EUROPEAN CENTRAL BANK
EUROSYSTEM

Working Paper Series

Víctor López-Pérez

Do professional forecasters behave as if they believed in the new Keynesian Phillips Curve for the euro area?

No 1763 / March 2015

Acknowledgements
The views expressed in this paper belong to the author and are not necessarily shared by the European Central Bank. I thank Oreste Tristani, Thomas Westermann and an anonymous referee for their helpful comments. The author is grateful for the support received from the European Central Bank, where the author was employed until October 2014.

Víctor López-Pérez
Universidad Politécnica de Cartagena; e-mail: victor.lopez@upct.es

Postal address	60640 Frankfurt am Main, Germany
Telephone	+49 69 1344 0
Internet	www.ecb.europa.eu

ISSN	1725-2806 (online)
ISBN	978-92-899-1576-2
DOI	10.2866/282501
EU catalogue number	QB-AR-15-003-EN-N

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.

Non-technical summary

The New Keynesian Phillips Curve (NKPC) is probably the most influential contemporaneous theory on the determination of inflation at the business-cycle frequency. According to this paradigm, inflation depends on indicators of current and future costs by firms, like labour costs. As labour costs have traditionally fluctuated with unemployment rates, an empirical relationship between inflation and unemployment emerged.

However, the International Monetary Fund recently suggested that inflation rates have become less responsive to unemployment during the current economic crisis. This phenomenon has been labelled the deflation puzzle: with unemployment rates as high as those experienced during the Great Recession, inflation could have been even lower. Unemployment rates in the euro area are close to historical highs now, between 11% and 12%, but inflation rates stay positive.

This paper investigates whether the NKPC remains valid during the current economic and financial crisis. To that end, panel data from the ECB's Survey of Professional Forecasters (SPF) is used. This survey asks highly-skilled macroeconomic forecasters throughout the European Union about their expectations of euro area inflation, GDP growth, unemployment, policy rates, compensation per employee, oil prices and the dollar/euro exchange rate for different forecast horizons. With this dataset, I estimated a forecast-based version of the NKPC for the euro area with post-crisis data and compared the results with estimations for the pre-crisis period.

The estimation results suggest that expected unemployment has become less important to explain inflation forecasts by SPF participants after 2007. This finding seems to support the IMF hypothesis. However, a second cost variable, the price of oil, replaces the role previously played by unemployment as an indicator of firms' costs. That is, during the financial crisis, the SPF panellists still provided forecasts that are consistent with a NKPC but as if oil-price inflation had become a better proxy for firms' costs than the unemployment rate.

Why may unemployment become a worse indicator of firms' costs during the crisis? The answer is probably related to the impact of downward wage rigidities: when the unemployment rate is relatively high, downward wage rigidities make declines in wages less likely because workers refuse to cut nominal wages or even accept wage increases below the expected inflation rate. In this scenario, the relationship between unemployment and inflation is likely to disappear. Interestingly, expectations of compensation per employee from the ECB's SPF seem to be consistent with professional forecasters assuming the existence of downward wage rigidities in the euro area.

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).[3] 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.

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, however, the inflation rate in the euro area has fallen to 0.3% in September 2014, but stays positive while unemployment remains very high (11.5% in September 2014).

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. Inflation would then be a function of the expected future path of *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 real marginal cost faced by entrepreneurs and mc_t^{ss} stands for the value of real marginal costs in the steady state.[5]

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,[6] would lead to more inflationary pressures than similar policies under a flatter curve.

[3] 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".

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

[5] Galí and Gertler (1999) popularised the NKPC when they published parameters 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.

[6] The unemployment gap (i.e. the difference between the unemployment rate from the non-accelerating-inflation rate of unemployment, NAIRU), the output gap and the rate of capacity utilisation are common

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.[7]

This paper estimates the NKPC for the euro area with post-crisis data and compares the results with estimations for the pre-crisis period. Its main objective is to investigate if the NKPC remains valid during the current economic and financial crisis. Unfortunately, from an econometric standpoint, the relatively short sample since the start of the crisis makes estimations of NKPC with post-crisis time-series data unreliable. I address this problem by using a panel of data from the ECB's Survey of Professional Forecasters (SPF), which collects expectations of several macroeconomic variables for the euro area submitted by professional forecasters.[8]

How could the SPF help estimating 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. It also publishes individual SPF expectations of some proxies for the marginal costs one year ahead.[9] This dataset allows the estimation of the parameters in equation [1] by a transformation: multiplying both sides of equation [1] by the lead operator,[10] 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

proxies for real marginal costs in many empirical specifications of [1] (see Linde, 2005, Mankiw, 2001 or Roberts, 2001, among many others).

