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Abstract

How does the additional debt issued by the government affect the term structure of interest rates? In this paper, we examine the impact of unexpected government bond issuance by leveraging the institutional framework of the UK government bond primary market. Our findings indicate that higher-than-expected issuance exerts upward pressure on interest rates by increasing both real term premia and inflation risk premia. These effects are amplified during periods of market stress, with some evidence suggesting that yield reactions become more localized under such conditions.

Key words: Government debt, primary market, term structure, bond risk premia.

JEL classification: E43, E44, E60.

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1 Introduction

How does the additional issuance of government debt affect the term structure of interest rates? This question is both significant and timely in the post-COVID-19 context, where elevated government deficits and debt levels coincide with interest rates that have risen from the historically low levels of the 2010s. In this environment, higher servicing and refinancing costs pose significant challenges for debt management offices. Consequently, these authorities require a clear understanding of the financial market implications of their issuance decisions.

The effect of government debt issuance on interest rates is not well established in the empirical literature. Surveys by Gale and Orszag (2003) and Engen and Hubbard (2004) report roughly equal numbers of studies finding positive and significant effects and studies finding no effect. Identification is difficult. While interest rates are observed continuously, budgetary variables are available only at annual or, at best, quarterly frequency. Reverse causality, common factors, and anticipation effects further complicate inference. For example, agents anticipate and price fiscal policy in advance, obscuring its timing and causal impact. Moreover, factors affecting both interest rates and fiscal deficits can generate spurious correlations: during recessions, central banks cut policy rates while deficits and debt issuance rise. Our goal is to address these challenges and identify the causal effect of debt supply on interest rates.

This question has direct policy relevance. Governments face persistently large financing needs, which tend to increase during crises and require markets to absorb a steady flow of public debt. New issuance affects asset prices and the funding costs of governments, firms, and households, with broader implications for economic activity. Fiscal authorities therefore need quantitative evidence on how their funding decisions affect financial markets. In this paper, we provide such evidence by estimating the impact of new government debt issuance on asset prices.

Our first contribution is to propose a method to identify government bond supply surprises. The identification exploits institutional features of the UK Debt Management Office

(DMO). We focus on announcements that reveal the quantity supplied in an upcoming auction. These announcements are informative because they convey news about issuance volumes only; other characteristics, such as maturity and coupon, are announced earlier. Their information content is high, as unlike other countries, the UK DMO does not pre-announce auction sizes.[1] We isolate the surprise component of these announcements by purging their anticipated part. The anticipated component is modelled using the DMO's prior issuance pattern and moving-average terms correcting for previous prediction errors. While announced quantities themselves do not predict yield changes, the estimated surprises do. Although the model's explanatory power is high, it may not fully capture all anticipated issuance. Therefore, our estimated asset price responses in the second stage are to be interpreted as a lower bound.

Our second contribution is to study how these supply surprises transmit to the term structure of interest rates. This complements a fiscal policy literature that typically examines debt and interest rates at quarterly or annual frequency. By contrast, we estimate the effects at daily frequency, which allows for cleaner identification. We find that debt issuance has a positive, statistically significant, and state-dependent effect on yields. A £1 billion increase in 10-year-equivalent bond supply raises nominal yields by about 1 basis point in calm markets, with longer maturities reacting more strongly. The impact is nearly three times larger during periods of market stress and shows some evidence of localisation in maturity space. The effects also spill over to corporate bond markets. Decomposing yields shows that the response is driven by higher risk premia rather than changes in short rate expectations. This reflects both higher real term premia and higher inflation risk premia, with stronger responses in stressed conditions, although statistical power here is limited.

Our empirical approach relates to a growing literature that exploits primary government bond markets to identify government bond demand and supply effects (see Simon (1991), Lou, Yan, and Zhang (2013), and Fleming and Liu (2016)). Ray, Droste, and Gorodnichenko

[1] This contrasts with e.g., the US Treasury, which releases quarterly estimates of future auction sizes.

(2024) identify Treasury demand shocks by large institutional investors using high-frequency yield responses to auction results, test preferred-habitat predictions, and document state-dependent localisation. Lengyel and Giuliodori (2022) extends this approach to the Euro-zone. Closest to our paper, Phillot (2025) interpret high-frequency futures price changes around US Treasury auctions as shocks to expected supply and find positive effects on yields and term premia. Our paper differs in several respects. We focus on UK issuance announcements, where a larger share of information is unexpected due to institutional design (see Section 2.1). We identify supply surprises directly in quantity space, rather than inferring them from price changes that already reflect equilibrium responses. Lastly, we disentangle the reaction of the inflation expectations and inflation risk premia components of nominal yields, to uncover that only the latter responds to debt issuance surprises.

Our empirical findings point to a clear economic mechanism: increases in government bond supply raise yields primarily by increasing risk premia, with stronger effects during periods of market stress and larger responses at longer maturities. This pattern suggests that the transmission of debt issuance to the yield curve operates through investors' limited risk-bearing capacity rather than through changes in expected short rates.

This mechanism is naturally interpreted through the preferred-habitat theory of interest rates, originating with Modigliani and Sutch (1966) and formalised by Vayanos and Vila (2021). In this framework, investors have maturity-specific preferences, while arbitrageurs integrate the yield curve but require compensation for bearing interest rate and inflation risks. Shocks to the demand or supply of bonds therefore affect yields mainly through changes in risk premia.

A prominent application of this framework to government debt supply is Greenwood and Vayanos (2014), who show that changes in the quantity and maturity structure of public debt influence term premia by altering the risk exposure of bondholders. Our empirical results are consistent with this mechanism, but also highlight dimensions that are not captured by existing models: the joint behaviour of nominal and inflation-linked bonds, and the role of

inflation risk premia. To interpret these findings, we develop an equilibrium term structure model with supply effects. The model extends Greenwood and Vayanos (2014) by introducing exogenous inflation and two bond types, nominal and inflation-linked. Shocks to the supply of nominal bonds are absorbed by risk-averse investors, increasing their exposure to interest rate and inflation risk. As investors hold more risk, they require greater compensation, raising both real term premia and inflation risk premia and, in turn, yields. Inflation-linked bonds are insulated from inflation risk premia and therefore respond less. Consistent with the empirical evidence, supply effects are stronger when investors' risk-bearing capacity is limited.

More broadly, our paper contributes to the literature linking fiscal policy to asset pricing. Prior work shows that bond risk premia depend on fiscal expenditure levels and uncertainty (Bretscher, Hsu, and Tamoni (2020), Horvath, Kaszab, and Marsal (2021), Kučera, Kočenda, and Maršál (2022), Bayer, Born, and Luetticke (2023)), government debt ratios (Alesina, De Broeck, Prati, and Tabellini (1992), Greenwood and Vayanos (2014), Nguyen (2022)), and debt maturity structure (Chadha, Turner, and Zampolli (2013), Greenwood and Vayanos (2014), Corhay, Kind, Kung, and Morales (2023)). Government debt also affects equity, credit, and liquidity premia (Gomes, Michaelides, and Polkovnichenko (2013), Liu (2023), Krishnamurthy and Vissing-Jorgensen (2012), Bayer, Born, and Luetticke (2023)). We contribute to this literature by providing evidence that government debt issuance affects both real term premia and inflation risk premia.

The rest of the paper is organised as follows. Section 2 describes the UK government bond primary market and the data. Section 3 analyses bond supply effects, with Section 3.1 detailing identification, Section 3.2 presenting yield responses, Section 3.3 decomposing yields, and Section 3.4 examining localisation. Section 3.5 presents the equilibrium model. Section 4 concludes. The Online Appendix provides details on the intraday yield data (Section A), the term structure model (Section B), model proofs (Section C), and additional figures and tables (Section D).

2 Institutional setup and data

In this section, we first briefly describe the institutional framework of the UK Debt Management Office and its bond issuance process, which we leverage to construct our bond supply surprises. For more details on the DMO's operations, see DMO (2021). Then, we describe the data sources we use in the paper, with more details on intraday yield data left for Online Appendix A.

2.1 Description of the UK primary bond market

The DMO is the institution responsible for the UK government's debt management policy. It carries out this duty by issuing debt securities denominated in pound sterling. Securities with original maturity within a year are called bills, while securities with maturity over a year are called "gilts". Gilts make up the largest proportion of government debt, around 86%. The DMO issues two types of gilts: conventional and index-linked. Conventional gilts are nominal bonds, i.e., interest payments and coupon repayments are fixed in nominal terms. They constitute around three-quarters of the gilts issued by the DMO. Index-linked gilts are securities with coupon and final redemption payments linked to inflation, more specifically to the UK Retail Price Index (RPI). They make up about a quarter of the debt issued by the DMO. The primary means of issuing gilts is through regular auctions, accounting for over 75% of total gilt sales. The remaining part is issued through syndicated gilt offerings or mini gilt tenders.

The annual financing remit, set by the UK Treasury, outlines the gilt sales required from the DMO for the upcoming financial year. The document specifies the total amount of gilt sales and the breakdown between index-linked gilts and conventional gilts in different maturity buckets. It is published in mid-March, as the financial year runs from April 1 to March 31. The remit is then revised in April when the central government's final net cash requirement for the previous financial year is published. Furthermore, the remit is

usually revised again in the autumn when the government updates the budget together with forecasts of public finances. Additionally, the Remit is sometimes revised when there are large unforeseen changes in fiscal policy, such as during the Covid-19 crisis. The Remit states the dates of the auctions and provides estimates of total issuance and its split across four categories. These categories are index-linked gilts; and short-, medium-, and long-maturity conventional gilts with 0-7, 7-15, and 15+ years to maturity, respectively. The DMO explicitly targets different auction sizes in each of these buckets.

The DMO announces its quarterly auction plan at the end of March, May, August, and November. This calendar fixes the dates of the coming auctions and specifies the maturity and coupon of the bonds. Importantly, it does not provide information about the issuance size. This is unlike the US, where the Treasury provides quarterly estimates of future auction sizes.[2] The announcement of the upcoming auction is published at 3:30 PM, usually, but not always, on the Tuesday in the week preceding the auction. This is the first time investors learn the exact size of the upcoming auction. These documents also include ISIN, SEDOL codes, coupon payments, and the terms and conditions of the auction. Figures in the Online Appendix show the debt issuance proportions across security types (C1), give example of the Financing Remit (C2), the Quarterly Plan (C3), and an example of a 20-year gilt auction announcement from October 2, 2012 (C4)

2.2 Data

We collect auction announcements from the DMO's website. For the few cases when this was not available, we use date and volume from the DMO's operations calendar and auction results publications. We restrict our attention to announcement days with a single conventional nominal gilt announcement (and no tenders, syndicated offerings or index-linked gilt auctions). Our sample runs from May 15, 2001 (the first conventional gilt announcement on the DMO's website) to December 31, 2019 and contains 317 announcements. Table 1 sum-

[2]See the US Quarterly Refunding Press Conference: `https://home.treasury.gov/policy-issues/financing-the-government/quarterly-refunding`.

marizes the descriptive statistics of the announced volumes in each bucket. The number of announcements is roughly balanced across the DMO's buckets, with 111 short, 92 medium, and 114 long maturity, with means around £3.5bn, £2.8bn and £2.1bn, and standard deviations of £0.7bn, £0.4bn and £0.3bn in the short, medium and long buckets, respectively. This is consistent with the DMO's goal of targeting larger auction sizes for shorter maturities, which are easier for the market to absorb.

To study the transmission of the supply surprises, we conduct our analysis in two frequencies. First, to analyse the immediate market reaction to the announcements and to test the localisation of the effect, we use 1-minute intraday yield data at eight available maturities (1, 2, 3, 5, 10, 15, 20, and 30 years) from LSEG Data & Analytics (formerly Refinitiv/Thomson Reuters). The yield surprises $z_t^{(m)} = y_{t,post}^{(m)} - y_{t,pre}^{(m)}$ on announcement day t at maturity m are measured as the yield difference from 5 minutes before to 15 minutes after the announcement, then standardised to zero mean and unit variance.[3] The time series of the eight yield surprise series $z_t^{(m)}$ is displayed in Figure 1, the ACF in Figure C5 while Table 2 reports their descriptive statistics. The means and autocorrelations are statistically zero, suggesting that the surprises are not systematic, while the standard deviations range from 0.6 to 1.2 basis points. More information on the series and how we ensure it only reflects bond supply information is in Online Appendix A.

To better understand how changes in the supply transmits to yields, we move to daily frequency and use the Bank of England's zero-coupon nominal and real yield curves, corporate bond indices, inflation swap rates and Credit Default Swap (CDS) rates from LSEG. Inflation swap rates are available from October 2004, CDS rates from July 2008, limiting the estimation sample when using these. To identify high financial stress regimes, we use the country-level composite indicator of systemic stress in the financial system (CISS) by Hollo, Kremer, and Lo Duca (2012), from the ECB. Our state variable equals one when the UK

[3]Our results are robust to both narrower and wider event window specifications. These results are available upon request.

CISS index is above its 75th percentile and zero otherwise (see Figure C8).[4]

3 Bond supply effects on the term structure of interest rates

In this section, we analyse how unexpected changes in issuance quantities transmit to the term structure of interest rates. We start by describing how we construct our measure of bond supply surprises in Section 3.1. Then, in Section 3.2, we first examine the immediate market reaction of high-frequency yields, then trace the effects across the nominal and real zero-coupon yield curves, corporate bonds, and, in Section 3.3, expectations, risk premia, and CDS rates. Lastly, in Section 3.4, we test whether the yield reaction localises in the maturity space. Throughout the analysis, we also examine how financial stress amplifies these effects, motivated by the fact that risk aversion is a key state-dependency in the preferred-habitat framework, and He and Krishnamurthy (2013) suggests that risk aversion is higher during financial market stress periods.

3.1 Isolating the surprise quantities

As described above, the first time the DMO discloses the size of an upcoming auction is in the auction announcement, published about a week before the auction, which contains new information only about the supplied quantity. To study the effects of debt issuance on interest rates, one can regress announced volume on daily yield changes. However, the estimated coefficient will be statistically insignificant and not capture the true causal effect, and because most of the impact is priced in before the announcement. Debt issuance policy generally aims to be predictable, which market participants factor in when forming expectations, so much of the issuance is anticipated and its effects already priced in.

[4]Our results are robust to a wide range of this threshold and they are available upon request.

Therefore, in the first step, we purge the anticipated portion from the supply announcements. To do so, we estimate a time-series model using issuance quantity information available before the announcement. The anticipated quantity for each bond is modeled using the previous auction size in the same maturity bucket, the average issuance size in that bucket according to the Remit in place at the time of the announcement,[5] and the size of the last auction of the same ISIN. We also include two moving average terms to correct for prediction errors from previous announcements. This specification was selected by first adding the average auction size implied by the Remit, then confirming that only one lag of the last bucket and last ISIN variables were significant, and finally selecting the MA lag order using the Bayesian Information Criterion (BIC), which has a clear preference unlike the Akaike Information Criterion (AIC), see Figure 2.

$$\begin{aligned}\text{Announced Supply}_t = \alpha + \beta_1 \text{Last Bucket}_t + \beta_2 \text{Last ISIN}_t + \beta_3 \text{Remit}_t \\ + \beta_4 \text{MA}(1)_t + \beta_5 \text{MA}(2)_t + \varepsilon_t \end{aligned} \tag{1}$$

The construction of these explanatory variables reduces the effective sample size to 254, mainly because we can only use re-issues so that a previous auctioned volume is already observed at the time of the announcement. The regression results are displayed in Table 3 and shows that all variables are strongly statistically significant and explain over 90% of the variation in the announced volume. We purge the Announced Supply series from this anticipated portion, which amounts to using the series of residuals from (1). These we label as the supply surprises and label as S_t. The descriptive statistics of S_t is shown in the last row of Table 1 and is plotted in Figure 3. The series show no autocorrelation (see Figure 4), ranges between £-840mn and £1.3bn and has a standard deviation of £250mn. The volatility is largest around the Global Financial Crisis, the Eurozone crisis and after the

[5] An alternative would be to use progress reports on the Remit, but these are inconsistently published and often outdated, making them unsuitable for our purpose.

Brexit referendum. Our results in the sections below are robust to omitting the two largest surprises or the intercept in equation (1), if anything, makes them stronger.

One might expect these surprises to reflect only shifts in the timing of the DMO's issuance plan, implying that a positive surprise would later be offset by negative surprises. However, the Remit is not executed perfectly, and both over- and under-issuances occur. Because the Remit is typically revised several times a year, surprises may also signal the direction of an upcoming revision. Figure 3 and its ACF in Figure 4 show no serial autocorrelation and no evidence that the size declines toward the end of the fiscal year. In other words, a smaller-than-expected supply announcement is not systematically followed by a larger one, and the informational content of announcements remains stable throughout the year.

Our method of purging announced volumes is unlikely to perfectly eliminate all the anticipated portion, as market participants form expectations using a broader information set than that spanned by our explanatory variables. Consequently, the series may still include some anticipated elements, implying that the true shocks might be smaller than S_t. The results presented below should therefore be viewed as a conservative lower bound.

3.2 Effect on nominal and real yields

We begin by examining the immediate market response to auction announcement surprises. Specifically, we regress S_t on high-frequency yield changes for short-, medium-, and long-maturity buckets. The dependent variables are changes in 5-, 10-, and 20-year benchmark yields within a 20-minute event window. Panel A of Table 4 reports these results, while Panel B includes interactions between S_t and our market stress indicator. Yields generally rise within minutes following higher-than-expected issuance. In the medium- and long-maturity buckets, this reaction is statistically significant only during periods of market stress. In those periods, however, the reaction is amplified suggesting that risk premia migth play an important role in the transmission, which point we return later.

To further examine the transmission mechanism, we analse the response of the zero-

coupon yield curve at the daily frequency. This allows us to capture effects that may unfold beyond the 20-minute event window and available across all maturities in both real and nominal terms. Importantly, this enables a decomposition of yields into expectations and risk premia components, both in real and nominal terms, which we look at in the next section.

