Inflation News and Euro-Area Inflation Expectations*

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The protracted period of low inflation in the euro area since 2013 has triggered a significant decline in long-term inflation compensation. This paper re-investigates changes in the sensitivity of inflation compensation to inflation and macroeconomic news and expands existing work in two key dimensions: (i) we analyze all available (advanced) inflation releases for country and euro-area-wide inflation; (ii) we use daily frequency, time-varying sensitivity, and intraday regressions to reach more robust conclusions. Our key findings are twofold. First, timeliness is crucial in inflation markets: it is the early inflation news (flash estimates) which led to revisions in long-term compensation. Second, the anchoring of euro-area inflation expectations has weakened significantly since 2013.

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1. Introduction

The monetary transmission mechanism is most effective when long-term inflation expectations are strongly anchored. The analysis of inflation expectations has therefore become a crucial element of modern monetary policy and a recurrent topic for research. Indeed, most central banks' public statements, speeches (e.g., Draghi 2014, 2019; Powell 2017), and specialized press and market commentary (e.g., *The Economist* 2014, 2017; *Financial Times* 2019) nowadays provide a detailed account on the evolution of inflation expectations.

In the case of the euro area, interest in long-term inflation expectations has become more relevant in recent years. Since early 2013, inflation has remained very low and almost constantly below the level consistent with the European Central Bank's (ECB's) quantitative definition of price stability. Such a protracted period of below-target inflation is unprecedented since the launching of the single monetary policy in 1999. While there is consensus that the ECB managed to attain a significant degree of credibility and long-term inflation expectations were strongly anchored ahead of the global financial crisis (GFC henceforth), and even during the subsequent European debt crisis, the situation over more recent years is less clear. Since well-anchored expectations are key for a sustained return of inflation to the ECB's goal, the anchoring of inflation expectations is a fundamental question for the current monetary policy discussion in the euro area.

Against this background, the goal of this paper is to investigate whether the anchoring of euro-area inflation expectations has weakened in recent years and which macroeconomic releases may help explain that weakening. We assess changes in the anchoring of inflation expectations by the reaction of long-term inflation expectations to macroeconomic data releases. The rationale is that if expectations are well anchored, far-ahead inflation expectations should not be significantly affected by news about the current state of the economy, which should only be informative to update the short-term outlook. As a measure of long-term inflation expectations we use forward inflation-linked (IL) swap rates, in line with existing literature for the euro area and other major economies (e.g., Gürkaynak, Levin, and Swanson 2010; Galati, Poelhekke, and Zhou 2011). The euro-area IL swap market is arguably the most liquid in the world, and IL swap rates offer some important advantages over other indicators of inflation expectations. IL swaps are actively traded at much higher frequency than survey data, which makes them more appropriate to analyze the reaction to macroeconomic news.[1] IL swap rates, however, comprise both the level of inflation expectations and the inflation risk premium associated with the risks surrounding them. Since inflation (and macroeconomic news in general) may trigger changes in the perceived risk surrounding future inflation, the inflation risk premium required by investors is also an important dimension of the anchoring of inflation expectations for monetary policy.

To assess whether the anchoring of euro-area inflation expectations has weakened since 2013, we implement three different but complementary pieces of analysis. First, we use daily regressions over different subsamples, expanding the approach in most previous literature on the anchoring of euro-area inflation expectations in two key aspects. By using data up to 2017, we have a sufficient sample for testing whether the below-target period since 2013 has triggered a change in the reaction of long-term inflation compensation. In order to find the relevant macroeconomic news, we provide a detailed description of the information flow of news about inflation in the specific case of the euro area. The euro-area market for IL products has the unique feature of comprising several different economies. As a result, country-specific data releases add a crucial additional dimension to the information flow, in particular regarding inflation, that needs to be taken into account when assessing market reactions to news. Specifically, advance estimates of inflation—the so-called *flash*

[1] Survey measures of euro-area long-term inflation expectations are available at quarterly frequency from the ECB's Survey of Professional Forecasters and Consensus Economics, while market-based inflation expectations are available at daily or even intraday frequency. In addition to lower frequency, survey measures have several additional shortcomings: panelists have litttle incentive to continuously update or even reveal their true forecasts, there is substantial disagreement among panelists, and survey responses may not be internally consistent (e.g., García and Manzanares 2007) and may have become disconnected from actual inflation over the most recent years (Coibion and Gorodnichenko 2015; Chan, Clark, and Koop 2018).

estimates for German, Spanish, and Italian inflation[2]—have been released since 2005 *ahead* of the euro-area-wide flash estimate, the advance estimate of euro-area-wide inflation, and Harmonised Index of Consumer Prices (HICP henceforth) inflation releases. Failing to account for those country flash estimates would therefore ignore the potentially most relevant pieces of news on euro-area inflation, and may cast doubts on the conclusions about the anchoring of inflation expectations. Indeed, we show that those early inflation releases—and in particular the first one of them, the German flash estimate—do have a statistically significant impact on euro-area short-term inflation compensation measures over our whole sample.

Second, we complement our subsample analysis with time-varying estimation of the response of inflation compensation to inflation news. As most inflation (and macroeconomic) releases only occur once a month, to overcome the short-sample problem inherent in rolling regressions we adapt the empirical approach recently used by Swanson and Williams (2014) to the flow of news on euro-area inflation, and gauge evidence of changes in sensitivity to news from rolling regressions at daily frequency.

Finally, we also provide novel additional evidence on intraday reactions of inflation compensation. Daily-frequency regressions are always subject to influences from additional news. Even a careful selection of release dates and additional control variables may not completely isolate that reaction in the data-rich environment in which financial markets nowadays operate. Intraday data on euro-area inflation compensation measures are unfortunately only available since late 2008, but we can cross-check their reaction to news since 2013 and compare it with the previous years using intraday trading evidence.

Our results point to a deterioration in the anchoring of euro-area long-term inflation expectations since 2013. Daily-frequency regressions corroborate existing findings of little sensitivity of long-term inflation compensation to news releases before 2013. In contrast, we find evidence of a significantly stronger reaction of long-term inflation compensation from 2013. Importantly, such statistically

[2]A flash estimate for HICP inflation in France has also been released since early 2016. However, as a crucial dimension of our analysis is the comparison with earlier periods, we exclude it from our analysis.

significant reactions mainly refer to the release of flash estimates at the country level, especially the German, and also the Spanish, flash estimates, which are, respectively, the first two pieces of news about each month's inflation in the euro area. These findings are robust to controlling for a wide range of additional market and macroeconomic information.

Empirical evidence allowing for time-varying sensitivity to news also points to a significantly higher reaction of long-term inflation compensation to inflation news since the euro-area inflation rate has remained below target. Moreover, while previous periods of high sensitivity—for example, over the spring of 2008 when surging oil prices pushed actual inflation to the highest levels in euro-area history—were short-lived, the responsiveness of long-term inflation compensation to inflation news has remained stubbornly high since mid-2013, even well after the launching of the ECB's sovereign bond purchases program in 2015.

Novel analysis using intraday data provides further support for a weakening in the anchoring of euro-area long-term inflation expectations in the later part of our sample. Between 2013 and 2017, we find statistically significant reactions to the release of the flash estimates for Germany and Spain, as well as of the Spanish HICP. While other variables—in particular, economic activity and monetary policy surprises—are also significant, inflation surprises display a much higher explanatory power. Over a shorter time window of 15 minutes, the reaction is somewhat more muted but nonetheless statistically significant for the German flash and also the euro-area flash estimate.

This paper belongs to a large literature investigating the reaction of financial markets to incoming news, and contributes to the analysis of euro-area long-term inflation compensation in three specific aspects: by employing a larger set of news than most existing papers and emphasizing the crucial role of the timing of inflation news (particularly the early releases of inflation at the country level), by combining three different pieces of empirical analysis (daily frequency, time-varying sensitivity, and for the first time intraday regressions) and by using a larger sample, January 2005–March 2017. Those three pieces of analysis provide substantial robustness to our key finding of a significant increase in the reaction of euro-area long-term inflation compensation to news in the latter part of our sample.

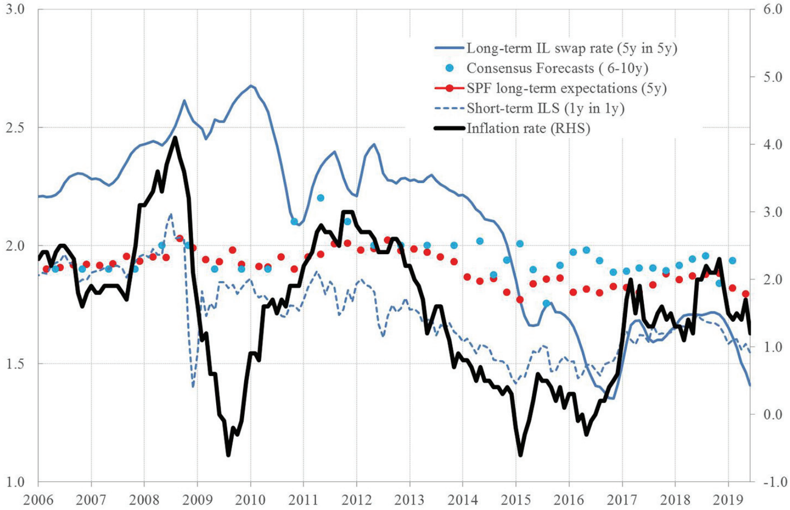
The remainder of the paper is organized as follows. Section 2 reviews the literature on the reaction of inflation compensation to news and motivates the case for further investigation after 2013. Section 3 discusses the data we employ in our analysis. Specifically, we explain in detail the flow of macroeconomic news in the euro area—in particular regarding inflation—and the characteristics of inflation compensation in the euro area. Section 4 reports our empirical results. The main findings from daily frequency, time-varying sensitivity, and intraday data regressions are described in detail. Section 5 contains an additional discussion and robustness checks of some of our findings, and places them in the context of the other results employing alternative indicators of inflation expectations. Finally, section 6 concludes.

2. Related Literature and the Case for Reassessing Euro-Area Inflation Expectations

This paper belongs to an extensive literature investigating the reaction of financial markets to incoming news, from the early works of Dwyer and Hafer (1989), Fleming and Remolona (1997), and Kuttner (2001), among others, on U.S. data. The significant development of financial markets for inflation-linked products and the lumpy manner in which inflation and other macroeconomic data are released provide a great opportunity for researchers to carry out those studies on inflation expectations, as the search for robust evidence of causality is significantly eased using (high-frequency) financial market data (e.g., Gürkaynak and Wright 2013). Indeed, the sensitivity of market-based indicators of inflation expectations, either based on bond-market instruments or from the inflation-linked (IL) swap market, to macroeconomic releases has already been studied using euro-area (and international) data following Gürkaynak, Sack, and Swanson (2005a), among others. This section provides a selective overview of that literature, focusing on the different questions it tried to address over time.

To help put those contributions in context, figure 1 depicts actual inflation dynamics in the euro area (black line), together with some key indicators of euro-area inflation expectations. Two survey measures of long-term inflation expectations, from Consensus Economics (6 to 10 years ahead) and from the ECB's Survey of Professional

Figure 1. Actual Inflation, Market-Based and Survey Long-Term Inflation Expectations



Notes: The figure shows developments in indicators of long-term inflation expectations (both market and survey based) and the actual year-on-year rates of inflation in the euro area. Specifically, it depicts two benchmark forward IL swap rates, for short-term (1-year forward IL swap rate in 1 year, dashed gray/blue line) and long-term expectations (5-year forward IL swap rate in 5 years, gray/blue line); 6–10 years Consensus Forecasts (gray/blue dots) and 5-year-ahead expectations from the ECB's Survey of Professional Forecasters (black/red dots), and the realized inflation rate (thick black line). (Figures in color can be seen in the online version of the paper at http://www.ijcb.org.)

Forecasters (SPF) (5 years ahead), are denoted by red and blue dots. Short-term (one-year in one year) and long-term (five-year in five years) forward IL swap rates are denoted by a dashed and a continuous gray/blue line, respectively, reflecting their availability at higher frequency.[3]

When interpreting differences between survey and market-based inflation expectations, it has to be borne in mind that IL swap rates, and break-even inflation rates (BEIRs), incorporate inflation risk

[3]Given the relatively limited issuance of IL bonds in the euro area—only France, Italy, Germany, Spain, and Greece have issued those bonds—the IL swap market has arguably become the most developed in the world, and can therefore provide a cleaner measure of inflation compensation than bond-based BEIRs.

premiums and should better be interpreted as the overall *inflation compensation* requested by investors, also including risks surrounding future inflation. Changes in inflation compensation could therefore reflect changes in the level of expected inflation, changes in the perceived risks about future inflation, or a combination of both. From a central bank's perspective, both components are of relevance. A credible commitment to price stability should anchor the level of expected inflation to its policy objective. In turn, the requested risk compensation provides relevant information about how firmly inflation expectations may be anchored. Inflation compensation measures are therefore a very rich source of information for testing the reaction to macroeconomic news.[4] In addition, survey measures of inflation expectations have in contrast remained relatively more stable in the euro area, but a decline away from (below but close to) the 2 percent level they exhibited since the early 2000s can also be observed at the end of our sample (e.g., Draghi 2019).

Before the GFC, the sensitivity of inflation expectations to macroeconomic news was mainly explored as a measure of the degree of anchoring of inflation expectations in the euro area (e.g., Coffinet and Frappa 2008; Beechey, Johannsen, and Levin 2011; Ehrmann et al. 2011).[5] Overall, no evidence of a significant reaction of long-term inflation compensation measures to news has been found for the euro area. Moreover, euro-area inflation expectations have been often found to be more strongly anchored than in other countries like the United States, the United Kingdom, or Sweden (e.g., Gürkaynak, Levin, and Swanson 2010).[6]

[4]We focus here on the literature looking at the reaction of inflation compensation measures; section 5 discusses alternative measures of inflation expectations.

[5]Coffinet and Frappa (2008) test the five-year forward IL rate in five years (our benchmark measure) against the euro-area flash estimate and country-specific HICP releases as well as various other macro announcements. Beechey, Johannsen, and Levin (2011) test for the one-year forward IL rate in nine years and similarly use country-specific HICP and other macroeconomic releases, but they do not consider flash estimates. Ehrmann et al. (2011) instead compare the mean and variance of one-year forward *nominal interest rates* in nine years in the nominal yield curves of euro-area countries prior to and during the single-currency period, and interpret the decline and convergence of both moments across countries as signaling a stronger anchoring of inflation expectations.

[6]Research on other countries (see De Pooter et al. 2014 for Brazil, Chile, and Mexico and Gürkaynak et al. 2007 for Canada, Chile, and the United States)

Assessing whether the anchoring of euro-area inflation expectations survived the GFC was the main question for a new wave of research (e.g., Galati, Poelhekke, and Zhou 2011, Autrup and Grothe 2014).[7] Those studies explicitly acknowledge that the analysis in the aftermath of the Lehman collapse years is subject to some uncertainty given the extreme market volatility and limited liquidity during the crisis, and added additional controls (e.g., VIX, bond market volatility, oil prices) to the regressions to account for them. Overall, their findings suggest that the anchoring of euro-area long-term inflation expectations remained strong after the GFC, despite the significant rise in long-term forward inflation compensation at the time.

