Did the crisis affect inflation expectations?

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Abstract

We investigate whether the anchoring properties of long-run inflation expectations in the United States, the euro area and the United Kingdom have changed around the economic crisis that erupted in mid-2007. We document that survey-based measures of long-run inflation expectations remained fairly stable around 2% in the euro area, fluctuated above 2% in the United States, and drifted up to about 2.5% in the United Kingdom. Expectations measures extracted from inflation-indexed bonds and inflation swaps became much more volatile in 2007. Moreover, structural break tests show that their sensitivity to news about inflation and other domestic macroeconomic variables – a measure of anchoring – increased during the crisis, and in particular during the heightened turmoil triggered by the collapse of Lehman Brothers. While liquidity premia and technical factors have significantly influenced the behaviour of inflation-indexed markets since the outburst of the crisis, we show that these factors did not contaminate the relationship between macroeconomic news and financial market-based inflation expectations at the daily frequency. While our evidence is consistent with the idea that long-run inflation expectations may have become less firmly anchored during the crisis, problems in measuring expectations accurately make it difficult to draw definitive conclusions.

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1. Introduction

> *"Ultimately, the firm anchoring of inflation expectations remains the best way to check the appropriateness of monetary policy in an uncertain environment." (Bini-Smaghi, 2009)*

After rising sharply in 2007 and the first half of 2008 – against the background of rallying commodity and food prices – global inflation went on a marked downward trend when the macroeconomic consequences of the financial crisis became visible in the fall of 2008. A few months later, price changes turned negative in a number of economies, sparking a debate on whether deflationary risks were likely to materialise and what steps would be needed to avoid deflation (Jeanne, 2009). The discussion centred to an important extent on how inflation expectations were affected by the crisis, and in particular on whether the anchoring properties of long-run inflation expectations changed as the crisis unfolded (Svensson, 2009). Our paper attempts to answer this question, by examining the behaviour of long-run inflation expectations in three major economies – the United States, the euro area and the United Kingdom – in the years around the crisis.

We proceed in two steps. We first examine the time series behaviour of two types of measures of long-run inflation expectations – survey-based measures and measures extracted from financial market instruments – for the United States, the euro area and the United Kingdom. In particular, we check whether the two measures vary around central banks' inflation objectives. The former are available at semi-annual frequency, while information extracted from financial markets is available at daily or even intra-day frequency. Markets for inflation-indexed bonds and swaps in these economies are the most liquid in "normal" times and have enjoyed sufficient liquidity to allow extracting "reasonable" measures of inflation compensation from them.

In the second step, we follow the approach of Gürkaynak et al. (2005), Gürkaynak et al. (forthcoming) and Beechey et al (forthcoming), and investigate the reaction of financial market-based measures to news about inflation and other domestic macroeconomic variables. The idea is that if long-run inflation expectations are perfectly anchored, they should not react to the arrival of news but rather be stable around the central banks target for inflation. In particular, we test whether the reaction of expectations to news has changed since the outburst of the crisis. Statistically and economically significant evidence of such a change would indicate that the crisis has influenced market participants' perception of the Fed, the Bank of England and the ECB's commitment to price stability. One conjecture is that the unprecedented monetary easing (through both conventional and non-standard monetary policies), coupled with the accumulation of a huge fiscal debt, may have undermined market participants' confidence in the ability of central banks to keep inflation at target in the longer-run.

We find two main empirical results. First, consistently with evidence presented in Beechey et al (forthcoming), survey-based measures of long-run inflation expectations point to different expectation dynamics in the three economies. In the euro area, they remained fairly stable within the central bank's comfort zone both during the oil price rally in 2006–07 and as the crisis unfolded, although their cross-sectional dispersion increased sharply in early 2009. In the United States, survey-based measures fluctuated above 2% during the whole sample period, and cross-sectional dispersion seems to have risen. In the United Kingdom, the level of survey-based measures of long-term inflation expectations has drifted up

Second, financial market-based measures became much more volatile around mid-2007. Moreover, structural break tests show that inflation expectations became more sensitive to macroeconomic news, particularly during the heightened turmoil triggered by the collapse of Lehman Brothers. We find that liquidity premia and technical factors appear to have significantly influenced the behaviour of measures of inflation expectations derived from inflation-indexed bonds or inflation swaps since the outburst of the crisis. At the same time, they do not appear to have contaminated the relationship at the daily frequency between macroeconomic news and financial market-based inflation expectations.

In the absence of formal tests, we cannot use our graphical evidence on survey-based measures of inflation expectations to draw firm conclusions on the anchoring of expectations during the crisis. This said, the observed trends in these variables are consistent with the idea that long-term expectations may have become somewhat less firmly anchored in the United States and the United Kingdom during the crisis, especially when compared to the euro area. Similarly, our understanding of day-to-day changes in financial markets remains far from perfect and shifts in liquidity are not easily distinguishable from shifts in long-run inflation expectations or perceived inflation risk. Hence, also results on financial-market based expectations measures have to be interpreted with great caution. Subject to these caveats, our evidence on structural breaks in regressions estimated with financial market-based expectations measures suggests that in all three areas, long-run inflation expectations may have become less firmly anchored during the crisis.

Differences between the dynamics of survey-based measures and measures derived from financial instruments – particularly for the euro area – may at least in part stem from difficulties in accurately measuring expectations but could also reflect important differences in the formation of expectations across different types of agents. Understanding this heterogeneity and its implications for policymakers is an important avenue for future research.

The remainder of the paper is organised as follows. Section 2 provides brief overview of the relevant literature on the anchoring of inflation expectations. Section 3 describes two measures of long-run inflation expectations: survey-based measures and measures backed out from financial instruments. Section 4 discusses the role of liquidity and technical factors in financial market measures of inflation expectations. Section 5 presents our empirical model and the main results. Section 6 concludes.

2. Literature review

While expectations and credibility play a central role in the theoretical literature, there is little theoretical work on the concept of "anchoring" of inflation expectations. In standard macroeconomic theory, if the central bank's objective function is known and constant, the rational expectations hypothesis implies that long-run inflation expectations do not change over time in response to the arrival of new information. In fact, Del Negro and Eusepi (2009) show that standard medium scale DSGE models have difficulties explaining the evolution of inflation expectations, and that the fit is even worse when the assumption of perfect information is relaxed. In recent years, a series of papers departed from the rational expectations hypothesis and the assumption of a known and constant central bank objective (e.g. Orphanides and Williams, 2005; Brazier et al., 2008; Demertzis et al., 2007, 2008). This approach allows more realistic models of the link between inflation expectations and underlying inflation.

A key element of these models is the relationship between inflation expectations and shocks to the economy. The higher the sensitivity of expectations to these shocks, the more successful monetary policy will be. In Orphanides and Williams (2005) agents do not know the true model of the economy but rather constantly update their estimates based on all information available to them. As a result, inflation expectations are sensitive to economic shocks. Orphanides and Williams (2005) introduced central bank communication in their model and find that with learning, successful communication reduces the sensitivity of inflation expectations to actual inflation.

A similar idea is found in Demertzis et al. (2007), who modelled monetary policy as an information game, in which individual agents have to interpret new (publicly available and private) information when they form *ex ante* expectations about future long-term inflation. In their model, *ex post* inflation is a function of the monetary policy chosen by the central bank to pursue its objectives and the average of all individual expectations. It is then optimal for agents to form expectations based on three elements: monetary authorities' objectives and their policy decisions; shocks that occur after these decisions; and the average of individual inflation expectations. Once the central bank communicates its inflation objective to the public — such as the ECB's operational definition of price stability as below but close to 2% – agents can either form their expectations based on the above three elements or, alternatively, coordinate their expectations on that target. They derived a time-varying parameter that captures the credibility of the central bank's target. If the target's credibility is sufficiently high, individual agents will focus their expectations on that target.

In a companion paper, Demertzis et al. (2008) took their model to the data and estimate the degree to which long-run inflation expectations in the United States have been anchored to the Fed's objective. They tested whether long-run inflation expectations– derived from the Fed's FRB model or quarterly survey-based measures – are influenced by short-run inflation dynamics. They estimated a time-varying parameter (λ) that measures the extent to which inflation expectations are anchored across quarters. They found that in recent years, the anchorness of expectations in the United States has weakened but only slightly, without compromising the Fed's credibility.

Agents may also use rules of thumb ("heuristics") to make inflation forecasts. Brazier et al. (2008) consider two heuristics: one is based on lagged inflation and the other on an inflation target announced by the central bank. In their model, agents switch between these two heuristics based on an imperfect assessment of how each has performed in the past.