[7] Schmidtt-Grohé and Uribe (2012) 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.

[8] Another approach for dealing with short time series is Dynamic Model Averaging, employed by Koop and Onorante (2012) to estimate Phillips curves for the euro area. They also use expectations from the ECB's SPF but at the aggregated level, not at the level of the individual forecasters.

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

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

a change in the NKPC or not. 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.

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, tyme-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:

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

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

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, 2013). Hence, our error term cannot be interpreted as deviations from rationality.

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:

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

with $\varepsilon_{it} = \upsilon_i + \eta_{it}$. The measurement error is thus decomposed into a persistent individual effect, υ_i, and a transitory shock, η_{it}, which is assumed to be *iid*. 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 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.[12]

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 in the previous month (June), the GDP growth rate from two quarters ago (2013 Q1) and the unemployment rate from two months ago (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]. These forecasts refer to year-on-year inflation rates, as quarter-on-quarter inflation forecasts are not surveyed in the ECB's SPF.[13] The average inflation expectations across participants for these two forecast horizons in each survey round since 2000 Q4 are shown on Figure 1.[14] Table 1 displays some summary statistics of the individual SPF forecasts used throughout this paper.

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, both from the SPF.[15] The time series for the average unemployment gap across forecasters is shown on Figure 2.

A complementary proxy used for the expected marginal cost is the year-on-year expected increase in oil prices four quarters ahead, due to the significant impact of volatile energy prices on HICP developments in the last decade. The SPF surveys the expected price of oil (Brent, in dollars) since 2002 Q1.[16] We could have included the expected increase in oil prices in euros because the SPF also surveys the dollar/euro exchange rate. We 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.[17] The time series of the average expected increase in oil prices four quarters ahead across SPF participants is shown on Figure 3.

As indicated in the previous section, there are measurement errors in our econometric model and instruments are needed to estimate its parameters. Lags of the regressors will

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

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

[14] The sample starts in 2000 Q4 to allow for a two-year "training period" for the SPF panellists as in Boero, Smith and Wallis (2014).

[15] 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.

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

[17] 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.

be used when appropriate. 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 and the average unemployment forecast five calendar years ahead from the SPF conducted in that quarter. Forecasters know the value of this variable by the time the following SPF is conducted.[18]
- the *labour-costs instrument* in any given quarter is the year-on-year percentage change in the quarterly labour cost index published by Eurostat. Due to its publication lag, forecasters do not know the value of this variable by the time the following SPF is conducted, but it will be available for the round after that.[19] 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.
- 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 when the following SPF round is conducted.[20]
- 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 last month of the quarter. It was obtained from the ECB´s Statistical Data Warehouse. Forecasters know the value of this variable when the following SPF round is conducted.[21]

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. 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-

[18] 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 will be part of the information set available to SPF panellists in 2013 Q4.
[19] In 2013 Q3, this instrument is the year-on-year percentage change in the labour cost index in 2013 Q3. It will be part of the information set available to SPF panellists in 2014 Q1.
[20] In 2013 Q3, this instrument is the year-on-year inflation rate in September 2013. It will be part of the information set available to SPF panellists in 2013 Q4.
[21] In 2013 Q3, this instrument is the year-on-year increase in the price of oil (in dollars) in September 2013. It will be part of the information set available to SPF panellists in 2013 Q4.

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.

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, and one lag of the inflation-rate instrument and the oil-price-inflation instrument.[22] Table 2 shows the estimation results for the sample period 2002 Q1 – 2013 Q3 because oil-price forecasts are not available before 2002. 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.[23] The coefficient of the expected unemployment gap has a negative sign, as expected, and is statistically significant.[24] The coefficient of the expected oil-price inflation 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).[25]

[22] The Sargan test does not reject the over-identifying restrictions of the model (p value: 0.503). Moreover, the null hypothesis of no correlation between each instrument and the residuals was not rejected: the p-values of the tests were 0.208 for the first lag of the expected inflation rate two years ahead, 0.201 for the second lag of expected unemployment gap, 0.662 for the first lag of the expected increase in the price of oil, 0.592 and 0.910 for the two lags of the unemployment-gap instrument, 0.219 for the second lag of the labour-costs instrument, 0.978 for the lag of the inflation-rate instrument and 0.631 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.

[23] The null hypothesis $\gamma+\beta=1$ is not rejected (p-value: 0.709).

[24] 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.

[25] 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.