In light of the weak inflation since 2013 and the launching of additional unconventional monetary policy (UMP) measures in 2015, interest in the reaction of inflation compensation to macroeconomic news has gained momentum again. Long-term forward inflation compensation has declined significantly in the euro area since actual inflation rates started surprising negatively in 2013. Inflation compensation declined in several markets, but the correction was particularly strong in the euro area, and attracted substantial attention among policymakers (Draghi 2014, 2019), as well as in specialized press and market commentary (e.g., *The Economist* 2014; *Financial Times* 2019).

To assess the reactions to news in the low inflation environment since 2013, Miccoli and Neri (2015) look at the reaction of IL swap spot and forward rates to the euro-area HICP releases at different maturities separately, comparing the average of 10 business days prior to and after the date of release. Speck (2016) instead jointly regresses the two-year spot, the three-year forward in two years, and

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also finds evidence of the crucial role of inflation targeting in the anchoring of inflation expectations.

[7] Galati, Poelhekke, and Zhou (2011) use a sample for the United States, the euro area, and the United Kingdom from June 2004 until March 2009, and find that the euro-area and the U.K. inflation expectations seem to be more stable than those in the United States. To gauge the reaction of inflation compensation, they only use country-specific HICP releases as inflation news and found that the VIX has a statistically significant impact on expectations. Autrup and Grothe (2014) use data up to 2012 and find the euro-area expectations to be more stable than the U.S. ones. The paper also shows that bond-based inflation compensation, in contrast to IL swap rates, is statistically influenced by a liquidity risk premium.

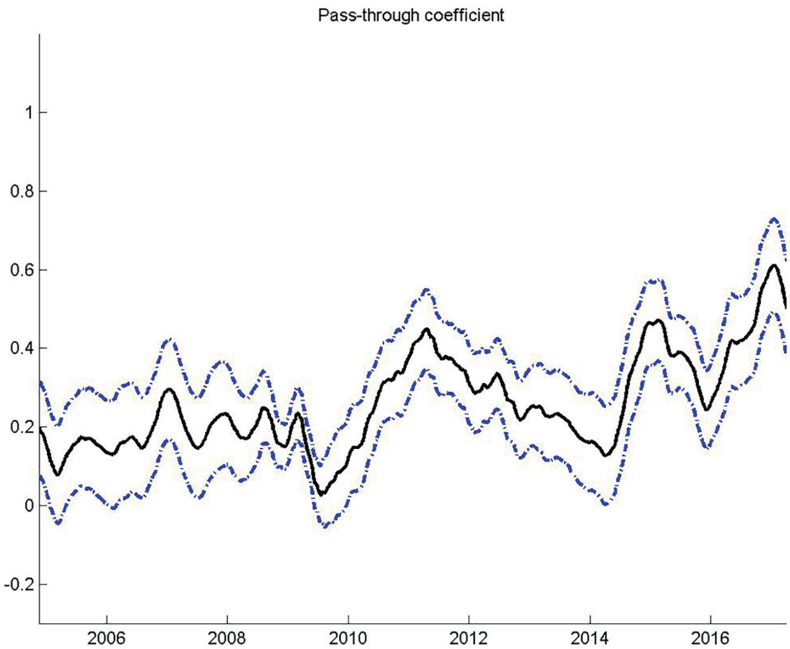
the five-year forward in five years, and tests the time variation in their sensitivity to news using country HICP estimates and a few confidence indicators. Both studies find a somewhat stronger sensitivity of medium-to-long-term inflation compensation to news in the latter part of their samples, but they disagree on whether that evidence is sufficiently conclusive for the presence of a significant de-anchoring of inflation expectations or just a longer-than-normal correction from low inflation levels. Other contributions (e.g., Fracasso and Probo 2017, Nautz, Pagenhardt, and Strohsal 2017) however find evidence of a de-anchoring of euro-area inflation expectations as early as late 2011 using structural break tests on the reaction of long-term inflation compensation to the euro-area HICP and some other macroeconomic data releases.

The lack of conclusive evidence on whether the anchoring of inflation expectations has weakened in the euro area since 2013 may reflect several potential shortcomings of existing literature. First, a challenge for the more recent research is that, for a weaker anchoring of euro-area inflation expectations to be economically meaningful, it has to be sustained over a certain period of time. A minimum sample is therefore required to reach robust conclusions. Figure 1 shows that the decline in long-term inflation compensation took place over a relatively long period between 2013 and most of 2016. Moreover, the decline was in line with that in actual inflation in a way that was not common before, suggesting there were important changes in the pricing of inflation compensation.

Additional evidence on the potential deterioration on the anchoring of inflation expectations can be found by assessing the pass-through from short-term to longer-term inflation compensation. If long-term inflation expectations are well anchored, revisions in short-term inflation expectations—reacting, for example, to actual inflation readings—should not trigger strong revisions in longer-term inflation expectations. Indeed, evidence from a standard linear regression[8] points to a much stronger impact of changes in

[8]Figure 2 reports the estimates of the posterior mean for β_t and standard quantiles in $\Delta y_t^{5y5y} = \alpha + \beta_t \Delta y_t^{1y1y} + \epsilon_t$, where β_t evolves following a random walk. The framework allows for stochastic volatility to account for potential changes in market conditions and varying volatility as suggested by Jochmann, Koop, and Potter (2010) and Chan (2013).

Figure 2. Pass-Through from Short- to Long-Term Inflation Expectations



Notes: The figure shows the estimates of the posterior median for β_t and the 16th and 84th percentiles of the confidence interval in the regression $\Delta y_t^{5y5y} = \alpha_t + \beta_t \Delta y_t^{1y1y} + \epsilon_t$, where β_t evolves following a random walk and allowing the stochastic volatility to account for potential changes in market conditions and time-varying volatility following Chan (2013) and Jochmann, Koop, and Potter (2019). For presentational purposes, results are smoothed over a 100-day window.

short-term expectations on longer-term expectations in the latter part of our extended sample (see figure 2).

The goal of this paper is to assess the extent to which the response to inflation news was responsible for that decline, and to provide robust evidence on whether the observed reactions have been different from those observed in the past. While our analysis is close in spirit to the literature reviewed above, this paper also expands existing evidence to help reach more robust conclusions. Specifically, we will analyze in greater detail the unique flow of inflation releases in the euro area, expanding the set of macroeconomic news—particularly inflation releases—with respect to most existing literature, and paying particular attention to their timing.

Evidence that news released earlier in time tends to have a greater impact on financial assets than news about the same variable released later has been supported by early event studies (e.g., Fleming and Remolona 1997; Andersen et al. 2003). More recently Hess and Niessen (2010) and Gilbert et al. (2017), among others, have explored the tradeoffs between timeliness and precision of macroeconomic news in asset price responses. Both studies find that financial markets favor timeliness to more comprehensive macroeconomic news.[9]

The timing of macroeconomic releases has not been emphasized in the literature on the reaction of inflation expectations to news. In most countries—for example, in the large number of studies employing U.S. data—inflation news mostly comes from two variables, namely CPI and PPI indexes. The relevant flow of inflation news in the euro area is instead much richer. We discuss in detail the flow of inflation releases in the euro area and how they can be incorporated into event studies on the euro area in the next section. In addition, we also provide novel evidence of euro-area inflation expectations on the reaction to news using intraday frequency data.

3. News and Inflation Expectations in the Euro Area

A central argument of this paper is that a meaningful event-study analysis on euro-area data requires taking into account some specific

[9]Gilbert et al. (2017) estimate the intrinsic value of a macroeconomic announcement (defined as the ability to nowcast gross domestic product (GDP) growth, inflation, and the federal funds target rate), in terms of relation to fundamentals, timing, and revision noise, and find that timing is the most important characteristic in explaining asset price impact. Similarly, Hess and Niessen (2010) test the tradeoff between early availability and information quality of news on two German economic indicators—the ZEW, with an earlier release, and the IFO indicator, comprising more information—and report evidence of a larger ZEW impact due to its earlier release. Similarly, Hess (2004) studies the impact of a large and diverse set of macroeconomic announcements using high-frequency analysis and finds that the response to announcements that are released in month $m{+}1$ is significantly stronger than the impact of announcements released a month later $(m + 2)$.

characteristics of euro-area macroeconomic data releases and financial market indicators. We therefore discuss them in detail before embarking on our empirical analysis.

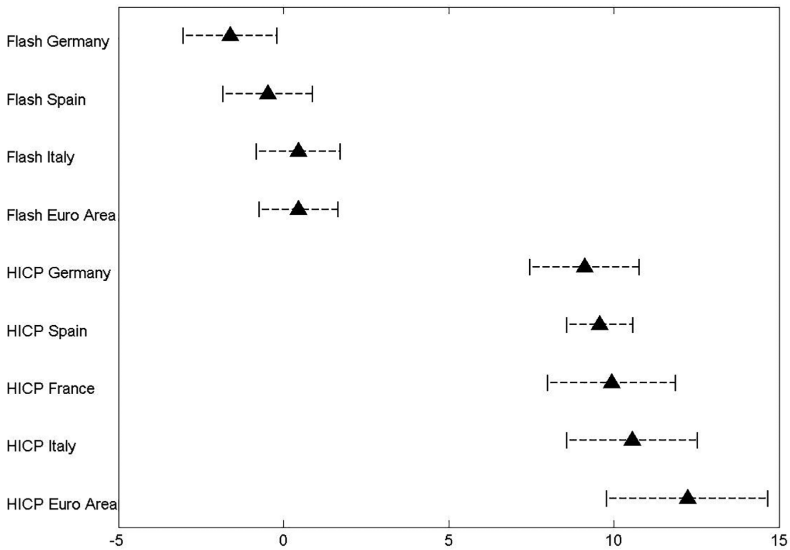
3.1 The Flow of Euro-Area Releases

Financial markets nowadays operate in a data-rich environment (Bernanke and Boivin 2003), and indeed evidence suggests that market participants do monitor a large number of macroeconomic indicators (e.g., Bartsch et al. 2014; Swanson and Williams 2014). Given its unique characteristics as a monetary union comprising a relatively large number of countries, the presence of a data-rich environment is even more extreme in the case of euro-area financial markets.

In addition to the amount of relevant news, the timing of the macroeconomic releases is another important aspect to take into account. In the euro area, particularly regarding inflation news but also some other macroeconomic series like the GDP, there are a number of key country-specific macroeconomic releases that precede the release of euro-area aggregates. A key motivation of our analysis is that failing to account for those country-specific releases can impair a proper understanding of the reaction of inflation expectations to macroeconomic news and bias the empirical results.

Since we will focus on the reaction of long-term inflation expectations, we will pay particular attention to the news surrounding the releases of the HICP. There are two kinds on HICP releases in the euro area. *Flash* releases are *advance* estimates regularly issued by statistical offices. Eurostat releases a flash estimate for the euro-area-wide HICP, but flash estimates are also issued for Germany, Spain, and Italy (and, since 2016, also France) by the corresponding national statistical institutes. Those country flash estimates are normally limited to a figure for the year-on-year inflation rate for the current month and do not contain a breakdown of the different subcomponents (e.g., service or non-energy industrial goods inflation rates). The euro-area flash estimate, for example, has only included a breakdown of main HICP subcomponents since late 2012. The relevance of the flash estimates instead stems from the fact that they are released well ahead of (about two to three weeks before) the full release for the corresponding country or the euro-area-wide HICP.

Figure 3. Schedule of Flash and HICP Inflation Data Releases



Notes: The figure shows the average calendar timing of flash estimates and HICP data releases for European countries (Germany, Spain, France, and Italy) and euro-area-wide inflation. The y-axis lists the inflation releases; the x-axis indicates the average release days (triangle), and whiskers reflect the distribution over our sample (January 2005 to March 2017). 0 denotes the last business day of the month to which the release refers; 5 denotes five business days after the last business day of the month, etc. For example, the flash HICP for Germany was released on average 2 days before the last business day of the corresponding month and the HICP release for Italy on average more than 10 working days into the following month.

Figure 3 illustrates the flow of inflation releases in the euro area. The x-axis indicates the number of business days between the release dates of the monthly flash estimates and the corresponding HICP with respect to the last business day of the reference month m. A zero on the x-axis therefore indicates that a flash estimate was released on the last day of the month to which it refers, negative numbers indicate the number of days before the end of the month (the most common situation), while positive numbers point to days into the month following the one to which the flash release refers. More formally, $flash_{m,t}$ provides an estimate of $HICP_{m,t}$ for the corresponding country or the euro area in month m.

Dates of the inflation releases nonetheless vary from month to month across countries, depending on data collection and processing. A triangle in figure 3 indicates the average business day of the inflation data release, and the whiskers represent the distribution in the release dates. For example, the German flash estimate is released on average two business days within the reference month, the earliest among the flashes (the Spanish flash is released one day before the end of the month, while the Italian and the euro-area-wide flash tend to be released one day into the following month). All flash estimates are nonetheless released ahead of the corresponding HICP inflation releases, which take place in the month following the reference period: 9 days for the German, 10 for the Spanish and the French, 11 for the Italian, and about 13 days for the euro area after the end of the reference month.

Besides inflation data releases, we also consider data releases for other major euro-area macroeconomic series to provide an ample coverage of factors shaping the macroeconomic situation in the euro area and therefore potentially having an impact on inflation expectations. Specifically, we consider producer price index (PPI), GDP, purchasing manager's index (PMI), consumer confidence, industrial confidence, and production (new orders, unemployment rate, growth of the monetary aggregate (M3), retail sales, and trade balance). A complete list of the series can be found in table 1.

Financial markets are forward-looking and, under the efficient market hypothesis, all available and expected information should be priced in, so only the "unexpected" or "news" component of macroeconomic data releases should lead to changes in inflation compensation. We measure the news component of each release using the macroeconomic expectations collected by Bloomberg L.P. from a selection of professional economists until soon before the release takes place. In our econometric analysis, we use euro-area-wide and some key country releases for which Bloomberg has collected expectations over a large part of our sample.