We estimate the following specification:

$$\Delta Y_t^{(m)} = a^{(m)} + b^{(m)} S_t + c^{(m)} X_t + \varepsilon_t^{(m)} \tag{2}$$

$\Delta Y_t^{(m)} = Y_t^{(m)} - Y_{t-1}^{(m)}$ is the change in the Bank of England zero-coupon curve at maturity m relative to the previous day, while X_t collects control variables we add at a later stage. We multiply S_t with the announced bond's maturity, following the maturity-weighted debt measure of Greenwood and Vayanos (2014), and scale it such that the estimated coefficients in Table 5 can be interpreted as the effects of an unexpected $1bn$ increase in 10-year-equivalent conventional bond supply. The results in Table 5 show that such an increase in supply raises nominal yields between 0.9 to 1.2 basis points. Rates at longer maturities respond stronger, implying an increase in the slope of the yield curve. Real yields also rise, though by a smaller and not statistically significant amount. Figure C9 illustrates these responses, while Figure 5 shows that the effect on nominal yields persists for a few days, particularly at the long end of the curve

In Panel B of Table 5 we split the sample into low-stress and high-stress regimes, and estimate a state-depend version of equation (2):

$$\Delta Y_t^{(m)} = (1 - I_t)\big[a_0^{(m)} + b_0^{(m)} S_t\big] + I_t\big[a_1^{(m)} + b_1^{(m)} S_t\big] + c^{(m)} X_t + \varepsilon_t^{(m)} \tag{3}$$

Where we use the CISS index to construct I_t, as described in Section 2.2. Yield responses to issuance surprises are estimated to be roughly three times larger during high-stress periods, reaching about 3 basis points, compared with low-stress periods. Real yields exhibit a similar

pattern, though the effect is smaller and statistically insignificant.

Our results are robust to including various control variables in X_t, such as weekday dummies, the short-term interest rate, and inflation (as implied by the model in Section 3.5), and Bloomberg GDP and CPI surprises. Table C1 in the Online Appendix reports these.

Next, we examine spillovers to the corporate bond market by estimating Equation (2) on corporate bond indices. Table 6 shows the reaction of these, first split into maturity buckets of 1–3, 3–5, 5–10, 10–15, and 15+ years in Panel A, and then categorized into AAA-, AA-, A-, and BBB-rated sub-indices in Panel B. Results show large spillovers to corporate bonds, with strongest response between 10 and 15 years and for AAA rating.

3.3 Supply effects on expected short rates and risk premia

What explains the reaction of the yield curve? According to the expectations hypothesis, the response of long-term rates could reflect either higher expected future short rates or higher risk premia. Ang, Bekaert, and Wei (2008) find that about 80% of the variation in U.S. nominal yields is attributable to changes in expected inflation and the inflation risk premium. Using quarterly data and recursive identification, Dai and Philippon (2005) find that risk premia account for one-third of the response of long-term nominal rates to a fiscal deficit shock. Similarly, Laubach (2011), using the same frequency, find that fiscal deficits primarily affect short rates and inflation, with only small movements in risk premia.

We decompose nominal and inflation-linked bond yields into expected real short rates, expected inflation, real term premium, and inflation risk premium, and assess how each component responds to unexpected changes in debt issuance. The decomposition is estimated with an affine term structure model following Abrahams, Adrian, Crump, Moench, and Yu (2016). We also control for relative liquidity effects between nominal and inflation-linked bonds using the breakeven inflation-linked swap spread as a liquidity proxy, following Pflueger and Viceira (2016). As this series begins in October 2004, our sample size for this

analysis is reduced to 233 observations. This further reduces to 229 for the real decomposition, due to gaps in the inflation-linked gilt data. The model assumes that bond yields, the short rate, and market prices of risk are affine functions of observable state variables. For details on the model and estimation, see Section B in the Online Appendix.

We re-estimate Equation (2), replacing the dependent variable with each yield component in turn. The top panel of Table 7 reports responses of the nominal term premium and expected short rates, while the bottom panels further decompose these into real and inflation components. The total yield response equals the sum of these coefficients, and Figure C10 illustrates these effects across the curve. Results indicate that issuance surprises affect yields through the risk premium channel rather than expectations. Point estimates show no response in real or inflation expectations, whereas both real term premia and inflation risk premia increase, albeit most coefficients lack statistical significance. Inflation risk premia is more important at shorter and intermediate maturities, while real term premia drive more the long end. Table 8 presents results by stress regime (albeit with very low statistical power), suggesting stronger risk premia responses under high stress, with expectations remaining unresponsive.

Apart from duration and inflation risks, the behaviour of credit risk is also of interest. Large-scale government issuance could raise concerns about repayment capacity, potentially increasing perceived credit risk. Such risk is typically reflected in CDS spreads, which represent the cost of insuring against default. However, Table 9 indicates that issuance surprises have no significant effect on CDS rates, suggesting that markets do not price higher sovereign credit risk following increased supply.

Overall, our results indicate that additional government debt issuance increases the compensation investors demand for bearing interest rate and inflation risks, while expectations of future short rates remain unaffected. This contrasts with Dai and Philippon (2005) and Laubach (2011), who report stronger responses of short rate expectations to fiscal deficits, based on lower-frequency data and deficit measures rather than bond supply shocks. Phillot

(2025) finds that supply shocks raise breakeven inflation rates; our results suggest this reflects higher inflation risk premia rather than higher inflation expectations. The positive response of short rates and term premia resembles the effect of a monetary policy tightening (see Hanson and Stein (2015); Abrahams et al. (2016)), except that inflation risk premia move in the opposite direction. This difference arises because monetary policy shocks are contractionary and disinflationary, whereas bond supply shocks signal a more expansionary and inflationary fiscal stance.

3.4 Localisation

We continue our analysis of the transmission of supply effects by testing the Localisation Hypothesis of the preferred habitat view of the term structure. In this framework, arbitrageurs smooth out shocks to specific segments of the curve by taking positions across maturities. Their ability to do so is tightly linked to their capacity to bear the risks associated with these positions. When this capacity is high, shocks are well smoothed and have global effects. The location of the shock does not influence relative yield responses. Conversely, when risk-bearing capacity is low, yield reactions become concentrated around the maturity where the shock occurs.

To test this, we measure the relative response of yields across the curve, conditional on a shock at a given point. We use the high-frequency yield changes $z_t^{(m)}$ recorded around to the auction announcements and estimate the regression specification of Ray, Droste, and Gorodnichenko (2024):

$$z_t^{(\tau)} = \alpha^{(\tau)} + \gamma_s^{(\tau)} I(m_t = \text{short}) z_t^{(\tau*)} + \gamma_l^{(\tau)} I(m_t = \text{long}) z_t^{(\tau*)} + \epsilon_t^{(\tau)} \tag{4}$$

where m_t is the maturity of the announced bond at time t, $I(m_t = \text{short})$ and $I(m_t = \text{long})$ are indicator functions that take the value of one if the auctioned bond's maturity at time t is short or long, respectively. The estimates of $\gamma_s^{(\tau)}$ and $\gamma_l^{(\tau)}$ measure the relative effect of the

supply shock for short-maturity and long-maturity bond issuance, respectively. Note, that $z_t^{(\tau*)}$ is not the true structural shock, but the reaction to the true unobserved shock, given that the shock convey valuable supply information, and the narrow high-frequency event window minimizes the risk of confounding influences in $z_t^{(\tau*)}$. Then, equation (4) can reveal the relative effects of shocks to the short and the long end of the curve.

Following Ray, Droste, and Gorodnichenko (2024), we estimate equation (4) for all maturities τ, holding a baseline maturity $\tau*$ fixed and splitting the sample into low- and high-stress periods based on the CISS index. Short-maturity announcements are defined as those with maturities below 7 years, and long-maturity announcements as those above 10 years. As Ray, Droste, and Gorodnichenko (2024) explains, the choice of baseline $\tau*$ is arbitrary; we set it to 5 years. Figure 6 displays the estimated γ coefficients, with the left panel showing low-stress periods and the right panel high-stress periods. While results for short maturities are inconclusive, we find clear evidence of localisation for long-maturity announcements. The null hypothesis of equality between $\hat{\gamma}_s^{(\tau)}$ and $\hat{\gamma}_l^{(\tau)}$ cannot be rejected in the low-stress sample but is strongly rejected in the high-stress sample at all conventional significance levels.

Table 10 reports an alternative test of localisation based on maturity-specific versions of equation (2), where supply surprises are interacted with indicators for short- and long-maturity issuance, which are below 7 years and above 10 years respectively. In Panel A, short-maturity issuance tends to have larger effects on short and intermediate yields, while long-maturity issuance loads increasingly on longer maturities and becomes statistically significant from the 10-year yield onwards. While the statistical significance is low, the pattern points towards some localisation of supply shocks along the curve. Panel B introduces state dependence, but the estimates are imprecise and do not allow us to verify the prediction that localisation intensifies during periods of financial stress. These results are therefore best viewed as corroborative evidence.

3.5 A term structure model of nominal and real bonds with supply effects

Our empirical findings suggest that an increase in nominal bond supply raises both nominal and real yields by increasing real term premia and inflation risk premia. Investors require higher compensation for bearing interest rate and inflation risks when faced with greater government bond supply. To provide theoretical context, we now examine this mechanism within a formal framework. The objective is illustrative rather than to deliver a full structural explanation. This section presents the model setup, equilibrium, and intuition, while proofs and derivations are provided in Online Appendix C. The model extends the Greenwood and Vayanos (2014) version of the Vayanos and Vila (2021) preferred-habitat framework by incorporating inflation risk and introducing two bond types: a continuum of nominal bonds and a continuum of inflation-linked (real) bonds. Related work includes Diez de los Rios (2020), who develop a discrete-time version of Greenwood and Vayanos (2014) with fixed nominal and real bond supply and endogenous inflation determined by a Taylor-rule equation, focusing on how higher bond supply can raise inflation. In contrast, our continuous-time model assumes exogenous inflation and emphasizes the transmission of bond supply shocks to yields through risk pricing.

Bonds in the model have maturities τ in the interval $(0; T]$. A nominal bond with maturity τ pays one unit of currency at time $t + \tau$. Its time t price is denoted by $P_t^{N,(\tau)}$. An inflation-linked bond with maturity τ pays one unit of consumption good at time $t + \tau$. Its time t price is denoted by $P_t^{R,(\tau)}$. The bond's spot yields are denoted by $y_t^{N,(\tau)}$ and $y_t^{R,(\tau)}$. They are related to prices by:

$$y_t^{N,(\tau)} = -\frac{\log P_t^{N,(\tau)}}{\tau} \tag{5a}$$

$$y_t^{R,(\tau)} = -\frac{\log P_t^{R,(\tau)}}{\tau} \tag{5b}$$

Bonds are issued by a government and traded by arbitrageurs and other investors that are

not modeled explicitly. Supply by the government and the demand of un-modeled investors as assumed to be exogenous and price inelastic.[6] The supply of nominal bonds $s_t^{N,(\tau)}$ is given by a one factor model, while inflation-linked bond supply $s_t^{R,(\tau)}$ is fixed for simplicity.

$$s_t^{N,(\tau)} = \zeta^N(\tau) + \theta^N(\tau)\beta_t \tag{6a}$$

$$s_t^{R,(\tau)} = \zeta^R(\tau) \tag{6b}$$

The functions $\zeta^N(\tau)$, $\zeta^R(\tau)$, and $\theta^N(\tau)$ are deterministic functions of the maturity of the bonds. The variable β_t is a stochastic nominal bond supply factor that follows the Ornstein-Uhlenbeck process. Its long-run mean is zero, similar to our empirical supply shock series.

$$d\beta_t = -\kappa_\beta \beta_t dt + \sigma_\beta dB_{\beta,t} \tag{7}$$

The function $\zeta^N(\tau)$ gives the average supply of bonds at maturity τ, while $\theta^N(\tau)$ measures the sensitivity of the nominal bond supply to the supply factor β_t. Assumption 1 in Online Appendix C postulates that $\theta^N(\tau)$ satisfies some properties ensuring that an increase in β_t makes arbitrageurs' equilibrium portfolios more sensitive to inflation and interest rate risks.

The instantaneous risk-free real rate is denoted by r_t. It is defined as $\lim_{\tau \to 0} y_t^{R,(\tau)} = r_t$ and it follows the Ornstein-Uhlenbeck process, as inflation:

$$dr_t = \kappa_r(\bar{r} - r_t)dt + \sigma_r dB_{r,t} \tag{8}$$

$$d\pi_t = \kappa_\pi(\bar{\pi} - \pi_t)dt + \sigma_\pi dB_{\pi,t} \tag{9}$$

where $\bar{r}, \kappa_r, \sigma_r, \bar{\pi}, \kappa_\pi, \sigma_\pi > 0$ are constants and $B_{r,t}$ and $B_{\pi,t}$ are independent Brownian motions.[7] Parameters $\bar{r}$ and $\bar{\pi}$ are the long-run means of the processes, κ_r and κ_π are the

[6] At the end of the section we briefly discuss how introducing price elastic habitat investors would change the model's predictions.

[7] Treating objects r_t and π_t independent in this partial equilibrium setting yields the main intuition of our model while keeping it tractable. In general equilibrium, they are of course jointly determined. For such treatment in a New Keynesian setting see Ray (2019). One can generalise our partial equilibrium setting by

mean-reverting parameters. The volatility parameters are σ_r and σ_π.

Marginal investors in the model are short-lived, risk averse arbitrageurs, who select their portfolio by solving:

$$\max_{\left\{x_t^{N,(\tau)},x_t^{R,(\tau)}\right\}_{\tau\in(0,T]}} E_t[dW_t] - \frac{a}{2}V_t[dW_t] \tag{10}$$

W_t denotes arbitrageurs' real wealth, a is the coefficient of risk aversion. $x_t^{N,(\tau)}$ and $x_t^{R,(\tau)}$ are the units of wealth invested in the nominal bond and the inflation-linked bond with maturity of τ. The corresponding budget constraint is given by:

$$\begin{aligned} dW_t = \int_0^T \left(x_t^{N,(\tau)} \frac{dP_t^{N,(\tau)}}{P_t^{N,(\tau)}} + x_t^{R,(\tau)} \frac{dP_t^{R,(\tau)}}{P_t^{R,(\tau)}} \right) d\tau - \Big(\int_0^T x_t^{N,(\tau)} d\tau \Big) \pi_t dt \\ + \Big(W_t - \int_0^T \big(x_t^{N,(\tau)} + x_t^{R,(\tau)} \big) d\tau \Big) r_t dt \end{aligned} \tag{11}$$

The first expression is the return from investing in bonds, as $\int_0^T x_t^{N,(\tau)} d\tau$ and $\int_0^T x_t^{R,(\tau)} d\tau$ represent the amount of wealth invested in nominal bonds and real bonds, respectively. The second term $\Big(\int_0^T x_t^{N,(\tau)} d\tau \Big) \pi_t dt$ deflates the return from nominal bonds.[8] Finally, the last expression is the return gained by investing the remaining wealth in the risk-free rate.

The model is solved by first conjecturing and later verifying that equilibrium spot rates are affine functions of the risk factors:

$$y_t^{N,(\tau)} = \frac{1}{\tau} \big[A_r^N(\tau) r_t + A_\beta^N(\tau) \beta_t + A_\pi^N(\tau) \pi_t + C^N(\tau) \big] \tag{12a}$$

$$y_t^{R,(\tau)} = \frac{1}{\tau} \big[A_r^R(\tau) r_t + A_\beta^R(\tau) \beta_t + C^R(\tau) \big] \tag{12b}$$

assuming correlated r_t and π_t in Equation 8: $dr_t = \kappa_r(\bar{r} - r_t - \gamma\pi_t)dt + \sigma_r dB_{r,t} + \sigma_{r\pi} dB_{\pi,t}$. Then, shocks to π_t will affect r_t. With $\gamma \neq 0$, inflation shocks affect expected future short rates, while with $\rho_{r\pi} \neq 0$, shocks to inflation affect the current short rate. Nevertheless, our key results carry over also to this case (for a derivation of correlated r_t and β_t see Appendix B of Greenwood and Vayanos (2014)).

[8] Note that the way we formulate price dynamics implies a (locally) non-stochastic price level $dP_t = P_t \pi_t dt$. A locally stochastic price level would introduce a real payoff risk for nominal bonds, which we abstract away from in this paper.

Online Appendix C shows that one can apply Ito's lemma to these, substitute back into the budget constraint (11), solve the arbitrageurs' optimization problem (10) and arrive to the first order conditions:

$$\mu_t^{N,(\tau)} - \pi_t - r_t = A_r^N(\tau)\lambda_{r,t} + A_\beta^N(\tau)\lambda_{\beta,t} + A_\pi^N(\tau)\lambda_{\pi,t} \tag{13a}$$

$$\mu_t^{R,(\tau)} - r_t = A_r^R(\tau)\lambda_{r,t} + A_\beta^R(\tau)\lambda_{\beta,t} \tag{13b}$$

where coefficients $\lambda_{i,t}$ are given by:

$$\lambda_{i,t} = a\sigma_i^2 \int_0^T x_t^{N,(\tau)} A_i^N(\tau) + x_t^{R,(\tau)} A_i^R(\tau) d\tau \qquad \text{for } i = r, \beta \tag{14}$$

$$\lambda_{\pi,t} = a\sigma_\pi^2 \int_0^T x_t^{N,(\tau)} A_\pi^N(\tau) d\tau \tag{15}$$

Equations (13a) and (13b) are also the no-arbitrage conditions in the model. No arbitrage requires the existence of prices for each risk factor. Then, the expected excess return of each asset over the short rate is equal to the asset's sensitivity to the risk factors times the risk factor's price, summed across all risk factors. The coefficients $\lambda_{i,t}$ are the prices of the risk factors, measuring the expected excess return per unit of sensitivity to each factor. They are determined through equilibrium conditions. Note that $\lambda_{i,t}$ are proportional to how sensitive the arbitrageurs' portfolio is to factor i. For example, the sensitivity of the portfolio to the short rate is $\int_0^T x_t^{N,(\tau)} A_r^N(\tau) + x_t^{R,(\tau)} A_r^R(\tau) d\tau$. As inflation-linked bonds shield investors from inflation, the inflation risk factor only loads on the nominal bond. Note that even the real return of the nominal bond is sensitive to inflation risk.

