# Monetary Policy and Corporate Debt Maturity* 

Andrea Fabiani ${ }^{\dagger}$ Luigi Falasconi ${ }^{\ddagger}$ Janko Heineken ${ }^{\S}$


#### Abstract

We show that a policy rate cut lengthens corporate debt maturity. A 1 standard deviation (10 basis points) interest rate cut raises the share of long-term debt by 87 basis points, explaining $20 \%$ of its variation. In the cross-section, large and bond-issuing firms drive this adjustment. We provide a theory to rationalize these findings. A policy rate cut increases demand for long-term bonds due to reach for yield. Financial frictions allow only large, unconstrained firms to benefit by refinancing at lower yields. Empirical evidence on corporate bond issuance and insurer debt security holdings supports our proposed mechanism.


Keywords: Monetary Policy, Corporate Debt, Maturity, Reach for yield, Financial Frictions
JEL codes: E50, G11, G30, G32

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## 1. Introduction

Rising corporate debt exerts notable influence on business cycle fluctuations (Giroud \& Mueller 2017, 2021, Ivashina et al. 2024). In this respect, among other features of corporate debt, the maturity structure affects crucially how companies react to both real and financial shocks (Almeida et al. 2009, Duchin et al. 2010, He \& Xiong 2012, Gomes et al. 2016, Fahlenbrach et al. 2021, Kalemli-Özcan et al. 2022, Jungherr \& Schott 2022, Jungherr et al. 2022).

Interestingly, the maturity structure of corporate debt varies substantially over time. Figure 1 plots a proxy for the US corporate debt maturity structure, namely the share of long-term debt - i.e. debt with maturity above 1 year over total outstanding debt - increasing from $55 \%$ in 1990 to roughly $65 \%$ in 2023, with large swings in between. The existing literature has discussed several drivers, including changes in the supply of long-term Treasuries (Greenwood et al. 2010), excess bond returns (Baker et al. 2003), and the occurrence of recessions (Chen et al. 2021). Yet, the influence of monetary policy has been ignored, despite its well-known implications for non-financial firms' funding costs and investment (see, e.g., Gertler \& Karadi 2015) and the strikingly negative correlation between the Effective Federal Funds Rate (EFFR) and the share of long-term debt in the post-WWII period (Figure 2). ${ }^{1}$

Our study aims to fill this gap. First, we ask whether conventional monetary policy, through interest rate changes, has any impact on the aggregate maturity structure of corporate debt. Put differently, we empirically assess whether monetary policy rates play a meaningful role in explaining the dynamics of the share of long-term debt of the US non-financial corporate sector. Second, we exploit crosssectional heterogeneity in order to identify which companies' debt maturity struc-

[^1]Figure 1: \% of Long-Term Debt - Aggregate Level


This figure shows the evolution of the aggregate percentage share of long-term debt (i.e. with outstanding maturity above 1 year). The black dashed line reports the series in levels, the gray line in first differences. Following Greenwood et al. (2010), long-term debt is defined as the sum of corporate bonds and mortgages and industrial revenues. The remaining short-term corporate debt is proxied by the sum of short-term loans (and advances) and commercial paper.
ture is most affected by monetary policy shocks. By answering these two questions, we uncover novel empirical facts and rationalize them through a theoretical model combining firm-level financial frictions and investors' reach for yield. We further gather empirical evidence supporting the model's main mechanism by analyzing the reaction to variation of monetary policy rates of firms' bond issuance decisions and insurers' debt security holdings.

Given our strict focus on changes in short-term interest rates, we exploit surprises in conventional monetary policy by the Fed over the period 1990-2017. In particular, we use high-frequency monetary policy surprises (Gürkaynak et al. 2005), identified through variations in Fed Funds Future rates in a tight interval after the policy announcement. Importantly, pre-QE conventional monetary policy (i.e., changes in short-term rates through open market operations) does not aim at altering the maturity structure of outstanding marketable Treasuries, differently from QE-era uncoventional monetary policy conducted mainly through the purchase of long-term

Treasuries. Thus, we isolate the direct effects of monetary policy rates from the indirect effects linked to the substitutability between sovereign and corporate long-term bonds (Greenwood et al. 2010), analyzed empirically by Foley-Fisher et al. (2016) in the context of QE. We gather data from various sources. The time-series analysis of corporate debt maturity takes advantage of quarterly data from FED Flows of Funds. At the firm level, we retrieve quarterly financial data for US listed companies from Compustat, complemented with information on the issuance of bonds and syndicated loans from Mergent FISD and Thomson Reuters Dealscan, respectively. Moreover, to analyze the dynamics of the demand for corporate bonds, we access data on holdings and transactions of corporate bonds by the largest class of investors in that market, namely insurance firms, provided by the National Association of Insurance Commissioners (NAIC). ${ }^{2}$

Figure 2: \% of Long-TERM Debt and the Effective Federal Funds Rate


This figure shows the evolution of the aggregate percentage share of long-term debt and the Effective Federal Funds Rate (EFFR). The black dashed line reports the \% of long-term Debt (in levels), measured on the left y-axis. The gray solid line depicts the EFFR, expressed in \% and measured on the right-axis.

Using local projections (Jordà 2005), we document a positive dynamic response of

[^2]the share of long-term (LT) debt to expansionary monetary policy surprises: a policy rate cut lengthens corporate debt maturity. Quantitatively speaking, a 1 standard deviation (10 basis points, b.p.) expansionary monetary policy surprise boosts the share of LT debt by $20 \mathrm{~b} . \mathrm{p}$. on impact. The effect is persistent and increases over time, amounting to $65 \mathrm{~b} . \mathrm{p}$. one year after the cut and peaking at $87 \mathrm{~b} . \mathrm{p}$. three years after the cut. For comparison, the s.d. of the quarterly variation in the LT debt share equals 64 b.p.; the corresponding s.d. figures for the 1-year and 3-year growth rate of the LT debt share are 240 b.p. and 413 b.p., respectively. Hence, at relevant horizons, exogenous innovations in monetary policy rates account for about $20 \%$ of the variation in the share of LT debt.

Next, we test whether such an effect is heterogeneous across companies. In a panel-data framework, we employ again local projections to pin down the relative response to monetary policy surprises of firms with heterogeneous balance-sheet characteristics (Jeenas 2023, Ottonello \& Winberry 2020, Cloyne et al. 2023). This approach allows controlling for firm time-invariant firm characteristics via firm fixed effects and restricting the cross-sectional comparison of the effects of monetary policy surprises within an industry by applying industry*time fixed effects. Practically speaking, we horse-race several balance sheet indicators which have been shown to influence firms' response to monetary policy shocks, including firm size (Gertler \& Gilchrist 1994, Caglio et al. 2021), leverage (Bernanke et al. 1999, Ottonello \& Winberry 2020, Caglio et al. 2021) and liquid assets (Jeenas 2023). It turns out that the influence of monetary policy surprises on corporate debt maturity is especially strong among very large firms, identified as those in the top asset-size quartile of their respective industry-level distribution. Interestingly, a closer inspection of the absolute firm-level response shows that only very large firms lengthen their debt maturity structure after a policy rate cut, whereas smaller firms do not significantly adjust. Hence, virtually the whole aggregate-level adjustment rests on large firms'
response to monetary policy surprises.
Established theories of monetary policy transmission do not account for our findings. In models of monetary policy transmission to firms employing financial frictions (Gertler \& Bernanke 1989, Kashyap \& Stein 1994, Bernanke et al. 1999, Adrian \& Shin 2010, Gertler \& Karadi 2011) an expansionary monetary policy shock predicts a greater relaxation of borrowing terms for small, constrained companies. Hence, implicitly, those models would predict a greater lengthening of the debt maturity structure for smaller companies (though in these theories maturity is generally not explicitly accounted for), contradicting our cross-sectional evidence. ${ }^{3}$ On the other hand, the corporate finance literature focuses on the term spread, delivering seemingly counterfactual predictions relative to our findings: a policy rate cut, widening the term spread (documented by, e.g., Adrian \& Shin 2010), would increase the relative convenience of short-term debt issuance.

We propose a simple theoretical framework that accounts for our aggregate and cross-sectional empirical facts. We augment a standard model featuring short and long-term debt and financial frictions due to moral hazard (Holmström \& Tirole 1998, 2000) with yield-seeking investors (Hanson \& Stein 2015). These investors take long-short positions and care about current portfolio yield rather than expected returns. As a result, in reaction to a policy rate drop, they rebalance their portfolios toward LT debt, to keep their portfolio yield up (i.e., they reach for yield). Hence, they create upward demand pressures on the price of LT debt and contribute to lower term premia, making LT debt a relatively cheaper financing option. Large NFCs accommodate the boost in demand for LT debt, as they have financial flexibility. Importantly, our model nests the standard "balance-sheet channel" of monetary policy, whereby smaller firms benefit from a relaxation of the monetary conditions

[^3]and reduce their reliance on short-term debt. Nonetheless, we provide conditions such that yield-seeking motives dominate the balance-sheet channel in equilibrium and therefore large, unconstrained companies account for the aggregate increase in LT debt.