The empirical literature on drivers of inflation expectations – surveyed carefully in a paper by Clark and Davig (2008) – has highlighted the role of macroeconomic variables.[1] The periodical announcements on the state of the economy and forecasts released by various (statistical) offices and agencies form a steady source of new information. To the extent that the new information is unanticipated, beliefs about future inflation may be updated. If expectations are perfectly anchored – i.e. the central bank credibly commits to its inflation objective – long-run inflation expectations should not be responsive to news about actual inflation, or more general about macroeconomic conditions.

A number of recent studies documented the anchoring of long-run inflation expectations in a number of countries. One strand of the literature relies on inflation surveys. Levin et al. (2004) analysed the behaviour of private-sector inflation forecasts at horizons up to ten years – measured by quarterly Consensus forecasts – in United States and the euro area over the period 1994–2003. They found that expectations were highly correlated with a three-year moving average of lagged inflation. By contrast, in industrial countries that have adopted inflation targeting (United Kingdom, Sweden, Canada, Australia and New Zealand), inflation expectations were found not to be sensitive to actual inflation. Levin et al. (2004) concluded that inflation targeting has played a significant role in anchoring long-run inflation expectations. Paloviita and Viren (2005) found that inflation expectations, proxied by OECD inflation forecasts, respond to changes in output and actual inflation. The results are based on a simple VAR model with inflation, inflation expectations, estimated with pooled annual data for euro area countries over the period 1979–2003. Clark and Nakata (2008) showed that in the United States, inflation expectations appeared to be slightly better anchored in recent years compared to 20 or more

[1] Another strand of literature relies on economic experiments (e.g. Marimon and Sunder, 1994; Hommes et al, 2005, 2007; Adam, 2007).

years ago. In particular, they found evidence of a declining impact of unexpected increases in inflation on long-term expectations.

A second strand of the literature extracted inflation expectations from inflation-indexed financial market instruments, and looked at the relationship between inflation expectations and macroeconomic variables at high (daily or intraday) frequency (Swanson, 2006). Gürkaynak et al. (2003, 2005) derived inflation expectations from inflation-indexed bonds and examined their sensitivity to surprises about macroeconomic announcements at the daily frequency. To test for anchoring of inflation expectations, they regressed daily inflation expectations on a set of macroeconomic news variables. They found that between 1990 and 2002, long-run inflation expectations in the United States were not perfectly anchored. This analysis was extended by Gürkaynak et al. (forthcoming) and Gürkaynak et al. (2006), who documented that long-run inflation expectations are more solidly anchored in the United Kingdom, Sweden and Canada – countries that adopted formal inflation targets. Consistently with Levin et al. (2004), they concluded that a numerical inflation target has helped anchoring long-term inflation expectations.

Beechey et al. (forthcoming) followed the same methodology to compare the anchoring properties of long-run inflation expectations in the United States and the euro area over the period 1 June 2003 – 31 December 2006. They found that surprises about monetary policy decisions and macroeconomic data releases – the core CPI but also indicators of economic activity such as the National Association of Purchasing Managers (NAPM) index or non-farm payrolls – have significant effects on US forward inflation compensation at different horizons. By contrast, long-term inflation compensation does not significantly react to any news about price or output developments in the euro area. A similar picture emerges from graphical evidence on medians and dispersions of survey-based inflation expectations in the United States and the euro area. Beechey et al. (forthcoming) concluded that long-run inflation expectations are more firmly anchored in the euro area than in the United States.

This paper builds on the literature by developing a method to assess possible changes in anchoring of inflation expectations around the recent crisis. Mounting inflationary pressure due to booming commodity prices in the run-up to the crisis might have caused inflation expectations to drift. Similarly, the crisis may well have led to drifting inflation expectations in the wake of unsurpassed monetary and fiscal expansion.

3. Measuring inflation expectations

Survey based measures

There are two main approaches to measuring long-run inflation expectations. A common approach relies on inflation surveys.[2] A frequently used data source, *Consensus Economics*, provides semi-annual data on expectations of a panel of some 30 professional forecasters six to ten years ahead, for a number of countries. There are also survey data for somewhat shorter horizons. For example, at a horizon of 5 years, the ECB Survey of Professional Forecasters (SPF) collects, on quarterly basis, forecasts by a panel of some 70 professional forecasters on euro area HICP.

Survey measures have several important shortcomings. First, given their low frequency, survey measures appear well-suited for analysing longer-run properties of inflation expectations but not for identifying the existence and timing of breaks in the expectation formation over a short horizon. Second, survey results may not be reliable to the extent that respondents do not have to act on the basis of their responses – i.e. "do not put their money where their mouth is".[3] Third, as shown by Van der Klaauw et al. (2008), survey results are sensitive to the wording of the questions. Fourth, different types of survey measures may produce very different results. Mankiw et al. (2003), for example, looked at 50 years of data on inflation expectations in the United States, and documented substantial disagreement among both consumers and professional economists about expected future inflation. They found that this disagreement varied substantially through time, depending on the level of inflation, the absolute value of the change in inflation, and relative price variability.

Figures 1–3 show survey-based measures of inflation expectations in the United States, the euro area and the United Kingdom around the crisis. For each area we included the measures with the longest horizon available. Each graph plots the mean value of inflation forecasts from Consensus Economics between 5 and 10 years ahead (labelled 'CE 5to10 mean'). For the United States and the euro Area we also included the SPF expectations at 5 years from now (for the Euro Area, labelled 'SPF at 5yr mean') and over the next 10 years (for the US, labelled 'SPF 10yr median'). For SPF expectations we also obtained a measure of the cross-sectional dispersion given by the standard deviation of respondents' answers, which is plotted as the dotted 'SPF sd' line.

As documented also in Beechey et al. (forthcoming), these graphs suggest that long-term inflation expectations followed different dynamics around the crisis in the three economies. In the euro area, expectations remained fairly stable within the ECB's comfort zone both during the oil price rally in 2006–07 and as the crisis unfolded. Their cross-sectional dispersion did also not change significantly. In the United States, survey-based measures always exceeded 2% and both survey sources started to

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[2] For a detailed analysis of the properties of these measures, see ECB (2006) and Clark and Davig (2008).

[3] This point is emphasised in the experimental economics literature on inflation expectations (Smith, 1982).

disagree on the long run level of inflation starting 2007. The dispersion of respondents of the SPF also started to rise somewhat during this period. In the United Kingdom, the level as expected by Consensus Forecasts respondents follows a remarkable upward trend which started in 2007. Moreover, while measures of dispersion of long-term inflation expectations are not available for the United Kingdon, recent issues of the Bank of England Inflation Reports show that the dispersion of 9-quarter-ahead inflation projections of professional forecasters has risen substantially over the past several years.

The limitations of survey-based expectations measures do not allow drawing strong conclusions from graphical evidence. Moreover, because of the low frequency of survey data, it is not possible to formally test changes in the anchoring properties of long-term expectations. This said, Figures 1–3 are consistent with the idea that the anchoring properties of long-term inflation expectations may have changed in the United States and the United Kingdom during the crisis. By contrast, they seem to point to no changes in the anchoring of long-term expectations in the euro area.

Financial market based measures

A second approach to measuring long-run expectations consists in using financial market-based measures. In a number of countries, bonds or interest rate swaps that are linked to some measure of domestic inflation are actively traded (Deacon et al., 2004, and JP Morgan, 2008). These instruments can be combined with nominal bonds or nominal interest rate swaps to back out financial markets' inflation expectations. The main advantage of this type of measure is that, given its high frequency, it allows examining more formally changes in the behaviour of expectations over a relatively short horizon. It is therefore most useful in investigating whether the financial crisis has affected the anchoring of inflation expectations to the Fed's, the ECB's, and the Bank of England's inflation objectives.

In this paper we derived inflation expectations from nominal and inflation-indexed financial instruments. In particular, we considered inflation-indexed bonds for the United States (Treasury Inflation-Protected Securities, TIPS) and the United Kingdom (inflation-linked Gilts), and inflation swaps for the euro area. These instruments are actively traded and the most liquid ones among inflation-indexed products.[4] For example, market commentary indicated that the monthly trading volume of 10-year euro area inflation swaps averaged around 6bn in 2007 (JP Morgan, 2008). Buyers of inflation swaps primarily include insurance companies and pension funds, which suffer a loss of

[4] Alternatively, we could have used nominal and inflation-indexed bonds also for the euro area, which in recent years have become increasingly liquid. Our preference for swaps is motivated by their greater liquidity along the whole maturity spectrum, particularly in the earlier years in our sample period. Beechey et al. (forthcoming) note that inflation compensation derived from swaps and TIPS are very similar. For a more extensive discussion of the difference between inflation-indexed bonds and inflation swaps, see Deacon et al. (2004) and JP Morgan (2008).

income if the actual inflation rate increases. Typical inflation swap sellers include firms whose income is linked to inflation while their expenses are not, or only to a lesser degree, such as public authorities, utilities, real estate companies, or distribution companies. By selling inflation in an inflation-linked swap, they are able to protect future income linked to inflation. Data for the yield curves of UK inflation-indexed bonds were taken directly from the Bank of England, while data on US TIPS and euro area swaps were obtained from Bloomberg.[5]

Here we describe in detail the method followed to extract inflation expectations from swaps markets but the main features apply also to the approach we followed for bond markets. Inflation expectations can be measured by one-year zero-coupon forward rates of inflation swaps. Since we focus on long-term inflation expectations, we chose one-year zero-coupon forward rates ending ten year ahead. Hence, at day t inflation expectation are measured by

(1) $$f_t = f_{t,10}^{is},$$

where $f_{t,10}^{is}$ denote the one-year zero-coupon forward rates ending ten year ahead for inflation swaps.