We compute the surprise component for each data release as the difference between the actual values of the official release and the Bloomberg survey expectations for each variable k for reference period t. Using the surprise component of the releases also removes any issues of endogeneity arising from inflation expectations feeding back to the macroeconomy, because any such effects, to the

Table 1. Long-Term Inflation Compensation (5y5y, Daily): Joint Linear Regression $\Delta y_t = \alpha + \beta X_t + \varepsilon_t$

	Full Sample		Pre-disinflation		Below Target	
	β	t-stats	β	t-stats	β	t-stats
Flash						
Germany	0.57**	(2.19)	0.34	(0.89)	1.08***	(4.91)
Spain	0.39*	(1.85)	0.62**	(2.09)	−0.04	(−0.16)
Italy	−0.06	(−0.30)	−0.08	(−0.30)	−0.17	(−0.56)
Euro Area	0.10	(0.38)	0.15	(0.38)	0.00	(0.00)
HICP						
Germany	0.10	(0.54)	0.05	(0.26)	0.43	(1.15)
Spain	0.23	(0.99)	0.24	(0.90)	0.11	(0.33)
France	−0.08	(−0.58)	−0.13	(−0.86)	0.14	(0.40)
Italy	0.19	(1.06)	0.38*	(1.81)	−0.40	(−1.27)
Euro Area	0.06	(0.43)	0.08	(0.45)	0.04	(0.24)
PPI	−0.05	(−0.25)	0.00	(−0.02)	−0.14	(−0.55)
GDP	0.40***	(2.69)	0.45***	(2.80)	0.19	(0.60)
PMI	−0.03	(−0.15)	−0.03	(−0.16)	−0.02	(−0.04)
Consumer Confidence	0.16	(1.15)	0.12	(0.93)	0.36	(0.83)
Industrial Confidence	−0.13	(−0.63)	−0.17	(−0.76)	−0.03	(−0.09)
Industrial Production	0.47***	(2.65)	0.48**	(2.27)	0.30	(1.27)
New Orders	−0.19	(−0.79)	−0.17	(−0.71)	—	—
Unemployment Rate	−0.28	(−1.16)	−0.19	(−0.58)	−0.55**	(−2.01)

(*continued*)

Table 1. (Continued)

	Full Sample		Pre-disinflation		Below Target	
	β	t-stats	β	t-stats	β	t-stats
M3	−0.13	(−0.80)	−0.14	(−0.81)	−0.06	(−0.17)
MRO	—	—	—	—	0.40***	(3.42)
Retail Sales	0.13	(0.81)	0.11	(0.62)	—	—
Trade Balance	−0.03	(−0.13)	−0.04	(−0.13)	−0.01	(−0.03)
Mon. Policy Shock	4.59	(0.63)	7.99	(0.93)	−13.33	(−0.80)
CB Info. Shock	25.75**	(2.18)	24.97*	(1.85)	13.52	(0.69)
Observations	1,767		1,209		525	
$\bar{R}^2$	0.02		0.00		0.03	
R^2	0.03		0.03		0.07	
$H_0 : \beta = 0$ (p-value)	0.00		0.06		0.00	
Chow Test (p-value)	0					

Notes: The table reports estimated β coefficients from regressions on days with data releases over three periods: from January 2005 to March 2017 (full sample), from January 2005 to December 2012 (pre-disinflation), and from January 2013 to March 2017 (below target), using surprises for all variables in a joint regression. Dependent variable is one-day change, $\Delta y_t \equiv y_t - y_{t-1}$, in long-term inflation compensation (5y5y, in basis points). The surprise component of macroeconomic releases, $\mathbf{X}_t$, is normalized by their historical standard deviations; coefficients there represent a basis point per standard deviation response. Heteroskedasticity-consistent t-statistics are in parentheses; ***, **, and * denote a 1, 5, and 10 percent level of significance, respectively. $H_0 : \beta = 0$ p-value is for the test that all elements of β are zero (joint significance test). Chow test (p-value) is for the H_0 of parameter constancy between the pre-disinflation and below-target period. (—) indicates an omission either due to the discontinuity of the expectations collection (e.g., new orders) or for lack of surprises over most of the period of the regression (e.g., MRO).

extent that they are predictable, should be already incorporated into market expectations.

Among the set of variables, monetary policy surprises deserve particular attention. Bloomberg L.P. also collects market expectations about the ECB's policy rates, but as the ECB was highly predictable, there were very few monetary policy surprises, particularly before 2013.[10] We nonetheless include them in our set of surprises. Recent literature has however emphasized that central banks' announcements reveal information not just about policy rates but also about the central bank's assessment of the economic outlook. Jarociński and Karadi (2018) uses a wider range of market price reactions to identify two structural shocks embodied in central bank announcements—a purged monetary policy shock and a central bank information shock—and find evidence of a significant impact of information shocks on macroeconomic data, including on long-term inflation expectations. Moreover, central bank information shocks appear to be more frequent in the ECB's announcements than in the U.S. Federal Reserve's Federal Open Market Committee's announcements, for example. In order to assess the response of long-term inflation compensation measures on inflation surprises, we also add those monetary policy and central bank information shocks to our set of surprises.

3.2 Measuring Inflation Expectations

Our empirical analysis will focus on long-term forward inflation compensation. Forward IL swap rates provide a very useful means of interpreting markets' inflation compensation at medium-to-long-term horizons. Standard macroeconomic models predict that inflation should return to its steady state between 5 and 10 years after a typical shock (e.g., Gürkaynak, Sack, and Swanson 2005b). We will focus our analysis on the five-year forward IL swap rate five years ahead, which has become the most widely used measure to assess

[10]The expectation collection was also extended beyond policy rates following the introduction of some UMP measures, but they are available over too short a period for our analysis.

developments in euro-area long-term inflation expectations for policy (e.g., Draghi 2014) and research purposes.[11]

While euro-area IL swap contracts have been very actively traded since 2004 over maturities from 1 to 30 years, available evidence and market intelligence suggest that the 5- and 10-year maturities have tended to concentrate around 50 percent of all trade activity in the market.[12] Using swap rates also alleviates difficulties in controlling for the sovereign and the liquidity risk embodied in the prices of IL bonds since the GFC, which includes the main period of interest in our analysis.

4. Assessing the Anchoring of Inflation Expectations

This section reports the main findings of our empirical analysis. To reach robust conclusions, we investigate the response of long-term inflation compensation to inflation and macroeconomic news by means of three different but complementary pieces of analysis.

4.1 Evidence from Daily Regressions

Table 1 reports estimation results from the standard linear regression:

$$\Delta y_t = \alpha + \beta \mathbf{X}_t + \epsilon_t, \tag{1}$$

[11] Formally, the price of a spot zero-coupon swap with a 10-year maturity, s_t^{10y}, reflects the average inflation compensation over the next 10 years. It is similar for the five-year spot rate, s_t^{5y}. In contrast, by construction, 5-year forward inflation compensation 5 years ahead, f_t^{5y5y}, reflects the average inflation compensation priced in between 5 and 10 years ahead, a medium-to-long-term period that captures well the movements in inflation compensation we are interested in. Formally, the long-term forward IL rates implicit in the term structure of IL swap rates can be calculated from the 5- and 10-year spot rates as $f_t^{5y5y} = \left[\frac{(1+s_t^{10y})^{10}}{(1+s_t^{5y})^5}\right]^{\frac{1}{10-5}} - 1$.

[12] Evidence from other indicators used in the related literature, like the 1-year forward in 9 years for example, involves using 9- and 10-year spot IL swap rates, with the former having only around 7 percent of trading activity of the latter. The forward rate should therefore be interpreted with caution. As a robustness check and link to previous literature, the appendix nonetheless shows that our main findings are robust to different measures of long-term inflation compensation (see tables A.1 and A.2 in the appendix).

where t indexes days, and $\Delta y_t \equiv y_t - y_{t-1}$ is the change in the inflation swap rate over a day; $\mathbf{X}_t$ is a vector of macroeconomic news in data releases, including all macroeconomic releases and the monetary policy surprises (as described above), and ϵ_t denotes the regression residual capturing the influence of other information or financial factors on inflation compensation that day. Surprises are normalized by its standard deviation, and the estimated coefficients can be interpreted as basis points per standard deviation in the release.

The estimated impact coefficients and their standard deviations are reported over three different subsamples to gain some insights about their changes over time. The first two columns report results over the full sample 2005–17. The releases of the German and the Spanish flash inflation estimates as well as the euro-area activity (GDP and industrial production) and the central bank information shock are statistically significant at standard significance levels. Inflation news in particular leads to a change in long-term inflation compensation of around 0.6 and 0.4 basis point on average on the day of their release. These results underscore the importance of accounting for the timing of the inflation news to assess the reaction of long-term inflation compensation: not only do flashes tend to trigger larger (and statistically significant) reactions than the HICP releases, but also the strongest reactions are those for the flash releases that take place earlier in time and are therefore the earliest pieces of news about euro-area inflation.

The remaining columns of table 1 report results over two subperiods: the pre-disinflation period, 2005–12, and the period over which actual euro-area inflation was consistently below target, 2013–17, the main period of interest here. Results point to some important changes in the news that triggered statistically significant reactions in long-term inflation compensation. Over the pre-disinflation period, the economic activity news (GDP and industrial production) and the monetary policy information shock are identified as the most relevant non-inflation news ones. Regarding inflation news, statistically significant reactions on inflation expectations were triggered by the Spanish flash and the Italian HICP before 2013, possibly reflecting the market focus on periphery countries during the European debt crisis period.

Over the below-target period, however, the relevant surprises are quite different. Among inflation news, it is the German flash estimate which leads to large reactions in long-term inflation compensation (its impact coefficient rises to 1.08, more than three times its value before 2013) and becomes strongly significant. In contrast, market reactions to economic activity news seemed to shift from the GDP and industrial production towards the unemployment rate. In addition, monetary policy shocks continue to be significant news, now captured by unexpected changes in the main refinancing rate, whose level changes did trigger significant surprises among market participants since 2013.[13]

Although the estimated responses per standard deviation to news reported in table 1 may appear to be limited and the R^2 low, it is important to bear in mind that daily changes in euro-area long-term inflation compensation have been, on average, relatively limited (around 2 basis points). Indeed, in normal circumstances—with a strong anchoring of inflation expectations—one should observe that macroeconomic data releases trigger no change at all in euro-area long-term inflation compensation. Therefore, low coefficient estimates are common in this type of analysis given the low signal-to-noise ratio of any single monthly data release for the true underlying state of economic activity and inflation, particularly over long horizons (Swanson and Williams 2014). The statistical significance of some coefficients in table 1, particularly inflation news, instead suggests that regressions like (1) are extremely informative to gauge the sensitivity of long-term inflation expectations to economic news and the anchoring of inflation expectations.

The comparison of relevant surprises in the pre-disinflation period (before 2013) and in the below-target period (after 2013) provides two main insights. First, while there are some differences between the significant news in the two periods, the overall joint test of no significance cannot be rejected before 2013 (p-value of 0.06 for the pre-disinflation period) but is strongly rejected for the below-target period (p-value of 0.00). Second, the Chow forecast

[13]The standard joint regression commonly used in the related literature, including table 1 here, associates a zero with the days on which there are no releases for the variable. This standard practice implies no distinction to the value associated to a fully anticipated release, that is a no-surprise in the announcement. This is particularly relevant for the monetary policy surprises regarding the ECB's main refinancing rate and its strong predictability before unconventional monetary policy measures were introduced (see Hartmann and Smets 2018).

test also rejects the null hypothesis of overall stability of the coefficients over the two periods.[14] The remainder of the section provides additional empirical evidence in support of the robustness of those conclusions.[15]

In our sample, releases of individual macroeconomic series only occur once a month. Within a joint regression for any given news surprise, all other news is set to zero, though strictly speaking no news release has occurred for any other variable. To gain additional evidence on the main drivers of changes in inflation compensation, we also estimate daily regressions using the surprises for individual variables:

$$\Delta y_t = \alpha + \beta x_t + \epsilon_t, \tag{2}$$

where x_t denotes the vector of releases for a single variable. This type of analysis allows for regression estimates to be based only on data for those days on which a release for each variable takes place therefore complementing joint regressions above (e.g., Kilian and Vega 2011).

Table 2 corroborates the results from the joint variable regressions. News on German and Spanish flash estimates has a significant impact over the full sample, leading to a change in long-term inflation compensation of around 0.6 and 0.5 basis point in the day of its release, respectively. However, subsample analysis suggests the statistical significance is mainly related to the latter part of the sample. Prior to 2013, there had been no strong significant reaction to inflation releases, while the reactions to surprises in flash estimates turned strongly significant over the below-target period 2013–17. Specifically, the reaction to the German flash estimate rose

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[14]The Chow forecast test selects a subsample to check the validity across the full sample. Since we assume that the below-target (BT) period has a more significant parameter vector, we use it to estimate the parameter vector $\widehat{\beta_{BT}} = (\mathbf{X}'_{t,BT}\mathbf{X}_{t,BT})^{-1}\mathbf{X}'_{t,BT}\Delta y_{t,BT}$ as well as obtain the residual sum of squares RSS_{BT}. The estimated parameter vector $\widehat{\beta_{BT}}$ and the surprises of the full sample $\mathbf{X}_{t,All}$ are used to estimate $\Delta y_{t,All}$ and subsequently obtain the RSS_{All}. The F statistic is then computed: $F(T_{PD}, T_{BT} - K) = \frac{RSS_{All} - RSS_{BT}/T_{PD}}{RSS_{BT}/(T_{BT}-K)}$.

[15]The appendix also confirms these results employing the same regression using a two-day window (tables A.4 and A.5) as well as using the 1y9y forward rate as an alternative measure of long-term inflation compensation (tables A.1 and A.2).

Table 2. Long-Term Inflation Compensation (5y5y, daily): Individual Variable Linear Regression $\Delta y_t = \alpha + \beta X_t + \varepsilon_t$

	Full Sample			Pre-disinflation			Below Target		
	β	t-stats	R^2	β	t-stats	R^2	β	t-stats	R^2
Flash									
Germany	0.60**	(2.36)	0.05	0.34	(0.91)	0.01	1.10***	(5.48)	0.26
Spain	0.48**	(2.35)	0.06	0.58*	(1.79)	0.08	0.33	(1.38)	0.03
Italy	0.04	(0.26)	0.00	0.00	(−0.02)	0.00	0.12	(0.37)	0.00
Euro Area	0.15	(0.82)	0.01	0.14	(0.52)	0.00	0.16	(0.72)	0.01
HICP									
Germany	0.09	(0.53)	0.00	0.06	(0.32)	0.00	0.46	(1.19)	0.02
Spain	0.26	(1.16)	0.02	0.31	(1.17)	0.02	0.02	(0.06)	0.00
France	−0.06	(−0.40)	0.00	−0.07	(−0.50)	0.00	0.08	(0.22)	0.00
Italy	0.19	(1.01)	0.01	0.36	(1.66)	0.03	−0.33	(−1.24)	0.03
Euro Area	0.06	(0.39)	0.00	0.06	(0.35)	0.00	0.03	(0.18)	0.00
PPI	−0.02	(−0.12)	0.00	0.02	(0.06)	0.00	−0.10	(−0.45)	0.00
GDP	0.42***	(2.98)	0.04	0.46***	(2.89)	0.06	0.26	(0.83)	0.01
PMI	−0.02	(−0.08)	0.00	−0.02	(−0.12)	0.00	0.34	(0.68)	0.00
Consumer Confidence	0.13	(0.93)	0.00	0.11	(0.74)	0.00	0.34	(0.80)	0.01
Industrial Confidence	−0.10	(−0.46)	0.00	−0.09	(−0.39)	0.00	0.01	(0.03)	0.00
Industrial Production	0.45***	(2.71)	0.05	0.46**	(2.38)	0.05	0.37	(1.57)	0.03
New Orders	−0.17	(−0.74)	0.01	−0.17	(−0.74)	0.01	—	—	—
Unemployment Rate	−0.30	(−1.28)	0.02	−0.25	(−0.82)	0.01	−0.30	(−1.11)	0.03
M3	−0.14	(−0.86)	0.01	−0.16	(−0.91)	0.01	−0.05	(−0.14)	0.00
MRO	—	—	—	—	—	—	0.35**	(2.50)	0.08
Retail Sales	0.13	(0.81)	0.00	0.09	(0.48)	0.00	—	—	—
Trade Balance	−0.02	(−0.09)	0.00	−0.02	(−0.05)	0.00	−0.02	(−0.09)	0.00
Mon. Policy Shock	−2.70	(−0.40)	0.00	4.37	(0.47)	0.00	−19.62**	(−2.36)	0.13
CB Info. Shock	25.31**	(2.42)	0.08	26.63*	(1.98)	0.08	25.22***	(3.19)	0.11

Notes: The table reports estimated β coefficients from regressions on days with data releases over three periods: from January 2005 to March 2017 (full sample), from January 2005 to December 2012 (pre-disinflation), and from January 2013 to March 2017 (below target), using surprises for each single variable in an individual regression. Dependent variable is one-day change, $\Delta y_t \equiv y_t - y_{t-1}$, in long-term inflation compensation (5y5y, in basis points). The surprise component of macroeconomic releases, x_t, is normalized by their historical standard deviations; coefficients β therefore represent a basis point per standard deviation response. Heteroskedasticity-consistent t-statistics are in parentheses; ***, **, and * denote a 1, 5, and 10 percent level of significance, respectively. R^2 reports the explanatory variable for each individual regression. (—) indicates an omission either due to the discontinuity of the expectations collection (e.g., new orders) or for lack of surprises over most of the period of the regression (e.g., MRO).

to 1.1 basis points—with a t-statistic of 5.48—almost twice as much as over the sample as a whole (0.6) and more than three times the average response (0.34) over the pre-disinflation period 2005–12. The impact of the Spanish flash estimate instead seems to be somewhat stronger before 2013, which is consistent with the early start of the decline in Spanish inflation after its post-GFC peak compared with the euro-area-wide inflation, and our interpretation that market participants paid more attention to inflation developments in periphery countries during the onset of the European debt crisis. Individual variable regressions also confirm the relevance of economic activity (GDP, industrial production) mainly before 2013, and monetary policy variables over the whole sample for long-term inflation compensation.

