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Michael Ehrmann, Gaetano Gaballo,
Peter Hoffmann, Georg Strasser

## Can more public information raise uncertainty? The international evidence on forward guidance

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**Abstract**

Central banks have used different types of forward guidance, where the forward guidance horizon is related to a state contingency, a calendar date or left open-ended. This paper reports cross-country evidence on the impact of these different types of forward guidance on the sensitivity of bond yields to macroeconomic news, and on forecaster disagreement about the future path of interest rates. We show that forward guidance mutes the response to macroeconomic news in general, but that calendar-based forward guidance with a short horizon counterintuitively raises it. Using a model where agents learn from market signals, we show that the release of more precise public information about future rates lowers the informativeness of market signals and, as a consequence, may increase uncertainty and amplify the reaction of expectations to macroeconomic news. However, when the increase in precision of public information is sufficiently large, uncertainty is unambiguously reduced.

*Keywords*: central bank communication, heterogeneous beliefs, forward guidance, disagreement, macroeconomic news

*JEL Codes*: D83, E43, E52, E58

# Non-technical Summary

In the aftermath of the global financial crisis, several central banks took policy rates to their effective lower bound, and resorted to forward guidance (FG) if they wanted to ease monetary policy further. Central banks have used different types of FG, where the horizon, over which policy rates are expected to remain at current levels, has been defined by means of i) a state-contingent threshold, ii) a calendar date, or has been iii) left open-ended.

By giving markets more precise information on the likely future path of interest rates, central banks aim to manage market expectations about the future course of policy. One implication of this is that market rates should be less responsive to incoming macroeconomic news, as documented by Swanson and Williams (2014a) and Feroli, Greenlaw, Hooper, Mishkin and Sufi (2017) for the United States, and by Swanson and Williams (2014b) for the UK and Germany.

In this paper we exploit a unique cross-country dataset (covering Canada, Germany, Italy, Japan, Sweden, the UK, and the United States) to show that the types of FG differ systematically in how they affect the link between bond yields and macroeconomic news. In line with the earlier literature, we find that time-contingent FG over a long horizon (above 1.5 years) and state-contingent FG mute the market responsiveness. Open-ended FG, in contrast, retains the original market responsiveness. More surprisingly, time-contingent FG over a short horizon (below or equal to 1.5 years) shows a perverse effect, in that it more than doubles the responsiveness to news.

The disagreement among professional forecasters about the future path of interest rates resembles the responsiveness of bond yields: Long-horizon time-contingent and state-contingent FG reduce disagreement, whereas open-ended and short-horizon time-contingent FG are ineffective.

We rationalize these findings as the unintended consequence of a market externality in information aggregation. To develop this point, we present a model where agents learn from market signals and show that the release of more precise public information about future rates can perversely increase uncertainty and the sensitivity of bond prices to public information.

In our model, agents receive noisy signals about the state of the economy, one private and one public: the latter mirrors the flow of macroeconomic news studied in the empirical analysis, the former accounts for the structural heterogeneity of agents. In addition, agents observe a noisy price signal which imperfectly aggregates expectations about the realization of the future return on a safe bond (which comoves with the policy rate). Stronger FG corresponds to a lower

dependence of policy rates on fundamentals, i.e. lowers prior uncertainty on policy rates.

In this model, the effect of FG on the sensitivity of bond prices to macroeconomic news and on forecaster disagreement depends crucially on the strength of FG. On the one hand, strengthening FG has the direct effect of reducing the prior uncertainty of agents. On the other hand, as more public information is provided to agents, there is less to learn from prices. This is due to the fact that - everything else being equal – agents' expectations become less sensitive to their private signals as the precision of public information increases. As expectations react less to private information, the price signal loses some of its informativeness. Therefore, while FG directly decreases prior uncertainty, it renders prices less informative, so that ex-post uncertainty can increase. When prices are a good source of information, this countervailing, indirect effect is stronger than the direct effect. Only if FG is sufficiently strong, the direct effect dominates.

# 1 Introduction

In the aftermath of the global financial crisis, several central banks embarked on an unprecedented easing of monetary policy. As policy rates approached the effective lower bound (ELB), they had to resort to unconventional policies. One such tool, which was employed by several central banks, is forward guidance (FG). With policy rates being constrained, further easing of monetary policy was attempted by managing expectations about the future course of policy, effectively turning communication into a central policy tool.

By providing FG, central banks strive to provide more precise information about their reaction function. The objective is typically twofold. On the one hand, FG clarifies that the central bank considers policy rates to be at, or close to, the ELB. On the other hand, it states that an *extended period* of loose policy with no rate hikes is likely to follow.

As a consequence, markets have more precise information on the likely future path of interest rates, implying that market rates should generally be less responsive to macroeconomic news. This has been documented by Swanson and Williams (2014a) and Feroli et al. (2017) for the United States, and by Swanson and Williams (2014b) for the UK and Germany. However, as we show in this paper, this need not necessarily be the case: By exploiting a unique cross-country dataset (covering Canada, Germany, Italy, Japan, Sweden, the UK, and the United States), we show that, depending on the form of FG adopted, interest rates can also become *more* responsive to macroeconomic news in comparison to a no-FG benchmark.

In practice, central banks have used different types of FG. The literature typically identifies three such types, depending on whether the FG horizon, i.e. the time period over which policy rates are expected to remain at current levels, has been defined by means of i) a state-contingent threshold, ii) a calendar date, or has been iii) left open-ended.

In this paper, we show that the types of FG differ systematically in how they affect the link between bond yields and macroeconomic news. In line with the earlier literature, we find that some forms of FG (which we consider to be "stronger") lower the reaction to macroeconomic news. This is the case, in particular, for time-contingent FG over a long horizon (above 1.5 years), which mutes the market responsiveness to macroeconomic news almost completely. Also state-contingent FG leads to a reduced responsiveness, but does not fully eliminate it. This is plausible, because markets should remain responsive to news about the macroeconomic indicators to which the FG relates (e.g. unemployment in the case of the Federal Reserve and the Bank of England).[1]

Open-ended FG, in contrast, retains the original market responsiveness, which can be interpreted as markets perceiving no change in the reaction function of the authority. More surprisingly, time-contingent FG over a short horizon (below or equal to 1.5 years) shows a perverse effect, in that it more than doubles the responsiveness to news. In general, we show that the shorter the horizon, the higher the response to macro news.

We furthermore study the effect of FG types on the disagreement among professional forecasters about the future path of interest rates, and report findings that are consistent with those for the responsiveness of bond yields. Long-horizon time-contingent and state-contingent FG effectively reduce disagreement,[2] whereas open-ended and short-horizon time-contingent FG are ineffective in this regard.

The second contribution of this paper is theoretical. We present a model where agents learn from market signals and show that the release of more precise public information about future rates can perversely increase uncertainty and the sensitivity of bond prices to public information. The key ingredient for generating this effect is learning from market prices.

To see this, consider a Bayesian model of expectation formation where agents receive exogenous signals, public and private, about the future realization of policy rates, but where there is no learning from market signals. This is the typical framework used recently in the context of the FG debate by, among others, Angeletos and Lian (2018) and Wiederholt (2014). FG can be interpreted as a decrease in prior uncertainty (unconditional volatility) about future rates, where the magnitude of this decrease depends on the strength (or type) of FG. In this setting, FG will unambiguously reduce the responsiveness of expectations to any signal (public or private) and lead to lower disagreement, independently of its strength. Accordingly, a model with exogenous signals, as typically adopted in the literature, cannot replicate our findings of increased macro-news sensitivity under short-horizon time-contingent FG.

However, the introduction of an endogenous market signal allows us to generate a non-monotonic effect of ex-ante uncertainty on the macro-news sensitivity of bond prices. In our model, agents receive noisy signals about the state of the economy, one private and one public: the latter mirrors the flow of macro news studied in the empirical analysis, the former accounts for the structural heterogeneity of agents. In addition, agents observe a noisy price

[1] Our findings are in line with the evidence provided by Femia, Friedman and Sack (2013) and Detmers (2016) for the United States.

[2] Andrade, Gaballo, Mengus and Mojon (2019) document a fall in disagreement across professional forecasters at the time of the introduction of time-contingent FG in the United States.

signal which imperfectly aggregates expectations about the realization of the future return on a safe bond (which comoves with the policy rate). We assume that the central bank determines the extent to which policy rates co-vary with the state of the economy. In this context, stronger FG corresponds to a lower dependence of policy rates on fundamentals, that is a lower prior uncertainty on policy rates. In the extreme case of "perfect" FG, policy rates and bond returns are completely detached from stochastic variations in the state of the economy and are purely deterministic (a case of no prior uncertainty).

The effect of FG on the sensitivity of bond prices to macro news and on forecaster disagreement crucially depends on the strength of FG. On the one hand, strengthening FG has a direct effect of reducing the prior uncertainty of agents. On the other hand, as more public information is made available to agents, there is less to learn from prices. This is due to the fact that agents' expectations become less sensitive to their private signals as the precision of public information increases. As expectations react less to private information, the price signal loses its informativeness.[3] Therefore, while FG directly decreases prior uncertainty, it renders prices less informative, so that ex-post uncertainty can increase. This countervailing indirect effect is stronger than the direct beneficial effect when prices are a good source of information, i.e. when the variance of noisy supply shocks is sufficiently small. However, if FG is sufficiently strong, the direct effect dominates and the predictions are identical to those obtained from a model without price signals.

These predictions of the model are in line with our empirical findings. A small decrease in prior uncertainty makes bond prices more reactive to public signals, as observed for weak forms of FG. On the contrary, a large decrease in prior uncertainty implies a lower responsiveness to public macro news, consistent with our findings for strong forms of FG.

The remainder of this paper is organized as follows. The next section provides an overview of the related literature. Section 3 presents the empirical findings, and section 4 the theoretical model. Section 5 concludes.

# 2 Literature Review

The empirical evidence on the effectiveness of FG is summarized in Moessner, Jansen and de Haan (2017). While FG is overall judged to be an effective tool, not all results are entirely

[3]Learning from prices requires the presence of private signals; public signals are observed by everyone, so there is nothing further to learn about them from prices.

conclusive. For instance, the FG employed by the U.S. Federal Reserve has been judged as effective by Campbell, Evans, Fisher and Justiniano (2012), Moessner (2013, 2015) and Woodford (2013), whereas Filardo and Hofmann (2014) cast a more cautious tone. That different studies come to different conclusions is not too surprising, for at least two reasons. First, identification is not trivial, given that central banks often employed a variety of unconventional tools together with FG. Second, while theory typically assumes that the central bank commits to a future path of policy rates (Eggertsson and Woodford, 2003), FG was in practice probably closer to what Campbell et al. (2012) call "delphic", where the central bank provides a forecast of its future policy rates and stresses the conditionality of the forecast. Such FG has been found to generate smaller effects (Adam and Billi, 2006, 2007; Eggertsson and Woodford, 2006; Nakov, 2008).

Despite the mixed evidence on its effectiveness, FG is generally considered to be an effective tool by central bankers and academic economists alike. The survey by Blinder, Ehrmann, de Haan and Jansen (2017) shows that more than 70% of central bank governors and more than 85% of academics think that FG should remain an instrument in the central banks' toolkit.[4]

FG is an essential ingredient of the optimal policy commitment at the ELB advocated by Krugman (1998) and Eggertsson and Woodford (2003). Campbell et al. (2012) refer to this type of FG as *odyssean*, as opposed to *delphic*.[5] In our paper, this distinction does not matter to the extent that FG announcements lower the covariance of states of the economy with future rates, independently of whether this guidance relies on a commitment or on a prolonged binding of the ELB.

The seminal work by Morris and Shin (2002) presents a case where a release of public information can reduce welfare. Angeletos and Pavan (2007) clarify that this arises because of misaligned incentives between individuals and the social planner in the use of information. In this class of models, in contrast to our model, public information unambiguously reduces the ex-post uncertainty of agents; however, the individual use of such enhanced knowledge is socially inefficient.

The paper closest to our theory is Amador and Weill (2010). They model an economy where

[4]This broad agreement on the overall merits of FG between practitioners and academics masks disagreement over the way FG should be implemented. Time-contingent FG is liked least by both groups. State-contingent FG is the preferred type among academics by a large margin. In contrast, the favorite type among central bank heads is open-ended FG.

[5]*Delphic* FG remains by far the most relevant case in practice. For a nice overview of the FG debate among scholars and practitioners see den Haan (2013). Andrade et al. (2019) present a model where agents are confused about the nature – delphic or odyssean – of public announcements.

agents learn from prices and where social welfare is inversely related to agents' ex-post uncertainty. They show that social welfare can decrease, i.e. uncertainty can increase, as a consequence of more precise prior information because, as in our model, public information crowds out the aggregation of private information. We are able to generate the same insight using a much simpler asset pricing model where agents learn from market signals. In our setting, the noisy processing of this endogenous signal captures cognitive limitations or costly information processing, in the vein of a recent literature on rational inattention (see Mackowiak, Matejka and Wiederholt, 2018; Vives and Yang, 2017). Our approach allows us to solve analytically for a threshold at which the effect of increasing the prior's precision on agents' uncertainty changes its sign. Moreover, in line with our empirical work, we focus on the price sensitivity to public signals, as well as disagreement among investors. The international evidence on FG represents a useful laboratory to test the effects of public information releases.

Gaballo (2016) studies the link between FG and rational inattention. He presents a dynamic model of learning from prices in which the precision of the *prior* – reflecting the entire prior information and not only the precision of the signals – is endogenous to the use of information. While public announcements induce agents' information sets to account for a larger share of price volatility, such announcements also increase the overall level of price volatility. When the latter effect dominates, agents commit larger – rather than smaller – forecast errors.

In response to the so-called "FG puzzle" put forward by Del Negro, Giannoni and Patterson (2015), several recent papers (Angeletos and Lian, 2018; Farhi and Werning, 2017; Gabaix, 2016; McKay, Nakamura and Steinsson, 2016; Wiederholt, 2014) introduce motives for discounting in the Euler equation to prevent the explosive forward-looking behavior predicted by standard models in the absence of monetary stabilization. Other work investigates the role of imperfect information at the ELB.[6] In particular, Wiederholt (2014) argues that FG can be detrimental because it reveals bad news to otherwise imperfectly informed agents. Angeletos and Lian (2018) show that informational frictions can solve the "FG puzzle". Our work, while sharing the approach of blurring agents' information, shows that FG can amplify the news-sensitivity of asset prices by introducing market signals with endogenous precision.

[6]For example Bianchi and Melosi (2017, 2018); Kiley (2016); Michelacci and Paciello (2017).

# 3 Forward Guidance, Macro News and Disagreement: Cross-Country Evidence

FG has been implemented in many different ways. In this section, we classify FG used by central banks into three types. We then introduce our methodology to study the effectiveness of FG, and by applying it to recent FG periods we then show that the effect of FG differs substantially between FG types.