In equilibrium, the supplied amount of bonds will be equal to the investment of the arbitrageurs.

$$x_t^{N,(\tau)} = s_t^{N,(\tau)} \tag{16a}$$

$$x_t^{R,(\tau)} = s_t^{R,(\tau)} \tag{16b}$$

We can use the market clearing (16a), (16b), the supply (6a), (6b) and (B6a), (B6b) to substitute into the first order conditions (13a), (13b), yielding two functions that are affine in the risk factors r_t, β_t and π_t, verifying our initial conjecture. Setting linear terms in r_t, β_t and π_t equal to zero gives five ordinary differential equations (ODEs) in $A_r^N(\tau)$, $A_\beta^N(\tau)$, $A_\pi^N(\tau)$, $A_r^R(\tau)$ and $A_\beta^R(\tau)$. The solutions to these ODEs are stated in Theorem 1 and derived in Online Appendix C.

Theorem 1. *The nominal bond sensitivities $A_r^N(\tau)$, $A_\beta^N(\tau)$ and $A_\pi^N(\tau)$ are given by*

$$A_r^N(\tau) = \frac{1 - e^{-\kappa_r \tau}}{\kappa_r} \tag{17a}$$

$$A_\beta^N(\tau) = \frac{Z_r}{\kappa_r}\Big[\frac{1 - e^{-\hat{\kappa}_\beta \tau}}{\hat{\kappa}_\beta} - \frac{e^{-\hat{\kappa}_\beta \tau} - e^{-\kappa_r \tau}}{\kappa_r - \hat{\kappa}_\beta}\Big] + \frac{Z_\pi}{\kappa_\pi}\Big[\frac{1 - e^{-\hat{\kappa}_\beta \tau}}{\hat{\kappa}_\beta} - \frac{e^{-\hat{\kappa}_\beta \tau} - e^{-\kappa_\pi \tau}}{\kappa_\pi - \hat{\kappa}_\beta}\Big] \tag{17b}$$

$$A_\pi^N(\tau) = \frac{1 - e^{-\kappa_\pi \tau}}{\kappa_\pi} \tag{17c}$$

The real bond sensitivities $A_r^R(\tau)$ and $A_\beta^R(\tau)$ are given by

$$A_r^R(\tau) = \frac{1 - e^{-\kappa_r \tau}}{\kappa_r} \tag{18a}$$

$$A_\beta^R(\tau) = \frac{Z_r}{\kappa_r}\Big[\frac{1 - e^{-\hat{\kappa}_\beta \tau}}{\hat{\kappa}_\beta} - \frac{e^{-\hat{\kappa}_\beta \tau} - e^{-\kappa_r \tau}}{\kappa_r - \hat{\kappa}_\beta}\Big] \tag{18b}$$

Where Z_r, Z_π and $\hat{\kappa}_\beta$ are given by equations (B12a), (B12b) *and* (B13) *respectively in Online Appendix C, together with functions $C^N(\tau)$ and $C^R(\tau)$.*

In our empirical analysis, we found that the supply shock raises both nominal and real yields due to an increase in real term premia and inflation risk premia. To demonstrate this in this setting we need to show that $\partial y_t^{N,(\tau)}/\partial \beta_t = A_\beta^N(\tau)/\tau$ and $\partial y_t^{R,(\tau)}/\partial \beta_t = A_\beta^R(\tau)/\tau$ are positive.

Proposition 1. *The effect of a shock to the supply factor β_t on nominal and real yields is positive.*

Proposition 2. *The effect of a shock to the supply factor on interest rate and inflation risk prices is positive.*

Proposition 3. *The effect of a shock to the supply factor on yields and risk prices increases with risk aversion.*

The proofs follow from Lemma A.1 and Lemma A.2 of Greenwood and Vayanos (2014) and are presented in Online Appendix C. The intuition is as follows. An increase in the supply factor raises nominal bond holdings, increasing portfolio exposure to interest rate and inflation risks. This elevates the prices of these risks, pushing up both nominal and real yields. Because inflation risk loads only on nominal bonds, the nominal-real spread widens. When risk aversion is high, arbitrageurs reduce risk-taking, amplifying these effects, which are all consistent with our empirical results.

The framework can be extended to incorporate price-elastic preferred-habitat demand, as in Vayanos and Vila (2021). Market clearing (Equations 16a and 16b) requires that arbitrageur and habitat demand sum to the total supply in equilibrium, which feeds into the arbitrageurs' first-order conditions (Equations 13a and 13b). These are also the no-arbitrage conditions implying that bond's expected excess returns equals the sum of their factor loadings times the corresponding risk prices $\lambda_{\cdot}$. Without habitat investors, risk prices depend on total outstanding supply; with habitat demand, they depend on net supply after habitat demand. Habitat demand increases with yield and demand elasticity, dampening the transmission of shocks to yields. See Sections 3.3 and 4.1 of Vayanos and Vila (2021) for more details.

4 Conclusion

This paper identifies UK government debt supply surprises from bond issuance announcements and examines their impact on the term structure of interest rates. We find that additional debt raises yields significantly, with effects amplified during periods of market

stress. We find some evidence that under such conditions yield responses become localized around the maturity of the announced bond for long-matuirty bonds. Decomposing yields into expected short rates and risk premia reveals that the transmission occurs through risk premia rather than expectations: both real term premia and inflation risk premia rise with supply. We rationalize these findings in an equilibrium term structure model where risk-averse investors absorb shocks to the supply of nominal bonds issued by the government, increasing portfolio exposure to interest rate and inflation risks. This raises the prices of these risks and, consequently, risk premia and yields in a state-dependent manner linked to risk aversion. Because real bonds are insulated from inflation risk, nominal yields rise more, widening the nominal-real spread via inflation risk premia.

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Figures

Figure 1: Time series of the high-frequency yield surprises $z_t^{(m)}$

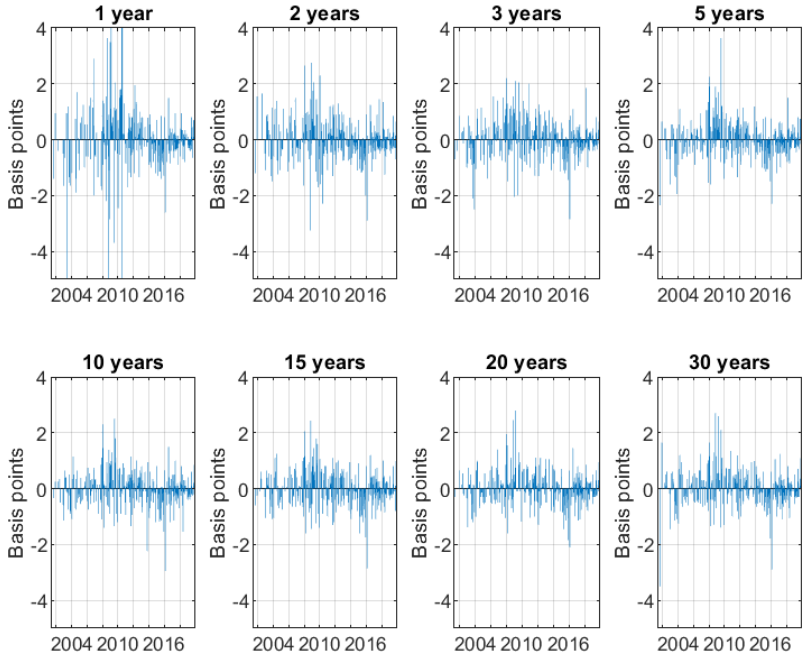


Figure 2: BIC (Bayesian Information Criterion) and AIC (Akaike Information Criterion) of the first stage model

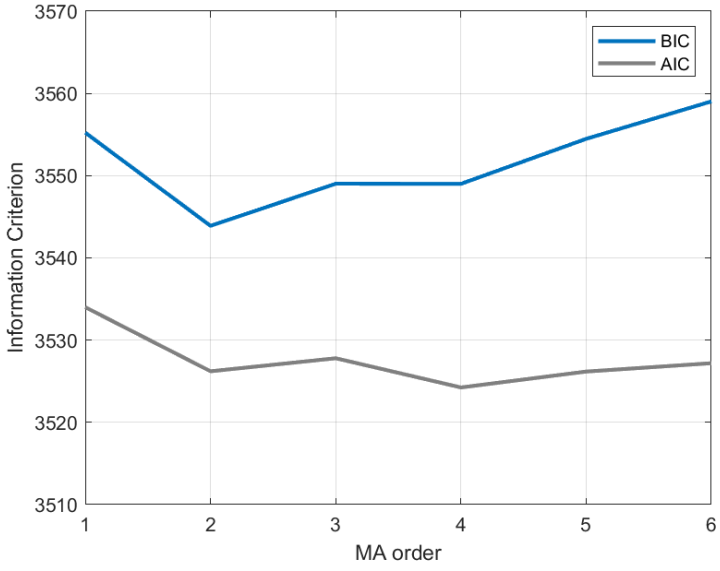


Figure 3: Time series of the announcement surprises S_t

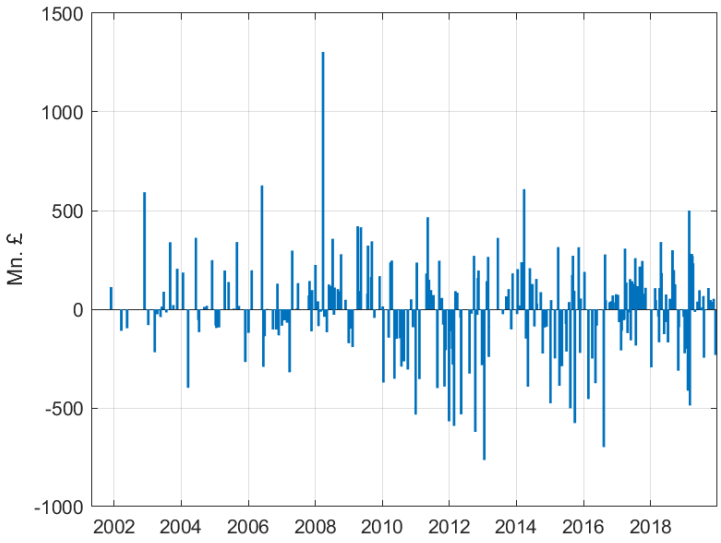
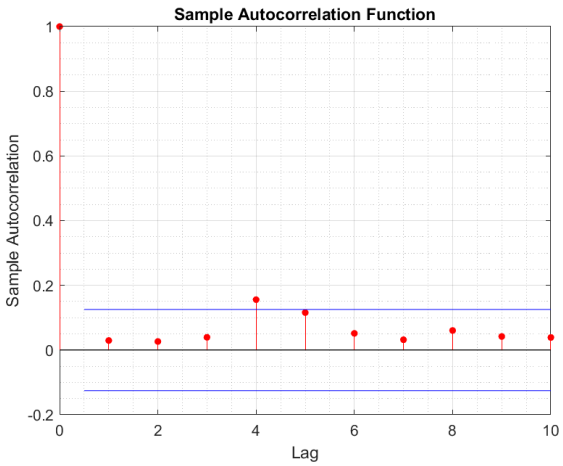
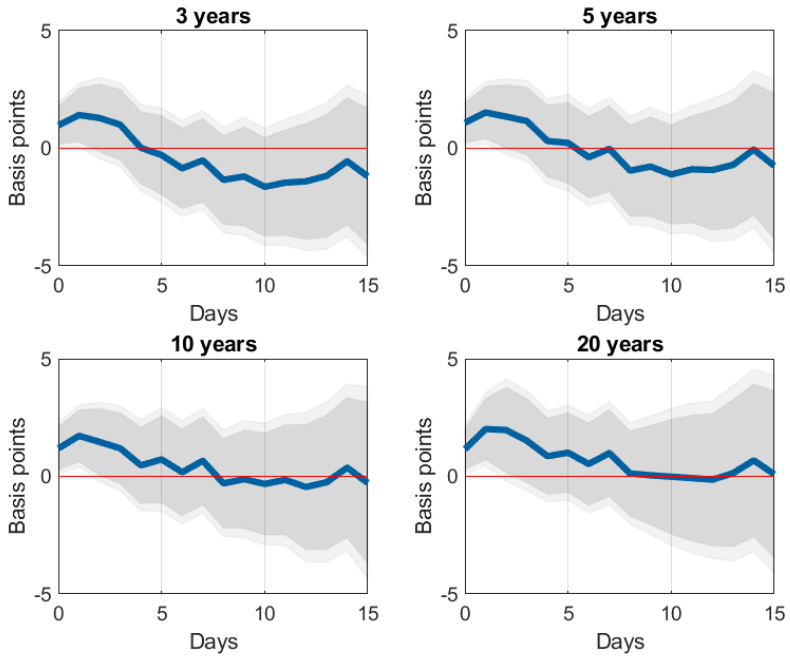


Figure 4: Sample autocorrelation function of the announcement surprises S_t



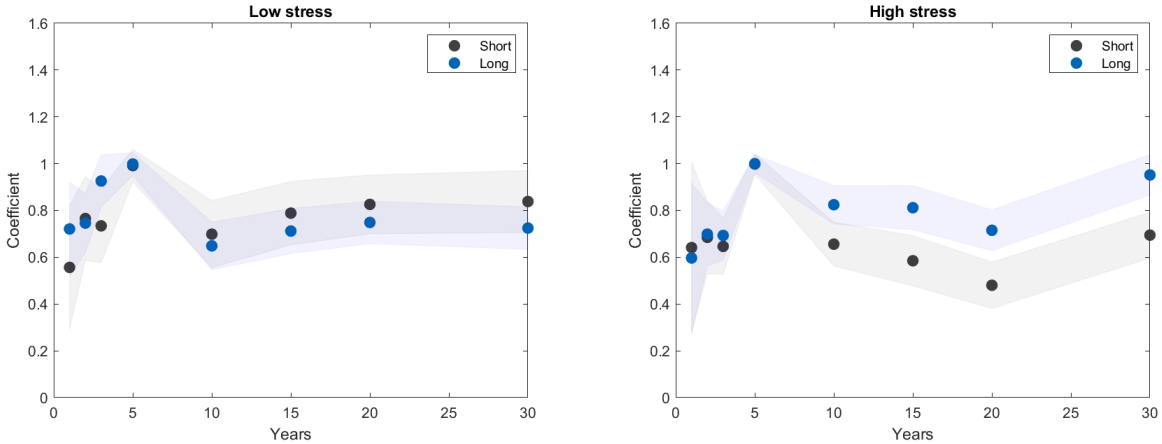
Note: Estimated autocorrelation function of the announcement surprises S_t up to ten lags. Blue lines represent two standard errors confidence intervals.

Figure 5: Impulse responses of 3, 5, 10 and 30-year nominal yields to announcement surprises S_t



Note: Impulse responses of 3, 5 and 10-year maturities from daily local projections $R^{(m)}_{t+h} - R^{(m)}_{t-1} = a^{(h,m)} + b^{(h,m)} S_t + \varepsilon^{(h,m)}_t$ for $h = 1, ..., 15$. Blue lines are the estimated $b^{(h,m)}$ coefficients, shaded areas are 90% and 95% Newey-West (10 lags) confidence intervals.

Figure 6: Localization Regression Results



Note: Plots of the regression coefficients from regression equation (4). The left panel compares the effects of short-, and long-maturity auctions during low-stress periods; the right panel compares the effects of short-, and long-maturity auctions during high-stress periods. Shaded areas are two standard error (10 lags) confidence intervals are included. Appendix Figure C11 reports P-values for the equality of estimates.

Tables

Table 1: Descriptive statistics of the auction announcement volumes (in £mn)

	N	Mean	Std.	Min	Max
Short	111	3461.712	730.969	2500	5250
Medium	92	2823.37	440.931	1750	4000
Long	114	2144.737	322.282	1250	2750
Pooled	317	2802.839	766.8	1250	5250
Surprises	254	-2.105	245.839	-763.854	1304.019

Table 2: Descriptive statistics of the high-frequency yield surprises $z_t^{(m)}$

Maturity	N	Mean	Std.	Correlations							
				1 y.	2 y.	3 y.	5 y.	10 y.	15 y.	20 y.	30 y.
1 year	317	-0.017	1.249	1							
2 years	317	-0.04	0.753	0.312	1						
3 years	317	0.002	0.697	0.221	0.652	1					
5 years	317	-0.014	0.673	0.332	0.645	0.733	1				
10 years	317	-0.036	0.625	0.32	0.639	0.725	0.79	1			
15 years	317	-0.029	0.638	0.308	0.657	0.725	0.779	0.881	1		
20 years	317	-0.01	0.604	0.292	0.616	0.715	0.78	0.84	0.863	1	
30 years	317	-0.008	0.655	0.298	0.677	0.711	0.816	0.806	0.847	0.855	1

Note: Means and standard deviations are in basis points.

Table 3: First stage regression

	Coeff	S.E.	T-stat	P-value
Constant	-66.336	61.552	-1.078	0.281
Last Bucket	0.210***	0.049	4.257	0.000
Last ISIN	0.118***	0.034	3.282	0.001
Remit	0.666***	0.050	13.381	0.000
MA(1)	0.189***	0.070	2.688	0.007
MA(2)	0.138**	0.069	2.013	0.044
R^2	91.770%		F-stat	459.034
N	254			

Note: Regression results for Equation (1). Standard errors in parenthesis. Stars *, ** and *** denote statistical significance at 10%, 5% and 1% respectively.

Table 4: Regression of the announced volume surprises on high-frequenct changes in yields

Panel A			
	Short	Medium	Long
S_t	0.488*	-0.128	0.052
	(0.275)	(0.334)	(0.286)
N	89	73	92
R^2	3.506%	0.208%	0.037%

Panel B			
	Short	Medium	Long
S_t	0.337	-0.426	-0.157
	(0.409)	(0.364)	(0.306)
$S_t \times$ Stress	0.276	1.601*	1.399*
	(0.552)	(0.845)	(0.79)
N	89	73	92
R^2	3.785%	5.083%	3.443%

Note: Results of regressing S_t on $z_t^{(m)}$ in Panel A. In Panel B S_t is also interracted with the stress indicator I_t. In column Short, the sample is restricted to short maturity announcements and $m = 5$ years; in column Medium, the sample is restricted to medium maturity announcements and $m = 10$ years. In column Long, the sample is restricted to long maturity announcements and $m = 20$ years. Estimated intercepts are omitted from the table. Standard errors in parenthesis. Stars *, ** and *** denote statistical significance at 10%, 5% and 1% respectively.