Our model delivers predictions in line with both the time-series and the crosssectional evidence. Moreover, the model is consistent with empirically grounded interest rate adjustments. The term spread mechanically increases in reaction to a policy rate descent (see, e.g., Adrian et al. 2010) but its term-premium component shrinks (see, among others, Gertler \& Karadi 2015, Hanson \& Stein 2015, Hanson et al. 2021); in our model, the fall in term premia is driven by an investor demand channel. ${ }^{4}$

We conclude with an empirical investigation of the mechanism proposed by the model. To start with, we look at the reaction of corporate bond issuance to monetary policy surprises. In line with our theory, a 1 s.d. expansionary monetary policy surprise boosts the likelihood of issuing bonds for large firms by 38 b.p., relative to smaller firms - an economically meaningful difference, equal to about $6 \%$ of the unconditional likelihood of issuing bonds. In addition, we find that large firms experience a larger reduction in bond spreads at issuance, indicating that investor demand drives the adjustment. In contrast, the reaction of syndicated loan issuance to expansionary monetary policy surprises indicates a relative credit supply expansion for smaller firms. It follows that adjustments in the bond market are mostly responsible for the influence of monetary policy on corporate debt maturity. Indeed, we show that the increase in the share of LT debt due to a policy rate cut is stronger, within large companies, for bond issuers (as compared to large firms inactive in the bond market). Additionally, we further investigate the reasons for larger bond issuance by large firms and find that large firms are more likely to refinance bonds

[^4]early (i.e. long before their scheduled maturity) after an expansionary monetary policy surprise. This is in line with large firms taking advantage of more favorable financing conditions, rather than with a general relaxation of financial constraints. ${ }^{5}$

To test our assumptions on investor behavior, we check if corporate bond holdings by the largest investor class in this market, i.e. insurance companies - holding 38 \% of all US corporate bonds in 2017 (Koijen \& Yogo 2022) — react to monetary policy surprises in line with our theoretical mechanism. In particular, we find that, following an exogenous policy rate cut, insurers increase their net purchases of corporate bonds substantially, i.e. by nearly $50 \%$ as compared to the unconditional average. In addition, the maturity and the share of purchases on the primary market also go up after an expansionary monetary policy surprise. The latter finding is especially important, as it suggests that investor demand is potentially associated with larger bond issuance, as in our model. Lastly, we test whether reach for yield drives these adjustments by insurance firms. To this end, we label as yield-seeking those insurers with relatively high corporate bond portfolio yield, in the spirit of Becker \& Ivashina (2015). In practical terms, starting from security*insurer level holdings data, we compute insurers' corporate bond portfolio yield as the value-weighted average yield across their bond holdings and define as yield-seeking those insurers with above-median corporate bond portfolio yield. It turns out that yield-seeking insurers disproportionately account for the increase in net purchases, both in general and on the primary market.

Our paper contributes to several strands of literature. Our study is the first to provide a systematic analysis of the relation between monetary policy and the maturity structure of corporate debt. Few other studies exploit the ex-ante heterogeneity in firms' debt maturity to explain the heterogeneous effects of monetary policy shocks across firms (Ippolito et al. 2018, Jungherr et al. 2022, Deng \& Fang 2022, Gürkaynak

[^5]et al. forthcoming). Likewise, De Fiore et al. (2011) and Gomes et al. (2016) show that ex-ante aggregate longer-term leverage may amplify the macroeconomic effects of monetary policy. ${ }^{6}$ However, differently from us, these papers do not endogenize the response of debt maturity itself to monetary policy shocks.

In a closely related paper to ours, Foley-Fisher et al. (2016) show a lengthening of corporate bond maturity following a specific unconventional policy by the Fed, namely the Maturity Expansion Program, which consisted in a rebalancing of the Fed portfolio towards longer-term Treasuries. We focus on short-term interest rate changes, allowing us to separate the effects of monetary policy from those associated with the outstanding volume of LT Treasuries (Greenwood et al. 2010), which varies over the business cycle for reasons unrelated to Fed policies. Bräuning et al. (2020) analyze the relation between the policy rate and the maturity of new bond issues and find that a rate cut is associated with lower maturity at issuance. These findings do not contradict our results since their focus lies only on the intensive margin, i.e. on the marginal adjustments in bond maturity, conditional on issuing a bond. Differently from Bräuning et al. (2020), we focus on the corporate debt maturity structure (i.e., the share of long-term debt to total debt outstanding), rather than solely on the maturity of new debt issues. Indeed, when we analyze bond issuance to test our model mechanism, we show that adjustments to the extensive margin in long-term bond (i.e., with maturity above 1-year) issuance are crucial to explaining both aggregate and cross-sectional responses of the corporate debt maturity structure to monetary policy surprises.

In addition, as we exploit a mechanism based on investors' reach for yield (Stein 2013, Becker \& Ivashina 2015), our work relates to papers highlighting the importance of those motives for the reallocation of investment across different securities after monetary policy shocks (Di Maggio \& Kacperczyk 2017, Lian et al. 2019, Daniel

[^6]et al. 2021, Hanson et al. 2021). We follow Hanson \& Stein (2015) and introduce yield-seeking investors, who tilt their portfolios towards LT debt securities after an interest rate decrease. ${ }^{7}$ Our main contribution to this literature is to link yieldoriented investors' reaction to monetary policy to firms' issuance of LT bonds and, ultimately, to the evolution of the maturity structure of corporate debt.

Several other papers look at financial channels for monetary policy different from a standard credit channel (Bernanke \& Gertler 1995), i.e., mechanisms other than bank intermediation. Among others, Foley-Fisher et al. (2016), Grosse-Rueschkamp et al. (2019), Giambona et al. (2020) investigate adjustments in the bond market in response to unconventional monetary policy. Darmouni et al. (2022) exploit frictions in the Eurozone bond markets and find, similarly to us, that access to the bond market is linked to greater firm sensitivity to interest rate surprises. We innovate by focusing on debt maturity and by highlighting a bond channel of conventional interest rate policy connected to reach for yield in financial markets.

Finally, our paper speaks to the literature on the determinants of the maturity structure of corporate debt (Diamond 1991, Barclay \& Smith Jr 1995, Baker et al. 2003, Berger et al. 2005, Faulkender 2005, Greenwood et al. 2010, Custódio et al. 2013, Diamond \& He 2014, He \& Milbradt 2016, Badoer \& James 2016, Choi et al. 2018, 2021) including cyclical factors (Xu 2018, Mian \& Santos 2018, Chen et al. 2021, Poeschl 2023). We contribute by discussing explicitly the role of monetary policy and its interaction with firm financial frictions and investors' reach for yield.

The rest of the paper is organized as follows. Section 2 describes the data. In Section 3, we report the baseline empirical findings. To explain them, we elaborate a model, presented in Section 4. Next, Section 5 empirically tests the model's mechanism. Section 6 briefly concludes.

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## 2. DATA

Our empirical analysis covers the period from 1990Q1 to 2017Q4. We employ several datasets, which we describe below. In addition, we report summary statistics in Appendix Table A1.