Importantly, the individual variable regressions underscore the strong role of the German flash estimate, the first piece of inflation news in the euro area every month, during the below-target period. Surprises in the German flash releases explain 26 percent of the variation of long-term inflation compensation in their release days, at least twice as much as any other individual variable, and almost four times as much as the R^2 for the joint regression shown above. Moreover, over the pre-disinflation period the explanatory power of all variables, including the statistically significant ones on economic activity, remain in single digits. Finally, the same pattern of stronger responses in the below-target period (2013–17) compared with the previous pre-disinflation period (2005–12) emerges after controlling for a large set of additional indicators of market conditions.[16] Joint and individual variable regressions at daily frequency provide evidence consistent with an increase in the sensitivity of long-term inflation compensation to inflation news since 2013. While actual inflation developments provide a strong case for splitting the sample in 2013, we investigate the time variation in the sensitivity of long-term inflation compensation to news in greater detail next.

[16]Table A.3 in the appendix consider an extended regression in which additional controls for market conditions have been included in the regression. Specifically, changes in the EA Citigroup surprise index, oil and gold prices, the EUR/USD exchange rate, VIX and VSTOXX, and CDS premiums are considered as controls.

4.2 Evidence on Time-Varying Sensitivity to News

The subsample analysis suggests that the impact of macroeconomic data releases on long-term inflation compensation has changed over time and, more specifically, has strongly risen since 2013. To obtain more precise evidence on how those reactions have evolved over time, equations (1) and (2) could be estimated over rolling windows. However, as most macroeconomic announcements only occur once a month, any rolling window based on individual variable release would suffer from small-sample problems. We therefore provide some evidence of time variation in the response of inflation compensation to macroeconomic releases using the approach recently introduced by Moessner and Nelson (2008) and Swanson and Williams (2014, SW henceforth) to study the evolution of the reaction of U.S. yields to macroeconomic news. The first step as proposed by Moessner and Nelson (2008) is to specify a nonlinear equation. A set of multiplicative time dummies δ allows for the time variation in the response of the variables of interest to macroeconomic news to be measured by the combined effect $\delta\beta$. Such a specification therefore measures whether the response to macroeconomic news has changed over time even if the relative magnitude of the elements of β are constant over time.

Formally, the regression (1) used in the previous section is extended to a nonlinear least squares specification as follows:

$$\Delta y_t = \gamma^{\tau_i} + \delta^{\tau_i}\beta\mathbf{X}_t + \epsilon_t^{\tau_i}, \tag{3}$$

where the parameters γ^{τ_i} and δ^{τ_i} take on different values in each calendar year $i = 2005, 2006, \ldots, 2017$. The second step is the key motivation for regression (3). It allows us to reduce the small-sample problem associated with allowing every element of X_t to vary over time. More specifically, by pooling information across data releases, we can obtain more detailed evidence on the time variation in the response of long-term inflation compensation to macroeconomic news through the estimation of daily rolling regressions of the form

$$\Delta y_t = \gamma^{\tau} + \delta^{\tau}\hat{\mathbf{X}}_t + \epsilon_t^{\tau}, \tag{4}$$

where $\hat{\mathbf{X}}_t \equiv \hat{\beta}\mathbf{X}_t$ is the generic surprise regressor defined using $\hat{\beta}$ from regression (3).

To make this approach operational, two additional assumptions are needed. First, a normalization for δ^{τ_i} needs to be chosen to separately identify β and δ^{τ_i}. The estimated response of long-term inflation compensation to macroeconomic news will then be relative to the response over the normalization period. Following SW, we normalize δ^{τ_i} to average 1 over a given period—in our baseline case, over the period of below-target period 2013–17. Intuitively, if there has been a deterioration in the anchoring of inflation expectations since 2013, by normalizing the response between 2013–17 to 1, we should observe a response of inflation compensation to macroeconomic news over the rest of our sample that is, on average, *below* 1.[17]

Second, τ in regression (4) denotes the period over which the rolling regressions are carried out. To identify a weakening of the anchoring of inflation expectations that is taking place over a substantial period of time rather than capturing a stronger reaction to news that may be triggered by just a few influential data releases in a few months in a row, we run regression (4) over a two-year window.[18] Following SW, we account for the two-stage sampling uncertainty by scaling the standard error σ_{δ^τ} of δ^τ in (4) by the weighted average of the derived standard errors of δ^{τ_i} in regression (3).[19]

Table 3 reports the estimated coefficients of nonlinear regression (3). The German flash estimate, the Spanish HICP, industrial production, and the central bank information shock are found to be significant at least at the 10 percent level and carry the expected sign, with the two inflation surprises significant at the 1 percent level and the 5 percent level, respectively. Figure 4 depicts the estimated sensitivity of long-term inflation compensation to macroeconomic news δ^{τ_i} together with their uncertainty bands using all variables

[17]Figure A.1 in the appendix offers results for different normalization periods and shows that our main findings do not depend on the specific period chosen for the normalization of δ^{τ_i}.

[18]The subsample analysis in the previous section points to an increase in the impact of inflation news since 2013. We look for supporting evidence by assessing the responses on rolling regressions over one-sided windows, while SW report results over two-sided windows centered around the business date.

[19]Specifically, σ^τ is adjusted as $\sigma^\tau_{adj} = \sigma^\tau \left(\sum_i^{I_\tau} w_i \zeta^{\tau_i}\right)$, where I_τ specifies the calendar years the rolling window covers and w_i the number of days in calendar year i devided by the total number of days of rolling window τ.

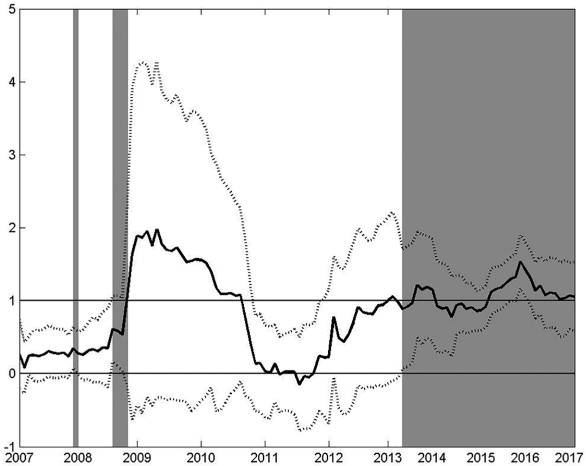
Table 3. Long-Term Inflation Compensation (5y5y, daily): Nonlinear Regression $\Delta y_t = \gamma^{\tau_i} + \delta^{\tau_i}\beta \mathbf{X}_t + \varepsilon_t$

	All Variables		Selected Variables	
	β	t-stats	β	t-stats
Flash				
Germany	0.89***	(4.14)	0.87***	(4.03)
Spain	0.12	(0.69)	0.11	(0.53)
Italy	−0.20	(−1.22)		
Euro Area	−0.05	(−0.16)		
HICP				
Germany	0.32	(1.34)		
Spain	0.68**	(2.43)	0.63**	(2.26)
France	−0.01	(−0.08)		
Italy	0.37	(1.62)	0.27	(1.49)
Euro Area	−0.01	(−0.11)		
PPI	−0.24	(−1.37)		
GDP	0.09	(0.54)		
PMI	0.05	(0.34)		
Consumer Confidence	0.08	(0.57)		
Industrial Confidence	−0.19	(−1.15)		
Industrial Production	0.50*	(1.85)	0.49*	(1.79)
New Orders	0.13	(0.92)		
Unemployment Rate	−0.16	(−0.62)		
M3	0.09	(0.69)		
Retail Sales	0.04	(0.20)		
Trade Balance	−0.07	(−0.49)		
Mon. Policy Shock	6.17	(1.14)		
CB Info. Shock	19.89*	(1.81)	19.96*	(1.87)
Observations	1,767		687	
R^2	0.05		0.11	
$H_0 : \beta$ Constant, p-value	0.99		0.99	
$H_0 : \delta$ Symmetric (p-value)	0.92		0.92	
$H_0 : \delta$ Constant (p-value)	0		0	

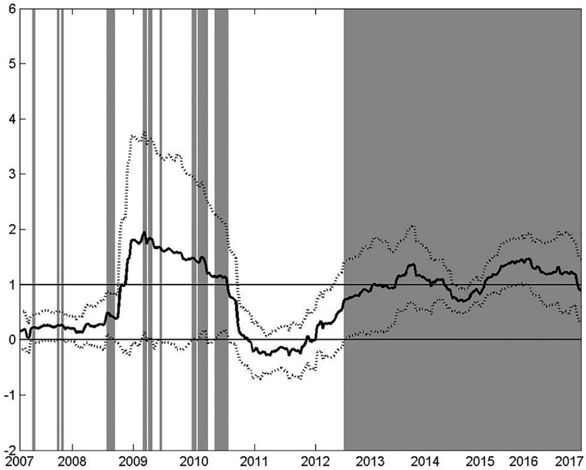
Notes: The table reports estimated β coefficients from a regression on days of releases from January 2005 to March 2017, using surprises for all variables in a joint (nonlinear) regression (Swanson and Williams 2014). Coefficients indexed τ_i may take on different values in different calendar years. Heteroskedasticity-consistent t-statistics are in parentheses; and ***, **, and * denote a 1, 5, and 10 percent level of significance, respectively. $H_0 : \beta$ *constant* implies that the coefficient vector β in the regression is constant over time and only the scalar coefficients δ^{τ_i} vary against an alternative in which every element of β is permitted to vary independently across calendar years. $H_0 : \delta$ *symmetric* assumes that the δ^{τ_i} in the regression are the same for positive and negative surprises $\beta \mathbf{X}_t$, against an alternative in which we allow separate coefficients $\delta_+^{\tau_i}$ and $\delta_-^{\tau_i}$ for positive and negative values of $\beta \mathbf{X}_t$ in each calendar year i. $H_0 : \delta$ *constant* assumes that the time-varying sensitivity coefficients δ^{τ_i} in the regression are constant over time. That is, we test whether $\delta^{\tau_i} = 1$ for each calendar year $i = 2005, \ldots, 2015$.

Figure 4. Long-Term Inflation Compensation (5y5y, daily): Time-Varying Sensitivity Coefficients δ^{τ} from Nonlinear Regression

A. δ^{τ} Using All Variable Regression Results



B. δ^{τ} Using Selected Variable Regression Results



Notes: This figure plots the coefficients δ^{τ} estimated from a regression $\Delta y_t = \gamma^{\tau} + \delta^{\tau}\hat{\mathbf{X}}_t + \epsilon_t^{\tau}$. Panels A and B depict δ^{τ} using the all and selected variable regression results (see table 3), respectively. Dotted gray lines depict heteroskedasticity-consistent $\pm$ two-standard-error bands, adjusted for two-stage sampling uncertainty in δ^{τ}. Shaded regions denote δ^{τ} being significantly above 0 and suggest a weaker anchoring of inflation expectations.

employed in the daily regressions above. While the sensitivity to news displays some variation in the responses between 2005 and 2017, and several episodes of responsiveness significantly above zero can be identified over the sample, the level and significance of the sensitivity of inflation compensation to news is (in relative terms) clearly higher, and more persistent, from 2013. Indeed, since euro-area inflation has remained (mostly) below target, the estimates of δ^{τ_i} rose from a level statistically not different from zero since mid-2010 to be most of the time above 1 since 2014.[20]

Our results also identify a sharp deterioration in the anchoring of inflation expectations in the second half of 2008. This is consistent with the substantial increase in oil prices, actual inflation, and inflation compensation over 2008 ahead of the collapse of Lehman Brothers. The results depicted in figure 4 suggest that the 2008 episode was more short-lived, but the period 2008–09 should be interpreted with caution because euro-area inflation markets were under significant stress during that time, and it cannot be ruled out that the stronger sensitivity to news may just be the result of higher volatility of inflation compensation measures. Indeed, there is a significant widening of uncertainty in the estimation around that time and δ^{τ} becomes statistically insignificant despite its relatively high level.

Including many non-significant macroeconomic news in (3) might however add unnecessary noise to the regression. As a robustness check, we also provide the estimation of δ^{τ_i} restricting the set of variables to those whose news is found to trigger significant reactions in regressions above (results for the restricted nonlinear specification for those six variables are reported in table 3). The regression results and the rolling parameter δ^{τ} are very much in line with those from the previous regression. Similarly, the estimates of the time-varying

[20]Following SW, tests for several additional dimensions of the time variation in the responses are reported at the bottom of table 3. First, we test whether the relative response coefficients β are constant over time (only δ^{τ_i} varies) against an alternative in which every element of β varies. p-values close to 1 support the restricted specification (3). Second, we test whether the δ^{τ_i} coefficients are the same for positive and negative surprises. With a p-value close to 1, the hypothesis of a symmetric response is not rejected. Finally, we test the hypothesis that the time-varying coefficient δ^{τ_i} is constant over time—that is, for each calendar year $i = 2005, \ldots, 2017$ $\delta^{\tau_i} = 1$. With a low p-value, the data clearly reject that restriction.

sensitivity of long-term inflation compensation to news further single out the relevance of the period beyond 2013 in terms of statistical and economic significance over the sample as a whole.

Overall, the time-varying regressions point to a deterioration in the anchoring of euro-area inflation expectations since 2013. While our findings are robust to different normalization periods and the inclusion of different inflation releases in the analysis, it is also true that their estimation requires nonetheless pooling together a number of variables. The next section provides additional evidence on the responses to news on specific data releases.