## 3.1 Forward Guidance Types

Central banks have used different types of FG. These can differ in how they affect the expectations of agents about the future course of policy, for example by signalling different degrees of commitment, or by differing in their clarity. We distinguish three types:

1. *Open-ended guidance* (*OG*): Purely qualitative statements about the policy path
2. *Time-contingent guidance* (*TG*): Statements about the policy path with an explicit reference to a calendar date
3. *State-contingent guidance* (*SG*): Statements about the policy path that are conditional on economic outcomes

An example of open-ended (or "purely qualitative") FG is the ECB's statement "we expect the key ECB interest rates to remain at present or lower levels for an extended period of time", used between July 2013 and January 2016. While such statements on the expected rate path imply less of a risk for the credibility of the central bank, they might also be less effective, as they can easily be interpreted as being vague or not containing any commitment. We would therefore expect that open-ended FG has only small effects.

Time-contingent (or "calendar-based") FG expresses the likely future path of the policy instrument as a function of calendar time. Within this category, different formulations have been used, varying the degree of commitment. The Bank of Canada, for example, used time-contingent FG from April 2009 until April 2010, with a relatively strong formulation stating that "conditional on the inflation outlook, it *commits* [emphasis added] to hold the current policy rate until the end of the second quarter of 2010". In contrast, in its statements between

August and December 2011, the Federal Open Market Committee (FOMC) said "The Committee currently *anticipates* [emphasis added] that economic conditions [...] are likely to warrant exceptionally low levels for the federal funds rate at least through mid-2013". Whereas the Canadian communication explicitly refers to a commitment (which, however, is not unconditional), the U.S. example does not. We classify both as time-contingent FG, because there is an explicit reference to a date for the expected lift-off of policy rates.

State-contingent (or "data-based") FG states how the policy path depends on economic conditions. For example, in its December 2012 statement the FOMC communicated that its low policy rates were "appropriate at least as long as the unemployment rate remains above 6 1/2 percent, inflation between one and two years ahead is projected to be no more than a half percentage point above the Committee's 2 percent longer-run goal, and longer-term inflation expectations continue to be well anchored." The beauty of this type of FG is that the expected time of lift-off responds endogenously to new economic developments. At the same time, this type of FG creates a trade-off between simplicity and accuracy: On the one hand, if the central bank provides a relatively simple state contingency that is easy to communicate, its message might turn out to be too simplistic in the end, requiring the bank to deviate from it. This was the case for the FOMC, which saw unemployment rates drop below 6.5% before it felt that an increase in policy rates was called for – partly due to a large decrease in labour force participation. This led the FOMC to remove the unemployment threshold from its FG. On the other hand, if the central bank lists a multitude of indicators to be considered, accurate and intelligible communication of the contingency might prove impossible, especially if the different indicators point in opposite directions.

Based on this classification, we collect and classify FG statements from central banks' monetary policy press releases, starting with Japan back in the year 2000, covering the time period until November 2016. The ECB, for example, used open-ended FG starting in July 2013. Since March 2016, the ECB's FG about policy rates has been explicitly linked to the duration of its asset purchase programme (APP). Because the APP itself has an explicitly stated expected minimum duration, we classify the ECB's FG since then as time-contingent. Central banks from nine countries used FG at some point in time during the past two decades. Table 1 shows that, at some point in time, three of them followed state-contingent FG (column 1), five open-end FG (column 2), and six followed time-contingent FG (column 3).

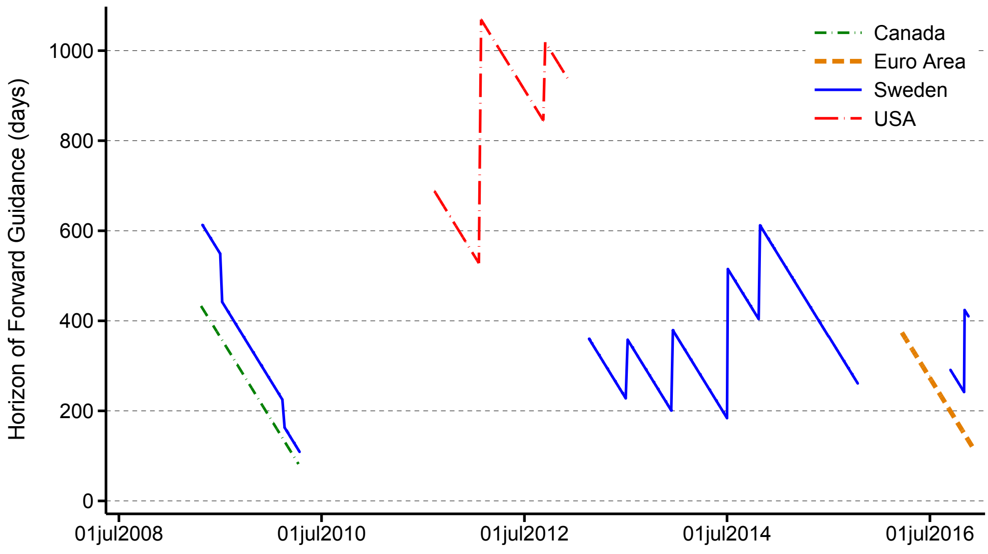
The average (remaining) horizon of time-contingent FG varies greatly across countries (Ta-

Table 1: FG episodes

| | (1) state cont. (periods) | (2) open end (periods) | (3) time contingent (periods) | (4) time contingent avg. horizon $\tilde{g}_t^c$ (# mts) |
|---|---|---|---|---|
| Canada | - | - | 09-10 | 8 |
| Euro area | - | 13-16 | 16+ | 8 |
| UK | 13-14 | 14+ | - | - |
| Sweden | - | - | 09-10, 13-15, 16+ | 12 |
| Japan | 10-13 16+ | 99-00 | - | - |
| United States | 12-14 | 08-11 14+ | 11-12 | 28 |
| Hungary | - | 14+ | - | - |
| Chile | - | - | 09-10 | 5 |
| Poland | - | - | 13-14 | 4 |

Notes: Years with FG in selected countries. A "+" indicates that the respective FG regime was active in November 2016, i.e. at the end of the sampling period.

Figure 1: Remaining guidance horizon of time-contingent FG




Notes: The graph shows the remaining horizon of the announced time contingency at each point in time. In the absence of time-contingent FG, no line is shown.

ble 1, column 4). On the one side of the spectrum are Chile and Poland with an average horizon of four and five months, respectively. On the other end of the spectrum are the United States with an average horizon of more than two years. Canada, the euro area, and Sweden are intermediate cases with eight and twelve months.

In what follows, we restrict our sample to advanced economies, and to the time periods when their policy rates are at the ELB.[7] Figure 1 plots the remaining guidance horizon of time-contingent FG in our sample. It shows that the remaining guidance horizon is often subject to revisions, which the graph shows as jumps. We observe two instances where the guidance horizon was shortened – both in Sweden during its first FG period –, as well as several episodes where the horizon was lengthened, and at times repeatedly so. Many of these extensions appear to reflect a readjustment in order to keep the remaining guidance close to a desired horizon. In the United States, for example, the guidance horizon oscillated around 950 days in 2012 and in Sweden around 300 days during 2013 and the first half of 2014. Furthermore, the chart reveals that the United States abandoned time-contingent FG long before the end of the previously announced guidance horizon. Likewise, in Sweden time-contingent FG was – ex-post temporarily – discontinued in October 2015 before the end of the guidance horizon, but resumed with a longer horizon later on.

Based on this information, we generate a binary indicator variable for each of the three types of FG. The indicator for state-contingent FG is $SG_t^c$, the indicator for time-contingent FG $TG_t^c$, and the one for open-ended FG $OG_t^c$. Each indicator is one if the respective FG regime is active in country $c$ at time $t$, and zero otherwise.

## 3.2 Data and Methodology

In this section we describe the key data series, in particular macroeconomic news surprises, bond yields, and forecaster disagreement. We also discuss the econometric models that we employ.

[7] We define these to be periods where the policy rate is at or below 1%, given that several central banks for some time perceived this to be the lower bound. The restriction to ELB periods does not change the results, as time-contingent FG was only implemented at the ELB. In appendix A.3, we show the robustness of our findings to including off-ELB periods.

### 3.2.1 Macroeconomic News

Our main analysis covers the seven advanced countries Canada, Germany, Italy, Japan, Sweden, UK, and United States. We define the macroeconomic news surprise $s_t^{c,i}$ at release time $t$ as the difference between actual and expected values, divided by the standard deviation of the time series for the respective country $c$, indicator $i$, and release step $r$.

$$s_t^{c,i} \equiv s_t^{c,i,r} = \frac{a_t^{c,i,r} - e_t^{c,i,r}}{\sigma^{c,i,r}} \tag{1}$$

The initially released values of macroeconomic indicators $a_t^{c,i,r}$ as well as the market expectations $e_t^{c,i,r}$ at that time are available from Bloomberg. Some macroeconomic indicators are announced in several pre-scheduled steps, $r$. The release of GDP, for example, typically follows a sequence of data releases based on increasingly comprehensive data. We treat each of these releases as a separate announcement event of macroeconomic indicator $i$, but standardize the surprises of each step separately. Further, we do not distinguish announcements based on their offset to their reference period. If multiple reference periods are announced simultaneously (which is uncommon), then we take the net surprise over all these reference periods.

We focus on nine macroeconomic indicators for which market expectations are available for most countries in our sample over a long time span. We exclude indicators without a significant asset price impact in any country during our sample period.[8] Our sample consists of business confidence indices (BCI), consumer confidence indices (CCI), consumer price indices (CPI), GDP growth (GDP), industrial production (IPI), non-farm payroll employment (NFP, available for the US only), purchasing manager indices (PMI), retail sales (RS), and unemployment rates (UR). Column 2 of Table 2 shows that, on average, we have about seven indicators per country.

The sovereign bonds traded in the euro area are bonds of individual euro area countries. These bonds respond more to national macroeconomic news than to euro area aggregates, because national statistics are often available first. For these reasons, we use national bonds and national macroeconomic news in the yield regressions. We choose two countries to represent the euro area, namely Germany and Italy, to reflect the diversity of the euro area and to have a similar number of news releases for the euro area as for the United States and Japan.

[8]Our criterion is the significance of $\beta^{c,i}$ in the auxiliary regressions $y_t^{c,i} = \alpha^{c,i} + \beta^{c,i} s_t^{c,i} + \varepsilon_t^{c,i}$, where $y_t^{c,i}$ is the yield change of bonds with a residual maturity of two years. In this regression we use both ELB and off-ELB periods. We exclude an indicator if $\beta^{c,i}$ is insignificant at the 5% level for all countries in both the daily and the intraday specification. Based on this criterion we exclude the indicators durable goods orders, housing starts and incoming orders.

Table 2: Summary statistics

| | | (1) ELB period | (2) Yield regressions # of indicators | (3) Yield regressions start | (4) Yield regressions # obs @ELB | (5) Disagreement start | (6) Disagreement # obs @ELB |
|---|---|---|---|---|---|---|---|
| Canada | | 2009+ | 6 | 2009 | 399 | 2009 | 93 |
| Euro area | Germany | 2009+ | 8 | 2009 | 573 | 2009 | 81 |
| | Italy | | 8 | 2009 | 601 | | |
| Japan | | 1995+ | 7 | 2000 | 1005 | 1999 | 214 |
| Sweden | | 2009-10, 12+ | 8 | 2009 | 478 | 2009 | 68 |
| UK | | 2009+ | 7 | 2009 | 594 | 2009 | 93 |
| United States | | 2003-04, 08+ | 9 | 2003 | 1092 | 2003 | 108 |
| | | | | | 4742 | | 657 |

Notes: The table reports summary statistics for the time periods when the various countries were at the ELB (given in column (1)). Columns (2)-(4) list the beginning of the sample period, the number of indicators and the number of observations underlying the daily bond yield regressions. Columns (5) and (6) list the beginning of the sample period and the number of observations underlying the disagreement regressions.

We sign surprises so that a positive surprise is good news about the economy, which (via a more restrictive monetary policy) is likely to imply higher yields. For this reason, we invert the sign of the surprise in the unemployment report.

We are interested in a causal effect of FG on the responsiveness of bond yields. However, in the vast majority of cases, FG has been implemented by central banks that had reached the ELB, or they had reached what they perceived the ELB to be at the time.[9] If interest rates are at their lower bound, the impact of macroeconomic surprises on bond yields might well be muted – if only because negative news cannot lead to a further downward move in policy rates. To separate the effects of the ELB from those stemming from FG, we restrict our sample to ELB periods. We define these to be periods where the policy rate is at or below 1%. Based on this definition, all countries listed in Table 2 were in this ELB regime at some time during the sample period. Column (1) of Table 2 lists the years at the ELB.

Column (4) of Table 2 reports the number of macroeconomic indicator releases that we use during ELB periods in each country. The observations for most countries are at or after the year 2009, with the two exceptions Japan (earliest observation in the year 2000) and United States (earliest observation in the year 2003).

43% of macroeconomic news releases in our daily sample occur in times when central banks

[9]Note that this perception changed over time, with estimates of the ELB moving downwards considerably.

are at the ELB, but have not provided FG. Open-ended FG was in place for 37% of observations, whereas time-contingent FG and state-contingent FG was in place for to 13% and 7% of observations, respectively.

#### 3.2.2 Bond Yields

We use daily as well as intradaily bond yield data. Daily bond yields are from Datastream. We use data of sovereign bonds with remaining maturities of one and two years. The daily yields are estimates taken from the yield curve. We calculate daily bond returns $y_t^c$. On Mondays we use the Friday-to-Monday return.

Yield changes at minute-by-minute frequency are based on mid-quotes. We obtain the underlying indicative bid and ask quotes for benchmark bonds from the Thomson Reuters Tick History, using the last tick for each minute. For our benchmark results, we use sovereign bonds with a residual maturity of two years. We calculate bond returns in a two-hour window from 60 minutes before until 60 minutes after the announcement release time.

#### 3.2.3 Forecaster Disagreement

To assess the disagreement among professional forecasters, we use one-year ahead forecasts of three-month interest rates, because these are a close proxy for expectations about policy rates. The data are provided by Consensus Economics. For robustness, we also study forecasts for 10-year government bond yields. Consensus Economics forecast data are particularly suited for our analysis, as they are available at a monthly frequency for a sufficiently long history in a comparable fashion across countries.[10]

To study disagreement, we follow Ehrmann (2015) and use the interdecile range of forecasts in a given country and month. The advantage of this measure over the standard deviation is that it is insensitive to outliers, which might be important in the analysis of survey data.[11] The individual forecaster data for the construction of our disagreement measure covers the same set of countries as the bond yield regressions. However, given the existence of a genuine euro area disagreement measure, we use this rather than data for Germany and Italy separately, as we do in the bond yield regressions. Table 2 shows that this analysis is based on 657 observations.

[10]The data have been used in several other studies, such as Crowe (2010), Dovern, Fritsche and Slacalek (2012), Ehrmann, Eijffinger and Fratzscher (2012), Davis and Presno (2014) or Ehrmann (2015).

[11]Furthermore, by using the interdecile range instead of the interquartile range (as in Mankiw, Reis and Wolfers (2004) or Dovern et al. (2012)), we potentially incorporate a broader range of views while still being robust to outliers.