Table 5: The effect of the supply surprises on yields

Panel A						
Nominal yields						
	2 years	3 years	5 years	10 years	15 years	20 years
S_t	0.863*	0.979**	1.083**	1.189**	1.199**	1.156**
	(0.446)	(0.495)	(0.521)	(0.558)	(0.536)	(0.523)
N	254	254	254	254	254	254
R^2	1.111%	1.247%	1.445%	1.563%	1.78%	1.8%
Real yields						
	2 years	3 years	5 years	10 years	15 years	20 years
S_t			0.438	0.695	0.69	0.63
			(0.498)	(0.484)	(0.444)	(0.417)
N			254	254	254	254
R^2			0.16%	0.528%	0.646%	0.601%

Panel B					
Nominal yields					
	2 years	5 years	10 years	15 years	20 years
$S_t\times$ Low stress	0.653	0.871	1.017*	1.021*	0.974*
	(0.433)	(0.548)	(0.608)	(0.59)	(0.578)
$S_t\times$ High stress	2.696	2.93*	2.685*	2.749**	2.741**
	(2.022)	(1.644)	(1.45)	(1.234)	(1.107)
$N^{\text{Low stress}}$	192	192	192	192	192
$N^{\text{High stress}}$	62	62	62	62	62
R^2	1.688%	1.929%	1.848%	2.123%	2.19%
Real yields					
	2 years	5 years	10 years	15 years	20 years
$S_t\times$ Low stress		0.301	0.597	0.569	0.496
		(0.527)	(0.526)	(0.485)	(0.453)
$S_t\times$ High stress		1.635	1.544	1.748	1.794
		(2.112)	(1.591)	(1.331)	(1.255)
$N^{\text{Low stress}}$		192	192	192	192
$N^{\text{High stress}}$		62	62	62	62
R^2		0.297%	0.619%	0.821%	0.838%

Note: Coefficients are the estimated $b^{(m)}$ coefficients from Equation (2) in Panel A, and $b_0^{(m)}$ (low stress) and $b_1^{(m)}$ (high stress) coefficients from Equation (3). Dependent variables are daily changes in the zero-coupon yield curve. Estimated intercepts are omitted from the table. High stress regime is when the CISS index is above its historical 75th percentile. Newey-West standard errors (10 lags) in parenthesis. Stars *, ** and *** denote statistical significance at 10%, 5% and 1% respectively.

Table 6: The effect of the supply surprises on corporate bond indices

	Panel A: Maturity split				
	1-3 years	3-5 years	5-10 years	10-15 years	15+ years
S_t	0.954**	0.674	1.01**	1.079**	0.83*
	(0.456)	(0.521)	(0.509)	(0.502)	(0.472)
N	252	252	252	252	252
R^2	1.223%	0.604%	1.422%	1.473%	0.977%
	Panel B: Rating split				
	AAA	AA	A	BBB	BBB-AAA
S_t	1.271**	1.072**	0.951**	0.982*	-0.289
	(0.5)	(0.465)	(0.451)	(0.517)	(0.262)
N	252	252	252	252	252
R^2	2.582%	1.904%	1.505%	1.307%	0.328%

Note: Coefficients are the estimated $b^{(m)}$ coefficients from Equation (2), where the dependent variables are corporate bond indices obtained from LSEG. Estimated intercepts are omitted from the table. Newey-West standard errors (10 lags) in parenthesis. Stars *, ** and *** denote statistical significance at 10%, 5% and 1% respectively.

Table 7: The effect of the supply surprises on yield components

	Panel A: Nominal decomposition					
	Nominal rate expectations			Nominal term premia		
	3 years	5 years	10 years	3 years	5 years	10 years
S_t	0.071	-0.036	-0.074	0.458	0.773	1.06*
	(0.492)	(0.403)	(0.228)	(0.503)	(0.552)	(0.618)
N	233	233	233	233	233	233
R^2	0.005%	0.002%	0.027%	0.238%	0.593%	0.947%

	Panel B: Real decomposition					
	Real rate expectations			Real term premia		
	3 years	5 years	10 years	3 years	5 years	10 years
S_t		-0.038	-0.081		0.134	0.674
		(0.368)	(0.231)		(0.494)	(0.549)
N		229	229		229	229
R^2		0.003%	0.032%		0.016%	0.457%
	Inflation expectations			Inflation risk premia		
	3 years	5 years	10 years	3 years	5 years	10 years
S_t		0.001	0.008		0.682**	0.481*
		(0.175)	(0.078)		(0.32)	(0.271)
N		229	229		229	229
R^2		0%	0.003%		0.925%	0.663%

Note: Coefficients are the estimated $b^{(m)}$ coefficients from Equation (2), where the dependent variables are the daily changes in the decomposed yield components. Estimated intercepts are omitted from the table. Newey-West standard errors (10 lags) in parenthesis. Stars *, ** and *** denote statistical significance at 10%, 5% and 1% respectively.

Table 8: The effect of the supply surprises on yield components in low and high financial stress

	Real rate expectations		Real term premia	
	5 years	10 years	5 years	10 years
$S_t\times$ Low stress	-0.085	-0.102	0.03	0.539
	(0.361)	(0.228)	(0.524)	(0.596)
$S_t\times$ High stress	0.367	0.092	1.032	1.832
	(1.619)	(0.978)	(1.692)	(1.6)
$N^{\text{Low stress}}$	171	171	171	171
$N^{\text{High stress}}$	58	58	58	58
R^2	0.04%	0.049%	0.102%	0.615%
	Inflation expectations		Inflation risk premia	
	5 years	10 years	5 years	10 years
$S_t\times$ Low stress	-0.004	0.007	0.594*	0.404
	(0.185)	(0.081)	(0.311)	(0.264)
$S_t\times$ High stress	0.049	0.009	1.44	1.149
	(0.482)	(0.261)	(1.211)	(1.028)
$N^{\text{Low stress}}$	171	171	171	171
$N^{\text{High stress}}$	58	58	58	58
R^2	0.003%	0.003%	1.059%	0.812%

Note: oefficients are the estimated $b^{(m)}$ coefficients from Equation (2) in Panel A, and $b_0^{(m)}$ (low stress) and $b_1^{(m)}$ (high stress) coefficients from Equation (3). Dependent variables are daily changes in the decomposed yield components. Estimated intercepts are omitted from the table. High stress regime is when the CISS index is above its historical 75th percentile. Newey-West standard errors (10 lags) in parenthesis. Stars *, ** and *** denote statistical significance at 10%, 5% and 1% respectively.

Table 9: Reaction of credit default swaps to the supply surprises

	2 years	5 years	10 years	30 years
S_t	-0.22	-0.191	-0.132	-0.068
	(0.182)	(0.177)	(0.176)	(0.184)
N	229	229	229	229
R^2	0.247%	0.202%	0.087%	0.019%

Note: Coefficients are the estimated $b^{(m)}$ coefficients from Equation (2), where the dependent variables are credit default swaps written on 2-years, 5-years, 10-years and 30-years UK government bonds, obtained from LSEG. Estimated intercepts are omitted from the table. Newey-West standard errors (10 lags) in parenthesis. Stars *, ** and *** denote statistical significance at 10%, 5% and 1% respectively.

Table 10: Localisation of the supply surprises

Panel A: Linear specification						
	2 years	3 years	5 years	10 years	15 years	20 years
$S_t\times$ Short	6.259	6.393*	5.277*	3.388	3*	2.731*
	(3.902)	(3.594)	(2.789)	(2.109)	(1.776)	(1.485)
$S_t\times$ Long	1.978	2.35*	2.928*	3.602**	3.647**	3.47**
	(1.226)	(1.387)	(1.546)	(1.625)	(1.53)	(1.475)
N	254	254	254	254	254	254
R^2	7.526%	7.091%	5.319%	3.284%	3.374%	3.212%
Panel B: State-dependent specification						
	2 years	3 years	5 years	10 years	15 years	20 years
$S_t\times$ Short $\times$ High stress	14.973	16.088	13.692	10.11	10.38*	10.123*
	(13.463)	(12.442)	(9.636)	(7.091)	(6.076)	(5.437)
$S_t\times$ Long $\times$ High stress	1.22	1.349	1.179	0.96	1.214	1.46
	(1.6)	(1.619)	(1.489)	(1.529)	(1.289)	(1.068)
$S_t\times$ Short $\times$ Low stress	3.567	4.023	3.518	0.692	0.329	0.921
	(2.614)	(3.199)	(3.825)	(4.212)	(4.229)	(4.285)
$S_t\times$ Long $\times$ Low stress	0.66	0.768	0.955	1.217*	1.235**	1.162*
	(0.455)	(0.52)	(0.586)	(0.636)	(0.615)	(0.607)
N	254	254	254	254	254	254
R^2	4.118%	4.286%	3.467%	2.545%	3.013%	3.099%

Note: The table reports estimated coefficients from regressions of the form $\Delta Y_t^{(m)} = a^{(m)} + \beta_t^{Short}\left(I_t^{Short} S_t\right) + \beta_t^{Long}\left(I_t^{Long} S_t\right) + c^{(m)} X_t + \varepsilon_t^{(m)}$, where m denotes maturity, S_t is the supply surprise, and I_t^{Short} and I_t^{Long} are indicators for short (< 7 years) and long-maturity (> 10 years) bond supply announcement respectively. In **Panel A**, coefficients are state-invariant, $\beta_t^{Short} = b^{Short}$ and $\beta_t^{Long} = b^{Long}$. In **Panel B**, coefficients are state-dependent and vary with financial stress, $\beta_t^j = b^{High,j} I_t^{High} + b^{Low,j} I_t^{Low}$. with $j \in \{\text{Short}, \text{Long}\}$, and I_t^{High} and I_t^{Low} denote high- and low-stress states. Intercepts are omitted from the table. Newey–West standard errors with 10 lags are reported in parentheses. *, ** and *** denote statistical significance at the 10%, 5% and 1% levels, respectively.

Online Appendix (Not for publication)

A Details of the high-frequency yield surprises

The time series of the eight yield surprise series $z_t^{(m)}$ is displayed in Figure 1, the ACF in Figure C5 while Table 2 reports their descriptive statistics. The means and autocorrelations are statistically zero, suggesting that the surprisess are not systematic, while the standard deviations range from 0.6 to 1.2 basis points. Maturities above 3 years are strongly correlated, while the 1-year and 2-year series are less so. To ensure that the information released in the event window is important for financial markets, we compare these statistics to the statistics of placebo announcements, i.e., surprises in the same event window recorded on non-announcement days. We record the price changes in the same event window on days with no auction announcements, Bank of England rate decisions, UK GDP or CPI data release or US ISM and Michigan Consumer Sentiment survey results which are released at 3:00 PM GMT. These descriptive statistics are displayed in Table C2 in the Online Appendix.[9] Price moves in the event window on days outside of event days have means close to zero with standard deviations that are mostly lower than on announcement days. This supports our argument that the auction announcements convey important information to market participants.

While the 15-minute event window is short enough to rule out potential factors contaminating the identification of the surprises, it is important to make sure that no other relevant information is released around the announcements. The most important drivers of Treasury yields are macroeconomic news releases, monetary policy decisions, and government bond auction results according to Fleming and Remolona (1997). The times of these events fall outside of our event window, but our results are robust to omitting announcement days that coincide with any of these events.[10] Our event window starts at 3.25 PM. Macroeconomic data releases are published at 7:30 AM or 9:00 AM by the statistical office. Monetary policy announcements are published at 12:00, with a press conference held at 12:30. The DMO is very transparent about releasing public announcements. Market-sensitive information is usually announced between 7.30 AM and 8.00 AM. On auction days the bidding process closes at 10.00 AM or 10.30 AM, with results published shortly after. Post Auction Option Facility[11] results are published at the end of the take-up window closure at 1.00 PM or 2.00 PM. For more information, see DMO (2021).

An illustrative example is the issuance announcement of a 20-year conventional gilt on 2 October 2012. The exact size of the auction was published at 3:30 pm (see the press release in Figure C4 in the Online Appendix). The volume was £1750 million, which was around £400 million lower than the average required auction size implied by the DMO's Remit in place. The release of this information about the relatively low supply was followed by a drop in yields across the curve, as shown in Figure C6. Figure C7 shows a similar example, with an upward jump in yields, following a 39-year bond auction announcement on 20 Mar 2018.

[9] The sample starts on 17 December 2008, due to the availability of the news surprise database.

[10] These results are available upon request.

[11] Since 1 of June 2009, all successful UK gilt auction bidders have the option to purchase up to 10-15% of the bond they have bought, at the published average auction price.

B The Abrahams et al. (2016) ATSM and its estimation

The model assumes that bond yields, the sort rate and the market prices of risks are affine functions of the state variables, which are in turn assumed to be observable. The log prices of a nominal ($P_t^{(\tau)}$) and an inflation-linked ($P_{t,R}^{(\tau)}$) zero-coupon risk-free bonds with remaining time to maturity τ follows:

$$\log P_t^{(\tau)} = A_\tau + B_\tau' X_t \qquad\qquad \log P_{t,R}^{(\tau)} = A_{\tau,R} + B_{\tau,R}' X_t$$

where bond prices and yields are related through

$$y_t^{(\tau)} = -\frac{\log P_t^{(\tau)}}{n} \qquad\qquad y_{t,R}^{(\tau)} = -\frac{\log P_{t,R}^{(\tau)}}{n}$$

The $K \times 1$ vector of state variables or pricing factors X_t assumed to follow an autoregression under the physical measure ($\mathbb{P}$):

$$X_{t+1} - \mu_X = \Phi(X_t - \mu_X) + \nu_{t+1}, \qquad\qquad \nu_{t+1} \sim i.i.d.N(0_{K\times 1}, \Sigma) \tag{B1}$$

The stochastic discount factor is written as:

$$M_{t+1} = exp\Big(- r_t - \frac{1}{2}\lambda_t'\lambda_t - \lambda_t'\Sigma^{-1/2}\nu_{t+1}\Big)$$

where r_t is the nominal short rate and λ_t is $K \times 1$ the vector of risk prices. These are related to the pricing factors as: $\lambda_t = \Sigma^{-1/2}(\lambda_0 + \lambda_1 X_t)$. The short rate follows $r_t = \delta_0 + \delta_1 X_t$. Abrahams et al. (2016) define the parameters of the pricing factor dynamics under the risk neutral measure ($\mathbb{Q}$) as:

$$\begin{aligned}\tilde{\mu} &= (I_K - \Phi)\mu_X - \lambda_0 \\ \tilde{\Phi} &= \Phi - \lambda_1\end{aligned}$$

The system of recursive bond pricing parameters can be obtained once no-arbitrage is imposed (see Ang and Piazzesi (2003)):

$$\begin{aligned}A_\tau &= A_{\tau-1} + B_{\tau-1}'\tilde{\mu} + \frac{1}{2}B_{\tau-1}'\Sigma B_{\tau-1} - \delta_0 \\ B_\tau' &= B_{\tau-1}'\tilde{\Phi} - \delta_1' \\ A_0 &= 0, \quad B_0 = 0_{K\times 0}\end{aligned}$$

Risk neutral counterparts are obtained by setting the price of risk parameters λ_0 and λ_1 to zero. Then, the pricing recursion modifies to:

$$\begin{aligned}
\tilde{A}_\tau &= \tilde{A}_{\tau-1} + \tilde{B}'_{\tau-1}(I_K - \Phi)\mu - \delta_0 \\
\tilde{B}'_\tau &= \tilde{B}'_{\tau-1}\Phi - \delta'_1 \\
\tilde{A}_0 &= 0, \quad \tilde{B}_0 = 0_{K\times 0}
\end{aligned}$$

The price of an inflation-linked bond under under the risk neutral measure ($\mathbb{Q}$) can be written as:

$$\begin{aligned}
\log P^{(\tau)}_{t,R} &= E^{\mathbb{Q}}_t\Big[\exp(-r_t - \cdots - r_{t+\tau-1})\frac{Q_{t+\tau}}{Q_t}\Big] \\
&= E^{\mathbb{Q}}_t\Big[\exp(-r_t - \cdots - r_{t+\tau-1} + \pi_{t+1} + \cdots + \pi_{t+\tau})\Big] \qquad \text{(B2)}
\end{aligned}$$

where $E^{\mathbb{Q}}$ is the expectation operator under the risk neutral measure. Q_t is the price index at time t, and $\pi_t = \ln(Q_t/Q_{t-1})$ is the one period log inflation, related to the pricing factors as $\pi_t = \pi_0 + \pi'_1 X_t$. The inflation-linked bond recursion can be obtained by writing Equation B2 in terms of an inflation-linked bond purchased one period ahead. Taking logs, calculating the expectation, and matching coefficients in the expression for the log bond price yields the recursion:

$$\begin{aligned}
A_{\tau,R} &= A_{\tau-1,R} + B^{\pi\prime}_{\tau-1,R}\tilde{\mu} + \frac{1}{2}B^{\pi\prime}_{\tau-1,R}\Sigma B^{\pi}_{\tau-1,R} - \delta_{0,R} \\
B^{\pi\prime}_{\tau,R} &= B^{\pi\prime}_{\tau-1,R}\tilde{\Phi} - \delta'_1 \\
A_{0,R} &= 0, \quad B_{0,R} = 0_{K\times 0}
\end{aligned}$$

where $\delta_{0,R} = \delta_0 - \pi_0$ and $B^{\pi}_{\tau,R} = B_{\tau,R} + \pi_1$. Similar to nominal bonds, the risk neutral counterparts are given by:

$$\begin{aligned}
\tilde{A}_{\tau,R} &= \tilde{A}_{\tau-1,R} + \tilde{B}^{\pi\prime}_{\tau-1,R}(I_K - \Phi)\mu - \delta_{0,R} \\
\tilde{B}^{\pi\prime}_{\tau,R} &= \tilde{B}^{\pi\prime}_{\tau-1,R}\Phi - \delta'_1 \\
\tilde{A}_{0,R} &= 0, \quad \tilde{B}_{0,R} = 0_{K\times 0}
\end{aligned}$$

where $\tilde{B}^{\pi}_{\tau,R} = \tilde{B}_{\tau,R} + \pi_1$.