We collected daily data on euro area swap markets for the period 23 June 2004 to 24 March 2009. For each day, swap rates are available for different maturities, allowing us to estimate a whole yield curve. The forward rates ending ten year ahead can be calculated from the swap rates with 9- and 10-year maturities as

(2) $$f_{t,10} = \frac{(1+y_{10})^{10}}{(1+y_9)^9} - 1,$$

where y_9 and y_{10} are the 9-year and 10-year swap rates, respectively.[6]

One concern is that since the market for swaps with a 9-year maturity may be less liquid than the market for 10-year maturity, estimates of forward rates based on 9-year swaps may exhibit a high level of noise. In order to filter out this noise, we followed a standard technique of the finance literature – the Nelson-Siegel method (see, e.g. Nelson and Siegel, 1987) – to estimate a smooth yield curve for each day. Details on this method are presented in Appendix 1. Söderlind and Svensson (1997) argued that estimating yield curves by a simple curve fitting, as we did, rather than by a structural model for interest rate dynamics is appropriate when the purpose is to extract market expectations about future interest rates without making additional assumptions about the model structure.

5 See http://www.bankofengland.co.uk/statistics/yieldcurve/index.htm.

6 The same formula to calculate the forward rates applies to both inflation swaps and interest swaps. This argument holds for the rest of the section.

We obtained a smoothed yield curve for each day for inflation swaps, from which we got the smoothed 9- and 10-year inflation swap rates. We then estimated 10-year ahead forward rates for the inflation swaps, from which we derived a measure of inflation expectations. Note that our smoothing procedure is applied for each specific day but it gives also a smoother measure of inflation expectations across time. This smoothing effect can be explained in terms of a smaller impact of liquidity effects. Throughout the paper, we used the smoothed series as a measure of long-run inflation expectations.

Figures 1–3 show that our measures of long-run inflation expectations - based either on nominal and inflation-indexed bonds or on inflation swaps - differ significantly from the survey-based measures of inflation expectations. The expectations derived from financial markets are plotted as the solid lines labelled ‘FM 9to10’. The difference is most noteworthy in the euro area, where over the whole sample period, survey measures were very stable and close to the ECB’s objective of medium-term inflation below but close to 2%. By contrast, inflation expectations derived from inflation swaps swung between 2% and 2.8% and became much more volatile after the onset of the crisis. Similarly, financial market and survey based measures of long-run inflation expectations also differ visibly for the United States and the United Kingdom.

4. The role of liquidity premia and technical factors

One potential reason for the visible difference between survey and financial market based measures of inflation expectations is that the latter may be contaminated by other factors, especially during the crisis.[7] Break-even rates, i.e. the difference between nominal and inflation-indexed bonds, can be decomposed into four main factors: expected inflation, inflation risk premia, liquidity premia, and technical factors (Hördahl, 2009). According to Hördahl (2009), the same applies to a much lesser extent to inflation-indexed swaps. One example of these technical factors are sudden portfolio shifts by leveraged investors that may affect nominal and inflation-indexed bond markets but are unrelated to changing views about future economic fundamentals. One critical assumption of exercises that back out measures of inflation expectations from financial instruments is that the last two factors do not respond to macroeconomic news at the daily frequency. Changes in far-horizon break-even rates can then be interpreted as a revision of market participants’ long-run inflation expectations or inflation risk in response to new information, indicating that inflation expectations are not solidly anchored.[8]

In “normal” times this assumption appears plausible. Dudley et al. (2009) documented the deepening of TIPS markets by looking at trading volumes, bid-ask spreads and estimates of illiquidity premia.

[7] See e.g. Barclays Capital Research (2008).

[8] Inflation risk reflects both the volatility of inflation expectations as well as market participants’ attitude towards risk. If inflation expectations are firmly anchored, none of these two components should react to new information.

Beechey et al. (forthcoming) documented that the announcement effects on inflation expectations measured using TIPS or inflation swaps persist for about one business week. They interpreted this as evidence that the reaction of the dependent variable is not driven primarily by liquidity effects. Beechey and Wright (2008) decomposed US nominal yields into three components – nominal yields, real yields, and the spread between these two (i.e. inflation compensation) – and then estimated the effect of news on these three components using intra-day data. They found that different types of news – about prices, the real economy or monetary policy – have quite different effects on real rates and rates of inflation compensation. In particular, only news about prices affect inflation compensation. They also tested whether the impact of news has changed over time, since the market for inflation-indexed bonds in the United States – Treasury Inflation Protected Securities (TIPS) – has deepened. Their evidence suggests that the reaction of long-term inflation compensation to inflation news has not changed between 17 February 2004 and 13 June 2008, implying no significant change in inflation expectations' anchoring properties during this period.[9]

During the current crisis, the assumption that changes in market liquidity and technical factors may contaminate the behaviour of financial market-based measures of inflation expectations appears much less innocuous. Financial markets, including bond and swap markets, experienced pronounced swings in volatility and liquidity. In these circumstances, there was evidence that yields on nominal and inflation-indexed bonds (and swaps) were driven not just by expectations about future inflation but also by high and volatile liquidity premia and technical factors related e.g. to hedging activity (Fender et al., 2009; Hördahl, 2009).

Note that by construction, our measure of inflation compensation filtered out part of the noise. In particular, by taking the difference between nominal and inflation-indexed bonds, we purged the effect of liquidity and technical factors that affect both markets in a similar way. For example, if on a particular day there is a sudden broad portfolio shift out of fixed income markets in general and into equity markets, nominal and inflation-indexed bond yields could both be affected in a similar fashion. Moreover, we focus on day-to-day changes in inflation compensation, and the relative liquidity of the nominal and inflation-indexed markets may not change substantially from day to day. The same might be true for inflation swaps.

This discussion suggests that simple graphical evidence on the behaviour of break-even rates can be misleading: the "true" inflation expectations may be anchored even though financial market based measures are not close to the central bank's objective if the influence of liquidity and technical factors

[9] By contrast, using daily data from early 1999 to early 2004, Beechey and Wright (2008) compared the performance of their model before and after 2004 and found evidence that inflation compensation became less sensitive to news after 2004. They interpreted this result as suggesting that the improved liquidity and functioning of the TIPS market since 2004 may have allowed TIPS yields to become more responsive to new information.

is not filtered out properly. In addition, even if our expectation measure is "accidentally" close to the central bank's objective, it is not sufficient to conclude that inflation expectations are firmly anchored. In this case, the movements of expectations are influenced by macroeconomic (and other) news, which happen to drive inflation expectations close to the central bank's objective. To get more accurate evidence on the anchoring of long-run inflation expectations, we therefore turn to examine their response to macroeconomic news.

5. Empirical results on breaks and anchoring

In order to test whether inflation expectations became unanchored during the crisis, we examined the impact of news on HICP inflation and other macroeconomic variables on our measures of inflation expectations. Our focus is on testing for structural breaks while improving estimation by dealing explicitly with liquidity effects. Our sample period is 23 June 2004 – 23 March 2009, and includes almost two crisis years during which liquidity effects may have been especially important.

Following the approach developed by Gürkaynak et al. (2006) and Beechey et al. (forthcoming) – who analysed the credibility of the Fed, the ECB and inflation targeting central banks – we captured news by the difference between actual releases of the main euro area macroeconomic variables and values anticipated by market participants according to surveys conducted by Bloomberg and JP Morgan. This is a common method in the literature, although Rigobon and Sack (2008) found that it tends to underestimate the responses to true news because of measurement errors. In fact, Bartolini et al. (2008) discussed shortcomings of survey-based measures of news but concluded that this approach may be the only one available in practice. We normalize the macro data surprises by the standard deviation of each series, which allows the coefficient estimates to be interpreted as the impact of a one-standard-deviation surprise in a given data release.

