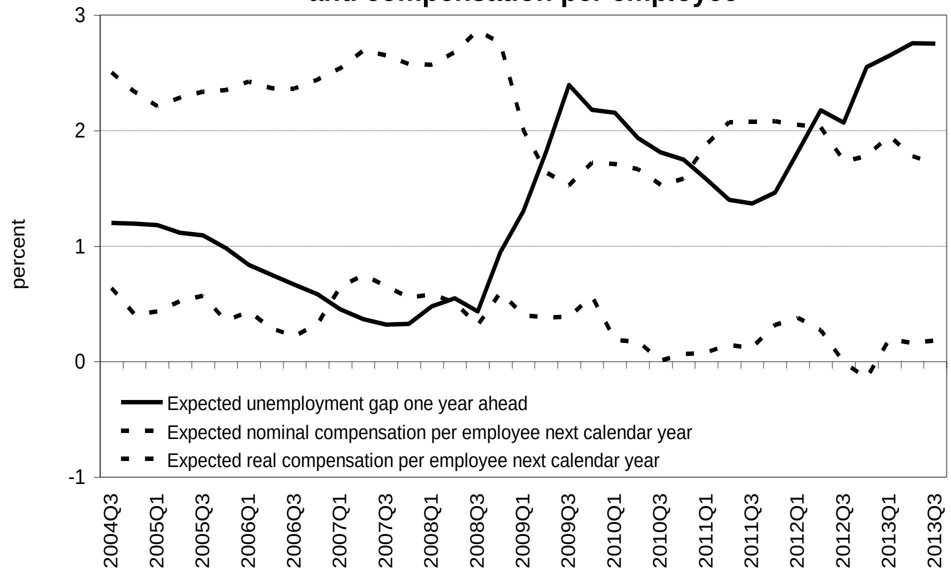
[46] Number of observations: 398. See footnote 35.

[47] Number of observations: 367. See footnote 36.

[48] The expected rate of increase in nominal compensation per employee is obtained from the ECB's SPF. As pointed out above, the ECB does not survey forecasts of compensation per employee one year ahead, just for calendar years. The ECB did not survey forecasts of compensation per employee before 2004 Q3. The expected increase in real compensation per employee is computed

faster rises in expected compensation per employee. In 2009, expected nominal compensation per employee decelerated with the increase in the expected unemployment gap but stayed in a range between 1.5% and 2%, which is consistent with the inflation objective of the ECB. Interestingly, the rate of growth of expected *real* compensation per employee, which has stayed slightly above zero since 2010 has only fallen to negative territory in one period and even then the decline was very small (2012 Q4: -0.14%).

Figure 5: Expected unemployment gap and compensation per employee



Looking at individual SPF data, the histogram of all expectations of real compensation per employee since 2004 Q3 seems to be consistent with SPF panellists taking into account the existence of downward real-wage rigidities (see Figure 6). The histogram is clearly asymmetric: it looks like some of the observations that should have been located in negative territory (black bars) under the assumption of symmetry have been moved to the [0, 0.2) interval.[49] Dickens *et al.* (2007) used this assumption of symmetry to suggest a measure of the relevance of real-wage rigidities from the distribution of wage changes.[50] The same measure

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by subtracting the expected inflation rate in the next calendar year, also available from the SPF, from the expected increase in nominal compensation per employee.