A τ-period nominal bond yield can be decomposed into the average expected real short rate, the average expected inflation, real term premium $\mathrm{TP}^{(\tau)}_{t,R}$ and inflation risk premium $\mathrm{IRP}^{(\tau)}_t$:

$$y^{(\tau)}_t = \frac{1}{\tau}\sum_{i=0}^{\tau} E_t(r_{t+i,R} + \pi_{t+i}) + \mathrm{TP}^{(\tau)}_{t,R} + \mathrm{IRP}^{(\tau)}_t \qquad \text{(B3)}$$

The interpretation of $\mathrm{TP}^{(\tau)}_{t,R}$ is the compensation investors require today to hold (real) interest rate risk for the next τ periods, while the interpretation of $\mathrm{IRP}^{(\tau)}_t$ is the compensation

investors require to hold inflation risk for the next τ periods.

The elements of yields can be obtained as the following. We use the risk adjusted counterparts of the pricing recursion coefficients $\tilde{A}_\cdot$ and $\tilde{B}_\cdot$ to calculate the risk adjusted fitted yields. These yields are interpreted as the time t expectation of average future short rates over the next τ periods:

$$\frac{1}{\tau}\sum_{i=0}^{\tau} E_t r_{t+i} = -\frac{1}{\tau}\Big[\tilde{A}_\tau + \tilde{B}'_\tau X_t\Big], \qquad \frac{1}{\tau}\sum_{i=0}^{\tau} E_t r_{t+i,R} = -\frac{1}{\tau}\Big[\tilde{A}_{\tau,R} + \tilde{B}'_{\tau,R} X_t\Big]$$

The difference between the nominal and the real expected short rates is the average expected future inflation over the next τ periods:

$$\frac{1}{\tau}\sum_{i=0}^{\tau} E_t(\pi_{t+i}) = -\frac{1}{\tau}\Big[\tilde{A}_\tau + \tilde{B}'_\tau X_t\Big] - \left(-\frac{1}{\tau}\Big[\tilde{A}_{\tau,R} + \tilde{B}'_{\tau,R} X_t\Big]\right)$$

Term premiums can be obtained by subtracting the expectation component from the fitted yields:

$$TP_t^{(\tau)} = -\frac{1}{\tau}\Big[A_\tau + B'_\tau X_t\Big] - \left(-\frac{1}{\tau}\Big[\tilde{A}_\tau + \tilde{B}'_\tau X_t\Big]\right)$$

for nominal term premium, and similarly for the real term premium $TP_{t,R}^{(\tau)}$. The inflation risk premium is obtained as the difference between the fitted breakeven inflation and the inflation expectation:

$$IRP_t^{(\tau)} = -\frac{1}{\tau}\Big[A_\tau + B'_\tau X_t\Big] - \left(-\frac{1}{\tau}\Big[A_{\tau,R} + B'_{\tau,R} X_t\Big]\right) - \frac{1}{\tau}\sum_{i=0}^{\tau} E_t(\pi_{t+i})$$

The model parameters are estimated following Adrian, Crump, and Moench (2013) and Abrahams et al. (2016). First, we estimate the risk neutral dynamics of the pricing factors by an autoregression. Then, we estimate the sensitivities of bond excess returns to current and past values of the pricing factors. Lastly, we do cross-sectional regressions of excess return sensitivities to lagged pricing factors onto excess return sensitivities to current pricing factors.

State variables are extracted principal components from yields. Following Abrahams et al. (2016), we extract three principal components from month-end zero coupon nominal yields and two principal components from orthogonalized real yields.[12] The factors are shown in Figure C12 in the Online Appendix. Additionally, we extend the state space with a liquidity factor, as Abrahams et al. (2016), where we follow Pflueger and Viceira (2016), Kaminska, Liu, Relleen, and Vangelista (2018) and Bekaert and Ermolov (2023) and use the 5-year inflation-swap spread ISS_t^{5Y} as our liquidity proxy.[13] We standardize it and add

[12]Orthogonalized yields are obtained by purging inflation-linked yields from the nominal principal components, to reduce collinearity among the pricing factors.

[13]This is constructed as the difference between the inflation swap rate ISR_t^{5Y} and the breakeven inflation rate $ISS_t^{5Y} = ISR_t^{5Y} - (y_t^{5Y} - y_{t,R}^{5Y})$. Liquidity premium in inflation swap rates is considered to be negligible,

to each observation the negative of the minimum of the series to ensure the positivity of the index. Then, before calculating inflation-related variables in the model, we subtract the effect associated with the liquidity proxy from yields. The availability of inflation swap data is from the 29th of June 2007 to the 31st of December 2019, and it is displayed in the bottom right panel of Figure C12 in the Online Appendix. It shows a steep increase during the financial crisis and elevated levels during the European debt crisis. The short rate is the Bank of England's official bank rate, inflation is calculated with the monthly RPI series from the Office of National Statistics. We calculate excess returns on eleven nominal maturities of $\tau = 6, 12, 24, \ldots, 120$ months and eight real maturities of $\tau = 60, 66, 72, \ldots, 120$ months. We do the decomposition up to 10 years, as the fit of the model deteriorates at higher maturities.

The ATSM parameters are estimated on monthly data, then, following Adrian, Crump, and Moench (2013) use the monthly parameters on factors extracted from daily yield curve data to obtain the daily yield decomposition. Model fit diagnostics are summarized in Tables C3 and C4 in the Online Appendix. The fit of the model at 10-years maturity is displayed in Figure C13 at the monthly frequency, and in Figure C14 at the daily frequency in the Online Appendix. The mean pricing errors are somewhat larger than in Abrahams et al. (2016), while the standard deviations are similar. Consistent with the relationship between yield and return pricing errors, we find a strong serial correlation in yield pricing errors but not in return pricing errors. The decomposed 10-year expected nominal short rate and nominal term premium are displayed in Figure C15. The decomposition of the 10-year breakeven inflation rate into inflation expectations and inflation risk premium are displayed in Figure C16. The trends in the estimated 10-year nominal term premium are in line with the estimates of Bianchi, Mumtaz, and Surico (2009), Malik and Meldrum (2016), Abrahams et al. (2016) and Kaminska et al. (2018). Expected inflation and inflation risk premium are similar to the estimates of Abrahams et al. (2016), Kaminska et al. (2018) and Bekaert and Ermolov (2023). Expected 10-year average RPI inflation is rather stable around the long run mean, while inflation risk premium shows more variation.

therefore, the rates only represent expected inflation and inflation risk premium (see ECB (2018) and Bekaert and Ermolov (2023). Results with the 10-year maturity are similar.). In the absence of liquidity risk premium in breakeven rates, the spread between the swap rate and the breakeven rate should be zero.

C Proofs of the equilibrium term structure model

Assumption 1. *We assume that the function that controls the sensitivity of bonds to the supply factor $\theta^N(\tau)$ satisfies:*
(i) $\int_0^T \theta^N(\tau) \geq 0$;
(ii) There exists $\tau^* \in [0;T)$ such that $\theta^N(\tau) < 0$ for $\tau < \tau^*$ and $\theta^N(\tau) > 0$ for $\tau > \tau^*$

The first point of the assumption ensures that an increase in β_t does not decrease the total value of bonds supplied to arbitrageurs. The second point allows the possibility that after an increase in β_t the supply of some shorter maturity bonds can decrease, while the total supply of bonds does not decrease. These assumptions ensure that an increase in β_t makes arbitrageurs' equilibrium portfolios more sensitive to inflation and duration risks. We assume $B_{\beta,t}$ is independent of $B_{r,t}$ and $B_{\pi,t}$. Greenwood and Vayanos (2014) considers a case where $B_{\beta,t}$ is correlated with $B_{r,t}$ which is reasonable given their empirical measure of supply which is in levels. In our case, assuming independence corresponds more our to daily frequency supply surprises (in changes space) in Section 3.1.

Lemma 1. *The dynamics of the nominal bond prices and the inflation-linked bond prices are given by*

$$\frac{dP_t^{N,(\tau)}}{P_t^{N,(\tau)}} = \mu_t^{N,(\tau)}dt - A_r^N(\tau)\sigma_r dB_{r,t} - A_\beta^N(\tau)\sigma_\beta dB_{\beta,t} - A_\pi^N(\tau)\sigma_\pi dB_{\pi,t} \tag{B4a}$$

$$\frac{dP_t^{R,(\tau)}}{P_t^{R,(\tau)}} = \mu_t^{R,(\tau)}dt - A_r^R(\tau)\sigma_r dB_{r,t} - A_\beta^R(\tau)\sigma_\beta dB_{\beta,t} \tag{B4b}$$

where instantaneous expected returns $\mu_t^{N,(\tau)}$ and $\mu_t^{R,(\tau)}$ are given by equations (B6a) *and* (B6b)*.*

Proof of Lemma 1. *Prices of the bonds are given by:*

$$P_t^{N,(\tau)} = e^{-[A_r^N(\tau)r_t + A_\beta^N(\tau)\beta_t + A_\pi^N(\tau)\pi_t + C^N(\tau)]} \tag{B5a}$$

$$P_t^{R,(\tau)} = e^{-[A_r^R(\tau)r_t + A_\beta^R(\tau)\beta_t + C^R(\tau)]} \tag{B5b}$$

Applying Ito's lemma to (B5a) *and* (B5b) *using* (8), (7) *and* (9)*, we get:*

$$\begin{aligned}
\frac{dP_t^{N,(\tau)}}{P_t^{N,(\tau)}} =& \big[\dot{A}_r^N(\tau)r_t + \dot{A}_\beta^N(\tau)\beta_t + \dot{A}_\pi^N(\tau)\pi_t + \dot{C}^N(\tau)\big]dt - A_r^N(\tau)dr_t - A_\beta^N(\tau)d\beta_t - A_\pi^N(\tau)d\pi_t \\
&+ \frac{1}{2}\big[(A_r^N(\tau)dr_t)^2 + (A_\beta^N(\tau)d\beta_t)^2 + (A_\pi^N(\tau)d\pi_t)^2\big] \\
=& \big[\dot{A}_r^N(\tau)r_t + \dot{A}_\beta^N(\tau)\beta_t + \dot{A}_\pi^N(\tau)\pi_t + \dot{C}^N(\tau)\big]dt + A_r^N(\tau)[\kappa_r(r_t-\bar{r})dt - \sigma_r dB_{r,t}] \\
&+ A_\beta^N(\tau)[\kappa_\beta\beta_t dt - \sigma_\beta dB_{\beta,t}] + A_\pi^N(\tau)[(\pi_t-\bar{\pi})dt - \sigma_\pi dB_{\pi,t}] \\
&+ \frac{1}{2}\big[(A_r^N(\tau))^2\sigma_r^2 + (A_\beta^N(\tau))^2\sigma_\beta^2 + (A_\pi^N(\tau))^2\sigma_\pi^2\big]dt \\
\frac{dP_t^{R,(\tau)}}{P_t^{R,(\tau)}} =& \big[\dot{A}_r^R(\tau)r_t + \dot{A}^R\beta(\tau)\beta_t + \dot{C}^R(\tau)\big]dt - A_r^R(\tau)dr_t - A_\beta^R(\tau)d\beta_t \\
&+ \frac{1}{2}\big[(A_r^R(\tau)dr_t)^2 + (A_\beta^R(\tau)d\beta_t)^2\big] \\
=& \big[\dot{A}_r^R(\tau)r_t + \dot{A}_\beta^R(\tau)\beta_t + \dot{A}_\pi^R(\tau)\pi_t + \dot{C}^R(\tau)\big]dt + A_r^R(\tau)[\kappa_r(r_t-\bar{r})dt - \sigma_r dB_{r,t}] \\
&+ A_\beta^R(\tau)[\kappa_\beta\beta_t dt - \sigma_\beta dB_{\beta,t}] + \frac{1}{2}\big[(A_r^R(\tau))^2\sigma_r^2 + (A_\beta^R(\tau))^2\sigma_\beta^2\big]dt
\end{aligned}$$

where we arrive to (B4a) *and* (B4b) *if we define* $\mu_t^{N,(\tau)}$ *and* $\mu_t^{R,(\tau)}$ *as:*

$$\begin{aligned}
\mu_t^{N,(\tau)} =& \dot{A}_r^N(\tau)r_t + \dot{A}_\beta^N(\tau)\beta_t + \dot{A}_\pi^N(\tau)\pi_t + \dot{C}^N(\tau) + A_r^N(\tau)\kappa_r(r_t-\bar{r}) + A_\beta^N(\tau)\kappa_\beta\beta_t \\
&+ A_\pi^N(\tau)\kappa_\pi(\pi_t-\bar{\pi}) + \frac{\sigma_r^2}{2}(A_r^N(\tau))^2 + \frac{\sigma_\beta^2}{2}(A_\beta^N(\tau))^2 + \frac{\sigma_\pi^2}{2}(A_\pi^N(\tau))^2 \qquad \text{(B6a)} \\
\mu_t^{R,(\tau)} =& \dot{A}_r^R(\tau)r_t + \dot{A}_\beta^R(\tau)\beta_t + \dot{C}^R(\tau) + A_r^R(\tau)\kappa_r(r_t-\bar{r}) + A_\beta^R(\tau)\kappa_\beta\beta_t \\
&+ \frac{\sigma_r^2}{2}(A_r^N(\tau))^2 + \frac{\sigma_\beta^2}{2}(A_\beta^N(\tau))^2 \qquad \text{(B6b)}
\end{aligned}$$

□

Having derived the price dynamics of the two assets, we can substitute into the budget constraint (11) and solve the arbitrageurs' optimization problem (10).

Lemma 2. *The first order conditions are given by:*

$$\mu_t^{N,(\tau)} - \pi_t - r_t = A_r^N(\tau)\lambda_{r,t} + A_\beta^N(\tau)\lambda_{\beta,t} + A_\pi^N(\tau)\lambda_{\pi,t} \qquad \text{(B7a)}$$

$$\mu_t^{R,(\tau)} - r_t = A_r^R(\tau)\lambda_{r,t} + A_\beta^R(\tau)\lambda_{\beta,t} \qquad \text{(B7b)}$$

where coefficients $\lambda_{i,t}$ are given by:

$$\lambda_{i,t} = a\sigma_i^2 \int_0^T x_t^{N,(\tau)} A_i^N(\tau) + x_t^{R,(\tau)} A_i^R(\tau) d\tau \qquad \text{for } i = r, \beta \tag{B8}$$

$$\lambda_{\pi,t} = a\sigma_\pi^2 \int_0^T x_t^{N,(\tau)} A_\pi^N(\tau) d\tau \tag{B9}$$

Proof of Lemma 2. *Simplifying terms yields* (B4a) *and* (B4b)*. Substituting these into the budget constraint* (11)*:*

$$\begin{aligned} dW_t =& \Big[W_t r_t + \int_0^T x_t^{N,(\tau)} [\mu_t^{N,(\tau)} - r_t - \pi_t] + x_t^{R,(\tau)} [\mu_t^{R,(\tau)} - r_t] d\tau \Big] dt \\ & - \sigma_r \int_0^T x_t^{N,(\tau)} A_r^N(\tau) + x_t^{R,(\tau)} A_r^R(\tau) d\tau dB_{r,t} - \sigma_\beta \int_0^T x_t^{N,(\tau)} A_\beta^N(\tau) + x_t^{R,(\tau)} A_\beta^R(\tau) d\tau dB_{\beta,t} \\ & - \sigma_\pi \int_0^T x_t^{N,(\tau)} A_\pi^N(\tau) d\tau dB_{\pi,t} \end{aligned}$$

Then, the optimization problem can be written as:

$$\begin{aligned} \max_{\left\{x_t^{N,(\tau)}, x_t^{R,(\tau)}\right\}_{\tau \in (0,T]}} & \int_0^T x_t^{N,(\tau)} [\mu_t^{N,(\tau)} - r_t] + x_t^{R,(\tau)} [\mu_t^{R,(\tau)} - r_t] d\tau - \frac{a\sigma_r^2}{2} \left(\int_0^T x_t^{N,(\tau)} A_r^N(\tau) + x_t^{R,(\tau)} A_r^R(\tau) d\tau \right)^2 \\ & - \frac{a\sigma_\beta^2}{2} \left(\int_0^T x_t^{N,(\tau)} A_\beta^N(\tau) + x_t^{R,(\tau)} A_\beta^R(\tau) d\tau \right)^2 - \frac{a\sigma_\pi^2}{2} \left(\int_0^T x_t^{N,(\tau)} A_\pi^N(\tau) d\tau \right)^2 \end{aligned}$$

Point-wise maximization gives the first order conditions (13a) *and* (13b)*.* □

Proof of Theorem 1. *Setting linear terms to zero in* (13a) *and* (13b) *yields ordinary differential equations that we solve with the initial conditions $A_r^N(0) = A_r^R(0) = A_\beta^N(0) = A_\beta^R(0) = A_\pi^N(0) = C^N(0) = C^R(0) = 0$. Identifying terms in r_t gives:*

$$\dot{A}_r^N(\tau) + \kappa_r A_r^N(\tau) - 1 = 0 \tag{B10a}$$

$$\dot{A}_r^R(\tau) + \kappa_r A_r^R(\tau) - 1 = 0 \tag{B10b}$$

Identifying terms in β_t gives:

$$\dot{A}_\beta^N(\tau) + \hat{\kappa}_\beta A_\beta^N(\tau) = Z_r A_r^N(\tau) + Z_\pi A_\pi^N(\tau) \tag{B11a}$$

$$\dot{A}_\beta^R(\tau) + \hat{\kappa}_\beta A_\beta^R(\tau) = Z_r A_r^R(\tau) \tag{B11b}$$

where

$$Z_r = a\sigma_r^2 \int_0^T \theta^N(\tau) A_r^N(\tau) d\tau \tag{B12a}$$

$$Z_\pi = a\sigma_\pi^2 \int_0^T \theta^N(\tau) A_\pi^N(\tau) d\tau \tag{B12b}$$

And $\hat{\kappa}_\beta$ *solves*

$$\hat{\kappa}_\beta = \kappa_\beta - a\sigma_\beta^2 \int_0^T \theta^N(\tau) A_\beta^N(\tau) d\tau \tag{B13}$$