### 3.2.4 Methodology

If FG were effective in managing expectations about the future course of monetary policy, markets would generally be less responsive to macroeconomic news. This has been documented by Swanson and Williams (2014a) and Feroli et al. (2017) for the United States, and by Swanson and Williams (2014b) for the UK and Germany. At the same time, the responsiveness to news might very well depend on the FG specification in place. On the one hand, some types of FG might be less credible than others, therefore leaving markets relatively more responsive than under a highly credible FG. On the other hand, state-contingent FG explicitly conditions the future path of interest rates on economic developments, therefore leaving expectations about future interest rates responsive to macroeconomic developments. In contrast, under credible time-contingent FG, bonds maturing during the FG horizon should in principle not respond to macroeconomic news at all.

To shed light on the effectiveness of different types of FG, we rely on an event-study setup (Balduzzi, Elton and Green, 2001). In our baseline specification we examine how the three types of FG differ in muting the impact of macroeconomic surprises $s_t^{c,i}$ on bond yields. We start from a first specification, which looks for the overall effect of FG on bond yield changes $y_t^{c,i}$:

$$y_t^{c,i} = \alpha^{c,i} + \alpha_{FG} FG_t^c + \beta s_t^{c,i} + \beta_{FG} FG_t^c \times s_t^{c,i} + \varepsilon_t^{c,i}. \tag{2}$$

The binary indicator $FG_t^c$ equals whenever some form of FG is provided in country $c$ at time $t$. In this fixed-effects specification, the error term $\varepsilon_t^{c,i}$ is assumed to be independent across country-indicator clusters. The coefficient $\beta$ captures the average impact of a macroeconomic surprise on bond yields outside of FG episodes. The coefficient $\alpha_{FG}$ captures possible bond market trends specific to the FG period. The coefficient of primary interest is $\beta_{FG}$. Its size relative to $\beta$ captures the differential effect of FG.

Subsequently, we expand this specification to allow for different types of FG:

$$\begin{aligned} y_t^{c,i} &= \alpha^{c,i} + \alpha_{SG} SG_t^c + \alpha_{OG} OG_t^c + \alpha_{TG} TG_t^c \\ &+ \beta s_t^{c,i} + \beta_{SG} SG_t^c \times s_t^{c,i} + \beta_{OG} OG_t^c \times s_t^{c,i} + \beta_{TG} TG_t^c \times s_t^{c,i} + \varepsilon_t^{c,i}. \end{aligned} \tag{3}$$

The coefficients of interest now are $\beta_{SG}$, $\beta_{TG}$ and $\beta_{OG}$. In addition, we test whether the horizon of time-contingent FG relative to the maturity of the bond is important: For bonds with $m$ years to maturity we define the time-to-maturity coverage ratio as $g_t^c = min(\frac{\tilde{g}_t^c}{365 \times m}, 1)$, where $\tilde{g}_t^c$ is the residual horizon of the time-contingent FG in country $c$ at time $t$, measured in calendar days. This leads to our third specification:

$$\begin{aligned} y_t^{c,i} &= \alpha^{c,i} + \alpha_{SG} SG_t^c + \alpha_{OG} OG_t^c + \alpha_{TG} TG_t^c \\ &+ \beta s_t^{c,i} + \beta_{SG} SG_t^c \times s_t^{c,i} + \beta_{OG} OG_t^c \times s_t^{c,i} + \beta_{TG} TG_t^c \times s_t^{c,i} \\ &+ \gamma TG_t^c \times g_t^c + \gamma_g TG_t^c \times g_t^c \times s_t^{c,i} + \varepsilon_t^{c,i}. \end{aligned} \tag{4}$$

Here, the coefficient $\beta_{TG} + \gamma_g$ measures the reduction of the asset price impact $\beta$ if FG covers the entire time to maturity (or more). Alternatively, in order to get results that are easier to interpret, we distinguish two ranges of time-contingent FG horizons. We define a long horizon to be above 1.5 years and denote this with the indicator variable $LTG_t^c$, and a short remaining horizon of less than or equal to 1.5 years (with indicator variable $STG_t^c$). With this, we arrive at our benchmark specification

$$\begin{aligned} y_t^{c,i} &= \alpha^{c,i} + \alpha_{SG} SG_t^c + \alpha_{OG} OG_t^c + \alpha_{STG} STG_t^c + \alpha_{LTG} LTG_t^c \\ &+ \beta s_t^{c,i} + \beta_{SG} SG_t^c \times s_t^{c,i} + \beta_{OG} OG_t^c \times s_t^{c,i} \\ &+ \beta_{STG} STG_t^c \times s_t^{c,i} + \beta_{LTG} LTG_t^c \times s_t^{c,i} + \varepsilon_t^{c,i}. \end{aligned} \tag{5}$$

Another way to test the effectiveness of FG in managing expectations is to study its impact on forecaster disagreement. Andrade et al. (2019) have shown that under FG, forecaster disagreement about future interest rates is reduced, although disagreement about the future macroeconomic outlook has increased. We extend their analysis to see whether these effects differ depending on the type of FG. For this purpose, we estimate the model

$$\Omega_t^c = \alpha^c + \alpha_t + \alpha_{FG} FG_t^c + \varepsilon_t^c, \tag{6}$$

where $\Omega_t^c$ is the interdecile range of one-year ahead forecasts of three-month interest rates in country $c$, as provided in the Consensus Economics forecast conducted in month $t$. $\alpha^c$ denotes country fixed effects, and $\alpha_t$ time fixed effects. As before, we restrict the sample to ELB periods (which we define to be periods where the policy rate is at or below 1%), because the disagreement and the impact of FG could well be different if interest rates are at the lower bound. We calculate Driscoll and Kraay (1998) standard errors, which allow for heteroscedasticity, autocorrelation up to a maximum lag order of 12, and cross-sectional correlation. $\alpha_{FG}$

is our parameter of interest, as it informs us how disagreement under FG compares to periods without FG. We would expect $\alpha_{FG} < 0$.

This regression model is extended in an analogous way to the regressions for bond yields by i) differentiating the different types of FG, ii) adding the horizon of time-contingent FG relative to the maturity of the forecast, and iii) subsequently differentiating long-horizon and short-horizon time-contingent FG, leading to the final specification

$$\Omega_t^c = \alpha^c + \alpha_t + \alpha_{SG} SG_t^c + \alpha_{OG} OG_t^c + \alpha_{STG} STG_t^c + \alpha_{LTG} LTG_t^c + \varepsilon_t^c. \tag{7}$$

## 3.3 The Effectiveness of Forward Guidance

In this section, we provide evidence on how the effectiveness of FG in anchoring expectations depends on the specification of FG. We first study the responsiveness of bond yields to macroeconomic surprises, and then disagreement among economic forecasters.

### 3.3.1 Bond Yields

The following tables report the net surprise impact of macroeconomic announcements based on the coefficient estimates of the regression models introduced in section 3.2.4. In Table 3, we first analyse our daily sample of two-year sovereign bonds, which allows us to explore a cross-section of seven advanced economies.

We start with the sobering observation that FG overall did not significantly change the impact of a surprising macroeconomic announcement. This observation is documented in column (1) of Table 3, which reports the estimates of our baseline regression (2). The selected macroeconomic indicators do significantly affect bond prices, but this does not change with FG, as the incremental effect of $\beta_{FG}$ is not statistically significant.

Distinguishing three types of FG, as in regression (3), allows us to provide a more nuanced perspective in column (2). Open-ended FG has no effect, as its overall coefficient ($\beta + \beta_{OG}$) is not statistically different from the coefficient of no FG ($\beta$). As expected, state-contingent FG reduces the asset price response, with the sensitivity decreasing by about one half. As the policy path is contingent on macroeconomic variables, bond prices remain sensitive to news, but to a lower degree. This occurs because other macroeconomic indicators, which a given SG does not condition on, remain indirectly relevant as predictors of the conditioning variable. However, their relevance decreases and with it the market response to them (see also Detmers, 2016).

Table 3: Net surprise impact (daily data)

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| No FG ($\beta$) | $0.465^{***}$ | $0.464^{***}$ | $0.465^{***}$ | $0.465^{***}$ |
| | *(0.168)* | *(0.168)* | *(0.168)* | *(0.168)* |
| FG ($\beta + \beta_{FG}$) | $0.518^{***}$ | | | |
| | *(0.159)* | | | |
| SG ($\beta + \beta_{SG}$) | | $0.226^{*}$ | $0.223^{*}$ | $0.223^{*}$ |
| | | *(0.122)* | *(0.123)* | *(0.123)* |
| OG ($\beta + \beta_{OG}$) | | $0.424^{*}$ | $0.423^{*}$ | $0.422^{*}$ |
| | | *(0.231)* | *(0.231)* | *(0.231)* |
| TG ($\beta + \beta_{TG}$) | | $0.920^{***}$ | $1.949^{***}$ | |
| | | *(0.212)* | *(0.497)* | |
| $g$ ($\gamma_g$) | | | $-1.815^{***}$ | |
| | | | *(0.583)* | |
| STG ($\beta + \beta_{STG}$) | | | | $1.252^{***}$ |
| | | | | *(0.256)* |
| LTG ($\beta + \beta_{LTG}$) | | | | 0.084 |
| | | | | *(0.089)* |
| # observations | 4742 | 4742 | 4742 | 4742 |
| $R^2$ | 0.01 | 0.01 | 0.01 | 0.01 |

Notes: Dependent variable is the daily change in two-year sovereign bond yields in basis points. Fixed effects model (2) in column (1), (3) in column (2), (4) in column (3), and (5) in column (4). Country-indicator fixed effects and FG fixed effects not reported. SG denotes state-contingent FG, OG open-ended FG, TG time-contingent FG, LTG time-contingent FG with a remaining guidance horizon of more than 1.5 years, STG time-contingent FG with a remaining guidance horizon of less than or equal to 1.5 years. $g$ measures the horizon of time-contingent FG relative to the maturity of the bond. Standard errors clustered at the country-indicator level in parentheses. Asterisks indicate the level of significance, (*) at the 10%, (**) at the 5%, and (***) at the 1% level.

Time-contingent FG, in contrast, *amplifies* the response of bond yields to macroeconomic news. The observation that $\beta_{TG}$ is positive and statistically significant begs the question "How can bond prices become more sensitive to news in presence of a guidance that should mute this link?"

As a step towards resolving this puzzle, we allow the effect to vary with the residual FG horizon following equation (4). The result in column (3) shows that a longer guidance horizon significantly reduces the announcement impact. For guidance horizons of 20 months or more, the yields respond less than under no FG.

Based on this observation, we distinguish short-horizon time-contingent FG from long-horizon time-contingent FG, with a cutoff at the 1.5 years guidance horizon, as in regression (5). Column (4) of Table 3 shows that long-horizon time-contingent FG does not only mute, but even entirely eliminates the asset price response. Short-horizon time-contingent FG, however, has a perverse effect. On average it triples the response of bond yields to macroeconomic news. This suggests that short-horizon time-contingent FG de-anchors expectations. In Section 4 we will rationalize this counterintuitive effect in a theoretical model that sheds light on the circumstances under which FG can effectively anchor expectations, and when not.

The results based on intraday data are very similar. Table 4 shows that the conclusions are identical to the daily sample, with lower standard errors due to the narrower event window allowing for more precise estimation. The net surprise impact of state-contingent FG, and of short-term and long-term time-contingent FG are significantly different from the surprise impact during no FG, whereas open-ended FG shows no significant difference. The muting effect of state-contingent FG is very strong.

In summary, the four types of FG differ systematically in their effectiveness. Long-horizon time-contingent FG mutes the market responsiveness to macroeconomic news (almost) completely. State-contingent FG lowers it, but does not fully eliminate it. Open-ended FG has no effect, whereas short-horizon time-contingent FG amplifies the responsiveness. This suggests that time-contingent FG with long horizons has been sufficiently credible to shift market perceptions about the central bank's reaction function, a finding that is in line with the evidence provided by Femia et al. (2013) for the United States. State-dependency seems to have gone some way in this direction: it muted the market impact of macroeconomic news, but not completely. Open-ended FG, in contrast, retains the original market responsiveness, which can be interpreted as markets perceiving this FG to be Delphic (i.e., the regular central bank reaction

Table 4: Net surprise impact (intraday)

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| No FG ($\beta$) | 0.497$^{***}$ | 0.497$^{***}$ | 0.497$^{***}$ | 0.497$^{***}$ |
| | *(0.135)* | *(0.135)* | *(0.135)* | *(0.135)* |
| FG ($\beta + \beta_{FG}$) | 0.485$^{***}$ | | | |
| | *(0.122)* | | | |
| SG ($\beta + \beta_{SG}$) | | 0.112 | 0.110 | 0.110 |
| | | *(0.070)* | *(0.070)* | *(0.070)* |
| OG ($\beta + \beta_{OG}$) | | 0.466$^{**}$ | 0.465$^{**}$ | 0.464$^{**}$ |
| | | *(0.183)* | *(0.183)* | *(0.183)* |
| TG ($\beta + \beta_{TG}$) | | 0.809$^{***}$ | 1.633$^{***}$ | |
| | | *(0.131)* | *(0.251)* | |
| $g$ ($\gamma_g$) | | | -1.453$^{***}$ | |
| | | | *(0.302)* | |
| STG ($\beta + \beta_{STG}$) | | | | 1.080$^{***}$ |
| | | | | *(0.141)* |
| LTG ($\beta + \beta_{LTG}$) | | | | 0.127$^{*}$ |
| | | | | *(0.070)* |
| # observations | 4739 | 4739 | 4739 | 4739 |
| $R^2$ | 0.03 | 0.03 | 0.03 | 0.03 |

Notes: Dependent variable is the 120-minute window change in two-year sovereign bond yields in basis points. Fixed effects model (2) in column (1), (3) in column (2), (4) in column (3), and (5) in column (4). Country-indicator fixed effects and FG fixed effects not reported. SG denotes state-contingent FG, OG open-ended FG, TG time-contingent FG, LTG time-contingent FG with a remaining guidance horizon of more than 1.5 years, STG time-contingent FG with a remaining guidance horizon of less than or equal to 1.5 years. $g$ measures the horizon of time-contingent FG relative to the forecast horizon. Standard errors clustered at the country-indicator level in parentheses. Asterisks indicate the level of significance, (*) at the 10%, (**) at the 5%, and (***) at the 1% level.