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 4 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 ahead. Indeed, there is significant variation across forecasts within every cross section, with possibly a few exceptions only.[26] Therefore, the panel of forecasts may add valuable information for the estimation of the NKPC parameters.

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.[27] 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 there were no individual effects and the parameters in equation [7] are constant across sub-samples, 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 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.[28]

[26] 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

[27] 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). 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 (López-Pérez, 2014).

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

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, where the estimated parameters are likely to be even more biased in the presence of measurement error.

Table 3 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 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.

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 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.[29] These findings do not strongly support the case for the "fixed effects" estimator.

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 4 shows the values of the GMM "random effects" estimators of the parameters of equation [7] for the two sub-samples and the full sample. 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.[30] More importantly, the expected unemployment gap is no longer significant. This finding supports the claim by the IMF about the Phillips curve being now flatter than before the crisis.[31]

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 that its magnitude remains unchanged is not rejected at conventional significance levels.[32] 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.

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 relationship between unemployment and inflation in "good times" (first sub-sample).

[29] 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.

[30] The null hypothesis Ho: $\gamma + \beta = 1$ cannot be rejected (p-value: 0.551).

[31] Note that the estimation results with the full sample are, as expected, a combination of the results from the two sub-samples. Interestingly, 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 smaller than in the first sub-sample.

[32] The null hypothesis Ho: $\kappa_2 = 0.007$ (the magnitude of the short-run effect of oil prices on inflation has not changed), has a p-value of 0.908. The null hypothesis Ho: $\kappa_2/(1-\gamma)= 0.007$ (the magnitude of the long-run effect of oil prices on inflation has not changed), has a p-value of 0.774.

In "bad times", when the unemployment rate is well above the NAIRU, 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 2 together with the average expected rate of increase in nominal and real compensation per employee for the next calendar year from the SPF.[33] From 2005 to 2008, the shrinking expected unemployment gap coincided with 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%).

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 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 (red bars) under the assumption of symmetry have been moved to the [0, 0.2) interval.[34] Dickens *et al.* (2007) suggested an index to measure the relevance of real-wage rigidities from the distribution of wage changes.[35] The same index 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.[36]

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. We cannot

[33] 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 by subtracting the expected inflation rate in the next calendar year, also available from the SPF.

[34] 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.

[35] 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:

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

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 6, and l is the fraction of forecasts below zero.

[36] The average real 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.

tell if the NKPC is dead or alive with this dataset but SPF participants behave as if they believed in it.

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 (ECB, 2009). 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 to deal with unobserved individual heterogeneity. It also mitigates the small-sample problems that arise from the use of time series that suffered from a structural break with the start of the financial crisis in 2007.

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 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*.

References

Acedo Montoya, L., and B. Döhring (2011), "The Improbable Renaissance of the Phillips Curve: the Crisis and Euro Area Inflation Dynamics", European Economy Economic Papers 446.

Arellano, M. (2003), *Panel Data Econometrics,* Oxford University Press.

Astrayuda, I., L. Ball and S. Mazumder (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.

Boero, G., J. Smith, and K.F. Wallis (2014), "The Measurement and Characteristics of Professional Forecasters' Uncertainty", forthcoming in the *Journal of Applied Econometrics*.

Brissimis, S.N., and N.S. Magginas (2008), "Inflation Forecasts and the New Keynesian Phillips Curve", *International Journal of Central Banking* 4, pp. 1-22.

Dickens, W.T., L. Goette, E.L. Groshen, S. Holden, J. Messina, M.E. Schweitzer, J. Turunen and M.E. Ward (2007), "How Wages Change: Micro Evidence from the International Wage Flexibility Project", Journal of Economic Perspectives 21, pp. 195-214.

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.

ECB (2012), "Euro Area Labour Markets and the Crisis", ECB Monthly Bulletin, October, pp. 69-80.

Fuhrer, J., and G. Moore (1995), "Inflation Persistence", *Quarterly Journal of Economics* 110, pp. 127-159.

Galí, J., and M. Gertler (1999), "Inflation Dynamics: a Structural Econometric Analysis", *Journal of Monetary Economics* 44, pp. 195-222.

Hamilton, J.D. (1989), "A New Approach to the Economic Analysis of Nonstationary Time Series and the Business Cycle", *Econometrica* 57, pp. 357-384.

IMF (2013), "The Dog that Didn't Bark: Has Inflation Been Muzzled or Was it Just Sleeping?", World Economic Outlook, April, Chapter 3.