4.3 Evidence from Intraday Data

We expand existing analysis on the reaction of euro-area inflation expectations to news by conducting regressions using high-frequency (intraday) data. Arguably, event-study analysis, as the one conducted in previous sections and the related literature, is preferably undertaken with intraday data: the smaller the window around the news arrival, the less likely anything but the news under investigation affects asset price changes, reinforcing the case for direct causality and making the event study resemble as close as possible a controlled natural experiment (Gürkaynak and Wright 2013).

To our knowledge, this is the first event-study analysis using intraday data for euro-area inflation compensation measures, and there is some uncertainty about the speed of reaction of the euro-area IL swap market to news. We consider the changes in the spot and forward inflation compensation over a benchmark window of 120 minutes after the data release, allowing for 10 minutes before the data release for the formation of expectations ahead of the release. Over that window, at least in our sample, there are no important additional releases that interfere with market trading of inflation compensation. Therefore, the changes should be directly attributable to the specific data release under study. International evidence on intraday data on inflation compensation is also limited. Beechey and Wright (2009), for example, look at the U.S. market but focus on bond-based BEIRs rather than IL swaps, which allows them to consider a shorter window of 15 minutes.

A shortcoming of our high-frequency data is that their collection is only available from late 2008. We use the specific time of the data release as reported in Bloomberg, and our intraday trades of euro-area IL swap rates from Reuters. While we cannot use our intraday data sample to investigate the anchoring of inflation expectations prior to the GFC, we can however assess the reaction of inflation compensation measures in the most important period, namely the severe disinflation that took place in the euro area since 2013, and compare it with the 2009–12 period.

Intraday evidence corroborates our previous findings of a deterioration in the anchoring of long-term inflation expectations in the later part of our sample. Between 2013 and 2017, over a 120-minute window from the data release, long-term inflation compensation has significantly reacted to the news component of the flash releases for Germany and Spain (1 percent and 10 percent), as well as of the Spanish HICP (5 percent) (see table 4). Significant responses (at 10 percent) were also found for PPI, the unemployment rate, and policy rate announcements, in line with the evidence from daily regressions. In the 2009–12 period instead we find some significant reactions to the German flash estimate (at 10 percent) and the PPI (with a negative sign), but those reactions seem to be quantitative and qualitatively weaker than those after 2013. First, before 2013 we cannot reject the joint hypothesis of nonsignificance (p-value 0.8), while since 2013 we can safely reject the joint nonsignificance with a p-value of 0.01. In addition, the regression coefficients indicate a structural break with the p-value of the Chow forecast test of 0.059.

As argued in section 4.1, the joint regressions might add additional noise to the regression by treating all other variables as zero surprise though only one variable has a news release in a given day. As a robustness check, we also run variable-by-variable regressions over the same window length (t–10 minutes, t+120 minutes) to gain additional insights on the variables driving changes in inflation compensation. German and Spanish flash news is strongly significant (at 1 percent) and explains 38 percent and 12 percent of the variability of long-term inflation compensation, significantly above the Spanish HICP, the PPI, and the policy rate surprise, which are also found to be significant. In stark contrast, no single inflation news surprise is found to trigger a significant

Table 4. Long-Term Inflation Compensation (5y5y, intraday): 120-Minute Window (minus 10 minutes)

	Joint Linear Regression				Individual-Variable Linear Regressions					
	2009–12		Below Target		2009–12			Below Target		
	β	t-stats	β	t-stats	β	t-stats	R^2	β	t-stats	R^2
Flash										
Germany	0.53*	(1.75)	0.56***	(3.63)	0.58	(1.58)	0.08	0.62***	(3.90)	0.38
Spain	0.09	(0.41)	0.2*	(1.67)	0.07	(0.24)	0.00	0.38**	(2.66)	0.12
Italy	−0.29	(−1.11)	−0.14	(−0.54)	−0.33	(−1.58)	0.09	0.02	(0.08)	0.00
Euro Area	−0.06	(−0.33)	−0.09	(−0.43)	−0.28	(−1.31)	0.04	0.01	(0.05)	0.00
HICP										
Germany	−0.14	(−0.89)	0.07	(0.16)	−0.17	(−0.98)	0.02	0.08	(0.19)	0.00
Spain	0.06	(0.40)	0.29**	(2.24)	0.20	(1.14)	0.01	0.27**	(2.51)	0.03
France	−0.07	(−0.45)	0.37*	(1.95)	−0.18	(−1.18)	0.03	0.32	(1.50)	0.03
Italy	−0.14	(−0.63)	0.04	(0.17)	−0.24	(−1.04)	0.07	0.02	(0.09)	0.00
Euro Area	−0.29	(−0.78)	0.05	(0.26)	−0.47	(−1.10)	0.10	0.01	(0.09)	0.00
PPI	−0.47**	(−2.36)	−0.18*	(−1.83)	−0.62**	(−2.05)	0.10	−0.20*	(−2.00)	0.05
GDP	0.13*	(1.69)	−0.24	(−1.51)	0.16**	(2.44)	0.04	−0.11	(−0.78)	0.00
PMI	−0.06	(−0.31)	−0.17	(−0.72)	−0.11	(−0.42)	0.00	−0.03	(−0.12)	0.00
Consumer Confidence	−0.01	(−0.14)	0.01	(0.06)	−0.07	(−0.69)	0.00	0.02	(0.17)	0.00
Industrial Confidence	−0.08	(−0.44)	0.01	(0.06)	−0.05	(−0.27)	0.00	0.05	(0.22)	0.00
Industrial Production	0.11	(0.92)	−0.1	(−0.72)	0.18	(1.34)	0.03	−0.14	(−0.97)	0.01
New Orders	0.26	(0.75)	—	—	0.39	(1.03)	0.05	—	—	—
Unemployment Rate	0.18	(0.34)	−0.42*	(−1.93)	0.00	(0.00)	0.00	−0.28	(−1.42)	0.04
M3	0.02	(0.16)	0.28	(1.33)	0.07	(0.61)	0.00	0.27	(1.24)	0.04
MRO	—	—	0.16*	(1.67)	—	—	—	0.23**	(2.48)	0.06
Retail Sales	−0.03	(−0.21)	—	—	−0.03	(−0.18)	0.00	—	—	—
Trade Balance	−0.11	(−0.87)	0.05	(0.73)	0.18	(1.57)	0.03	0.10	(0.75)	0.01
Observations	516		562							
$\bar{R}^2$	0.00		0.03							
R^2	0.03		0.07							
$H_0 : \beta = 0$ (p-value)	0.80		001							
Chow Test (p-value)	0									

Notes: The table reports estimated β coefficients from regressions on days with data releases over two periods: from April 2009 to December 2012 (pre-disinflation) and from January 2013 to March 2017 (below target), using surprises for each variable in a joint regression (left columns) and for each single variable in an individual regression (right columns). Changes in long-term inflation compensation are calculated over a 120-minute window. The surprise component of macroeconomic releases k is normalized by their historical standard deviations; coefficients represent a 10^{-2} basis point per standard deviation response. Heteroskedasticity-consistent t-statistics are in parentheses; ***, **, and * denote a 1, 5, and 10 percent level of significance, respectively. Chow test (p-value) is for the H_0 of parameter constancy between the pre-disinflation and below-target period. (—) indicates an omission either due to the discontinuity of the expectations collection (e.g., new orders) or for lack of surprises over most of the period of the regression (e.g., MRO).

reaction before 2013, when only the PPI and the GDP are found to be significant.

We also investigate the robustness of our conclusions to the speed of reaction to news in the euro-area IL swap market by looking at two different windows around our benchmark, 15 minutes and 240 minutes after the release, and computing changes from 10 minutes before the release. See tables 5 and 6. Results over a shorter time of a 15-minute window confirm the strong impact of the German flash release after 2013, and also of other inflation news (Italian, with a wrong sign, and euro-area flashes, and the French CPI, at 10 percent significance levels) as well as the unemployment rate and M3 releases. Overall, there is evidence of statistical significance at individual and joint level after 2013, while, despite some significant impact, the hypothesis of joint insignificance of news before 2013 cannot be rejected (p-value 0.46), while the stability of the regression coefficients can be rejected (p-value 0).

Reactions over a longer window of t+240 minutes after the release (from t–10 minutes) confirm that impact identified over shorter windows persists over time, and further corroborate the main message found so far. Specifically, while very few significant reactions to news are found prior to 2013, since then German and Spanish flashes, the Spanish HICP, the PPI (with a negative sign though), and news about the ECB's main refinancing rate are found to trigger statistically significant reactions. Intraday analysis therefore provides strong support to the importance of early releases, as well as the presence of a weakening of the anchoring in inflation expectations since 2013.

5. Discussion

Previous sections have reported robust evidence of an increase in the response of long-term inflation compensation to inflation and macroeconomic news in the euro area in recent years by means of three different but complementary pieces of analysis, namely daily, time-varying, and intraday regressions. This section highlights some additional dimensions of our findings, and relates them to some additional evidence and recent literature.

Table 5. Long-Term Inflation Compensation (5y5y, intraday): 15-Minute Window (minus 10 minutes)

	Joint Linear Regression				Individual-Variable Linear Regressions					
	2009–12		Below Target		2009–12			Below Target		
	β	t-stats	β	t-stats	β	t-stats	R^2	β	t-stats	R^2
Flash										
Germany	0.08	(0.94)	0.21**	(2.50)	0.10	(0.83)	0.03	0.19**	(2.11)	0.16
Spain	0.07	(0.52)	−0.01	(−0.10)	0.04	(0.27)	0.01	0.07	(0.87)	0.02
Italy	0.05	(0.47)	−0.18**	(−1.98)	−0.07	(−0.70)	0.02	−0.04	(−0.56)	0.00
Euro Area	−0.23**	(−2.35)	0.19*	(1.77)	−0.18*	(−1.94)	0.09	0.19*	(1.85)	0.07
HICP										
Germany	0.01	(0.05)	0.06	(0.52)	0.02	(0.13)	0.00	0.09	(0.74)	0.01
Spain	−0.17	(−0.78)	0	(0.08)	−0.24	(−0.86)	0.02	0.01	(0.11)	0.00
France	0.02	(0.26)	0.12*	(1.74)	0.01	(0.16)	0.00	0.13	(1.55)	0.02
Italy	0.03	(0.22)	0.05	(0.50)	0.01	(0.08)	0.00	0.07	(0.62)	0.01
Euro Area	−0.12	(−1.19)	0.04	(0.35)	−0.11	(−0.91)	0.02	0.04	(0.43)	0.01
PPI	−0.18	(−1.41)	0.06	(1.25)	−0.21	(−1.23)	0.02	0.06	(1.10)	0.01
GDP	−0.06	(−0.87)	−0.09	(−0.63)	−0.11	(−0.93)	0.03	−0.04	(−0.34)	0.00
PMI	−0.15**	(−1.97)	−0.13	(−0.68)	−0.18**	(−2.05)	0.07	−0.16	(−0.81)	0.00
Consumer Confidence	−0.03	(−0.41)	0.02	(0.22)	−0.04	(−0.40)	0.00	0.01	(0.12)	0.00
Industrial Confidence	0.02	(0.36)	0.21	(1.47)	0.01	(0.18)	0.00	0.26	(1.67)	0.10
Industrial Production	−0.06	(−0.76)	−0.06	(−0.63)	−0.08	(−0.92)	0.02	−0.08	(−0.74)	0.01
New Orders	0.24	(1.37)	—	–	0.28	(1.43)	0.06	—	—	—
Unemployment Rate	−0.08*	(−0.58)	−0.17*	(−1.78)	−0.17	(−1.44)	0.05	−0.17**	(−2.27)	0.04
M3	0.07	(0.56)	−0.18**	(−1.96)	0.10	(−0.76)	0.01	−0.19**	(−2.14)	0.09
MRO	—	—	0.04	(0.79)	—	—	—	0.04	(0.77)	0.01
Retail Sales	0.08	(1.70)	—	—	0.08	(1.42)	0.04	—	—	—
Trade Balance	−0.11	(−0.87)	0.05	(0.73)	−0.15	(−0.92)	0.02	0.05	(0.72)	0.01
Observations	562		516							
$\bar{R}^2$	0		0.02							
R^2	0.04		0.05							
$H_0 : \beta = 0$ (p-value)	0.46		0.04							
Chow Test (p-value)	0									

Notes: The table reports estimated β coefficients from regressions on days with data releases over two periods: from April 2009 to December 2012 (pre-disinflation) and from January 2013 to March 2017 (below target), using surprises for each variable in a joint regression (left columns) and for each single variable in an individual regression (right columns). Changes in long-term inflation compensation are calculated over a 15-minute window. The surprise component of macroeconomic releases k is normalized by their historical standard deviations; coefficients represent a 10^{-2} basis point per standard deviation response. Heteroskedasticity-consistent t-statistics are in parentheses; ***, **, and * denote a 1, 5, and 10 percent level of significance, respectively. Chow test (p-value) is for the H_0 of parameter constancy between the pre-disinflation and below-target period. (—) indicates an omission either due to the discontinuity of the expectations collection (e.g., new orders) or for lack of surprises over most of the period of the regression (e.g., MRO).

