# Bond Supply, Yield Drifts, and Liquidity Provision Before Macroeconomic Announcements[*]

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### Abstract

UK government bond yields tend to rise in a two-day window before labor market data releases and monetary policy news. This effect, particularly pronounced during UK bond issuances, is linked to higher term premia. Financial intermediary constraints play a role as dealers avoid accumulating inventory in pre-news windows with issuances. The composition of liquidity providers also shifts: hedge funds buy a larger share of the bond issuance outside pre-news windows, but more passive investors, such as foreign central banks and pension funds, provide liquidity in pre-news windows. We outline a simple model to rationalize these findings.

*Keywords*: Macroeconomic Announcements, Yield Drift, Bond Supply, Liquidity Provision

*JEL codes*: G11, G12, G14, D83, D84

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# 1 Introduction

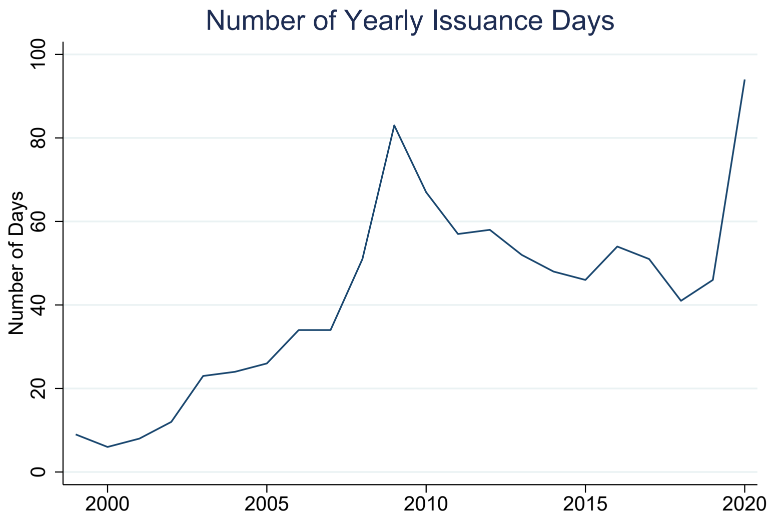
Two of the most important drivers of nominal interest rates are the arrival of macroeconomic news (Gürkaynak, Sack, and Swanson, 2005a,b) and changes in government bond supply (Greenwood and Vayanos, 2014). Given the rising levels of government debt and increased primary issuance in developed markets in recent years, it is natural to presume that news and bond supply effects on yields have become increasingly important and intertwined. Figure 1 illustrates the sharp rise in the frequency of government bond issuance and the increased co-occurrences of bond issuance and macroeconomic announcements in the UK. For example, since 2005, the majority of the Bank of England's Monetary Policy Committee (MPC) announcements have been preceded by government bond issuance in the two days preceding. Yet, how the interaction between these two forces affect investor behaviour, bond market liquidity, and ultimately interest rates is not well understood. Our paper uses both aggregate and granular, transaction-level data to empirically study this interaction and its effect on interest rates. We also present a simple model to illustrate the mechanics of secondary market trading following bond issuance and prior to news events such as the arrival of labour market data releases and monetary policy announcements.

We start our empirical analysis by documenting that long-term bond yields systematically rise in a two-day window before the scheduled arrival of macroeconomic news such as labour market data releases and monetary policy announcements, which we refer to as 'pre-news windows'. We refer to the yield changes in pre-news windows as 'pre-news yield drift'. Over our sample period of 1997-2021, this pre-news drift pushed up yields by about two percentage points in 10-20 year maturities, which is non-negligible compared to a total fall of 6-7 percentage points since 1997. The effect concentrates in pre-news windows that coincide with new issuance of government bonds. For example, as a baseline, the average daily change in 10- and 20-year yields during pre-news windows is 0.3-0.5 bps larger than yield changes outside pre-news windows. But this difference rises to 0.6-1.1 bps when the pre-news window coincides with primary issuances. Decomposing the pre-news drift into a term premium component and changes in expectations about future short-term interest rates, we find that term premia play a dominant role in our results.

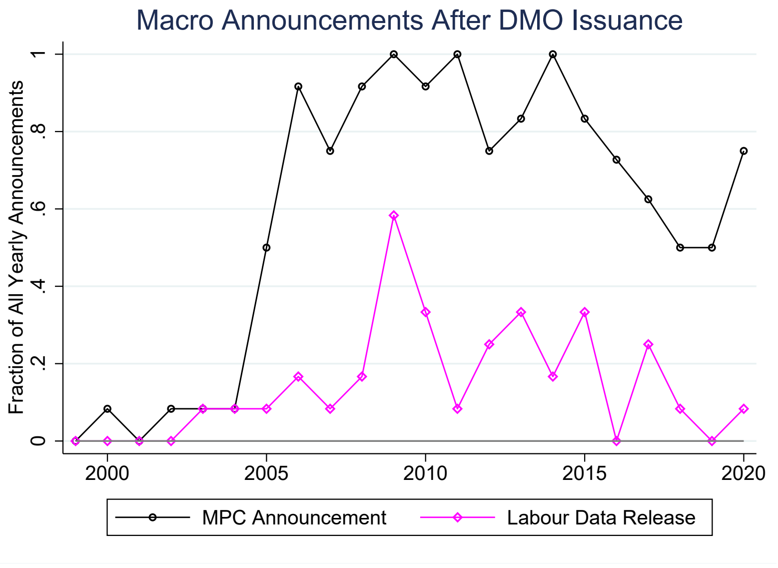
To analyse the mechanisms underlining the pre-news yield drift and its interaction with primary issuances, we study the behaviour of both primary dealers and clients. The explanation we explore, which we also formalise in a theoretical model, relates to the limited risk-bearing capacity of primary dealers during government bond issuance, which becomes more pronounced when issuance is closely followed by an informationally sensitive period such as a macroeconomic announcement.

Figure 1: Primary Issuance and Macroeconomic Announcements: 1999-2020

(a) Trading Days with Bond Issuance



(b) Fraction of Macro-news Days with Preceding Issuance Days



Notes: Panel A of this figure shows the total number days (in a given year) on which new nominal or inflation-linked government debt was issued. Panel B shows the fraction of pre-news windows (in a given year) that coincided with new government bond issuance. Pre-new windows are defined as trading days that are either one or two days before days of monetary policy announcements (black line) or labour market data releases (magenta line).

To empirically analyse the relevance of dealer constraints we employ various empirical proxies. First, we use our transaction-level data to construct a measure of inter-dealer price dispersion as an alternative proxy for dealer constraints (Eisfeldt, Herskovic, and Liu, 2023), and show that pre-news yield drift has a larger interaction with primary issuances during periods of higher inter-dealer price dispersion. Second, we use our transaction-level data to construct a measure of inventory imbalance at the dealer-day level, and show that dealers with larger imbalances prior to issuance end up selling more after the issuance during pre-news windows than those with smaller imbalances.[1]

Regarding the behaviour of clients, we find evidence on changes in the composition of liquidity providers. Hedge funds tend to buy a large share of the new issue outside pre-news windows, but their relative importance in liquidity provision decreases in pre-news windows, with more passive investors such as foreign central banks and pension funds increasing their share in liquidity provision. Further analyses confirm that hedge funds are better informed – measured by their ability to forecast future price movements – on announcements days, possibly because they are better at understanding and/or interpreting public news and the implications for the broader economy and markets. Based on this observation, we propose a simple model to rationalize both the change in the composition of liquidity providers before information events and the pre-news yield drift observed in the data. The model explores the idea that soon-to-be-informed clients, such as hedge funds, choose to refrain from liquidity provision as a way to mitigate their expected price impact when they acquire an information advantage.

Our model features two trading days, the pre-news day and the announcement day, and three types of agents, dealers, uninformed clients, and informed clients, where the information status is relevant only for the second trading day. While informed clients are fully strategic, other non-strategic (price-taking) agents impute informed clients' demand imperfectly by observing the market price. Naturally, when trading on the announcement day, informed clients face a price impact per share traded. In turn, not only do informed clients cut down their demand on the announcement day, but they also cut down their demand in the pre-news day so as to enter the next trading day with a moderate asset position. Put differently, the soon-to-be-informed clients effectively face a price impact in the pre-news day as well, because they trade in a way to mitigate their expected price impact in the second period.

Our model therefore points to a novel conflict between liquidity provision and an anticip-

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[1] We also include in Appendix section A.3.6 results conditioned on high and low yield curve noise using Hu, Pan, and Wang (2013)'s measure as a proxy for financial intermediary constraints. Ashtari-Tafti, Guimaraes, Pinter, and Wijnandts (2024) show its use in documenting liquidity-dependent monetary policy transmission to long-term yields, linked to arbitrage capital availability. Their results suggest that the noise measure relates more to hedge funds than dealers, whereas our focus is on dealer-based measures.

ated information advantage. Contrary to what liquidity provision would require, (soon-to-be) informed clients are less willing to acquire extreme positions before information events. On average, liquidity provision by informed clients falls in these periods, and uninformed clients step in to supply liquidity. However, these passive clients are risk averse and demand high risk compensation to buy the newly issued bonds from the dealers especially during times of heightened uncertainty, generating a significant price drop in equilibrium.

Moreover, the price effects of the interaction between monetary policy announcements and government bond issuance are not specific to the UK; they are also present in other markets such as the US. In a recent paper, Hillenbrand (2020) shows that movements in a narrow window around monetary policy meetings of the Federal Reserve Board's Federal Open Market Committee ('FOMC windows') explain the secular decline in nominal long-term interest rates over the last three decades.[2] At first sight, this appears to be counter to the findings of our paper. However, we argue that the effect our paper identifies is present in the US as well. We show that all of the yield drift that Hillenbrand (2020) identifies concentrates in FOMC windows that do not coincide with issuance of US treasuries longer than four years maturity. These FOMC windows that do not coincide with long-term debt issuance account for about two thirds of all FOMC windows in the sample. In the remaining one third of FOMC windows featuring long-term debt issuance, there is no significant change in yields. Moreover, we show that the term premium component actually increases during FOMC windows featuring bond issuance, consistent with the UK evidence, and this effect concentrates in the period after the Great Recession of 2008 during which dealers in the US Treasury market have become increasingly constrained (Duffie, 2020; Du, Hebert, and Li, 2023).[3]

**Related Literature** Our paper is related to the recent literature on asset price movements before central bank announcements.[4] The majority of this literature has focused on stock markets (Lucca and Moench, 2015; Bernile, Hu, and Tang, 2016; Ai and Bansal, 2018; Neuhierl and Weber, 2018; Cieslak, Morse, and Vissing-Jorgensen, 2019; Laarits, 2020; Ai, Bansal, and

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[2] Contrary to our setting, Hillenbrand (2020) uses a window which includes, in addition to the announcement day, the day before and the day after the announcement. We also report how the US results change when we apply our focus on two-day windows before announcements. We find an increase in bond yields (and risk premia) in the sample of FOMC announcements that coincide with long-term Treasury issuance, which is consistent with the UK results in this paper.

[3] In other words, the learning effect emphasized in Hillenbrand (2020) is to some extent offset by the liquidity effect, induced by the interaction between monetary policy announcements and changes in government debt supply, which our paper focuses on.

[4] This is complementary to the literature that focuses on asset price movements after central bank announcements (Cochrane and Piazzesi, 2002; Gertler and Karadi, 2015; Hanson and Stein, 2015; Nakamura and Steinsson, 2018; Albagli, Ceballos, Claro, and Romero, 2019; Hanson, Lucca, and Wright, 2021; Miranda-Agrippino and Ricco, 2021; Kroencke, Schmeling, and Schrimpf, 2021; Pflueger and Rinaldi, 2022; Karnaukh and Vokata, 2022).

Han, 2021; Hu, Pan, Wang, and Zhu, 2022). A small set of papers have looked at how bond prices move around monetary policy announcements (Savor and Wilson, 2013; Brooks, Katz, and Lustig, 2018; Hillenbrand, 2020).

Our contribution to this literature is threefold. First, we highlight the role of bond supply effects (Lou, Yan, and Zhang, 2013; Greenwood and Vayanos, 2014; Vayanos and Vila, 2021) in shaping asset price dynamics around central bank announcements. Second, we exploit transaction-level data to study the role of investor behaviour in explaining the observed price patterns around monetary policy announcements. Most papers that study pre-announcement price drifts use aggregate data, which limits their ability to identify the mechanisms at play. Third, we develop a simple theoretical framework to capture both the changing composition of liquidity providers before informational events and how prices are affected by the interaction between asset supply and informational events. Existing theoretical work is typically silent on the drivers of uncertainty before central bank announcements and how it is resolved (e.g., Hu, Pan, Wang, and Zhu, 2022). Ai, Bansal, and Han (2021) and our paper highlight the role of informed traders before central bank announcements. Differently from Ai, Bansal, and Han (2021), we model dealers and other uninformed traders separately. This allows us to analyse how dealers' increased risk-sharing need after primary issuances contributes to the pre-news drifts, when post-issuance and pre-news periods coincide.

Our paper is also related to the literature that highlights the role of dealers' balance sheet constraints in liquidity provision, and analyses changes in the tightness of these constraints since the Great Recession (Duffie, 2020; Augustin, Chernov, Schmid, and Song, 2021; He, Nagel, and Song, 2022; Du, Hebert, and Li, 2023).[5,6] Our paper's contribution to this literature is to demonstrate that periods in which dealers are desperate to sell to the client sector and in which clients are particularly averse to buying may overlap, such as during pre-announcement windows after a primary issuance. Our results therefore call for communication and strategic interactions between monetary and fiscal authorities to determine the optimal timing/mechanism of bond issuance around monetary policy announcements.

Similar to our paper, Kekre, Lenel, and Mainardi (2024) also highlight the importance of intermediaries' balance sheets in the dynamics of term premia. In our work, balance sheet constraints amplify price effects *prior to announcements*, particularly when bond issuance coincides

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[5] For related analysis on corporate bond markets, see Adrian, Boyarchenko, and Shachar (2017); Bao, O'Hara, and Zhou (2018); Bessembinder, Jacobsen, Maxwell, and Venkataraman (2018) and references therein. See also the theoretical papers by Gertler and Kiyotaki (2010); He and Krishnamurthy (2013); Brunnermeier and Sannikov (2014); Drechsler, Savov, and Schnabl (2018) among others that explore business cycle implications.

[6] There is also a literature that highlights the role of dealers' market power (An and Song, 2020; Pinter and Uslu, 2022; Eisenschmidt, Ma, and Zhang, 2024). We instead model perfectly competitive and risk-averse dealers where risk aversion stems from their balance sheet constraints, which are tighter during times of increased uncertainty.

with these events. In contrast, Kekre, Lenel, and Mainardi (2024) study the *post-announcement response*, whereby intermediary positioning and equity influence how yields adjust after the realisation of monetary policy news. An additional difference in the theoretical framework is that informational effects are absent in Kekre, Lenel, and Mainardi (2024), as they focus solely on the post-announcement yield response due to dealers' wealth effects.

The remainder of the paper is organised as follows. Section 2 describes the sources for our aggregate and transaction-level data; Section 3 presents the baseline results based on aggregate data; Section 4 provides empirical evidence on the mechanism; Section 5 describes our theoretical model; Section 6 concludes.

# 2 Data

## *2.1 Macroeconomic Announcements*

Our sample of monetary policy announcements starts in 1997 when the Bank of England gained operational independence. Since then, the Monetary Policy Committee (MPC) has been responsible for conducting monetary policy in the UK.[7] The MPC comprises 9 members: the Governor, three Deputy Governors, the Chief Economist, and four external members appointed by the Chancellor of the Exchequer. Scheduled MPC meetings in the first part of our sample were organised every month. This was followed by a reduction in the frequency of scheduled meetings in recent years, to 8 times per year. After each meeting, monetary policy decisions, including the decision on the latest policy rate, are announced at 12:00pm. This typically occurs on the first or the second Thursday of the month. Our sample of all (scheduled) monetary policy announcements sample includes 273 (270) release days over the period from May 1997 to July 2021.

Our sample of labour market data releases starts in 1998. This is later than the monetary policy announcements sample because the monthly Labour Market statistics data release by the UK's Office for National Statistics was first published in April 1998.[8] The releases typically occur on Wednesdays following the week when monetary policy announcements occur. The full labour market data releases sample includes 280 scheduled release days over the period from from April 1998 to July 2021.[9]

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[7] The laws governing the creation of the MPC were laid out in the Bank of England Act 1998.

[8] Prior to April 1998 there was no integrated monthly release and the Labour Force Survey estimates were published separately, and on different dates, from other labour market statistics.

[9] The 270 monetary policy meeting dates are presented in Table A.15 of the Appendix. Further information on historical MPC meetings can be found on the Bank of England website. The 280 scheduled labour market data release dates are presented in Table A.16 of the Appendix.

## *2.2 Government Bond Issuance*

Data on government bond issuance are from the Gilt Issuance Calendars published by the UK Debt Management Office (DMO). We use the historical yearly reports, corresponding to a given financial year, that include information on the operation date, instrument name, nominal amount issued, cash raised, and issuance method.[10] In our sample of around 6000 trading days, almost 1000 days coincide with primary issuances of government bonds. A key motivation behind our study is the fact that a significant fraction of monetary policy announcements and labour market data releases occur 1-2 days after issuance of government bonds. Specifically, out of the 540 days that fall in pre-news windows ahead of monetary policy announcements, 296 days coincide with the issuance of government bonds. Out of the 560 days that fall in pre-news windows ahead of labour market data releases, 78 days coincide with the issuance of government bonds.[11]

[Table 1]

The summary statistics in Panel A of Table 1 show that both the mean and median issued amount is around £2.4 billion. Looking at the subsamples reveals that both the average issue size and issuance frequency have increased markedly over time. The mean issued amount grew from £1.6 billion in 1997-2007 to £2.6 billion in 2008-2021. The table also reveals that newly issued government debt in the UK tends to have a long-term maturity structure.

The co-occurrence of pre-news windows and issuance has increased as the frequency of government bond issuance has risen. To show this, Figure 1a shows the time-series of the number of days in a given calendar year that coincided with new bond issuance. The number has steadily increased from approximately 10 issuance days in 2000 to around 50 annual issuance days in recent years (exceptions include the crisis years of 2009 and 2020). Given the increased issuance activity, the co-occurrence of pre-news windows and bond issuance has increased markedly, especially after 2004 (Figure 1b). For example, the years 2009, 2011, and 2014 saw all pre-news windows ahead of monetary policy announcements coinciding with bond issuance.

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[10] The data can be downloaded from the DMO's website. Data on issuance pertaining to the period before April 1998 were obtained from the Bank of England.

[11] Note that the settlement date (date on which transfer of gilt and payment occur) in the gilt market is by convention the next business day after the trade is conducted (T+1). (However, other settlement dates may be negotiated bilaterally.) This may put pressure on dealers to off-load on the day after a primary issuance occurred.

## *2.3 Bond Yields*

We use data at daily frequency on zero coupon bond yields on UK government bonds, as constructed by the Bank of England.[12] This includes nominal and real yield curves and the implied inflation term structure for the UK, that are derived using spline-based techniques (Anderson and Sleath, 2001). In our baseline regressions, we use data for yields at maturities of 5, 10, 15, and 20 years. Panel B of Table 1 summarises our dataset on daily yield changes, which will be used as dependent variables in the analysis below. Daily changes in yields average around -0.1 bps in our sample, which amounts to a decline of more than 6 percentage points in our sample, consistent with the secular decline in interest rates during this period.

To decompose daily yields into term premia and expectations components, we use a dynamic no-arbitrage affine term structure model, estimated by linear regression techniques (Adrian, Crump, and Moench, 2013; Malik and Meldrum, 2016) as summarised by Section A.2 of the Appendix. We use as factors the first five principal components of the yield curve. We estimate the factor loadings at monthly frequency, and combine these estimates with the daily time-series of the factors to obtain daily estimates of term premia and expectations components. The obtained decomposition is similar to recent estimates of the UK term structure (Moench, 2019).

## *2.4 Transaction-level Data*

To study the microstructure of UK bond markets around macroeconomic announcements and debt issuance, we use a detailed transaction-level dataset which contains information on the identity of both sides of a trade. The ZEN database covers the period between August 2011 and December 2017, and MIFID II database covers the period from January 2018 to December 2019. Both datasets are sourced by the UK Financial Conduct Authority, and contain information on client and dealer identities along with information on the transaction time, the transaction price and quantity, the International Securities Identification Number, the account number, and buyer-seller flags.[13]

Our analysis focuses on transactions that occur between clients and designated market makers, called Gilt-Edged Market Makers (GEMMs). GEMMs are the primary dealers in the UK government bond market, hence are allowed to participate in the primary auctions organised

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[12] The data can be downloaded from the Bank of England's website.

[13] For further details on the Zen dataset, see the Transaction Reporting User Pack: https://www.fca.org.uk/publication/finalised-guidance/fg15-03.pdf. For further details on the MIFID II dataset, see the Reporting Guidelines: https://www.esma.europa.eu/sites/default/files/library/2016-1452_guidelines_mifid_ii_transaction_reporting.pdf. Recent applications of the datasets can be found in Czech, Huang, Lou, and Wang (2021); Kondor and Pinter (2022); Pinter and Uslu (2022); Pinter, Wang, and Zou (2024) among others.

by the DMO. Moreover, the majority of client-dealer trades are intermediated by GEMMs.[14] After filtering out all duplicates and erroneous entries, we are left with approximately 3.5 million observations for government bond market trades, nominal bond transactions making up for about two thirds of theses trades and inflation-linked bond trades accounting for the remaining one third. We identify around 600 clients that cover the majority of trading volume between clients and dealers in both segments of the government bond markets. We classify these clients by various types as detailed further below. Section A.1 in the Appendix provides further details and presents summary statistics.

# 3 Pre-news Yield Drift

## 3.1 *Baseline Results*

### 3.1.1 *Yield Drift before Macroeconomic Announcements*

This section starts by documenting that long-term UK government bond yields tend to rise during 2-day windows before the arrival of monetary policy announcements or labour market data releases. To estimate the baseline effect, we run the following regression at daily frequency:

$$\Delta_{t-1,t} r_k = \beta_0 + \beta_1 D_t^{News} + \varepsilon_t, \tag{3.1}$$

where $\Delta_{t-1,t} r_k$ is the daily change in bond yield with maturity $k$, $\beta_0$ is a constant and $D_t^{News}$ is an indicator variable which takes value of one during the pre-news window (i.e. on days that are either one or two days before the macroeconomic announcement) and zero otherwise. The estimated value of $\beta_1$ is the coefficient of interest, which captures the pre-news drift. Table 2 presents the results separately for yields with 5, 10, 15, and 20 years of maturities with panel A and B showing the estimates for pre-news windows corresponding to monetary policy announcements and labour market data releases, respectively.

[Tables 2–3]

As shown in Panel A, the average daily increase in 10, 15, and 20-year yields during the pre-news windows ahead of monetary policy announcements is about 0.5 bps larger than yield changes outside of these windows. Yield drift seems weaker at shorter maturities. The constant is estimated to be about -0.15 bps, which is consistent with the secular decline in interest rates

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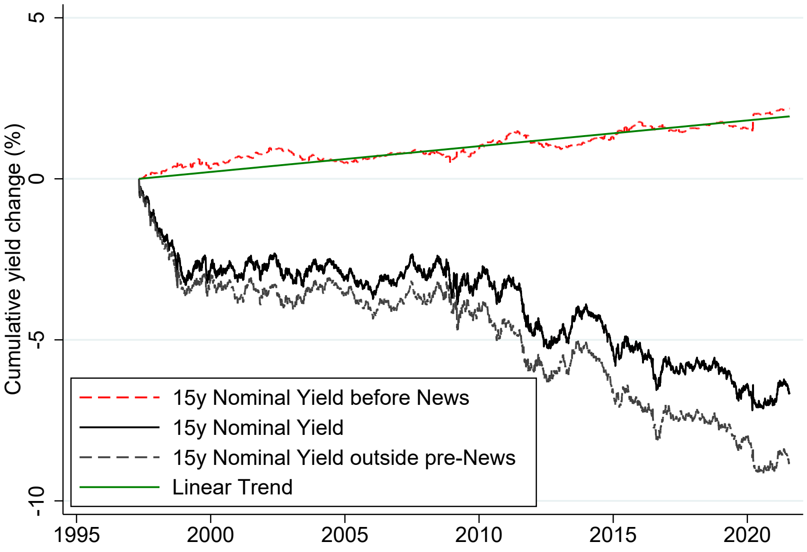
[14] For further details on the identities of GEMMs, see https://www.dmo.gov.uk/responsibilities/gilt-market/market-participants/.

during this period.[15] As shown in Panel B of Table 2, we find a similar yield drift before scheduled labour market data releases, which amount to about 0.4-0.5 bps.

Figure 2: A Decomposition of Long-term Gilt Yields: pre-MPC Yield Drift



Note: this figure documents that 15-year UK nominal bond yields tend to rise in the 2-day window before MPC meetings. This 2-day window includes for every MPC meeting the day prior to the meeting, and the day that is two days before the meeting. The black line shows the actual evolution of the yield. The red line shows a hypothetical time series that is constructed by taking into account only the yield changes that were realised in the 2-day window before MPC meetings; the yield changes that occurred on all days outside of this window are set to zero. The green line is an estimated linear trend associated with the red line. The gray line shows a hypothetical time series that is constructed by taking into account only the yield changes that were realised outside the 2-day window before MPC meetings. The analysis includes all 273 MPC meetings from May 1997 to June 2021.

To visually illustrate the pre-news drift, we construct hypothetical time-series based on 15-year yield changes that realised during the 2-day period before macroeconomic announcements, assuming that yield movements outside these periods were zero. Similarly, we construct hypothetical time-series that take into account changes in the 15-year yield only outside pre-news windows. The two hypothetical time-series along with the cumulative changes in the realised 15-year yield are presented in Figures 2–3. The solid green line is linear trend fitted on the pre-news drift series. The results highlight that the pre-news drift is sizeable, amounting to a cumulative effect of around 2% over the 1997-2021 period. As a comparison, cumulative changes in realised yields amount to -6.5% during this period.

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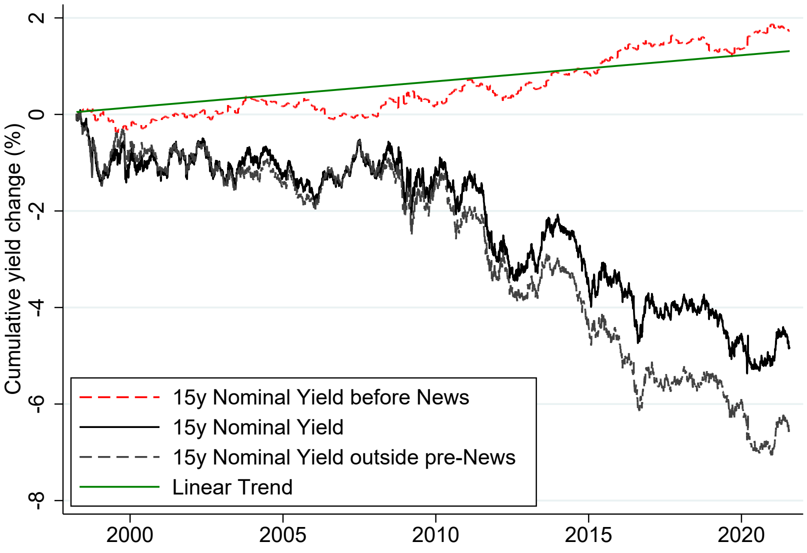
[15] These results are based on scheduled monetary policy announcements. During our sample period, there were three unscheduled monetary policy announcements that typically occurred during high-volatility periods (September 18, 2001; March 11, 2020; March 19, 2020). To explore the effect of unscheduled monetary policy announcements on our results, we extend our sample of windows accordingly, and re-estimate our regressions. Appendix Table A.18 shows that our baseline results are statistically and economically stronger, which is driven by the announcements during the COVID-19 crisis period (see Hauser (2020)).