Table 5: FG and forecaster disagreement

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| FG ($\alpha_{FG}$) | -0.105$^{*}$ | | | |
| | *(0.057)* | | | |
| SG ($\alpha_{SG}$) | | -0.184$^{*}$ | -0.209$^{**}$ | -0.245$^{**}$ |
| | | *(0.104)* | *(0.100)* | *(0.099)* |
| OG ($\alpha_{OG}$) | | -0.031 | -0.058 | -0.079 |
| | | *(0.121)* | *(0.114)* | *(0.108)* |
| TG ($\alpha_{TG}$) | | -0.117 | 0.353 | |
| | | *(0.153)* | *(0.278)* | |
| $g$ ($\gamma_g$) | | | -0.632 | |
| | | | *(0.435)* | |
| STG ($\alpha_{STG}$) | | | | 0.044 |
| | | | | *(0.090)* |
| LTG ($\alpha_{LTG}$) | | | | -0.544$^{***}$ |
| | | | | *(0.148)* |
| # observations | 657 | 657 | 657 | 657 |
| $R^2$ | 0.76 | 0.77 | 0.77 | 0.79 |
| $\Omega^*$ | 0.546 | 0.546 | 0.546 | 0.546 |

Notes: The table shows the effect of FG on Consensus forecaster disagreement regarding one-year-ahead forecasts for 3-month interest rates, as measured by the interdecile range. Country and time fixed effects not reported. SG denotes state-contingent FG, OG open-ended FG, TG time-contingent FG, LTG time-contingent FG with a remaining guidance horizon of more than 1.5 years, STG time-contingent FG with a remaining guidance horizon of less than or equal to 1.5 years. $g$ measures the horizon of time-contingent FG relative to the forecast horizon, $\Omega^*$ the sample average of the interdecile range in the absence of FG. Driscoll and Kraay (1998) standard errors are given in parentheses. Asterisks indicate the level of significance, (*) at the 10%, (**) at the 5%, and (***) at the 1% level.

applies). Finally, the increased market responsiveness under short-horizon time-contingent FG is puzzling: the central bank announces that it will keep short-term rates stable for a while, yet interest rates become more responsive to incoming news about the economy.

### 3.3.2 Forecaster Disagreement

Another way to test the effectiveness of FG in managing expectations is to study its impact on forecaster disagreement at monthly frequency. Andrade et al. (2019) have shown that under time-contingent FG, forecaster disagreement about future interest rates is reduced, although disagreement about the future macroeconomic outlook has increased. We extend their analysis to see whether the effects on interest rate expectations differ depending on the type of FG.

Column (1) of Table 5 replicates the results of Andrade et al. (2019): in the presence of FG, there is less disagreement across professional forecasters about the future path of interest rates. While only marginally significant, the impact is estimated to be economically important.

To see this, look at the bottom row of the table, which reports the average disagreement that prevails in our sample in the absence of FG, denoted by $\Omega^*$. The estimate of $\Omega^*$ of 0.546 is our reference point for the following results. Under FG, this disagreement is reduced by 20%. While this is sizeable, it also implies that professional forecasters still have different views about the future path of interest rates. Splitting up the various types of FG in column (2), it is apparent that the statistical significance stems from state-contingent FG, which also shows the largest reduction. While column (3) suggests that the effects of time-contingent FG do not move step in step with the guidance horizon, it is still useful to split time-contingent FG into short-horizon and long-horizon guidance as done in the previous analysis for the responsiveness of bond yields to macroeconomic news. Results of this exercise are reported in Column (4). Consistent with the previous results, we find that long-horizon time-contingent FG effectively "cements" expectations – they become entirely unresponsive to news, and they lead to full agreement among professional forecasters: the average disagreement in the absence of FG of 0.546 is reduced by virtually the same amount, fully eliminating any disagreement. Also in line with the earlier results on the responsiveness of bond yields, we find that state-contingent FG roughly halves the disagreement across forecasters. Finally, we find open-ended FG and time-contingent FG over a short horizon to be ineffective, in the sense that they do not affect disagreement in any meaningful manner: the estimated coefficients are close to zero, and not statistically significant.

In the Appendix, we show that the central bank can enhance the credibility of its FG by putting "skin in the game" by embarking on an asset purchase programme (APP). This implies that the central bank is willing to take losses on its balance sheet if it does not follow up on its own FG. In this case, the differences across FG types become less pronounced, and all FG types do lead to the desired effect. Time-contingent FG in particular benefits from the existence of an APP: it mutes the perverse effect of short-horizon time-contingent FG on the responsiveness of bond yields to macro news and strengthens the effect of long-horizon time-contingent FG. Also, once the central bank embarks on an APP, FG always leads to less disagreement, regardless of its type.

To summarise, our empirical findings show that FG generally reduces the responsiveness of bond yields to macro news and the disagreement across forecasters. However, there are certain types of FG where this is not the case, namely open-ended FG and time-contingent FG over a short horizon. In the latter case, responsiveness to news counterintuitively increases, whereas disagreement is barely affected.

# 4 A Model of Learning from Market Signals

In our empirical analysis, we have provided evidence that FG can be effective in anchoring expectations about the future path of policy. However, we have also identified cases where FG does not lead to the intended effect, failing to reduce disagreement across professional forecasters, and even raising the responsiveness of yields to macroeconomic news.

In what follows we develop a model that rationalizes the perverse effect of public announcements as the consequence of a market externality in information aggregation.

We present a static asset pricing model where agents trade a bond that has a payoff which is related to the state of the economy. While the payoff is unknown to agents at the time of trading, they receive a noisy private signal as well as a noisy public signal about the state of the economy, and also imperfectly observe market information. We assume that the central bank determines the extent to which the bond payoff varies with the state of the economy. In particular, we interpret FG as generating a lower pass-through of fundamentals to bond payoffs.

We abstract from issues concerning credibility and time-inconsistency and focus on how the central bank affects the way agents form expectations by altering the reaction of returns to fundamentals. Intuitively, FG announcements reduce ex-ante uncertainty because the asset payoff fluctuates less with the state of the economy. More precisely, it is less likely that rates will go down in response to news of a bad state as a FG announcement signals that the ELB is close; it is less likely that rates will go up in response to news of an improved outlook as the FG announcement indicates that rates will be low for some time.

Thus, in principle, FG announcements reduce the usefulness of private and public information for predicting returns and implies a lower news-sensitivity of asset prices and lower disagreement. We show that this is indeed always the case when studying the case where agents do not learn from market signals.

In contrast, the presence of endogenous information introduces an externality that can explain our empirical findings by generating a second, countervailing effect. As FG reduces ex-ante uncertainty, expectations tend to react less to private information. As a consequence, the market price aggregates less information and becomes less informative. This loss of information can make agents' expectations more sensitive to macro news.

The combination of these two economic forces implies a non-linearity, so that a moderate strengthening of FG can lead to an increase, rather than a decrease, of uncertainty about future rates, which exacerbates the news-sensitivity of expected future rates and generates disagreement across forecasters. However, for a sufficiently strong implementation of FG, the market signals become less relevant, so that one always obtains a decrease in uncertainty. We derive an analytical expression for the threshold at which the marginal effect of strengthening FG changes sign.

## 4.1 Setup

**Financial market.** There is a continuum of agents with mass one, indexed by $i \in (0,1)$. They can invest in bonds with a stochastic final payoff $\tilde{\theta} \sim N\left(\bar{\theta}, \tau_\theta^{-1}\right)$, where $\tau_\theta$ denotes the precision of $\tilde{\theta}$. Agent $i$ solves the optimization problem

$$\max_{Q_i} \left[ \left( E\left[\tilde{\theta}|\Omega_i\right] - P \right) Q_i - \frac{1}{2} Q_i^2 \right]$$

where $E\left[\cdot|\Omega_i\right]$ the expectations operator conditional on the information set of agent $i$, $Q_i$ is her investment into the treasury bond, $Q_i^2/2$ represents a quadratic cost of portfolio management, and P denotes the bond price. Let $\tilde{\kappa} \sim N\left(\bar{\kappa}, \tau_\kappa^{-1}\right)$ denote the noisy aggregate net supply of treasury bonds. Market clearing implies $\int_i Q_i di = \tilde{\kappa}$. The optimal individual demand is

$$Q_i = E\left[\tilde{\theta}|\Omega_i\right] - \tilde{p},$$

which, combined with market clearing, gives the equilibrium price

$$p = \int E\left[\theta|\Omega_i\right] di + \kappa,$$

where $\kappa = \bar{\kappa} - \tilde{\kappa}$, $p = P - \bar{\theta} + \bar{\kappa}$ and $\theta = \tilde{\theta} - \bar{\theta}$. As usual, the price of the bonds increases in the expected payoff and decreases in the net supply.

**Central Bank.** The macroeconomic state $\tilde{\pi}$, which is related to the central bank's mandate, follows the distribution $\tilde{\pi} \sim N\left(\bar{\pi}, \tau_\pi^{-1}\right)$. The random component is therefore $\pi = \tilde{\pi} - \bar{\pi}$. We will refer to the inverse of the precision $\tau_\pi$ as the ex-ante uncertainty on the state of the economy. The central bank observes $\pi$ and sets the bond payoff (i.e. the policy rate) $\theta$ according to the rule

$$\theta = \alpha\pi, \tag{8}$$

where $\alpha \geq 0$. The parameter $\alpha$ represents the systematic component of monetary policy and is publicly announced.[12]

**Information.** By varying $\alpha$, the central bank also affects the ex-ante uncertainty that agents have about $\theta$. In fact, the prior distribution of $\theta$ is given by

$$\theta \sim N\left(0, \alpha^2 \tau_\pi^{-1}\right), \tag{9}$$

where the ex-ante uncertainty about the bond payoff $\theta$ is given by $\alpha^2 \tau_\pi^{-1}$. We interpret announcements of a lower $\alpha$ as stronger forms of FG, as they make rates less dependent on fundamentals. Note that stronger (weaker) forms of FG imply lower (higher) volatility on policy rates because $\tau_\theta = \alpha^{-2}\tau_\pi$.

In contrast to the central bank, the public only observes a noisy signal of $\pi$, which is given by

$$y = \pi + \varepsilon, \tag{10}$$

with $\varepsilon \sim N\left(0, \tau_\varepsilon^{-1}\right)$. This signal can be viewed as the surprise component of a macroeconomic news release. Besides this public signal, each agent observes a private signal

$$s_i = \pi + \eta_i, \tag{11}$$

where $\eta_i \sim N\left(0, \tau_\eta^{-1}\right)$ is i.i.d. across agents. For future reference, it is useful to refer to the heterogeneity captured by the latter as structural. Taken together, (9), (10) and (11) constitute the exogenous component of agents' information set. We use the term exogenous for pieces of information with a precision that is independent of equilibrium relations.

Finally, agents also observe endogenous signals, i.e signals with a precision that depends on equilibrium relations. More specifically, we assume that agents' market signals consist of private signals of the equilibrium price in the bond market. Formally, these are given by

$$x_i = p + \xi_i = \int E\left[\theta|\Omega_i\right] di + \kappa + \xi_i, \tag{12}$$

[12]For instance, the central bank might typically set policy rates according to a Taylor rule, with a Taylor rule coefficient $\alpha_{TR}$. Under FG, it would set $0 \leq \alpha < \alpha_{TR}$.

where $\xi_i \sim N\left(0, \tau_\xi^{-1}\right)$ is i.i.d. across agents. Note that the price signal provides information about the aggregate expectation subject to two noise components: a common supply disturbance $\kappa$, and differences in the way individuals perceive the market price $\xi_i$. The private noise may be interpreted in several ways. It can capture different interpretations of the same market evidence, due to, for example, the use of different market indexes, or the presence of cognitive limitations in processing information. This second interpretation is in the spirit of a growing body of literature on rational inattention, intended as a behavioral model of expectation formation (for a recent survey see Mackowiak et al. (2018). See also Vives and Yang (2017) for an approach that is similar in spirit.) We will refer to the heterogeneity arising from the noise in market signals as behavioral, in contrast to the structural heterogeneity induced by private signals.

All stochastic variables $\pi$, $\varepsilon$, $\eta_i$, $\xi_i$ and $\kappa$ are assumed to be mutually independent. Their probability distributions and corresponding moments are public knowledge.

## 4.2 Equilibrium

To solve for the equilibrium, we analyze how agents form expectations about the policy rate $\theta$. There are three sources of information, so that agent $i$ forms her expectation according to the linear rule

$$E\left[\theta|\Omega_i\right] = a\alpha s_i + b\alpha y + cx_i, \tag{13}$$

where $a$, $b$, and $c$ are equilibrium weights. Aggregating across agents and substituting the various signals yields

$$\int E\left[\theta|\Omega_i\right] di = \frac{a}{1-c}\alpha\pi + \frac{b}{1-c}\alpha\left(\pi+\varepsilon\right) + \frac{c}{1-c}\kappa, \tag{14}$$

which can be substituted into the market price to obtain

$$p = \frac{a\alpha}{1-c}\pi + \frac{b\alpha}{1-c}y + \frac{1}{1-c}\kappa. \tag{15}$$

The coefficient $\phi \equiv \frac{b\alpha}{1-c}$ represents the sensitivity of market prices to public information. Note from equation (14) that $\phi$ also represents the news-sensitivity of the average expectation. Structural disagreement is given by $\Delta_s = a^2\alpha^2\tau_\eta^{-1}$, whereas behavioral disagreement is $\Delta_b = c^2\tau_\xi^{-1}$. Overall, disagreement obtains as $\Delta = \Delta_b + \Delta_s$.

The optimal weights $a$, $b$ and $c$ which characterize the rational expectation equilibrium must be such that forecast errors are orthogonal to available signals. Formally, the weights in equation (13) have to satisfy the following three orthogonality conditions:

$$
\begin{aligned}
E\left[\alpha s_i\left(\theta - E\left[\theta|\Omega_i\right]\right)\right] &= 0, && (16)\\
E\left[\alpha y\left(\theta - E\left[\theta|\Omega_i\right]\right)\right] &= 0, && (17)\\
E\left[x_i\left(\theta - E\left[\theta|\Omega_i\right]\right)\right] &= 0. && (18)
\end{aligned}
$$

The following proposition states the equilibrium weights.

**Proposition 1** *In equilibrium, we have*

$$
\begin{aligned}
a(c) &= \frac{\tau_\eta}{\frac{1}{1-c}\tau_\eta + \tau_\varepsilon + \tau_\pi}, && (19)\\
b(c) &= \frac{(1-c)\tau_\varepsilon}{\frac{1}{1-c}\tau_\eta + \tau_\varepsilon + \tau_\pi}, && (20)
\end{aligned}
$$

*and so*

$$\phi = \frac{\alpha\tau_\varepsilon}{\frac{1}{1-c}\tau_\eta + \tau_\varepsilon + \tau_\pi} \quad \text{and} \quad \Delta_s = \left(\frac{\alpha}{\frac{1}{1-c}\tau_\eta + \tau_\varepsilon + \tau_\pi}\right)^2 \tau_\eta \quad \text{and} \quad \Delta_b = \frac{c^2}{\tau_\xi} \tag{21}$$

*with $c$ being a real root of the fixed-point equation*

$$\frac{\frac{1}{1-c}\tau_\eta}{\left(\frac{1}{1-c}\tau_\eta + \tau_\varepsilon + \tau_\pi\right)^2}\alpha^2 - c\left[\frac{1}{(1-c)^2}\frac{1}{\tau_\kappa} + \frac{1}{\tau_\xi}\right] = 0. \tag{22}$$

The proof is given in the Appendix. The fixed-point equation (22) pins down the equilibrium value of $c$. While there is no closed-form solution, we can state following:

**Proposition 2** *In equilibrium, $c \in (0,1)$. Moreover, $c \to 0^+$ as $\alpha \to 0^+$ or $\tau_\eta \to 0^+$.*

The intuition for why $c$ approaches zero as $\alpha$ declines towards zero is straightforward. A lower $\alpha$ reduces the uncertainty on the asset payoff, and the usefulness of any information. For $\alpha$ equal to zero, agents do not need any information; as $\alpha$ marginally increases above zero, the sensitivity to any signal has to increase. Explaining why $c \to 0^+$ as $\tau_\eta \to 0^+$ is slightly more subtle. The price signal is informative about the aggregate expectation and not about $\theta$ directly. Therefore, when the only informative signal is public (because the private signal is uninformative), the price signal can only be a noisy version of that public information (see (15) with $a = 0$), and thus becomes redundant.