Table 6. Long-Term Inflation Compensation (5y5y, intraday): 240-Minute Window (minus 10 minutes)

	Joint Linear Regression				Individual-Variable Linear Regressions					
	2009–12		Below Target		2009–12			Below Target		
	β	t-stats	β	t-stats	β	t-stats	R^2	β	t-stats	R^2
Flash										
Germany	0.98	(1.61)	0.91***	(3.19)	1.05	(1.42)	0.07	1.04***	(3.46)	0.32
Spain	0.10	(0.33)	0.42*	(1.71)	0.09	(0.21)	0.00	0.69**	(2.62)	0.10
Italy	−0.63	(−1.49)	−0.10	(−0.19)	−0.59*	(−1.76)	0.10	0.09	(0.16)	0.00
Euro Area	0.10	(0.29)	−0.38	(−0.92)	−0.38	(−0.96)	0.02	−0.17	(−0.58)	0.00
HICP										
Germany	−0.29	(−1.15)	0.08	(0.09)	−0.37	(−1.32)	0.03	0.07	(0.08)	0.00
Spain	0.28	(0.98)	0.57**	(2.33)	0.65	(1.60)	0.04	0.54**	(2.52)	0.04
France	−0.15	(−0.52)	0.62	(1.62)	−0.37	(−1.28)	0.04	0.52	(1.25)	0.02
Italy	−0.3	(−0.84	0.02	(0.05)	−0.49	(−1.28)	0.09	−0.03	(−0.08)	0.00
Euro Area	−0.46	(−0.71)	0.06	(0.22)	−0.83	(−1.10)	0.10	−0.01	(−0.04)	0.00
PPI	−0.75**	(−1.96)	−0.43**	(−2.29)	−1.03*	(−1.82)	0.09	−0.45**	(−2.50)	0.10
GDP	0.31*	(1.94)	−0.38	(−1.52)	0.43**	(2.73)	0.08	−0.19	(−0.75)	0.00
PMI	0.03	(0.09)	−0.22	(−0.63)	−0.04	(−0.07)	0.00	0.10	(0.28)	0.00
Consumer Confidence	0.01	(0.05)	0	(−0.02)	−0.09	(−0.44)	0.00	0.03	(0.15)	0.00
Industrial Confidence	−0.19	(−0.49)	−0.19	(−0.77)	−0.12	(−0.31)	0.00	−0.17	(−0.53)	0.01
Industrial Production	0.28	(1.24)	−0.14	(−0.50)	0.43*	(1.75)	0.05	−0.20	(−0.70)	0.01
New Orders	0.28	(0.47)	—	—	0.50	(0.77)	0.03	—	—	—
Unemployment Rate	0.43	(0.46)	−0.66	(−1.60)	0.17	(0.12)	0.00	−0.40	(−1.02)	0.02
M3	−0.03	(−0.11)	0.75*	(1.86)	0.03	(0.12)	0.00	0.74*	(1.74)	0.09
MRO	—	—	0.29	(1.42)	—	—	—	0.42**	(2.26)	0.06
Retail Sales	−0.14	(−0.52)	—	—	−0.14	(−0.43)	0.01	—	—	—
Trade Balance	0.35	(1.25)	0.15	(0.71)	0.50*	(1.95)	0.04	0.15	(0.69)	0.01
Observations	562		516							
$\bar{R}^2$	0.00		0.03							
R^2	0.03		0.07							
$H_0 : \beta = 0$ (p-value)	0.82		0.02							
Chow Test (p-value)	0									

Notes: The table reports estimated β coefficients from regressions on days with data releases over two periods: from April 2009 to December 2012 (pre-disinflation) and from January 2013 to March 2017 (below target), using surprises for each variable in a joint regression (left columns) and for each single variable in an individual regression (right columns). Changes in long-term inflation compensation are calculated over a 240-minute window. The surprise component of macroeconomic releases k is normalized by their historical standard deviations; coefficients represent a 10^{-2} basis point per standard deviation response. Heteroskedasticity-consistent t-statistics are in parentheses; ***, **, and * denote a 1, 5, and 10 percent level of significance, respectively. Chow test (p-value) is for the H_0 of parameter constancy between the pre-disinflation and below-target period. (—) indicates an omission either due to the discontinuity of the expectations collection (e.g., new orders) or for lack of surprises over most of the period of the regression (e.g., MRO).

5.1 The Impact of Inflation News and Its Timing

The statistical evidence of a significant increase in the impact of inflation surprises on inflation compensation since 2013 points to a weakening of the anchoring of euro-area long-term inflation expectations. Our analysis sheds additional evidence on the role of the timing of the inflation news to assess that de-anchoring: not only does inflation news have a stronger impact in the below-target period, but that evidence is consistent with a stronger role of the early inflation releases, particularly the German but also the Spanish flash estimates, instead of the HICP releases (country and euro-area-wide). We believe this evidence is fully consistent with an optimal pricing of news among market participants: earlier inflation news is fully priced in, making additional information released later on somewhat less likely to cause a significant impact.

The importance of early inflation releases in the pricing of inflation compensation is indeed a crucial contribution of this paper. We have emphasized their role in explaining changes of long-term inflation compensation, but it is important to note that flash estimates also have a crucial role in the formation of inflation expectations at shorter horizons. Indeed, regression results along the lines of the ones reported in the previous section confirm the relevance of the flash estimates for short-term measures of inflation compensation at one-year (spot IL swap rate) and the one-year forward IL swap rate in one year (see tables A.6 and A.7 in the appendix). Specifically, early inflation releases—in particular, the German flash but also the Italian and the euro-area flashes—are found to be statistically significant and have a much stronger impact than the corresponding HICP releases. Moreover, among the HICP releases only the French HICP, for which there was no flash estimate before 2016, appears to trigger strong significant reactions in short-term inflation compensation. This evidence provides strong support for the crucial importance of early released inflation news in the expectation formation in inflation markets.

That observation is consistent with recent literature investigating the tradeoff between early availability and information quality, namely that it is difficult to achieve substantial improvements that outweigh the disadvantage of having formed price expectations using the early, even if less precise, indicator (see Hess and Niessen 2010, Gilbert et al. 2017). Hence, once there is a statistically significant reaction to the early-released flash estimates, the fact that the euro-area flash estimate or the HICP releases have not triggered significant market reactions should not be interpreted as signaling strong anchoring. The crucial statistical role of the early-released country flashes in the variable-by-variable and joint regressions reported above is clear evidence of their economic relevance. Moreover, evidence from different windows for intraday movements, daily regressions, and two-day regressions does confirm that the effects of early news are long-lasting.

Previous sections also report statistically significant reactions to other macroeconomic surprises both before (e.g., GDP, industrial production) and after (e.g., unemployment rate) 2013. Their statistical significance, however, tends to be less consistent in joint and variable-by-variable regressions. It changes over subsamples, and—in contrast to inflation news in the below-target period—tests for joint significance tend to be rejected at standard significance levels. Similarly, while monetary policy surprises may also have an impact on long-term inflation expectations, their inclusion in the joint regressions does not diminish the significance of early inflation news. Indeed, the joint regressions point to a lower relevance of central bank communication (as measured by the central bank information shock) beyond surprises in the interest rate announcement themselves in the below-target period.

The strong relevance of inflation news is also consistent with empirical evidence pointing at the superior performance of univariate models when forecasting inflation (e.g., Faust and Wright 2013 or Chan, Clark, and Koop 2018 in two very different settings). Non-inflation surprises, which rather relate to the current state of the business cycle, may be arguably more relevant for the anchoring of inflation expectations in countries like the United States, where the Federal Reserve has a double mandate of inflation stability and maximum employment, than to the euro area, where the ECB's target is solely defined in terms of price stability.[21]

[21] Speck (2016), for example, also finds evidence of significant surprises using German and Italian business confidence as well as French manufacturing confidence, but concludes there is no clear evidence of a de-anchoring of inflation expectations over his sample.

5.2 What Is Behind the Reaction of Inflation Compensation?

The two components of inflation compensation, the level of long-term inflation expectations and the inflation risk premium, reflect different dimensions of anchoring, and are therefore very relevant for our goal in this paper. While it is difficult to disentangle these two (unobserved) components,[22] we argue that the stronger sensitivity of long-term forward inflation compensation to macroeconomic news since 2013 is indicative of a weaker anchoring of inflation expectations. In this section we use additional information from survey data to discuss whether that weakening may reflect changes in expected inflation or the inflation risk premium, or both.

It is also important to note that our analysis of inflation compensation does not rely on the expectations theory of the term structure. Since long-term forward IL swap rates capture the compensation that investors demand both for expected inflation and for the risks associated with inflation at that horizon, changes in inflation compensation need not be due solely to shifts in the conditional expected level of inflation at long horizons: if the anchoring of inflation expectations weakens, economic news might well also shift the inflation risk premium in long-term inflation compensation, either because investors' perceptions regarding the distribution of long-run inflation outcomes changes or because the economic news triggered a change in pricing of long-run inflation risks.

The significant decline in long-term inflation compensation and its very low levels since 2013 (see figure 1) suggest that a decline in the level of inflation expectations embodied in long-term inflation compensation measures, above and beyond a potential decline in the inflation risk premiums,[23] is the most plausible interpretation.

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[22] Approaches to correct BEIRs for liquidity premium—based on relative traded volumes or asset-swap spreads, for example—necessarily involve some crucial assumptions on model specification or its presence across maturities. In turn, inflation risk premium estimation usually takes place using term structure models. Estimates vary significantly across specifications even for a single country, but they generally point to significant variation in inflation risk premium over time and across maturities (see, for example, Kim, Walsh, and Wei 2019).

[23] Term structure models regularly monitored by the Federal Reserve or the ECB point to a compression in inflation risk premium in the United States and the euro area, but cannot fully explain the decline of long-term inflation compensation measures (Board of Governors of the Federal Reserve System 2015).

Additional results beyond the main focus of this paper are consistent with that interpretation. For example, trend inflation estimates along the lines of Stock and Watson (2007) point to a significant decline in long-term optimal conditional forecast since 2013. Moreover, that finding holds even allowing for forward-looking information into the term structure estimation, For example, García and Poon (2018) show that the protracted decline in long-term inflation compensation since 2013 is explained by both a significant decline in the level of long-term inflation expectations—from around 2 percent to lows around 1.5 percent—and a decline in inflation risk premiums—from around 40 basis points in mid-2013 to zero and even temporarily negative levels in the second half of 2016.

Survey data do not incorporate inflation risk premiums, and offer a somewhat longer history (since the start of the euro area in 1999) but at much lower frequencies (quarterly; see figure 1). Yet, Łyziak and Paloviita (2017) tested the sensitivity of long-term inflation expectations to changes in short-term ones and, by splitting the sample at the second quarter of 2008, found evidence of a de-anchoring of long-term inflation expectations based on the significant impact of short-term expectations on longer-term ones over the second part of their sample.

5.3 Evidence from Additional Inflation Derivatives

We have focused on the literature investigating the reaction of inflation compensation to news because it is the closest to the analysis carried out in this paper. A growing literature has however investigated changes in the anchoring of euro-area long-term inflation expectations using alternative data sources from the expanding IL derivatives market, mainly the growing inflation options market.[24] Here we review its findings, which overall tend to support a weakening in the anchoring of inflation expectations in the latter part of our sample.

[24] Inflation options, in similar fashion to standard options for stocks or interest rates, provide financial protection when inflation, the underlying asset here, moves above or below (cap and floor options, respectively) a given threshold (i.e., the strike price or rate). As in the case of IL swap rates, inflation options are nowadays actively traded over the counter, and their market is more developed in the euro area than in most other countries (Gimeno and Ibañez 2018).

Galati et al. (2016) investigate the reaction the deflation risk implied by risk-neutral densities (RNDs) on year-on-year inflation to oil price changes, and find evidence consistent with a subtle, but nonetheless statistically significant, weakening in the anchoring of long-term inflation expectations. Gimeno and Ibañez (2018) estimate inflation RNDs using IL swap and inflation options (caps and floors) across a large number of horizons—including at the five-year forward in five years benchmark reference for monetary policy—and find that the priced probability associated with negative inflation values at that long horizon, after declining from 2012, rose significantly again from early 2014 and almost doubled in less than two years to around 12 percent (above 30 percent at the two-year forward in two years). While those probabilities may not appear to be very high, their pricing at fairly long horizons suggests that they do not reflect just temporary factors, which is consistent with a weakening of the anchoring of euro-area long-term inflation expectations since 2013. Those perceived risks can explain why investors could be willing to forgo some return in the form of risk premiums in circumstances in which nominal bonds may turn from inflation bets to "deflation hedges" (Campbell, Sunderam, and Viceira 2017). A downward revision in inflation risk premium in light of a rising probability of persistent low inflation, even if the perceived deflation risks somewhat recede, is also a potential concern for any central bank, and part of a de-anchoring of long-term expectations (and associated risks) that the long-term inflation compensation measures used in this paper are able to capture.

Natoli and Sigalotti (2017) investigate tail correlations using inflation swaps and options data and introduce a new indicator based on the odds that strong negative shocks to short-term expectations are connected to large declines in long-term expectations using a logistic regression framework. That paper reports an increase in the risk of de-anchoring since late 2014 for the euro area, while evidence for the United States and the United Kingdom instead points to firmly anchored inflation expectations. Scharnagl and Stapf (2015) also find low sensitivity in euro-area inflation options to news before early 2013, and conclude that inflation expectations remained well anchored in their sample.

5.4 Additional Robustness Checks

Our main findings are also robust to increasing the number of controls in the joint regressions (see table A.3 in the appendix). Interestingly, the relevance of market conditions also changes from the pre-disinflation period to the below-target period. While bond and stock market volatility, and index of CDS premiums in euro-area sovereign bond markets, or even an index of economic surprises for the U.S. economy, triggered statistically significant reactions in euro-area long-term inflation compensation before 2013, in the below-target period in contrast there was an important shift in the relevant news for inflation compensation, with early inflation news (i.e., the German flash estimate) emerging as the main driver, above and beyond oil price changes as the most relevant variable.

Liquidity considerations have led our choice of the five-year IL forward rate in five years as a benchmark inflation compensation measure. Alternative measures employed in the related literature, such as the one-year forward in nine years, are also broadly supportive of our main findings (see tables A.1 and A.2 in the appendix). While the strongly significant reactions triggered by the German flash estimate in the below-target period are corroborated, results over the full sample period are instead somewhat puzzling. Limited liquidity in the euro-area IL swap market at this maturity however calls for some caution when interpreting these empirical results.

The linear specification of our regressions may also trigger some concerns about the irrelevance of changes in the distribution of data surprises, in particular if the distribution of data surprises X_t in the period 2013–17 were very different from that in the earlier part of the sample. In our empirical regressions, the surprise component of data releases X_t is regarded as strictly exogenous, as the expectations data used in their calculation are assumed to incorporate all relevant information available prior to the release. The inflation surprises used in our analysis are therefore considered independent of all past and future values of the changes in long-term inflation compensation on the left-hand side of these regressions, and strict exogeneity implies that the empirical distribution of the data surprises is irrelevant for our estimates of the response coefficients in our linear empirical regressions (1), (2), and (3). The distribution of inflation surprises—in particular, those for the surprise component in country flash estimates—has however been quite similar across time. Such similarity in the size of surprises may be puzzling at first sight, but it is consistent with a gradual downward revision to short-term inflation expectations in line with the decline in actual inflation among financial market participants. In those circumstances, the surprises in inflation releases should not look very different from those over earlier periods. Moreover, given the significant downward revision in short-term inflation expectations since 2013, it is plausible that the protracted period of low inflation also led to downward revisions in the level of long-term inflation compensation, as shown in figure 1, and its responsiveness to inflation news as shown in previous sections.

6. Conclusions

Following the unprecedented period of low (and mostly below-target) inflation in the euro area since 2013, whether the anchoring of euro-area inflation expectations has weakened in recent years is a crucial question. The monetary transmission mechanism is most effective when long-term inflation expectations are strongly anchored, and, indeed, evidence on earlier samples has found that euro-area inflation expectations were well anchored around the ECB's quantitative definition of price stability of (below but close to) 2 percent.

This paper provides two main findings to shed new light on those considerations. First, by expanding the standard analysis to all available early inflation releases for country and euro-area-wide inflation, we find a crucial role for the timing of inflation releases. Indeed the main reactions in inflation compensation are triggered by early data releases, the country flash estimates regularly released ahead of the euro-area-wide data, and official country HICP releases. Second, taking into account those early releases, we find evidence of stronger reactions of long-term inflation compensation measures to news since 2013. Evidence from daily, time-varying, and intraday regressions consistently shows a significant impact of inflation news over recent years that had not been observed before in the euro area, and points to a significant deterioration in the anchoring of inflation expectations in the euro area.