### 2.1. Time-SERIES DATA

For the time-series analysis of the LT debt share, we use the Federal Reserve Flow of Funds (FoF), tracking financial flows throughout the U.S. economy. We use quarterly data on the credit market liabilities of the non-farm, non-financial, corporate business sector. We focus on the share of total corporate debt with maturity above 1 year, which we label as the share of LT debt. Following Greenwood et al. (2010), we define short-term debt as the sum of commercial paper and loans with maturity no longer than 1 year. On the other hand, LT debt is given by the sum of corporate bonds, mortgages and industrial revenue bonds. We depict the resulting series, given by the fraction of LT debt over the sum of short-term and LT debt, in Figure 1. The black line, referring to such variable in levels, displays an upward trend over the period of interest, with the share of LT debt increasing from roughly $55 \%$ to $65 \%$. Throughout the paper, we look at the impact of policy rate surprises on the dynamics of the LT debt share and therefore we compute its growth over different horizons. In Figure 1, the solid gray line shows the evolution of the quarterly growth rate; its mean equals 0.15 p.p., as can be seen from the summary statistics in Appendix Table A1. We also report summary statistics for the cumulative growth rate of the LT debt share over longer horizons, used for pinning down impulse response functions through local projections. In general, the variable $\Delta L T-D e b t_{t+h}$ is computed as the difference between the LT debt share as of year-quarter $t+h$ and $t-1$, for $h=0,1,2, \ldots, 20$. For brevity, we show summary statistics only for up to 1-year growth of the LT debt share. Both the mean and the volatility of the growth
rate increase along with the length of the horizon over which they are computed.
Our first proxy of changes in the policy rate is the quarterly variation in the effective federal funds rate (EFFR), $\triangle E F F R_{t}$, gathered from FRED, reflecting the evolution of business cycle conditions and the associated policy response by the Fed. We additionally lever an alternative conventional proxy for unexpected monetary (interest rate) policy surprises. We borrow data from Gürkaynak et al. (2005), who build a widely-used measure of interest rate surprises based on the change in Fed Funds Futures rate in 30-minute windows around the policy announcement. We label the resulting quarterly series of monetary surprises, obtained summing up all the surprises occurring in a given quarter, as $\varepsilon_{t}$. We plot both series in Appendix Figure A1. The post-2009 period is characterized by lower variation in interest rates, a well-known fact linked to the implementation of the zero lower bound (ZLB) after the Great Financial Crisis. For this reason, whenever possible, we check that our results survive the exclusion of the post-2009 period. Moreover, while the two series display a large extent of correlation, the surprises are in general an order of magnitude smaller than $\triangle E F F R_{t}$. Concretely, while a 1 s.d. change in $\triangle E F F R_{t}$ equals 45 b.p., a 1 s.d. change in $\varepsilon_{t}$ amounts to 10 b.p. (see Appendix Table A1).

Finally, we collect from FRED several other macro-economic indicators that we use as controls, including: the annual GDP growth rate and inflation rate; an indicator for recession periods; the $10 \mathrm{y}-3 \mathrm{~m}$ US sovereign term spread, and the BAA-AAA corporate spread. We borrow term premia estimates at different horizons (1, 2, 5, and 10 years) from Adrian et al. (2013). From Thomson Reuters Datastream, we also download information on the share of Treasuries with maturity above 20 years and compute its quarterly growth rate $\left(\Delta L T-\right.$ Treas $\left._{t-1}\right)$. We use this variable to control for gap-filling driven variations in the corporate LT debt share (Greenwood et al. 2010, Badoer \& James 2016).

### 2.2. Firm-LEVEL DATA

Our primary source for firm-level data is Compustat, containing balance sheet information on the universe of US listed companies. Compustat couples well with monetary policy surprises, since it provides quarterly balance sheet information, whereas most other firm-level datasets contain annual balance sheets. On the other hand, the information on debt is rather limited. In fact, we can only distinguish the fraction of total debt with maturity above 1 year - in line with our macroeconomic data from FoF - without additional information on the maturity profile of existing liabilities.

Our sample includes 12,655 companies. We are mainly interested in the variation over time of the share of debt with maturity above 1 year. The variable $\Delta L T-$
 to $t+h$. In the Appendix Table A1 we report summary statistics for $h=0,1, \ldots, 4$. Across the different horizons, the distribution is centered around 0 , as suggested by the median value. Nonetheless, the extent of heterogeneity is remarkable.

An important variable throughout our analysis is firms' asset size, our preferred proxy for financial constraints, i.e. access to bond financing. There are large differences in firms' asset size (expressed in logs of 1990q1 millions of US\$). From the unconditional summary statistics in Appendix Table A1, one interquartile variation reflects an increase in asset size by nearly 358 p.p.. Clearly, this figure conflates cross-sectional and time-series variation. However, our interest in asset size is mostly cross-sectional. To this end, we look at the within-industry (3-digit SIC) time-varying distribution of total asset size and define a dummy variable, Large $_{f, t-1}$, with value 1 if a company is in the fourth quartile and 0 otherwise. The focus on the fourth quartile is justified by the fact that - as shown in Section 3.2 - it is within this class of firms that the debt maturity structure responds to interest rate surprises. Interestingly, Figure 3 Panel A shows that the LT debt share is unevenly distributed
between firms and increases in firm size. As a matter of fact, for firms in the first asset-size quartile, LT debt accounts for roughly $50 \%$ of total debt. For firms in the fourth quartile, it increases to nearly $80 \%$. Throughout the rest of the paper, we refer to companies in the top-size quartile of their industry distribution as "large" companies. We also gather additional information from Compustat on other firm-level controls such as leverage (debt to assets), liquid assets, and sales growth.

Figure 3: \% of LT Debt \& Debt Issuance across Firms


We sort firms according to the quartiles of their (3-digit SIC industry-level) asset size distribution. Panel A shows the average share of LT Debt across companies in different asset-size quartiles (black bars); the complement to 1 gives the average share of short-term debt (gray bars). Panel B and C depict, respectively, the number of bonds and syndicated loan issues by firms in the top-size quartile (gray bars) and by all other firms in lower quartiles (black bars).

To test our mechanism, we retrieve data on the issuance of bonds and syndicated loans (both with maturity above 1 year) from Mergent Fixed Income Secu-
rities Database (FISD) ${ }^{8}$ and Thomson Reuters Dealscan, ${ }^{9}$ respectively. Our sample comprises 2,858 bond issuers, labeled by the dummy variable Bond $d_{f, t}$, with value 1 from the first period a firm issues a bond onward, and value 0 before. First, we analyze bond issuance on the extensive margin, tracked by the dummy $\mathbb{1}$ (BondIssue) $f_{f, t+h}$, equal to 1 if a firm $f$ issues bonds in year-quarter $t+h$ and 0 otherwise, $h=0,1, \ldots, 20$. On average, the likelihood of a current year-quarter new issuance is $6.77 \%$, suggesting that bond issuance is relatively lumpy and infrequent. This average increases slightly but steadily over future horizons, reflecting the fact that older and/or larger companies tend to issue bonds relatively more frequently. Relatedly, Figure 3 Panel B plots the number of bond issuances per year-quarter and splits them depending on whether they are conducted by a large company, or not. We note that the share of new issuances by large companies is disproportionately large. That is, while such firms account (by construction) for roughly $1 / 4$ of the firms in our sample, they represent about $60 \%$ of new bond issuances. This is prima-facie evidence that large companies are much more active in the corporate bond market. Next, as we are interested in the response of financing costs to monetary policy, we retain data on the annualized bond spread at issuance, BondSpread $_{f, t+h}$, i.e. the difference between the bond yield at issuance and the Treasury yield at comparable tenor, which equals 6 p.p., on average. Regarding syndicated loans, our analysis focuses on 7,603 firms active in this market and covers 30,723 new loan issuances. The likelihood of a current year-quarter new loan issuance (described by the dummy variable LoanIssue $_{f, t}$ ) is $9.3 \%$. Interestingly, in Figure 3 Panel C, the disproportion

[^8]in syndicated loan issuance by large firms (as compared to smaller ones) is less pronounced than for bonds. Moreover, the mean firm-level loan all-in-drawn spread, AIDSpread $_{f, t}$ is about $1.74 \% .{ }^{10}$

### 2.3. Insurance Firms' Bond Holding Data

Insurance firms play a key role in bond markets. Koijen \& Yogo (2022) show that since 1945 they have been the largest investors in publicly-traded and privatelyheld corporate bonds. In 2017, insurance firms held around $40 \%$ of US corporate bonds. Public corporate bonds constitute around $40 \%$ of life insurers' portfolios, with private corporate bonds making up another $40 \%$. For property and casualty insurers, public corporate bonds constitute approximately $30 \%$ of their portfolios, with private corporate bonds making up another $20 \%$.

We collect data on US insurers' holdings and transactions of corporate bonds from the National Association of Insurance Commissioners (NAIC), available between 2006 and 2022. The original data on bond transactions are available at the daily level and describe the par value of the corporate bonds sold and bought by insurers, respectively, and their maturity. We collapse this information at the quarterly frequency, by taking the sum of the net purchases (i.e., purchases minus sales) and computing their value-weighted average maturity in a quarter. We also retrieve information on the share of purchases in the primary market. Moreover, we employ yearly insurer-security holdings data to categorize yield-seeking investors. In practice, we compute the weighted-average corporate bond portfolio yield as $r_{i, y}=\sum_{j} \omega_{i, j, y} * r_{j, y}$, where $\omega_{i, j, y}$ is the par-value weight of security $j$ in the end-of-the-year corporate bond portfolio of insurer $i$ and $r_{j, y}$ is its yield (retrieved from Thomson Reuters Datastream). ${ }^{11}$ Next, we define as yield-seeking those insurers

[^9]with above-median corporate bond portfolio yield. Formally, a yield-seeking investor is labeled with a unitary value for the dummy $H Y_{i, y}=\mathbb{1}\left(r_{i, y}>r_{y}^{50}\right)$, where $r_{y}^{50}$ is the average corporate bond yield in year $y$. Detailed variable definitions can be found in the notes for Appendix Table A1.