We expect that data releases on inflation variables are most important. However, without empirical guidance from the literature, we have few other priors on what type of information influences inflation expectations. Other macroeconomic variables – such as GDP growth, business confidence indicators, the unemployment rate or wage growth – may also give indications about possible inflationary pressures. We used the same macro-announcements as in Beechey et al. (forthcoming) with two exceptions. In the regressions where the dependent variable is euro area or UK inflation expectations, we do not include US news. Moreover, in our regressions with US data, the macro news do not include oil futures. [10] For the euro area, we concentrated on data releases for the three main economies – Germany, France, and Italy – since these are most likely to have a primary influence on views on

[10] Differently from Beechey et al. (forthcoming), we measure changes in inflation compensation in percentage points rather than in in basis points.

future euro area inflation. Considerably more news variables are available through Bloomberg and JP Morgan for the United Kingdom and the United States. Appendix 2 lists all variables that were used.

We regressed our measure of long-run inflation expectations on a constant, a set of macroeconomic news variables and a set of control variables, according to the following model:

$$(3) \qquad \Delta f_t = \alpha + \beta \mathbf{X}_t + \gamma \mathbf{Z}_t + \varepsilon_t$$

where the dependent variable $\Delta f_t = f_t - f_{t-1}$ is the change, from closing of the markets at day *t-1* to closing on day *t*, in one-year inflation compensation ten years ahead.[11] The explanatory variables $\mathbf{X}_t$ are a vector of news variables on various measures of the state of the economy. Most macroeconomic news arrives at 8:30 am, before stock markets open.[12] Our expectations variable therefore measures the change in inflation expectations between the end of the trading day before macro news arrive (*t-1*), and the end of the trading day on which the news arrive (*t*).

$\mathbf{Z}_t$ is a vector of control variables intended to capture the influence that shorter-term changes in liquidity premia and technical factors unrelated to inflation expectations may have on inflation swap rates, or on the difference between nominal and inflation-index bonds. In particular, our control variables are useful in purging the effect of liquidity premia and technical factors that are related to shocks that broadly hit financial markets. For example, a sudden increase in financial stress on a particular day due to news about the collapse of a major financial player may induce a broad flight to liquidity. This could benefit nominal bonds at the expense of other assets such as inflation-index bonds. Our preferred control variable is the implied volatility of bond yields but we checked that our results were robust to using four alternative variables that measure market liquidity: the Chicago Board Options Exchange Volatility Index (VIX), a widely used measure of the implied volatility of S&P 500 index options; the euro bund implied volatility; the on-the-run off-the-run spread (a commonly measure of bond market liquidity); and analogously for the euro area, the KfU-bund spread.[13] All these non-news controls were first differenced as well, because a change in implied volatility should lead to an additional change in our measure of inflation expectations through a liquidity premium. We also controlled for day-of-the-week effects but these turned out not to be statistically significant.[14]

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11 Since heteroskedasticity-consistent standard errors can be misleading when the explanatory variable only has few non-zero values, we use conventional OLS standard errors. Using these rather than heteroskedasticity-consistent standard errors does not affect our results substantially.

12 See for example JP Morgan (2009).

13 For a detailed description of this variable, see Hördahl (2009).

14 For reasons of space, the results for these dummies are not reported here. All variables used in the regressions are stationary.

We interpret a high R^2 – to the extent that it is driven by significant coefficients on the variables X_t and that the level of inflation expectations differs significantly from the central bank's comfort zone – as evidence that expectations are weakly anchored. A low R^2 conversely implies well anchored inflation expectations. A change in explanatory power of the model during the sample period would then indicate that anchoring properties of inflation expectations have changed. In particular, we verified whether the sensitivity of inflation expectations to news about inflation and other macroeconomic variables increased in 2006, when first commodity and food prices started to rally, and in 2007, when the financial crisis erupted.

To detect such changes, we need to test for a structural break. We used the Chow test, which verifies whether the structure of the model changes on a certain day. However, this is an ex post test in the sense that the timing of the break must be chosen with some prior knowledge. Chow tests for dates that are close to the real structural break will also give statistically significant results. To get a more precise indication of the timing of the structural break, we therefore used a rolling version of the Chow test.[15] Changing the date at which we split the sample gives us a measure of when, if at all, the model starts to predict changes in inflation expectations. If we find that the model fit increases after the break we may conclude that expectations have become less anchored from that point on. Note that we adjusted the Chow test statistic so that we capture only the change in model performance due to news variables.[16]

As a further step, we considered the most conservative dating of a structural break, given by the maximum of the test statistics from the rolling Chow tests (see Zeileis et al., 2003). We examined the change over time in both the F-test value and p-value to assess the timing of a structural break. [17]

Our results – summarised separately for the United States, the euro area and the United Kingdom in Figures 4–6 – indicate that the sensitivity of inflation expectations to news has indeed changed during the sample period. Each figure plots on the left axis the p-value of the Chow test for each day, and on the right axis the actual value of the test statistic. Several results stand out. First, the test statistics (and, inversely, the p-values) all increase towards end-2008 and early 2009, which coincides with the period of highest financial turmoil . The closer we get to the period of heightened financial stress – as highlighted by the dotted vertical line, the date at which Lehman Brothers filed bankruptcy – the more

[15] Note that the rolling Chow test does not rely on standard errors. Our analysis for the break dates is therefore unaffected by the choice between conventional OLS or heteroskedasticity-consistent standard errors (see footnote 13).

[16] If liquidity premia are important, then the implied volatility variables may also cause the Chow test to flag a break, just because volatility increased during the crisis period. To prevent this, we replaced the sum of squared residuals from the three regressions (entire period, before and after potential break point) in the test statistic with one minus their partial R-squared – where the explanatory effect of the control variables was partialled out – multiplied by their respective total sum of squares.

[17] We truncate the sample by 100 observations on each side to allow for feasible testing near the beginning and end of the sample.

different the model performs in the later sample period compared to the first period. This indicates that in all three areas, the sensitivity of inflation expectations to macroeconomic news changed during the crisis.

Secondly, we found a break in each economic area. The dates at which the Chow test becomes significant for the first time at the 99% confidence level are highlighted in the graphs by vertical dashed lines, and are reported in Table 1. We identified a first statistically significant change in the behaviour of inflation expectations in July 2006 for the United States, in December 2007 for the euro area, and in April 2008 in the United Kingdom.

Thirdly, the test statistics show a hump shaped pattern around the crisis period in all three economies, suggesting strong evidence for a break at the height of the crisis and decreasing evidence in the following months.

Table 1 shows, in addition to the timing of the first break identified by Chow tests, dates on which the Chow test statistics is maximized. Using this more conservative method of dating structural breaks in the relationship between inflation expectations and news we find more similarities across the United States, the euro area and the United Kingdom. For all three economies, the maximum Chow test statistic can be found between September and November of 2008, a period that is commonly considered as the height of the financial crisis, and encompasses the collapse of Lehman Brothers on 15 September 2008.

As a robustness check, we used the method developed by Andrews (1993), which tests whether a structural break exists within a short period without knowing the break point.[18] The essence of these tests is to look at the test statistics yielded by Chow-tests on each day within the period under consideration, and construct new unified testing statistics from them. This helps to narrow down the location of the structural break. Once we identified the day on which the F-test statistic of our rolling Chow tests reaches its maximum, we applied the Andrews test to verify whether it identifies significant structural breaks around that day. The Andrews test indeed confirmed that at a 99% confidence level, a structural break exists on the days on which the Chow test statistics are maximized for all three economic areas.

Having identified structural breaks in the relationship between inflation expectations and inflation news and macroeconomic news, we looked at the fit of the model in each sub-sample to get evidence on whether inflation expectations became more or less anchored after those breaks. Tables 2– 4 show

[18] See also Andrews and Ploberger (1994). Our approach is very similar to that used by Beechey and Wright (2008) to test for parameter instability.

the regression results for respectively the United States, the euro area and the United Kingdom, and distinguish three different sample periods.[19] In each table, the first column refers to the entire sample period, while columns 2 and 3 refer sub-sample before and after the first break point identified by rolling Chow tests. Columns 4 and 5 report results for the sub-samples before and after the day on which the Chow statistic reached its maximum.

The tables highlight that in all three economic areas, the sensitivity of inflation expectations increased in the run-up to and during the crisis. In particular, the reaction of inflation expectations to news on US inflation was much higher in July 2006–March 2009 than in June 2004–July 2006.[20] The partial R^2, which excludes the effect of our measures of liquidity and technical factors, increased from 1.9% to 5.1% between the first and the second sub-sample. In the euro area, the partial R^2 rose to 6.4% in the period December 2007–March 2009, from 2.9% during the period June 2004–December 2007. The difference is even starker in the United Kingdom: the partial R^2 increased almost six-fold, from 1% to 5.7%.