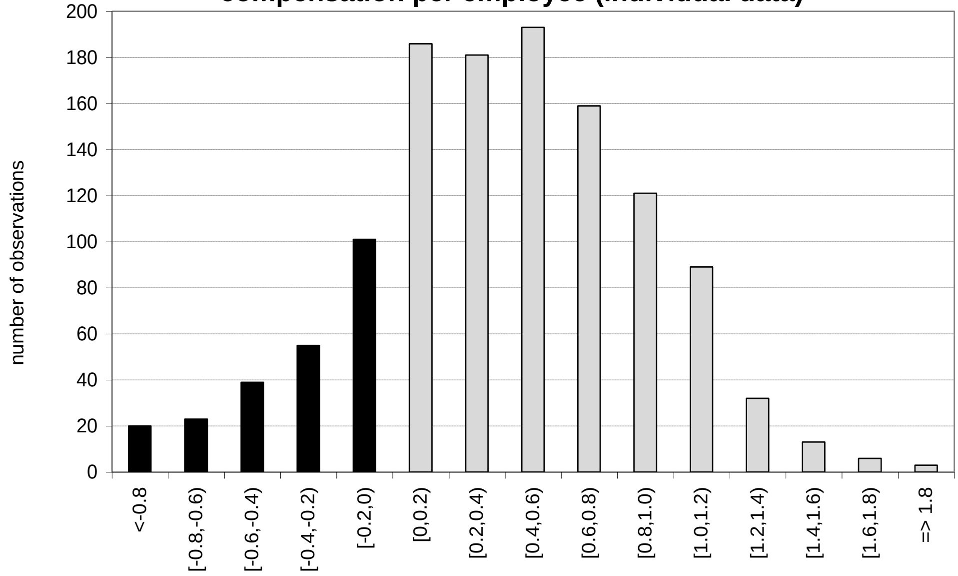
[49] For evidence of downward wage rigidities in the euro area, see the results of the euro-area Wage Dynamics Network: http://www.ecb.europa.eu/home/html/researcher_wdn.en.html.

[50] The real-wage-rigidity measure proposed by Dickens et al. (2007) is the number of workers with real wage freezes divided by the number potentially affected. In our context:

$$r = \frac{2(u-l)}{u}$$

can be applied here to the distribution of expectations of compensation per employee and takes the value 0.20. Put differently, 20% forecasts of real compensation per employee below zero are estimated to be affected by the downward real-wage-rigidity constraint.[51]

Figure 6: Expected rate of growth of real compensation per employee (individual data)



Overall, it seems that SPF panellists have been providing forecasts of inflation, unemployment and oil prices that are consistent with a predominantly forward-looking NKPC for the euro area and the existence of downward real-wage rigidities.[52] We cannot tell if the NKPC is dead or alive with this dataset but SPF participants behave as if they believed in it.

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where *r* is the real-wage-rigidity index, *u* is the fraction of expectations of real compensation per employee above twice the median of the distribution shown on Figure 5, and *l* is the fraction of forecasts below zero. More recent findings obtained by Daly and Hobijn (2014) imply that *r* may underestimate the true degree of rigidity because downward wage rigidities could push downwards not only the left tail but also the right tail of the distribution of wage changes.

[51] The average real-wage rigidity measure reported by Dickens *et al.* (2007) is 0.26, with the following countries included in the average: Ireland, Denmark, France, Belgium, UK, Switzerland, Austria, Germany, Italy, Netherlands, Finland, Norway, Greece, Sweden, US and Portugal.

[52] Downward nominal wage rigidities could also potentially explain a weaker link between inflation and unemployment (Reitz and Slopek, 2014, Daly and Hobijn, 2014). In the SPF dataset used in this paper there are very few instances where a forecaster submitted an expected growth rate of compensation per employee lower than zero (3 out of 1245 forecasts for the horizon "current calendar year" and 3 out of 1241 for the "next calendar year"). While this could be an indication of downward nominal wage rigidities, there is not a clustering of expectations around zero to support this hypothesis: in particular, only 11 out of 1245 forecasts expected a growth rate of compensation per employee of less than 0.5% for the "current calendar year" (15 out of 1241 for the "next calendar year") .

Where does this behaviour come from? Is it that forecasters maybe use a version of the NKPC model to compute their expectations? It does not seem to be the case: only around 5% of the SPF panellists that participated in a special questionnaire conducted by the ECB in September 2008 reported the use of Dynamic-Stochastic General-Equilibrium (DSGE) models.[53] This compares to 70% of professional forecasters reporting the use of traditional, backward-looking macroeconomic models and 65% of them declaring the use of single-equation time series models.