Overall, our quarterly panel is populated by 4,526 insurers. The average quarterly net purchase amounts to $1.9 \%$ of total assets, with an average maturity of approximately 10 years. The average share of total purchases in the primary market is $15 \%$. In addition, there is significant variation across insurers' weighted-average corporate bond portfolio yield, $r_{i, y}$. The median and mean values are $1.9 \%$ and $2.2 \%$, respectively, with a one s.d. of about $10 \%$. Hence, our yield-seeking dummy $H Y_{i, y}$ reflects notable differences in corporate bond portfolio yield across insurers. Finally, we also report the values of insurers' $\log$ assets and regulatory capital ratio, which we use as control variables.

## 3. Empirical Analysis

This section presents the baseline empirical findings. First, we present the aggregatelevel analysis. Next, we investigate cross-sectional differences across firms.

### 3.1. Time-Series Analysis

We apply local projections (Jordà 2005) to study the response of the share of LT debt to changes in monetary policy interest rate. In particular, we estimate separately the following regressions through OLS:

$$
\begin{equation*}
\Delta_{h} y_{t+h}=\beta_{1, h} \Delta M P_{t}+\Gamma_{h} X_{t-1}+u_{t, h} \tag{1}
\end{equation*}
$$

for $h=0,1, \ldots, 20$. The dependent variable, $\Delta_{h} y_{t+h}$, is given by the cumulative variation in the share of LT debt between year-quarters $t-1$ and $t+h . \Delta M P_{t}$ is the
chosen proxy of changes in monetary policy, namely either the quarterly variation in the EFFR, $\triangle E F F R_{t}$, or the high-frequency surprises from Gürkaynak et al. (2005), $\varepsilon_{t}$. Hence, the coefficient $\beta_{1, h}$ provides the effect of a time- $t$ unitary change in $\Delta M P_{t}$ on the cumulative $h$-quarter-ahead cumulative growth in the share of LT debt. $X_{t-1}$ is a vector of lagged macro-controls, including variables that might simultaneously have an influence on $\Delta_{h} y_{t+h}$ and on the current policy rate variation. In particular, $X_{t-1}$ includes: the annual GDP growth rate and inflation rate; the quarterly variation in the $10 \mathrm{y}-3 \mathrm{~m}$ term spread, in the corporate spread and in the share of Treasuries with maturity above 20 years; a recession dummy. Finally, $u_{t, h}$ is a robust error term.

Figure 4 reports the impulse response function (IRF) obtained from the OLS estimation of the coefficients $\beta_{1, h}$ in Equation 1. In particular, Panel A shows the results from the estimation exercise employing the simple variation in the EFFR, $\triangle E F F R_{t}$. The IRF is calibrated to a $25 \mathrm{~b} . \mathrm{p}$. cut - i.e., a decrease of the short-term policy rate - and displays the $90 \%$ confidence interval around the point estimates. Clearly, an interest rate cut boosts the share of LT debt. The effect is persistent and, while effective on impact, peaks up 3 years after the shock, and does not fade away throughout the considered 5-year time window. Such large degree of persistence might reflect the significant persistence of $\triangle E F F R_{t}$ along monetary policy cycles (see, e.g. Adrian et al. 2010). That said, the effect is economically meaningful. For instance, a 25 b.p. interest rate descent in year-quarter $t$ implies a cumulative increase in the share of LT debt by 42 b.p. over a one-year horizon.

The result holds as well when using the high-frequency monetary policy surprises. Figure 4 Panel B shows the impulse-response functions resulting from the estimation of Equation 1 with $\Delta M P_{t}=\varepsilon_{t}$, calibrated to a 1 s.d. expansionary surprise (i.e., a 10 b.p. cut). The analysis validates the positive effect of an interest rate decrease on the LT debt share. Again, the effect is economically significant. It turns out that a 1 s.d. expansionary monetary policy surprise boosts the share of LT debt

Figure 4: Monetary Policy and Debt Maturity Structure: Aggregate RESPONSE
(a) Quarterly Variation in Effective Federal Funds Rate: $\Delta E F F R_{t}$

(B) GÜRKAYNAK ET AL. (2005) SURPRISES: $\varepsilon_{t}$


This figure depicts the IRF of the aggregate-level share of LT debt to a change in monetary policy. Both Panel A and Panel B show the coefficients $\beta_{1, h}$ from the estimation of Equation $1, h=0,1, \ldots, 20$, with the proxy of monetary policy change (i.e., $\Delta M P_{t}$ ) given, respectively, by the quarterly variation in the Effective Federal Funds Rate, $\triangle E F F R_{t}$ and the Gürkaynak et al. (2005) monetary policy surprises, $\varepsilon_{t}$. In Panel A, the IRF is calibrated to a $\triangle E F F R_{t}=-25$ b.p.. In Panel B, the IRF is calibrated to a $1 \mathrm{~s} . \mathrm{d}$. expansionary monetary policy surprise, i.e. $\varepsilon_{t}=-10 \mathrm{~b}$.p.. The solid line reports the point estimates for $\beta_{1, h}$; the dashed lines the $90 \%$ confidence intervals. In both panels, we apply robust standard errors.
by 20 b.p. on impact. The effect is persistent and increases over time, amounting to $65 \mathrm{~b} . \mathrm{p}$. 1 year after the impact and peaking at 87 p.p. 3 years after the impact. For comparison, the s.d. of the quarterly variation in the LT debt share equals 64 b.p.; the corresponding s.d. figures for the 1-year and 3-year growth rate of the LT
debt share are 240 b.p. and 413 b.p., respectively. Hence, at relevant horizons, monetary policy surprises account for roughly $20 \%$ of the variation in the share of LT debt. Further, the high-frequency shocks exert a less persistent effect (as compared to $\triangle E F F R_{t}$ ), decaying in 14 quarters.

Finally, we check in Appendix Figure A2 that our findings are robust to restricting the sample to the period between 1990q1 and 2008q4, which we label as pre-crisis. In fact, if anything, trimming the sample to exclude the post-crisis (i.e., ZLB) period magnifies the effect of monetary policy shocks on the share of LT debt; this is reassuring given that most of the variation in the policy rate occurs before 2008.

### 3.2. Firm-LEVEL ANALYsis

In order to investigate cross-sectional differences in the relation between monetary policy and corporate debt maturity structure, we employ a panel version of Jordà (2005)'s local projections. In practical terms, we estimate by OLS the following set of equations:

$$
\begin{align*}
\Delta_{h} y_{f, t+h}= & \beta_{1, h} \text { Large }_{f, t-1}+\beta_{2, h} \text { Large }_{f, t-1} * \Delta M P_{t}+  \tag{2}\\
& +\Gamma_{h} X_{f, t-1}+\mu_{f}+\mu_{s, t}+q_{f, t}+u_{f, t+h}
\end{align*}
$$

for $h=0,1, \ldots, 20$. The dependent variable, $\Delta_{h} y_{f, t+h}$, is given by the cumulative variation of the share of LT debt of firm $f$ between year-quarters $t-1$ and $t+h$. Most importantly, the model includes the full interaction of the monetary policy change, $\Delta M P_{t}$, and a dummy for large companies, Large $_{f, t-1}$, with value 1 for companies in the top quartile of the respective industry asset-size distribution (and with value 0 otherwise). The main coefficient of interest is $\beta_{2, h}$, capturing the relative response of large companies (as compared to smaller ones) to policy rate changes.

We augment the model with a vector of firm controls, $X_{f, t-1}$, comprising the (lagged) share of liquid assets, leverage, and the quarterly growth of sales. These
control variables are also fully interacted with $\Delta M P_{t}$. By doing so, we horse-race our channel (based on firm size as a proxy for bond financing constraints) against other layers of heterogeneity that have been found to influence firms' response to monetary policy shocks. Using an empirical approach similar to ours, Jeenas (2023) shows that companies with a relatively lower share of liquid assets respond more to monetary policy shocks, whereas Ottonello \& Winberry (2020) find that distance to default (proxied by financial leverage) matters as well for firms' reaction to monetary policy. Sales growth is meant to capture variation in firms' profitability. Furthermore, we interact $\operatorname{Large}_{f, t-1}$ with the usual set of macroeconomic controls to avoid that $\beta_{2, h}$ reflects the contemporaneous response of large companies to other variations in macroeconomic conditions, which may correlate with the interest rate policy decisions by the Fed. We saturate the model with firm and industry*year-quarter fixed effects, i.e. $\mu_{f}$ and $\mu_{s, t}$, respectively. The former set of dummies controls for all time-invariant heterogeneity at the level of the firm; the latter absorbs time-varying (observed and unobserved) shocks that are common to firms in a given (3-digit SIC) industry. The application of such fixed effects implies that our coefficient of interest $\beta_{3, h}$ is identified by: i) within-firm variation over time, i.e., changes in the response of the share of LT debt by an otherwise identical firm when it is large, as compared to when it was small; ii) cross-sectional variation across firms in a given industry and point in time. $q_{f, t}$ is a categorical variable indicating the firm-specific fiscal quarter of the year (Ottonello \& Winberry 2020). Finally, $u_{f, t+h}$ is an error term, which we double-cluster at the firm and industry*year-quarter level.