Kirchner, M., J. Cimadomo and S. Hauptmeier (2010), "Transmission of Government Spending Shocks in the Euro Area: Time Variation and Driving Forces", ECB Working Paper No. 1219.

Koop, G., and L. Onorante (2012), "Estimating Phillips Curves in Turbulent Times Using the ECB´s Survey of Professional Forecasters", ECB Working Paper 1422.

López-Pérez, V. (2014), "Measures of Macroeconomic Uncertainty for the ECB's Survey of Professional Forecasters" *mimeo*.

Matheson, T., and E. Stavrev (2013), "The Great Recession and the Inflation Puzzle", IMF Working Paper 13/124.

Mavroeidis, S., M. Plagborg-Møller and J. Stock (2013), "Empirical Evidence on Inflation Expectations in the New Keynesian Phillips Curve", forthcoming in the *Journal of Economic Literature*.

Mazumder, S. (2011), "The Empirical Validity of the New Keynesian Phillips Curve Using Survey Forecasts of Inflation", *Economic Modelling* 28, pp. 2439-2450.

New York Times (2011), "The Bank Run We Knew so Little about", 2 April, available in http://www.nytimes.com/2011/04/03/business/03gret.html?_r=0

Raftery, A., M. Karny and P. Ettler (2010), "Online Prediction under Model Uncertainty via Dynamic Model Averaging: Application to a Cold Rolling Mill", *Technometrics* 52, pp. 52-66.

Sbordone, A. (2006), "US Wage and Price Dynamics: a Limited Information Approach", *International Journal of Central Banking* 2, pp. 155-191.

Schmitt-Grohé, S., and M. Uribe (2012), "The Case for Temporary Inflation in the Eurozone", presented at the seminar *Inflation, Unemployment and Monetary Policy* organised by the Sveriges Riksbank in February 2013.

Sims, C.A., and T. Zha (2006), "Were There Regime Switches in U.S. Monetary Policy", *American Economic Review* 96, pp. 54-81.

Woodford, M. (2003), *Interest and Prices: Foundations of a Theory of Monetary Policy*, Princeton University Press.

Tables and Figures

Table 1: Summary statistics of individual SPF forecasts (sample: 2000 Q4 – 2013 Q3)

	Expected inflation one year ahead	Expected inflation two years ahead	Expected unemployment gap one year ahead	Expected oil-price inflation rate one year ahead
Observations	2637	2333	1888	2248
Mean	1.76	1.82	1.32	-1.90
Median	1.80	1.80	1.20	-1.97
Standard deviation	0.38	0.28	0.93	13.56
5th percentile	1.10	1.40	0.00	-22.50
95th percentile	2.30	2.20	3.00	20.01
IQ range	0.49	0.30	1.10	16.25

Table 2: 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 (*HAC standard errors*)	-1.094*** (*0.298*)	0.043* (*0.026*)	1.565*** (*0.174*)	-0.072*** (*0.025*)	0.000 (*0.002*)
GMM. Sample: 2002 Q1 – 2013 Q3					
Point estimates (*HAC standard errors*)	-0.245 (*0.485*)	0.127*** (*0.044*)	0.973*** (*0.305*)	-0.046*** (*0.016*)	0.002 (*0.002*)

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

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					
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.

Table 4: 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"	0.098 (0.308)	0.004 (0.085)	1.027*** (0.140)	-0.129*** (0.037)	0.007* (0.004)
Sub-sample 2007 Q4 – 2013 Q3					
GMM[38] "Random effects"	-0.037 (0.204)	0.164*** (0.032)	0.770*** (0.204)	-0.005 (0.026)	0.007** (0.003)
Full sample 2002 Q1 – 2013 Q3					
GMM[39] "Random effects"	-0.090 (0.144)	0.186*** (0.026)	0.763*** (0.083)	-0.069*** (0.019)	0.010*** (0.002)

Note: Cross-section SUR standard errors in parentheses. *** denotes significance at the 1% level. ** denotes significance at the 5% level. * denotes significance at the 10% level.

[37] Number of observations: 441. The list of instruments for this estimation includes the first and second lag of the expected inflation rate two years ahead, the expected unemployment gap, the expected increase in the price of oil, the unemployment-gap instrument, the inflation-rate instrument and the oil-price-inflation instrument. It was also used as instrument the second lag of the labour-costs instrument. The Sargan test does not reject the over-identifying restrictions of the model (p value: 0.265).