Figure 3: A Decomposition of Long-term Gilt Yields: pre-Labour Market Data Release Drift



Note: this figure documents that 15-year UK nominal bond yields tend to rise in the 2-day window before scheduled labour market data releases. This 2-day window includes for every release date the day prior to the meeting, and the day that is two days before the meeting. The black line shows the actual evolution of the yield. The red line shows a hypothetical time series that is constructed by taking into account only the yield changes that were realised in the 2-day window before release dates; the yield changes that occurred on all days outside of this window are set to zero. The green line is an estimated linear trend associated with the red line. The gray line shows a hypothetical time series that is constructed by taking into account only the yield changes that were realised outside the 2-day window before data release dates. The analysis includes all 280 labour market data release days from April 1998 to June 2021.

Given possible intra-week seasonalities in bond markets (Flannery and Protopadakis, 1988), we also check how these results change when we include in regression (3.1) dummy variables corresponding to weekdays.[16] As shown in Table 3, we obtain qualitatively similar results to our baseline, though the pre-news drift is quantitatively weaker (stronger) for monetary policy announcements (labour market data releases). Inspecting the estimates of the dummy variables reveals that yields tend to fall on Mondays, and Tuesdays and Wednesdays are associated with yield increases. The estimated intra-week pattern of yield changes, coupled with the fact that most labour data releases and monetary policy announcements occur on Wednesdays and Thursdays, respectively, explains why the inclusion of intra-week dummies weakens the pre-news drift ahead of monetary policy announcement and strengthens the pre-news drift ahead of labour market data releases.[17]

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[16] Specifically, we include four dummy variables corresponding to Tuesdays, Wednesdays, Thursdays and Fridays. In this modified regression, the constant $\beta_0$ captures the average daily yield change on Mondays and the dummy variables capture the average daily yield change relative to Mondays.

[17] Appendix Section A.3.3 provides further details on the links among intra-week seasonality, yield dynamics and our baseline results.

### *3.1.2 Pre-News Windows and Primary Issuance*

To analyse the role of debt issuance, we assign all available pre-news windows in our dataset into two groups: one that coincides with new issuance of either nominal or indexed-linked government debt and the remaining pre-news windows without debt issuance. We thereby extend regression (3.1) as follows:

$$\Delta_{t-1,t} r_k = \beta_0 + \beta_1 D^{News}_{ISS,t} + \beta_2 D^{News}_{noISS,t} + \beta_3 D^{noNews}_{ISS,t} + \varepsilon_t, \tag{3.2}$$

where $D^{News}_{ISS,t}$ is an indicator variable which takes value one when the pre-news windows coincide with debt issuance and zero otherwise. Similarly, $D^{News}_{noISS,t}$ indicates pre-news windows without debt issuance. $D^{noNews}_{ISS,t}$ serves as a control variable indicating days with issuance and without macroeconomic news.

[Tables 4–5]

Table 4 presents the estimated coefficients for model (3.2) for the four different maturities, with Panel A and B presenting the results for pre-news windows associated with monetary policy announcements and labour market data releases, respectively. As shown by Panel A, the effects tend to be economically and statistically significant during pre-news windows ahead of monetary policy announcements with debt issuance. During these windows, changes in 15 and 20-year yields tend to be about 0.65 bps larger compared to days that are outside pre-news windows and do not coincide with primary issuances. As shown by Panel B, the effects are economically stronger in pre-news windows ahead of labour market data releases with debt issuance, with yield changes being about 1.1 bps larger across the four maturities we study. As shown in Table 5, we obtain qualitatively similar results when we include in regression (3.2) dummy variables corresponding to weekdays, although the interaction between primary issuance and pre-news windows is quantitatively weaker (stronger) for monetary policy announcements (labour market data releases).[18]

### *3.1.3 A Decomposition: Term Premia vs Expectations*

Next, we check whether the pre-news yield drift can be linked to an increase in term premia or to higher expectations for the future path of the short-term interest rate. To that end, we use a dynamic no-arbitrage affine term structure model, estimated by linear regression techniques (Adrian, Crump, and Moench, 2013; Malik and Meldrum, 2016), and decompose the 10-year

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[18] Given the increasing trend in issuance activity (see Figure 1), we also check whether the baseline results are robust to including year fixed effects. Table A.17 shows that including these controls does not materially change the baseline results.

yield into estimated term premium and expectations components.[19] We then employ these time-series as left-hand side variables in our baseline regressions (3.1)–(3.2). The results are presented in Table 6.

[Table 6]

Panel A of Table 6 shows that the majority of pre-news drift ahead of monetary policy announcements in the 10-year yield is driven by the term premium component (0.42 bps), which is strongly statistically significant. The expectations component is negative (-0.10 bps) and statistically insignificant. When we decompose the pre-news drift ahead of labour market data releases, we still find a significant term premium component (0.31 bps), as shown in Panel B.

Panels C and D of Table 6 present the results for the case when we interact the pre-news windows with debt issuance. We find that most of the term premium effect concentrates in pre-news windows that coincide with debt issuance. We find that in these periods the term premium tends to rise by about 0.67 bps (0.76 bps) when we consider pre-news windows ahead of monetary policy announcements (labour market data releases).

### *3.1.4 Impact on Monetary Policy*

In this section we analyse the relationship between debt issuance and monetary policy surprises. King (2000) argues that "A transparent monetary policy reaction function means that the news should be in developments of the economy not in the announcements of decisions by the central bank (...) Hence a successful central bank should be boring... (p. 6)" According to this interpretation, a metric for successful monetary policy could be the size of monetary policy surprises, i.e. the lower the surprise, the more predictable (hence more successful) the central bank policy.

Given that debt issuance tends to generate a yield drift during the pre-news window ahead of monetary policy announcements, it could be that this also increases the volatility of interest rates *after* monetary policy announcements.[20] High-frequency changes in interest rates around monetary policy announcements are increasingly used by macroeconomists to extract the surprise component in the monetary policy reaction function and to estimate the causal effect of monetary policy.[21] To estimate the effect of primary issuances in pre-news windows

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[19] See Section A.2 of the Appendix for a description of the term structure model.

[20] This could be because bond issuance absorbs market liquidity, amplifying price movements post-news due to less liquid market conditions.

[21] See Bernanke and Kuttner (2005), Gürkaynak, Sack, and Swanson (2005a), Gertler and Karadi (2015) and many others.

on the volatility of monetary policy surprises is interesting in its own right, as it could change estimates of the monetary policy transmission.

We follow the high frequency approach to the identification of monetary policy surprises of Braun, Miranda-Agrippino, and Saha (2022), which implements the methodology of Gürkaynak, Sack, and Swanson (2005a) for the UK.[22] We use price changes in a 30-minute window around the monetary policy announcements for five different assets: the 3-month futures contracts (3m), and gilt yields with one year (1Y), two years (2Y) , five years (5Y) and ten years (10Y) maturity.[23] We then estimate the following time-series regression:

$$\left|MPShock_t^k\right| = \alpha_{year} + \beta \times PreMPCIssuance_t + \varepsilon_{i,t}, \tag{3.3}$$

where $\left|MPShock_t^k\right|$ is the absolute value of the price changes of asset $k \in \{3mF, 1Y, 2Y, 5Y, 10Y\}$ around the monetary policy announcements; the term $\alpha_{year}$ is a year fixed effect which controls for lower frequency changes in issuance policy and in the interest rate environment (e.g. the zero lower bound on interest rates after the Great Financial Crisis); the term $PreMPCIssuance_t$ is a dummy taking value one if there was a primary issuance either one or two days before the monetary policy announcements.

[Table 7]

Table 7 shows the results with panel A presenting the results for the whole sample (1997 to 2021) and panel B presenting the results for the subsample (2006 to 2021) when the frequency of issuances in pre-news windows is more elevated. When considering the full sample, we find that surprise movements in interest rate across the maturity spectrum are around 0.57-0.83 bps higher when the monetary policy announcements are preceded by bond issuance. When considering the subsample (2006 to 2021), monetary policy surprises are about 0.85-1.39 bps higher after pre-news windows that coincide with debt issuance compared to other monetary policy announcements.[24]

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[22] This high frequency approach has been followed by the rapidly expanding monetary economics literature (Gertler and Karadi, 2015; Gerko and Rey, 2017; Cesa-Bianchi, Thwaites, and Vicondoa, 2020; Bahaj, Foulis, Pinter, and Surico, 2022).

[23] The interest rate on 3-month sterling futures is typically used as a proxy for short-term interest rates (e.g. Gerko and Rey (2017); Cesa-Bianchi, Thwaites, and Vicondoa (2020)) given the relative shortage of short-term maturity government bonds in the UK.

[24] As an additional robustness check, we also experimented with the shock measures of Swanson (2021), which extended Gürkaynak, Sack, and Swanson (2005a) to separately identify surprise changes in the policy rate, forward guidance, and quantitative easing (QE). Appendix Table A.23 presents the results, highlighting that issuance mostly affects the volatility of the policy rate shock.

### *3.2 Extensions and Robustness Checks*

**Timing and Size of Issuances** Given the importance of bond issuance in driving the pre-news drift, an interesting question is whether the issuance decision is strategically timed with respect to scheduled announcements. Appendix Section A.3.1 discusses these issues and find no evidence of disproportionately larger issuances around announcements, nor does there seem to be evidence that the DMO has strategically changed the timing of issuances around announcements.

**Daily Yield Changes around Macroeconomic Announcements** In our baseline estimation, we compare yield changes during two-day windows before macroeconomic announcements to daily yield changes on other days. Appendix Section A.3.2 takes a closer look at individual days around announcements, and finds that the yield drift occurs one or two days prior to the announcement days without any significant reversal subsequently.

**Inflation Data Releases** We also check whether our results hold for macroeconomic announcements associated with inflation data releases. Appendix Section A.3.4 shows that during the majority of our sample there are no bond issuances in these pre-news windows. Consistent with this, we find no visible pre-news drift associated with inflation data releases.

**Evidence from the US** A recent paper, Hillenbrand (2020) documents that a narrow window around monetary policy meetings of the Fed captures the secular decline in nominal long-term interest rates over the last three decades. This appears to be counter to the results we find in this paper. Appendix Section A.3.5 performs a consistency check and argues that the effect we identify using UK data is present in the US as well.

# 4 Inspecting the Mechanism

The interpretation we propose to explain the empirical facts is related to the limited risk-bearing capacity of dealers during government bond issuances, which becomes more pronounced when the issuance is closely followed by informationally sensitive events such as macroeconomic announcements. Primary dealers in the UK are obliged to play an active role in the issuance and distribution of UK government bonds (DMO, 2021).[25] Fulfilling this obligation before

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[25] Specifically, primary dealers should aim to purchase at least 2% of gilt issuance by sector, conventional and index-linked, on a six-month rolling average basis. Moreover, it is expected that each wholesale dealer's bids would amount to the equivalent of at least 5% of the amounts issued, calculated on a six-month rolling average basis (p. 4 of DMO (2021)).

macroeconomic announcements poses an additional risk to the dealer sector insofar as these announcements may imminently move interest rates and therefore bond prices.[26,27]

A fresh supply of bonds in pre-news windows thereby increases dealers' risk exposure to interest rate changes, which could incentivise dealers to distribute the new issuance among clients as quickly as possible. If clients have elastic demand for these bonds as well as limited risk-bearing capacity, then they would require a risk premium in exchange of providing liquidity to the dealer sector. This could explain why bond prices fall in the pre-news window when there is a concurrent issuance of government bonds and that a significant portion of the effect on yields loads on the term premium component, consistent with the evidence in Table 6.

## *4.1 Constraints on Primary Dealers*

To underscore the relevance of constraints on financial intermediaries, we employ a two-pronged approach. First, we use our transaction-level data to construct a measure of inter-dealer price dispersion as an alternative proxy for dealer constraints (Eisfeldt, Herskovic, and Liu, 2023), and show that the pre-news yield drift has a more pronounced interaction with debt issuance during periods of higher inter-dealer price dispersion. Second, we use our transaction-level data to construct a measure of inventory imbalance at the dealer-day level, and show that dealers with larger imbalances prior to the primary issuance end up selling more after issuance during the pre-news window. In addition, Appendix section A.3.6 also presents results using a yield curve noise measure (Hu, Pan, and Wang, 2013; Ashtari-Tafti, Guimaraes, Pinter, and Wijnandts, 2024).

### *4.1.1 Aggregate Evidence using Inter-dealer Price Dispersion*

Since Jankowitsch, Nashikkar, and Subrahmanyam (2011), aggregate price dispersion has often been used as a measure of market illiquidity.[28] The recent literature shows that price dispersion in the inter-dealer market proxies inventory constraints on the dealer sector (Eisfeldt, Herskovic, and Liu, 2023). Building on this literature, we measure inter-dealer price dispersion, $D_{ID}$, is as follows:

$$D_{ID} = \sqrt{\frac{1}{N}\sum_{v}^{N}\left((P_v^\star) - \left(\overline{P}_v\right)\right)^2}, \tag{4.1}$$

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[26] Note that balance sheet constraints have tightened since the Great Recession (Adrian, Boyarchenko, and Shachar, 2017; Bessembinder, Jacobsen, Maxwell, and Venkataraman, 2018; Bao, O'Hara, and Zhou, 2018).

[27] Table A.19 confirms that interest rate volatility (especially in the case of monetary policy announcements) tends to be significantly higher on days of macroeconomic announcements.

[28] See Friewald, Jankowitsch, and Subrahmanyam (2012); Uslu (2019); Uslu and Velioglu (2019); Pinter and Uslu (2022) among many others.

where $P_v^\star$ is the transaction price corresponding to trade $v$ in the inter-dealer market, and $\overline{P}_v$ is the average hourly transaction price in a given bond. We then compute the monthly average of the trade-specific deviations (4.1), and then sort months into high and low-dispersion groups based on the yearly medians.

[Tables 8–9]

A limitation of this analysis is that the estimation period covers a narrower period from September 2011 to July 2021, which is dictated by the availability of the transaction-level data. This implies that we lose a large part of the time-series variation of our sample that was used to generate our baseline result of the pre-news drift. An implication of this is that the pre-news drift ahead of monetary policy announcements is now statistically weaker. However, we show that the point estimates appear to be higher during periods of higher inter-dealer price dispersion.[29]

Table 8 and Table 9 present the results for monetary policy announcements and labour market data releases, respectively. We continue to find a positive estimate on the interaction between pre-news windows and debt issuance that is higher and statistically significant in periods when inter-dealer price dispersion is higher compared to periods with lower price dispersion.[30]

### *4.1.2 Evidence from Transaction-Level Data*

While the results above are suggestive of the role of constraints on primary dealers in generating the yield drift induced by debt issuance that occurs before macroeconomic announcements, we go further and use our transaction-level data to measure more directly the role of dealer inventories in the mechanism. We explore the idea that some dealers have higher inventory imbalances than other dealers ahead of primary issuance, which would generate cross-dealer heterogeneity in the pressure to off-load inventory after the primary auction and during a pre-news window.

As a proxy for inventory pressures, we build on the previous literature of dealer inventories (Comerton-Forde, Hendershott, Jones, Moulton, and Seasholes, 2010; Friewald and Nagler, 2019) and use the cumulative orderflow of dealers against clients prior to issuance as a proxy for inventory imbalance. We then check its correlation with the orderflow after the issuance

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[29] As a robustness check, we also estimate regression models without splitting the sample into high-noise and low-noise periods. Instead, we include a dummy variable to capture high- and low-noise periods. As shown in Tables A.21–A.22, we obtain very similar results to our baseline.

[30] Appendix Table A.20 presents the p-values corresponding to testing for the equalities between the estimated coefficients on the pre-news drift with and without primary issuances. Even though the differences are not statistically significant at the 10% level in the entire sample, the difference become statistically more significant during periods of lower market liquidity.

during the pre-news window. Intuitively, if inventory constraints are relevant to our mechanism, then we should find a negative correlation: the more positive the dealer's inventory imbalance is before the issuance, the more the dealer is expected to sell during the pre-news window after the issuance.

Formally, we estimate the following panel-regression for each dealer $i$ and pre-news window (featuring primary issuances) $t$:

$$DealerFlow_{i,t}^{preNewspostDMO} = \alpha_i + \mu_t + \beta \times DealerFlow_{i,t-1}^{preDMO} + \varepsilon_{i,t}, \tag{4.2}$$

where $DealerFlow_{i,t}^{preNewspostDMO}$ is the cumulative dealer flow after issuances in the pre-news window; $\alpha_i$ and $\mu_t$ are dealer and time fixed effects; and $DealerFlow_{i,t-1}^{preDMO}$ is the orderflow at dealer $i$ the day before the primary issuance.

[Table 10]

Panel A of Table 10 shows that a positive inventory imbalance of the average dealer the day before the primary issuance is associated with stronger bond sales by dealers after the issuance during pre-news windows. However, dealers' willingness to participate in the auction may depend on the inventory imbalance prior to the auction.[31] We therefore estimate regression (4.2) separately on the subsample consisting of dealers participating in the auction in the given pre-news window (Panel B) and dealers that do not participate (Panel C). We find that conditional on participating in the auction, dealers with larger positive inventory imbalances have larger post-issuance sales, whereas the results for non-participating dealers are statistically insignificant. Importantly, as shown in column (3) of Table 10, all the coefficients are statistically insignificant without impending news, underscoring the importance of the interaction between macro news and bond issuance in driving dealers' inventory dynamics.[32]

Taken together, the evidence presented in this section supports the hypothesis that constraints on gilt market dealers contribute to explaining the pre-news yield drift.

## *4.2 The Composition of Liquidity Providers*

To better understand the dynamics of bond prices during the pre-news window, we also study how different types of clients trade after bond issuance during and outside pre-news windows.

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[31] As explained by DMO (2021), "The DMO also sees it as important to encourage wholesale GEMMs to participate effectively in the price formation process at auctions, to which end it envisages that each wholesale GEMM's bids would be the equivalent of at least 5% of the amounts issued, measured on a six-month rolling average basis."

[32] In addition, Appendix Section A.3.7 we also check a specification where we pool both types of news as well as pool all observations together (including those before news and no news).

### *4.2.1 Liquidity Provision After Primary Issuances*

To analyse clients' trading behaviour, we estimate the average daily orderflow of each of the six client types on different types of trading days. We consider the following six types of trading days: issuances days that are two days before macroeconomic announcements; days that are one day after issuance days and one day before announcements; issuance days that are one day before announcements; issuance days without impending news; days after issuance days without impending news; and all other trading days.

Table 11 presents the average net gilt purchases with the columns and rows showing the estimates for different client types and day types, respectively. Panel A and B of the Table show the results for the case when we consider monetary policy announcements and labour market data releases, respectively. Inspecting Panel A reveals that clients tend to buy more gilts from dealers during monetary policy announcement pre-news windows and on issuance days (by around £538 million) or on the day after the issue (by around £175 million) compared with all other days. This increase in gilt purchases is more muted on issuance days or days after issuance without impending monetary policy announcements. Moreover, the majority of the increased liquidity provision outside pre-news windows is done by hedge funds (around £340 million) and asset managers (£190 million). In contrast, liquidity provision during pre-news windows is done predominantly by other clients such as pension funds, foreign central banks and other clients.

[Table 11]

Panel B of Table 11 paints a qualitatively similar picture when pre-news windows are defined with respect to scheduled labour market data releases.[33] Clients tend to buy more gilts from dealers during pre-news windows and on issuance days (by around £687 million) compared with all other days. This increase in gilt purchases is more muted on issuance days or days after the issue without impending labour market data releases. Moreover, the majority of the increased liquidity provision outside pre-news windows is done by hedge funds (around £295 million) and asset managers (£187 million). In contrast, the relative share of these liquidity providers falls as other clients – most notably pension funds – step up their purchases in pre-news windows.

### *4.2.2 Trading Activity of Hedge Funds*

Why does the relative share of hedge funds in clients' liquidity provision fall after primary issuances, when there is an impending macroeconomic announcement? A natural explanation

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[33] Note that the first two rows of panel B are missing because there are no trading days in our sample that satisfy the given criteria. Specifically, there are virtually no issuances on Mondays and Fridays, and most labour market data releases happen on Wednesdays and Tuesdays (Appendix Table A.16 has the list of relevant dates in sample).

we explore is that hedge funds have a weaker incentive to commit to a long position (by buying the new bonds) before the arrival of news compared to issuance without impending news.

As shown by the recent literature (Czech, Huang, Lou, and Wang, 2021; Kondor and Pinter, 2022), this could happen for two reasons. First, hedge funds may have informational signals about the news by the time the issuance occurs. Assuming that the signal is negative half the time, hedge funds would on average buy less after the auction given that their signals would (half the time) require them to take short positions. Second, hedge funds may not yet have informational signals about the upcoming announcement by the time the issuance occurs, but they are expecting to receive such signals shortly. Assuming that there are costs to making large portfolio adjustment after the issuance and before the news because of information asymmetry-induced illiquidity, for example, hedge funds may choose to participate less actively to provide liquidity during issuance.

To disentangle these two explanations, we take a closer look at the timing and performance of hedge fund trades in pre-news windows that coincide with bond issuance. To measure trading performance, we follow Di Maggio, Franzoni, Kermani, and Sommavilla (2018) by computing the $T$-day-horizon return on each hedge fund trade on day $t$, measured as the percentage difference between the transaction price and a benchmark price $T$ days after the transaction date.[34] Formally, for each trade $j$, we construct the measure $Performance_j^T$ as follows:

$$Performance_j^T = \left[\ln\left(P^T\right) - \ln\left(P_j^\star\right)\right] \times \mathbf{1}_{B,S}, \tag{4.3}$$

where $P_j^\star$ is the transaction price, $P^T$ is the $T$-day ahead median transaction price of the corresponding bond, and $\mathbf{1}_{B,S}$ is an indicator function equal to 1 when the transaction is a buy trade, and equal to $-1$ when it is a sell trade. All transaction-specific returns are then averaged within day $t$ for the hedge fund sector. We compute both unweighted average as well as weighted average using the pound sterling volume of the trades as weights.

[Table 12]

Table 12 presents the unweighted and weighted performance measures over 1-, 3-, and 6-day horizons during pre-news windows as well as on announcement days with Panels A and B showing the results for monetary policy announcements and labour market data releases, respectively. We find that during these periods, hedge fund trades that predict future price movements

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[34] The $T$-day horizon starts at the start of each day and ends after $T$ days. We use overlapping time windows. For example, to compute one-day performance measures ($T = 1$), we compare all trades on day 1 to the volume-weighted average price on day 2, and compare all trades on day 2 to the volume-weighted average price on day 3, and so on.

occur predominantly on the day of the announcement.[35] For example, the unweighted 6-day performance measure of trades executed on the day of monetary policy announcements (labour market data releases) is around 13 (7.8) bps. The unweighted performance measures on these days are economically and statistically stronger than the weighted counterparts, suggestive of smaller hedge fund trades performing better than larger ones during this period. Importantly, hedge fund trades in the 2-day period before monetary policy announcements and after issuance seem less informative to the extent that both weighted and unweighted performance measures are statistically insignificant, with point estimates often negative.[36] Overall, these results support the idea that hedge funds receive informational signals about the nature and the effect of macroeconomic news after the issuance takes place. In the next section, we propose a theoretical model which can rationalise that informed clients could refrain from providing liquidity to the dealer sector after new issuance of bonds when there is an upcoming macroeconomic announcement.

# 5 Theory

In this section, we provide a theoretical model to illustrate the mechanics of secondary-market trading following a government bond issuance and prior to the revelation of a macroeconomic shock. Importantly, we are not after presenting a full-fledged, structural model amenable to realistic calibration. Instead, we develop an illustrative framework so that the reader can make sense of our empirical findings by looking at them through the lens of our theoretical framework.

## *5.1 Model Environment*

There are two trading dates, $t \in \{1, 2\}$, and one divisible risky asset whose random payoff, $\tilde{v}$, realizes after trading at $t = 2$. There are three types of rational agents: a unit-mass continuum of dealers ($D$), a unit-mass continuum of uninformed clients ($UC$), and $N > 2$ atomic hedge funds ($HF$). Each dealer and uninformed client are of zero measure, while each hedge fund has a normalized measure of one, and so, the total mass of agents in the economy equals $N + 2$. Dealers and uninformed clients have constant-absolute-risk-aversion (CARA) preferences with

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[35] Appendix Section A.3.8 complements this price-based analysis with analysing hedge fund quantities after issuance and before and during announcements. Overall, the evidence suggests that hedge funds cut back on liquidity provision to the dealers until the realisation of uncertainty associated with the macro announcement. In addition, Section A.3.8 explores how hedge fund positioning before labour market data releases and BoE announcements – and so their revealed expectation for the sign of the news – affects their trading activity in the days leading up to those announcements in weeks with issuance.

[36] An exception is the 1-day performance measure on issuance days (two days prior to monetary policy announcements) which is positive and statistically significant using the volume weighted measure. This is suggestive of some hedge funds "riding" the price drift that occurs before monetary policy announcements.

potentially different CARA coefficients. Hedge funds are risk neutral. None of the agents discount the future.

We think of the risky asset as a government bond subject to interest rate risk, which is issued shortly before trading at $t = 1$. Each dealer obtains $z \in \mathbb{R}_{++}$ shares of the asset in the primary market. Hedge funds and uninformed clients do not participate in the primary market. Hence, once secondary-market trading starts at $t = 1$, each dealer has an endowment of $z$ shares, and hedge funds and uninformed clients 0. We assume that these endowments are public information and that there is no further issuance of the asset.

The payoff of the asset consists of two parts,

$$\tilde{v} = \tilde{\theta} + \tilde{\varepsilon}. \tag{5.1}$$

Random variable $\tilde{\theta}$ is characterized by a normal distribution with mean $\theta$ and variance $\sigma_\theta^2$. Random variable $\tilde{\varepsilon}$ also follows a normal distribution with mean 0 and variance $\sigma_\varepsilon^2$. These random variables, $\tilde{\theta}$ and $\tilde{\varepsilon}$, are independent. All agents have homogeneous expectations about the asset payoff at $t = 1$. Just prior to trading at $t = 2$, hedge funds privately learn in advance the true realization, $\tilde{\theta}$, of the first component of the asset payoff but other agents, dealers and uninformed clients, do not receive such a private information. However, dealers and uninformed clients understand that the equilibrium price at $t = 2$ reflects information about hedge funds' demand, and in turn, about $\tilde{\theta}$. We will describe the details of this information revelation by the market price at $t = 2$ in the next section as part of the equilibrium definition.

Finally, to prevent the market price at $t = 2$ from perfectly revealing $\tilde{\theta}$, we assume that a noisy demand shock $\tilde{d}$ hits the market at $t = 2$. Random variable $\tilde{d}$ has a normal distribution with mean 0 and variance $\sigma_d^2$ and is pair-wise independent from both $\tilde{\theta}$ and $\tilde{\varepsilon}$.

## *5.2 Equilibrium Definition*

In this subsection, we define a dynamic trading equilibrium for the economy described above. Our equilibrium concept and the economics behind it build on Grossman and Stiglitz (1980), Kyle (1985), and Kyle (1989). First, the equilibrium price at $t = 2$ has two simultaneous roles as in Grossman and Stiglitz (1980): clearing the market and acting as a public signal about the informed parties' actions. Second, because the informed parties in our economy, hedge funds, are large players, they internalize their impact on the market price when making their trading decisions as in Kyle (1985) and Kyle (1989).