The relative adjustment of large companies estimated through Equation 2 does not allow us to understand the overall response of both large and smaller companies. In fact, Equation 2 is saturated with industry*year-quarter fixed effects, which span out time-series variation common across all firms. Hence, we additionally estimate the following model separately for firms in different size quartiles:

$$
\begin{equation*}
\Delta_{h} y_{f, t+h}=\beta_{1, h} \Delta M P_{t}+\Psi_{h} X_{f, t-1}+\mu_{f}+q_{f, t}+v_{f, t+h} \tag{3}
\end{equation*}
$$

In practice, we estimate a model that exploits only time variation and hence describes the absolute change in the share of LT debt after a policy rate change. In fact, we do not use year-quarter fixed effects (nor any subtler version of them), while we keep using firm fixed effects to control for unobserved time-invariant firm heterogeneity. $X_{f, t-1}$ is the usual vector of macro and firm-level controls and $q_{f, t}$ absorbs differences in fiscal calendars across firms.

The trade-off between the two models is clear: Equation 2 precisely estimates the cross-sectional differences across firms, as it controls for time-varying common heterogeneity within narrowly defined industries. On the other hand, Model 3 pins down the absolute variation in LT debt share due to changes in the policy rate. Hence, it serves the purpose of better understanding the connection between firmlevel and time-series findings.

### 3.2.1. Results

Figure 5 plots the impulse response function (IRF) obtained from the estimation of the parameters $\beta_{2, h}$ - for $h=0,1, \ldots ., 20$ - from Equation 2. Irrespective of whether one looks at the model employing the simple variation in the EFFR (Panel A) or the high-frequency surprises (Panel B), large companies expand LT debt more when the policy rate goes down, relative to smaller corporations. Put differently, large companies react more in line with the aggregate-level evidence shown above. For understanding the level of the absolute response, however, we additionally estimate Equation 3 within different size quartiles. In Figure 6, we report the resulting IRFs. We plot estimates obtained using the exogenous high-frequency surprises $\varepsilon_{t}$, a con-
vention we maintain throughout the paper in the interest of brevity. ${ }^{12}$ The plots indicate that only large companies adjust, whereas smaller firms' LT debt share is generally insensitive to monetary policy.

We focus on asset size as it turns out to be the key firm-level attribute to explain cross-sectional differences across firms. We report in Appendix Tables A2 and A3 additional coefficients from the estimation of the baseline firm-level model using either $\varepsilon_{t}$ or $\Delta E F F R_{t}$, respectively. In particular, we show the horse race with the other balance-sheet characteristics employed as firm-level controls. First, companies tend to increase the share of LT debt when sales jump; nonetheless, the interaction of such dynamics with monetary policy is not significant. Moreover, the share of LT debt goes down when firms hold relatively more liquid assets, intuitively reflecting maturity matching of assets and liabilities. Also in this case, however, the share of held liquid assets does not influence the relation between debt maturity structure and monetary policy. Likewise, the interaction between the different proxies of changes in the policy rate and leverage is generally insignificant.

To interpret the economic significance of these effects, note that the IRFs in Figures 5 and 6 are calibrated to a 1 s.d. (i.e., 10 b.p.) expansionary monetary policy surprise. The jump in the share of LT debt by large companies peaks up 8 quarters after the impact when it amounts to roughly 45-75 b.p. (depending on whether one takes as a reference the adjustment in Figure 5 or in Figure 6, respectively). ${ }^{13}$ Interestingly, the described size of the effect is comparable - at relevant horizons with that observed at the aggregate level.

Finally, for robustness, we estimate Equation 2 over the pre-crisis, pre-ZLB period (i.e., from 1990 to 2008) so to restrict our analysis to a time window $w$ th substantial

[^10]Figure 5: Monetary Policy and Debt Maturity Structure - Relative Response of Large Companies
(a) Quarterly Variation in Effective Federal Funds Rate: $\Delta E F F R_{t}$

(B) GÜRKAYNAK ET AL. (2005) SURPRISES: $\varepsilon_{t}$


This figure depicts the relative response of large firms to a monetary policy shock (as compared to smaller firms). Both Panel A and Panel B show the coefficients $\beta_{2, h}$ from the estimation of Equation $2, h=0,1, \ldots, 20$, with the proxy of monetary policy change (i.e., $\Delta M P_{t}$ ) given, respectively, by the quarterly variation in the Effective Federal Funds Rate, $\triangle E F F R_{t}$ and the Gürkaynak et al. (2005) monetary policy surprises, $\varepsilon_{t}$. In Panel A, the IRF is calibrated to a $\triangle E F F R_{t}=-25$ b.p.. In Panel B, the IRF is calibrated to a $1 \mathrm{~s} . \mathrm{d}$. expansionary monetary policy surprise, i.e. $\varepsilon_{t}=-10$ b.p.. The solid lines report the point estimates; the dashed lines the $90 \%$ confidence intervals. In both panels, we doublecluster standard errors at the firm and industry*year-quarter level.
interest rate variations. Results in Appendix Figure A3 - using either $\Delta E F F R_{t}$ or $\varepsilon_{t}$ - confirm that the baseline findings are both qualitatively and quantitatively nearly identical.

Figure 6: Monetary Policy and Debt Maturity Structure - Absolute FIRM-LEVEL RESPONSE


This figure depicts the absolute response of companies in different size quartiles to a $1 \mathrm{~s} . \mathrm{d}$. cut in the Gürkaynak et al. (2005) monetary policy surprise. Formally, the gray solid line reports the size-quartile-specific point estimates for $\beta_{1, h}$ from Equation 3; the dashed gray line the respective $90 \%$ confidence intervals. Standard errors are double-clustered at the firm and industry*year-quarter level.

## 4. MODEL

### 4.1. SETUP

The model builds on Holmström \& Tirole (1998, 2000), amended by introducing a continuum of firms, instead of one single firm, and yield-seeking behavior on the investor side. In the section IB. 1 of the Internet Appendix, we discuss thoroughly the different assumptions and modeling choices, which we briefly present here.

Our economy lasts three periods $(t=0,1,2)$ and is populated by a continuum of firms and investors. Each firm is endowed with capital $A$ - a proxy for firm size - and a project. $A$ is heterogeneous across firms and distributed uniformly on the interval $[0, I]$, with $I$ denoting the fixed investment scale for a firm's project in period 0 . Each project is subject to a reinvestment shock $\rho$ in period 1 , drawn from an exponential distribution $f(\rho)=\chi e^{-\chi \rho}$, for $\rho \in[0, \infty)$. If the reinvestment need is
not met, the project is liquidated and does not generate any payoff in period $2 .{ }^{14}$
Each project generates a riskless short-term pay-out $r$ in period 1. In period 2, conditional on reinvestment, the project yields $R$ in case of success, and zero in case of failure. The likelihood of success depends on firms' behavior: if they exert effort, it equals $p_{h}$ (without loss of generality, we set $p_{h}=1$ ); if they shirk, success materializes with probability $p_{l}<1$, but the owner enjoys private benefits $B$. Additionally, there is aggregate risk: with probability $1-\delta$, all firms get a pay-out of zero in period 2. Firms do not have access to storage technology and are protected by limited liability.

Investors are perfectly competitive; a subset of them "reach for yield", in the sense of Hanson \& Stein (2015). They care about current portfolio returns, as opposed to other (rational) investors who consider the whole stream of current and future portfolio returns. Therefore, yield-seeking investors take more risk than rational investors when interest rates are low to generate higher yields.

The short-term interest rates between periods 0 and 1 , and 1 and 2, are set exogenously by a monetary authority and denoted as $i_{1}$ and $i_{2}$, respectively. Investors borrow at rates $i_{1}$ and $i_{2}$, while firms receive funds intermediated by the investors. In the following, we lay out formally how we model firms and investors.