[38] Number of observations: 334. The list of instruments for this estimation includes the first lag of the expected inflation rate two years ahead, the third and fourth lag of the expected increase in the price of oil, the first and second lag of the expected unemployment gap, the unemployment-gap instrument, the inflation-rate instrument and the oil-price-inflation instrument and the second and third lag of the labour-costs instrument. The Sargan test does not reject the over-identifying restrictions of the model (p value: 0.550). The second lag of the expected inflation two years ahead and the first and second lag of the expected increase in the price of oil are not on the instrument list because they led to the rejection of the over-identifying restrictions.

[39] Number of observations: 963. 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 for this estimation includes the first and second lag of the expected inflation rate two years ahead, the expected unemployment gap, the unemployment-gap instrument, the inflation-rate instrument and the oil-price-inflation instrument. The first lag of the expected increase in the price of oil and the second lag of the labour-costs instrument were also used as instruments. The Sargan test does not reject the over-identifying restrictions of the model (p value: 0.306). The second lag of the expected increase in the price of oil and the third lag of the labour-costs instrument are not on the instrument list because they led to the rejection of the over-identifying restrictions.

Figure 1: Average inflation expectations from the SPF

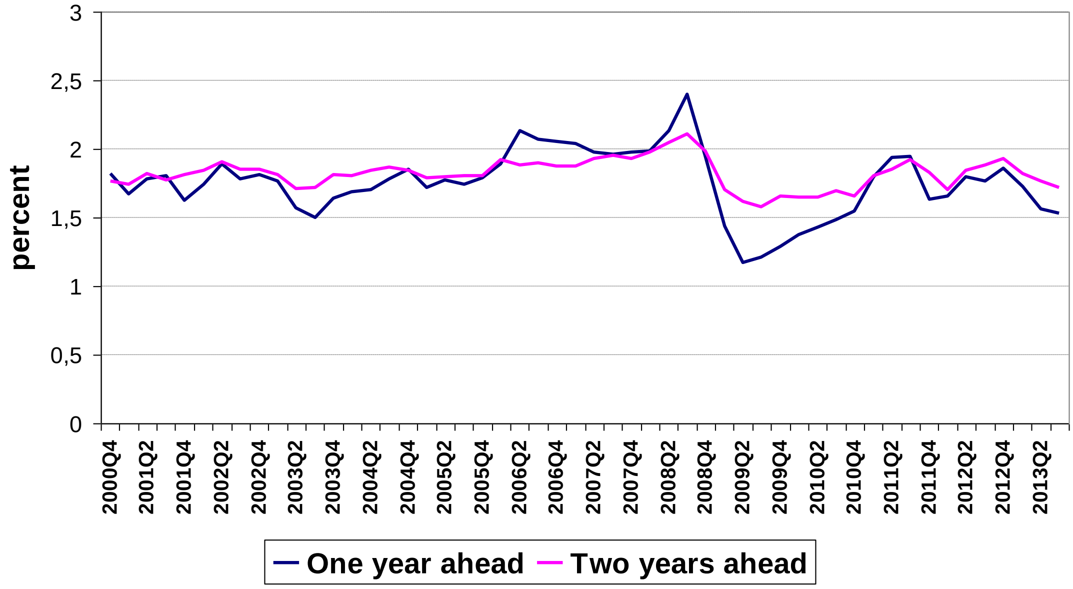


Figure 2: Average expected unemployment gap from the SPF

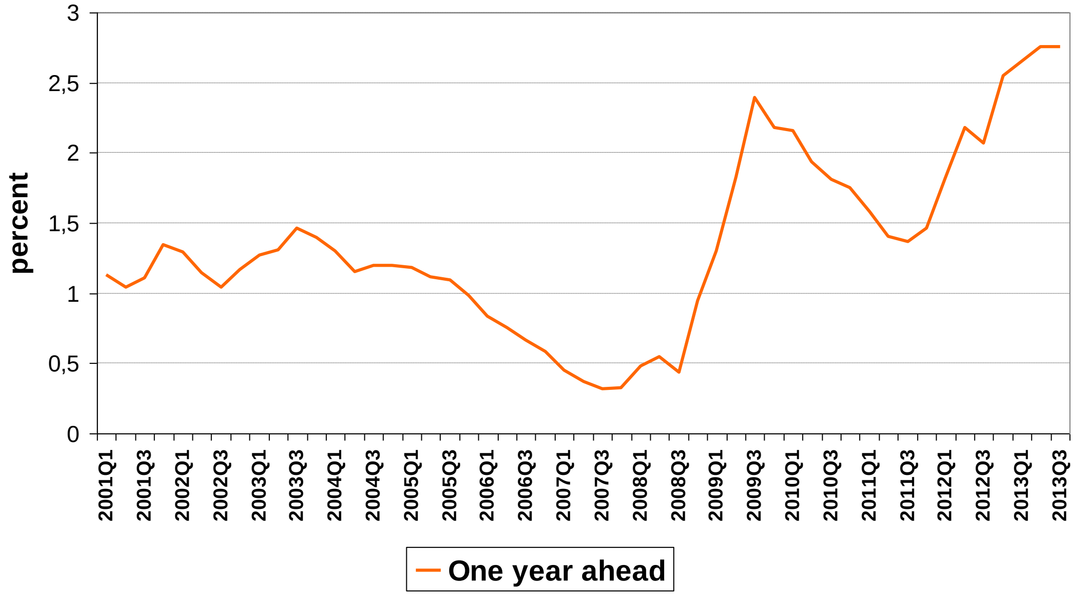


Figure 3: Average year-on-year expected increase in oil prices from the SPF

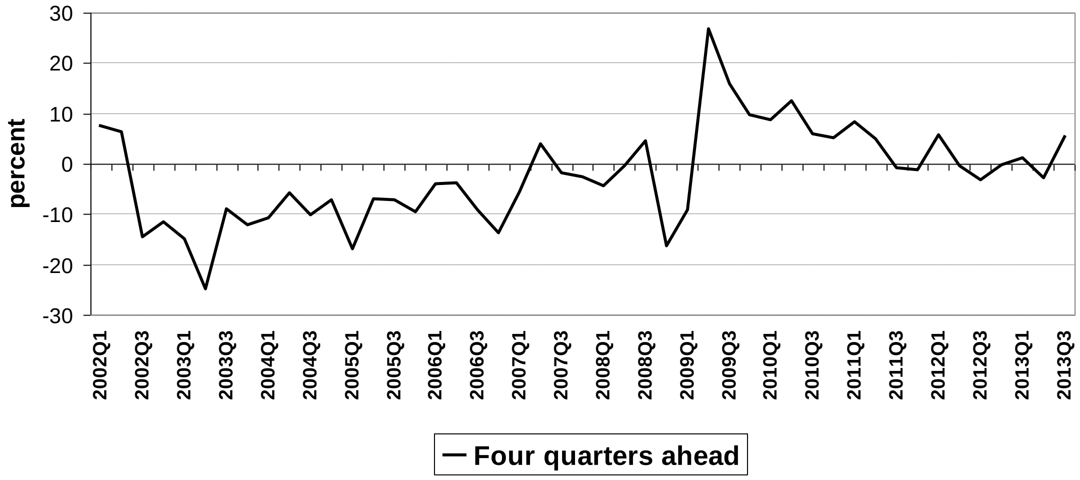
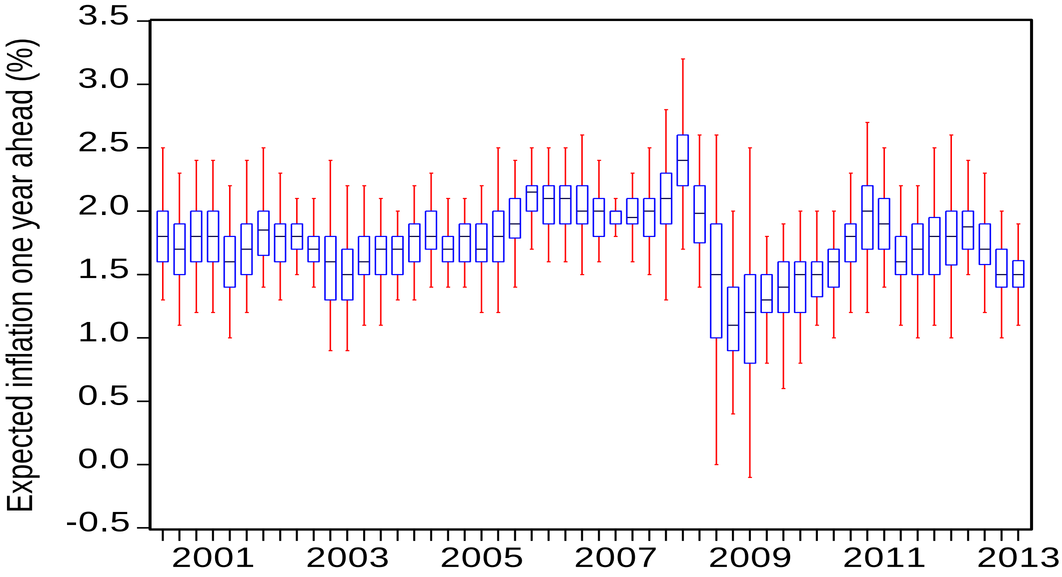
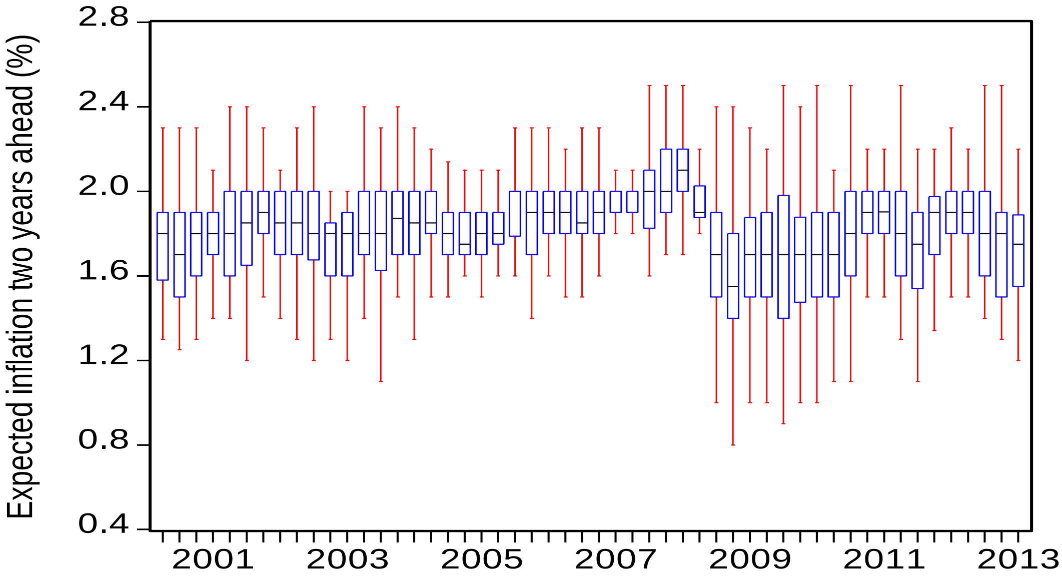