In what follows, the term "agent" refers to an infinitesimal competitive representative of the particular type for dealers and uninformed clients, while each hedge fund is a unique agent with a positive mass. Agent $i \in \{D, UC\}$ makes decisions to maximize her expected utility defined as

$$\mathbb{E}_t^i\left[Utility^i\right] = \mathbb{E}_t^i\left[-e^{-\gamma_i W_2^i}\right], \tag{5.2}$$

while hedge fund $i \in \{1, 2, ..., N\}$ maximizes

$$\mathbb{E}_t^i\left[Utility^i\right] = \mathbb{E}_t^i\left[W_2^i\right], \tag{5.3}$$

where $\mathbb{E}_t^i$ denotes the expectation with respect to agent $i$'s information set at time $t$, $W_2^i$ agent $i$'s post-trade wealth at $t = 2$, and $\gamma_i$ agent $i$'s CARA coefficient. Thus, we assume that dealers and uninformed clients are risk-averse, while hedge funds are risk-neutral. With this assumption, we aim to capture that dealers and uninformed clients such as mutual funds and index funds may be subject to regulatory capital requirements, collateral requirements, as well as sudden withdrawals, which may lead to a costly liquidation of their inventory, while hedge funds are less subject to such costs.

The dynamic nature of trading in our model stems from the fact that agents make their trading decisions at $t = 1$ by anticipating the equilibrium trading decisions at $t = 2$. Because all agents have the same information set at $t = 1$, they choose their post-trade holdings, $D_1^i$ for $i \in \{D, UC, 1, 2, ..., N\}$, to maximize (5.2) and (5.3) under symmetric information. While small agents, dealers and uninformed clients, do their optimization by taking the price, $P_1$, as given, hedge funds internalize their price impact. That is, hedge fund $i$ chooses $D_1^i$ to maximize (5.3) by taking a pricing function as given: $P_1 = \hat{P}_1^0 + \hat{P}_1^1\left(D_1^{-i} + D_1^i\right)$ for constants $\hat{P}_1^0$ and $\hat{P}_1^1$ and given the sum, $D_1^{-i} \equiv \sum\limits_{j\in\{1,2,...,N\}\setminus\{i\}} D_1^j$, of demands of all other hedge funds. In turn, the market-clearing condition pins down $\hat{P}_1^0$ and $\hat{P}_1^1$.

As the economy moves to the second trading stage, which is at $t = 2$, the per-capita post-trade holdings at $t = 1$, $D_1^{HF}$, $D_1^D$, and $D_1^{UC}$, become the economy's state variables. That is, as is standard in dynamic models, trading strategies at $t = 2$ are, in principle, functions of those state variables inherited from $t = 1$. In addition to those three variables, hedge funds also have the realization of $\tilde{\theta}$ as their fourth state variable. Hence, this informational advantage of hedge funds is an additional trading motive for them that was not present in the previous round.

Similar to the first trading round at $t = 1$, we guess (and later verify) that the market-clearing price at $t = 2$ takes the form $P_2 = \hat{P}_2^0 + \hat{P}_2^1\left(D_2^{-i} + D_2^i\right)$, where $\hat{P}_2^1$ is constant and $\hat{P}_2^0$ is, in principle, a function of hedge funds' private information $\tilde{\theta}$ and the noisy demand $\tilde{d}$. Naturally, the optimally chosen $D_2^{-i}$ and $D_2^i$ are also functions of $\tilde{\theta}$ and $\tilde{d}$ because hedge funds already observe $\tilde{\theta}$, and then, the equilibrium price reveals $\tilde{d}$ to them as in Grossman and Stiglitz (1980). In turn, by observing $P_2$, dealers and uninformed clients infer a noisy signal about $\tilde{\theta}$. Again, we guess (and later verify) that hedge funds' demand is linear in the realization of random variables $\tilde{\theta}$ and $\tilde{d}$, which implies that dealers and uninformed clients use Bayesian updating in forming their demand by taking as given the following linear price function,

$$\tilde{P}_2 = \beta_c + \beta_\theta \tilde{\theta} + \beta_d \tilde{d}, \tag{5.4}$$

for information extraction purposes.

Taking stock, at $t = 2$, each hedge fund $i \in \{1, 2, ..., N\}$ chooses its post-trade holding $D_2^i$ by taking as given the pricing function $P_2 = \hat{P}_2^0 + \hat{P}_2^1 \left(D_2^{-i} + D_2^i\right)$, and so, by internalizing its impact on the market price. Dealers and uninformed clients choose their post-trade holdings, $D_2^i$ for $i \in \{D, UC\}$, by taking as given the market price, $P_2 = \tilde{P}_2$, and by learning and incorporating the implication of the market price for the true value of the asset via Equation (5.4). Then, the implied market-clearing condition, together with the rational expectations condition ($P_2 = \tilde{P}_2$), pins down $\hat{P}_2^0$, $\hat{P}_2^1$, $\beta_c$, $\beta_\theta$, and $\beta_d$.

## *5.3 Equilibrium Characterisation and Properties*

We characterize the equilibrium in two steps by using a backward induction. At the first step, we pin down the equilibrium objects at $t = 2$ for an arbitrary collection of "endowments," $D_1^i$ for $i \in \{HF, D, UC\}$. In the second step, we determine the equilibrium objects at $t = 1$, including $D_1^i$ for $i \in \{HF, D, UC\}$, by allowing agents to anticipate the equilibrium strategies that will prevail at $t = 2$.

Employing the "improper prior trick" (i.e., taking limit as $\sigma_\theta \to \infty$), we summarize the agents' trading behavior on the equilibrium path in Proposition 1 (with the proof presented in Section A.5.1 of the Appendix).[37]

**Proposition 1** *Assume $N > 2\bar{\gamma}\sigma_\varepsilon\sigma_d$, where $\bar{\gamma} = \left(\frac{1}{\gamma_D} + \frac{1}{\gamma_{UC}}\right)^{-1}$ is the harmonic sum of the dealers' and the uninformed clients' risk aversion parameters. Let $\Phi = \frac{N - \sqrt{N^2 - 4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2}}{2\bar{\gamma}\sigma_d^2}$. Then, on the equilibrium path, the agents' signed trade volumes, $q_t^{HF}$, $q_t^D$, and $q_t^{UC}$, and the market-clearing prices, $P_t$, are as follows. As $\sigma_\theta \to \infty$,*

$$\sigma_\theta^2 \left(q_1^{HF} - \frac{z}{N}\right) \to -\frac{N+1}{N^2}\frac{\Phi z}{\bar{\gamma}}, \quad \sigma_\theta^2 \left(q_1^D + z\right) \to \frac{\Phi z}{\gamma_D}\left(1 + \frac{1}{N}\right), \quad \sigma_\theta^2 q_1^{UC} \to \frac{\Phi z}{\gamma_{UC}}\left(1 + \frac{1}{N}\right), \tag{5.5}$$

*and*

$$P_1 \to \theta - \Phi z \tag{5.6}$$

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[37] The use of improper prior is not necessary for our analysis but it simplifies the exposition by reducing the number of exogenous parameters by two.

*at* $t = 1$ *and*

$$\sigma_\theta^2 \left( q_2^{HF} + \frac{\tilde{d}}{N} \right) \to \frac{\Phi z}{N\bar{\gamma}}, \ \ \sigma_\theta^2 q_2^D \to -\frac{\Phi z}{\gamma_D}, \ \ \sigma_\theta^2 q_2^{UC} \to -\frac{\Phi z}{\gamma_{UC}}, \tag{5.7}$$

*and*

$$P_2 \to \tilde{\theta} + \Phi \tilde{d} \tag{5.8}$$

*at* $t = 2$*. Informational inefficiency, measured by the conditional variance of the price signal at* $t = 2$*, is*

$$\mathbb{V}\left[\tilde{\theta} + \Phi\tilde{d} \,|\, \tilde{\theta}\right] = \Phi^2 \sigma_d^2. \tag{5.9}$$

As can be seen from Proposition 1, the endogenous object $\Phi$ is a main determinant for all equilibrium outcomes. Equation (5.8) implies that $\Phi$ is the coefficient of the noisy demand inside the second period's trading price. That is, it reflects the informational inefficiency of the market price. As agents make their trading decisions at $t = 1$ by anticipating the outcomes of $t = 2$, $\Phi$ becomes the risk premium inside the first period's trading price as revealed by (5.6). That is, it reflects the inverse of the market's effective risk bearing capacity. As an intermediate step for further interpreting the results in Proposition 1, we next derive in Lemma 1 comparative statics for this endogenous object $\Phi$.

**Lemma 1** *Assume* $N > 2\bar{\gamma}\sigma_\varepsilon\sigma_d$*. Let* $\Phi = \frac{N - \sqrt{N^2 - 4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2}}{2\bar{\gamma}\sigma_d^2}$*. Then,*

$$(i)\frac{\partial \Phi}{\partial N} < 0, \ (ii)\ \frac{\partial \Phi}{\partial \bar{\gamma}} > 0, \ (iii)\ \frac{\partial \frac{\Phi}{\bar{\gamma}}}{\partial \bar{\gamma}} > 0, \ (iv)\ \frac{\partial \Phi}{\partial \sigma_\varepsilon} > 0, \ \textit{and}\ (v)\ \frac{\partial \Phi}{\partial \sigma_d} > 0. \tag{5.10}$$

Hence, Lemma 1 implies that informational efficiency at $t = 2$ and the market's effective risk bearing capacity at $t = 1$ increase with $N$ and decrease with $\bar{\gamma}$, $\sigma_\varepsilon$, and $\sigma_d$.

The first type of implications of Proposition 1 we discuss are on liquidity provision in the secondary market following a bond issuance. Dealers arrive at the market at $t = 1$ with a large inventory they have obtained in the primary issuance. Thus, dealers are natural sellers and hedge funds and uninformed clients are natural buyers at $t = 1$. In the first-best allocation, which would obtain in a market with perfectly competitive hedge funds, each dealer sells $z$ shares, each hedge fund buys $z/N$, and uninformed clients do not trade. That is, this full liquidity provision by hedge funds would maximize all gains from trade as hedge funds are risk neutral, while dealers and uninformed clients are risk averse. Equations in (5.5) show the deviations from this first-best.

Looking at (5.5) and using Lemma 1, one can see that the hedge fund's liquidity provision at $t = 1$ increases with $N$ and decreases with $\sigma_\varepsilon$, $\sigma_d$, and $\bar{\gamma}$.[38] The main economic channel behind

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[38] That is, the absolute value of the deviation from perfect liquidity provision, $\left| -\frac{N+1}{N^2}\frac{\Phi z}{\bar{\gamma}} \right|$, decreases with $N$ and increases with $\sigma_\varepsilon$, $\sigma_d$, and $\bar{\gamma}$.

these comparative statics is hedge funds' price-impact avoidance incentive, which depresses the market's aggregate effective risk bearing capacity at $t = 1$. If there is a smaller number of hedge funds, if the small players are more risk averse, or if there is more uncertainty regarding the upcoming macro announcement (either because of information asymmetry or noise trading risk), each hedge fund has a larger impact on the price, both when they trade with payoff-relevant information at $t = 2$ and when the previous trading round takes place at $t = 1$ in anticipation of the outcomes of $t = 2$. Therefore, hedge funds' strategic price-impact avoidance at $t = 1$ makes them deviate from the perfect liquidity provision allocations accordingly. One can also see from (5.5) that uninformed clients partially compensates the lost liquidity provision from the hedge funds, as $\sigma_\theta^2 q_1^{UC}$ is an increasing function of $\Phi$, and so, it decreases with $N$ and increases with $\sigma_\varepsilon$, $\sigma_d$, and $\bar{\gamma}$—exactly in the opposite direction of the hedge funds' liquidity provision.

Next, we discuss the implications of Proposition 1 for the price drift before macro announcements. Our normalized benchmark price is $P_0 = \theta$, which we interpret as the asset's consensus fundamental value before any issuance. Indeed, when the supply is 0, the quantity of risk is 0 as well, implying no risk premium or no compensation for inventory costs. Then, the implied drift is

$$P_1 - P_0 = -\Phi z = -\frac{N - \sqrt{N^2 - 4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2}}{2\bar{\gamma}\sigma_d^2} z. \tag{5.11}$$

The following corollary highlights important properties of this price drift.

**Corollary 1** *Assume* $N > 2\bar{\gamma}\sigma_\varepsilon\sigma_d$. *Then,* $(i)\, P_1 - P_0 < 0$, $(ii)\, \frac{\partial|P_1-P_0|}{\partial z} > 0$, $(iii)\, \frac{\partial|P_1-P_0|}{\partial\bar{\gamma}} > 0$, $(iv)\, \frac{\partial|P_1-P_0|}{\partial\sigma_\varepsilon} > 0$, $(v)\, \frac{\partial|P_1-P_0|}{\partial\sigma_d} > 0$, *and* $(vi)\, \frac{\partial|P_1-P_0|}{\partial N} < 0$.

In terms of the terminology used in our empirical analysis, the Corollary implies that (i) bond yields rise after issuance, (ii) the rise is larger after a large issuance, (iii) the rise is larger if small players are more risk averse, (iv) the rise is larger if the announcement-related information asymmetry is larger, (iv) the rise is larger if the noise trading risk is larger, and (v) the rise is smaller if the competition among hedge funds is more severe. There results mainly follow from the fact that hedge funds' price impact, $\hat{P}_1^1$, at $t = 1$ given by (A.21) and the risk premium inside the $t = 1$ price (5.6) both increase with $\Phi$. This situation leads to the general-equilibrium observation that as hedge funds refrain from liquidity provision at $t = 1$, the price drift gets stronger.

**Interpreting Our Empirical Findings** We view the post-issuance, pre-news period as an event window that leads to an increase in $\sigma_\varepsilon$ in our model.[39] That is, under the heightened uncertainty about future interest rates, dealers and uninformed clients face larger information asymmetry because this is a period when speculative traders have stronger incentive to acquire information. With only one parameter changing, our model is quite parsimonious because it offers an explanation for two sets of empirical regularities, one set of aggregate results on yield and another set of more granular results on agent-specific trading behavior.[40] Hence, the set of our empirical findings is more than sufficiently rich to discipline our theory.

First, considering Table 4, the term, $\Phi$, inside the price drift (5.11) must be larger during a pre-news window compared to post-issuance times without macro announcements. Lemma 1 and Corollary 1 imply that this holds if information asymmetry is heightened (i.e., $\sigma_\varepsilon$ is larger) during pre-news windows. As explained above, this follows from hedge funds' price-impact avoidance incentive, which depresses the market's aggregate effective risk bearing capacity at $t = 1$.

Second, considering Table 11, the hedge fund's share in liquidity provision must be lower during a pre-news window compared to times outside pre-news windows, which means the object $\left|-\frac{N+1}{N^2}\frac{\Phi z}{\bar{\gamma}}\right|$ must be higher during a pre-news window according to (5.5). Again, Lemma 1 implies that this holds if $\sigma_\varepsilon$ is larger during pre-news windows. The intuition is as follows. As our model highlights, if dealers and uninformed clients face worse information asymmetry problem, hedge funds face a larger marginal cost of liquidity provision because of their increased price impact per share held. Hence, hedge funds refrain from liquidity provision so as to mitigate their price impact.

# 6 Concluding Remarks

**Summary** A rapidly expanding literature in macroeconomics has studied the role of monetary policy announcements and government bond supply in affecting long-term interest rates. In this paper we have documented a sizeable pre-news yield drift ahead of macroeconomic announcements such as monetary policy announcements or labour market data releases. This pre-news drift concentrates in periods during which these announcements are preceded by government bond issuance. We argue that this pre-news drift is linked to the limited risk-bearing capacity of primary dealers, which becomes more pronounced when issuance is closely followed

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[39] One could also argue that the noise trading risk, $\sigma_d$, must increase at the same time. Assuming an increase in $\sigma_d$ would only reinforce the current implications of our model because $\sigma_\varepsilon$ and $\sigma_d$ affect our main endogenous variable $\Phi$ qualitatively the same way.

[40] Most of the literature relies only on aggregate datasets on yields or prices and does not analyze transaction-level data. Thus, their analyses lack the dimension of who provides liquidity and when.

by an informationally sensitive period that could move bond yields. A main message of our paper is that macroeconomic announcements and bond issuance do not operate in a vacuum, and the interaction of these two factors can generate a sizeable impact on interest rates as well as on bond market liquidity (even before the announcement takes place).

**Policy Implications** Our paper has relevant implications for both monetary and fiscal policymakers. First, we have shown that debt issuance in pre-news windows ahead of monetary policy announcements amplifies monetary policy surprises, which could be interpreted as an additional constraint on successful monetary policy in the sense of King (2000). Second, debt issuance ahead of macroeconomic announcements is associated with pre-news drift, which could affect the signal that monetary policymakers should draw from market interest rates and might require them to alter the policy decisions as a result. An interesting normative question for future research is to determine the socially optimal way of timing bond issuances around macroeconomic announcements.

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# Tables

Table 1: Summary Statistics

(a) Primary Issuances

| Time period | (1)<br>Mean | (2)<br>Std. Dev. | (3)<br>25th | (4)<br>50th | (5)<br>75th | (6)<br>N |
|---|---|---|---|---|---|---|
| Panel A.1: Amount (£) Issued (in millions) | | | | | | |
| 1997m5-2021m7 | 2385.26 | 1389.28 | 1200.00 | 2474.60 | 3250.00 | 1075 |
| 1997m5-2007m12 | 1566.80 | 1031.70 | 450.00 | 2000.00 | 2500.00 | 211 |
| 2008m1-2021m7 | 2585.13 | 1392.56 | 1392.30 | 2500.00 | 3500.00 | 864 |
| Panel A.2: Years to Maturity | | | | | | |
| 1997m5-2021m7 | 19.27 | 13.41 | 8.53 | 15.55 | 29.89 | 1075 |
| 1997m5-2007m12 | 20.59 | 12.54 | 10.52 | 19.90 | 30.52 | 211 |
| 2008m1-2021m7 | 18.94 | 13.60 | 7.61 | 14.24 | 29.80 | 864 |

(b) Daily Changes in Nominal Bond Yields

| Maturity | (1)<br>Mean | (2)<br>Std. Dev. | (3)<br>25th | (4)<br>50th | (5)<br>75th | (6)<br>N |
|---|---|---|---|---|---|---|
| 5 years | -0.111 | 4.637 | -2.854 | -0.151 | 2.520 | 6128 |
| 10 years | -0.109 | 4.842 | -3.041 | -0.135 | 2.68 | 6128 |
| 15 years | -0.108 | 4.611 | -2.898 | -0.165 | 2.572 | 6128 |
| 20 years | -0.107 | 4.490 | -2.713 | -0.106 | 2.439 | 6128 |

Note: Panel A reports summary statistics of UK government bond issuances. The statistics are computed for the whole sample (1997m5-2021m7) as well as for two subperiods. Panel B reports summary statistics of daily nominal yield changes (expressed in basis points) for four different maturities, using the whole sample.

## Table 2: Yield Changes in Pre-news Windows

| | $5Y$-yield<br>(1) | $10Y$-yield<br>(2) | $15Y$-yield<br>(3) | $20Y$-yield<br>(4) |
|---|---|---|---|---|
| (A) Yield Changes before MPC Meetings | | | | |
| Pre-News window | 0.28 | 0.45** | 0.50** | 0.51*** |
| | (1.42) | (2.20) | (2.51) | (2.58) |
| Constant | -0.14** | -0.15** | -0.15** | -0.15** |
| | (-2.18) | (-2.30) | (-2.46) | (-2.52) |
| $N$ | 6128 | 6128 | 6128 | 6128 |
| (B) Yield Changes before Labour Market Data Release | | | | |
| Pre-News window | 0.50** | 0.48** | 0.43** | 0.39** |
| | (2.46) | (2.35) | (2.23) | (2.06) |
| Constant | -0.15** | -0.13** | -0.12* | -0.12* |
| | (-2.31) | (-1.99) | (-1.94) | (-1.90) |
| $N$ | 5896 | 5896 | 5896 | 5896 |

Note: Panel A (B) of this table regresses daily changes in 5-year, 10-year, 15-year and 20-year nominal gilt yields on an indicator variable that takes value one for days that are either one or two days before scheduled monetary policy announcements (labour market data releases) and zero otherwise. The estimation period in Panel A covers 1997m5-2021m7 and includes 270 MPC announcement windows. The estimation period in Panel B covers 1998m4-2021m7 and includes 280 pre-news windows corresponding to scheduled labour market data releases. T-statistics in parentheses are based on robust standard errors. Asterisks denote significance levels (* p<0.1, ** p<0.05, *** p<0.01).

## Table 3: Yield Changes in Pre-news Windows

| | $5Y$-yield<br>(1) | $10Y$-yield<br>(2) | $15Y$-yield<br>(3) | $20Y$-yield<br>(4) |
|---|---|---|---|---|
| (A) Yield Changes before MPC Meetings | | | | |
| Pre-News window | 0.09 | 0.34 | 0.38* | 0.37* |
| | (0.43) | (1.51) | (1.77) | (1.77) |
| Monday | -0.29** | -0.29** | -0.24** | -0.20* |
| | (-2.29) | (-2.32) | (-2.04) | (-1.73) |
| Tuesday (relative to Monday) | 0.36** | 0.21 | 0.15 | 0.12 |
| | (1.98) | (1.10) | (0.83) | (0.67) |
| Wednesday (relative to Monday) | 0.34* | 0.33* | 0.29 | 0.27 |
| | (1.75) | (1.68) | (1.56) | (1.49) |
| Thursday (relative to Monday) | 0.18 | 0.28 | 0.21 | 0.14 |
| | (0.97) | (1.46) | (1.18) | (0.80) |
| Friday (relative to Monday) | -0.06 | -0.07 | -0.15 | -0.22 |
| | (-0.31) | (-0.35) | (-0.80) | (-1.19) |
| $N$ | 6128 | 6128 | 6128 | 6128 |
| (B) Yield Changes before Labour Market Data Release | | | | |
| Pre-News window | 0.60*** | 0.63*** | 0.55*** | 0.48** |
| | (2.79) | (2.93) | (2.67) | (2.33) |
| Monday | -0.45*** | -0.46*** | -0.39*** | -0.33** |
| | (-3.35) | (-3.35) | (-2.94) | (-2.51) |
| Tuesday (relative to Monday) | 0.46** | 0.37** | 0.33* | 0.30* |
| | (2.56) | (2.03) | (1.89) | (1.77) |
| Wednesday (relative to Monday) | 0.49*** | 0.56*** | 0.52*** | 0.48*** |
| | (2.59) | (2.86) | (2.77) | (2.64) |
| Thursday (relative to Monday) | 0.38* | 0.48** | 0.38** | 0.28 |
| | (1.93) | (2.37) | (1.98) | (1.51) |
| Friday (relative to Monday) | 0.13 | 0.13 | 0.02 | -0.08 |
| | (0.69) | (0.65) | (0.08) | (-0.43) |
| $N$ | 5896 | 5896 | 5896 | 5896 |

Note: Panel A (B) of this table regresses daily changes in 5-year, 10-year, 15-year and 20-year nominal gilt yields on an indicator variable that takes value one for days that are either one or two days before scheduled monetary policy announcements (labour market data releases) and zero otherwise. The estimation period in Panel A covers 1997m5-2021m7 and includes 270 MPC announcement windows. The estimation period in Panel B covers 1998m4-2021m7 and includes 280 pre-news windows corresponding to scheduled labour market data releases. All regressions include a weekday fixed effect. T-statistics in parentheses are based on robust standard errors. Asterisks denote significance levels (* p<0.1, ** p<0.05, *** p<0.01).

Table 4: Yield Changes in Pre-news Windows: the Role of Bond Issuance

| | $5Y$-yield<br>(1) | $10Y$-yield<br>(2) | $15Y$-yield<br>(3) | $20Y$-yield<br>(4) |
|---|---|---|---|---|
| (A) Yield Changes before MPC Meetings and after Bond Issuance | | | | |
| Pre-News window # Issuance | 0.20 | 0.54* | 0.65** | 0.68** |
| | (0.73) | (1.91) | (2.39) | (2.53) |
| Pre-News window # No issuance | 0.46 | 0.41 | 0.37 | 0.36 |
| | (1.63) | (1.44) | (1.36) | (1.29) |
| No Pre-News # Issuance | 0.23 | 0.20 | 0.19 | 0.17 |
| | (1.32) | (1.10) | (1.07) | (1.01) |
| Constant | -0.17** | -0.18** | -0.18*** | -0.18*** |
| | (-2.50) | (-2.55) | (-2.69) | (-2.72) |
| $N$ | 6128 | 6128 | 6128 | 6128 |
| (B) Yield Changes before Labour Market Data Release and after Bond Issuance | | | | |
| Pre-News window # Issuance | 1.06** | 1.17** | 1.09** | 1.12** |
| | (1.96) | (2.26) | (2.30) | (2.45) |
| Pre-News window # No issuance | 0.46** | 0.42* | 0.38* | 0.33 |
| | (2.07) | (1.90) | (1.82) | (1.61) |
| No Pre-News # Issuance | 0.21 | 0.25 | 0.28 | 0.28* |
| | (1.28) | (1.39) | (1.64) | (1.72) |
| Constant | -0.18*** | -0.18** | -0.17** | -0.17** |
| | (-2.61) | (-2.40) | (-2.47) | (-2.47) |
| $N$ | 5896 | 5896 | 5896 | 5896 |

Note: this table regresses daily changes in 5-year, 10-year, 15-year and 20-year nominal gilt yields on an indicator variables on indicator variables capturing (i) pre-news windows with new bond issuance, (ii) pre-news windows without new bond issuance and (iii) all trading days with issuance and without announcements. The estimation period in Panel A covers 1997m5-2021m7 and includes 270 MPC announcement windows. The estimation period in Panel B covers 1998m4-2021m7 and includes 280 pre-news windows corresponding to scheduled labour market data releases. T-statistics in parentheses are based on robust standard errors. Asterisks denote significance levels (* p<0.1, ** p<0.05, *** p<0.01).

## Table 5: Yield Changes in Pre-news Windows: the Role of Bond Issuance

|  | $5Y$-yield<br>(1) | $10Y$-yield<br>(2) | $15Y$-yield<br>(3) | $20Y$-yield<br>(4) |
|---|---|---|---|---|
| (A) Yield Changes before MPC Meetings and after Bond Issuance |  |  |  |  |
| Pre-News window # Issuance | -0.02 | 0.41 | 0.52* | 0.53* |
|  | (-0.05) | (1.35) | (1.79) | (1.86) |
| Pre-News window # No issuance | 0.27 | 0.29 | 0.25 | 0.22 |
|  | (0.90) | (0.97) | (0.87) | (0.76) |
| No Pre-News # Issuance | 0.09 | 0.08 | 0.07 | 0.06 |
|  | (0.47) | (0.38) | (0.37) | (0.31) |
| Weekday fixed effects | YES | YES | YES | YES |
| $N$ | 6128 | 6128 | 6128 | 6128 |
| (B) Yield Changes before Labour Market Data Release and after Bond Issuance |  |  |  |  |
| Pre-News window # Issuance | 1.09** | 1.27** | 1.15** | 1.14** |
|  | (2.00) | (2.42) | (2.39) | (2.46) |
| Pre-News window # No issuance | 0.53** | 0.55** | 0.48** | 0.40* |
|  | (2.32) | (2.39) | (2.17) | (1.81) |
| No Pre-News # Issuance | 0.06 | 0.11 | 0.15 | 0.15 |
|  | (0.35) | (0.60) | (0.81) | (0.84) |
| Weekday fixed effects | YES | YES | YES | YES |
| $N$ | 5896 | 5896 | 5896 | 5896 |

Note: this table regresses daily changes in 5-year, 10-year, 15-year and 20-year nominal gilt yields on an indicator variables on indicator variables capturing (i) pre-news windows with new bond issuance, (ii) pre-news windows without new bond issuance and (iii) all trading days with issuance and without announcements. The estimation period in Panel A covers 1997m5-2021m7 and includes 270 MPC announcement windows. The estimation period in Panel B covers 1998m4-2021m7 and includes 280 pre-news windows corresponding to scheduled labour market data releases. All regressions include a weekday fixed effect. T-statistics in parentheses are based on robust standard errors. Asterisks denote significance levels (* p<0.1, ** p<0.05, *** p<0.01).