To overcome the difficulty of not having a closed form solution for $c$, we will proceed by studying the signal extraction problem in particular cases, to highlight some general properties.

## 4.3 The Case of No Market Signals

In this subsection, we will focus on a case where market signals are absent. We will show that our two empirical facts established in the previous section cannot be rationalized in a typical Bayesian model of expectation formation in the absence of market signals.

Expressing the optimal weights that agents attach to the private and the public exogenous signals as a function of the weight on the market signal helps understanding how the presence of market signals affects the signal extraction problem. The case without market signals is characterized by $c = 0$. In equilibrium, this is consistent with $\tau_\kappa \to 0$ or $\tau_\xi \to 0$, i.e. the cases in which market signals do not transmit any information. The first observation is that $a(0)$ and $b(0)$ given by (19) and (20) are equal to the precision of public and private signals, respectively, divided by the total precision (i.e. the sum of the precision of the prior $\tau_\pi$, of the public signals $\tau_\varepsilon$ and of the private signals $\tau_\eta$). In particular, note that $E[\pi|s_i, y] = a(0)s_i + b(0)y$ is the conditional expectation about $\pi$ given $s_i$ and $y$ and that $E[\theta|s_i, y] = \alpha E[\pi|s_i, y]$. Moreover, one can immediately check by substitution in (21) the following:

**Corollary 3** *In the absence of an endogenous signal – which is equivalent to imposing $c = 0$ – both the sensitivity to public signals, $\phi$, and the disagreement, $\Delta$, strictly increase in $\alpha$.*

This Corollary establishes an important benchmark in which agents do not observe any price signal and can only access information with exogenous precision. In such a case, the only possibility is that stronger FG (lower $\alpha$) dampens the sensitivity of prices and expectations to news and reduces disagreement. We have thus established that in the absence of market signals a standard model of Bayesian updating cannot replicate the empirical facts presented in the previous section.

## 4.4 The Case with Market Signals

We now return to the complete model with market signals. While it is not possible to obtain a closed-form solution for the equilibrium coefficients, it is still possible to deduce additional statements about the evolution of the news-sensitivity $\phi$ and of disagreement $\Delta$ in the limiting case where $\tau_\kappa \to \infty$. In this case, the fixed-point equation collapses to

$$\frac{\frac{1}{1-c}\tau_\eta}{\left(\frac{1}{1-c}\tau_\eta + \tau_\varepsilon + \tau_\pi\right)^2}\alpha^2 = c\tau_\xi^{-1}, \tag{23}$$

and we can the prove the following result:

**Proposition 4** *For $\tau_\kappa \to \infty$, $\phi$ and $\Delta_s$ are a non-monotonic function of $\alpha$. In particular, it achieves a maximum at $c = 1/2$ which corresponds to*

$$\alpha^* = \sqrt{\frac{\tau_\eta}{\tau_\xi} + \frac{1}{2}\frac{\tau_\varepsilon + \tau_\pi}{\sqrt{\tau_\eta \tau_\xi}}}.$$

From the proposition, we know that as the variance of the private noise in the price signal $\tau_\xi^{-1}$ becomes small, $\alpha^*$ becomes small. This means that the range of $\alpha$, in which sensitivity to macroeconomic news is decreasing in $\alpha$, gets wider. Intuitively, a decrease in $\alpha$ decreases ex-ante uncertainty about $\theta$, which normally should decrease reliance on other sources of information. However, when $\alpha$ decreases, agents react less to private signals so that market signals become less informative. It is exactly when market signals are powerful aggregators of information - i.e. with $\tau_\xi$ large - that the second effect may dominate and agents may become overall more (rather than less) uncertain about rates.

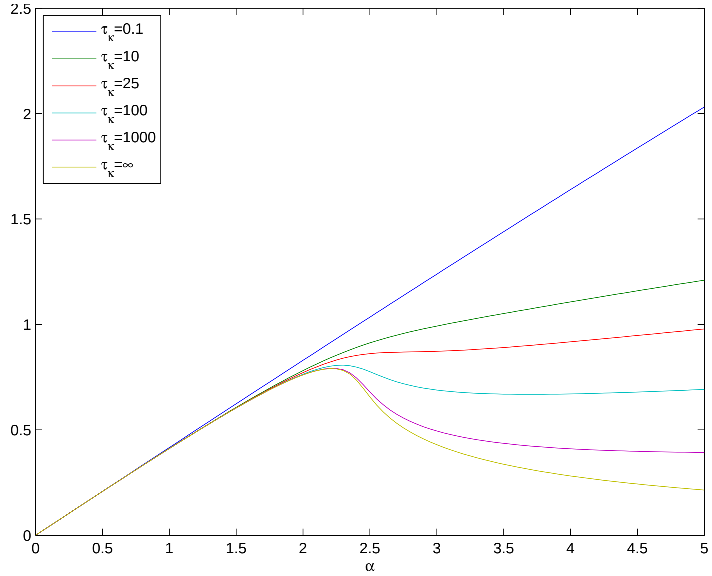
While Proposition 4 proves that the model can generate a higher news-sensitivity $\phi$ for a strengthening of FG (a decrease in $\alpha$) in the limiting case where $\tau_\kappa \to \infty$, numerical simulations show that this result is actually more general. Figure 2 plots $\phi$ as a function of $\alpha$ for various values of $\tau_\kappa$, holding the remaining parameters of the model fixed. As $\tau_\kappa$ increases, $\phi$ changes from a monotone function into a hump-shaped function, so that for $\alpha$ sufficiently high, a small decrease in $\alpha$ — akin to a change from no FG to short-term FG — leads to a higher sensitivity of asset prices to public information.

Figure 3 illustrates the underlying economic mechanisms. In the first panel, we plot ex-post uncertainty, given by

$$V(\theta|\Omega_i) = \left(\frac{1}{1-c}\tau_\eta + \tau_\varepsilon + \tau_\pi\right)^{-1},$$

as a function of $\alpha$. As we see, marginally lower values of $\alpha$ increase uncertainty for sufficiently high values of $\alpha$. In particular, a marginal decrease of $\alpha$ for values above $\alpha^* = 2.21$, i.e. a drop in ex-ante uncertainty, may then cause an increase in ex-post uncertainty in the presence

Figure 2: News-sensitivity $\phi$ as a function of $\alpha$ for different values of $\tau_\kappa$




Notes: This figure depicts the news sensitivity $\phi$ as a function of $\alpha$ for different values of $\tau_\kappa$. We set $\tau_\eta = 0.4$ and $\tau_\varepsilon = \tau_\pi = \tau_\xi = 1$.

of market signals.

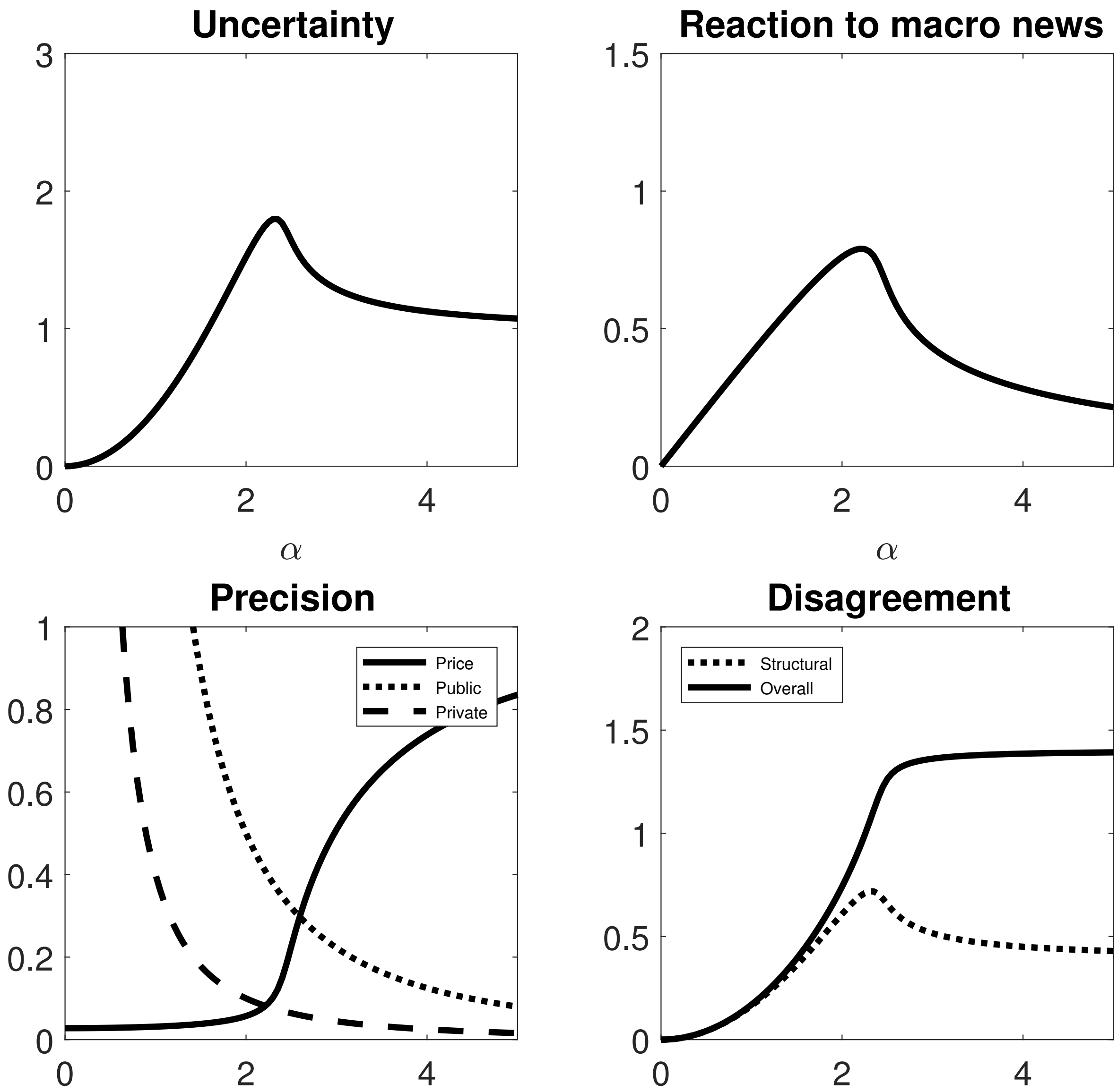
To get an intuition of the mechanism at work, it is instructive to look at the panel below. We plot with a dashed line the precision of the private signal $\alpha^{-2}\tau_\eta$ about $\theta$ as a function of $\alpha$. We also plot with a dotted line the overall precision of public information, i.e. prior precision plus precision of the public signals $\alpha^{-2}(\tau_\pi + \tau_\varepsilon)$. Both curves are decreasing in $\alpha$, meaning that for stronger FG, the precision of these signals increases. On the contrary, the precision of the market signal about $\theta$, which is given by

$$\left(\frac{a}{1-c}\right)^2 \tau_\xi$$

and is depicted with a solid line, is increasing in $\alpha$. What happens is that, as $\alpha$ increases, $a$ decreases more slowly than $c$ increases. In particular, despite agents being in principle less reactive to their private signals, the amplification loop (generated by the $1-c$ term in the denominator) makes expectations increasingly reactive to private signals, thereby increasing the precision of the market signal. Coming from the opposition direction, the precision of the market signal therefore drops drastically as we move only marginally towards lower values of $\alpha$, and, as a consequence, ex-post uncertainty may increase despite an increase in precision of exogenous information.

Finally, the reaction to macro news $\phi$ and structural disagreement $\Delta_s$ follows the same evo-

Figure 3: Model results for different values of $\alpha$

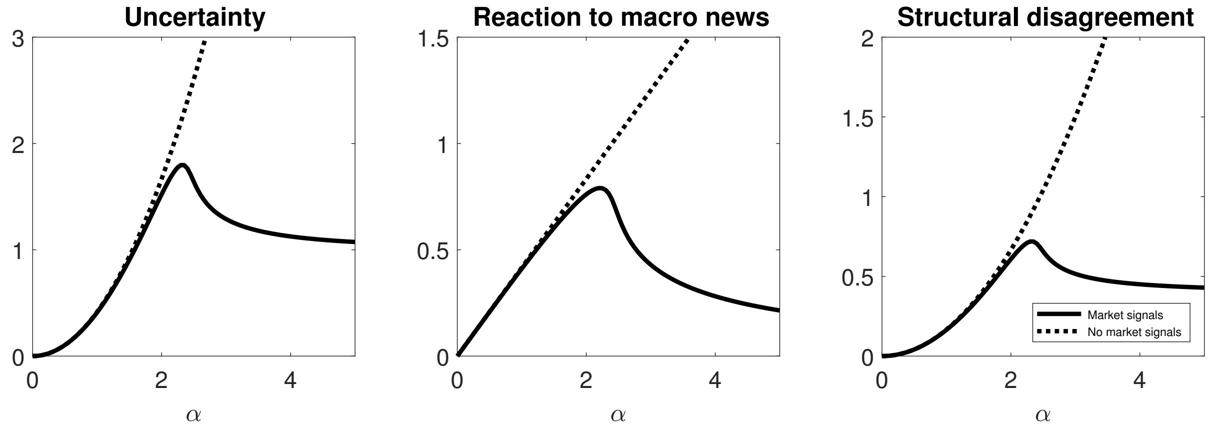



Notes: This figure shows how the model results vary with $\alpha$. The figure is obtained with $\tau_\eta = 0.4$, $\tau_\varepsilon = \tau_\pi = \tau_\xi = 1$, which yields $\alpha^* = 2.21$.

lution of ex-post uncertainty. This is intuitive by looking at their expressions in Proposition 1. Because of higher ex-post uncertainty, agents tend to put higher weights on market signals that amplify the reaction to macro news and idiosyncratic differences in signals. This is how the sensitivity to macro-news and structural disagreement may increase with $\alpha$.

The last panel of Figure 3 shows that, even if structural disagreement exhibits a non-monotonic behavior, this need not be the case for overall disagreement, given the monotonic evolution of the behavioral component $\Delta_b$ of disagreement. In particular, the rather flat derivative of overall disagreement for values of $\alpha$ above $\alpha^*$ is consistent with our finding that an increasing responsiveness to macro news is associated with mild changes in disagreement, whereas decreased responsiveness to macro news is coupled with more substantial reductions in disagreement.