Appendix

Table A.1. Alternative Long-Term Inflation Compensation (1y9y, daily): Joint Linear Regression $\Delta y_t = \alpha + \beta X_t + \varepsilon_t$

	Full Sample		Pre-disinflation		Below Target	
	β	t-stats	β	t-stats	β	t-stats
Flash						
Germany	0.34	(0.89)	−0.11	(−0.21)	1.38***	(3.27)
Spain	0.92**	(2.58)	1.34**	(2.56)	0.10	(0.26)
Italy	0.11	(0.27)	0.26	(0.50)	−0.14	(−0.22)
Euro Area	−0.74	(−1.61)	−1.22*	(−1.83)	0.14	(0.28)
HICP						
Germany	0.40	(1.25)	0.32	(0.95)	1.22	(1.41)
Spain	0.11	(0.16)	0.07	(0.08)	0.47	(0.58)
France	−0.64	(−0.88)	−0.77	(−0.96)	0.56	(0.97)
Italy	0.41	(1.09)	0.42	(0.90)	0.42	(0.89)
Euro Area	−0.05	(−0.13)	−0.04	(−0.09)	−0.06	(−0.16)
PPI	−0.26	(−0.79)	−0.28	(−0.68)	−0.23	(−0.44)
GDP	0.71**	(2.52)	0.71**	(2.24)	0.76	(1.45)
PMI	0.02	(0.06)	0.08	(0.26)	−1.36*	(−1.85)
Consumer Confidence	0.28	(0.98)	0.32	(1.04)	0.17	(0.23)
Industrial Confidence	−0.60*	(−1.71)	−0.72*	(−1.79)	−0.15	(−0.24)
Industrial Production	0.88**	(2.30)	1.03**	(2.20)	0.40	(0.91)
New Orders	−0.69	(−1.28)	−0.67	(−1.25)	—	—

(*continued*)

Table A.1. (Continued)

	Full Sample		Pre-disinflation		Below Target	
	β	**t-stats**	**β**	**t-stats**	**β**	**t-stats**
Unemployment Rate	−0.58	(−1.31)	−0.50	(−0.79)	−0.58	(−1.32)
M3	−0.19	(−0.60)	−0.01	(−0.04)	−0.97	(−1.62)
MRO	—	—	—	—	0.04	(0.35)
Retail Sales	0.05	(0.17)	0.30	(0.73)	—	—
Trade Balance	−0.05	(−0.17)	−0.12	(−0.29)	0.07	(0.19)
Mon. Policy Shock	1.51	(0.13)	5.17	(0.39)	−16.52	(−0.47)
CB Info. Shock	21.59	(1.52)	19.75	(1.21)	9.10	(0.29)
Observations	1,767		1,209		525	
$\bar{R}^2$	0.01		0.01		0.01	
R^2	0.02		0.03		0.05	
$H_0 : \beta = 0$ (p-value)	0.02		0.10		0.30	
Chow Test (p-value)	0					

Notes: The table reports estimated β coefficients from regressions on days with data releases over three periods: from January 2005 to March 2017 (full sample), from January 2005 to December 2012 (pre-disinflation), and from January 2013 to March 2017 (below target), using surprises for all variables in a joint regression. Dependent variable is one-day change, $\Delta y_t \equiv y_t - y_{t-1}$, in long-term inflation compensation (1y9y, in basis points). The surprise component of macroeconomic releases, X_t, is normalized by their historical standard deviations; coefficients represent a basis point per standard deviation response. Heteroskedasticity-consistent t-statistics are in parentheses; ***, **, and * denote a 1, 5, and 10 percent level of significance, respectively. $H_0 : \beta = 0$ p-value is for the test that all elements of β are zero (joint significance test). Chow test (p-value) is for the H_0 of parameter constancy between the pre-disinflation and below-target period. (—) indicates an omission either due to the discontinuity of the expectations collection (e.g., new orders) or for lack of surprises over most of the period of the regression (e.g., MRO).

Table A.2. Alternative Long-Term Inflation Compensation (1y9y, daily): Individual Variable Linear Regression $\Delta y_t = \alpha + \beta X_t + \varepsilon_t$

	Full Sample			Pre-disinflation			Below Target		
	β	t-stats	R^2	β	t-stats	R^2	β	t-stats	R^2
Flash									
Germany	0.40	(1.08)	0.01	−0.13	(−0.26)	0.00	1.35***	(3.41)	0.14
Spain	0.89**	(2.55)	0.04	1.08*	(1.98)	0.05	0.59	(1.59)	0.03
Italy	−0.20	(−0.75)	0.00	−0.39	(−1.30)	0.01	0.29	(0.52)	0.01
Euro Area	−0.54	(−1.57)	0.02	−0.99**	(−2.13)	0.04	0.44	(1.17)	0.03
HICP									
Germany	0.39	(1.24)	0.01	0.32	(0.98)	0.01	1.18	(1.29)	0.04
Spain	0.15	(0.20)	0.00	0.17	(0.19)	0.00	0.34	(0.44)	0.01
France	−0.54	(−0.76)	0.01	−0.67	(−0.86)	0.02	0.64	(1.10)	0.02
Italy	0.35	(0.93)	0.01	0.35	(0.74)	0.01	0.30	(0.67)	0.01
Euro Area	−0.07	(−0.18)	0.00	−0.09	(−0.20)	0.00	0.02	(0.06)	0.00
PPI	−0.24	(−0.75)	0.00	−0.26	(−0.67)	0.00	−0.19	(−0.39)	0.00
GDP	0.72***	(2.67)	0.04	0.71**	(2.36)	0.04	0.88	(1.56)	0.04
PMI	0.04	(0.12)	0.00	0.11	(0.32)	0.00	−0.35	(−0.50)	0.00
Consumer Confidence	0.19	(0.66)	0.00	0.23	(0.72)	0.00	0.20	(0.26)	0.00
Industrial Confidence	−0.57	(−1.58)	0.01	−0.67	(−1.63)	0.02	0.03	(0.05)	0.00
Industrial Production	0.86**	(2.31)	0.05	0.96**	(2.18)	0.06	0.30	(0.63)	0.01
New Orders	−0.68	(−1.29)	0.02	−0.68	(−1.29)	0.02	—	—	—
Unemployment Rate	−0.69	(−1.62)	0.02	−0.73	(−1.23)	0.02	−0.40	(−1.00)	0.02
M3	−0.21	(−0.65)	0.00	−0.05	(−0.12)	0.00	−0.94	(−1.58)	0.04
MRO	—	—	—	—	—	—	0.09	(0.42)	0.00
Retail Sales	0.04	(0.12)	0.00	0.25	(0.58)	0.00	—	—	—
Trade Balance	−0.06	(−0.19)	0.00	−0.11	(−0.25)	0.00	0.02	(0.06)	0.00
Mon. Policy Shock	−4.64	(−0.41)	0.00	2.51	(0.17)	0.00	−20.97	(−1.07)	0.07
CB Info. Shock	21.84*	(1.69)	0.01	20.10	(1.25)	0.01	25.41	(1.45)	0.05

Notes: The table reports estimated β coefficients from regressions on days with data releases over three periods: from January 2005 to March 2017 (full sample), from January 2005 to December 2012 (pre-disinflation), and from January 2013 to March 2017 (below target), using surprises for each single variable in an individual regression. Dependent variable is one-day change, $\Delta y_t \equiv y_t - y_{t-1}$, in long-term inflation compensation (1y9y, in basis points). The surprise component of macroeconomic releases, x_t, is normalized by their historical standard deviations; coefficients represent a basis point per standard deviation response. Heteroskedasticity-consistent t-statistics are in parentheses; ***, **, and * denote a 1, 5, and 10 percent level of significance, respectively. R^2 reports the explanatory variable for each individual regression. (—) indicates an omission either due to the discontinuity of the expectations collection (e.g, new orders) or for lack of surprises over most of the period of the regression (e.g., MRO).

Table A.3. Long-Term Inflation Compensation (5y5y, daily): Joint Linear Regression (with controls) $\Delta y_t = \alpha + \beta \mathbf{X}_t + \beta \mathbf{Z}_t + \varepsilon_t$

	Full Sample		**Pre-disinflation**		**Below Target**	
	β	**t-stats**	β	**t-stats**	β	**t-stats**
Flash						
Germany	0.60**	(2.33)	0.41	(1.12)	0.99***	(4.61)
Spain	0.41**	(2.10)	0.64**	(2.32)	−0.01	(−0.03)
Italy	−0.04	(−0.22)	−0.08	(−0.31)	−0.17	(−0.58)
Euro Area	0.10	(0.40)	0.19	(0.48)	−0.01	(−0.03)
HICP						
Germany	0.12	(0.66)	0.08	(0.41)	0.37	(0.79)
Spain	0.24	(1.05)	0.23	(0.85)	0.25	(0.74)
France	−0.08	(−0.49)	−0.15	(−0.88)	0.20	(0.59)
Italy	0.12	(0.67)	0.28	(1.44)	−0.38	(−1.28)
Euro Area	0.05	(0.34)	0.05	(0.29)	0.02	(0.12)
PPI	−0.02	(−0.09)	0.05	(0.18)	−0.14	(−0.52)
GDP	0.38**	(2.55)	0.42**	(2.54)	0.13	(0.39)
PMI	−0.06	(−0.30)	−0.05	(−0.28)	−0.01	(−0.03)
Consumer Confidence	0.10	(0.67)	0.03	(0.18)	0.36	(0.88)
Industrial Confidence	−0.14	(−0.73)	−0.19	(−0.85)	0.14	(0.41)
Industrial Production	0.41**	(2.57)	0.41**	(2.17)	0.33	(1.34)
New Orders	−0.27	(−1.17)	−0.25	(−1.06)	—	—
Unemployment Rate	−0.26	(−1.08)	−0.19	(−0.57)	−0.55*	(−1.84)
M3	−0.24	(−1.60)	−0.33*	(−1.94)	−0.04	(−0.14)

(*continued*)

Table A.3. (Continued)

	Full Sample		Pre-disinflation		Below Target	
	β	t-stats	β	t-stats	β	t-stats
MRO	—	—	—	—	0.35**	(2.48)
Retail Sales	0.11	(0.69)	0.09	(0.50)	—	—
Trade Balance	0.04	(0.19)	0.08	(0.26)	0.06	(0.21)
Mon. Policy Shock	7.00	(0.90)	11.52	(1.22)	−10.86	(−0.63)
CB Info. Shock	18.77	(1.57)	17.02	(1.24)	13.81	(0.66)
Citigroup EA Surprise	0.00	(0.55)	0.01	(0.93)	−0.02*	(−1.68)
Citigroup U.S. Surprise	0.02*	(1.85)	0.03*	(1.93)	0.00	(0.20)
Oil Price	0.06	(1.49)	0.00	(0.08)	0.23***	(3.17)
Gold Price	0.00	(−0.17)	0.00	(−0.09)	0.00	(−0.71)
FX EUR/USD	0.05	(0.29)	0.03	(0.14)	0.28	(0.66)
EA Bond Market Vol.	0.40*	(1.88)	0.31	(1.29)	0.35	(0.81)
VSTOXX	−0.10**	(−2.01)	−0.10*	(−1.77)	−0.06	(−0.64)
VIX	0.04	(0.80)	0.07	(1.22)	−0.11	(−1.05)
CDS	−0.09***	(−5.71)	−0.10***	(−5.61)	−0.04	(−1.21)
Observations	1,767		1,209		525	
$\bar{R}^2$	0.06		0.06		0.06	
R^2	0.08		0.09		0.11	
$H_0 : \beta = 0$ (p-value)	0.00		0.00		0.00	
Chow Test (p-value)	0.1					

Notes: The table reports estimated β coefficients from regressions on days with data releases over three periods: from January 2005 to March 2017 (full sample), from January 2005 to December 2012 (pre-disinflation), and from January 2013 to March 2017 (below target). Dependent variable is the change (basis points) in long-term inflation compensation (5y5y); the covariates consist of the surprise component of macroeconomic releases, $\mathbf{X}_t$, which is normalized by their historical standard deviations, and control variables $\mathbf{Z}_t$ that are observed at daily frequency; coefficients represent a basis point per standard deviation response. Heteroskedasticity-consistent t-statistics are in parentheses; ***, **, and * denote a 1, 5, and 10 percent level of significance, respectively. $H_0 : \beta = 0$ p-value is for the test that all elements of β are zero. Chow test (p-value) is for the H_0 of parameter constancy between the pre-disinflation and below-target period. (—) indicates an omission either due to the discontinuity of the expectations collection (e.g., new orders) or for lack of surprises over most of the period of the regression (e.g., MRO).

Table A.4. Long-Term Inflation Compensation (5y5y, two days): Joint Linear Regression $\Delta y_t = \alpha + \beta X_t + \varepsilon_t$

	Full Sample		Pre-disinflation		Below Target	
	β	t-stats	β	t-stats	β	t-stats
Flash						
Germany	0.62**	(2.34)	0.56	(1.56)	0.82**	(2.19)
Spain	0.33	(1.10)	0.62	(1.45)	−0.12	(−0.34)
Italy	0.02	(0.05)	0.07	(0.15)	−0.25	(−0.46)
Euro Area	−0.07	(−0.15)	−0.16	(−0.23)	−0.04	(−0.09)
HICP						
Germany	−0.02	(−0.07)	−0.09	(−0.30)	0.56***	(3.09)
Spain	0.17	(0.63)	0.04	(0.14)	0.91	(1.62)
France	−0.12	(−0.47)	−0.26	(−1.00)	1.05**	(2.28)
Italy	0.39	(1.45)	0.58**	(2.11)	−0.14	(−0.24)
Euro Area	0.08	(0.36)	0.13	(0.44)	0.03	(0.15)
PPI	−0.37	(1.35)	−0.44	(−1.35)	−0.18	(−0.39)
GDP	0.49*	(1.91)	0.48	(1.62)	0.58	(1.38)
PMI	0.50	(1.53)	0.47	(1.37)	1.31**	(2.35)
Consumer Confidence	0.11	(0.47)	0.15	(0.61)	−0.14	(−0.36)
Industrial Confidence	−0.25	(−0.53)	−0.27	(−0.46)	−0.31	(−0.68)
Industrial Production	0.54	(2.08)	0.64**	(2.13)	0.20	(0.48)
New Orders	0.09	(0.26)	0.10	(0.28)	—	—

(*continued*)

Table A.4. (Continued)

	Full Sample		Pre-disinflation		Below Target	
	β	t-stats	β	t-stats	β	t-stats
Unemployment Rate	−0.30	(−0.86)	−0.04	(−0.07)	−0.99***	(−2.92)
M3	−0.20	(−0.73)	−0.52*	(−1.67)	1.21**	(2.52)
MRO	—	—	—	—	0.76***	(4.37)
Retail Sales	0.04	(0.14)	−0.25	(−0.76)	—	—
Trade Balance	0.10	(0.40)	0.11	(0.30)	0.11	(0.31)
Mon. Policy Shock	20.26	(1.80)	28.27**	(2.25)	−2.86	(−0.13)
CB Info. Shock	44.83***	(2.62)	40.20***	(2.08)	44.46	(1.69)
Observations	1,767		1,209		525	
$\bar{R}^2$	0.01		0.01		0.02	
R^2	0.03		0.03		0.06	
$H_0 : \beta = 0$ (p-value)	0.00		0.04		0.00	
Chow Test (p-value)	0.00					

Notes: The table reports estimated β coefficients from regressions on days with data releases from January 2005 to March 2017 (full sample), from January 2005 to December 2012 (pre-disinflation), and from January 2013 to March 2017 (below target), using surprises for all variables in a joint regression. Dependent variable is changes (basis points) in long-term inflation compensation (5y5y) across a two-day window, $\Delta y_t \equiv y_{t+1} - y_{t-1}$. The surprise component of macroeconomic releases, $\mathbf{X}_t$, is normalized by their historical standard deviations; coefficients therefore represent a basis point per standard deviation response. Heteroskedasticity-consistent t-statistics are in parentheses; ***, **, and * denote a 1, 5, and 10 percent level of significance, respectively. $H_0 : \beta = 0$ p-value is for the test that all elements of β are zero (joint significance test). Chow test (p-value) is for the H_0 of parameter constancy between the pre-disinflation and below-target period. (—) indicates an omission either due to the discontinuity of the expectations collection (e.g., new orders) or for lack of surprises over most of the period of the regression (e.g., MRO).