Equilibria in the model exist if the arbitrageurs' risk-aversion coefficient a is below a threshold $\bar{a} > 0$*. As in Greenwood and Vayanos (2014), we focus on that case, and select the equilibrium corresponding to the largest solution for* $\hat{\kappa}$*. For more details see Greenwood and Vayanos (2014). Identifying terms in* π *leads to:*

$$\dot{A}_\pi^N(\tau) + \kappa_\pi A_\pi^N(\tau) - 1 = 0 \tag{B14}$$

The solutions to is (B10a) *and* (B10b) *are* (17a) *and* (18a)*. The solutions to* (B11a) *and* (B11b) *are given by* (17b) *and* (18b)*, with* Z_r*,* Z_π *and* $\hat{\kappa}_\beta$ *are given by equations* (B12a)*,* (B12b) *and* (B13)*. The solution to* (B14) *is* (17c)*. Identifying constant terms in* (13a) *yields*

$$\begin{aligned}
&\dot{C}^N(\tau) - A_r^N(\tau)\kappa_r\bar{r} - A_\pi^N(\tau)\kappa_\pi\bar{r} + \frac{\sigma_r^2}{2}(A_r^N(\tau))^2 + \frac{\sigma_\beta^2}{2}(A_\beta^N(\tau))^2 + \frac{\sigma_\pi^2}{2}(A_\pi^N(\tau))^2 \\
&\quad = a\sigma_r^2 A_r^N(\tau) \int_0^T \zeta^N(\tau) A_r^N(\tau) + \zeta^R(\tau) A_r^R(\tau) d\tau \\
&\quad + a\sigma_\beta^2 A_\beta^N(\tau) \int_0^T \zeta^N(\tau) A_\beta^N(\tau) + \zeta^R(\tau) A_\beta^R(\tau) d\tau \\
&\quad + a\sigma_\pi^2 A_\pi^N(\tau) \int_0^T \zeta^N(\tau) A_\pi^N(\tau)
\end{aligned}$$

The solution to $\dot{C}^N(\tau)$ *is*

$$\begin{aligned}
C^N(\tau) =& \hat{Z}_r \int_0^\tau A_r^N(\tau')d\tau' + \hat{Z}_\beta \int_0^\tau A_\beta^N(\tau')d\tau' + \hat{Z}_\pi \int_0^\tau A_\pi^N(\tau')d\tau' \\
&- \frac{\sigma_r^2}{2} \int_0^\tau (A_r^N(\tau'))^2 d\tau' - \frac{\sigma_\beta^2}{2} \int_0^\tau (A_\beta^N(\tau'))^2 d\tau' - \frac{\sigma_\pi^2}{2} \int_0^\tau (A_\pi^N(\tau'))^2 d\tau'
\end{aligned} \tag{B15}$$

with $\hat{Z}_r$*,* $\hat{Z}_\beta$ *and* $\hat{Z}_\pi$ *given by*

$$\begin{aligned}
\hat{Z}_r &= \kappa_r\bar{r} + a\sigma_r^2 \int_0^T \zeta^N(\tau) A_r^N(\tau) + \zeta^R(\tau) A_r^R(\tau) d\tau \\
\hat{Z}_\beta &= a\sigma_\beta^2 \int_0^T \zeta^N(\tau) A_\beta^N(\tau) + \zeta^R(\tau) A_\beta^R(\tau) d\tau \\
\hat{Z}_\pi &= \kappa_\pi\bar{\pi} + a\sigma_\pi^2 \int_0^T \zeta^N(\tau) A_\pi^N(\tau)(\tau) d\tau
\end{aligned}$$

Identifying constant terms in (13b) *yields*

$$\dot{C}^R(\tau) = A_r^R(\tau)\hat{Z}_r + A_\beta^R(\tau)\hat{Z}_\beta - \frac{\sigma_r^2}{2}(A_r^R(\tau))^2 - \frac{\sigma_\beta^2}{2}(A_\beta^R(\tau))^2$$

with the solution

$$\begin{aligned} C^R(\tau) = \hat{Z}_r \int_0^\tau A_r^R(\tau')d\tau' + \hat{Z}_\beta \int_0^\tau A_\beta^R(\tau')d\tau' \\ - \frac{\sigma_r^2}{2}\int_0^\tau (A_r^R(\tau'))^2 d\tau' - \frac{\sigma_\beta^2}{2}\int_0^\tau (A_\beta^R(\tau'))^2 d\tau' \end{aligned} \tag{B16}$$

□

Proof of Proposition 1. *The effect of a shock to the supply factor to nominal yields is given by:*

$$\frac{\partial y_t^{N,(\tau)}}{\partial \beta_t} = \frac{A_\beta^N(\tau)}{\tau} = \frac{Z_r}{\tau\kappa_r}\Big[\frac{1-e^{-\hat{\kappa}_\beta\tau}}{\hat{\kappa}_\beta} - \frac{e^{-\hat{\kappa}_\beta\tau} - e^{-\kappa_r\tau}}{\kappa_r - \hat{\kappa}_\beta}\Big] + \frac{Z_\pi}{\tau\kappa_\pi}\Big[\frac{1-e^{-\hat{\kappa}_\beta\tau}}{\hat{\kappa}_\beta} - \frac{e^{-\hat{\kappa}_\beta\tau} - e^{-\kappa_\pi\tau}}{\kappa_\pi - \hat{\kappa}_\beta}\Big]$$

First, we show that Z_r and Z_π are positive. Then we show that the expression in the brackets are positive.

From Equation (B10a) (B14) *$A_r^N(\tau)$ and $A_\pi^N(\tau)$ are positive and they are increasing as:*

$$\frac{\partial A_r^N(\tau)}{\partial \tau} = e^{-k_r\tau} > 0$$

Then, we show that $\int_0^T A_r^N(\tau)\theta^N(\tau)d\tau > 0$ as it can be written an:

$$\begin{aligned} \int_0^T A_r^N(\tau)\theta^N(\tau)d\tau &= \int_0^{\tau^*} A_r^N(\tau)\theta^N(\tau)d\tau + \int_{\tau^*}^T A_r^N(\tau)\theta^N(\tau)d\tau \\ &> A_r^N(\tau^*)\int_0^{\tau^*}\theta^N(\tau)d\tau + A_r^N(\tau^*)\int_{\tau^*}^T \theta^N(\tau)d\tau \\ &= A_r^N(\tau^*)\int_0^T \theta^N(\tau)d\tau \geq 0, \end{aligned}$$

where we used Part (ii) of Assumption 1 in the second step and Part (i) if Assumption 1 in the third step. Therefore $Z_r = a\sigma_r^2\int_0^T A_r^N(\tau)\theta^N(\tau)d\tau > 0$ and analogously for Z_π. Then, we can write $A_\beta^N(\tau)$as:

$$A_\beta^N(\tau) = Z_r\int_0^\tau \frac{1-e^{-\kappa_r\hat{\tau}}}{\kappa_r}e^{-\hat{\kappa}_\beta(\tau-\hat{\tau})}d\hat{\tau} + Z_\pi\int_0^\tau \frac{1-e^{-\kappa_\pi\hat{\tau}}}{\kappa_\pi}e^{-\hat{\kappa}_\beta(\tau-\hat{\tau})}d\hat{\tau}$$

which is positive as Z_r and Z_π are both positive. The proof for $A_\beta^R(\tau)$ is analogous.
□

Proof of Proposition 2. *The effect of the supply factor on duration risk is given by:*

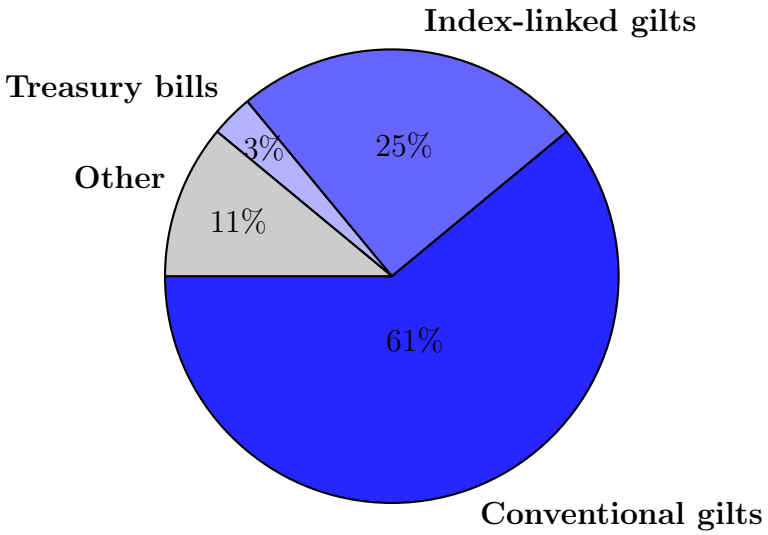
$$\frac{\partial \lambda_{r,t}}{\partial \beta_t} = a\sigma_r^2 \int_0^T \theta^N(\tau) A_r^N(\tau) d\tau$$

which is positive as $\int_0^T \theta^N(\tau) A_r^N(\tau) d\tau > 0$*, as shown in Proof of Proposition 1. The effect of the supply factor on inflation risk is analogous.*
□

Proof of Proposition 3. *This can be seen immediately from* (14), (15), (B11a) *and* (B11b) *as* Z_r *and* Z_π *both increase in* a*.*
□

D Figures and tables

Figure C1: Composition of UK central government sterling debt in December 2019



Source: UK DMO

Figure C2: DMO Financing Remit for the 2015-16 financial year, published the 18th of March 2015

DMO FINANCING REMIT 2015-16: 18 MARCH 2015

1. The DMO's financing remit for 2015-16 has been published today as part of the Budget 2015 announcements. The main points are summarised below.

A) Debt issuance by the DMO

2. The DMO plans to raise £140.4[1] billion in 2015-16, split as follows:

- Outright gilt sales: £133.4 billion.
- Net Treasury bill sales (via tenders): £7.0 billion.

B) Planned gilt sales

3. It is intended that the gilt sales plans will be met through a combination of:

   - £105.2 billion of issuance in 39 auctions; and
   - additional supplementary gilt sales of £28.2 billion (21.1% of total issuance) via a combination of syndicated offerings and, subject to demand, mini-tenders. This will comprise a minimum £24.2 billion via a syndication programme. Any additional sales via syndication can only be of long conventional or index-linked gilts but mini-tenders can be used for issuance of conventional and index-linked gilts across the curve.

4. The planned split of issuance by maturity and type of gilt to be sold via auctions and syndicated offerings is as follows:

   Conventional:

   Short: £33.9 billion (25.4%) in 8 auctions

   Medium: £26.7 billion (20.0%) in 8 auctions

   Long: £37.4 billion (28.0%) in 12 auctions and via syndicated offerings (aiming to raise £28.1 billion by auctions and a current planning assumption of a minimum of £9.3 billion via syndication).

   Index-linked: £31.4 billion (23.5%) in 11 auctions and via syndicated offerings (aiming to raise £16.5 billion by auctions and a current planning assumption of a minimum of £14.9 billion via syndication).

5. The issuance methods to achieve the syndication and mini-tender plans are based on current assumptions. In particular, total financing achieved through each supplementary issuance method will be dependent on market and demand conditions at the time transactions are conducted.

[1] Sales figures in this announcement are in cash terms unless otherwise indicated.

Figure C3: Gilts Operations Calendar for April-May 2015, published the 31st of March 2015

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31 March 2015

PRESS NOTICE

GILT OPERATIONS CALENDAR: APRIL- JUNE 2015

PLANNED SYNDICATED OFFERING OF AN INDEX-LINKED GILT WITH A MATURITY IN THE 30 YEAR AREA OR LONGER IN JUNE 2015

The UK Debt Management Office ("the DMO") is announcing today that the first syndicated offering of the 2015-16 programme will be the sale of an Index-linked gilt with a maturity in the 30 year area or longer. The DMO expects that, subject to market conditions, the sale will take place in the second half of June 2015. Further details of the sale, including the composition of the syndicate, will be announced in due course.

The DMO also announces that in the period April-June 2015 it plans to hold ten outright gilt auctions as well as the syndicated offering, as set out below.

Auction date	**Gilt**	**Further details announced**[1]
Wednesday 8 April	2% Treasury Gilt 2020	Tuesday 31 March
Thursday 16 April	0⅝% Index-linked Treasury Gilt 2040	Tuesday 7 April
Tuesday 21 April	3½% Treasury Gilt 2045	Tuesday 14 April
Wednesday 29 April	2% Treasury Gilt 2025	Tuesday 21 April
Thursday 14 May	2% Treasury Gilt 2020	Tuesday 5 May
Thursday 21 May	4¾% Treasury Gilt 2030	Tuesday 12 May
Wednesday 27 May	0⅛% Index-linked Treasury Gilt 2058	Tuesday 19 May
Tuesday 2 June	2% Treasury Gilt 2025	Tuesday 26 May
Tuesday 9 June	0⅛% Index-linked Treasury Gilt 2024	Tuesday 2 June
Thursday 11 June	3½% Treasury Gilt 2045	Tuesday 2 June

[1] Further to the announcement on 29 January 2015, as of 31 March 2015 the DMO will no longer be declaring a "When Issued" (WI) trading period in cases where the stock being auctioned is a re-opening of an existing line of gilts with a pre-existing ISIN code. Accordingly, the DMO will no longer be issuing a separate WI ISIN code for new tranches of existing gilts. The 29 January announcement can be found at: http://www.dmo.gov.uk/documentview.aspx?docName=/gilts/press/sa290115.pdf

Figure C4: Announcement of the auction of £1,750 million of 20-year Treasury Gilt published on 2 October 2012

United Kingdom
Debt Management Office

Eastcheap Court
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2 October 2012

PRESS NOTICE

AUCTION OF BRITISH GOVERNMENT STOCK

Auction Details

Auction Date	Tuesday, 9 October 2012
Issue and Settlement Date	Wednesday, 10 October 2012
Bidding Convention	Fully paid Bid Price (see Note 1)
Accrued Interest payable with bid	£1.451503 per £100 nominal
Auction Close	10:30am London Time

Details of Security

Title	4¼% Treasury Stock 2032
Amount (nominal) for auction	£1,750 million (fungible with previous issues) (see Note 3)
Nominal outstanding after auction	£32,341.2 million
Maturity Date	7 June 2032 at par
Interest Dates	7 June – 7 December
Parent ISIN Code	GB0004893086
Parent SEDOL Code	0-489-308
Strippable	Yes, since 27 September 2000
Interest Payable	Gross (see Note 2)
Next Interest Date	7 December 2012 (Standard Coupon)

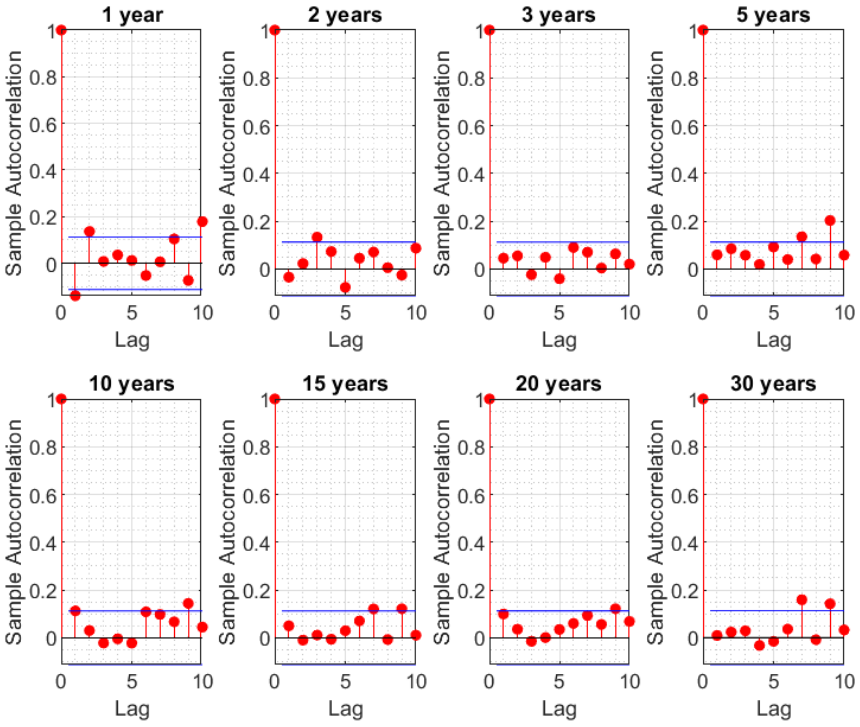
"When Issued" Trading under London Stock Exchange Rule 1530

Commences	3:30pm London Time: 2 October 2012
Closes	Close of business: 9 October 2012
"When issued" ISIN Code	GB00B8J76215
"When issued" SEDOL Code	B-8J7-621
TIDM (for trade reporting)	AUCA

Note 1: Bids may be made on either a competitive or a non-competitive basis. Details of the bidding procedures are set out in the prospectus and in the Information Memorandum. Gilt-edged Market Makers may bid by means of the Bloomberg Bond Auction System to the DMO not later than 10.30 am on Tuesday, 9 October 2012.

Note 2: Holders may elect to have United Kingdom income tax deducted from interest payments, should they so wish, on application to the Registrar, Computershare Investor Services PLC.

Figure C5: Sample autocorrelation function of the high-frequency yield surprises



Note: Sample autocorrelation function of the high-frequency yield surprises series $z_t^{(m)}$ up to ten lags. Blue lines represent two standard errors confidence intervals.