Figure 4: The importance of market signals




Notes: This figure shows how the presence of market prices affects the key quantities of the model. The figure is obtained with $\tau_\eta = 0.4$, $\tau_\varepsilon = \tau_\pi = \tau_\xi = 1$, which yields $\alpha^* = 2.21$.

Figure 4 emphasizes the central role of market signals in our proposed mechanism. In the three panels we show ex-post uncertainty, the reaction to macroeconomic news and structural disagreement as a function of $\alpha$. The solid line repeats the results for an economy with market signals, which we presented in Figure 3. We contrast this in Figure 4 with the results for an economy without market signals – plotted as dotted line – which obtains by fixing $c = 0$, everything else being equal. The non-monotonic behavior of all three variables and their maximum at the same value $\alpha^*$ are striking. But it only obtains in the presence of market signals. Absent market signals, all three variables monotonically converge to zero as $\alpha$ decreases.

The figure also highlights that weak FG might not just elevate the responsiveness of asset prices to macroeconomic news. Rather, the elevated responsiveness of asset prices is symptomatic of the increased uncertainty and the larger structural disagreement in the economy.

# 5 Conclusion

Intuitively, one would expect FG to reduce the impact of macroeconomic news on asset prices. However, we show that this crucially depends on the type of guidance adopted. We study the impact of different types of FG on the responsiveness of bond yields to macroeconomic news, and on forecaster disagreement about the future path of interest rates. Time-contingent FG over long horizons eliminates both the asset price response to incoming news and disagreement across forecasters. State-contingent FG works in the same direction, but preserves some responsiveness and disagreement, because future policy continues to depend on a subset of macroeconomic information. In contrast, time-contingent FG over short horizons and open-ended FG are ineffective at best, and add uncertainty about the policy path at worst.

Rational expectations model with noisy market information explain these findings. In such a model, more public information can hamper the aggregation of private information in prices. Thus, when the market is an important source of information, the release of more precise public information can increase, instead of decrease, uncertainty and can amplify, instead of reduce, the reaction of expectations to macroeconomic news.

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# A Appendices

## A.1 Data Appendix

All macroeconomic announcement series come from Bloomberg, and were collected during November 2016 and January 2017. For the euro area we use the macroeconomic news announcements of Germany and Italy. We use the nine macroeconomic indicators **business confidence indices** (BCI) in Canada (CA), Germany (DE), Italy (IT), Japan (JP), Sweden (SE), United States (US), **consumer confidence indices** (CCI) in DE, IT, JP, SE, UK, US, **consumer price indices** (CPI) in DE, IT, JP, SE, UK, US, **GDP growth** (GDP) in CA, DE, IT, JP, SE, UK, US, **industrial production** (IPI) in DE, IT, JP, SE, UK, US, **non-farm payroll employment** (NFP) in US, **purchasing manager indices** (PMI) in CA, DE, IT, SE, UK, US, **retail sales** (RS) in CA, DE, IT, JP, SE, UK, US, and **unemployment rates** (UR) in CA, DE, IT, JP, SE, UK, US.

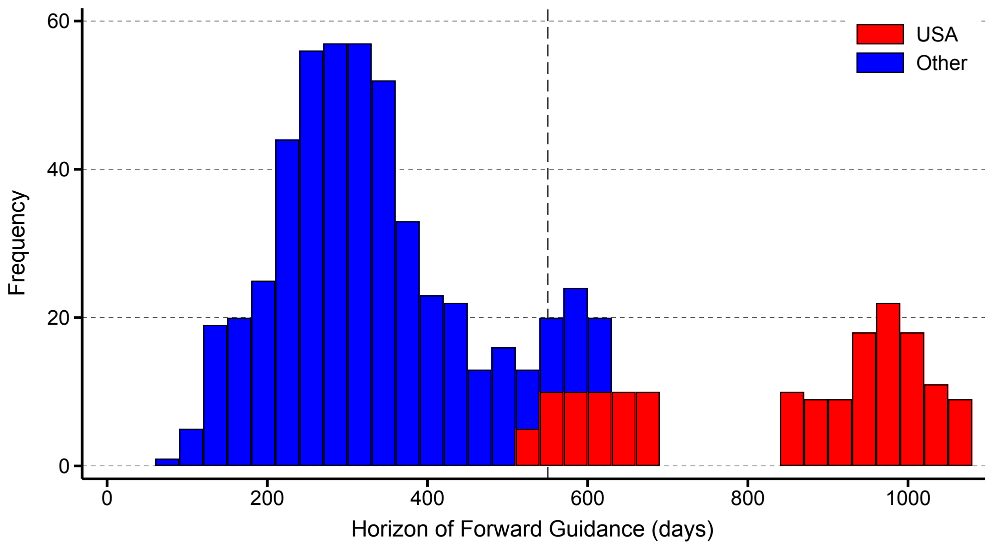
Table A.1: Observations (daily sample) by country, FG and APP regime

| | No FG | SG | OG | STG | LTG | total |
|---|---|---|---|---|---|---|
| Canada | 359 / 0 | 0 / 0 | 0 / 0 | 40 / 0 | 0 / 0 | 399 / 0 |
| Euro area | 585 / 0 | 0 / 0 | 259 / 210 | 0 / 120 | 0 / 0 | 844 / 330 |
| Japan | 548 / 264 | 0 / 179 | 14 / 0 | 0 / 0 | 0 / 0 | 562 / 443 |
| Sweden | 76 / 78 | 0 / 0 | 0 / 0 | 218 / 75 | 31 / 0 | 325 / 153 |
| UK | 0 / 325 | 0 / 39 | 0 / 230 | 0 / 0 | 0 / 0 | 0 / 594 |
| United States | 132 / 5 | 0 / 154 | 0 / 640 | 0 / 8 | 0 / 153 | 132 / 960 |
| total | 1700 / 672 | 0 / 372 | 273 / 1080 | 258 / 203 | 31 / 153 | 2262 / 2480 |

Notes: The first number is the number of observations outside of APP periods in our sample, followed by the number of observations during APP periods. An observation is a release of a macroeconomic indicator during the respective FG regime in the respective country. ELB periods only.

Table A.1 reports the number of observations during ELB periods in the daily sample. The number of observations during APP periods is about the same as the number of observations outside of APP periods. There are no APP observations for CA, and only APP observations for the UK. Figure A.1 shows the frequency of FG horizons at the macroeconomic news announcement times in our sample.

Figure A.1: Frequency of guidance horizons of time-contingent FG




Notes: The figure shows the frequency of time-contingent FG horizons at the macroeconomic news announcement times in our sample. The countries covered are Canada, the euro area (Germany and Italy), Sweden, and the United States. The subset of observations for the United States is shaded in red.

## A.2 The Effects of Forward Guidance in Presence of an Asset Purchase Programme

We have conjectured that different types of FG might signal different levels of commitment on behalf of the central bank to keep policy rates at the ELB, and are accordingly more or less effective in anchoring expectations about the future path of policy. Especially short-term time-contingent and open-ended FG reveal a strong "delphic" element. At short FG horizons, the time-contingent FG announcement might have been primarily perceived as the central bank projecting a further deterioration of the economy. This bad signal dominates the short-term fixing of the policy rate. But even longer-term guidance is not enough to fully mute the surprise impact. There might have been credibility issues, and an inability of central banks to commit.

FG is in principle subject to a time-consistency problem: once central banks have managed to affect expectations about the future path of policy, they have an incentive to deviate from the communicated path earlier than announced. Of course, agents are aware of this inherent time inconsistency, and do therefore not necessarily see FG as credible. In addition, with FG having been a novel tool for most central banks in our sample, there is no reputational capital or track record that agents can rely on to understand whether the central bank would follow up on its communication. In the light of this, the credibility of FG can possibly be enhanced substantially if the central bank also has "skin in the game". One such way is to embark on an APP, which implies that the central bank is willing to take losses on its balance sheet if it does not follow up on its own FG (Clouse, Henderson, Orphanides, Small and Tinsley, 2003; Eggertsson and Woodford, 2003).

To test this hypothesis, we distinguish the effects during and outside of APP periods. We create an APP indicator variable $APP_t^c$, which is equal to one if country $c$ has an active APP at time $t$ and is zero otherwise. Among the economies in our sample, and for the sample period that we cover, the U.S. Federal Reserve started its APP on 25 November 2008, the Bank of England on 19 January 2009, the Bank of Japan on 05 October 2010, the ECB on 22 January 2015 and the Swedish Riksbank on 11 February 2015. We consider an APP as active as long as any principal is reinvested at maturity; therefore, all APP programmes were still active at the end of our sample period. By adding the APP indicator variable and the respective interaction terms, our regression expands to

$$\begin{aligned}
y_t^{c,i} &= \alpha^{c,i} + \alpha_{SG}SG_t^c + \alpha_{OG}OG_t^c + \alpha_{STG}STG_t^c + \alpha_{LTG}LTG_t^c + \gamma APP_t^c \\
&+ \gamma_{SG}SG_t^c APP_t^c + \gamma_{OG}OG_t^c APP_t^c + \gamma_{STG}STG_t^c APP_t^c + \gamma_{LTG}LTG_t^c APP_t^c \\
&+ \beta s_t^{c,i} + \beta_{SG}SG_t^c \times s_t^{c,i} + \beta_{OG}OG_t^c \times s_t^{c,i} + \beta_{STG}STG_t^c \times s_t^{c,i} + \beta_{LTG}LTG_t^c \times s_t^{c,i} \\
&+ \delta APP_t^c \times s_t^{c,i} + \delta_{SG}SG_t^c APP_t^c \times s_t^{c,i} + \delta_{OG}OG_t^c APP_t^c \times s_t^{c,i} \\
&+ \delta_{STG}STG_t^c APP_t^c \times s_t^{c,i} + \delta_{LTG}LTG_t^c APP_t^c \times s_t^{c,i} + \varepsilon_t^{c,i}. \qquad \text{(A.1)}
\end{aligned}$$

In this regression, we are able to tell whether the responsiveness of bond yields to macro news in the presence of a certain FG type is different in the absence or in the presence of an APP. For instance, for state-contingent FG, we need to compare $\beta + \beta_{SG}$ to $\beta + \beta_{SG} + \delta + \delta_{SG}$.

For the disagreement analysis, the extended model becomes

$$\begin{aligned}
\Omega_t^c &= \alpha^c + \alpha_t + \alpha_{SG}SG_t^c + \alpha_{OG}OG_t^c + \alpha_{STG}STG_t^c + \alpha_{LTG}LTG_t^c + \gamma APP_t^c \\
&+ \gamma_{SG}SG_t^c APP_t^c + \gamma_{OG}OG_t^c APP_t^c + \gamma STG_t^c APP_t^c + \gamma_{LTG}LTG_t^c APP_t^c + \varepsilon_t^c. \qquad \text{(A.2)}
\end{aligned}$$

### A.2.1 Bond Yields

Indeed, as can be seen in Table A.2, an APP in combination with FG turns out to be highly effective in anchoring expectations. The coefficients of each regression are listed in two adjacent columns. For daily data, column (1) lists the coefficients in periods without APP, whereas

Table A.2: FG and net surprise impact with and without an APP

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| | daily data | | intraday data | |
| | without APP | with APP | without APP | with APP |
| No FG | 0.490** | 0.407** | 0.489*** | 0.509** |
| | *(0.242)* | *(0.168)* | *(0.172)* | *(0.200)* |
| SG | n.a. | 0.223* | n.a. | 0.111 |
| | | *(0.123)* | | *(0.071)* |
| OG | 0.008 | 0.516* | 0.056 | 0.561** |
| | *(0.213)* | *(0.277)* | *(0.169)* | *(0.215)* |
| STG | 1.907*** | 0.250 | 1.453*** | 0.510*** |
| | *(0.288)* | *(0.300)* | *(0.163)* | *(0.183)* |
| LTG | 0.194 | 0.046 | 0.248 | 0.124* |
| | *(0.212)* | *(0.091)* | *(0.238)* | *(0.071)* |
| # observations | 4742 | | 4739 | |
| $R^2$ | 0.02 | | 0.03 | |

Notes: Dependent variable in columns (1) and (2): Daily change in two-year sovereign bond yields in basis points. Dependent variable in columns (3) and (4): Intraday (120 minute window) change in two-year sovereign bond yields in basis points. Fixed effects model (A.1). The reported values are – analogous to Tables 3 and 4 – the sums of the no-FG response and the extra FG effect, i.e. e.g. $\beta + \beta_{OG}$, $\delta + \delta_{OG}$. Country-indicator fixed effects, FG and APP fixed effects not reported. Only periods at or below the ELB of 1%. Standard errors clustered at the country-indicator level in parentheses. Asterisks indicate the level of significance, (*) at the 10%, (**) at the 5%, and (***) at the 1% level.

column (2) list the coefficients for APP periods. For intraday data, the analogous regression coefficients are listed in columns (3) and (4).

The asset price response without FG is barely changed by APP, neither in daily nor in intraday data. Open-ended FG without APP, most of which has been implemented by the ECB, mutes the asset price response almost completely. During an APP, however, open-ended FG becomes ineffective and markets are even more responsive to news then. The open-endedness of FG might be seen as contradicting the commitment conveyed by APP. A contradictory message from the central bank might heighten uncertainty about monetary policy, possibly combined with a perceived heightened sensitivity of monetary policy to macroeconomic indicators.

Time-contingent FG, however, strongly benefits from the existence of an APP. APP mutes the perverse effect of short-horizon time-contingent FG and strengthens the effect of long-horizon time-contingent FG. In daily data, short-horizon time-contingent FG combined with APP can even shrink the asset price response to insignificance. But even long-horizon FG becomes more effective in combination with APP, getting very close to a full credibility – mirrored bond prices immune to macroeconomic news. This supports the view that if the central bank embarks on an APP, this can act as a commitment device to keep interest rates low for an extended period, thereby adding credibility to the FG about policy rates. When both an APP and long-horizon time-contingent FG point in the same direction, this sends a very clear signal.

Comparing the coefficients for the no-APP regime with the overall coefficients reported in Tables 3 and 4 shows that the perverse effect of short-horizon time-contingent FG is due to the non-APP periods. Without APP, markets are looking through short-horizon FG and are highly responsive to news.

### A.2.2 Forecaster Disagreement

We have also tested to what extent the effect of FG on forecaster disagreement is affected by the presence of an APP. Table A.3 reports the corresponding results. The first result to note is that forecaster disagreement differs substantially in our sample between periods with and without APP. $\Omega^*$, i.e. disagreement in the presence of the ELB but in the absence of FG is mich higher in the cases where the central bank also resorted to APP than otherwise. Interestingly, and in line with the bond yield results, we find that the effectiveness of FG is substantially improved in the presence of APP: without asset purchases, no type of FG manages to reduce disagreement, and open-ended as well as short-horizon time-contingent FG even raise disagreement, mirroring the perverse effect of the latter on bond yields. In contrast, once the central bank also embarks on an APP, FG is always leading to less disagreement, regardless of its type. As in our earlier tests, we find that long-horizon time-contingent FG completely eliminates disagreement and state-contingent cuts it roughly by half.