Table 6: 10Y Yield Changes before MPC Meetings: Term Premium vs Expectations

| | Yield (1) | Fitted yield (2) | Term pr. (3) | Expect. (4) |
|---|---|---|---|---|
| (A) 10Y Yield Changes before MPC Meetings | | | | |
| Pre-News window | 0.33 | 0.32 | 0.42*** | -0.10 |
| | (1.45) | (1.42) | (2.77) | (-0.70) |
| Weekday fixed effects | YES | YES | YES | YES |
| *N* | 5983 | 5983 | 5983 | 5983 |
| (B) 10Y Yield Changes before Labour Market Data Release | | | | |
| Pre-News window | 0.66*** | 0.67*** | 0.31** | 0.35** |
| | (3.01) | (3.03) | (2.21) | (2.38) |
| Weekday fixed effects | YES | YES | YES | YES |
| *N* | 5750 | 5750 | 5750 | 5750 |
| (B) 10Y Yield Changes before MPC Meetings and during Bond Issuance | | | | |
| Pre-News window # Issuance | 0.40 | 0.40 | 0.67*** | -0.27 |
| | (1.31) | (1.29) | (3.21) | (-1.42) |
| Pre-News window # No issuance | 0.31 | 0.31 | 0.17 | 0.14 |
| | (1.02) | (1.01) | (0.81) | (0.68) |
| No News # Issuance | 0.14 | 0.14 | 0.05 | 0.09 |
| | (0.70) | (0.69) | (0.34) | (0.70) |
| Weekday fixed effects | YES | YES | YES | YES |
| *N* | 5983 | 5983 | 5983 | 5983 |
| (C) 10Y Yield Changes before Labour Market Data Release and during Bond Issuance | | | | |
| Pre-News window # Issuance | 1.34** | 1.35** | 0.76* | 0.58 |
| | (2.50) | (2.50) | (1.70) | (1.35) |
| Pre-News window # No issuance | 0.59** | 0.59** | 0.27* | 0.32** |
| | (2.48) | (2.50) | (1.86) | (2.07) |
| No News # Issuance | 0.16 | 0.16 | 0.14 | 0.02 |
| | (0.82) | (0.81) | (0.99) | (0.17) |
| Weekday fixed effects | YES | YES | YES | YES |
| *N* | 5750 | 5750 | 5750 | 5750 |

Note: this table regresses daily changes in the 10-year nominal gilt yield and fitted yields, term premia and expectation components from a dynamic no-arbitrage affine term structure model, estimated by linear regression techniques (Adrian, Crump, and Moench, 2013; Malik and Meldrum, 2016) on a sample covering 1991m1-2020m12. The regressor in Panel A (B) is an indicator variable that takes value one for days that are either one or two days before scheduled monetary policy announcements (labour market data releases) and zero otherwise. The regressors in Panels C-D are indicator variables capturing (i) pre-news windows with new bond issuance, (ii) pre-news windows without new bond issuance and (iii) all trading days with issuance and without announcements. The estimation period in Panel A (B) covers 1997m5-2020m12 (1998m4-2020m12). All regressions include a weekday fixed effect. T-statistics in parentheses are based on robust standard errors. Asterisks denote significance levels (* $p<0.1$, ** $p<0.05$, *** $p<0.01$).

Table 7: The Relationship between Pre-MPC Bond Issuance and the Volatility of Monetary Policy Surprises

|  | (1)<br>3-month futures | (2)<br>1-year yield | (3)<br>2-year yield | (4)<br>5-year yield | (5)<br>10-year yield |
|---|---|---|---|---|---|
| (A) Sample: 1997-2021 |  |  |  |  |  |
| Pre-MPC Issuance | 0.57 | 0.83** | 0.59 | 0.67* | 0.59* |
|  | (1.18) | (2.16) | (1.44) | (1.71) | (1.67) |
| $N$ | 268 | 268 | 268 | 268 | 268 |
| $R^2$ | 0.161 | 0.180 | 0.156 | 0.186 | 0.232 |
| (A) Sample: 2006-2021 |  |  |  |  |  |
| Pre-MPC Issuance | 1.38*** | 1.39*** | 1.15** | 1.05** | 0.85* |
|  | (2.66) | (3.16) | (2.48) | (2.11) | (1.89) |
| $N$ | 168 | 168 | 168 | 168 | 168 |
| $R^2$ | 0.168 | 0.172 | 0.172 | 0.180 | 0.214 |

Note: This table presents the estimation results for regression 3.3. The regressands are taken from Braun, Miranda-Agrippino, and Saha (2022) and measure price changes in a 30-minute window around the interest rate announcement of the Monetary Policy Committee of the Bank of England for five different assets: the 3-month futures contracts (3m), and gilt yields with one year (1Y), two years (2Y) , five years (5Y) and ten years (10Y) maturity. T-statistics in parentheses are based on robust standard errors. Asterisks denote significance levels (* p<0.1, ** p<0.05, *** p<0.01).

## Table 8: Pre-news Drift and Bond Issuance: High vs Low Price Dispersion Months

| | 5Y<br>(1) | 10Y<br>(2) | 15Y<br>(3) | 20Y<br>(4) | 5Y<br>(5) | 10Y<br>(6) | 15Y<br>(7) | 20Y<br>(8) |
|---|---|---|---|---|---|---|---|---|
| (A) Pre-MPC Drift & Lower Dispersion | | | | | (B) Pre-MPC Drift & Higher Dispersion | | | |
| Pre-News # Issue | 0.10 | 0.11 | 0.10 | 0.15 | 1.30** | 1.45* | 1.42* | 1.34* |
| | (0.20) | (0.19) | (0.18) | (0.30) | (2.30) | (1.92) | (1.82) | (1.73) |
| Pre-News # No issue | 0.01 | 0.04 | 0.11 | 0.28 | -0.31 | -0.52 | -0.60 | -0.92 |
| | (0.02) | (0.05) | (0.15) | (0.38) | (-0.35) | (-0.49) | (-0.55) | (-0.84) |
| No News # Issue | -0.16 | -0.35 | -0.35 | -0.31 | -0.34 | -0.38 | -0.27 | -0.17 |
| | (-0.52) | (-0.98) | (-1.02) | (-0.95) | (-1.02) | (-0.99) | (-0.73) | (-0.47) |
| Constant | -0.07 | -0.09 | -0.09 | -0.09 | -0.07 | -0.11 | -0.16 | -0.19 |
| | (-0.53) | (-0.64) | (-0.69) | (-0.74 | (-0.45) | (-0.58) | (-0.91) | (-1.13) |
| *N* | 1144 | 1144 | 1144 | 1144 | 1110 | 1110 | 1110 | 1110 |
| (C) Pre-MPC Drift & Lower Dispersion | | | | | (D) Pre-MPC Drift & Higher Dispersion | | | |
| Pre-News # Issue | -0.34 | -0.29 | -0.29 | -0.24 | 1.12* | 1.33* | 1.29 | 1.19 |
| | (-0.65) | (-0.48) | (-0.51) | (-0.43) | (1.87) | (1.67) | (1.57) | (1.45) |
| Pre-News # No issue | -0.42 | -0.36 | -0.28 | -0.11 | -0.50 | -0.66 | -0.75 | -1.07 |
| | (-0.55) | (-0.42) | (-0.37) | (-0.14) | (-0.56) | (-0.61) | (-0.68) | (-0.96) |
| No News # Issue | -0.36 | -0.59 | -0.60 | -0.55 | -0.60 | -0.57 | -0.45 | -0.36 |
| | (-1.06) | (-1.49) | (-1.58) | (-1.55) | (-1.58) | (-1.27) | (-1.05) | (-0.86) |
| Weekday FE | YES | YES | YES | YES | YES | YES | YES | YES |
| *N* | 1145 | 1145 | 1145 | 1145 | 1110 | 1110 | 1110 | 1110 |

Note: This table regresses daily changes in yields on indicator variables capturing (i) pre-news windows (associated with monetary policy announcements) and new bond issuance, (ii) pre-news windows without new bond issuance and (iii) all trading days with issuance and without announcements. Panel A (Panel B) presents the results without weekday fixed effects using the subsample when the market has lower (higher) inter-dealer price dispersion (Eisfeldt, Herskovic, and Liu, 2023). Panels C-D present the results with weekday fixed effects. The classification of months with lower and higher inter-dealer price dispersion is based on the median value in the given year. The estimation period covers 2011m9-2021m7, which is dictated by the availability of the transaction-level data. T-statistics in parentheses are based on robust standard errors. Asterisks denote significance levels (* $p<0.1$, ** $p<0.05$, *** $p<0.01$).

## Table 9: Pre-news Drift and Bond Issuance: High vs Low Price Dispersion

| | 5Y<br>(1) | 10Y<br>(2) | 15Y<br>(3) | 20Y<br>(4) | 5Y<br>(5) | 10Y<br>(6) | 15Y<br>(7) | 20Y<br>(8) |
|---|---|---|---|---|---|---|---|---|
| (A) Pre-Labour Drift & Lower Dispersion | | | | | (B) Pre-Labour Drift & Higher Dispersion | | | |
| Pre-News # Issue | 0.33 | 0.54 | 0.76 | 0.87 | 2.04*** | 1.93** | 1.58* | 1.40* |
| | (0.49) | (0.68) | (0.95) | (1.08) | (2.65) | (2.13) | (1.84) | (1.70) |
| Pre-News # No issue | 0.41 | 0.49 | 0.56 | 0.56 | 0.43 | 0.57 | 0.52 | 0.45 |
| | (1.08) | (1.15) | (1.39) | (1.52) | (0.86) | (0.98) | (0.88) | (0.74) |
| No News # Issue | -0.26 | -0.46 | -0.46 | -0.40 | -0.32 | -0.30 | -0.17 | -0.08 |
| | (-0.88) | (-1.34) | (-1.38) | (-1.25) | (-1.01) | (-0.79) | (-0.48) | (-0.22) |
| Constant | -0.08 | -0.11 | -0.12 | -0.12 | -0.07 | -0.12 | -0.17 | -0.21 |
| | (-0.60) | (-0.71) | (-0.82) | (-0.89) | (-0.43) | (-0.63) | (-0.92) | (-1.16) |
| *N* | 1144 | 1144 | 1144 | 1144 | 1110 | 1110 | 1110 | 1110 |
| (C) Pre-Labour Drift & Lower Dispersion | | | | | (D) Pre-Labour Drift & Higher Dispersion | | | |
| Pre-News # Issue | 0.19 | 0.42 | 0.65 | 0.78 | 2.26*** | 2.30** | 1.89** | 1.65* |
| | (0.27) | (0.51) | (0.79) | (0.93) | (2.85) | (2.45) | (2.12) | (1.94) |
| Pre-News # No issue | 0.42 | 0.54 | 0.61 | 0.62 | 0.71 | 0.98 | 0.88 | 0.75 |
| | (1.06) | (1.21) | (1.45) | (1.60) | (1.35) | (1.60) | (1.42) | (1.18) |
| No News # Issue | -0.44 | -0.68* | -0.68* | -0.61* | -0.59 | -0.48 | -0.35 | -0.25 |
| | (-1.37) | (-1.80) | (-1.88) | (-1.77) | (-1.64) | (-1.12) | (-0.85) | (-0.64) |
| Weekday FE | YES | YES | YES | YES | YES | YES | YES | YES |
| *N* | 1145 | 1145 | 1145 | 1145 | 1110 | 1110 | 1110 | 1110 |

Note: This table regresses daily changes in yields on indicator variables capturing (i) pre-news windows (associated with labour market data releases) and new bond issuance, (ii) pre-news windows without new bond issuance and (iii) all trading days with issuance and without announcements. Panel A (Panel B) presents the results without weekday fixed effects using the subsample when the market has lower (higher) inter-dealer price dispersion (Eisfeldt, Herskovic, and Liu, 2023). Panels C-D present the results with weekday fixed effects. The classification of months with lower and higher inter-dealer price dispersion is based on the median value in the given year. The estimation period covers 2011m9-2021m7, which is dictated by the availability of the transaction-level data. T-statistics in parentheses are based on robust standard errors. Asterisks denote significance levels (* p<0.1, ** p<0.05, *** p<0.01).

Table 10: Evidence from Dealers' Inventory Imbalance

| | Before News | | No News |
|---|---|---|---|
| | MPC | Labour | |
| | (1) | (2) | (3) |
| (A) All Dealers | | | |
| DealerFlow | -0.011* | -0.021** | -0.004 |
| | (-1.89) | (-2.09) | (-1.46) |
| $N$ | 928 | 228 | 4467 |
| $R^2$ | 0.128 | 0.112 | 0.104 |
| (B) Dealers Participating in the Auction | | | |
| DealerFlow | -0.017** | -0.034*** | -0.006 |
| | (-2.49) | (-3.72) | (-1.54) |
| $N$ | 386 | 80 | 1746 |
| $R^2$ | 0.219 | 0.270 | 0.195 |
| (C) Dealers Not Participating in the Auction | | | |
| DealerFlow | -0.003 | 0.015 | -0.003 |
| | (-0.32) | (0.72) | (-0.96) |
| $N$ | 538 | 139 | 2709 |
| $R^2$ | 0.166 | 0.167 | 0.114 |

Note: This table presents the estimation results corresponding to regression 4.2. Columns (1)-(2) show, respectively, the estimates for pre-MPC windows and pre-news windows associated with labour market data releases. Column (3) shows the results for periods without impending news. T-statistics in parentheses are based on robust standard errors. Asterisks denote significance levels (* $p<0.1$, ** $p<0.05$, *** $p<0.01$).

Table 11: Composition of Orderflow after Primary Issuance Days *outside* and *during* pre-MPC Windows

| | (1) All Clients | (2) Comm. Banks | (3) Pens&Ins. | (4) For. Central Banks | (5) Others | (6) Hedge Funds | (7) Asset Mng. |
|---|---|---|---|---|---|---|---|
| Panel A: Monetary Policy Announcements | | | | | | | |
| On Issuance Day # 2 Days Before News | -10.00 | 24.41 | -6.86 | 52.46 | 23.58 | -73.53 | -30.06 |
| After Issuance Day # Day Before News | 175.54 | -33.61 | 5.81 | 147.20*** | 28.12 | -16.00 | 44.01 |
| On Issuance Day # Day Before News | 538.35** | 56.88 | 132.75 | 66.92 | 112.26** | 63.26 | 106.29 |
| After Issuance Day # Not Before News | 126.55*** | 5.30 | 4.71 | -2.46 | -4.08 | 76.76*** | 46.33** |
| On Issuance Day # Not Before News | 504.92*** | 17.20 | 71.07*** | 11.25 | -2.51 | 262.85*** | 145.06*** |
| Other Days | 84.55*** | -7.99** | 36.15*** | 68.30*** | 7.56*** | -52.77*** | 33.30*** |
| Panel B: Scheduled Labour Market Data Releases | | | | | | | |
| On Issuance Day # 2 Days Before News | - | - | - | - | - | - | - |
| After Issuance Day # Day Before News | - | - | - | - | - | - | - |
| On Issuance Day # Day Before News | 686.79*** | 25.67 | 259.76** | 8.36 | -35.43** | 249.27*** | 179.16** |
| After Issuance Day # Not Before News | 123.02*** | 2.71 | 2.47 | 14.39 | 1.31 | 51.58** | 50.57** |
| On Issuance Day # Not Before News | 491.11*** | 21.74** | 59.78*** | 20.19 | 6.92 | 244.77*** | 137.71*** |
| Other Days | 86.57*** | -8.37** | 36.51*** | 68.85*** | 7.41*** | -50.68*** | 32.85*** |
| *N* | 2105 | 2105 | 2105 | 2105 | 2105 | 2105 | 2105 |

Notes: the table presents the estimated mean orderflows (net purchases) of all clients (1) and different client sectors (2-7). The sample, covering the period from Aug 2011 to December 2019, includes 2105 trading days. We consider the following six types of trading days: issuances days that are two days before macroeconomic announcements; days that are one day after issuance days and one day before announcements; issuance days that are one day before announcements; issuance days without impending news; days after issuance days without impending news; and all other trading days. Panel A (B) shows the results for monetary policy announcements (labour market data releases). Asterisks denote significance levels (* $p<0.1$, ** $p<0.05$, *** $p<0.01$).

Table 12: Average Hedge Fund Returns (over 1-6 Day Horizons) Around Debt Issuance and Macroeconomic Announcements

| | Unweighted Performance | | | Volume-Weighted Performance | | |
|---|---|---|---|---|---|---|
| | $T = 1d$ | $T = 3d$ | $T = 6d$ | $T = 1d$ | $T = 3d$ | $T = 6d$ |
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Panel A: Monetary Policy Announcements | | | | | | |
| t-2 | 2.43 | -1.06 | -4.97 | 2.80** | -0.44 | -0.66 |
| t-1 | -1.16 | 0.63 | -0.31 | -0.97 | -1.12 | 1.24 |
| t [news] | 2.93* | 7.68*** | 13.22*** | 2.15* | 1.30 | 4.99* |
| Panel B: Scheduled Labour Market Data Releases | | | | | | |
| t-2 | - | - | - | - | - | - |
| t-1 | -0.29 | 0.66 | -2.03 | -1.37 | -2.99 | -3.80 |
| t [news] | 3.60** | 4.37 | 7.78* | 1.49 | 1.50 | 5.62* |
| $N$ | 2105 | 2105 | 2105 | 2105 | 2105 | 2105 |

Notes: the table presents average performance (measured in basis points) of hedge fund trades over 1-, 3- and 6-day horizons (based on the measure 4.3) in pre-MPC windows that coincide with primary issuance. The rows capture the days on which the trades are executed ($t$ denotes day of MPC meeting, $t-1$ denotes the day before MPC meetings and $t-2$ denotes the days that are two days before MPC meetings). The columns capture the horizon ($T$) over which the performance measure is computed. Columns 1-3 present unweighted average performance measures for the hedge funds sector and columns 4-6 present average measures that are weighted by pound values of the transactions. Panel A (B) shows the results for monetary policy announcements (labour market data releases). The asterisks indicate whether the returns are different from zero at 1%, 5% and 10% significance levels.

# A Online Appendix

“BOND SUPPLY, YIELD DRIFTS, AND LIQUIDITY PROVISION
BEFORE MACROECONOMIC ANNOUNCEMENTS”

29th October 2024

## *A.1 Summary Statistics in our Sample of Transaction-Level Data*

Using all available transactions from the client-dealer segment of the market, we group clients into six sectors following the classification of Czech, Huang, Lou, and Wang (2021): commercial banks, pension fund and insurance companies, foreign central banks, hedge funds, asset managers and other services.

Table A.1: Summary Statistics across Different Client Types

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | in £ millions | | N | % of All Clients | | |
| | Turnover | Orderflow | Trades | Turnover | Orderflow | Trades |
| All Clients | 7076.85 | 229.76 | 1011.12 | 100% | 100% | 100% |
| Comm.Banks | 484.54 | -6.61 | 88.77 | 6.8% | -2.9% | 8.8% |
| Pens&Ins. | 1160.40 | 61.78 | 178.58 | 16.4% | 26.9% | 17.7% |
| Foreign Central Banks | 820.49 | 81.19 | 27.84 | 11.6% | 35.3% | 2.8% |
| Other Services | 579.96 | 13.54 | 94.60 | 8.2% | 5.9% | 9.4% |
| Hedge Funds | 1693.97 | -9.55 | 99.11 | 23.9% | -4.2% | 9.8% |
| Asset Managers | 2337.48 | 89.41 | 522.22 | 33.0% | 38.9% | 51.6% |

Notes: the table provides summary statistics on daily turnover, orderflow (buy volume net sell volume) and number of transactions for all clients as well as the six client sectors. The sample covers the period from Aug 2011 to December 2019.

Table A.1 summarises the average daily activity of all clients as well the activity of the six sectors, using three different measures: turnover, orderflow (buy volume net of sell volume) and number of transactions. Average client turnover is about £7 billion on a trading day, with the majority of turnover generated by asset managers (£2.3 billion) and hedge funds (£1.7 billion) that together account for 56.9% of daily turnover in our sample. The client sector as a whole is a net buyer of gilts with a daily order flow of £230 million, that is mainly driven by purchases of asset managers (£89 million), foreign central banks (£81 million), pension funds and insurance companies (£62 million). There are about 1,011 daily transactions in the dealer-client segment of our sample with the majority generated by asset managers. It is interesting to note that foreign central banks trade very infrequently (2.8% of all client trades), but they generate more than a third of total client orderflow. This is consistent with foreign central banks trading in extremely large quantities.

## *A.2 Description of the Term Structure Model*

### *A.2.1 Excess Returns*

This section summarises the dynamic no-arbitrage affine term structure model, based on Adrian, Crump, and Moench (2013) and Malik and Meldrum (2016), that we use to decompose long-term bond yields into expectations of future short-term interest rates and the term premia that compensate investors for the risk associated with holding long-term bonds. The dynamics of a $K \times 1$ of state variables evolve according to a Gaussian vector autoregression (VAR(1)):

$$X_{t+1} = \mu + \Phi X_t + v_{t+1}, \tag{A.1}$$

where the shocks $v_{t+1} \sim N\left(0, \Sigma\right)$ are conditionally Gaussian, homoscedastic and independent across time. The price of a zero coupon bond with maturity $n$ at time $t$ is denoted by $P_t^{(n)}$. The assumption of no-arbitrage implies the existence of a pricing kernel $M_{t+1}$ such that:

$$P_t^{(n)} = \mathbb{E}\left[M_{t+1}, P_{t+1}^{(n-1)}\right]. \tag{A.2}$$

The pricing kernel is assumed to be exponentially affine:

$$M_{t+1} = \exp\left(-r_t - \frac{1}{2}\lambda_t'\lambda_t - \lambda_t'\Sigma^{-1/2}v_{t+1}\right), \tag{A.3}$$

where $r_t = \ln P_t^{(1)}$ is the continuously compounded risk-free interest rate. Further, market prices of risk are assumed to be affine in the state variables (Duffee, 2002):

$$\lambda_t = \Sigma^{-1/2}\left(\lambda_0 + \lambda_1 X_t\right). \tag{A.4}$$

The natural logarithm of excess one-period holding return of a bond maturing in $n$ periods is written as:

$$rx_{t+1}^{(n-1)} = \ln P_{t+1}^{(n-1)} - \ln P_t^{(n)} - r_t. \tag{A.5}$$

Combining A.2, A.3 and A.5 yields:

$$1 = \mathbb{E}_t\left[\exp\left(rx_{t+1}^{(n-1)} - \frac{1}{2}\lambda_t'\lambda_t - \lambda_t'\Sigma^{-1/2}v_{t+1}\right)\right]. \tag{A.6}$$

Assuming that $\left\{rx_{t+1}^{(n-1)}, v_{t+1}\right\}$ are jointly normally distributed, we can write:

$$\mathbb{E}_t\left[rx_{t+1}^{(n-1)}\right] = Cov_t\left[rx_{t+1}^{(n-1)}, v_{t+1}'\Sigma^{-1/2}\lambda_t\right] - \frac{1}{2}Var_t\left[rx_{t+1}^{(n-1)}\right]. \tag{A.7}$$

Denoting $\beta_t^{(n-1)} = Cov_t\left[rx_{t+1}^{(n-1)}, v_{t+1}'\right]\Sigma^{-1}$ and using A.4, A.7 can be written as:

$$\mathbb{E}_t\left[rx_{t+1}^{(n-1)}\right] = \beta_t^{(n-1)\prime}\left(\lambda_0 + \lambda_1 X_t\right) - \frac{1}{2}Var_t\left[rx_{t+1}^{(n-1)}\right].$$

One can then decompose unexpected excess returns into a components that is correlated with $v_{t+1}$ and another component that is conditionally orthogonal:

$$rx_{t+1}^{(n-1)} - \mathbb{E}_t\left[rx_{t+1}^{(n-1)}\right] = \beta_t^{(n-1)\prime} v_{t+1} + e_{t+1}^{(n-1)}, \tag{A.8}$$

where $e_{t+1}^{(n-1)} \sim N\left(0, \sigma^2\right)$. The return generating process for log excess holding period returns is then written as:

$$\begin{aligned} rx_{t+1}^{(n-1)} =& \beta_t^{(n-1)\prime}\left(\lambda_0 + \lambda_1 X_t\right) - \frac{1}{2}\beta_t^{(n-1)\prime}\Sigma\beta_t^{(n-1)} + \sigma^2 \\ &+ \beta_t^{(n-1)\prime} v_{t+1} + e_{t+1}^{(n-1)}. \end{aligned} \tag{A.9}$$

Stacking this system across maturities and time periods, excess returns can be rewritten as:

$$rx = \boldsymbol{\beta}'\left(\lambda_0 \iota_T' + \lambda_1 X_-\right) - \frac{1}{2}\left(B^\star vec\left(\Sigma\right) + \sigma^2 \iota_N\right)\iota_T' + \boldsymbol{\beta}' V + E, \tag{A.10}$$

where $rx$ is an $N \times T$ matrix of excess return, $\boldsymbol{\beta} = \left[\beta^{(1)}\ \beta^{(2)}\ \ldots \beta^{(N)}\right]$ is a $K \times N$ matrix of factor loadings, $\iota_T$ and $\iota_N$ are a $T \times 1$ and $N \times 1$ vectors of ones, $X_- = \left[X_0\ X_1\ \ldots X_{T-1}\right]$ is a $K \times T$ matrix of lagged pricing factors, $B^\star = \left[vec\left(\beta^{(1)}\beta^{(1)\prime}\right)\ldots vec\left(\beta^{(N)}\beta^{(N)\prime}\right)\right]$ is an $N \times K^2$ matrix, and $V$ and $E$ are matrices of $K \times T$ and $N \times T$ dimensions.

### *A.2.2 Estimation*

Given A.10, the three-step regression-based estimator for the parameters of the model is summarised as follows:

1. Estimate A.1 with ordinary least squares, and use the residuals, $\hat{V}$, to estimate the variance-covariance matrix of the state variables, $\hat{\Sigma} = \hat{V}\hat{V}'/T$.

2. Regress excess returns on a constant, lagged pricing factors and contemporaneous pricing factor innovations, $rx = a\iota_T' + \beta'\hat{V} + cX_- + E$, yielding estimates $\hat{a}$, $\hat{\beta}$ and $\hat{c}$ as well as $\hat{\sigma}^2 = tr\left(\hat{E}\hat{E}'\right)/NT$.

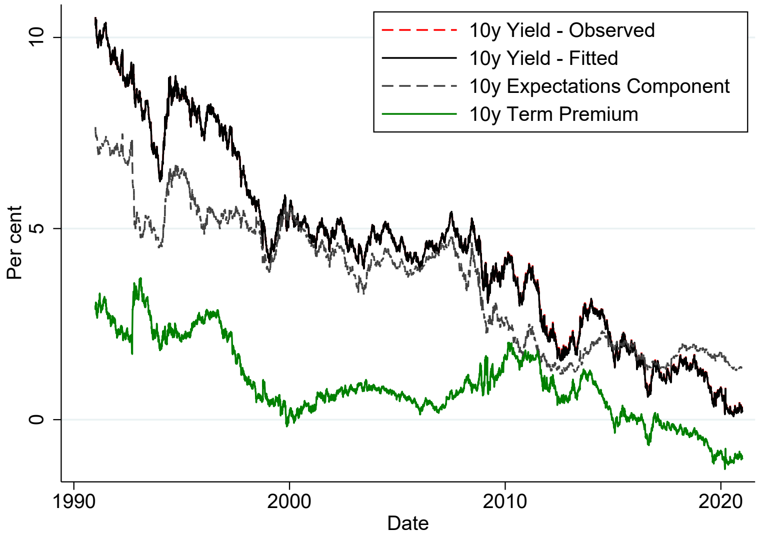
3. Using $\hat{B}^\star$ constructed from $\hat{\beta}$ and noting that $a = \beta'\lambda_0 - \frac{1}{2}\left(B^\star vec\left(\Sigma\right) + \sigma^2\iota_N\right)$ and $c = \beta'\lambda_1$,we can estimate the parameters of the price or risk using cross-sectional regressions:

$$\begin{aligned} \hat{\lambda}_0 &= \left(\hat{\beta}\hat{\beta}'\right)^{-1}\hat{\beta}\left(\hat{a} + \frac{1}{2}\left(\hat{B}^\star vec\left(\hat{\Sigma}\right) + \hat{\sigma}^2\iota_N\right)\right) \\ \hat{\lambda}_1 &= \left(\hat{\beta}\hat{\beta}'\right)^{-1}\hat{\beta}\hat{c}. \end{aligned} \tag{A.11}$$

### A.2.3 Implementation

We use as factors the first five principal components of the yield curve. We estimate the factor loadings at monthly frequency, and combine these estimates with the daily time-series of the factors to obtain daily estimates term premia and expectations components. We thereby follow section 4.4 of Adrian, Crump, and Moench (2013). The estimation sample covers Jan 1991 – Dec 2020, including 7584 daily observations.