Figure C6: Yield movements (in %) around the announcement a 20-year bond on 2 October 2012

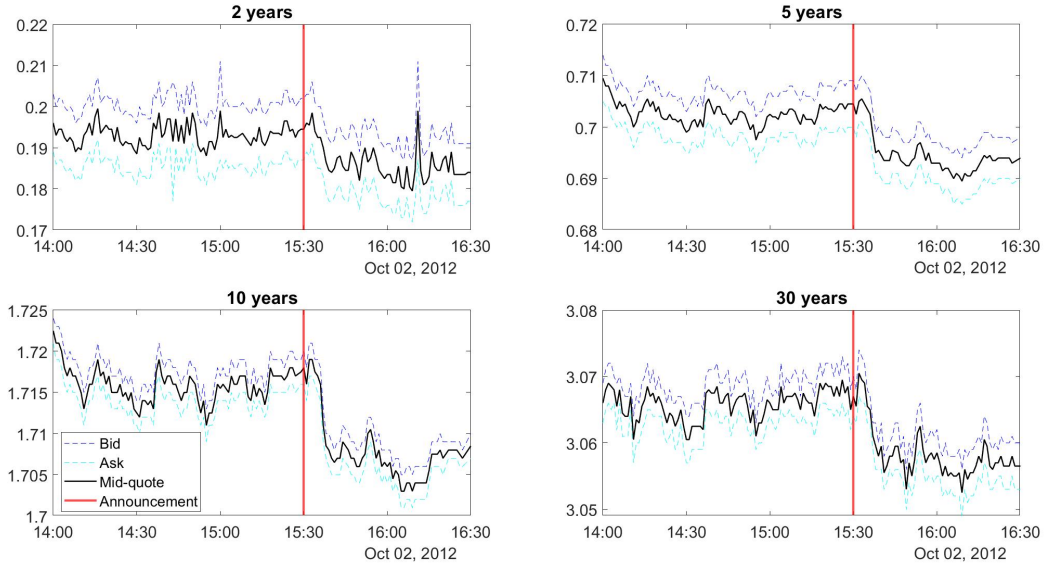


Figure C7: Yield movements (in %) around the announcement a 39-year bond on 20 Mar 2018

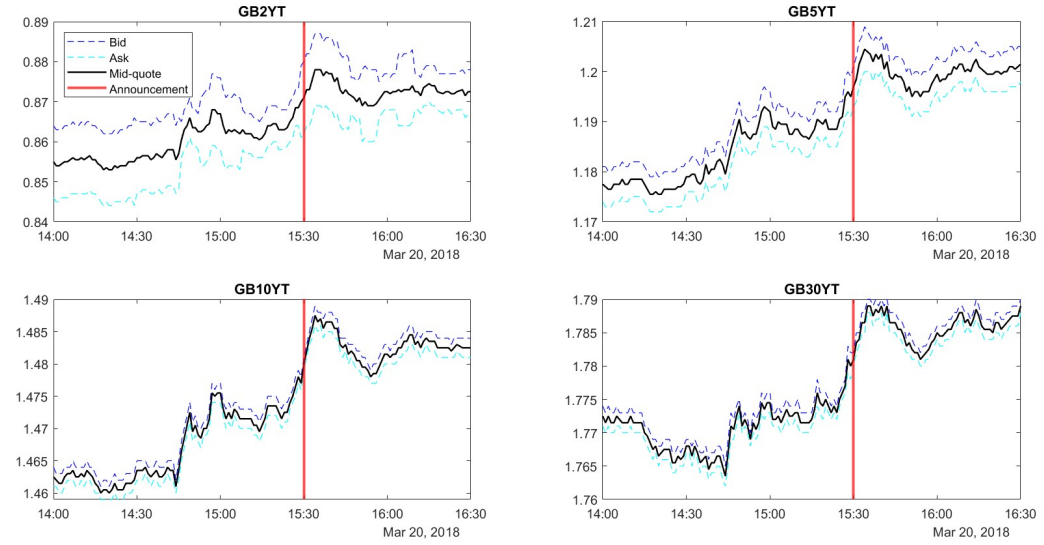
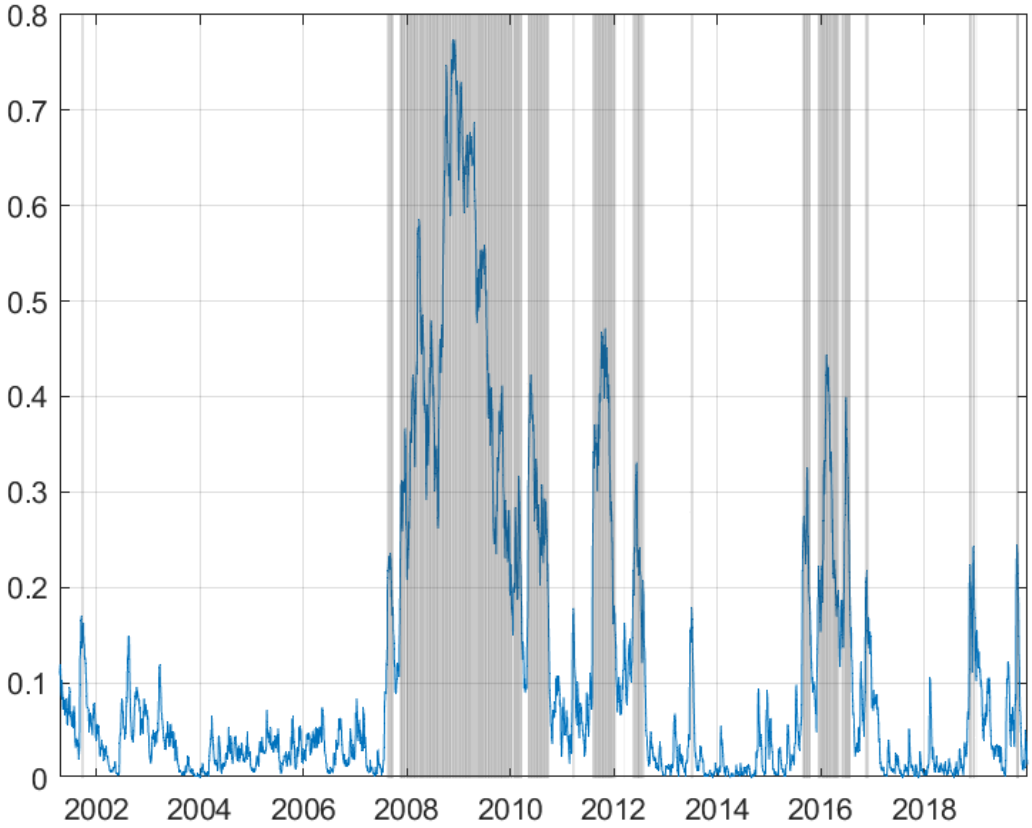
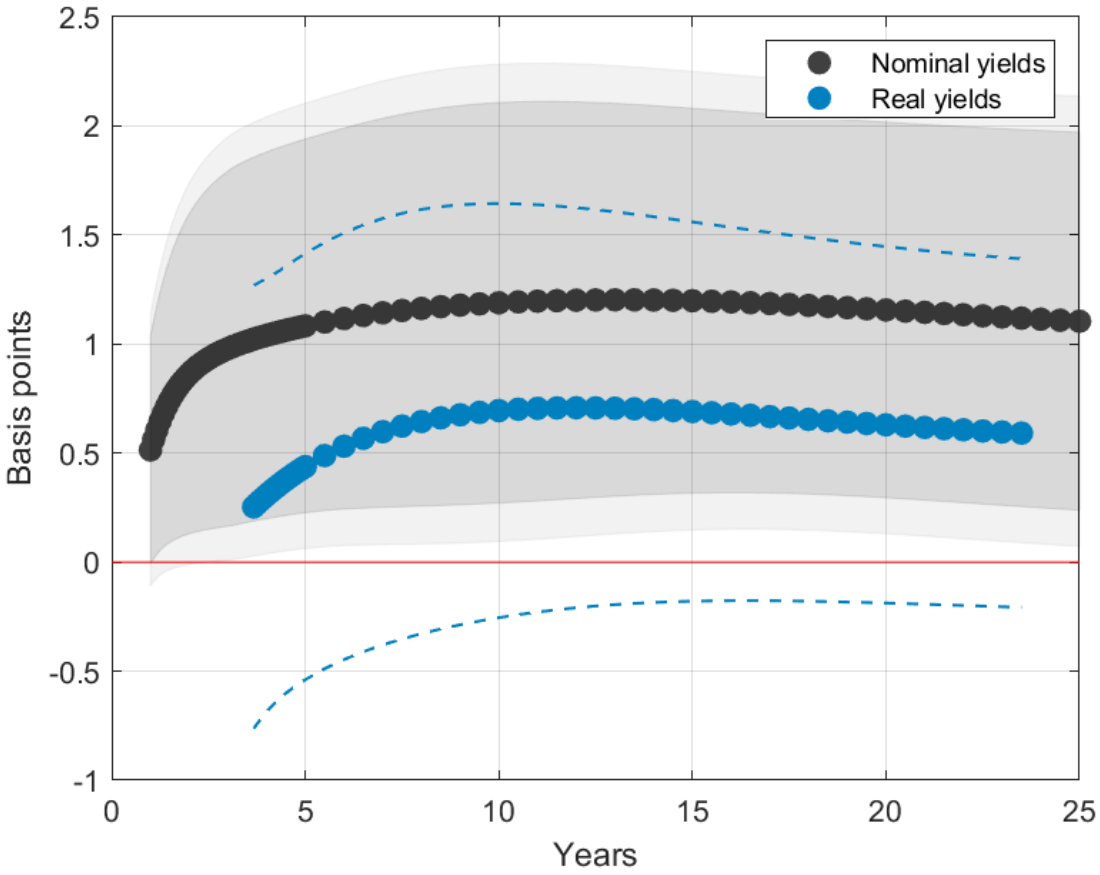


Figure C8: Times series of the UK CISS index with the high stress states shaded in gray



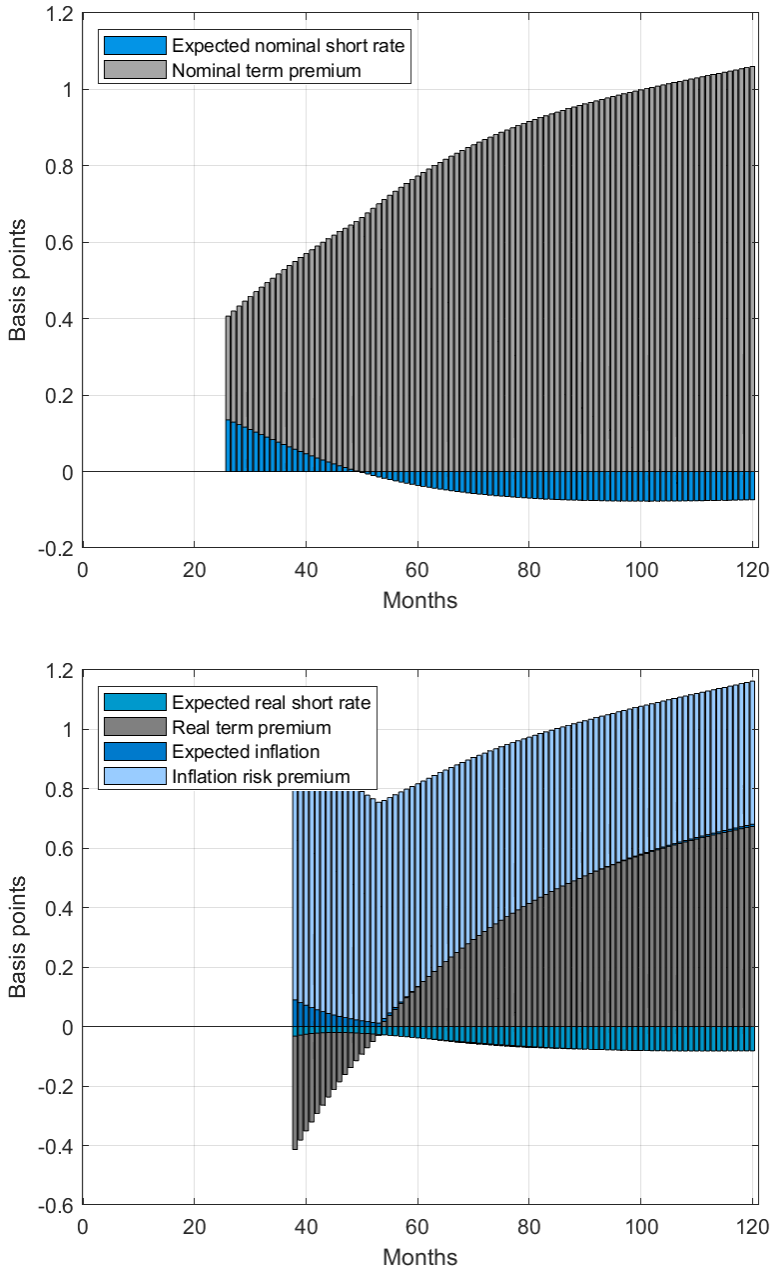
Note: The figure displays the CISS index for the UK. Shaded areas are when the index is above its historical 75th percentile, which is the baseline threshold value to construct the stress indicator I_t.

Figure C9: Reactions of the term structure to the bond supply surprises



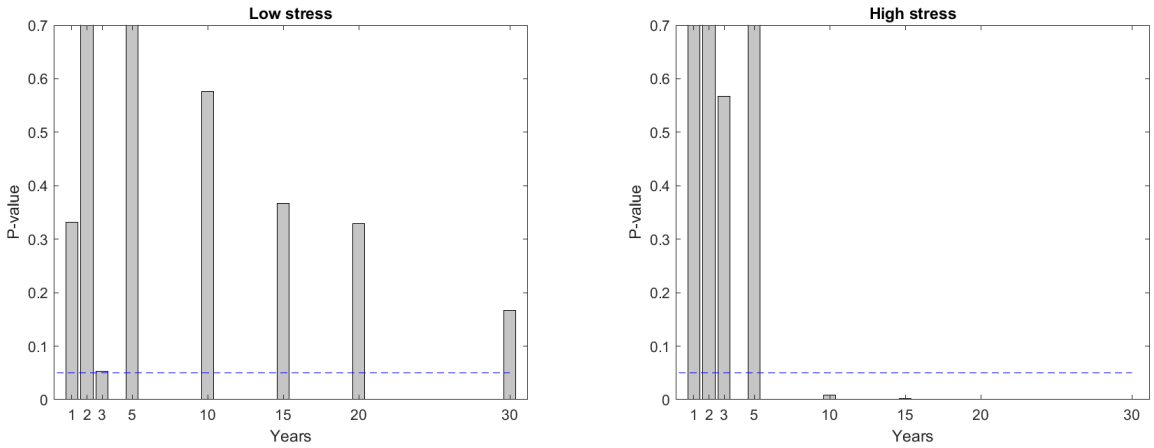
Note: Nodes are the estimated $b^{(m)}$ coefficients from Equation (2), where the dependent variables are the BoE's yield curve series. Dashed lines and shaded areas are 95% (Newey-West, 10 lags) confidence intervals. Sample: 17.04.2001-31.12.2019.

Figure C10: Reaction of yield components to the supply surprises



Note: Bars are the estimated $b^{(m)}$ coefficients stacked from Equation (2). Dependent variables in the top panel are the average expected nominal short rates and the nominal term premium. In the bottom panel: the average expected real short rate, expected inflation, the real term premium and inflation risk premium, all obtained via the ATSM.

Figure C11: Localization results - P-values of coeffcient equality test



Note: Plots of the P-values of coefficient equality tests from regression equation (4). The left panel compares the effects of short-, and long-maturity auctions during low-stress periods; the right panel compares the effects of short-, and long-maturity auctions during high-stress periods. Dashed line denotes the 0.05 value. equality of estimates.

Figure C12: Time Series of the ATSM Pricing Factors

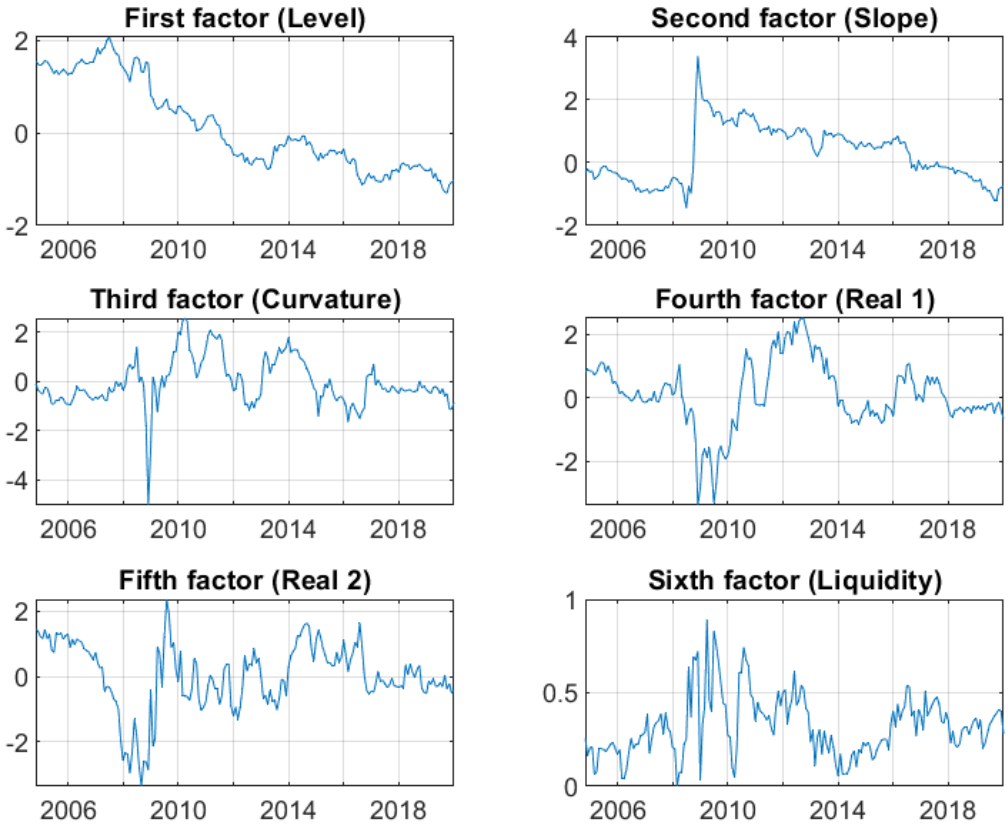


Figure C13: ATSM model fit at 10-years, monthly frequency - nominal yield (left), real yield (right)