The change in model fit over time is even more evident when we split the sample period on the dates on which the Chow test statistics are maximized (columns 4 and 5 in Tables 2–4). The increase in model performance – measured by the partial R^2 – is most visible for the United States, where economic news explained only 2.5% the of changes in inflation expectations before 30 October 2008 but 27.2% of the variation after that date! The partial R^2 surged also in the regressions for the euro area, from 2.8% before 23 October 2008 to 19.7% after. The model fit also increases from 2% to 11.0% for the United Kingdom. We interpret these result as a warning sign that around the end of 2008, and probably earlier, long-run inflation expectations might have started to be less perfectly anchored. This is true irrespective of the monetary policy framework and the type of strategy that the Federal Reserve, the ECB and the Bank of England followed in reaction to the crisis.

To get a sense of the magnitude of the sensitivity of inflation expectations to news, we compared our results to those reported by Gürkaynak et al. (2006) and Gürkaynak et al. (forthcoming) for the United Kingdom and United States for data up to 2005.[21] They found that economic news was able to explain

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19 The results for the pre-crisis period are broadly consistent with the findings by Beechey et al. (forthcoming). The main difference is that we also find statistically significant effects on euro area inflation expectations of news for France PPI, the German current account, Italian business confidence, and the Italian PPI. This difference is likely to come from the inclusion of the period 31 July 2003 – 23 March 2004 in the Beechey et al. (forthcoming) regressions, and possibly from the different source of survey data on US macro announcements – Beechey et al. (forthcoming) use Money Market Services, while we rely on Bloomberg.

20 For the euro area and the United Kingdom, these results are not sensitive to the presence of outliers, which is noteworthy since surprises in macro data releases tend to have heavy-tailed distributions. For the United States, excluding four outliers causes the partial R^2 to no longer increase after the first break, although the partial R^2 does increase substantially after the maximum break. We conclude is that for the United States, outliers have some influence on the results of break tests but they do not drive the main results.

21 Beechey et al. (forthcoming) do not report measures of model fit, so we could not compare what we found to their results.

4% of the variation of US inflation compensation between 1998 and 2005.[22] The results in Table 3 suggest that the sensitivity of inflation expectations to news increased significantly during the crisis. The same holds for the United Kingdom. Gürkaynak et al. (forthcoming) report that after the Bank of England became independent, economic news only explained up to 3% of the variation in inflation compensation (between 1998 and 2005).[23] We found that after September 2008 our model fit rose to 10.9%, which lies about half in-between our model fit for the pre-crisis years (2004-2006) and the model fit that Gürkaynak et al. (2006) found for the period before the Bank of England gained independence (1993-1997).

6. Conclusions

To what extent have inflation expectations been affected by the economic crisis that erupted in mid-2007? In particular, have anchoring properties of long-run inflation expectations changed as the crisis unfolded? In our paper we addressed these questions by examining long-run inflation expectations in the United States, the euro area and the United Kingdom between June 2004 and March 2009. We considered two types of measures of long-run inflation expectations: survey-based measures and measures extracted from inflation-linked financial market instruments.

In the period 2004–2009, the dynamics of survey-based measures of long-run inflation expectations differed across the three economies. They remained fairly stable within the ECB’s comfort zone in the euro area, fluctuated above 2% in the United States, and drifted up in the United Kingdom. In addition, the cross-sectional dispersion of survey-based measures rose markedly in the United States and the United Kingdom.

Expectations measures extracted from inflation-indexed bonds and inflation swaps show a marked increase in volatility since 2007. Moreover, we found evidence that their sensitivity to news about inflation and other domestic macroeconomic variables has increased since 2006. The reactivity to news increased particularly during the period of heightened turmoil triggered by the collapse of Lehman Brothers in September 2008, when central banks significantly eased monetary policy using both standard and non-conventional tools.

Although it has been argued that liquidity premia and technical factors have significantly influenced the behaviour of inflation-indexed markets since the outburst of the crisis, we found that they did not contaminate the relationship between macroeconomic news and financial market-based inflation expectations at the daily frequency.

[22] This includes the statistically significant effect of several first-business-day-of-the-year dummies, which we did not include in our regressions.

[23] The explanatory power is not higher when also international macro news were included.

While our graphical evidence on survey-based measures of inflation expectations does not allow definitive conclusions on the anchoring of expectations during the crisis, it is consistent with the idea that long-term expectations may have become somewhat less firmly anchored in the United States and the United Kingdom. Similarly, our results on financial-market based expectations measures have to be interpreted with great caution given that our understanding of day-to-day changes in financial markets remains far from perfect and shifts in liquidity are not easily distinguishable from shifts in long-run inflation expectations or perceived inflation risk. With these caveats in mind, our evidence on structural breaks suggests that in the United States, the euro area and the United Kingdom, long-run inflation expectations may have become less firmly anchored during the crisis.

One interpretation of our results is that at the height of the crisis, market participants viewed monetary authorities as focusing mainly on fixing the monetary transmission mechanism and on softening the impact of financial instability on the real economy. This appears to have affected market participants' views on the commitment of central banks to fighting inflation. In the midst of a crisis, this seemed less of a problem given the absence of inflationary pressures. However, once the economy regains traction, central banks need to carefully devise exit strategies from their expansionary monetary stance.

Whether the extent to which anchoring properties have changed during the crisis depends on monetary frameworks – such as an inflation targeting regime – or the strategies that were followed to counter the impact of the crisis, is an important topic for future research. In terms of methodology, we conclude that the approach developed by Gürkaynak et al. (2005) and others is flexible enough to be used to investigate the properties of expectations during crisis times.

Differences between the dynamics of survey-based measures and those of measures based on financial instruments may at least in part stem from difficulties in measuring expectations accurately. They could also reflect important heterogeneities in the expectation formation mechanism across different types of agents. Understanding these differences and their implications for policymakers is an important avenue for future research.

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Tables

Table 1: Break point dating

		first date significant	maximum test value
Euro Area	**Break date:**	**11dec2007**	**23oct2008**
	Chow F-test stat.	2.064	5.081
	p-value	0.005	7.95e-12
	Andrews' test stat.		101.62
	1% critical value		50.76
UK	**Break date:**	**16apr2008**	**09sep2008**
	Chow F-test stat.	2.302	5.615
	p-value	0.005	4.93e-10
	Andrews' test stat.		84.23
	1% critical value		42.05
US	**Break date:**	**17jul2006**	**30oct2008**
	Chow F-test stat.	2.154	13.200
	p-value	0.001	1.35e-27
	Andrews' test stat.		184.8
	1% critical value		40.10

Regressions are run as specified in Tables 2–4.

Table 2: Regression results for the United States

	(1)	(2)	(3)	(4)	(5)
		1st significance		**maximum significance**	
Break date (99% confidence level):		**17jul2006**		**30oct2008**	
	23jun2004 - 23mar2009	pre-break	post-break	pre-break	post-break
overall R^2	0.033	0.024	0.065	0.032	0.289
partial R^2	0.023	0.019	0.051	0.025	0.272
Observations	1189	516	673	1092	97
ISM Manufacturing PMI SA (value; NAPM)	0.005	0.009	0.001	0.014***	-0.063
	(0.007)	(0.007)	(0.011)	(0.005)	(0.046)
US Personal Consumption Expenditure Core Price Index MoM SA	0.006	-0.007	0.015	0.005	-0.017
	(0.008)	(0.009)	(0.011)	(0.006)	(0.082)
US Capacity Utilization % of Total Capacity SA	0.020*	0.000	0.024	0.010	0.066
	(0.012)	(0.015)	(0.017)	(0.011)	(0.057)
Conference board consumer confidence SA 1985=100	0.008	-0.003	0.016	-0.006	0.079**
	(0.007)	(0.007)	(0.010)	(0.006)	(0.039)
US Industrial Production MoM 2002=100 SA (rate)	-0.027**	0.002	-0.034**	-0.023*	-0.026
	(0.012)	(0.018)	(0.016)	(0.012)	(0.044)
US initial jobless claims SA	-0.009***	0.000	-0.016***	-0.001	-0.030**
	(0.003)	(0.004)	(0.005)	(0.003)	(0.014)
Conference board US leading index MoM	0.016**	-0.006	0.030***	-0.007	0.126***
	(0.007)	(0.008)	(0.010)	(0.006)	(0.037)
Federal Funds Target Rate	0.009	no	0.010	0.007	0.040
	(0.008)	news	(0.009)	(0.007)	(0.077)
US new privately owned housing units started by structure total SAAR (units/thou	-0.003	-0.004	-0.002	-0.005	-0.065
	(0.007)	(0.006)	(0.014)	(0.005)	(0.157)
US Employees on Nonfarm payrolls total MoM Net Change SA (thousands)	0.001	0.011	-0.010	0.005	-0.023
	(0.007)	(0.007)	(0.013)	(0.005)	(0.038)
Adjusted retail & food services SA total monthly % change	0.014**	-0.015*	0.028***	-0.014**	0.070**
	(0.007)	(0.008)	(0.010)	(0.006)	(0.027)
US unemployment rate total in labor force SA	-0.004	-0.015	-0.004	-0.008	0.023
	(0.007)	(0.012)	(0.009)	(0.006)	(0.043)
Δ Implied Volatility (Euro-bund future continuous call)	0.001	0.003	0.000	-0.000	0.005
	(0.001)	(0.002)	(0.002)	(0.001)	(0.009)
Δ VIX (CBOE SPX Volatility (New) – price index)	-0.005***	0.001	-0.005**	-0.004***	-0.004
	(0.002)	(0.004)	(0.002)	(0.001)	(0.006)
Constant	-0.000	-0.001	0.001	0.000	-0.011
	(0.001)	(0.002)	(0.002)	(0.001)	(0.013)

Standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

Table 3: Regression results for Euro Area

	(1)	(2)	(3)	(4)	(5)
		1st significance		**maximum significance**	
Break date (99% confidence level):		**11dec2007**		**23oct2008**	
	23jun2004 - 12mar2009	pre-break	post-break	pre-break	post-break
overall R^2	0.028	0.030	0.073	0.076	0.231
partial R^2	0.024	0.029	0.064	0.028	0.197
Observations	1232	904	328	1131	101
France Business Confidence overall indicator	0.001	0.001	0.002	-0.004	0.012
	(0.004)	(0.004)	(0.009)	(0.004)	(0.015)
France GDP QoQ	-0.006	-0.003	-0.023	-0.006	0.000
	(0.007)	(0.004)	(0.030)	(0.005)	(0.000)
France Industrial Production MoM SA 2000=100	-0.008**	0.000	-0.027***	-0.003	-0.040**
	(0.004)	(0.003)	(0.010)	(0.003)	(0.019)
France PPI MoM 2000=100	0.002	-0.013**	0.004	-0.003	0.005
	(0.004)	(0.006)	(0.007)	(0.007)	(0.009)
France Unemployment rate SA	-0.000	-0.000	0.000	-0.000	0.000
	(0.004)	(0.003)	(0.000)	(0.003)	(0.000)
France CPI MoM European harmonized NSA	0.005	0.004*	0.001	0.004	-0.025
	(0.004)	(0.002)	(0.013)	(0.003)	(0.069)
Bundesbank Germany Current Account EUR SA	-0.004	-0.006**	0.002	-0.006**	0.016
	(0.004)	(0.003)	(0.010)	(0.003)	(0.026)
Germany HICP MoN 2005=100	-0.001	0.000	-0.028	-0.001	-0.039
	(0.004)	(0.003)	(0.028)	(0.003)	(0.062)
IFO pan Germany business climate 2000=100	0.001	0.002	0.000	-0.001	0.012
	(0.004)	(0.003)	(0.010)	(0.003)	(0.021)
Germany Industrial production MoM SA 2000=100	-0.001	0.000	0.001	0.004	-0.022
	(0.004)	(0.003)	(0.010)	(0.003)	(0.032)
Germany PPI MoM 1995=100	0.003	-0.005	0.009	0.004	-0.014
	(0.004)	(0.004)	(0.008)	(0.003)	(0.016)
Germany Unemployment rate SA	-0.005	-0.003	-0.016	-0.002	-0.074*
	(0.004)	(0.002)	(0.013)	(0.003)	(0.040)
ZEW Germany assessment of current situation	0.009**	0.005	0.016*	0.010***	-0.011
	(0.004)	(0.004)	(0.009)	(0.003)	(0.046)
Italy Business confidence 2000=100	0.001	-0.004	0.008	-0.004	0.018
	(0.004)	(0.003)	(0.010)	(0.003)	(0.016)
Italy HICP MoM NSA 2005=100	0.002	-0.000	0.028	0.000	0.035
	(0.004)	(0.002)	(0.018)	(0.003)	(0.035)
Italy Industrial Production MoM SA 2000=100	-0.002	-0.003	0.010	-0.004	0.083*
	(0.004)	(0.002)	(0.014)	(0.003)	(0.045)
Italy PPI manufacturing MoM 2000=100	0.014***	0.007**	0.028***	0.008**	0.035*
	(0.004)	(0.003)	(0.010)	(0.003)	(0.018)
Italy Real GDP QoQ SA WDA	-0.001	0.005	-0.005	0.003	0.006
	(0.007)	(0.006)	(0.018)	(0.006)	(0.062)
Δ Implied Volatility (Euro-bund future continuous call)	0.000	-0.000	0.001	-0.000	0.004
	(0.001)	(0.001)	(0.002)	(0.001)	(0.004)
Δ VIX (CBOE SPX Volatility (New) – price index)	-0.002*	0.001	-0.002	-0.006***	0.008**
	(0.001)	(0.001)	(0.002)	(0.001)	(0.003)
Constant	-0.001	-0.000	-0.000	-0.000	0.002
	(0.001)	(0.001)	(0.003)	(0.001)	(0.007)

Standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

Table 4: Regression results for the United Kingdom

	(1)	(2)	(3)	(4)	(5)
		1st significance		maximum significance	
Break date (99% confidence level):		**16apr2008**		**09sep2008**	
	23jun2004 - 24mar2009	pre-break	post-break	pre-break	post-break
overall R^2	0.023	0.014	0.058	0.028	0.111
partial R^2	0.022	0.010	0.057	0.020	0.108
Observations	1204	965	239	1066	138
UK Manufacturing PMI Markit survey ticker	-0.003	0.002	-0.007	-0.001	-0.006
	(0.004)	(0.004)	(0.011)	(0.004)	(0.016)
UK industrial production MoM SA	0.007**	0.004	0.013	0.004	0.025*
	(0.003)	(0.002)	(0.009)	(0.002)	(0.014)
UK CPI EU harmonized MoM NSA	0.009*	0.000	0.016	0.004	0.018
	(0.005)	(0.004)	(0.014)	(0.004)	(0.017)
UK retail prices index MoM NSA	-0.000	-0.002	0.013	-0.002	0.025
	(0.005)	(0.004)	(0.017)	(0.004)	(0.027)
UK Nationwide consumer confidence Index SA	0.028***	0.000	0.028	-0.047***	0.057**
	(0.011)	(0.000)	(0.020)	(0.014)	(0.029)
UK unemployment claimant count monthly change SA	-0.005	-0.004	-0.000	-0.003	0.000
	(0.004)	(0.003)	(0.013)	(0.004)	(0.018)
UK claimant count (unemployment) rate SA	-0.000	0.004	-0.012	0.005	-0.017
	(0.004)	(0.003)	(0.016)	(0.003)	(0.023)
BoE official bank rate	-0.005	-0.010*	-0.005	-0.010	-0.005
	(0.004)	(0.006)	(0.007)	(0.006)	(0.009)
UK chained GDP at market prices QoQ	0.006	-0.003	0.021	0.002	0.019
	(0.006)	(0.004)	(0.020)	(0.005)	(0.029)
UK Retail Sales All Retailing	0.005	0.000	0.008	0.001	0.020
	(0.003)	(0.003)	(0.009)	(0.003)	(0.018)
UK PPI Manufactured Products M	-0.001	-0.002	0.002	-0.004*	0.044
	(0.004)	(0.003)	(0.010)	(0.002)	(0.028)
UK Avg Earnings Whole Economy	-0.001	-0.000	-0.006	-0.000	-0.007
	(0.004)	(0.003)	(0.010)	(0.003)	(0.013)
Δ Implied Volatility (Euro-bund future continuous call)	0.001	0.001*	0.000	0.001*	0.001
	(0.001)	(0.001)	(0.002)	(0.001)	(0.004)
Δ VIX (CBOE SPX Volatility (New) – price index)	-0.000	-0.001	-0.000	-0.002**	0.000
	(0.001)	(0.001)	(0.002)	(0.001)	(0.002)
Constant	0.000	0.001**	-0.003	0.001**	-0.006
	(0.001)	(0.001)	(0.003)	(0.001)	(0.006)

Standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

Figure 1: US Inflation Expectations

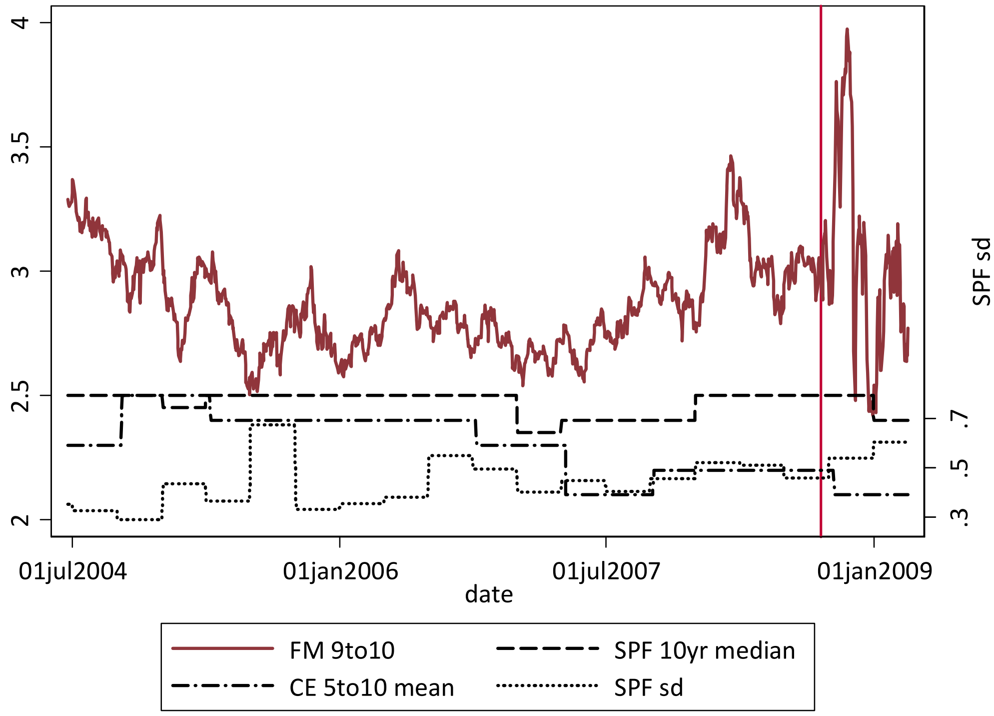


Figure 2: Euro Area Inflation Expectations

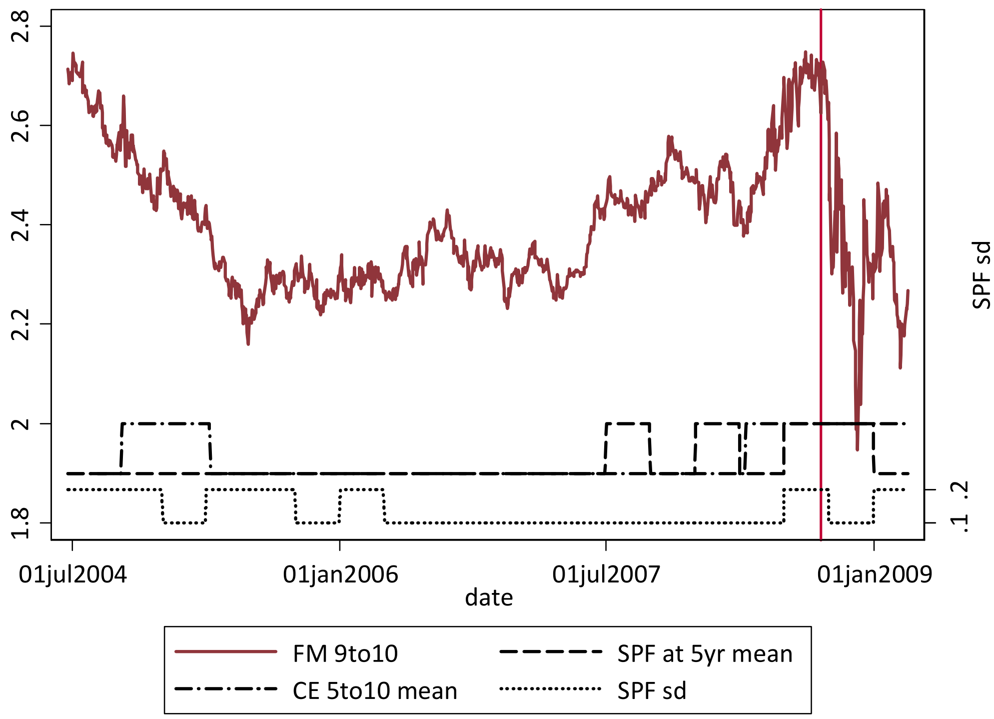


Figure 3: UK Inflation Expectations

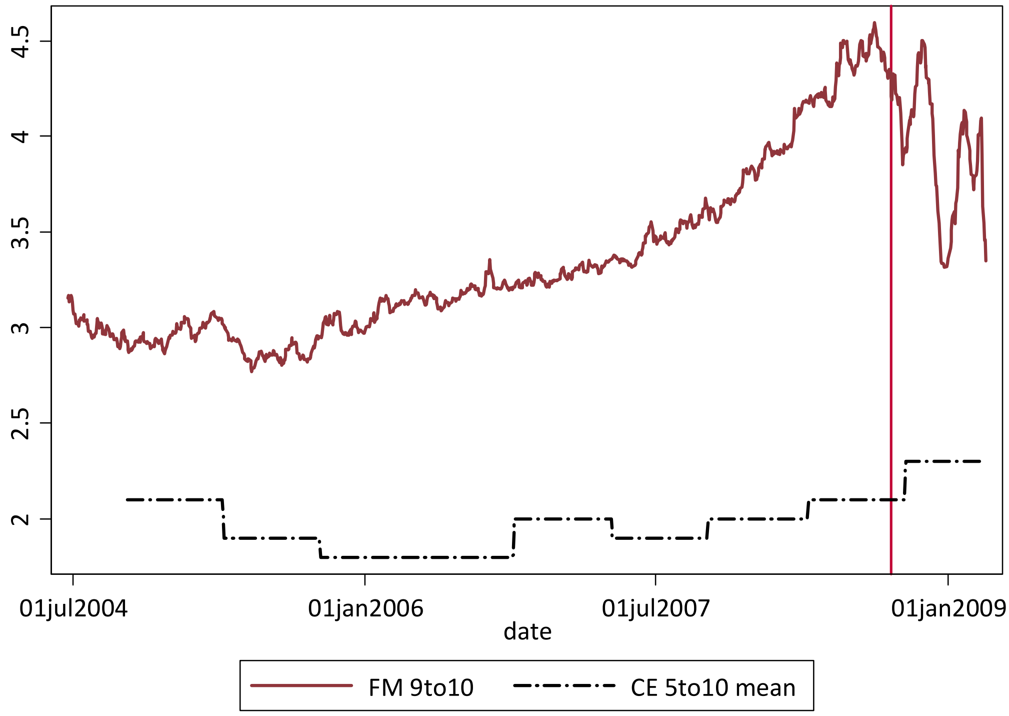


Figure 4: Chow Test results for the United States

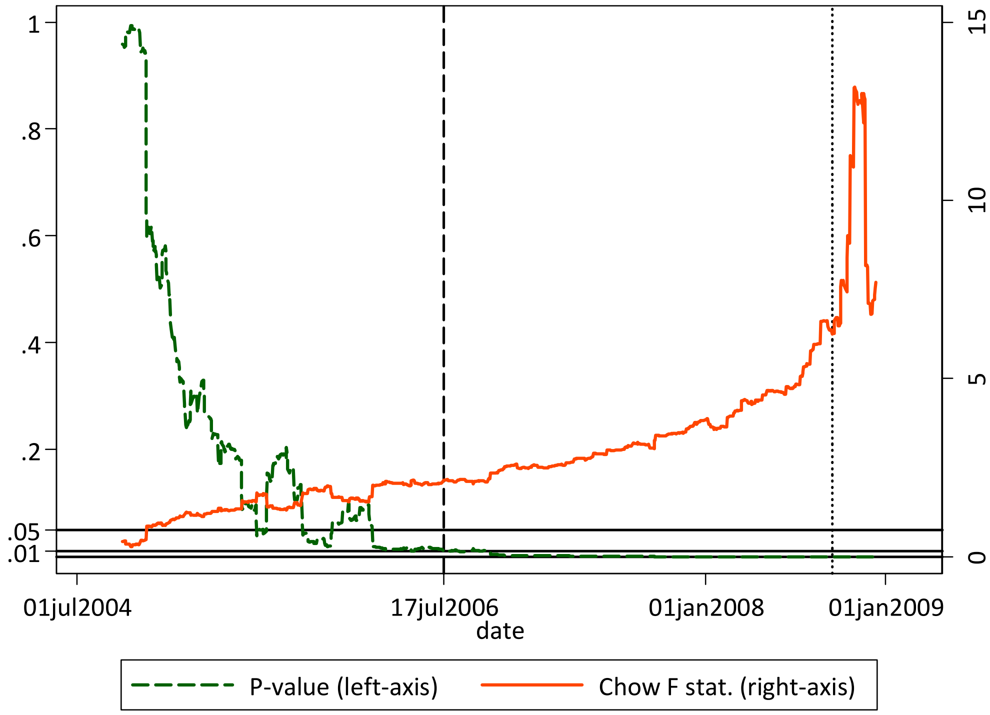


Figure 5: Chow Test results for the euro area

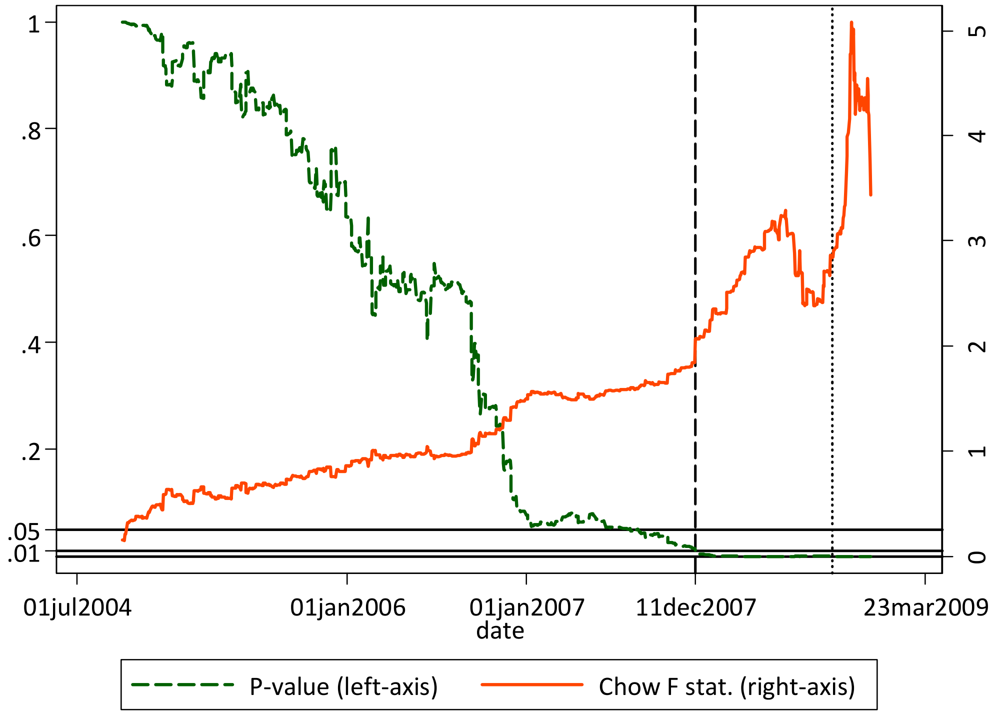
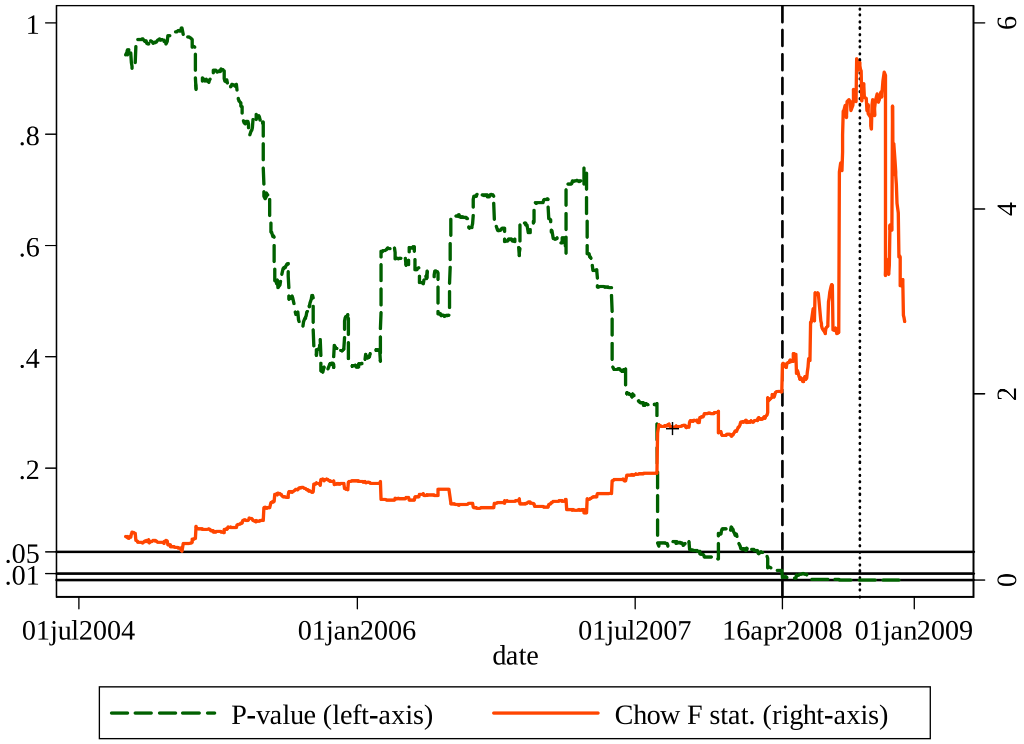


Figure 6: Chow Test results for the United Kingdom



Appendix 1: The Nelson-Siegel method

There is an extensive literature on modeling the term structure of interest rates or inflation swap rates. The seminal paper by Nelson and Siegel (1987) proposes a three-factor model that captures the level, slope and curvature of the yield curve. Empirical studies shows that it fits well for yield curves with different shapes.[24] At each time point t, the model is given as follows:

(A1) $$y_t^{(m)} = \beta_{1,t} + \beta_{2,t}\left(\frac{1-e^{-\lambda m}}{\lambda m}\right) + \beta_{3,t}\left(\frac{1-e^{-\lambda m}}{\lambda m} - e^{-\lambda m}\right) + \varepsilon_t^{(m)}.$$

Here $y_t^{(m)}$ is the yield with maturity m, $(\beta_{1,t}, \beta_{2,t}, \beta_{3,t})$ is the vector of parameters for the three factors, which indicate level, slope and curvature of the yield curve, λ is a decay parameter, which usually assumed to be constant across time, and $\varepsilon_t^{(m)}$ is an error term.

To estimate this model is by no means an easy task. In particular, taking the dynamics of the beta-parameters into account always involves advanced techniques such as Kalman filter (see Diebold et al.(2006). However, by fixing the decay parameter λ, the estimation becomes a simple OLS regression. An example of this approach is in Diebold and Li (2006).

We did not choose a particular value for the decay parameter ex ante. Rather, we allowed the decay parameter to vary on a certain interval, while estimating the Nelson-Siegel model for each specific value in this interval. We then chose the value of the decay parameter at which the total mean squared error is minimized.

When estimating the Nelson-Siegel model for a specific value of λ, the observations are the yields at different maturities. In our data set, the available maturities are 1,2,3,4,5,6,7,8,9,10,12,15,20 and 30