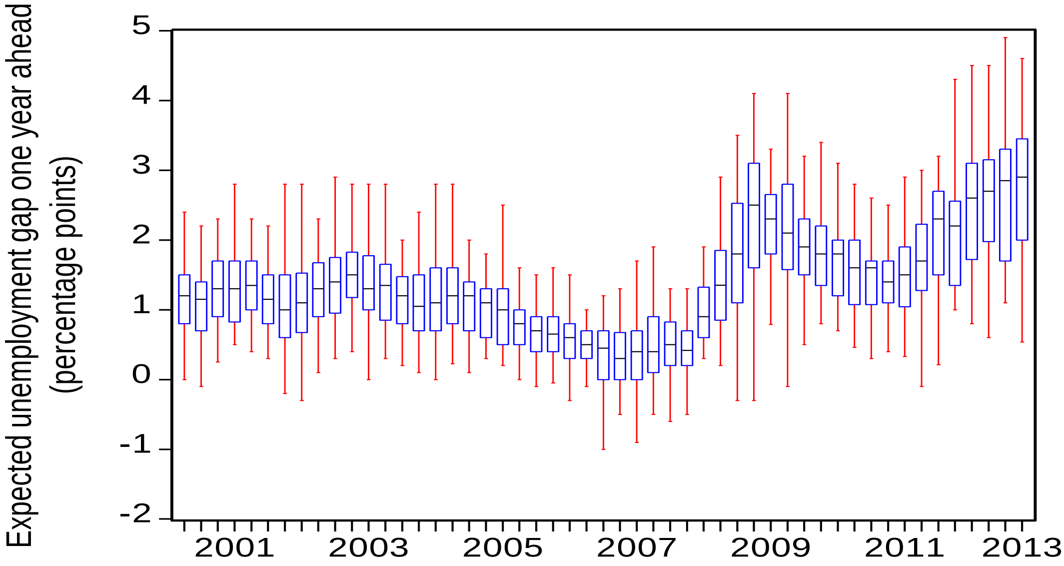
Figure 4: Cross-sectional variation of ECB-SPF forecasts