Firms. We follow Holmström \& Tirole $(1998,2000)$ in modeling firm financing with moral hazard. We assume that firms can credibly commit to a contract stating that the project is continued into period 2 whenever the stochastic reinvestment is sufficiently small, i.e. $\rho \leq \rho^{*}$, and terminated otherwise. The continuation threshold $\rho^{*}$ is chosen by the firm.

To finance the gap between the initial investment and the endowment, firms issue short-term and long-term (LT) debt. A riskless short-term bond is sold at price $P_{s}=$

[^11]$1 /\left(1+i_{1}\right)$ at time 0 and is promised to yield 1 in expectation in period 1 , while a long-term bond is sold at price $P_{l}$ in time 0 and yields 1 in period 2 if the project is successful. The amounts of issued short-term and LT debt are $d_{s}$ and $d_{l}$, respectively.

The timing is as follows. In period 1, short-term creditors are compensated out of earnings $r$, as the firm must repay $d_{s}$. Next, the firm draws a reinvestment need from $f(\rho)$. If the decision is not to invest - i.e., if $\rho>\rho^{*}$ - then the firm abandons the project and the owner consumes what is left, whereas long-term bondholders do not receive any compensation. If the project instead continues and succeeds in period 2 , the firm gets $R-d_{l}$, while long-term bondholders receive their compensation $d_{l}$. In case of failure the firm is liquidated at zero value and neither bondholders, nor the firm, receive anything. Hence, to induce the firm to exert effort, the following condition must hold: ${ }^{15}$

$$
\begin{equation*}
R-\frac{B}{\Delta p} \geq d_{l}(A) \tag{IC}
\end{equation*}
$$

where $\Delta p=p_{h}-p_{l}$. This incentive compatibility constraint states that the repayment cannot be too large, or else the firm will shirk.

Moreover, limited liability and riskless short-term debt imply:

$$
\begin{equation*}
\rho^{*}(A) \leq r-d_{s}(A) \tag{LL}
\end{equation*}
$$

i.e., the firm cannot be asked to meet the investment shock with other funds than those stemming from the project returns. Finally, the firm must raise enough funding to finance the project in the first place:

$$
\begin{equation*}
\frac{d_{s}(A)}{1+i_{1}}+P_{l}(A) d_{l}(A) \geq I-A \tag{IR}
\end{equation*}
$$

Investors. Firms borrow in bond markets populated by a continuum of investors. Investors are heterogeneous and we denote their type as $j$. Investors have zero initial

[^12]wealth and construct long-short positions to maximize:
\[

$$
\begin{equation*}
E\left[w^{j}\right]-\frac{\gamma}{2} \operatorname{Var}\left[w^{j}\right] \tag{4}
\end{equation*}
$$

\]

where $w^{j}$ is wealth of an investor of type $j$ as of $t=2$. They purchase a portfolio of LT debt, issued by the firms, and finance this position with short-term borrowing. As a result, $w^{j}$ equals (in terms of the future value at $t=2$ ):

$$
w^{j}=d_{l}^{* j}-\iota^{j}\left(i_{1}, i_{2}\right) \int_{0}^{I} P_{l}(A) d_{l}^{j}(A) d A \cdot{ }^{16}
$$

Here, $d_{l}^{j *}$ is the realized payoff from holding a portfolio comprising LT debt of all firms, whereas $i^{j}\left(i_{1}, i_{2}\right)$ is the individual (compound) discount factor that each investor uses to judge her financing costs. $\iota^{j}\left(i_{1}, i_{2}\right)$ is heterogeneous across investor types $j \in\{R, Y\}$. A fraction $1-\alpha$ of the investors are "rational" and their compounded discount rate $l^{R}\left(i_{1}, i_{2}\right)$ is $\left(1+i_{1}\right)\left(1+i_{2}\right)$. On the other hand, a fraction $\alpha$ of the investors is of the "yield-seeking" type. Yield-seeking investors compare their expected returns to the current interest rate instead of the stream of expected interest rates. Their $\iota^{Y}\left(i_{1}, i_{2}\right)$ is $\left(1+i_{1}\right)^{2} .{ }^{17}$

Investors maximize their wealth by optimally choosing a LT debt portfolio including all firms' debt and, due to the mean-variance utility assumption, they have limited risk-bearing capacity. Finally, we assume an inelastic demand $g$, originating from preferred-habitat investors into LT debt, defined in terms of expected bond payments in $t=2 .{ }^{18}$

[^13]
### 4.2. EQUILIBRIUM

In order to characterize the equilibrium we start from the solution of the investors' problem. ${ }^{19}$ The downward sloping inverse demand curve for LT debt reads as:

$$
\begin{equation*}
P_{l}(A)=\frac{\delta F\left(\rho^{*}(A)\right)-\gamma F\left(\rho^{*}(A)\right) \delta(1-\delta)\left(\int_{0}^{I} F\left(\rho^{*}(A)\right) d_{l}(A) d A-g\right)}{\left(1+i_{1}\right)\left[\alpha\left(1+i_{1}\right)+(1-\alpha)\left(1+i_{2}\right)\right]} . \tag{5}
\end{equation*}
$$

The denominator in (5) is a weighted average of the discount factors of the two types of investors who compete in the same market to buy LT debt and face the same price. However, their demand differs due to their different attitude towards interest rates. In the numerator, the first right-hand-side term is the expected payoff from holding firm $A^{\prime}$ s LT bonds. Importantly, for such an expected payoff, yield-seeking investors are willing to pay a premium on LT debt if $i_{1}<i_{2}$. The second term suggests that investors, being risk-averse, are compensated (through a lower price) for holding risky LT debt.

Before describing firms' optimal plans, we first define the notion of unconstrained and constrained firms: ${ }^{20}$ a firm is "unconstrained" whenever all three constraints IC, LL, IR are slack and "constrained" if the opposite is true. Let $\bar{A}$ denote the lowest endowment compatible with the optimal continuation threshold. For unconstrained firms, i.e. for firms endowed with $A>\bar{A}$, continuation is at the optimal value, namely $\rho^{*}(A)=\delta R /\left(1+i_{2}\right)$. For these firms, LL does not bind, so they are indifferent to taking on any amount of short-term debt. Moreover, they collectively invest in LT debt until its price equals their valuation. ${ }^{21}$

[^14]In contrast, a constrained firm piles up as much short-term and LT debt as it requires to fund its project, namely for $A<\bar{A}$ :

$$
\begin{aligned}
d_{l}(A) & =R-\frac{B}{\Delta P} \\
d_{s}(A) & =r-\rho^{*}(A)
\end{aligned}
$$

This implies that the continuation cutoff of constrained firms is set below the optimal level, as shown in Appendix IB.3. Intuitively, firms below $\bar{A}$ need a large amount of debt to start their projects. High reliance on short-term financing tightens LL, and thus continuation $\rho^{*}$ is compromised.

### 4.3. EfFECTS OF A POLICY RATE CHANGE ON THE MATURITY STRUCTURE

We study how changes to the short-term policy rate $i_{1}$ affect firms' debt maturity structure. Recall that our empirical findings indicate that unconstrained firms increase LT debt in response to an interest rate cut, thereby lengthening their debt maturity structure, and contributing to an analogous adjustment at the aggregate level. To rationalize this, we need a sufficiently strong reaction of investor demand for long-term bonds to variations in the short-term rate. In our model, this is possible because of the presence of yield-seeking investors. ${ }^{22}$

We denote the strength of the yield-seeking motive relative to risk aversion as the ratio $\varkappa=\alpha / \gamma \in[0, \infty)$. We characterize the effect of a policy rate change as our main theoretical result in Proposition 1.

## Proposition 1. If the monetary authority decreases $i_{1}$ :

1. Unconstrained firms do not change $\rho^{*}(A)$ and, for high enough $\varkappa$, they increase their
[^15]LT debt issuance.
2. The aggregate change of LT debt of unconstrained firms generalizes to individual firm behavior. Furthermore, as $\varkappa$ approaches infinity, the change in $d_{l}(A)$ approaches infinity, while the change in $d_{s}(A)$ does not depend on $\varkappa$.
3. Constrained firms increase $\rho^{*}(A)$ and reduce short-term debt while their LT debt is unchanged.