Figure A.1: Decomposition of the 10-year UK Nominal Government Bond Yield



Note: The Figure shows the decomposition of the UK 10-year yields into expectation (gray dashed line) and term premium (solid green line) components, implied by the term structure model of Adrian, Crump, and Moench (2013) and Malik and Meldrum (2016) over the sample 1991m1-2020m12; the black solid (red dash) line depicts the realised (fitted) yields.

Figure A.1 shows the decomposition of the UK 10-year yields into expectation (gray dashed line) and term premium (solid green line) components; the black solid (red dash) line depicts the realised (fitted) yields. The obtained decomposition is similar to recent estimates of the UK term structure decompositions (Moench, 2019). Daily changes in the estimated expectation and term premium components of the 10-year yields are then used as left-hand side variables in our baseline regressions 3.1–3.2, with the results presented in Table 6.

## A.3 Extensions and Robustness Checks

### A.3.1 Timing and Size of Primary Issuances

Given the importance of bond issuance in driving the pre-news drift, an interesting question is whether the issuance decision is strategically timed with respect to scheduled announcements. For example, there is an obvious interaction in that almost none of the primary issuances are scheduled by the Debt Management Office on days with MPC meetings. However, as discussed in Section 2.2, a non-negligible number of issuances occur in the pre-MPC windows. The question arises whether the amount of issuances might be systemically different on days around MPC meetings or scheduled labour market data releases compared to other periods. It might be that the risk premium effect we find in pre-news windows might be the result of disproportionately larger issuances in these periods.

To check this, we compute total weekly issuances (in nominal bonds, inflation-linked bonds and both types of bonds) and estimate whether the issued amount is statistically different on weeks with scheduled scheduled announcements compared to weeks without announcements. Panel A of Table A.2 presents the results for monetary policy announcements, indicating that total issuance in the whole sample is about £411 million smaller during weeks with MPC meetings. As shown in Panel B of Table A.2, we find no statistically significant difference in issuance size during weeks with scheduled labour market data releases compared to other weeks.

Table A.2: Average Issuance during Weeks with and without Scheduled Macroeconomic Announcements

| | (1) Nominal | (2) Linker | (3) All Bonds |
|---|---|---|---|
| Panel A: Weekly Issuance During MPC Weeks | | | |
| Change in News Weeks | -272.07* | -139.48*** | -411.55*** |
| | (-1.75) | (-3.60) | (-2.63) |
| Weeks without News | 1769.99*** | 343.18*** | 2113.17*** |
| | (21.96) | (12.81) | (25.41) |
| *N* | 1266 | 1266 | 1266 |
| Panel B: Weekly Issuance During Labour Market Data Release Weeks | | | |
| Change in News Weeks | 168.97 | -168.65*** | 0.32 |
| | (0.98) | (-4.42) | (0.00) |
| Weeks without News | 1725.81*** | 362.34*** | 2088.15*** |
| | (21.21) | (12.74) | (24.87) |
| *N* | 1218 | 1218 | 1218 |

Note: the table reports average issuances of nominal bonds (column 1), inflation-linked bonds (column 2) and all bonds (column 3) on weeks without scheduled MPC meetings as well as the difference in issuances on weeks with scheduled MPC meetings. The values are in £ millions, and the stars indicate whether the values are statistically different from zero (using robust standard errors) at 1%, 5 and 10% significance levels. The whole sample includes 1270 weeks, and the subsamples 1997-2004, 2005-2012 and 2013-2021 include 403, 418 and 449 weeks, respectively.

### *A.3.2 Daily Yield Changes around Macroeconomic Announcements*

In our baseline estimation, we compare yield changes during a two-day period before macroeconomic announcements to daily yield changes on other days. Here we take a closer look at whether individual days around announcements feature significant yield changes by estimating variants of regressions 3.1–3.2. Table A.3 shows how significantly different daily changes in 20-year yields are on selected days around MPC meetings compared to other days. Within each panel of the table, each column shows regression coefficients together with $t$-statistics based on robust standard errors. For example, column 4 shows the regression results corresponding to changes in yields based on end-of-day closing prices on the day before MPC meetings ($t - 1$) and closing prices on MPC days ($t$).

A shown by Table A.3, we find that the majority of the yield drift is concentrated in periods that are one day and two days prior to MPC days. Yield changes on all other selected days are not statistically different from yield changes in the rest of the trading days (Panel A). This suggests that there is no clear reversal of the pre-MPC drift after the MPC announcement. Similar to our baseline results, the inclusion of weekday fixed effects weakens these results, suggestive of within-week seasonalities (Panel B). Moreover, the interaction between pre-MPC windows and issuance is strongest the day before the announcement (Panels C-D). A shown by Table A.4, the results are qualitatively similar but economically and statistically more significant when we focus on scheduled labour market data releases instead of monetary policy announcements.

Table A.3: Daily Yield Changes around MPC Meetings

| | $\Delta_{t-4,t-3}$ (1) | $\Delta_{t-3,t-2}$ (2) | $\Delta_{t-2,t-1}$ (3) | $\Delta_{t-1,t}$ (4) | $\Delta_{t,t+1}$ (5) | $\Delta_{t+,1t+2}$ (6) |
|---|---|---|---|---|---|---|
| Panel A: Around MPC Announcements (without weekday fixed effects) | | | | | | |
| Days Around News | -0.24 | 0.50* | 0.58** | 0.18 | -0.27 | 0.28 |
| | (-0.93) | (1.77) | (2.04) | (0.53) | (-0.70) | (1.02) |
| Constant | -0.10 | -0.13** | -0.13** | -0.11** | -0.09* | -0.12** |
| | (-1.63) | (-2.20) | (-2.26) | (-1.98) | (-1.65) | (-2.02) |
| *N* | 6125 | 6126 | 6127 | 6128 | 6128 | 6127 |
| Panel B: Around MPC Announcements (with weekday fixed effects) | | | | | | |
| Days Around News | -0.12 | 0.48 | 0.39 | 0.15 | 0.04 | 0.44 |
| | (-0.42) | (1.56) | (1.28) | (0.43) | (0.11) | (1.46) |
| Weekday fixed effects | YES | YES | YES | YES | YES | YES |
| *N* | 6125 | 6126 | 6127 | 6128 | 6128 | 6127 |
| Panel C: Around MPC Announcements and after Issuance (without weekday fixed effects) | | | | | | |
| Days Around News # Bond issuance | -0.11 | 0.62 | 0.83** | 0.14 | -0.60 | 0.52 |
| | (-0.33) | (1.54) | (2.05) | (0.30) | (-1.06) | (1.41) |
| Days Around News # No Bond issuance | -0.40 | 0.36 | 0.28 | 0.23 | 0.15 | -0.02 |
| | (-1.01) | (0.93) | (0.72) | (0.48) | (0.31) | (-0.06) |
| Constant | -0.10 | -0.13** | -0.13** | -0.11** | -0.09* | -0.12** |
| | (-1.63) | (-2.20) | (-2.26) | (-1.98) | (-1.65) | (-2.02) |
| *N* | 6125 | 6126 | 6127 | 6128 | 6128 | 6127 |
| Panel D: Around MPC Announcements and after Issuance (with weekday fixed effects) | | | | | | |
| Days Around News # Bond issuance | 0.02 | 0.60 | 0.64 | 0.11 | -0.28 | 0.70* |
| | (0.04) | (1.41) | (1.51) | (0.23) | (-0.49) | (1.75) |
| Days Around News # No Bond issuance | -0.29 | 0.34 | 0.10 | 0.20 | 0.44 | 0.14 |
| | (-0.69) | (0.85) | (0.26) | (0.41) | (0.91) | (0.34) |
| Weekday fixed effects | YES | YES | YES | YES | YES | YES |
| *N* | 6125 | 6126 | 6127 | 6128 | 6128 | 6127 |

Note: This table shows average daily changes in the 20-year yield on UK nominal government bonds on days around MPC meetings (subpanel A), around MPC meetings with issuances one or two days before the meeting (subpanel B) and around MPC meetings with large issuances one or two days before the meeting. Panel A uses all MPC meetings and Panel B uses only scheduled MPC meetings. All regressions include a constant (not shown). The estimation period covers 1997m5-2021m7. The t-statistics are based on robust standard errors.

Table A.4: Daily Yield Changes around Labour Market Data Release

| | $\Delta_{t-4,t-3}$<br>(1) | $\Delta_{t-3,t-2}$<br>(2) | $\Delta_{t-2,t-1}$<br>(3) | $\Delta_{t-1,t}$<br>(4) | $\Delta_{t,t+1}$<br>(5) | $\Delta_{t+,1t+2}$<br>(6) |
|---|---|---|---|---|---|---|
| Panel A: Around Labour Market Data Releases (without weekday fixed effects) | | | | | | |
| Days Around News | 0.38 | 0.48* | 0.27 | 0.26 | 0.23 | -0.49* |
| | (1.20) | (1.85) | (1.02) | (0.97) | (0.78) | (-1.86) |
| Constant | -0.10* | -0.10* | -0.09 | -0.09 | -0.09 | -0.05 |
| | (-1.65) | (-1.70) | (-1.55) | (-1.53) | (-1.51) | (-0.91) |
| *N* | 5893 | 5894 | 5895 | 5896 | 5895 | 5894 |
| Panel B: Around Labour Market Data Releases (with weekday fixed effects) | | | | | | |
| Days Around News | 0.73** | 0.77*** | 0.19 | 0.04 | 0.20 | -0.27 |
| | (2.17) | (2.79) | (0.66) | (0.13) | (0.63) | (-0.93) |
| Weekday fixed effects | YES | YES | YES | YES | YES | YES |
| *N* | 5893 | 5894 | 5895 | 5896 | 5895 | 5894 |
| Panel C: Around Labour Market Data Releases and after Issuance (without weekday fixed effects) | | | | | | |
| Days Around News # Bond issuance | 0.68 | 1.41*** | 0.73* | 0.15 | -0.06 | -0.77 |
| | (1.25) | (2.99) | (1.94) | (0.39) | (-0.11) | (-1.59) |
| Days Around News # No Bond issuance | 0.22 | -0.03 | 0.03 | 0.32 | 0.38 | -0.34 |
| | (0.57) | (-0.09) | (0.08) | (0.91) | (1.10) | (-1.13) |
| Constant | -0.10* | -0.10* | -0.09 | -0.09 | -0.09 | -0.05 |
| | (-1.65) | (-1.70) | (-1.55) | (-1.53) | (-1.51) | (-0.91) |
| *N* | 5893 | 5894 | 5895 | 5896 | 5895 | 5894 |
| Panel D: Around Labour Market Data Releases and after Issuance (with weekday fixed effects) | | | | | | |
| Days Around News # Bond issuance | 0.95* | 1.69*** | 0.67* | -0.06 | -0.11 | -0.58 |
| | (1.74) | (3.50) | (1.75) | (-0.16) | (-0.20) | (-1.17) |
| Days Around News # No Bond issuance | 0.60 | 0.25 | -0.08 | 0.09 | 0.37 | -0.10 |
| | (1.50) | (0.83) | (-0.21) | (0.26) | (1.01) | (-0.29) |
| Weekday fixed effects | YES | YES | YES | YES | YES | YES |
| *N* | 5893 | 5894 | 5895 | 5896 | 5895 | 5894 |

Note: This table shows average daily changes in the 20-year yield on UK nominal government bonds on days around MPC meetings (subpanel A), around MPC meetings with issuances one or two days before the meeting (subpanel B) and around MPC meetings with large issuances one or two days before the meeting. Panel A uses all MPC meetings and Panel B uses only scheduled MPC meetings. All regressions include a constant (not shown). The estimation period covers 1997m5-2021m7. The t-statistics are based on robust standard errors.

### *A.3.3 Intra-Week Seasonalities in Yields*

Tables 2–3 in the main text show that the inclusion of weekday fixed effects strengthens the results for labour market news but weakens the results for monetary policy announcements. To shed light on the differential effect of weekday fixed effects on our baseline results, this section explores whether there are intra-week seasonalities in yields. As shown in Figure A.3, yields tend to fall unconditionally on Mondays and Fridays in our sample. These days tend to feature in two-day pre-news periods for labour market data releases. (In our sample 236 labour market data release days fall on Wednesdays and just 39 fall on Tuesdays). In contrast, Figure A.3 shows that yields tend to rise unconditionally on Wednesdays in our sample. Wednesdays tend to feature in two-day pre-news periods for BoE monetary policy announcements (as almost all scheduled MPC meetings are on Thursdays), but they never feature in two-day pre-news periods for labour market data releases.

Without the inclusion of weekday fixed effects, some of the upwards drift in yields in the two-day pre-news period for labour market data releases will be offset by the unconditional fall in yields on Mondays relative to other days of the week. This will lead to an underestimation of the drift. The inclusion of weekday fixed effects will effectively eliminate the cross weekday comparison, and will compare Mondays with impending labour market data releases to the other three Mondays of the month without impending labour market data releases.

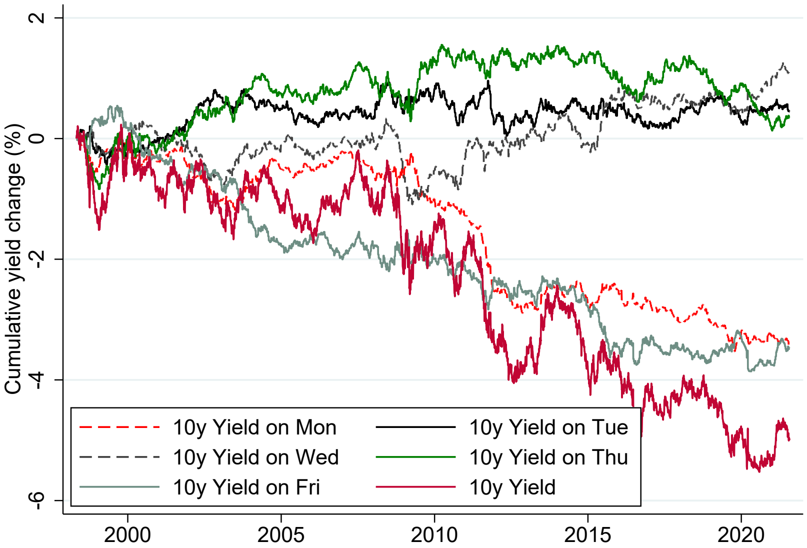
Based on a similar argument, in the case of the two-day pre-news period for BoE monetary policy announcements,the unconditional rise in yields on Wednesdays means the inclusion of weekday fixed effects will lower the estimate of drift compared to the regression without fixed effects.

Another way of presenting the argument is to check how movements in yields on different weekdays change when there is impending news. To that end, we decompose the cumulative yield changes on Mondays (shown by Figure A.3) into those Mondays that are followed by labour market data releases versus other Mondays. Figure A.3 shows that the unconditional fall in yields on Mondays occurs on those Mondays that are not followed by labour market data releases on Tuesday or Wednesday (black line). In contrast, yields actually tend to rise on Mondays that are followed by labour market data releases (red dashed line).

Similarly, we decompose the seasonal yield movements on Wednesdays into those Wednesdays that are followed by BoE monetary policy announcements the next day and those Wednesdays that are not followed by announcements. Figure A.4 shows the results, which indicate a more pronounced drift on Wednesdays without impending announcements.

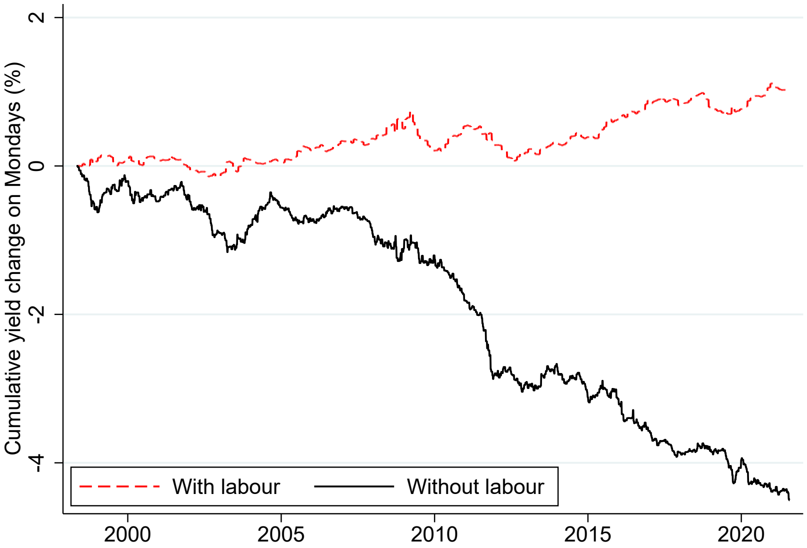
Importantly, we also check how the inclusion of weekday fixed affects the results when we pool monetary policy announcements and labour market data releases together and treat these two events the same in our regressions. As discussed above, the inclusion of weekday fixed weakens the monetary policy results and strengthens the labour market news results. These offsetting effects suggest that pooling the types of macroeconomic news would make the inclusion of weekday fixed effects have a smaller effect on the baseline results. This is indeed the case, as shown by Tables A.5–A.6 below.

Figure A.2: Cumulative yield changes on different weekdays



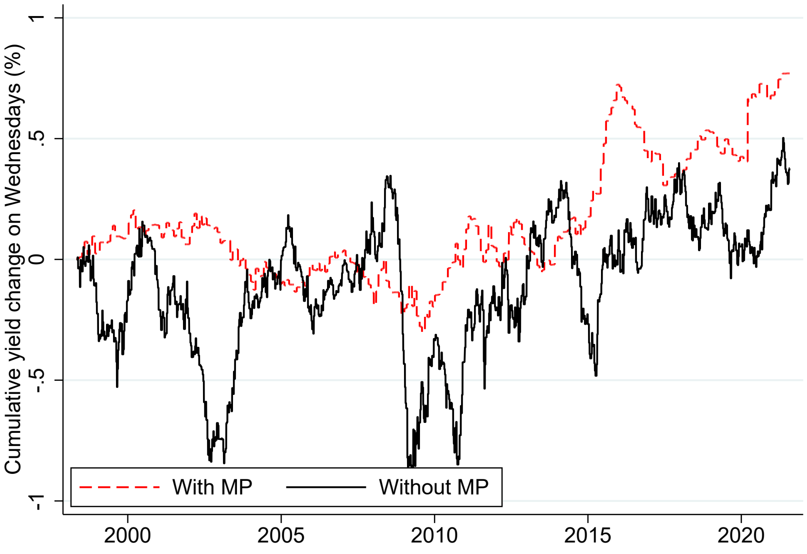
Note: this figure shows the cumulative 10-year yield changes on different weekdays along with the realise yield change (red solid line) over the period 1998m5-2021m7..

Figure A.3: Cumulative yield changes on Mondays: With and Without Upcoming Labour Data News



Note: this figure shows the cumulative 10-year yield changes on Mondays with (without) impending labour market data releases, shown by the red (black) line, over the period 1998m5-2021m7.

Figure A.4: Cumulative yield changes on Wednesdays: With and Without Upcoming MPC



Note: this figure shows the cumulative 10-year yield changes on Wednesdays with (without) impending MPC news, shown by the red (black) line, over the period 1998m5-2021m7..

Table A.5: Yield Changes before Monetary or Labour Market News

| | 5$Y$-yield<br>(1) | 10$Y$-yield<br>(2) | 15$Y$-yield<br>(3) | 20$Y$-yield<br>(4) |
|---|---|---|---|---|
| (A) Without Weekday Fixed Effects | | | | |
| Pre-News window | 0.39** | 0.47*** | 0.47*** | 0.46*** |
| | (2.52) | (2.99) | (3.16) | (3.14) |
| Constant | -0.17** | -0.17** | -0.17** | -0.16** |
| | (-2.52) | (-2.44) | (-2.51) | (-2.53) |
| $N$ | 5896 | 5896 | 5896 | 5896 |
| (B) With Weekday Fixed Effects | | | | |
| Pre-News window | 0.35** | 0.50*** | 0.47*** | 0.43*** |
| | (2.18) | (3.01) | (2.98) | (2.74) |
| Weekday fixed effects | YES | YES | YES | YES |
| $N$ | 5896 | 5896 | 5896 | 5896 |

Note: Panel A (B) of this table regresses daily changes in 5-year, 10-year, 15-year and 20-year nominal gilt yields on an indicator variable that takes value one for days that are either one or two days before scheduled monetary policy announcements or labour market data releases and zero otherwise. The estimation period in Panel A covers 1998m5-2021m7 and includes 540 announcements. Panel A (B) presents the results without (with) weekday fixed effects. T-statistics in parentheses are based on robust standard errors. Asterisks denote significance levels (* p<0.1, ** p<0.05, *** p<0.01).

Table A.6: Yield Changes before Monetary or Labour Market News: the Role of Bond Issuance

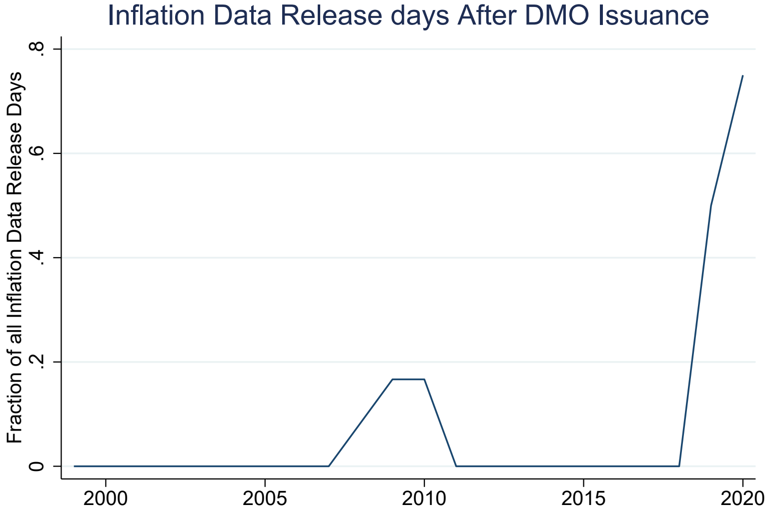
| | $5Y$-yield<br>(1) | $10Y$-yield<br>(2) | $15Y$-yield<br>(3) | $20Y$-yield<br>(4) |
|---|---|---|---|---|
| (A) Without Weekday Fixed Effects | | | | |
| Pre-News window # Issuance | 0.42* | 0.71*** | 0.78*** | 0.80*** |
| | (1.72) | (2.80) | (3.20) | (3.37) |
| Pre-News window # No issuance | 0.45** | 0.41** | 0.38** | 0.35** |
| | (2.44) | (2.23) | (2.16) | (1.99) |
| No Pre-News # Issuance | 0.30* | 0.29 | 0.28 | 0.25 |
| | (1.66) | (1.48) | (1.50) | (1.42) |
| Constant | -0.22*** | -0.22*** | -0.21*** | -0.20*** |
| | (-2.98) | (-2.86) | (-2.92) | (-2.91) |
| $N$ | 5896 | 5896 | 5896 | 5896 |
| (B) With Weekday Fixed Effects | | | | |
| Pre-News window # Issuance | 0.31 | 0.67** | 0.70*** | 0.69*** |
| | (1.16) | (2.47) | (2.75) | (2.77) |
| Pre-News window # No issuance | 0.44** | 0.47** | 0.41** | 0.34* |
| | (2.30) | (2.42) | (2.19) | (1.84) |
| No Pre-News # Issuance | 0.18 | 0.17 | 0.16 | 0.13 |
| | (0.91) | (0.83) | (0.81) | (0.70) |
| Weekday fixed effects | YES | YES | YES | YES |
| $N$ | 5896 | 5896 | 5896 | 5896 |

Note: this table regresses daily changes in 5-year, 10-year, 15-year and 20-year nominal gilt yields on an indicator variables on indicator variables capturing (i) pre-news windows with new bond issuance, (ii) pre-news windows without new bond issuance and (iii) all trading days with issuance and without announcements. The estimation period in Panel A covers 1998m5-2021m7 and includes 540 announcements. Panel A (B) presents the results without (with) weekday fixed effects. T-statistics in parentheses are based on robust standard errors. Asterisks denote significance levels (* p<0.1, ** p<0.05, *** p<0.01).

### *A.3.4 Inflation Data Releases*

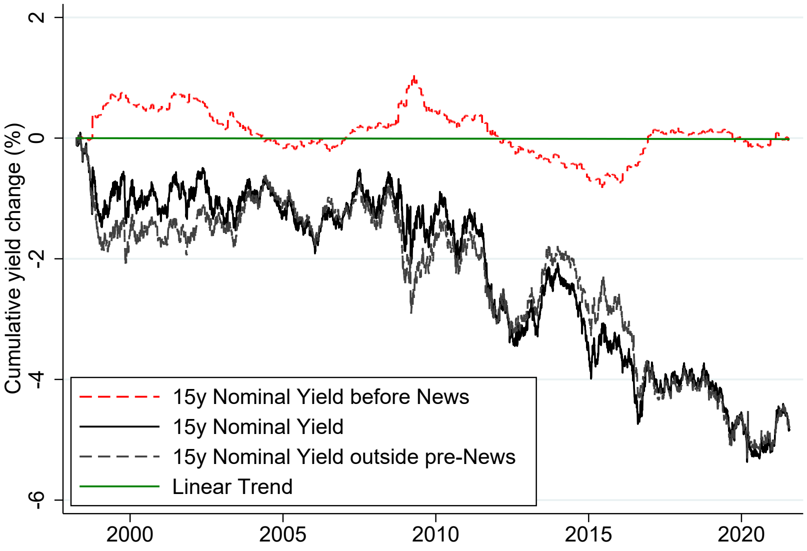
We also check whether our results hold for macroeconomic announcements associated with inflation data releases. Figure A.5 shows that during the majority of our sample there are no bond issuances in these pre-news windows. Consistent with this and with our argument in the paper, Figure shows A.6 that there is no visible pre-news drift associated with these types of announcements.

Figure A.5: MPC Days and Issuance Days over Time: 1999-2020



Notes: this figure shows the fraction of pre-inflation news windows (in a given year) that coincided with new government bond issuance.

Figure A.6: A Decomposition of Long-term Gilt Yields: The Role Inflation Data Releases



Note: this figure documents that 15-year UK nominal bond yields tend to rise in the 2-day window before scheduled inflation data releases. This 2-day window includes for every release date the day prior to the meeting, and the day that is two days before the meeting. The black line shows the actual evolution of the yield. The red line shows a hypothetical time series that is constructed by taking into account only the yield changes that were realised in the 2-day window before release dates; the yield changes that occurred on all days outside of this window are set to zero. The green line is an estimated linear trend associated with the red line. The gray line shows a hypothetical time series that is constructed by taking into account only the yield changes that were realised outside the 2-day window before data release dates. The analysis includes all 280 labour market data release days from April 1998 to June 2021.