In other words, APP renders all types of FG more "Odyssean". An APP appears to convince markets that once the economy eventually recovers, the central bank will not increase policy rates as quickly or early as it did historically. In contrast, without APP a guidance of one year appears to be too short to convince the markets, and the "Delphic"bad news effect might dominate.

### A.2.3 Rationalising these Findings: The Case of No Private Noise in the Price Signal

The situation of FG in the presence of an APP can be cast in our theoretical model as follows: We interpret APP interventions by the central bank as the introduction of common disturbances to the net supply of the asset available in the market, i.e. demand shocks not related to the fundamental value of the asset. This could be achieved via a stochastic and publicly unknown

Table A.3: FG and forecaster disagreement with and without an APP

| | (1) without APP | (2) with APP |
|---|---|---|
| SG | n.a. | -0.356*** |
| | | *(0.094)* |
| OG | 0.269** | -0.537*** |
| | *(0.136)* | *(0.178)* |
| STG | 0.219* | -0.394*** |
| | *(0.126)* | *(0.123)* |
| LTG | 0.106 | -0.985*** |
| | *(0.155)* | *(0.207)* |
| APP | 0.228*** | |
| | *(0.049)* | |
| # observations | 657 | |
| $R^2$ | 0.82 | |
| $\Omega^*$ | 0.455 | 0.779 |

Notes: The table provides estimates of the effect of FG on Consensus forecaster disagreement with and without an APP. Column (1) contains the coefficients that result in the absence of an APP, column (2) those that result in the presence of an APP. $\Omega^*$ denotes the sample average of the interdecile range in the absence of FG. Driscoll and Kraay (1998) standard errors are given in parentheses. Asterisks indicate the level of significance, (*) at the 10%, (**) at the 5%, and (***) at the 1% level.

order flow, even if the total stock of purchases is known.

In the limit, this can be viewed as a situation where $\tau_\xi \to \infty$, i.e. a case where the price signal is blurred only by common supply disturbances. In Appendix section A.4.4, we prove the following result:

**Proposition 5** *For $\tau_\xi \to \infty$, the news-sensitivity $\phi$ is a monotonic function of $\alpha$ for any given $\{\tau_\eta, \tau_\xi, \tau_\varepsilon, \tau_\pi\}$.*

This proposition tells us that when disturbances in the price signal are common, the sensitivity to macro news always decreases when $\alpha$ falls. The prominence of supply shocks decouples prices from agents' evaluations, and prices become less informative. Learning from prices is thus less important, and stronger FG always yields a lower news-sensitivity of asset prices, consistent with our empirical findings.

## A.3 Robustness

We have tested our empirical analyses for robustness in various ways. The results of our robustness tests are reported in Table A.4 for the bond yield regressions and in Table A.5 for the

Table A.4: FG and net surprise impact (robustness)

| | (1) benchmark | (2) bonds with 1yr TTM | (3) benchmark | (4) SE, UK, US only | (5) incl. obs. off the ELB |
|---|---|---|---|---|---|
| | daily | daily | intraday | intraday | intraday |
| No FG | 0.465*** | 0.389*** | 0.497*** | 1.080*** | 0.495*** |
| | *(0.168)* | *(0.114)* | *(0.135)* | *(0.301)* | *(0.135)* |
| SG | 0.223* | 0.066 | 0.110 | 0.208 | 0.113 |
| | *(0.123)* | *(0.050)* | *(0.070)* | *(0.155)* | *(0.070)* |
| OG | 0.422* | 0.119 | 0.464** | 0.707*** | 0.477** |
| | *(0.231)* | *(0.119)* | *(0.183)* | *(0.257)* | *(0.182)* |
| STG | 1.252*** | 0.644*** | 1.080*** | 1.155*** | 1.072*** |
| | *(0.256)* | *(0.222)* | *(0.141)* | *(0.162)* | *(0.138)* |
| LTG | 0.084 | -0.080 | 0.127* | 0.129* | 0.151** |
| | *(0.089)* | *(0.066)* | *(0.070)* | *(0.071)* | *(0.063)* |
| No FG off ELB | | | | | 1.062*** |
| | | | | | *(0.180)* |
| # obs. | 4742 | 4740 | 4739 | 2166 | 8695 |
| $R^2$ | 0.01 | 0.01 | 0.03 | 0.09 | 0.05 |

Notes: Dependent variable in columns (1) and (2) is the daily change in two-year sovereign bond yields in basis points, whereas in columns (3), (4) and (5) it is the intraday (120 minute window) change. Fixed effects model (5). Columns (1) and (3) repeat the respective benchmark specifications. Column (2) uses bonds with one year time-to-maturity (TTM). Column (4) restricts the sample to Sweden, UK and the United States. Column (5) expands the sample to periods outside the ELB. Country-indicator fixed effects and FG fixed effects not reported. Only periods at or below the ELB of 1%. Standard errors clustered at the country-indicator level in parentheses. Asterisks indicate the level of significance, (*) at the 10%, (**) at the 5%, and (***) at the 1% level.

disagreement regressions. For easy reference we repeat the daily benchmark results in column (1) and the intraday benchmark results in column (3) of Table A.4.

**Bond Yields at Other Maturities** We have argued that credible FG mutes the price response of bonds maturing within the guidance horizon. The guidance might be perceived as somewhat weaker towards the end of the guidance period, with no guidance for the period thereafter. In the analysis so far, we have studied bonds with two years to maturity. With rare exceptions, the two-year period was covered only partially by time-contingent FG. This gives room for bond prices to react to macro news relevant for the period between the end of the guidance horizon and the maturity of the bond.

In this section we study bonds whose time-to-maturity is covered more completely by the time-contingent FG episodes observed. More specifically, we study one-year bonds. Our conjecture is that time-contingent FG, in particular short-term time-contingent FG, mutes their price response more than that of two-year time-to-maturity bonds. Because there is only incomplete intraday coverage of one-year yields, we limit this analysis to daily data.

Comparing the benchmark results for two-year bonds in column (1) with the results for one-year bonds in column (2) of Table A.4 shows that this conjecture is indeed true. Absent FG, one-year bond prices react to news about as much as two-year bond prices, but they differ in their behaviour under FG. No matter what type of FG is provided, one-year bonds respond less to news. The already fully muted response under long-term time-contingent FG remains muted. Most pronounced is the reduction in short-horizon time-contingent FG. Note, however, that although the guidance period now covers a bigger share of the time-to-maturity, the perverse effect persists: Bond prices react (insignificantly) more strongly to news under short-horizon time-contingent FG than without any FG.

**Bond Yields for Sweden, UK, US Only** Our full sample covers a very diverse set of countries, with large differences in central bank governance and track record. In this subsection we look at a set of countries with somewhat more similar track record and data coverage. We focus here on the Sweden, UK and United States only, because these countries have each been subject to at least two types of FG during the sample period. In this subsample, both Sweden and the United States have used time-contingent FG, the later with the longest horizons worldwide. Our results from the full sample also apply to this small subset of countries, shown in column (4) of Table A.4. Interestingly, macroeconomic news in these three countries have on average a higher surprise impact than in the other countries. Short-horizon time-contingent FG (insignificantly) elevates the responsiveness of asset prices relative to a situation without guidance. Open-ended FG slightly reduces the asset price impact, but this reduction is only marginally significant. State-contingent FG and long-term time-contingent FG, however, are again highly (and significantly) effective in muting the surprise impact.

**Bond Yields off the ELB** We now compare the responsiveness of asset prices during ELB periods with their responsiveness during more ordinary times. For this purpose we include a dummy variable for the ELB, defined as periods where the policy rate is at or below 1%, both in levels as well as interacted with the surprise. No country provided FG away from the ELB. As one would expect, including the off-ELB periods does not affect the coefficients during ELB. The results for the benchmark specification in column (3) and for the modified specification in column (5) in Table A.4 are almost identical.

Comparing the ELB and off-ELB coefficient in column (5) of Table A.4 shows that the surprise impact of announcements shrinks to about one half during ELB periods.[A.1]

**Bond Yields during State-Contingent FG** The end of state-contingent FG explicitly depends on macroeconomic indicators. Therefore, the surprise impact of macroeconomic indicators overall should not be completely muted in this regime. If state-contingent FG does not fully reduce the surprise impact of an indicator policy conditions on, this means that the state triggering a return to normal monetary policy is seen as a possible outcome.

The effect on individual macroeconomic indicators is hard to identify for several reasons. First, only three countries in our sample implemented state-contingent FG. Second, although state-contingent FG in our definition is explicit on the criterion it conditions on, this criterion is not necessarily a single indicator. Third, the macroeconomic indicators most closely related to the conditioning quantity are often released with a large delay, so that their asset price impact is small in all environments.

A lucky exception is United States non-farm payroll employment (NFP). The Federal Reserve's state-contingent FG explicitly mentioned unemployment, and in the United States a closely watched indicator of the labor market, NFP, is released very early in the announcement cycle and has one of the largest impacts on asset prices among all macroeconomic news announcements. For this reason, we take a closer look at FG in the US with our intraday sample, distinguishing the response to surprises in NFP from surprises in all other announcements.

We find that state-contingent FG in the United States fully muted the bond price response to the group of announcements in our sample excluding NFP. The response to NFP, however, is reduced by only about 75%. It remains significant. In fact, the response is more than two-and-a-half times the response to time-contingent FG, which in the United States is mostly long-horizon time-contingent FG.

Despite the small sample and specificity to the US, this supports the notion that state-contingent FG does not eliminate the asset price response completely, because the indicator conditioned on remains relevant for the duration of FG. We find no evidence, however, that asset prices might respond more to news that monetary policy conditions on.

[A.1] Swanson and Williams document this decline in responsiveness for Eurodollar futures and U.S. Treasury (Swanson and Williams, 2014a) and for UK Gilts and German Bunds (Swanson and Williams, 2014b). They study each country separately, without an explicit distinction between FG and ELB episodes.

Table A.5: FG and forecaster disagreement (robustness)

| | (1) benchmark | (2) interqrt. range | (3) 3mo-ahead forecasts | (4) forecasts of 10yr bond rates | (5) incl. obs. off the ELB |
|---|---|---|---|---|---|
| SG | -0.245** | -0.092** | -0.113*** | -0.109** | -0.248*** |
| | *(0.099)* | *(0.040)* | *(0.025)* | *(0.047)* | *(0.075)* |
| OG | -0.079 | 0.034 | -0.063 | 0.026 | 0.006 |
| | *(0.108)* | *(0.052)* | *(0.044)* | *(0.075)* | *(0.079)* |
| STG | 0.044 | 0.018 | -0.034 | -0.074 | 0.034 |
| | *(0.090)* | *(0.066)* | *(0.063)* | *(0.050)* | *(0.088)* |
| LTG | -0.544*** | -0.239*** | -0.127*** | 0.034 | -0.595*** |
| | *(0.148)* | *(0.058)* | *(0.047)* | *(0.098)* | *(0.144)* |
| ELB | | | | | -0.004 |
| | | | | | *(0.079)* |
| # obs. | 657 | 657 | 657 | 657 | 1237 |
| $R^2$ | 0.77 | 0.74 | 0.65 | 0.82 | 0.74 |
| $\Omega^*/NoFG$ | 0.546 | 0.254 | 0.224 | 0.673 | 0.731 |

Notes: The table provides robustness tests regarding the effect of FG on Consensus forecaster disagreement. Column (1) repeats the benchmark specification. Column (2) uses the interquartile range as a measure of disagreement. Column (3) looks at 3-month ahead forecasts of 3-month interest rates, column (4) at one-year ahead forecasts of 10-year rates. Column (5) expands the sample to periods outside the ELB. Country and time fixed effects not reported. $\Omega^*$ denotes the sample average of the disagreement measure in the absence of FG. Driscoll and Kraay (1998) standard errors are given in parentheses. Asterisks indicate the level of significance, (*) at the 10%, (**) at the 5%, and (***) at the 1% level.

**Forecaster Disagreement** For the disagreement regressions, results of the robustness tests are provided in Table A.5. Column (1) repeats the benchmark results for ease of comparison. Column (2) measures disagreement by the interquartile range (as in Mankiw et al. (2004) or Dovern et al. (2012)) instead of the interdecile range, i.e. uses a more restrictive definition of outliers. Results are extremely robust: under long-horizon time-contingent FG, disagreement disappears entirely, whereas it is roughly halved under state-contingent FG, and remains unchanged under open-ended FG and short-horizon time-contingent FG. Results are also robust to using 3-month ahead forecasts for 3-month rates, i.e. a shorter forecast horizon than the 12-month ahead forecasts used in the benchmark; see column (3). The next set of results, reported in column (4), suggests that FG affect disagreement about interest rates with longer maturities much less: when we study disagreement about 1-year ahead forecasts for 10-year government bond yields, only state-contingent FG seems to reduce disagreement, but by a small amount. Finally, in column (5), we show that results are not dependent on restricting the sample to periods at the ELB.

## A.4 Proofs

### A.4.1 Proof of Proposition 1

We solve here the system of conditions (16)-(17)-(18). Substituting from equation (13), condition (16) yields

$$\sigma_\theta^2 - a\alpha^2\sigma_\eta^2 - \frac{a+b}{1-c}\sigma_\theta^2 = 0$$

and

$$a(c,b) = \frac{(1-c)\sigma_\theta^2 - b\sigma_\theta^2}{\sigma_\theta^2 + (1-c)\alpha^2\sigma_\eta^2}, \tag{A.3}$$

Similarly, equation (17) yields

$$\sigma_\theta^2 - \frac{a+b}{1-c}\sigma_\theta^2 - \frac{b}{1-c}\alpha^2\sigma_\varepsilon^2 = 0$$

and

$$b(a,c) = \frac{(1-c)\sigma_\theta^2 - a\sigma_\theta^2}{\sigma_\theta^2 + \alpha^2\sigma_\varepsilon^2}. \tag{A.4}$$

Combining (A.3) and (A.4), we get:

$$a(c) = \frac{\tau_\eta}{\frac{1}{1-c}\tau_\eta + \tau_\varepsilon + \tau_\pi} \tag{A.5}$$

$$b(c) = \frac{(1-c)\tau_\varepsilon}{\frac{1}{1-c}\tau_\eta + \tau_\varepsilon + \tau_\pi} \tag{A.6}$$

Substituting (A.5) and (A.6) into the endogenous signal $x_i$ and into the individual expectation $E[\theta|\Omega_i]$ yields

$$x_i = \frac{\frac{1}{1-c}\tau_\eta + \tau_\varepsilon}{\frac{1}{1-c}\tau_\eta + \tau_\varepsilon + \tau_\pi}\alpha\pi + \frac{\tau_\varepsilon}{\frac{1}{1-c}\tau_\eta + \tau_\varepsilon + \tau_\pi}\alpha\varepsilon + \frac{1}{1-c}\kappa + \xi_i$$

and

$$\begin{aligned} E[\theta|\Omega_i] &= \frac{\frac{1}{1-c}\tau_\eta + \tau_\varepsilon}{\frac{1}{1-c}\tau_\eta + \tau_\varepsilon + \tau_\pi}\alpha\pi + \frac{\tau_\varepsilon}{\frac{1}{1-c}\tau_\eta + \tau_\varepsilon + \tau_\theta}\alpha\varepsilon \\ &+ \frac{\frac{1}{1-c}\tau_\eta}{\frac{1}{1-c}\tau_\eta + \tau_\varepsilon + \tau_\pi}\alpha\eta_i + c\left(\frac{1}{1-c}\kappa + \xi_i\right). \end{aligned}$$

Using these two expressions, the third orthogonality condition (18) can be written as

$$\frac{\frac{1}{1-c}\tau_\eta + \tau_\varepsilon}{\frac{1}{1-c}\tau_\eta + \tau_\varepsilon + \tau_\pi}\alpha^2\tau_\pi^{-1} - \left(\frac{\frac{1}{1-c}\tau_\eta + \tau_\varepsilon}{\frac{1}{1-c}\tau_\eta + \tau_\varepsilon + \tau_\pi}\right)^2 \alpha^2\tau_\pi^{-1}$$
$$- \left(\frac{\tau_\varepsilon}{\frac{1}{1-c}\tau_\eta + \tau_\varepsilon + \tau_\pi}\right)^2 \alpha^2\tau_\varepsilon^{-1} - c\left(\frac{1}{(1-c)^2}\tau_\kappa^{-1} + \tau_\xi^{-1}\right) = 0.$$

Simplifying this yields the fixed-point equation stated in Proposition 1.