## Proof. See Appendix IB.3.

Results 1 and 2 require that yield-seeking motives are sufficiently strong relative to risk aversion. Intuitively, this means that an interest rate decrease affects the discount factors of yield-seeking investors relatively more, and thus they demand more LT bonds. This implies a large elasticity of the LT bond demand with respect to changes in the policy rate. Under these conditions, the upward demand shift due to an interest rate loosening creates an excess demand for LT bonds. Constrained firms issue LT debt at their limit; hence, only unconstrained (large) firms can accommodate the demand shift and find it optimal to do so (Result 1). Result 2 states that, under $\varkappa$ sufficiently large, any variations of the LT debt share of large companies are larger than those of small companies, thereby matching the cross-sectional empirical evidence. Finally, Result 3 indicates that constrained firms reduce their short-term debt in response to lower rates to achieve a more efficient continuation probability for their projects. Their LT debt choice is unaffected, as it is pinned down by the IC constraint. A corollary from Proposition 1 is that all firms make adjustments that result in a lengthening of their debt maturity structure, therefore the model aligns with the observed aggregate lengthening of the corporate debt maturity in reaction to an expansionary monetary policy shock.

## 5. EMPIRICAL EVIDENCE ON THE MODEL'S MECHANISM

Our model is consistent with the empirical results in Section 3. The mechanism entails that: i) yield-seeking investors demand more LT debt after a policy rate cut; ii) large, financially unconstrained companies accommodate this demand shift by issuing LT bonds, hence benefiting from lower financing costs. In this section, we bring such mechanism to the data.

Before moving to the empirical tests in the next subsections, we remark that throughout our sample period monetary policy shocks are associated with interest rate adjustments consistent with the model. First, expansionary monetary policy shocks widen the term spread (see Appendix Figure IA1, also consistent with Adrian \& Shin 2010), in line with our modeling of monetary policy shocks as a onetime change in short-term rates. Nonetheless, in our model, the term premium falls because yield-seeking investors are more willing to hold long-term bonds, as they aim at maximizing current portfolio yields. Indeed, Figure IA2 shows that expansionary and exogenous monetary policy shocks are associated to lower term premia at long (5- and 10-year) horizons (consistent with empirical evidence in Gertler \& Karadi 2015, Hanson \& Stein 2015, Hanson et al. 2021).

### 5.1. ISSUANCE OF CORPORATE BONDS AND SYNDICATED LOANS

First, we check whether the likelihood of issuing new LT bonds and/or loans is differently affected by interest rate changes across small and large companies. To this end, we use a model otherwise identical to that in Equation 2, but with the dependent variable given by either $\mathbb{1}(\text { BondIssue })_{f, t+h}$ or $\mathbb{1}(\text { LoanIssue })_{f, t+h}$. These are dummy variables with value 1 if firm $f$ issues LT bonds or loans, respectively, in year-quarter $t+h$ and with value 0 if it does not. Hence, at horizon $h$, the coefficient $\beta_{2, h}$ measures the relative difference in large companies' probability of issuing bonds
(bank loans) as of year-quarter $t+h$, induced by a policy rate shock at $t .{ }^{23}$
An adjustment in debt issuance may be driven by firms or financiers (either bond investors and/or banks). In a first effort to isolate their relative contribution, we investigate the reaction of financing costs at issuance. The idea behind this test is that an increase in debt issuance accompanied by lower (higher) financing rates would suggests that financiers (firms) contribute more decisively to the equilibrium adjustments.

One issue with this analysis is that only a tiny subset of companies ever issues bonds in two consecutive quarters. In operative terms, this means that if we were to apply a first-difference model, we would be left with very few observations and, ultimately, a meaningless cross-sectional comparison in a very small set of companies (highly skewed towards large firms). Therefore, we rather resort to the following model in levels:

$$
\begin{align*}
y_{f, t+h}= & \beta_{1, h} \operatorname{Large}_{f, t-1}+\beta_{2, h} \operatorname{Large}_{f, t-1} * \Xi_{t}+  \tag{6}\\
& +\Phi_{h} X_{f, t-1}+\mu_{f}+\mu_{s, t}+q_{f, t}+\xi_{f, t+h}
\end{align*}
$$

As dependent variable, we use standard proxies of bond and loan spreads, such as respectively BondSpread $_{f, t+h}$ - i.e. the spread between the bond yield at issuance and the Treasury yield at comparable maturity - and the all-in-drawn spread on syndicated loans AIDSpread $_{f, t+h}$. The main regressor of interest is a proxy of the current level of the policy rate, $\Xi_{t}$, eventually interacted with the dummy for large firms, $\operatorname{Large}_{f, t-1}$. We build $\Xi_{t}$ as the cumulative sum of the high-frequency monetary surprises: $\Xi_{t}=\sum_{t} \varepsilon_{t}$ (in the spirit of, among others, Romer \& Romer 2004, Coibion 2012, Ramey 2016). The vector of controls $X_{f, t-1}$ includes, as usual, the full interaction of $\operatorname{Large}_{f, t-1}$ with the macro controls ${ }^{24}$ and of the remaining firm-level

[^16]controls (sales growth, liquid assets and leverage) with $\Xi_{t}$. As usual, we augment the model with firm and industry*year-quarter fixed effects ( $\mu_{f}$ and $\mu_{s, t+h}$ ) and control for differences in fiscal calendar $q_{f, t}$. The error term, $\xi_{f, t+h}$, is double-clustered at the firm and industry*year-quarter level.

## Figure 7: Monetary Policy and Bond Issuance - Relative Response of Large Companies



This figure depicts the relative response in the likelihood of issuing bonds and in the related yield spread to an expansionary monetary policy surprise for large firms (as compared to smaller firms). Panel A shows the coefficients $\beta_{2, h}$ from the estimation of Equation $2, h=0,1, \ldots, 20$, with $\mathbb{1}\left(\right.$ BondIssue $_{f, t+h}$ as dependent variable, i.e. a dummy variable with value 1 if firm $f$ issues LT bonds in year-quarter $t+h$ and with value 0 if it does not. We calibrate a 1 s.d. expansionary Gürkaynak et al. (2005) monetary policy surprise $\left(\varepsilon_{t}=-10 b . p\right.$.). Panel B shows the coefficients $\beta_{2, h}$ from the estimation of Equation $6, h=0,1, \ldots, 20$, with $\mathbb{1}$ (BondSpread) $f_{f, t+h}$ as dependent variable. We calibrate to a 1 s.d. variation in the cumulative Gürkaynak et al. (2005) monetary policy shocks, $\Xi_{t}=-90$ b.p.. In both panels, the solid line represents the point estimates, and the dashed lines the $90 \%$ confidence interval. Standard errors doubleclustered at the firm and industry*year-quarter level.

### 5.1.1. Results

We display results for bond issuance in Figure 7. In Panel A, a 1 s.d. expansionary policy rate surprise boosts large firms' likelihood of issuing LT bonds by roughly 38 b.p. (as compared to smaller corporations). This increase corresponds to an additional $6 \%$ jump relative to the average likelihood of issuing LT bonds as of time $t$. The effect extends (and accumulates) over time and reverts to zero in roughly 5 quarters.

To understand the absolute impact of monetary policy shocks on the likelihood of issuing LT bonds, we report in Appendix Table A4 regressions for $h=0,1$ which do
not include time-varying fixed effects (columns 1 and 4, respectively). The results suggest that smaller companies also increase the likelihood of issuing bonds when the policy rate falls, however by a factor that is about half that observed for larger firms.

Panel B of Figure 7 reports the results from the estimation of Equation 6 for bond spread. An expansionary monetary policy shock triggers a bigger spread reduction for large companies, relative to smaller ones. In concrete terms, a 1 s.d. (i.e., 90 b.p.) descent in $\Xi_{t}$ triggers a relative fall in bond spread by about 20 b.p. (i.e., a $10 \%$ reduction compared to the average bond spread). Appendix Table A5 describes the background regressions for $h=0,1$, also shown with different sets of fixed effects that allow to evaluate the absolute response of the coupon rate. Appendix Table A5 also shows that there is a significant loss of observations due to the application of (3-digit SIC)-industry*year-quarter fixed effects. Bond issuance at the firm level is indeed quite lumpy across time. Therefore, narrowing the comparison within granular industries implies the loss of many (within-industry) singletons. Hence, the within-industry cross-sectional comparison comprehends few firms. For robustness purposes, we apply looser industry definitions and check that our findings go through. In Appendix Table A6, in columns 1-4, we replicate our analysis, but this time comparing companies along the whole cross-section. In this case, the Large $_{f, t-1}$ dummy captures those companies in the top quartile of the entire sample of NFCs in the US stock market. The result that large companies' bond spread at issuance descents along with $\Xi_{t}$ still goes through. If anything, the magnitude of the relative adjustment is even larger. A similar pattern emerges when applying increasingly more granular industry definitions (sectoral level in columns 5-8 and 2-digit SIC industry in columns 9-12). Finally, results are also robust to excluding the post-ZLB period (see Appendix Figure A4).