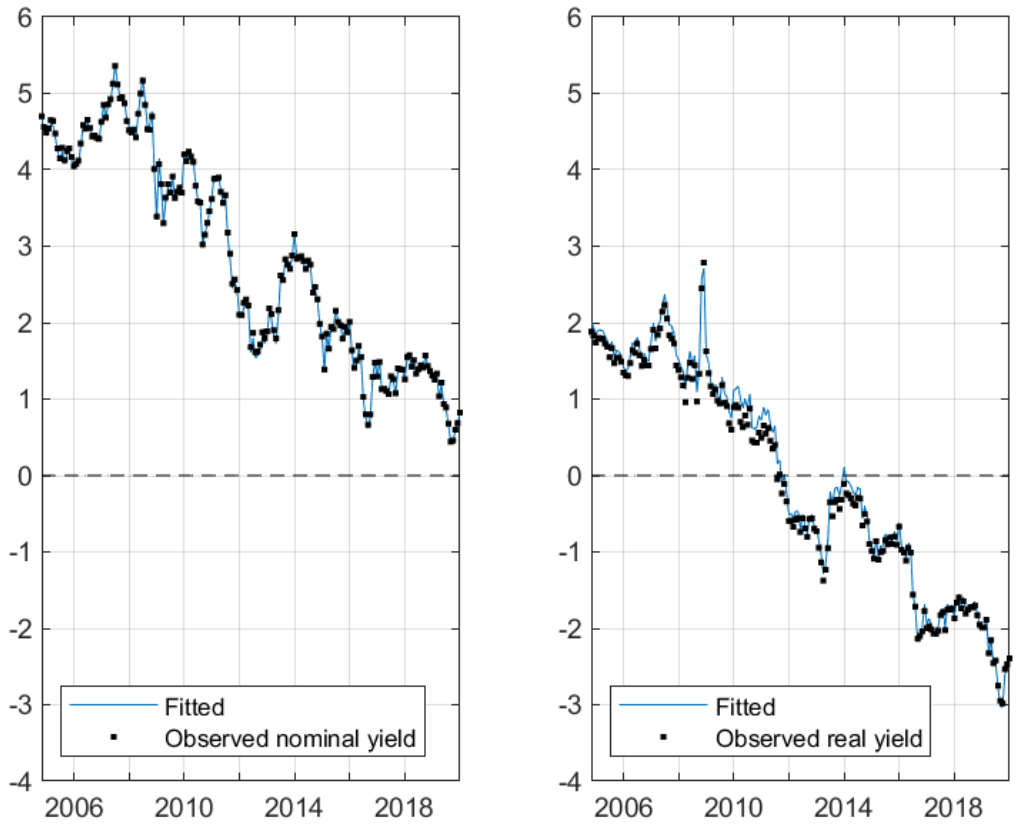


Figure C14: ATSM model fit at 10-years, daily frequency - nominal yield (left), real yield (right)

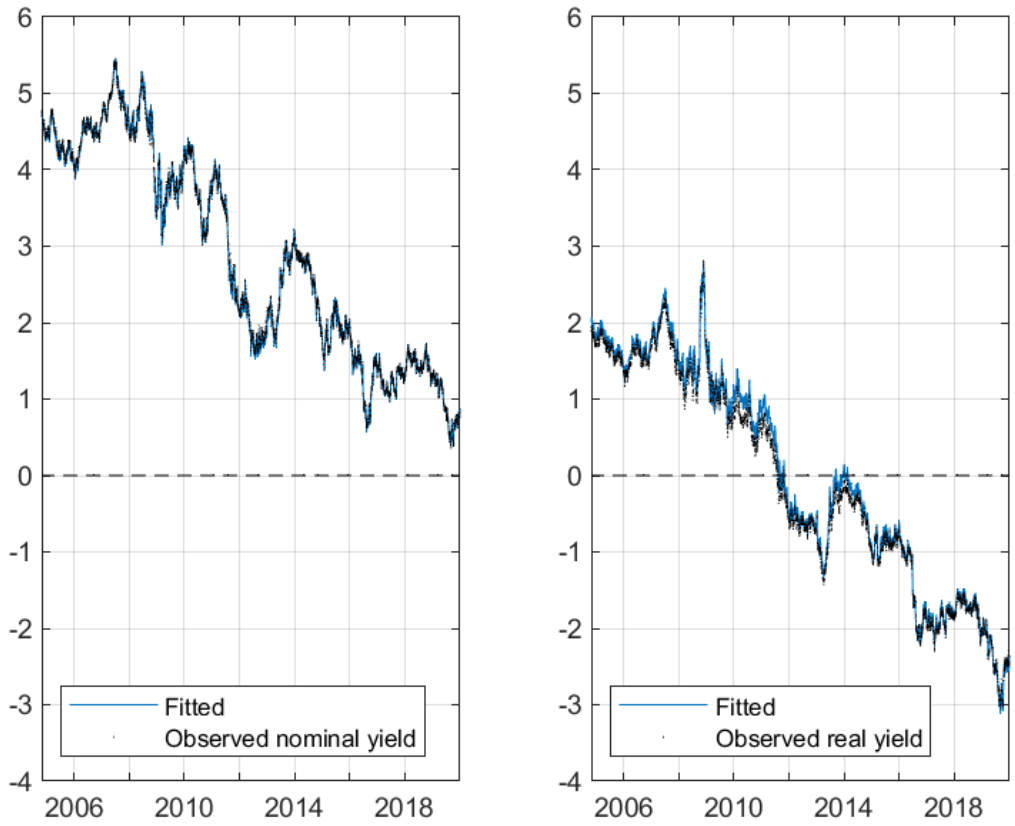


Figure C15: 10-year nominal yield decomposition (left) and real yield decomposition (right) at the daily frequency

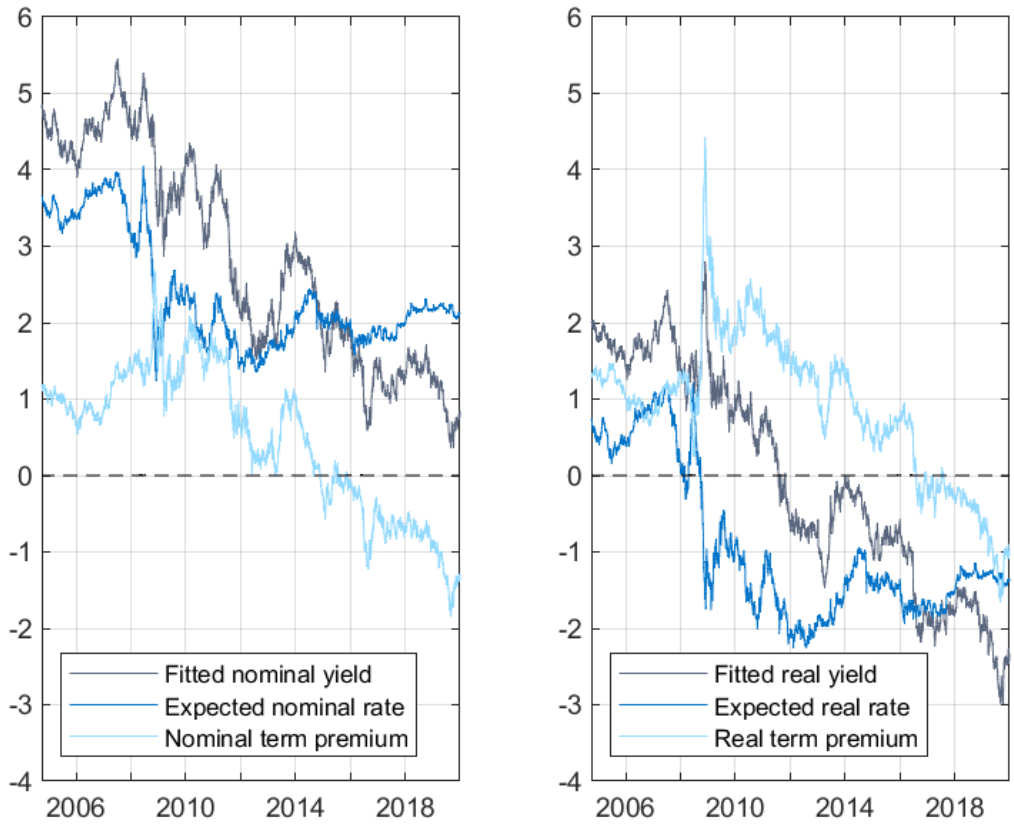


Figure C16: 10-year breakeven inflation rate decomposition at daily frequency

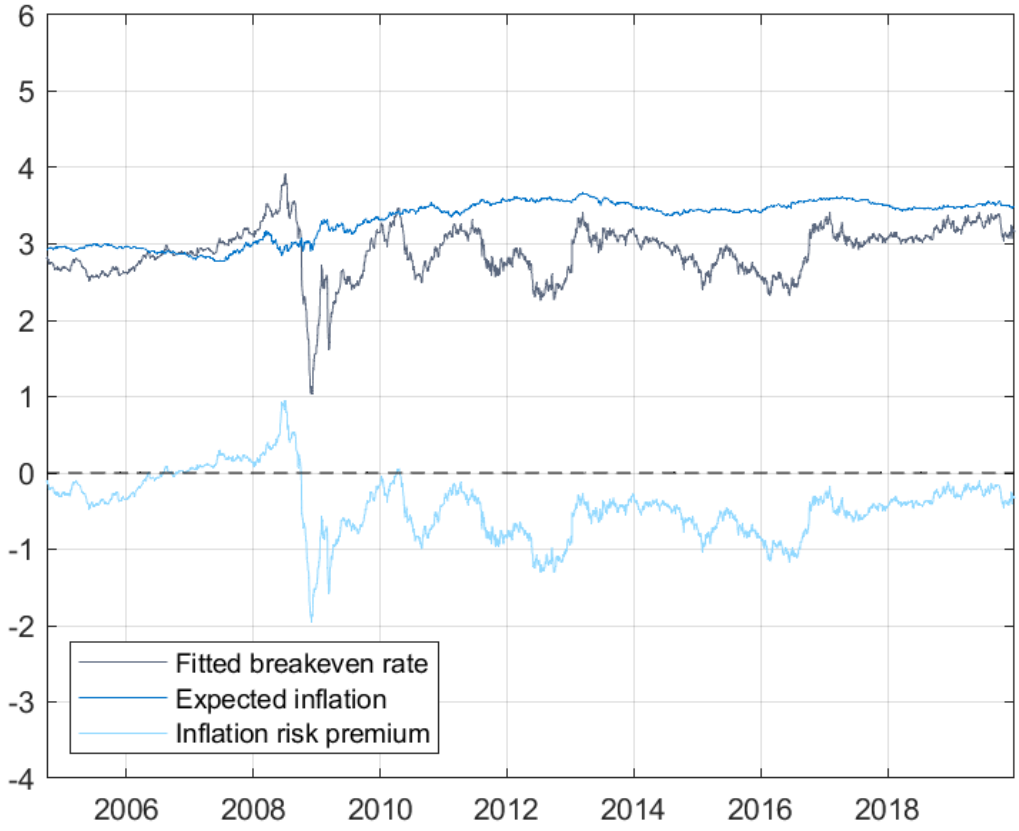
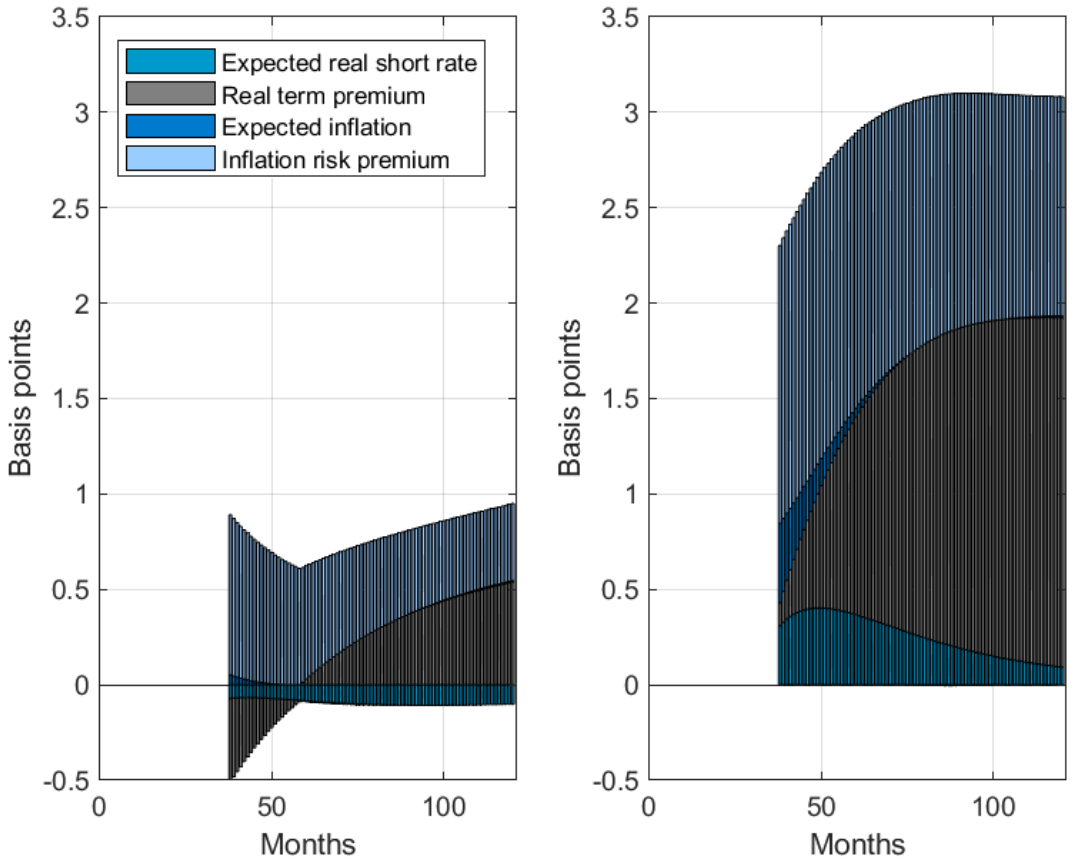


Figure C17: Reactions of yield components in normal times and in high-stress periods



Note: Bars are estimated $b_1^{(m)}$ and $b_0^{(m)}$ coefficients from Equation (3) on yield components obtained via the ATSM. I_t indicates periods with the CISS index above its 75th percentile. Dependent variables are the average expected real short rate, the expected inflation, the real term premium, and the inflation risk premium.

Table C1: The effect of the supply surprises on yields after including control variables

Panel A: Nominal yields - control vars.						
	2 years	3 years	5 years	10 years	15 years	20 years
Coeff (bps)	0.361	0.525	0.764	1.025	1.189*	1.261**
	(0.389)	(0.453)	(0.54)	(0.658)	(0.634)	(0.608)
N	183	183	183	183	183	183
R^2	10.789%	9.248%	6.995%	4.532%	5.081%	5.973%
Panel B: Real yields - control vars.						
	2 years	3 years	5 years	10 years	15 years	20 years
Coeff (bps)	NaN	NaN	-0.516	0.109	0.298	0.363
	(NaN)	(NaN)	(0.759)	(0.676)	(0.574)	(0.509)
N			183	183	183	183
R^2	NaN%	NaN%	4.043%	4.465%	6.257%	7.635%
Panel A: Nominal yields - matching sample						
	2 years	3 years	5 years	10 years	15 years	20 years
Coeff (bps)	0.668	0.834*	1.039*	1.257*	1.368**	1.409**
	(0.412)	(0.482)	(0.57)	(0.674)	(0.653)	(0.63)
N	183	183	183	183	183	183
R^2	1.15%	1.259%	1.369%	1.455%	1.872%	2.133%
Panel B: Real yields - matching sample						
	2 years	3 years	5 years	10 years	15 years	20 years
S_t			-0.272	0.289	0.44	0.478
			(0.741)	(0.662)	(0.591)	(0.557)
N			183	183	183	183
R^2			0.075%	0.095%	0.254%	0.317%

Note: Coefficients are the estimated $b^{(m)}$ coefficients from Equation (2), where the dependent variables are the zero-coupon yields of the Bank of England at different maturities. Independent variables are the supply surprises, an intercept, weekday dummies, the Bank Rate, month-on-month inflation, and Bloomberg GDP and CPI surprises (which are all omitted from the table). Newey-West standard errors (10 lags) in parenthesis. Stars *, ** and *** denote statistical significance at 10%, 5% and 1% respectively. Sample 01.01.2009 in both panels, due to the availability of the economic surprise data.

Table C2: Descriptive statistics of $z_t^{(m)}$ and price changes in the event window on non-event days

	Panel A: Announcement days							
	1 year	2 years	3 years	5 years	10 years	15 years	20 years	30 years
Mean	0.039	-0.082	-0.015	-0.024	-0.052	-0.042	-0.018	-0.019
Std.	1.217	0.692	0.645	0.620	0.630	0.623	0.570	0.589
N	220	220	220	220	220	220	220	220

	Panel B: Non-event days							
	1 year	2 years	3 years	5 years	10 years	15 years	20 years	30 years
Mean	0.017	0.003	0.012	0.004	0.006	0.011	0.012	0.003
Std.	0.941	0.707	0.640	0.612	0.574	0.585	0.560	0.570
N	1971	1971	1971	1971	1971	1971	1971	1971

Note: Descriptive statistics of yield changes in the event window on non-event days. Means and standard deviations (Std.) are expressed in basis points. Non-event days are business days without gilt auction announcements, BoE policy announcements, UK GDP or CPI data releases or US ISM and Michigan Consumer Sentiment Index survey, which are released shortly before our event window. Sample start 01.01.2009 in both panels, due to the availability of the economic surprise data.

Table C3: Fit Diagnostics of the ATSM model on monthly data

	Panel A: Nominal yields			
	3 years	5 years	7 years	10 years
Mean	0.018	0.019	0.009	0.026
Std.	0.031	0.022	0.014	0.032
ρ^y	0.928	0.935	0.858	0.896
ρ^{xy}	0.453	0.227	0.185	0.067

	Panel B: Real yields		
	5 years	7 years	10 years
Mean	0.012	-0.025	-0.058
Std.	0.056	0.062	0.053
ρ^y	0.921	0.94	0.927
ρ^{xy}	0.017	0.003	0.002

Note: Time series properties of the ATSM pricing errors implied by the monthly estimation. "Mean" and "std" refers to the sample mean and standard deviation of yield pricing errors; ρ^y denotes first order sample autocorrelation coefficient of the yield pricing errors, ρ^{xr} denotes first order sample autocorrelation coefficient of the excess return pricing errors.

Table C4: Fit Diagnostics of the ATSM model on daily data

	Panel A: Nominal yields			
	3 years	5 years	7 years	10 years
Mean	0.018	0.02	0.009	0.026
Std.	0.031	0.022	0.014	0.033
ρ^y	0.997	0.996	0.993	0.995

	Panel B: Real yields		
	5 years	7 years	10 years
Mean	0.012	-0.024	-0.057
Std.	0.056	0.062	0.052
ρ^y	0.996	0.997	0.996

Note: Time series properties of the ATSM pricing errors implied by the daily decomposition. "Mean" and "std" refers to the sample mean and standard deviation of yield pricing errors; ρ^y denotes first order sample autocorrelation coefficient of the yield pricing errors, ρ^{xr} denotes first order sample autocorrelation coefficient of the excess return pricing errors.