### *A.3.5 Evidence from the US*

In a recent paper, Hillenbrand (2020) documents that a narrow window around monetary policy meetings of the Fed captures the secular decline in nominal long-term interest rates over the last three decades. This appears to be counter to the results we find in this paper. This section performs a consistency check and argues that the effect we identify using UK data is present in the US as well.

Figure A.7a carries out the analysis of Hillenbrand (2020) for the period 1989m6-2019m12, confirming his baseline result that a 3-day window around FOMC meetings fully captured the decline in interest rates of the last decades.[41] The interpretation put forth is that FOMC meetings provide information to the market about the long-term interest rates, even though the Fed does not have direct control over these rates. This information effect would suggest that the market learned about the secular decline in interest rates from the Fed.

At first sight, this seems in contrast with the pre-MPC drift documented for the UK by Figure 2. However, we argue that the effect our paper studies, pertaining to the interaction between bond issuance and central bank announcements, is present in the US as well. To that end, we first note that out of the 268 FOMC meeting windows in the sample, 88 of them coincide with issuance of nominal government bonds with more than four year of maturity, and the remaining 180 FOMC windows have no concurrent issuance of such Treasuries. Following our analysis above, Figure A.7b splits the FOMC meeting windows into two groups: one set of meetings that coincide with such new issuances and the remaining set of windows without issuances. We find that all of the downward drift in long-term yields concentrates in FOMC windows without issuances, as shown by the red line. In contrast, there is no visible change in long-term yields during one third of all FOMC windows in our sample that coincide with new issuances of longer term bonds. These results provide an important extension to the findings of Hillenbrand (2020), suggesting that the interaction between FOMC windows and bond supply effects seems to generate a countervailing force which offsets some of the proposed learning effect associated with the downward drift in interest rates.

Given that the mechanism highlighted in our paper works via risk premia due to the interaction between agents' limited risk bearing capacity and supply effects in the vicinity of information events, it is natural to decompose the term premium component of the FOMC drift. Hillenbrand (2020) finds little cumulative decline in the term premium estimates around the 3 days around FOMC meetings. We corroborate this finding in Figure A.8a, and add to this evidence by decomposing the drift in the term premium that are attributed to FOMC meetings with and without concurrent Treasury issuances. Figure A.8b shows that there is a positive

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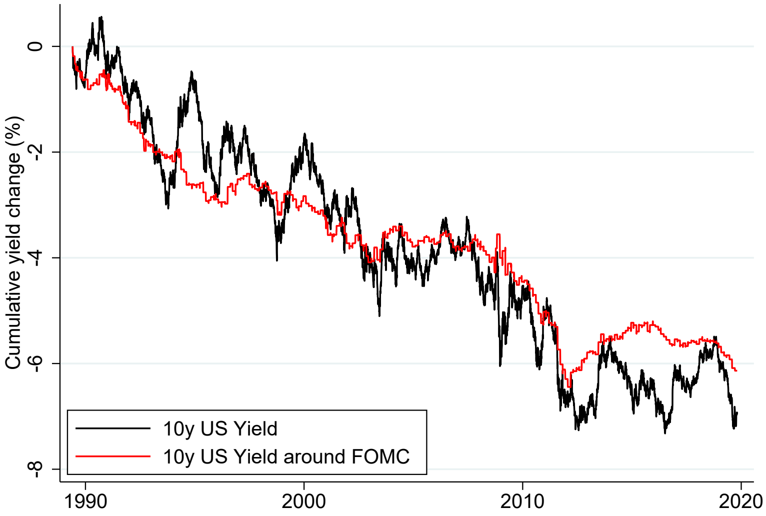
[41] The data on nominal US treasury yields are obtained from the updated dataset of Gurkaynak, Sack, and Wright (2007).

cumulative change in term premia during FOMC windows that coincide with Treasury issuance (consistent with the UK evidence), and the cumulative change has been negative during FOMC windows without issuance. The divergence in the two hypothetical series occurs after the Great Recession when primary dealers' risk-bearing capacity became more limited (Adrian, Boyarchenko, and Shachar, 2017; Bessembinder, Jacobsen, Maxwell, and Venkataraman, 2018; Bao, O'Hara, and Zhou, 2018) and the total amount of marketable Treasuries started outgrowing dealers' intermediation capacity (Duffie, 2020).

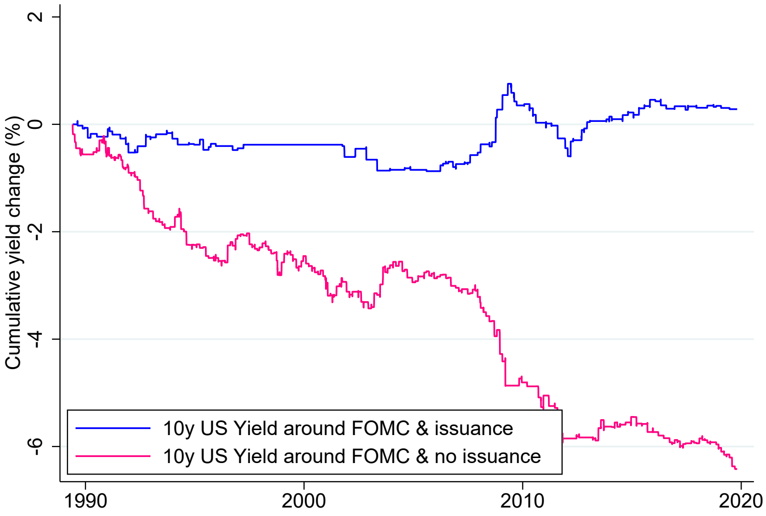
We also check how the results change when applying our timing assumption (by using the two-day window before FOMC meetings) instead of using the timing assumption of Hillenbrand (2020). Figure A.9 in the Appendix shows that the overall downward drift in yields is weaker when focusing on two-day windows before FOMC meetings. Importantly, however, long-term US yields continue to exhibit a non-negative drift during pre-FOMC windows that coincide with new issuances of long-term government bonds – this is largely driven by the term premium component of long-term yields, as shown by Figure A.10. Overall, these results are consistent with our findings for the UK.

Figure A.7: A Decomposition of Long-term US Yields: 1980m1-2019m10

(a) Specification of Hillenbrand (2020)



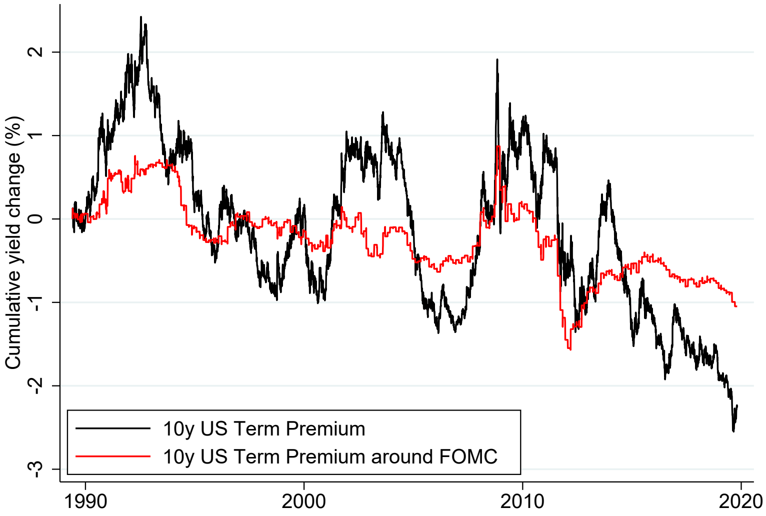
(b) The FOMC-drift with and without New Bond Issuance



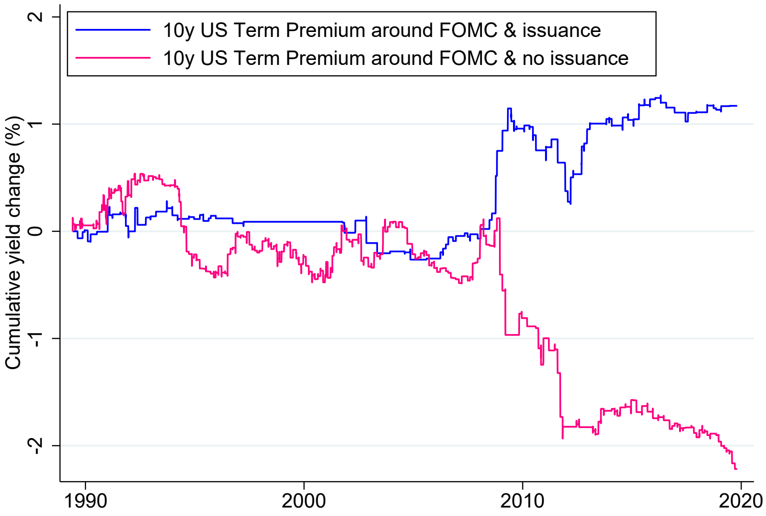
Note: Panel A of this figure is constructed following Hillenbrand (2020) and documents that the 3-day window around FOMC meetings captures the secular decline of the 10-year U.S. Treasury yield. This 3-day window includes for every FOMC meeting the day prior to the meeting, the day of the meeting and the day after the meeting. The black line shows the actual evolution of the 10-year U.S. Treasury yield. The red line in Panel A shows a hypothetical time series that is constructed by taking into account only the yield changes that were realized in the 3-day window around FOMC meetings; the yield changes that occurred on all days outside of this window are set to zero. Panel B decomposes the red line in Panel A by distinguish between FOMC meeting windows with and without new issuances of US Treasuries (with more than four years of maturity). Out of the 268 FOMC meeting windows in the sample, 88 of them coincide with issuances, and the remaining 180 FOMC windows have no concurrent issuance of such Treasuries. The blue (magenta) line shows the hypothetical time series based on FOMC windows with (without) new bond issuances. The analysis includes all FOMC meetings from June 1989 to Oct 2019.

Figure A.8: A Decomposition of Long-term US Term Premium: 1980m1-2019m10

(a) Specification of Hillenbrand (2020)



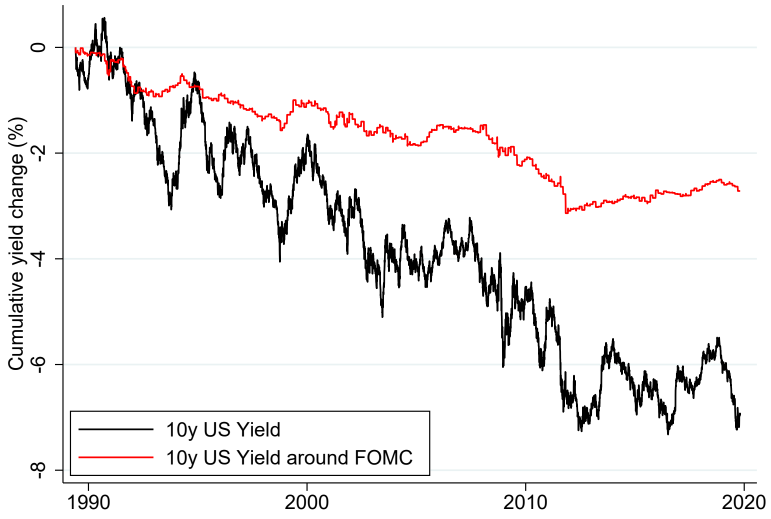
(b) The FOMC-drift with and without New Bond Issuance



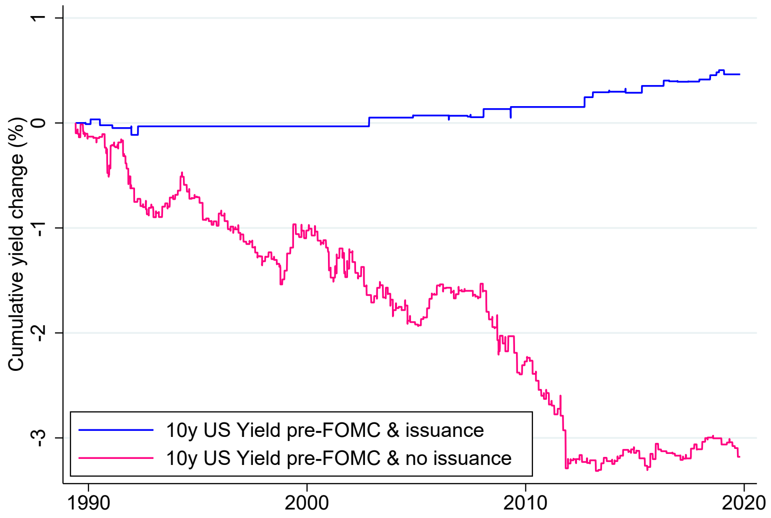
Note: Panel A of this figure is constructed following Hillenbrand (2020) and shows the realised and counterfactual time-series of the 10-year US term premium, obtained from the Federal Reserve Bank of New York website. The counterfactual time-series (red line) are based on changes in the premium during the 3-day FOMC window: the day prior to the meeting, the day of the meeting and the day after the meeting. The black line shows the actual evolution of the estimated term premium. Panel B decomposes the red line in Panel A by distinguish between FOMC meeting windows with and without new issuances of US Treasuries (with more than four years of maturity). Out of the 268 FOMC meeting windows in the sample, 88 of them coincide with issuances, and the remaining 180 FOMC windows have no concurrent issuance of such Treasuries. The blue (magenta) line shows the hypothetical time series based on FOMC windows with (without) new bond issuances. The analysis includes all FOMC meetings from June 1989 to Oct 2019.

Figure A.9: A Decomposition of Long-term US Yields: 1980m1-2019m10

(a) Two-day Window before FOMC



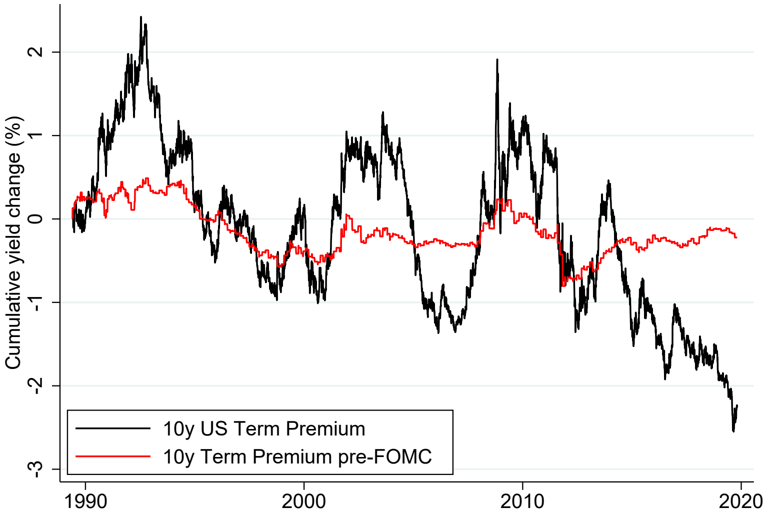
(b) Two-day Window before FOMC: the Role of New Bond Issuance



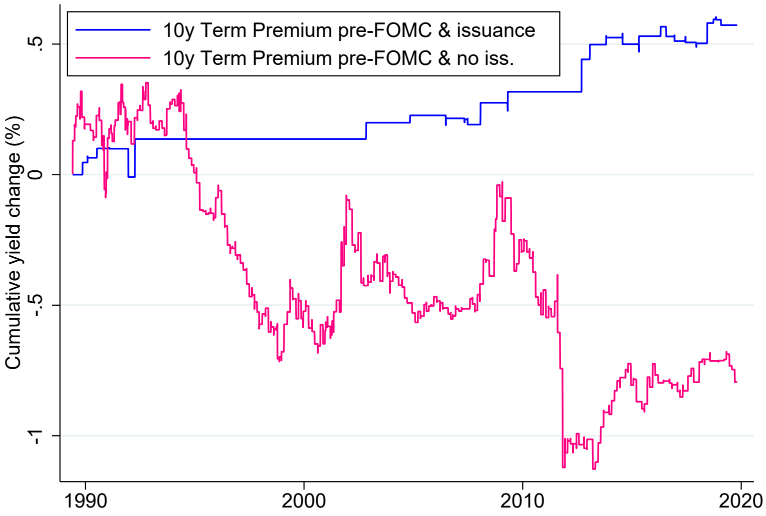
Note: Panel A of this figure shows cumulative US yield drift in a two-day window before FOMC meetings, thereby adopting the definition of pre-MPC windows used in our baseline. The black line shows the actual evolution of the 10-year U.S. Treasury yield. The red line in Panel A shows a hypothetical time series that is constructed by taking into account only the yield changes that were realized in the two-day window before FOMC meetings; the yield changes that occurred on all days outside of this window are set to zero. Panel B decomposes the red line in Panel A by distinguish between FOMC meeting windows with and without new issuances of US Treasuries (with more than four years of maturity). The blue (magenta) line shows the hypothetical time series based on FOMC windows with (without) new bond issuances. The analysis includes all FOMC meetings from June 1989 to Oct 2019.

Figure A.10: A Decomposition of Long-term US Term Premium: 1980m1-2019m10

(a) Two-day Window before FOMC



(b) Two-day Window before FOMC: the Role of New Bond Issuance



Note: Panel A of this figure is constructed following Hillenbrand (2020) and shows the realised and counterfactual time-series of the 10-year US term premium, obtained from the Federal Reserve Bank of New York website. The counterfactual time-series (red line) are based on changes in the premium during the two-day windwo before FOMC meetings. The black line shows the actual evolution of the estimated term premium. Panel B decomposes the red line in Panel A by distinguish between FOMC meeting windows with and without new issuances of US Treasuries (with more than four years of maturity). The blue (magenta) line shows the hypothetical time series based on FOMC windows with (without) new bond issuances. The analysis includes all FOMC meetings from June 1989 to Oct 2019.

### *A.3.6 Aggregate Evidence using Yield Curve Noise*

As an additional robustness check, we use the noise measure of Hu, Pan, and Wang (2013) – constructed from deviations of individual bond yields from a fitted yield curve – which is often interpreted as a measure of constraints on financial intermediaries (Goldberg, 2020). Given this interpretation, we divide our sample into periods of lower and higher noise periods by sorting each month into two groups based on the yearly median value of the noise measure. We then estimate the effect of the interaction between primary issuances and pre-news windows on the two subsamples separately.

Table A.7 presents the estimation results corresponding to model (3.2), during lower noise (panel A, C) and higher noise (panel B, D), when we focus on monetary policy announcements. We find that the majority of the price effect, induced by the interaction between bond issuance and the pre-MPC drift, concentrates in periods of lower liquidity. For example, column 4 in Panels A-B of Table A.7 implies that the average daily 20-year yield drift in pre-MPC windows that coincide with bond issuances is 0.39 bps during liquid market conditions, whereas the pre-MPC drift is 1.09 bps when it coincides with bond issuances and more illiquid conditions.

Table A.8 presents the results for the case when we focus on labour market data releases. We find that the interaction between pre-news windows and primary issuances during illiquid periods are even larger: the corresponding estimates suggest a change in 20-year yields of about 2.4 bps during illiquid periods (Panels B and D) compared to virtually no yield change during periods if higher liquidity (Panels A and C).

As robustness check, we also estimate regression without splitting the sample into high-noise and low-noise periods, and instead include a dummy variable capturing high-low and noise periods.

As a robustness check, we also estimate regression models without splitting the sample into high-noise and low-noise periods. Instead, we include a dummy variable to capture high- and low-noise periods. As shown in Tables A.9–A.10, we obtain very similar results.

## Table A.7: Pre-news Drift and Bond Issuance: the Role of Bond Issuance and Gilt Market Noise

| | 5Y | 10Y | 15Y | 20Y | 5Y | 10Y | 15Y | 20Y |
|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| (A) Pre-MPC Drift & Lower Noise | | | | | (B) Pre-MPC Drift & Higher Noise | | | |
| Pre-News # Issue | 0.17 | 0.31 | 0.37 | 0.39 | 0.35 | 0.94** | 1.08*** | 1.09** |
| | (0.47) | (0.75) | (0.92) | (1.01) | (0.87) | (2.28) | (2.60) | (2.53) |
| Pre-News # No issue | 0.10 | 0.17 | 0.33 | 0.49 | -0.17 | -0.00 | 0.13 | 0.15 |
| | (0.28) | (0.44) | (0.88) | (1.31) | (-0.37) | (-0.01) | (0.29) | (0.33) |
| No News # Issue | 0.33 | 0.24 | 0.21 | 0.17 | 0.11 | 0.17 | 0.21 | 0.26 |
| | (1.43) | (0.94) | (0.88) | (0.75) | (0.42) | (0.62) | (0.81) | (1.00) |
| Constant | -0.05 | -0.08 | -0.09 | -0.09 | -0.25** | -0.23** | -0.21* | -0.20* |
| | (-0.55) | (-0.79) | (-0.98) | (-1.04) | (-2.32) | (-2.07) | (-1.96) | (-1.90) |
| *N* | 2761 | 2761 | 2761 | 2761 | 2688 | 2688 | 2688 | 2688 |
| (C) Pre-MPC Drift & Lower Noise | | | | | (D) Pre-MPC Drift & Higher Noise | | | |
| Pre-News # Issue | 0.11 | 0.26 | 0.26 | 0.22 | 0.01 | 0.65 | 0.81* | 0.84* |
| | (0.30) | (0.58) | (0.61) | (0.53) | (0.02) | (1.47) | (1.82) | (1.84) |
| Pre-News # No issue | 0.05 | 0.12 | 0.23 | 0.32 | -0.47 | -0.26 | -0.11 | -0.07 |
| | (0.13) | (0.29) | (0.57) | (0.83) | (-0.99) | (-0.53) | (-0.23) | (-0.15) |
| No News # Issue | 0.22 | 0.11 | 0.08 | 0.03 | -0.09 | -0.03 | 0.02 | 0.09 |
| | (0.86) | (0.38) | (0.29) | (0.11) | (-0.33) | (-0.09) | (0.09) | (0.32) |
| Weekday FE | YES | YES | YES | YES | YES | YES | YES | YES |
| *N* | 2761 | 2761 | 2761 | 2761 | 2688 | 2688 | 2688 | 2688 |

Note: This table regresses daily changes in yields on indicator variables capturing (i) pre-news windows (associated with monetary policy announcements) and new bond issuance, (ii) pre-news windows without new bond issuance and (iii) all trading days with issuance and without announcements. Panel A (Panel B) presents the results without weekday fixed effects using the subsample when the market is less (more) liquid. Panels C-D present the results with weekday fixed effects. We define high (low) noise periods as those when the yield curve noise measure in a given month is above (below) the median value in the given year. T-statistics in parentheses are based on robust standard errors. Asterisks denote significance levels (* $p<0.1$, ** $p<0.05$, *** $p<0.01$).

Table A.8: Pre-news Drift and Bond Issuance: the Role of Bond Issuance and Gilt Market Noise

| | 5Y | 10Y | 15Y | 20Y | 5Y | 10Y | 15Y | 20Y |
|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| (A) Pre-Labour Drift & Lower Noise | | | | | (B) Pre-Labour Drift & Higher Noise | | | |
| Pre-News # Issue | 0.44 | 0.43 | 0.16 | 0.07 | 1.85** | 2.12*** | 2.26*** | 2.44*** |
| | (0.59) | (0.66) | (0.28) | (0.12) | (2.41) | (2.60) | (3.04) | (3.44) |
| Pre-News # No issue | 0.23 | 0.15 | 0.07 | -0.01 | 0.81** | 0.86** | 0.85*** | 0.81** |
| | (0.80) | (0.51) | (0.26) | (-0.03) | (2.32) | (2.52) | (2.65) | (2.55) |
| No News # Issue | 0.28 | 0.22 | 0.20 | 0.16 | 0.14 | 0.29 | 0.39 | 0.45* |
| | (1.29) | (0.88) | (0.85) | (0.72) | (0.57) | (1.15) | (1.57) | (1.86) |
| Constant | -0.06 | -0.07 | -0.07 | -0.06 | -0.33*** | -0.30*** | -0.28** | -0.27** |
| | (-0.68) | (-0.75) | (-0.75) | (-0.64) | (-3.08) | (-2.65) | (-2.55) | (-2.53) |
| *N* | 2761 | 2761 | 2761 | 2761 | 2688 | 2688 | 2688 | 2688 |
| (C) Pre-Labour Drift & Lower Noise | | | | | (D) Pre-Labour Drift & Higher Noise | | | |
| Pre-News # Issue | 0.52 | 0.55 | 0.23 | 0.10 | 1.89** | 2.21*** | 2.31*** | 2.47*** |
| | (0.69) | (0.82) | (0.39) | (0.17) | (2.46) | (2.71) | (3.10) | (3.45) |
| Pre-News # No issue | 0.32 | 0.28 | 0.17 | 0.06 | 0.92** | 1.03*** | 0.98*** | 0.90*** |
| | (1.08) | (0.90) | (0.56) | (0.20) | (2.51) | (2.85) | (2.88) | (2.68) |
| No News # Issue | 0.18 | 0.10 | 0.07 | 0.02 | -0.06 | 0.10 | 0.19 | 0.28 |
| | (0.76) | (0.39) | (0.28) | (0.07) | (-0.23) | (0.36) | (0.74) | (1.07) |
| Weekday FE | YES | YES | YES | YES | YES | YES | YES | YES |
| *N* | 2761 | 2761 | 2761 | 2761 | 2688 | 2688 | 2688 | 2688 |

Note: This table regresses daily changes in yields on indicator variables capturing (i) pre-news windows (associated with labour market data releases) and new bond issuance, (ii) pre-news windows without new bond issuance and (iii) all trading days with issuance and without announcements. Panel A (Panel B) presents the results without weekday fixed effects using the subsample when the market is less (more) liquid. Panels C-D present the results with weekday fixed effects. We define high (low) noise periods as those when the yield curve noise measure in a given month is above (below) the median value in the given year. T-statistics in parentheses are based on robust standard errors. Asterisks denote significance levels (* p<0.1, ** p<0.05, *** p<0.01).

Table A.9: Pre-MPC Drift and Bond Issuance: the Role of Bond Issuance and Gilt Market Noise

| | 5Y | 10Y | 15Y | 20Y | 5Y | 10Y | 15Y | 20Y |
|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| Pre-News # Issue # Lower Noise | 0.27 | 0.39 | 0.43 | 0.44 | 0.07 | 0.21 | 0.23 | 0.23 |
| | (0.75) | (0.94) | (1.08) | (1.15) | (0.18) | (0.50) | (0.56) | (0.57) |
| Pre-News # Issue # Higher Noise | 0.25 | 0.86** | 1.02** | 1.03** | 0.05 | 0.69* | 0.83** | 0.82* |
| | (0.64) | (2.14) | (2.50) | (2.45) | (0.12) | (1.65) | (1.97) | (1.89) |
| Pre-News # No issue # Lower Noise | 0.20 | 0.25 | 0.39 | 0.54 | 0.01 | 0.08 | 0.21 | 0.34 |
| | (0.55) | (0.64) | (1.06) | (1.47) | (0.04) | (0.21) | (0.54) | (0.90) |
| Pre-News # No issue # Higher Noise | -0.27 | -0.08 | 0.07 | 0.10 | -0.44 | -0.23 | -0.10 | -0.09 |
| | (-0.60) | (-0.17) | (0.16) | (0.22) | (-0.97) | (-0.49) | (-0.22) | (-0.19) |
| No News # Issue # Lower Noise | 0.43* | 0.32 | 0.27 | 0.23 | 0.27 | 0.15 | 0.11 | 0.07 |
| | (1.92) | (1.28) | (1.16) | (1.00) | (1.12) | (0.58) | (0.43) | (0.28) |
| No News # Issue # Higher Noise | 0.01 | 0.09 | 0.15 | 0.21 | -0.15 | -0.07 | -0.01 | 0.04 |
| | (0.05) | (0.36) | (0.62) | (0.84) | (-0.56) | (-0.28) | (-0.05) | (0.17) |
| Weekday FE | NO | NO | NO | NO | YES | YES | YES | YES |
| *N* | 5454 | 5454 | 5454 | 5454 | 5454 | 5454 | 5454 | 5454 |

Note: This table regresses daily changes in yields on indicator variables capturing high vs low noise periods interacted with indicator variables capturing (i) pre-news windows (associated with monetary policy announcements) and new bond issuance, (ii) pre-news windows without new bond issuance and (iii) all trading days with issuance and without announcements. Columns 1-4 (5-8) present the results without (without) weekday fixed effects. We define high (low) noise periods as those when the yield curve noise measure in a given month is above (below) the median value in the given year. T-statistics in parentheses are based on robust standard errors. Asterisks denote significance levels (* p<0.1, ** p<0.05, *** p<0.01).