24 For a recent discussion of the empirical performance of the Nelson-Siegel method, see e.g. de Pooter (2007).

years. The observations are therefore concentrated at the shorter end of the yield curve, making it more difficult to accurately represent the shape of the yield curve at longer maturities. In order to balance this effect, we introduced other maturities – such as 11,13,14,16,17,18,19,21,22 years – . To obtain yields on those maturities, we used the bootstrapping method in Fama and Bliss (1987).

We followed this procedure to estimate the yield curves for inflation swaps and interest rate swaps for each day t. Based on the estimated yield curve for inflation swaps, we obtained the yield with 9- and 10-year maturities. These were used to calculate the 10-year ahead forward rates.

Appendix 2: Macroeconomic data releases

United States
ISM Manufacturing PMI SA (value; NAPM)
Personal Consumption Exp. CPI MoM SA
Capacity Utilization % of Total Capacity SA
Conference board consumer confidence SA
Industrial Production MoM SA (rate)
Initial jobless claims SA
Conference board US leading index MoM
Federal Funds Target Rate
New privately owned housing units started by structure total SAAR
Employees on Nonfarm payrolls MoM SA
Adjusted retail & food services SA MoM
Unemployment rate total in labor force SA

United Kingdom
Manufacturing PMI Markit survey ticker
Industrial production MoM SA
CPI EU harmonized MoM NSA
Retail prices index MoM NSA
Nationwide consumer confidence Index
Unemployment claimant count MoM SA
Claimant count (unemployment) rate SA
BoE official bank rate
Chained GDP at market prices QoQ
Retail Sales All Retailing
PPI Manufactured Products M
Avg Earnings Whole Economy

Euro area
France
France Business confidence overall indicator
France GDP QoQ
France Industrial production MoM SA 2000=100
France PPI MoM 2000=100
France Unemployment rate SA
France CPI MoM European harmonized NSA

Germany
Germany Current Account EUR SA
Germany HICP MoN 2005=100
IFO pan Germany business climate 2000=100
Germany Industrial production MoM SA 2000=100
Germany PPI MoM 1995=100
Germany Unemployment rate SA

Italy
Business confidence 2000=100
Italy HICP MoM NSA 2005=100
Italy Industrial Production MoM SA 2000=100
Italy PPI manufacturing MoM 2000=100
Italy Real GDP QoQ SA WDA