### A.4.2 Proof of Proposition 2

The Proposition follows from the fact that, in equilibrium we have

$$\alpha^2 = (1-c)c\tau_\eta^{-1}\left(\frac{1}{1-c}\tau_\eta + \tau_\varepsilon + \tau_\pi\right)^2\left(\frac{1}{(1-c)^2}\tau_\kappa^{-1} + \tau_\xi^{-1}\right).$$

with $\tau_\eta, \tau_\varepsilon, \tau_\pi, \tau_\kappa, \tau_\xi$ being all positive.

### A.4.3 Proof of Proposition 4

The fixed-point equation (23) implies that, in equilibrium,

$$\alpha(c) = \left(\frac{1}{1-c}\tau_\eta + \tau_\varepsilon + \tau_\pi\right)\sqrt{\frac{(1-c)c}{\tau_\eta\tau_\xi}}.$$

Note that

$$\frac{\partial\alpha}{\partial c} = \frac{1}{2c(1-c)^2}\sqrt{\frac{(1-c)c}{\tau_\eta\tau_\xi}}\left[\tau_\eta + (1-2c)(1-c)(\tau_\varepsilon + \tau_\pi)\right]$$

and that $\partial c/\partial\alpha = (\partial\alpha/\partial c)^{-1}$. Therefore we get

$$\begin{aligned}
\frac{\partial\phi}{\partial\alpha} &= \alpha\frac{\partial\left(\frac{\tau_\varepsilon}{\frac{1}{1-c}\tau_\eta+\tau_\varepsilon+\tau_\pi}\right)}{\partial c}\frac{\partial c}{\partial\alpha} + \frac{\tau_\varepsilon}{\frac{1}{1-c}\tau_\eta + \tau_\varepsilon + \tau_\pi} \\
&= \frac{\tau_\varepsilon}{\frac{1}{1-c}\tau_\eta + \tau_\varepsilon + \tau_\pi} - \frac{\alpha\tau_\varepsilon\tau_\eta}{(\tau_\eta + (1-c)(\tau_\pi + \tau_\varepsilon))^2} \\
&\times \left[\frac{1}{2c(1-c)^2}\sqrt{\frac{(1-c)c}{\tau_\eta\tau_\xi}}(\tau_\eta + (1-2c)(1-c)(\tau_\varepsilon + \tau_\pi))\right]^{-1},
\end{aligned}$$

which is **weakly negative** if and only if

$$
\begin{aligned}
&\frac{\tau_\varepsilon}{\frac{1}{1-c}\tau_\eta+\tau_\varepsilon+\tau_\pi} \\
&\leq \quad \frac{\alpha\tau_\varepsilon\tau_\eta}{\left(\tau_\eta+(1-c)\left(\tau_\pi+\tau_\varepsilon\right)\right)^2}\left[\frac{1}{2c\left(1-c\right)^2}\sqrt{\frac{(1-c)c}{\tau_\eta\tau_\xi}}\left(\tau_\eta+\left(1-2c\right)\left(1-c\right)\left(\tau_\varepsilon+\tau_\pi\right)\right)\right]^{-1}.
\end{aligned}
$$

Applying the fixed-point equation gives

$$
\begin{aligned}
&\frac{\tau_\varepsilon}{\frac{1}{1-c}\tau_\eta+\tau_\varepsilon+\tau_\pi} \\
&\leq \quad \frac{\tau_\eta}{\left(1-c\right)^2}\frac{\alpha\tau_\varepsilon}{\left(\frac{1}{1-c}\tau_\eta+\left(\tau_\pi+\tau_\varepsilon\right)\right)^2}\left[\frac{1}{2c\left(1-c\right)^2}\sqrt{\frac{(1-c)c}{\tau_\eta\tau_\xi}}\left(\tau_\eta+\left(1-2c\right)\left(1-c\right)\left(\tau_\varepsilon+\tau_\pi\right)\right)\right]^{-1} \\
1 &\leq \quad \frac{\tau_\eta}{\left(1-c\right)^2}\frac{\alpha}{\frac{1}{1-c}\tau_\eta+\tau_\pi+\tau_\varepsilon}\left[\frac{1}{2c\left(1-c\right)^2}\sqrt{\frac{(1-c)c}{\tau_\eta\tau_\xi}}\left(\tau_\eta+\left(1-2c\right)\left(1-c\right)\left(\tau_\varepsilon+\tau_\pi\right)\right)\right]^{-1} \\
1 &\leq \quad \frac{\tau_\eta}{\left(1-c\right)^2}\sqrt{\frac{(1-c)c}{\tau_\eta\tau_\xi}}\left[\frac{1}{2c\left(1-c\right)^2}\sqrt{\frac{(1-c)c}{\tau_\eta\tau_\xi}}\left(\tau_\eta+\left(1-2c\right)\left(1-c\right)\left(\tau_\varepsilon+\tau_\pi\right)\right)\right]^{-1} \\
1 &\leq \quad \frac{2c\tau_\eta}{\tau_\eta+\left(1-2c\right)\left(1-c\right)\left(\tau_\varepsilon+\tau_\pi\right)}. \qquad \text{(A.7)}
\end{aligned}
$$

For this inequality to hold we need, first,

$$\tau_\eta \geq -\left(1-2c\right)\left(1-c\right)\left(\tau_\varepsilon+\tau_\pi\right),$$

and second,

$$
\begin{aligned}
\tau_\eta+\left(1-2c\right)\left(1-c\right)\left(\tau_\varepsilon+\tau_\pi\right) &\leq 2c\tau_\eta, \\
\left(1-2c\right)\left(1-c\right)\left(\tau_\varepsilon+\tau_\pi\right) &\leq -(1-2c)\tau_\eta,
\end{aligned}
$$

which is satisfied for

$$\frac{1}{2} \leq c \leq 1+\frac{\tau_\eta}{\tau_\varepsilon+\tau_\pi}.$$

Therefore we conclude that a maximum in the reaction to news obtains for $c = 1/2$, which corresponds to

$$\alpha^* \equiv \alpha\left(1/2\right) = \sqrt{\frac{\tau_\eta}{\tau_\xi}}+\frac{1}{2}\frac{\tau_\varepsilon+\tau_\pi}{\sqrt{\tau_\eta\tau_\xi}}.$$

We can check that for the structural dispersion the same inequality holds.

$$\begin{aligned}
\frac{\partial \Delta_s}{\partial c}\frac{\partial c}{\partial \alpha} &= \frac{2\alpha\tau_\eta}{\left(\frac{1}{1-c}\tau_\eta+\tau_\varepsilon+\tau_\pi\right)^2} - \frac{2\alpha^2\tau_\eta^2(1-c)}{\left(\tau_\eta+(1-c)\left(\tau_\varepsilon+\tau_\pi\right)\right)^3} \\
&\times \left[\frac{1}{2c(1-c)^2}\sqrt{\frac{(1-c)c}{\tau_\eta\tau_\xi}}\left(\tau_\eta+(1-2c)(1-c)(\tau_\varepsilon+\tau_\pi)\right)\right]^{-1} = \\
&= \frac{2\alpha\tau_\eta}{\left(\frac{1}{1-c}\tau_\eta+\tau_\varepsilon+\tau_\pi\right)^2} - \frac{2\alpha^2\tau_\eta^2}{(1-c)^2\left(\frac{1}{1-c}\tau_\pi+\tau_\varepsilon+\tau_\eta\right)^3} \\
&\times \left[\frac{1}{2c(1-c)^2}\frac{\alpha}{\frac{1}{1-c}\tau_\pi+\tau_\varepsilon+\tau_\eta}\left(\tau_\eta+(1-2c)(1-c)(\tau_\varepsilon+\tau_\pi)\right)\right]^{-1} = \\
&= \frac{2\alpha\tau_\eta}{\left(\frac{1}{1-c}\tau_\eta+\tau_\varepsilon+\tau_\pi\right)^2} - \frac{2\alpha\tau_\eta^2}{\left(\frac{1}{1-c}\tau_\pi+\tau_\varepsilon+\tau_\eta\right)^2}\left[\frac{1}{2c}\left(\tau_\eta+(1-2c)(1-c)(\tau_\varepsilon+\tau_\pi)\right)\right]^{-1}.
\end{aligned}$$

This is weakly negative if and only if

$$-\tau_\eta\left[\frac{1}{2c}\left(\tau_\eta+(1-2c)(1-c)(\tau_\varepsilon+\tau_\pi)\right)\right]^{-1}+1\leq 0$$

or – slightly rewritten –

$$1\leq\frac{2c\tau_\eta}{\tau_\eta+(1-2c)(1-c)(\tau_\varepsilon+\tau_\pi)}$$

which is identical to the inequality (A.7) studied above.

#### A.4.4 Proof of Proposition 5

With $\tau_\xi\to\infty$ the fixed-point (22) simplifies to

$$\frac{\frac{1}{1-c}\tau_\eta}{\left(\frac{1}{1-c}\tau_\eta+\tau_\varepsilon+\tau_\pi\right)^2}\alpha^2-c\left[\frac{1}{(1-c)^2}\tau_\kappa^{-1}\right]=0. \tag{A.8}$$

From this, we can deduce that, in equilibrium, we have

$$\alpha=\left(\frac{1}{1-c}\tau_\eta+\tau_\varepsilon+\tau_\pi\right)\sqrt{\frac{c}{(1-c)\,\tau_\kappa\tau_\eta}}.$$

Note that

$$\frac{\partial\alpha}{\partial c}=\frac{1}{2c\,(1-c)^2}\sqrt{\frac{c}{\tau_\kappa\tau_\eta\,(1-c)}}\left[(1+2c)\tau_\eta+(1-c)\left(\tau_\pi+\tau_\varepsilon\right)\right]$$

and that $\partial c/\partial\alpha = (\partial\alpha/\partial c)^{-1}$. As before, we get

$$\begin{aligned}
\frac{\partial\phi}{\partial\alpha} &= \alpha\frac{\partial\left(\frac{\tau_\varepsilon}{\frac{1}{1-c}\tau_\eta+\tau_\varepsilon+\tau_\pi}\right)}{\partial c}\frac{\partial c}{\partial\alpha}+\frac{\tau_\varepsilon}{\frac{1}{1-c}\tau_\eta+\tau_\varepsilon+\tau_\pi}\\
&= -\frac{\alpha\tau_\varepsilon\tau_\eta}{(\tau_\eta+(1-c)(\tau_\pi+\tau_\varepsilon))^2}\left[\frac{1}{2c(1-c)^2}\sqrt{\frac{c}{\tau_\kappa\tau_\eta(1-c)}}((1+2c)\tau_\eta+(1-c)(\tau_\pi+\tau_\varepsilon))\right]^{-1}\\
&+ \frac{\tau_\varepsilon}{\frac{1}{1-c}\tau_\eta+\tau_\varepsilon+\tau_\pi}
\end{aligned}$$

which is **weakly negative** if and only if

$$\begin{aligned}
&\frac{\tau_\varepsilon}{\frac{1}{1-c}\tau_\eta+\tau_\varepsilon+\tau_\pi}\\
&\le \frac{\alpha\tau_\varepsilon\tau_\eta}{(\tau_\eta+(1-c)(\tau_\pi+\tau_\varepsilon))^2}\left[\frac{1}{2c(1-c)^2}\sqrt{\frac{c}{\tau_\kappa\tau_\eta(1-c)}}((1+2c)\tau_\eta+(1-c)(\tau_\pi+\tau_\varepsilon))\right]^{-1}.
\end{aligned}$$

Using the fixed-point equation this gives

$$\begin{aligned}
&\frac{\tau_\varepsilon}{\frac{1}{1-c}\tau_\eta+\tau_\varepsilon+\tau_\pi}\\
&\le \frac{\tau_\eta}{(1-c)^2}\frac{\alpha\tau_\varepsilon}{\left(\frac{1}{1-c}\tau_\eta+(\tau_\pi+\tau_\varepsilon)\right)^2}\left[\frac{1}{2c(1-c)^2}\sqrt{\frac{c}{\tau_\kappa\tau_\eta(1-c)}}((1+2c)\tau_\eta+(1-c)(\tau_\pi+\tau_\varepsilon))\right]^{-1}\\
1 &\le \frac{\tau_\eta}{(1-c)^2}\frac{\alpha}{\frac{1}{1-c}\tau_\eta+\tau_\pi+\tau_\varepsilon}\left[\frac{1}{2c(1-c)^2}\sqrt{\frac{c}{\tau_\kappa\tau_\eta(1-c)}}((1+2c)\tau_\eta+(1-c)(\tau_\pi+\tau_\varepsilon))\right]^{-1}\\
1 &\le \frac{\tau_\eta}{(1-c)^2}\sqrt{\frac{c}{(1-c)\tau_\kappa\tau_\eta}}\left[\frac{1}{2c(1-c)^2}\sqrt{\frac{c}{\tau_\kappa\tau_\eta(1-c)}}((1+2c)\tau_\eta+(1-c)(\tau_\pi+\tau_\varepsilon))\right]^{-1}\\
1 &\le \frac{2c\tau_\eta}{(1+2c)\tau_\eta+(1-c)(\tau_\pi+\tau_\varepsilon)},
\end{aligned}$$

and therefore

$$\tau_\eta+(1-c)(\tau_\pi+\tau_\varepsilon)\le 0,$$

which is never true.

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**Michael Ehrmann** (corresponding author)
European Central Bank, Frankfurt am Main, Germany; email: michael.ehrmann@ecb.europa.eu

**Gaetano Gaballo**
Banque de France, Paris, France; PSE and CEPR; email: gaetano.gaballo@gmail.com

**Peter Hoffmann**
European Central Bank, Frankfurt am Main, Germany; email: peter.hoffmann@ecb.europa.eu

**Georg Strasser**
European Central Bank, Frankfurt am Main, Germany; email: georg.strasser@ecb.europa.eu





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