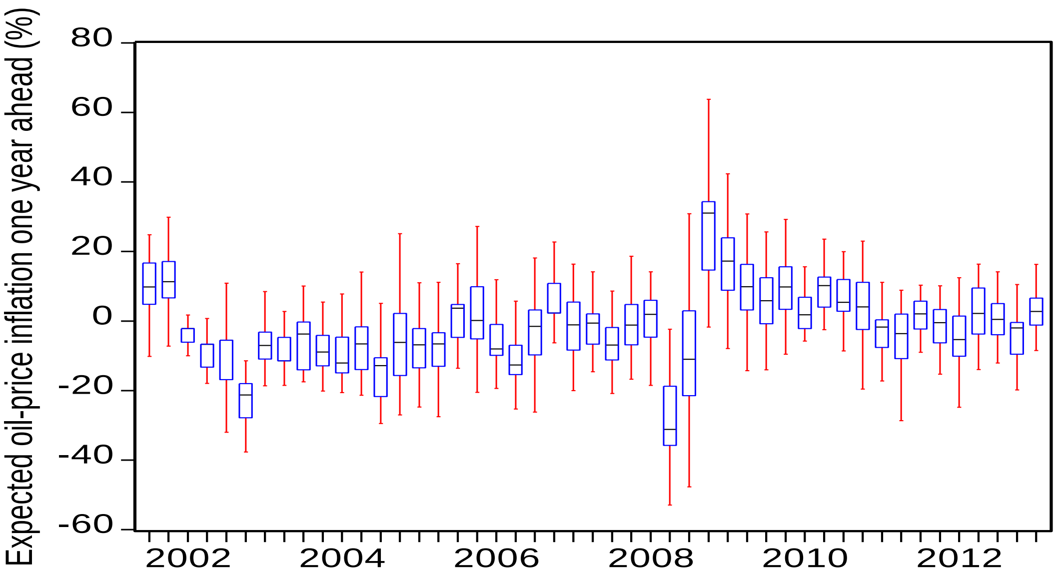




Note: Each blue box represents the inter-quartile range (IQR) from each cross section. The line in the box denotes the median observation from each cross section. The upper red lines represent the range of observations in each cross section between the first quartile and the first quartile minus 1.5 times the IQR (sometimes known as the upper *whiskers*). The lower red lines 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 very small number of forecasts fall outside the range depicted by the box and the whiskers.

Figure 4: Cross-sectional variation of ECB-SPF forecasts (cont.)





Note: Each blue box represents the inter-quartile range (IQR) for each cross section. The line in the box denotes the median observation from each cross section. The upper red lines 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 lower red lines 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 very small number of forecasts fall outside the range depicted by the box and the whiskers.

Figure 5: Expected unemployment gap and compensation per employee

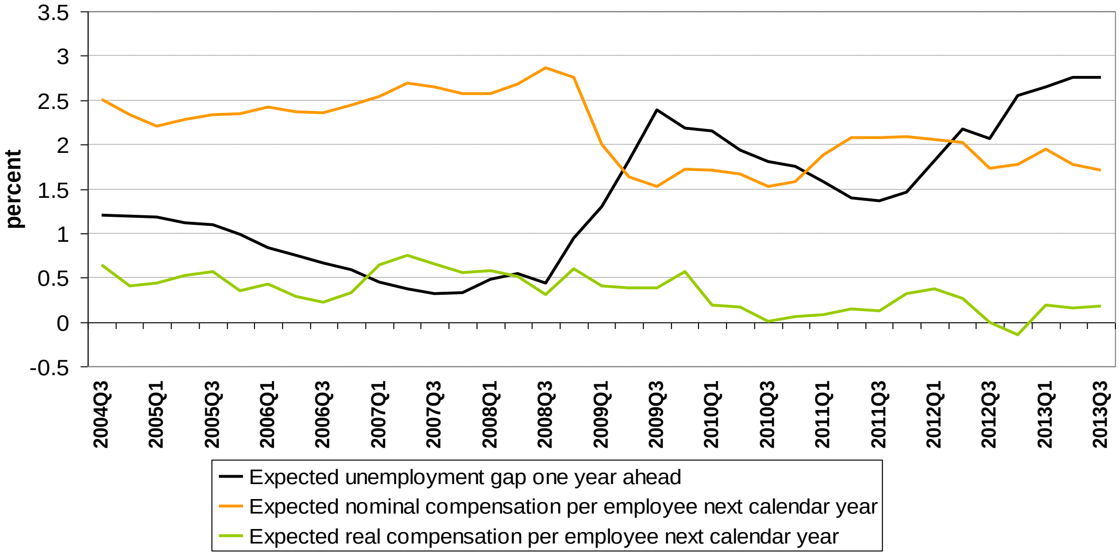


Figure 6: Expected rate of growth of real compensation per employee (individual data)