Table A.10: Pre-Labour News Drift and Bond Issuance: the Role of Bond Issuance and Gilt Market Noise

| | 5Y<br>(1) | 10Y<br>(2) | 15Y<br>(3) | 20Y<br>(4) | 5Y<br>(5) | 10Y<br>(6) | 15Y<br>(7) | 20Y<br>(8) |
|---|---|---|---|---|---|---|---|---|
| Pre-News # Issue # Lower Noise | 0.57 | 0.55 | 0.27 | 0.17 | 0.64 | 0.66 | 0.34 | 0.21 |
| | (0.78) | (0.84) | (0.46) | (0.31) | (0.85) | (0.99) | (0.57) | (0.37) |
| Pre-News # Issue # Higher Noise | 1.72** | 2.00** | 2.15*** | 2.34*** | 1.77** | 2.09*** | 2.20*** | 2.35*** |
| | (2.25) | (2.47) | (2.91) | (3.31) | (2.32) | (2.59) | (2.98) | (3.32) |
| Pre-News # No issue # Lower Noise | 0.37 | 0.27 | 0.18 | 0.10 | 0.46 | 0.41 | 0.29 | 0.17 |
| | (1.30) | (0.91) | (0.63) | (0.36) | (1.61) | (1.36) | (0.99) | (0.62) |
| Pre-News # No issue # Higher Noise | 0.68** | 0.75** | 0.75** | 0.70** | 0.78** | 0.89*** | 0.86*** | 0.78** |
| | (1.99) | (2.26) | (2.40) | (2.29) | (2.24) | (2.64) | (2.68) | (2.47) |
| No News # Issue # Lower Noise | 0.42** | 0.33 | 0.30 | 0.27 | 0.27 | 0.18 | 0.15 | 0.11 |
| | (1.97) | (1.39) | (1.34) | (1.23) | (1.21) | (0.73) | (0.62) | (0.49) |
| No News # Issue # Higher Noise | 0.00 | 0.18 | 0.28 | 0.34 | -0.15 | 0.02 | 0.12 | 0.18 |
| | (0.01) | (0.74) | (1.22) | (1.51) | (-0.62) | (0.08) | (0.49) | (0.77) |
| Weekday FE | NO | NO | NO | NO | YES | YES | YES | YES |
| *N* | 5454 | 5454 | 5454 | 5454 | 5454 | 5454 | 5454 | 5454 |

Note: This table regresses daily changes in yields on indicator variables capturing high vs low noise periods interacted with indicator variables capturing (i) pre-news windows (associated with labour market data releases) and new bond issuance, (ii) pre-news windows without new bond issuance and (iii) all trading days with issuance and without announcements. Columns 1-4 (5-8) present the results without (without) weekday fixed effects. We define high (low) noise periods as those when the yield curve noise measure in a given month is above (below) the median value in the given year. T-statistics in parentheses are based on robust standard errors. Asterisks denote significance levels (* p<0.1, ** p<0.05, *** p<0.01).

### *A.3.7 Further Evidence on Dealers' Inventory Dynamics*

In addition to the analysis presented in Section 4.1.2 of the main text, we also check a specification where we pool both types of news as well as pool all observations together (including those before news and no news). We then interact the previous period's orderflow with a news dummy, which takes value one in two-day pre-news periods for either labour market data releases or BoE monetary policy announcements. This enables us to test whether impending news induces post-issuance inventory dynamics in a statistically significant way. Table A.11 reiterates that inventory rebalancing by dealers participating in the auction occurs before news. The p-values hint at statistically significant differences (at the 15% level) for dealers participating in the auction compared to the rest of the dealers.

Table A.11: Evidence from Dealers' Inventory Imbalance

| | All Dealers<br>(1) | Dealers Participating in the Auction<br>(2) | Dealers Not Participating in the Auction<br>(3) |
|---|---|---|---|
| DealerFlow # No News | -0.004 | -0.006 | -0.004 |
| | (-1.49) | (-1.63) | (-1.02) |
| DealerFlow # News | -0.010** | -0.016** | -0.006 |
| | (-2.17) | (-2.51) | (-0.80) |
| p-values | 0.22 | 0.15 | 0.79 |
| $N$ | 5604 | 2213 | 3378 |
| $R^2$ | 0.103 | 0.191 | 0.115 |

Note: This table presents the estimation results corresponding to a modified version of 4.2, where we pool both types of news as well as pool all observations together (including those before news and before no news). T-statistics in parentheses are based on robust standard errors. Asterisks denote significance levels (* p<0.1, ** p<0.05, *** p<0.01).

### *A.3.8 Hedge Fund Volume around Announcements and Positioning*

**Hedge Fund Volume** To complement the analysis of hedge fund performance in Section 4.2.2, this section analyses the trading activity of hedge funds in the days leading up to either labour market data releases or BoE monetary policy announcements after issuance. We compute their total orderflow, gross volume and the total number of trades they make. As shown in Table A.12, we find evidence that hedge funds buy fewer bonds and provide less liquidity in two-day pre-news periods after issuance and on the day of news after issuance. Specifically, hedge funds provide less liquidity in two-day pre-news periods ahead of BoE monetary policy announcements and on the day of the announcements. This is probably because BoE monetary policy announcements come in the middle of the trading day, so there is still some pre-news uncertainty on the day of the announcements. In contrast, we find an increase in hedge fund liquidity provision on the day of labour market announcements. This is likely because labour market data releases occur at the start of the day, so the observed increase in trade activity occurs after the realisation of uncertainty.

Table A.12: Hedge Fund Trading after Issuance Day and around News

| | Order flow<br>(1) | Number of Trades<br>(2) |
|---|---|---|
| (A) Monetary Policy Announcements | | |
| After Issuance Day # 1 or 2 Days Before News | -58.41* | -11.72* |
| After Issuance Day # Day Of News | -142.81*** | -11.11 |
| Other Days | 40.63*** | 158.46*** |
| (B) Labour Market Data Releases | | |
| After Issuance Day # 1 or 2 Days Before News | -17.71 | -10.04* |
| After Issuance Day # Day Of News | 51.61 | 25.38*** |
| Other Days | 28.62** | 157.47*** |
| *N* | 1044 | 1044 |

Note: The table presents regression estimates with different columns for different measures of hedge fund liquidity provision in pre-news windows that coincide with primary issuance. The rows separately show trades that occurred 1 or 2 days before the news, and on the day of the news. Column 1 presents the total net order flow in value terms. Column 2 presents the total number of trades. Panel A (B) shows the results for monetary policy announcements (labour market data releases). The asterisks indicate whether the returns are different from zero at 1%, 5% and 10% significance levels.

**Hedge Fund Positioning** We also explore how hedge fund positioning before scheduled macro announcements – and so their revealed expectation for the sign of the news – affects their trading activity in the days leading up to those announcement in weeks after issuance. To proxy for hedge fund positioning on the day of the announcements, we compute their cumulative purchases of bonds in the preceding three days. (As a robustness check, we also check how the results change when we cumulate bond purchases in the preceding five trading days.) We then construct a dummy variable that takes value one if hedge funds are long bonds on the day of the announcement ('HFs Long Bonds' in the table) and zero if they are short ('HFs Short Bonds' in the table). We then check how this dummy correlates with hedge funds' trading activity after issuance and before the news. As above, the measures of trading activity we analyse are total orderflow and the total number of trades they make.

As shown in table A.13, our evidence suggests that hedge funds provide more liquidity after issuance in two-day pre-news periods ahead of labour market data releases and BoE monetary policy announcements if they expect news that would lead to an appreciation in bond prices. In other words, when hedge funds are building up to a long bond position on the day of the announcements they provide more liquidity to the market than when they are not. This is particularly true for periods ahead of labour market announcements across all three measures of trading activity. For periods ahead of BoE monetary policy announcements, the evidence is strongest for orderflow. (As shown in table A.14, these results are robust to extending the window of cumulating bond purchases to five trading days.)

Table A.13: Liquidity Provision after Issuance Day, 1 or 2 days before News: The Role of Hedge Fund Positioning

| | Order flow<br>(1) | Number of trades<br>(2) |
|---|---|---|
| (A) Monetary Policy Announcements | | |
| After Issuance Day # 1 or 2 Days Before News # HFs Long Bonds | 319.12*** | 10.56 |
| After Issuance Day # 1 or 2 Days Before News # HFs Short Bonds | -206.08*** | -17.07* |
| Other Days | 0.92 | 161.97*** |
| (B) Labour Market Data Releases | | |
| After Issuance Day # 1 or 2 Days Before News # HFs Long Bonds | 219.28*** | 39.09*** |
| After Issuance Day # 1 or 2 Days Before News # HFs Short Bonds | -105.50*** | -29.13*** |
| Other Days | -14.43 | 168.06*** |
| *N* | 1044 | 1044 |

Note: The table presents regression estimates that show the total net order flow (column 1) and number of trades (column 2) of hedge funds in pre-news windows that coincide with primary issuance. The rows show trades that occurred 1 or 2 days before the news, split by whether hedge funds had a long or short position in bonds on the day of the news. The long or short position is measured based on cumulative signed order flow in the three days leading up to the news. Panel A (B) shows the results for monetary policy announcements (labour market data releases). The asterisks indicate whether the returns are different from zero at 1%, 5% and 10% significance levels.

Table A.14: Liquidity Provision after Issuance Day, 1 or 2 days before News: Alternative Measure of Hedge Fund Positioning

| | Order flow<br>(1) | Number of trades<br>(2) |
|---|---|---|
| (A) Monetary Policy Announcements | | |
| After Issuance Day # 1 or 2 Days Before News # HFs Long Bonds | 133.45** | 3.40 |
| After Issuance Day # 1 or 2 Days Before News # HFs Short Bonds | -114.98** | -13.01 |
| Other Days | -16.16 | 159.18*** |
| (B) Labour Market Data Releases | | |
| After Issuance Day # 1 or 2 Days Before News # HFs Long Bonds | 170.52*** | 21.97** |
| After Issuance Day # 1 or 2 Days Before News # HFs Short Bonds | -109.90*** | -21.71*** |
| Other Days | -37.90* | 163.58*** |
| *N* | 1044 | 1044 |

Note: The table presents regression estimates that show the total net order flow (column 1) and number of trades (column 2) of hedge funds in pre-news windows that coincide with primary issuance. The rows show trades that occurred 1 or 2 days before the news, split by whether hedge funds had a long or short position in bonds on the day of the news. The long or short position is measured based on cumulative signed order flow in the five days leading up to the news. Panel A (B) shows the results for monetary policy announcements (labour market data releases). The asterisks indicate whether the returns are different from zero at 1%, 5% and 10% significance levels.

## *A.4 Additional Tables*

Table A.15: Scheduled MPC days

| year | Jan | Feb | Mar | Apr | May | Jun | Jul | Aug | Sep | Oct | Nov | Dec |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| 1997 | - | - | - | - | - | 06-Jun | 10-Jul | 07-Aug | 11-Sep | 09-Oct | 06-Nov | 04-Dec |
| 1998 | 08-Jan | 05-Feb | 05-Mar | 09-Apr | 07-May | 04-Jun | 09-Jul | 06-Aug | 10-Sep | 08-Oct | 05-Nov | 10-Dec |
| 1999 | 07-Jan | 04-Feb | 03-Mar | 08-Apr | 06-May | 10-Jun | 08-Jul | 05-Aug | 08-Sep | 07-Oct | 04-Nov | 09-Dec |
| 2000 | 13-Jan | 10-Feb | 09-Mar | 06-Apr | 04-May | 07-Jun | 06-Jul | 03-Aug | 07-Sep | 05-Oct | 09-Nov | 07-Dec |
| 2001 | 11-Jan | 08-Feb | 08-Mar | 05-Apr | 10-May | 06-Jun | 05-Jul | 02-Aug | 06-Sep | 04-Oct | 08-Nov | 05-Dec |
| 2002 | 10-Jan | 07-Feb | 07-Mar | 04-Apr | 09-May | 06-Jun | 04-Jul | 01-Aug | 05-Sep | 10-Oct | 07-Nov | 05-Dec |
| 2003 | 09-Jan | 06-Feb | 06-Mar | 10-Apr | 08-May | 05-Jun | 10-Jul | 07-Aug | 04-Sep | 09-Oct | 06-Nov | 04-Dec |
| 2004 | 08-Jan | 05-Feb | 04-Mar | 08-Apr | 06-May | 10-Jun | 08-Jul | 05-Aug | 09-Sep | 07-Oct | 04-Nov | 09-Dec |
| 2005 | 13-Jan | 10-Feb | 10-Mar | 07-Apr | 09-May | 09-Jun | 07-Jul | 04-Aug | 08-Sep | 06-Oct | 10-Nov | 08-Dec |
| 2006 | 12-Jan | 09-Feb | 09-Mar | 06-Apr | 04-May | 08-Jun | 06-Jul | 03-Aug | 07-Sep | 05-Oct | 09-Nov | 07-Dec |
| 2007 | 11-Jan | 08-Feb | 08-Mar | 05-Apr | 10-May | 07-Jun | 05-Jul | 02-Aug | 06-Sep | 04-Oct | 08-Nov | 06-Dec |
| 2008 | 10-Jan | 07-Feb | 06-Mar | 10-Apr | 08-May | 05-Jun | 10-Jul | 07-Aug | 04-Sep | 08-Oct | 06-Nov | 04-Dec |
| 2009 | 08-Jan | 05-Feb | 05-Mar | 09-Apr | 07-May | 04-Jun | 09-Jul | 06-Aug | 10-Sep | 08-Oct | 05-Nov | 10-Dec |
| 2010 | 07-Jan | 04-Feb | 04-Mar | 08-Apr | 10-May | 10-Jun | 08-Jul | 05-Aug | 09-Sep | 07-Oct | 04-Nov | 09-Dec |
| 2011 | 13-Jan | 10-Feb | 10-Mar | 07-Apr | 05-May | 09-Jun | 07-Jul | 04-Aug | 08-Sep | 06-Oct | 10-Nov | 08-Dec |
| 2012 | 12-Jan | 09-Feb | 08-Mar | 05-Apr | 10-May | 07-Jun | 05-Jul | 02-Aug | 06-Sep | 04-Oct | 08-Nov | 06-Dec |
| 2013 | 10-Jan | 07-Feb | 07-Mar | 04-Apr | 09-May | 06-Jun | 04-Jul | 01-Aug | 05-Sep | 10-Oct | 07-Nov | 05-Dec |
| 2014 | 09-Jan | 06-Feb | 06-Mar | 10-Apr | 08-May | 05-Jun | 10-Jul | 07-Aug | 04-Sep | 09-Oct | 06-Nov | 04-Dec |
| 2015 | 08-Jan | 05-Feb | 05-Mar | 09-Apr | 11-May | 04-Jun | 09-Jul | 06-Aug | 10-Sep | 08-Oct | 05-Nov | 10-Dec |
| 2016 | 14-Jan | 04-Feb | 17-Mar | 14-Apr | 12-May | 16-Jun | 14-Jul | 04-Aug | 15-Sep | . | 03-Nov | 15-Dec |
| 2017 | . | 02-Feb | 16-Mar | . | 11-May | 15-Jun | . | 03-Aug | 14-Sep | . | 02-Nov | 14-Dec |
| 2018 | . | 08-Feb | 22-Mar | . | 10-May | 21-Jun | . | 02-Aug | 13-Sep | . | 01-Nov | 20-Dec |
| 2019 | . | 07-Feb | 21-Mar | . | 02-May | 20-Jun | . | 01-Aug | 19-Sep | . | 07-Nov | 19-Dec |
| 2020 | 30-Jan | . | 26-Mar | . | 07-May | 18-Jun | . | 06-Aug | 17-Sep | . | 05-Nov | 17-Dec |
| 2021 | . | 04-Feb | 18-Mar | . | 06-May | 24-Jun | - | - | - | - | - | - |

Note: the table shows the dates of scheduled MPC meetings from May 1997 to July 2021. These dates represent the days when monetary policy actions or non-actions after scheduled meetings became known to the public.

## Table A.16: Scheduled Labour Market Data Release Days

| year | Jan | Feb | Mar | Apr | May | Jun | Jul | Aug | Sep | Oct | Nov | Dec |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| 1998 | - | - | - | 22-Apr | 13-May | 17-Jun | 15-Jul | 12-Aug | 16-Sep | 16-Oct | 11-Nov | 16-Dec |
| 1999 | 13-Jan | 17-Feb | 17-Mar | 21-Apr | 19-May | 16-Jun | 14-Jul | 11-Aug | 15-Sep | 13-Oct | 17-Nov | 15-Dec |
| 2000 | 19-Jan | 16-Feb | 15-Mar | 19-Apr | 17-May | 14-Jun | 12-Jul | 16-Aug | 13-Sep | 18-Oct | 15-Nov | 13-Dec |
| 2001 | 17-Jan | 14-Feb | 14-Mar | 11-Apr | 16-May | 13-Jun | 18-Jul | 15-Aug | 12-Sep | 17-Oct | 14-Nov | 12-Dec |
| 2002 | 16-Jan | 13-Feb | 20-Mar | 17-Apr | 15-May | 14-Jun | 17-Jul | 14-Aug | 11-Sep | 16-Oct | 13-Nov | 18-Dec |
| 2003 | 15-Jan | 12-Feb | 19-Mar | 16-Apr | 14-May | 11-Jun | 16-Jul | 13-Aug | 17-Sep | 15-Oct | 12-Nov | 17-Dec |
| 2004 | 14-Jan | 11-Feb | 17-Mar | 16-Apr | 12-May | 16-Jun | 14-Jul | 11-Aug | 15-Sep | 13-Oct | 17-Nov | 15-Dec |
| 2005 | 19-Jan | 16-Feb | 16-Mar | 13-Apr | 18-May | 15-Jun | 13-Jul | 17-Aug | 14-Sep | 12-Oct | 16-Nov | 14-Dec |
| 2006 | 18-Jan | 15-Feb | 15-Mar | 12-Apr | 17-May | 14-Jun | 12-Jul | 16-Aug | 13-Sep | 18-Oct | 15-Nov | 13-Dec |
| 2007 | 17-Jan | 14-Feb | 14-Mar | 18-Apr | 16-May | 13-Jun | 18-Jul | 15-Aug | 12-Sep | 17-Oct | 14-Nov | 12-Dec |
| 2008 | 16-Jan | 13-Feb | 19-Mar | 16-Apr | 11-May | 11-Jun | 16-Jul | 13-Aug | 17-Sep | 15-Oct | 12-Nov | 17-Dec |
| 2009 | 21-Jan | 11-Feb | 18-Mar | 22-Apr | 12-May | 17-Jun | 15-Jul | 12-Aug | 16-Sep | 14-Oct | 11-Nov | 16-Dec |
| 2010 | 20-Jan | 17-Feb | 17-Mar | 21-Apr | 12-May | 16-Jun | 14-Jul | 18-Aug | 15-Sep | 13-Oct | 17-Nov | 15-Dec |
| 2011 | 19-Jan | 16-Feb | 16-Mar | 13-Apr | 18-May | 15-Jun | 13-Jul | 17-Aug | 14-Sep | 12-Oct | 16-Nov | 14-Dec |
| 2012 | 18-Jan | 15-Feb | 14-Mar | 18-Apr | 16-May | 20-Jun | 18-Jul | 15-Aug | 12-Sep | 17-Oct | 14-Nov | 12-Dec |
| 2013 | 23-Jan | 20-Feb | 20-Mar | 17-Apr | 15-May | 12-Jun | 17-Jul | 14-Aug | 11-Sep | 16-Oct | 13-Nov | 18-Dec |
| 2014 | 22-Jan | 19-Feb | 19-Mar | 16-Apr | 14-May | 11-Jun | 16-Jul | 13-Aug | 17-Sep | 15-Oct | 12-Nov | 17-Dec |
| 2015 | 21-Jan | 18-Feb | 18-Mar | 17-Apr | 13-May | 17-Jun | 15-Jul | 12-Aug | 16-Sep | 14-Oct | 11-Nov | 16-Dec |
| 2016 | 20-Jan | 17-Feb | 16-Mar | 20-Apr | 18-May | 15-Jun | 20-Jul | 17-Aug | 14-Sep | 19-Oct | 16-Nov | 14-Dec |
| 2017 | 18-Jan | 15-Feb | 15-Mar | 12-Apr | 17-May | 14-Jun | 12-Jul | 16-Aug | 13-Sep | 18-Oct | 15-Nov | 13-Dec |
| 2018 | 24-Jan | 21-Feb | 21-Mar | 17-Apr | 15-May | 12-Jun | 17-Jul | 14-Aug | 11-Sep | 16-Oct | 13-Nov | 11-Dec |
| 2019 | 22-Jan | 19-Feb | 19-Mar | 16-Apr | 14-May | 11-Jun | 16-Jul | 13-Aug | 10-Sep | 15-Oct | 12-Nov | 17-Dec |
| 2020 | 21-Jan | 18-Feb | 17-Mar | 21-Apr | 19-May | 16-Jun | 16-Jul | 11-Aug | 15-Sep | 13-Oct | 10-Nov | 15-Dec |
| 2021 | 26-Jan | 23-Feb | 23-Mar | 20-Apr | 18-May | 15-Jun | 17-Jul | - | - | - | - | - |

Note: the table shows the dates of scheduled labour market data release days from April 1998 to July 2021.

Table A.17: Yield Changes before MPC Meetings: Adding Year Fixed Effects

| | $\Delta 5y$ yield<br>(1) | $\Delta 10y$ yield<br>(2) | $\Delta 15y$ yield<br>(3) | $\Delta 20y$ yield<br>(4) |
|---|---|---|---|---|
| (A) 10Y Yield Changes before MPC Meetings | | | | |
| Pre-News window | 0.29 | 0.46** | 0.50** | 0.51** |
| | (1.45) | (2.21) | (2.51) | (2.57) |
| Year fixed effects | YES | YES | YES | YES |
| *N* | 6128 | 6128 | 6128 | 6128 |
| (B) 10Y Yield Changes before Labour Market Data Release | | | | |
| Pre-News window | 0.50** | 0.48** | 0.43** | 0.39** |
| | (2.45) | (2.34) | (2.22) | (2.06) |
| Year fixed effects | YES | YES | YES | YES |
| *N* | 5896 | 5896 | 5896 | 5896 |
| (B) 10Y Yield Changes before MPC Meetings and during Bond Issuance | | | | |
| Pre-News window # Issuance | 0.20 | 0.54* | 0.64** | 0.67** |
| | (0.75) | (1.88) | (2.33) | (2.46) |
| Pre-News window # No issuance | 0.47 | 0.42 | 0.38 | 0.37 |
| | (1.61) | (1.42) | (1.36) | (1.31) |
| No News # Issuance | 0.23 | 0.20 | 0.19 | 0.17 |
| | (1.27) | (1.03) | (1.02) | (0.96) |
| Year fixed effects | YES | YES | YES | YES |
| *N* | 6128 | 6128 | 6128 | 6128 |
| (C) 10Y Yield Changes before MPC Meetings and during Large Bond Issuance | | | | |
| Pre-News window # Issuance | 0.99* | 1.08** | 1.01** | 1.05** |
| | (1.79) | (2.05) | (2.09) | (2.26) |
| Pre-News window # No issuance | 0.47** | 0.43** | 0.40* | 0.35* |
| | (2.13) | (1.97) | (1.90) | (1.68) |
| No News # Issuance | 0.23 | 0.27 | 0.30* | 0.31* |
| | (1.37) | (1.45) | (1.70) | (1.79) |
| Year fixed effects | YES | YES | YES | YES |
| *N* | 5896 | 5896 | 5896 | 5896 |

Note: Panel A of this table regresses daily changes in the 5-year, 10-year, 15-year and 20-year yields on UK nominal government bonds on an indicator variable that takes value one for days that are either one or two days before MPC days. Panel B regresses daily changes in yields on indicator variables capturing (i) pre-MPC windows with new (nominal or inflation-linked) bond issuance, (ii) pre-MPC windows without new bond issuance and (iii) all trading days with issuance and without MPC meetings. Panel C regresses daily changes in yields on indicator variables capturing (i) pre-MPC windows with new large (nominal or inflation-linked) bond issuance (i.e. issuance larger than the median issuance in the given year), (ii) pre-MPC windows without new large bond issuance and (iii) all trading days with large issuances and without MPC meetings. All regressions include a constant and year fixed effects. The estimation period covers 1997m5-2021m7 and includes 270 MPC announcement windows.

Table A.18: Yield Changes before MPC Meetings: Adding Unscheduled MPC Meetings

| (A) Yield Changes before MPC Meetings | | | | |
|---|---|---|---|---|
| | $\Delta 5y$ yield (1) | $\Delta 10y$ yield (2) | $\Delta 15y$ yield (3) | $\Delta 20y$ yield (4) |
| pre-MPC window | 0.32 | 0.51** | 0.56*** | 0.57*** |
| | (1.62) | (2.43) | (2.72) | (2.76) |
| Constant | -0.14** | -0.16** | -0.16** | -0.16*** |
| | (-2.25) | (-2.39) | (-2.56) | (-2.62) |
| $N$ | 6128 | 6128 | 6128 | 6128 |
| (B) Yield Changes before MPC Meetings and during Bond Issuance | | | | |
| Pre-MPC window # Issuance | 0.28 | 0.66** | 0.76*** | 0.77*** |
| | (1.05) | (2.23) | (2.61) | (2.67) |
| Pre-MPC window # No issuance | 0.44 | 0.40 | 0.37 | 0.37 |
| | (1.56) | (1.39) | (1.36) | (1.33) |
| Outside MPC window # Issuance | 0.22 | 0.19 | 0.18 | 0.16 |
| | (1.26) | (1.03) | (1.00) | (0.94) |
| Constant | -0.17** | -0.18*** | -0.18*** | -0.18*** |
| | (-2.54) | (-2.62) | (-2.76) | (-2.79) |
| $N$ | 6128 | 6128 | 6128 | 6128 |

Note: Panel A of this table regresses daily changes in the 5-year, 10-year, 15-year and 20-year yields on UK nominal government bonds on an indicator variable that takes value one for days that are either one or two days before MPC days. Panel B regresses daily changes in yields on indicator variables capturing (i) pre-MPC windows with new (nominal or inflation-linked) bond issuance, (ii) pre-MPC windows without new bond issuance and (iii) all trading days with issuance and without MPC meetings. Panel C regresses daily changes in yields on indicator variables capturing (i) pre-MPC windows with new (nominal or inflation-linked) large bond issuance (i.e. issuance larger than the median issuance in the given year), (ii) pre-MPC windows without new large bond issuance and (iii) all trading days with large issuances and without MPC meetings. All regressions include a constant. The estimation period covers 1997m5-2021m7 and includes 273 MPC announcement windows (270 scheduled and 3 unscheduled announcements).