The NKPC-like behaviour then could come from the judgmental adjustments that professional forecasters make to the output of their models. The participants in the SPF special questionnaire indicated that a substantial component of their short and medium-term inflation forecasts is judgement. The results of this paper suggest that maybe they have the NKPC in mind when they adjust their forecasts. Or maybe not, but their behaviour seems to be consistent with it.

5. Conclusion

This paper has tried to help answering the question: do professional forecasters behave as if they believed in the NKPC for the euro area? With that aim, it presents parameter estimates of the New Keynesian Phillips Curve model for the euro area with panel data from the ECB's SPF. Panel data helps dealing with unobserved individual heterogeneity. It also mitigates the small-sample problems that arise when using time series that suffered from a structural break with the start of the financial crisis.

The main finding of the paper is that professional forecasters in the euro area submitted inflation, unemployment and oil-price forecasts that are consistent with a mostly forward-looking Phillips curve for the euro area. While expectations of oil prices and unemployment have been important determinants of inflation forecasts before the financial crisis, the statistical impact of unemployment on inflation expectations seems to have diminished drastically during the crisis.

This result is consistent with the claim made by the IMF that the Phillips curve is flatter now and that unemployment matters less to explain inflation developments. The IMF argued that better-anchored inflation

[53] See ECB (2009), "Results of a Special Questionnaire for Participants in the ECB Survey of Professional Forecasters", available at http://www.ecb.europa.eu/stats/prices/indic/forecast/shared/files/quest_summary.pdf?8063c5fb1c8002823e72f92c1ecbcd98.

expectations and increases in structural unemployment may be behind this finding. But this paper shows that, according to the forecasts submitted by SPF panellists:

(i) the estimated impact of oil prices on inflation remains as important as before the crisis. If the Phillips curve had become flatter due to better-anchored inflation expectations, the contributions from oil-price expectations to inflation expectations should have also been more muted.

(ii) the subdued estimated response of inflation to unemployment during the crisis appears even after controlling from increases in the expected rate of unemployment five years ahead (a proxy for structural unemployment).

A more plausible explanation for the apparently broken relationship between inflation and unemployment during the crisis may be based on the existence of downward wage rigidities, which prevent wages from falling as much as one might expect when the unemployment rate is very high. These rigidities obscure the link between unemployment and real marginal costs, reducing the empirical validity of the former as a proxy for the latter.

The origin of the smaller slope of the Phillips Curve matters for policy recommendations. The IMF argued that a flatter curve caused by better-anchored inflation expectations would prevent inflation from rising rapidly if unemployment falls. This paper, however, has found that the lack of responsiveness of inflation to unemployment may come from the existence of asymmetric rigidities in the labour market instead. Therefore, increases in unemployment may not be affecting inflation much because of the binding wage-rigidity constraint, but decreases in unemployment might move the economy away from the binding constraint and bring the slope of the Phillips curve back to where it was before the crisis.

Further research may be directed to investigate the relevance of downward real wage rigidities in euro-area labour markets during the crisis, and to what extent this type of rigidities may be responsible for the relatively muted response of real marginal costs to unemployment. As the euro-area labour market remains fragmented in national labour markets, estimations of NKPCs at the national level may also be useful to help solving the *deflation puzzle*.

Additionally, the analysis presented in this paper may be replicated with data from the Survey of Professional Forecasters maintained by the Federal Reserve Bank of Philadelphia in the United States. The

ECB's SPF dataset was preferred for this exercise because long-term unemployment forecasts are available since 2001, while they are only available since October 2008 in the US SPF. This constraint would not allow the comparison between estimates of the NKPC before and after the start of the financial crisis. Having said this, there is enough data from the US SPF to at least estimate a version of the NKPC for the post-crisis period.

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