Table A.5. Long-Term Inflation Compensation (5y5y, two days): Individual Variable Linear Regression$\Delta y_t = \alpha + \beta X_t + \varepsilon_t$

	Full Sample			Pre-disinflation			Below Target		
	β	t-stats	R^2	β	t-stats	R^2	β	t-stats	R^2
Flash									
Germany	0.68**	(2.58)	0.05	0.56	(1.54)	0.03	0.91**	(2.66)	0.09
Spain	0.41	(1.38)	0.02	0.45	(0.96)	0.02	0.32	(0.98)	0.02
Italy	0.04	(0.15)	0.00	0.00	(0.01)	0.00	0.12	(0.24)	0.00
Euro Area	0.05	(0.16)	0.00	−0.08	(−0.15)	0.00	0.22	(0.62)	0.01
HICP									
Germany	−0.04	(−0.15)	0.00	−0.09	(−0.31)	0.00	0.46***	(2.91)	0.01
Spain	0.22	(0.84)	0.01	0.14	(0.49)	0.00	0.68	(1.17)	0.03
France	−0.08	(−0.33)	0.00	−0.19	(−0.70)	0.00	0.73	(1.53)	0.03
Italy	0.37	(1.53)	0.02	0.53**	(2.07)	0.03	0.00	(0.01)	0.00
Euro Area	0.08	(0.36)	0.00	0.10	(0.36)	0.00	−0.05	(−0.22)	0.00
PPI	−0.33	(−1.24)	0.01	−0.41	(−1.24)	0.02	−0.12	(−0.28)	0.00
GDP	0.48**	(2.02)	0.03	0.45	(1.59)	0.02	0.76*	(1.79)	0.04
PMI	0.51	(1.53)	0.02	0.47	(1.37)	0.02	1.64**	(2.50)	0.04
Consumer Confidence	0.08	(0.30)	0.00	0.13	(0.45)	0.00	−0.10	(−0.27)	0.00
Industrial Confidence	−0.24	(−0.50)	0.01	−0.18	(−0.30)	0.00	−0.42	(−0.83)	0.01
Industrial Production	0.50	(1.99)	0.03	0.58	(2.03)	0.04	0.19	(0.43)	0.00
New Orders	0.06	(0.17)	0.00	0.06	(0.17)	0.00	—	—	—
Unemployment Rate	−0.33	(−0.92)	0.01	−0.17	(−0.37)	0.00	−0.64*	(−1.93)	0.06
M3	−0.23	(−0.80)	0.01	−0.56	(−1.80)	0.04	1.24**	(2.63)	0.12
MRO	—	—	—	—	—	—	0.56	(1.94)	0.06
Retail Sales	−0.01	(−0.02)	0	−0.45	(−1.38)	0.02	—	—	—
Trade Balance	0.12	(0.46)	0.00	0.12	(0.34)	0.00	0.10	(0.32)	0.00
Mon. Policy Shock	8.17	(0.75)	0.01	22.92*	(1.71)	0.04	−27.18*	(−1.96)	0.11
CB Info. Shock	38.20**	(2.41)	0.10	39.42*	(1.92)	0.10	42.57	(2.95)	0.14

Notes: The table reports estimated β coefficients from regressions on days with data releases over three periods: from January 2005 to March 2017 (full sample), from January 2005 to December 2012 (pre-disinflation), and from January 2013 to March 2017 (below target), using surprises for each single variable in an individual regression. Dependent variable is changes across a two-day window, $\Delta y_t \equiv y_{t+1} - y_{t-1}$, in long-term inflation compensation (5y5y, in basis points). The surprise component of macroeconomic releases, x_t, is normalized by their historical standard deviations; coefficients β therefore represent a basis point per standard deviation response. Heteroskedasticity-consistent t-statistics are in parentheses; and ***, **, * denote a 1, 5, and 10 percent level of significance, respectively. R^2 reports the explanatory variable for each individual regression. (—) indicates an omission either due to the discontinuity of the expectations collection (e.g., new orders) or for lack of surprises over most of the period of the regression (e.g., MRO).

Table A.6. Short-Term Inflation Compensation (1-year spot IL swap rate, daily): Joint Linear Regression $\Delta y_t = \alpha + \beta X_t + \varepsilon_t$

	Full Sample		Pre-disinflation		Below Target	
	β	t-stats	β	t-stats	β	t-stats
Flash						
Germany	2.14***	(5.67)	1.58***	(4.44)	3.20***	(3.83)
Spain	0.80*	(1.87)	0.71	(1.64)	0.51	(0.55)
Italy	1.21***	(2.90)	0.85**	(1.96)	1.29	(1.48)
Euro Area	1.70***	(2.43)	1.67***	(2.22)	1.06	(0.91)
HICP						
Germany	0.25	(1.36)	0.25	(1.23)	0.03	(0.06)
Spain	−0.52	(−1.63)	−0.43	(−1.25)	−0.99*	(−1.70)
France	0.87***	(3.69)	0.83***	(3.28)	1.11	(1.46)
Italy	−0.05	(−0.22)	0.15	(0.57)	−0.76	(−1.37)
Euro Area	0.52**	(2.23)	0.62**	(2.45)	0.13	(0.26)
PPI	−0.22	(−0.61)	−0.36	(−0.76)	0.38	(0.84)
GDP	0.28	(1.32)	0.26	(1.08)	−0.03	(−0.06)
PMI	3.56***	(2.84)	3.69***	(2.89)	1.25	(0.35)
Consumer Confidence	−0.41*	(−1.71)	−0.44	(−1.63)	−0.06	(−0.12)
Industrial Confidence	−0.23	(−0.72)	−0.09	(−0.25)	−1.14	(−1.43)
Industrial Production	0.29	(0.72)	0.33	(0.69)	−0.06	(−0.10)
New Orders	0.03	(0.06)	0.07	(0.14)	—	—

(continued)

Table A.6. (Continued)

	Full Sample		**Pre-disinflation**		**Below Target**	
	β	**t-stats**	β	**t-stats**	β	**t-stats**
Unemployment Rate	0.76	(0.91)	2.60***	(3.04)	−3.31**	(−2.08)
M3	0.20	(0.90)	0.19	(0.91)	0.20	(0.29)
MRO	—	—	—	—	1.09	(1.19)
Retail Sales	−0.35	(−1.41)	−0.59*	(−1.68)	—	—
Trade Balance	−0.27	(−0.93)	−0.36	(−0.94)	−0.15	(−0.38)
Mon. Policy Shock	−30.91	(−1.22)	−38.83	(−1.34)	−26.48	(−0.68)
CB Info. Shock	27.05	(1.38)	37.87*	(1.93)	8.60	(0.22)
Observations	1,767		1,209		525	
$\bar{R}^2$	0.06		0.08		0.07	
R^2	0.08		0.09		0.01	
$H_0 : \beta = 0$ (p-value)	0		0		0	
Chow Test (p-value)	0					

Notes: The table reports estimated β coefficients from regressions on days with data releases over three periods: from January 2005 to March 2017 (full sample), from January 2005 to December 2012 (pre-disinflation), and from January 2013 to March 2017 (below target), using surprises for all variables in a joint regression. Dependent variable is one-day change, $\Delta y_t \equiv y_t - y_{t-1}$, in short-term inflation compensation (1-year spot IL swap rate, in basis points). The surprise component of macroeconomic releases, $\mathbf{X}_t$, is normalized by their historical standard deviations; coefficients represent a basis point per standard deviation response. Heteroskedasticity-consistent t-statistics are in parentheses; ***, **, and * denote a 1, 5, and 10 percent level of significance, respectively. $H_0 : \beta = 0$ p-value is for the test that all elements of β are zero. Chow test (p-value) is for the H_0 of parameter constancy between the pre-disinflation and below-target period. (—) indicates an omission either due to the discontinuity of the expectations collection (e.g., new orders) or for lack of surprises over most of the period of the regression (e.g., MRO).

Table A.7. Short-Term Inflation Compensation (1-year forward IL swap rate 1-year ahead, daily): Joint Linear Regression $\Delta y_t = \alpha + \beta X_t + \varepsilon_t$

	Full Sample		Pre-disinflation		Below Target	
	β	t-stats	β	t-stats	β	t-stats
Flash						
Germany	1.35***	(5.32)	1.32***	(4.17)	1.37***	(3.27)
Spain	0.20	(0.67)	0.29	(0.73)	0.05	(0.10)
Italy	0.59**	(2.06)	0.40	(1.06)	0.86**	(1.96)
Euro Area	0.81**	(2.14)	0.90*	(1.87)	0.49	(0.85)
HICP						
Germany	0.33*	(1.65)	0.32	(1.49)	0.34	(0.78)
Spain	−0.57*	(−1.82)	−0.60	(−1.63)	−0.45	(−1.11)
France	0.87***	(3.92)	0.89***	(3.75)	0.58	(0.82)
Italy	−0.07	(−0.29)	−0.01	(−0.03)	−0.31	(−0.64)
Euro Area	−0.02	(−0.09)	0.05	(0.16)	−0.26	(−0.64)
PPI	−0.13	(−0.52)	0.06	(0.17)	−0.46	(−1.28)
GDP	−0.38	(−1.57)	−0.46*	(−1.83)	0.09	(0.23)
PMI	0.03	(0.08)	0.00	(−0.01)	0.66	(0.65)
Consumer Confidence	−0.63**	(−1.99)	−0.75**	(−2.07)	0.02	(0.05)
Industrial Confidence	−0.40	(−1.10)	−0.44	(−1.03)	−0.27	(−0.62)
Industrial Production	0.09	(0.23)	0.12	(0.27)	−0.04	(−0.08)
New Orders	−0.14	(−0.50)	−0.13	(−0.46)	—	—

(*continued*)

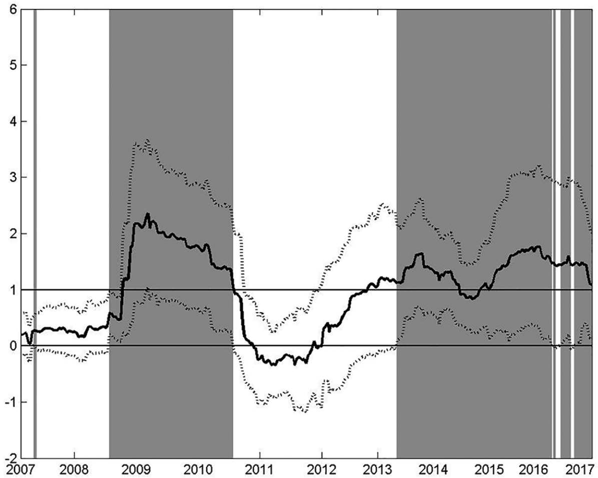
Table A.7. (Continued)

	Full Sample		Pre-disinflation		Below Target	
	β	t-stats	β	t-stats	β	t-stats
Unemployment Rate	−0.26	(−0.89)	0.16	(0.52)	−1.15**	(−2.29)
M3	0.21	(0.93)	0.21	(0.85)	0.17	(0.34)
MRO	—	—	—	—	0.38**	(2.29)
Retail Sales	−0.19	(−0.75)	−0.31	(−0.86)	—	—
Trade Balance	−0.34	(−1.39)	−0.47	(−1.42)	−0.09	(−0.31)
Mon. Policy Shock	−24.87	(−0.97)	−29.37	(−0.99)	−22.12	(−0.74)
CB Info. Shock	25.27	(1.12)	30.89	(1.36)	13.25	(0.48)
Observations	1,767		1,209		525	
$\bar{R}^2$	0.04		0.04		0.04	
R^2	0.05		0.05		0.08	
$H_0 : \beta = 0$ (p-value)	0		0		0	
Chow Test (p-value)	0					

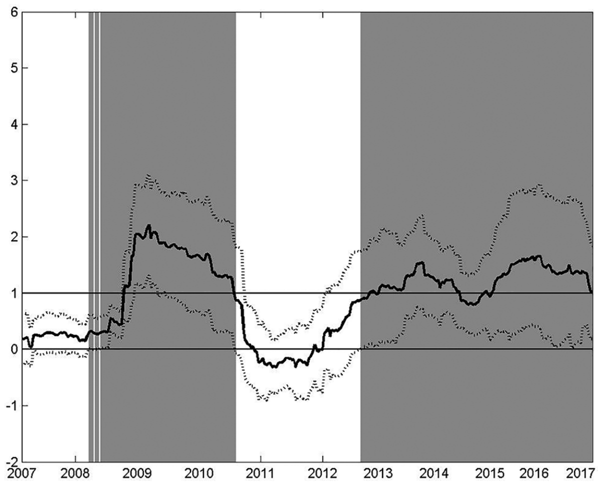
Notes: The table reports estimated β coefficients from regressions on days with data releases over three periods: from January 2005 to March 2017 (full sample), from January 2005 to December 2012 (pre-disinflation), and from January 2013 to March 2017 (below target), using surprises for all variables in a joint regression. Dependent variable is one-day change, $\Delta y_t \equiv y_t - y_{t-1}$, in short-term inflation compensation (one-year forward IL swap rate one year ahead, in basis points). The surprise component of macroeconomic releases, $\mathbf{X}_t$, is normalized by their historical standard deviations; coefficients represent a basis point per standard deviation response. Heteroskedasticity-consistent t-statistics are in parentheses; ***, **, and * denote a 1, 5, and 10 percent level of significance, respectively. $H_0 : \beta = 0$ p-value is for the test that all elements of β are zero. Chow test (p-value) is for the H_0 of parameter constancy between the pre-disinflation and below-target period. (—) indicates an omission either due to the discontinuity of the expectations collection (e.g., new orders) or for lack of surprises over most of the period of the regression (e.g., MRO).

Figure A1. Long-Term Inflation Compensation (5y5y, daily, alternative normalization): Time-Varying Sensitivity Coefficients δ^{τ} from Nonlinear Regression

A. δ^{τ} Normalized across 2005–08



B. δ^{τ} Normalized across 2008–12



Notes: This figure plots the coefficients δ^{τ} estimated from a regression $\Delta y_t = \gamma^{\tau} + \delta^{\tau}\hat{\mathbf{X}}_t + \epsilon_t^{\tau}$. δ^{τ_i} has been normalized to an average value of 1 in panel A for the years $i = 2005, \ldots, 2008$ and in panel B for the years $i = 2008, \ldots, 2012$. Dotted gray lines depict heteroskedasticity-consistent $\pm$ two-standard-error bands, adjusted for two-stage sampling uncertainty in δ^{τ}. Shaded regions denote δ^{τ} being significantly above 0 and suggest a weaker anchoring of inflation expectations.

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