## Table A.19: Yield Volatility on Scheduled Announcement Days

| | 2Y<br>(1) | 5Y<br>(2) | 7Y<br>(3) | 10Y<br>(4) | 20Y<br>(5) |
|---|---|---|---|---|---|
| (A) Monetary Policy Announcements | | | | | |
| Absolute Value of | 0.93*** | 0.90*** | 0.84*** | 0.74*** | 0.51** |
| Yield Changes | (4.14) | (4.09) | (3.73) | (3.20) | (2.14) |
| Constant | 3.07*** | 3.42*** | 3.58*** | 3.61*** | 3.29*** |
| | (78.71) | (85.29) | (87.00) | (87.35) | (84.08) |
| *N* | 6128 | 6128 | 6128 | 6128 | 6128 |
| (B) Labour Market Data Releases | | | | | |
| Absolute Value of | 0.62*** | 0.43** | 0.17 | 0.02 | -0.03 |
| Yield Changes | (2.68) | (1.97) | (0.80) | (0.11) | (-0.19) |
| Constant | 3.06*** | 3.42*** | 3.60*** | 3.63*** | 3.30*** |
| | (77.19) | (84.30) | (85.96) | (86.07) | (82.40) |
| *N* | 5896 | 5896 | 5896 | 5896 | 5896 |

Note: This table regresses the absolute daily changes in yields on an indicator variable taking value 1 if the trading day coincides with a scheduled monetary policy announcement (Panel A) or a labour market data release (Panel B). T-statistics in parentheses are based on robust standard errors. Asterisks denote significance levels (* p<0.1, ** p<0.05, *** p<0.01).

## Table A.20: P-Values Corresponding to Testing for Equality of Coefficients

| | 5Y<br>(1) | 10Y<br>(2) | 15Y<br>(3) | 20Y<br>(4) |
|---|---|---|---|---|
| (A) Pre-Labour News Drift # Issuance | | | | |
| All days | 0.333 | 0.196 | 0.192 | 0.134 |
| High Noise Days | 0.242 | 0.172 | 0.094 | 0.041 |
| Low Noise Days | 0.806 | 0.709 | 0.924 | 0.953 |
| High ID Price Disp. Days | 0.085 | 0.213 | 0.325 | 0.372 |
| Low ID Price Disp. Days | 0.758 | 0.897 | 0.961 | 0.863 |
| (B) Pre-MPC Drift # Issuance | | | | |
| All days | 0.634 | 0.537 | 0.331 | 0.312 |
| High Noise Days | 0.416 | 0.133 | 0.123 | 0.137 |
| Low Noise Days | 0.895 | 0.806 | 0.953 | 0.837 |
| High ID Price Disp. Days | 0.115 | 0.121 | 0.120 | 0.085 |
| Low ID Price Disp. Days | 0.922 | 0.945 | 0.988 | 0.883 |

Notes: The table presents the p-values associated with testing whether the coefficients on the interaction between pre-news windows with and without issuances are equal. Panel (A) and Panel (B) show the results corresponding to labour market data releases and to pre-MPC windows, respectively. In each panel, we also tabulate the p-values separately for days with lower and higher market liquidity. As explained in section 4 of the revised manuscript, liquidity is proxied by gilt market noise (Hu, Pan, and Wang 2013) as well as by inter-dealer price dispersion (Eisfeldt, Herskovic, and Liu, 2023).

# Table A.21: Pre-MPC Drift and Bond Issuance: High vs Low Price Dispersion Months

| | 5Y | 10Y | 15Y | 20Y | 5Y | 10Y | 15Y | 20Y |
|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| Pre-News # Issue # Lower Dispersion | 0.10 | 0.12 | 0.13 | 0.20 | -0.21 | -0.15 | -0.13 | -0.07 |
| | (0.20) | (0.21) | (0.25) | (0.39) | (-0.42) | (-0.25) | (-0.24) | (-0.12) |
| Pre-News # Issue # Higher Dispersion | 1.30** | 1.44* | 1.39* | 1.29* | 1.00* | 1.19 | 1.13 | 1.03 |
| | (2.35) | (1.95) | (1.81) | (1.69) | (1.76) | (1.57) | (1.44) | (1.31) |
| Pre-News # No issue # Lower Dispersion | 0.01 | 0.05 | 0.15 | 0.33 | -0.30 | -0.21 | -0.12 | 0.06 |
| | (0.02) | (0.06) | (0.20) | (0.45) | (-0.39) | (-0.25) | (-0.16) | (0.08) |
| Pre-News # No issue # Higher Dispersion | -0.31 | -0.53 | -0.63 | -0.97 | -0.58 | -0.75 | -0.86 | -1.19 |
| | (-0.36) | (-0.50) | (-0.59) | (-0.90) | (-0.65) | (-0.71) | (-0.80) | (-1.10) |
| No News # Issue # Lower Dispersion | -0.19 | -0.38 | -0.36 | -0.30 | -0.40 | -0.57 | -0.55 | -0.50 |
| | (-0.67) | (-1.10) | (-1.08) | (-0.95) | (-1.27) | (-1.54) | (-1.54) | (-1.46) |
| No News # Issue # Higher Dispersion | -0.34 | -0.39 | -0.30 | -0.22 | -0.55 | -0.58 | -0.50 | -0.41 |
| | (-1.09) | (-1.09) | (-0.89) | (-0.67) | (-1.63) | (-1.51) | (-1.35) | (-1.17) |
| Weekday FE | NO | NO | NO | NO | YES | YES | YES | YES |
| *N* | 2255 | 2255 | 2255 | 2255 | 2255 | 2255 | 2255 | 2255 |

Note: This table regresses daily changes in yields on indicator variables capturing high vs low inter-dealer price dispersion (Eisfeldt, Herskovic, and Liu, 2023) periods interacted with indicator variables capturing (i) pre-news windows (associated with monetary policy announcements) and new bond issuance, (ii) pre-news windows without new bond issuance and (iii) all trading days with issuance and without announcements. Columns 1-4 (5-8) present the results without (without) weekday fixed effects. We define high (low) dispersion periods as those when dispersion in a given month is above (below) the median value in the given year. The estimation period covers 2011m9-2021m7, which is dictated by the availability of the transaction-level data. T-statistics in parentheses are based on robust standard errors. Asterisks denote significance levels (* p<0.1, ** p<0.05, *** p<0.01).

## Table A.22: Pre-Labour News Drift and Bond Issuance: High vs Low Price Dispersion

| | 5Y | 10Y | 15Y | 20Y | 5Y | 10Y | 15Y | 20Y |
|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| Pre-News # Issue # Lower Dispersion | 0.32 | 0.55 | 0.79 | 0.92 | 0.34 | 0.65 | 0.87 | 0.98 |
| | (0.48) | (0.69) | (0.99) | (1.14) | (0.48) | (0.78) | (1.04) | (1.18) |
| Pre-News # Issue # Higher Dispersion | 2.05*** | 1.92** | 1.55* | 1.35* | 2.10*** | 2.06** | 1.66* | 1.43* |
| | (2.69) | (2.16) | (1.83) | (1.67) | (2.72) | (2.26) | (1.92) | (1.74) |
| Pre-News # No issue # Lower Dispersion | 0.41 | 0.50 | 0.58 | 0.61* | 0.55 | 0.72* | 0.79* | 0.79** |
| | (1.09) | (1.19) | (1.49) | (1.67) | (1.45) | (1.68) | (1.94) | (2.11) |
| Pre-News # No issue # Higher Dispersion | 0.43 | 0.56 | 0.49 | 0.40 | 0.57 | 0.78 | 0.69 | 0.58 |
| | (0.90) | (1.00) | (0.86) | (0.68) | (1.15) | (1.35) | (1.17) | (0.95) |
| No News # Issue # Lower Dispersion | -0.27 | -0.46 | -0.43 | -0.35 | -0.49 | -0.65* | -0.63* | -0.55* |
| | (-0.94) | (-1.36) | (-1.34) | (-1.13) | (-1.63) | (-1.83) | (-1.83) | (-1.66) |
| No News # Issue # Higher Dispersion | -0.32 | -0.31 | -0.20 | -0.12 | -0.54* | -0.51 | -0.40 | -0.32 |
| | (-1.06) | (-0.88) | (-0.61) | (-0.39) | (-1.71) | (-1.38) | (-1.15) | (-0.95) |
| Weekday FE | NO | NO | NO | NO | YES | YES | YES | YES |
| $N$ | 2255 | 2255 | 2255 | 2255 | 2255 | 2255 | 2255 | 2255 |

Note: This table regresses daily changes in yields on indicator variables capturing high vs low inter-dealer price dispersion (Eisfeldt, Herskovic, and Liu, 2023) periods interacted with indicator variables capturing (i) pre-news windows (associated with labour market data releases) and new bond issuance, (ii) pre-news windows without new bond issuance and (iii) all trading days with issuance and without announcements. Columns 1-4 (5-8) present the results without (without) weekday fixed effects. We define high (low) dispersion periods as those when dispersion in a given month is above (below) the median value in the given year. The estimation period covers 2011m9-2021m7, which is dictated by the availability of the transaction-level data. T-statistics in parentheses are based on robust standard errors. Asterisks denote significance levels (* p<0.1, ** p<0.05, *** p<0.01).

## Table A.23: The Relationship between Pre-MPC Bond Issuance and the Volatility of Monetary Policy Surprises

| | (1) | (2) | (3) |
|---|---|---|---|
| | Target | Path | QE |
| Pre-MPC Issuance | 0.47** | 0.28 | 0.59 |
| | (2.03) | (0.85) | (1.13) |
| $N$ | 132 | 132 | 132 |
| $R^2$ | 0.219 | 0.229 | 0.218 |

Note: This table presents the estimation results for regression 3.3. The regressands are taken from Braun, Miranda-Agrippino, and Saha (2022) and measure price changes in a 30-minute window around the interest rate announcement of the Monetary Policy Committee of the Bank of England for three different shock measures, following Swanson (2021): shocks related to the central bank target (1), forward guidance (2) and quantitative easing (3). T-statistics in parentheses are based on robust standard errors. Asterisks denote significance levels (* p<0.1, ** p<0.05, *** p<0.01).

## *A.5 Theoretical Appendix*

### *A.5.1 Proof of Proposition 1*

We start with the problem of small, price-taking agents at $t=2$. Given $\tilde{P}_2$ and the inferred signal $\tilde{\theta}+\Phi\tilde{d}\equiv\tilde{\theta}+\frac{\beta_d}{\beta_\theta}\tilde{d}$, agent $i$'s optimal post-trade holding at $t=2$, $D_2^i\left(D_1^i\right)$, solves

$$\max_{D_2^i}\mathbb{E}\left[-e^{-\gamma_i\left\{D_2^i\tilde{v}-\left(D_2^i-D_1^i\right)\tilde{P}_2\right\}}\mid\tilde{P}_2\right]$$

for $i\in\{D,UC\}$. Then, using Bayesian updating for normally distributed random variables and employing the "improper prior trick" (i.e., taking limit as $\sigma_\theta\rightarrow\infty$), the FOC implies that the solution for $i\in\{D,UC\}$,

$$D_2^i\left(D_1^i\right)=\frac{\tilde{\theta}+\Phi\tilde{d}-\tilde{P}_2}{\gamma_i\left(\sigma_\varepsilon^2+\Phi^2\sigma_d^2\right)},\tag{A.12}$$

is actually independent of $D_1^i$, which implies the lack of endowment effects—a standard result for price-taking agents with CARA utility. Hedge funds are, however, large, price-making agents. Hence, hedge fund $i\in\{1,2,...,N\}$ makes its choice of $D_2^i$ by keeping in mind that its choice will have a material impact on the market price $P_2$. This means that hedge fund $i$ uses the following market-clearing condition

$$D_2^i+D_2^{-i}+\sum_{j\in\{D,UC\}}\frac{\tilde{\theta}+\Phi\tilde{d}-P_2}{\gamma_j\left(\sigma_\varepsilon^2+\Phi^2\sigma_d^2\right)}+\tilde{d}=z$$

to figure out its price impact $\frac{\partial P_2}{\partial D_2^i}$. Using the standard linear price impact assumption, $P_2\equiv\hat{P}_2^0+\hat{P}_2^1\left(D_2^{-i}+D_2^i\right)$, the market-clearing condition becomes

$$D_2^i+D_2^{-i}+\frac{\tilde{\theta}+\Phi\tilde{d}-\hat{P}_2^0-\hat{P}_2^1\left(D_2^{-i}+D_2^i\right)}{\bar{\gamma}\left(\sigma_\varepsilon^2+\Phi^2\sigma_d^2\right)}+\tilde{d}=z,$$

where $\bar{\gamma}\equiv\left(\frac{1}{\gamma_D}+\frac{1}{\gamma_{UC}}\right)^{-1}$ is the harmonic sum of the dealers' and the uninformed clients' risk aversion parameters. Because this condition must hold for any $D_2^{-i}+D_2^i$, the implied pricing coefficients are

$$\hat{P}_2^0=\tilde{\theta}+\left[\Phi+\bar{\gamma}\left(\sigma_\varepsilon^2+\Phi^2\sigma_d^2\right)\right]\tilde{d}-\bar{\gamma}\left(\sigma_\varepsilon^2+\Phi^2\sigma_d^2\right)z\text{ and }\hat{P}_2^1=\bar{\gamma}\left(\sigma_\varepsilon^2+\Phi^2\sigma_d^2\right).\tag{A.13}$$

This implies that hedge fund $i$'s price impact per share held is $\frac{\partial P_2}{\partial D_2^i}=\hat{P}_2^1=\bar{\gamma}\left(\sigma_\varepsilon^2+\Phi^2\sigma_d^2\right)$: if the small players are more risk averse or if there is more uncertainty regarding the upcoming macroeconomic shock (either because of information asymmetry or noise trading risk), hedge funds have a larger impact on the price.

Given (A.13), hedge fund $i$'s problem is to choose $D_2^i\left(\tilde{\theta}, \tilde{d}, D_1^i, D_2^{-i}\right)$ that solves

$$\max_{D_2^i} \mathbb{E}\left[D_2^i\left(\tilde{\theta}+\tilde{\varepsilon}\right)-\left(D_2^i-D_1^i\right)\left\{\hat{P}_2^0+\hat{P}_2^1\left(D_2^i+D_2^{-i}\right)\right\} \mid \tilde{\theta}, \tilde{d}\right]. \tag{A.14}$$

Then, the FOC implies

$$D_2^i\left(\tilde{\theta}, \tilde{d}, D_1^i, D_2^{-i}\right)=\frac{D_1^{HF}}{2}-\frac{D_2^{-i}}{2}+\frac{\tilde{\theta}-\hat{P}_2^0\left(\tilde{\theta}, \tilde{d}\right)}{2\hat{P}_2^1}, \tag{A.15}$$

where we emphasize the dependence of $\hat{P}_2^0$ on $\tilde{\theta}$ and $\tilde{d}$ as shown in (A.13). The first term on the RHS of (A.15) is the familiar endowment effect term that arises in the linear price impact models. Being endowed with more shares from earlier trading rounds effectively reduces the marginal cost of holding more shares in the current round, which feeds positively back to the optimal post-trade asset demand of a large player.[42] The second term is the strategic demand reduction term typical of models with imperfect competition. That is, if other hedge funds demand to hold a larger share of the asset, their resulting price impact makes purchasing the asset more costly, which reduces hedge fund $i$'s demand.

So far, equations (A.12), (A.13), and (A.15) characterize the equilibrium outcomes at $t=2$ up to an endogenous object $\Phi$. Then, the rational expectations condition, $P_2=\tilde{P}_2$, that the market-clearing price and the information-revealing price are equal to each other in equilibrium pins down $\Phi$:

$$\Phi=\frac{N \pm \sqrt{N^2-4\bar{\gamma}^2\sigma_{\varepsilon}^2\sigma_d^2}}{2\bar{\gamma}\sigma_d^2}. \tag{A.16}$$

Although there are two roots of the equation that pins down $\Phi$, we choose the one with the negative sign because it is the "economically reasonable" root. As the number of informed agents, hedge funds, approaches infinity, the root we choose approaches zero, while the other one approaches infinity. Because $\Phi$ measures informational inefficiency, it is reasonable that $\Phi$ approaches zero as the number of atomic informed agents approaches infinity.

Then, the equilibrium price and the equilibrium post-trade holding of small players are

$$P_2\left(\tilde{\theta}, \tilde{d}, D_1^{HF}\right)=\tilde{\theta}+\Phi\tilde{d}+\frac{N}{N+1}\Phi\left(-z+ND_1^{HF}\right) \tag{A.17}$$

---

[42] More precisely, that stems from the total price impact term which obtains when one takes the derivative of (A.14) with respect to the last $D_2^i$ in the application of the chain rule: $-\left(D_2^i-D_1^i\right)\hat{P}_2^1$.

and

$$D_2^j\left(D_1^{HF}\right) = \frac{\bar{\gamma}}{\gamma_j}\frac{1}{N+1}\left(z - ND_1^{HF}\right), \tag{A.18}$$

for $j \in \{D, UC\}$, respectively. In the first trading round, which is at $t = 1$, agents will make their decisions by anticipating (A.15), (A.17), and (A.18).

Now we are ready to move to the next step in our backward induction. As in the previous round, we start with the problem of price-taking agents and the resulting market-clearing condition. Then, hedge funds, the price-making agents of our model, figure out their price impact and makes their decisions accordingly. Given $P_1$ and the inferred $D_1^{HF}$, agent $j$'s optimal post-trade holding at $t = 1$, $D_1^j$, solves

$$\max_{D_1^j} \mathbb{E}\left[-e^{-\gamma_i\left\{D_2^j\left(D_1^{HF}\right)\tilde{v} - \left[D_2^j\left(D_1^{HF}\right) - D_1^j\right]P_2\left(\tilde{\theta},\tilde{d},D_1^{HF}\right) - \left(D_1^j - \mathbb{I}_{\{j=D\}}z\right)P_1\right\}}\right],$$

for $j \in \{D, UC\}$. The FOC for this problem is

$$D_1^j = \frac{\mathbb{E}\left[P_2\left(\tilde{\theta},\tilde{d},D_1^{HF}\right)\right] - P_1}{\gamma_j\left(\sigma_\theta^2 + \Phi^2\sigma_d^2\right)} + D_2^j\left(D_1^{HF}\right). \tag{A.19}$$

This FOC implies that dealers and uninformed clients would carry different shares of the asset to the second round only if $\gamma_D \neq \gamma_{UC}$. That dealers are endowed with $z$ shares and uninformed clients 0 at the beginning of trade at $t = 1$ has nothing to do with their optimal asset demand—again, thanks to the lack of endowment effect for price-taking agents with CARA utility. Substituting (A.17) and (A.18) into (A.19),

$$D_1^j = \frac{\theta - P_1}{\gamma_j\left(\sigma_\theta^2 + \Phi^2\sigma_d^2\right)} + \frac{\bar{\gamma}}{\gamma_j}\frac{ND_1^{HF} - z}{N+1}\frac{\sigma_\varepsilon^2 - \sigma_\theta^2}{\sigma_\theta^2 + \Phi^2\sigma_d^2} \tag{A.20}$$

for $j \in \{D, UC\}$. Hedge fund $i$, then, will make its own trading decision at $t = 1$ by taking as given the following market-clearing condition:

$$D_1^i + D_1^{-i} + \sum_{j\in\{D,UC\}}\left\{\frac{\theta - P_1}{\gamma_j\left(\sigma_\theta^2 + \Phi^2\sigma_d^2\right)} + \frac{\bar{\gamma}}{\gamma_j}\frac{D_1^i + D_1^{-i} - z}{N+1}\frac{\sigma_\varepsilon^2 - \sigma_\theta^2}{\sigma_\theta^2 + \Phi^2\sigma_d^2}\right\} = z,$$

or equivalently,

$$D_1^i + D_1^{-i} + \frac{\theta - P_1}{\bar{\gamma}\left(\sigma_\theta^2 + \Phi^2\sigma_d^2\right)} + \frac{D_1^i + D_1^{-i} - z}{N+1}\frac{\sigma_\varepsilon^2 - \sigma_\theta^2}{\sigma_\theta^2 + \Phi^2\sigma_d^2} = z.$$

Because this must hold for any $D_1^i + D_1^{-i}$, the equilibrium pricing function is $P_1 \equiv \hat{P}_1^0 +$

$\hat{P}_1^1\left(D_1^i + D_1^{-i}\right)$ with

$$\hat{P}_1^0 = \theta - z\bar{\gamma}\left(\frac{N\sigma_\theta^2 + \sigma_\varepsilon^2}{N+1} + \Phi^2\sigma_d^2\right) \text{ and } \hat{P}_1^1 = \bar{\gamma}\left(\frac{N\sigma_\theta^2 + \sigma_\varepsilon^2}{N+1} + \Phi^2\sigma_d^2\right), \tag{A.21}$$

which implies that hedge funds' price impact per share held at $t = 1$ is $\frac{\partial P_1}{\partial D_1^i} = \hat{P}_1^1 = \bar{\gamma}\left(\frac{N\sigma_\theta^2+\sigma_\varepsilon^2}{N+1} + \Phi^2\sigma_d^2\right)$.

Given (A.13) and (A.21), hedge fund $i$ chooses $D_1^i$ that solves

$$\max_{D_1^i} \mathbb{E}\Big[D_2^i\left(D_1^i, D_1^{-i}\right)\tilde{v} - D_1^i\left\{\hat{P}_1^0 + \hat{P}_1^1\left(D_1^i + D_1^{-i}\right)\right\} \\ - \left(D_2^i\left(D_1^i, D_1^{-i}\right) - D_1^i\right)\left\{\hat{P}_2^0 + \hat{P}_2^1\left[D_2^i\left(D_1^i, D_1^{-i}\right) + D_2^{-i}\left(D_1^i, D_1^{-i}\right)\right]\right\}\Big].$$

Using (A.15) and (A.17) and after algebra, the FOC of hedge fund $i$'s problem is

$$D_1^i = \frac{\theta - \frac{\Phi}{(N+1)^2}\left[\left(N^2 + 2N - 1\right)z - \left(N^2 + N - 2\right)D_1^{-i}\right] - \left\{\hat{P}_1^0 + \hat{P}_1^1\left(D_1^i + D_1^{-i}\right)\right\}}{\frac{\Phi}{(N+1)^2}\left(N^3 - N + 2\right)},$$

which, in turn, implies that hedge fund $i$'s symmetric equilibrium post-trade holding at $t = 1$ is

$$D_1^i = \frac{z}{N}\left[1 - \frac{\Phi}{\frac{2\Phi}{(N+1)^2} + \bar{\gamma}\left(\frac{N\sigma_\theta^2+\sigma_\varepsilon^2}{N+1} + \Phi^2\sigma_d^2\right)}\right]. \tag{A.22}$$

Then, we substitute this into (A.21) and (A.20) to find the equilibrium price and the equilibrium post-trade holding of small players at $t = 1$. Finally, we calculate the limiting results as $\sigma_\theta$ approaches infinity to be consistent with our improper prior assumption at $t = 2$.

Let $q_t^i = D_t^i - D_{t-1}^i$ for $t \in \{1, 2\}$ and $i \in \{HF, D, UC\}$, where $D_0^{HF} = D_0^{UC} = 0$ and $D_0^D = z$. Then, $q_1^{HF}$ follows from (A.22); $q_1^D$ and $q_1^{UC}$ follow from (A.19); and $P_1$ is a re-statement of (A.21). Turning to the equilibrium objects at $t = 2$, $P_2$ follows from (A.17) and (A.22); $q_2^{HF}$ follows from (A.15), the equilibrium level of $D_1^{HF}$ given by (A.22), and the fact that hedge funds make symmetric decisions; and finally $q_2^D$ and $q_2^{UC}$ follow from (A.18) and (A.22).

### *A.5.2 Proof of Lemma 1*

For part (i), take the first derivative of $\Phi$ with respect to $N$:

$$\frac{\partial\Phi}{\partial N} = \frac{1 - N\left(N^2 - 4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2\right)^{-1/2}}{2\bar{\gamma}\sigma_d^2} < 0,$$

where the inequality follows because $N\left(N^2 - 4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2\right)^{-1/2} > 1$ when $N > 2\bar{\gamma}\sigma_\varepsilon\sigma_d$.

For part (ii), take the first derivative of $\Phi$ with respect to $\bar{\gamma}$:

$$\begin{aligned}\frac{\partial \Phi}{\partial \bar{\gamma}} &= \frac{4\bar{\gamma}\sigma_\varepsilon^2\sigma_d^2\left(N^2-4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2\right)^{-1/2}2\bar{\gamma}\sigma_d^2-\left(N-\sqrt{N^2-4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2}\right)2\sigma_d^2}{4\bar{\gamma}^2\sigma_d^4}\\
&= \frac{4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2\left(N^2-4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2\right)^{-1/2}-\left(N-\sqrt{N^2-4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2}\right)}{2\bar{\gamma}^2\sigma_d^2}\\
&= \frac{4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2-\left(N\sqrt{N^2-4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2}-N^2+4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2\right)}{2\bar{\gamma}^2\sigma_d^2\sqrt{N^2-4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2}}>0,\end{aligned}$$

where the inequality follows because $N\sqrt{N^2-4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2}-N^2<0$ when $N>2\bar{\gamma}\sigma_\varepsilon\sigma_d$.

For part (iii), take the first derivative of $\Phi/\bar{\gamma}$ with respect to $\bar{\gamma}$:

$$\begin{aligned}\frac{\partial \frac{\Phi}{\bar{\gamma}}}{\partial \bar{\gamma}} &= \frac{4\bar{\gamma}\sigma_\varepsilon^2\sigma_d^2\left(N^2-4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2\right)^{-1/2}2\bar{\gamma}^2\sigma_d^2-\left(N-\sqrt{N^2-4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2}\right)4\bar{\gamma}\sigma_d^2}{4\bar{\gamma}^4\sigma_d^4}\\
&= \frac{2\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2\left(N^2-4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2\right)^{-1/2}-\left(N-\sqrt{N^2-4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2}\right)}{\bar{\gamma}^3\sigma_d^2}\\
&= \frac{2\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2-\left(N\sqrt{N^2-4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2}-N^2+4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2\right)}{\bar{\gamma}^3\sigma_d^2\sqrt{N^2-4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2}}>0,\end{aligned}$$

where the inequality follows because $N\sqrt{N^2-4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2}-N^2+2\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2<0$ when $N>2\bar{\gamma}\sigma_\varepsilon\sigma_d$.

For part (iv), take the first derivative of $\Phi$ with respect to $\sigma_\varepsilon$:

$$\frac{\partial \Phi}{\partial \sigma_\varepsilon} = \frac{4\bar{\gamma}^2\sigma_\varepsilon\sigma_d^2\left(N^2-4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2\right)^{-1/2}}{2\bar{\gamma}\sigma_d^2}>0,$$

where the inequality follows because $\left(N^2-4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2\right)^{-1/2}>0$ when $N>2\bar{\gamma}\sigma_\varepsilon\sigma_d$.

For part (v), take the first derivative of $\Phi$ with respect to $\sigma_d$:

$$\begin{aligned}\frac{\partial \Phi}{\partial \sigma_d} &= \frac{4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d\left(N^2-4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2\right)^{-1/2}2\bar{\gamma}\sigma_d^2-\left(N-\sqrt{N^2-4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2}\right)4\bar{\gamma}\sigma_d}{4\bar{\gamma}^2\sigma_d^4}\\
&= \frac{2\bar{\gamma}\sigma_\varepsilon^2\sigma_d^2\left(N^2-4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2\right)^{-1/2}-\left(N-\sqrt{N^2-4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2}\right)}{\bar{\gamma}\sigma_d^3}\\
&= \frac{2\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2-\left(N\sqrt{N^2-4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2}-N^2+4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2\right)}{\bar{\gamma}\sigma_d^3\sqrt{N^2-4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2}}>0,\end{aligned}$$

where the inequality follows because $N\sqrt{N^2 - 4\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2} - N^2 + 2\bar{\gamma}^2\sigma_\varepsilon^2\sigma_d^2 < 0$ when $N > 2\bar{\gamma}\sigma_\varepsilon\sigma_d$.