Adjustments in the bond market are consistent with the firm-level patterns. It
remains to be understood whether, beyond such bond channel, also bank credit supply reacts consistently. This might not be the case, given a prominent literature (dating back to, e.g. Gertler \& Gilchrist 1994) suggesting a greater sensitivity to monetary policy shocks for bank-dependent small and medium (as opposed to large) enterprises. Our results confirm this intuition. Panel A of Figure 8 shows that, in reaction to an expansionary monetary policy shock, large firms' likelihood to issue a syndicated loan drops, as compared to small firms'. Moreover, in Panel B of Figure 8, large firms' all-in-drawn spread increases significantly relative to small firms'. Hence, evidence from syndicated loans suggests a relative credit supply tightening by banks for larger corporations, which is inconsistent with our cross-sectional evidence.

Figure 8: Monetary Policy and (Syndicated) Loan Issuance - Relative Response of Large Companies


This figure depicts the relative response in the likelihood of issuing loans and in the related spread to an expansionary monetary policy surprise for large firms (as compared to smaller firms). Formally, Panel A shows the coefficients $\beta_{2, h}$ from the estimation of Equation $2, h=0,1, \ldots, 20$, with $\mathbb{1}(\operatorname{LoanIssue})_{f, t+h}$ as dependent variable, i.e. a dummy variable with value 1 if firm $f$ issues LT bonds in year-quarter $t+h$ and with value 0 if it does not. The IRF is calibrated to a 1 s.d. expansionary Gürkaynak et al. (2005) monetary policy surprise $\left(\varepsilon_{t}=-10 b\right.$.p.). Panel B shows the coefficients $\beta_{2, h}$ from the estimation of Equation $6, h=0,1, \ldots, 20$, with $\mathbb{1}(\text { AIDSpread })_{f, t+h}$ as dependent variable. The IRF is calibrated to a $1 \mathrm{~s} . \mathrm{d}$. variation in the cumulative Gürkaynak et al. (2005) monetary policy surprises, $\Xi_{t}=-90 \mathrm{~b} . \mathrm{p}$. . In both panels, the solid line represents the point estimates, and the dashed lines the $90 \%$ confidence interval. Standard errors double-clustered at the firm and industry*year-quarter level.

Put differently, reactions in the bond markets to a monetary policy shock are the primary avenue for understanding the ultimate implications for corporate debt maturity. To test this hypothesis further, we check whether the effects of monetary
policy on corporate debt maturity reverberate more powerfully on bond issuers. In particular, we run the following panel local projection regressions, for $h=0,1, \ldots, 20$ :

$$
\begin{align*}
\Delta_{h} y_{f, t+h}= & \beta_{1, h} \text { Large }_{f, t-1}+\beta_{2, h} \text { Bond }_{f, t}+\beta_{3, h} \text { Large }_{f, t-1} * \Delta M P_{t}+ \\
& +\beta_{4, h} \text { Bond }_{f, t-1} * \Delta M P_{t}+\beta_{5, h} \text { Bond }_{f, t-1} * \text { Large }_{f, t-1}+  \tag{7}\\
& +\beta_{6, h} \text { Bond }_{f, t-1} * \text { Large }_{f, t-1} * \Delta M P_{t}+\Gamma_{h} X_{f, t-1}+\mu_{f}+\mu_{s, t}+u_{f, t+h} .
\end{align*}
$$

The key coefficient of interest is $\beta_{6, h}$, describing the relative response to an expansionary monetary policy shock by large bond issuers (as compared to large firms that do not issue bonds). A positive dynamic response would confirm that bond financing is a key channel amplifying the response of debt maturity to monetary policy by large firms (highlighted in section 3.2). We show the results in Figure 9, where, as usual by now, we employ the high-frequency monetary policy surprise $\varepsilon_{t}$. More in detail, the lower panel displays the IRF pinned down by the coefficients $\beta_{6, h} h=0,1,2, \ldots, 20$. The IRF suggests that, at relevant horizons, having access to the bond market reinforces the pass-through of monetary policy shocks for large firms. Actually, it nearly doubles such pass-through, as clear from the comparison with the baseline IRF for large firms that do not issue bonds in the upper-left panel (reporting the coefficients $\beta_{3, h}$ ). Finally, the upper-right panel suggests that for small bond issuers, there is not necessarily a significant effect of monetary policy shocks on corporate debt maturity.

### 5.2. BOND REFINANCING

Why do large firms issue more bonds when monetary policy is expansionary? To answer this question, we first look at the effect of monetary policy surprises on investment. The results (not shown for brevity though available upon request) show an insignificant relative reaction of large firms' investment to monetary policy sur-

Figure 9: Monetary Policy and Corporate Debt Maturity - Relative Response of Large \& Bond-Issuing Companies


This figure depicts the relative response of large firms and bond issuers to a monetary policy surprise (as compared to smaller, non-bond issuers firms). Formally, Panel A and Panel B show the coefficients $\beta_{3, h}$ and $\beta_{4, h}$ from the estimation of Equation 7, i.e. the relative response of large firms and bond issuers to a 1 s.d. expansionary monetary policy surprise $\left(\varepsilon_{t}=-10 b . p\right.$.) $h=0,1, \ldots, 20$. Panel $C$ shows coefficients $\beta_{6, h}$ from the estimation of Equation 7 , i.e. the relative response of large bond-issuing companies to an identical s.d. surprise. In each panel, the solid line represents the point estimates, and the dashed lines the $90 \%$ confidence intervals. Standard errors double-clustered at the firm and industry*year-quarter level.
prises. Indeed, in our model, large firms are unconstrained, and new bond issuance in the aftermath of an expansionary monetary policy shock does not alter their ability to invest. On the contrary, large firms exploit the discount on yield at issuance associated with investors' reach for yield. Hence, a natural hypothesis following our model predictions is that large firms time the market to refinance their longterm (bond) debt at relatively cheap rates. ${ }^{25}$ To test this conjecture, we run versions of Model 2 with bond-refinancing dummies as dependent variables. In practical terms, following Xu (2018), we classify a bond as having been refinanced if the same firm issues another bond within a 3 -month window around the bond's retirement and the new bond issuance is of a comparable amount relative to the retired amount (i.e. if it falls in a $70-130 \%$ of the retired amount interval). Moreover, we label refi-

[^17]nancing as "early" if it occurs more than 6 months before the bond is scheduled to mature.

Figure 10 displays the effect of expansionary monetary policy surprises on refinancing activity. In Panel A, large firms are significantly more likely than small firms to refinance a bond. In Panel B, in which we focus on the likelihood of earlyrefinancing bonds, the same result applies. Quantitatively speaking, a 1 s.d. expansionary monetary policy shock raises the likelihood of refinancing bonds for large firms in the following quarter by 6 p.p. (as compared to small firms). This is an economically large effect, corresponding to $25 \%$ of the unconditional probability of refinancing a bond.

Figure 10: Monetary policy and Corporate Bond Refinancing
(A) Refinancing x Large
(B) Early Refinancing x Large



This figure shows the relative effect of a $1 \mathrm{~s} . \mathrm{d}$. expansionary monetary policy surprise $\left(\varepsilon_{t}=-10 b . p\right.$.) on the likelihood of firms in the largest $25 \%$ of their three-digit SIC-industries to refinance bonds (Panel A) or early-refinance bonds (Panel B). Formally, it shows the coefficients $\beta_{2, h}$ from the estimation of a local projection model akin to Equation 2, in which we proxy monetary policy changes $\Delta M P_{t}$ with the Gürkaynak et al. (2005) monetary policy surprises, $\varepsilon_{t}$. The dependent refinancing dummies are defined as in Xu (2018), and are summarized in Appendix Table A1 as Refinancing $f_{f, t}$ and EarlyRefinancing $f, t$. As bonds are flagged as refinanced by matching them with previously and subsequently issued bonds, the effective sample runs from 1996 to 2013 (out of the full sample from 1990 to 2017).

### 5.3. Corporate Bond Purchases by Insurance Firms

We exploit data on insurance firms' bond holdings and purchases to test whether their reaction to monetary policy surprises aligns with our model predictions. We focus on insurers due to the fact that they represent a key class of investors in corpo-
rate bonds, holding around $40 \%$ of total US corporate bonds. We begin by running regressions exploiting time variation only, therefore pinning the average response of insurers' purchases of corporate bonds to monetary policy surprises:

$$
\begin{equation*}
y_{i, t+h}=\beta_{1, h} \Delta M P_{t}+\Gamma_{h} X_{i, t-1}+\mu_{i}+u_{i, t, h} \tag{8}
\end{equation*}
$$

We focus on three different dependent variables $y_{i, t+h}{ }^{26}$ First, we analyze the par value of new corporate bond purchases (net of the par value of sold bonds and rescaled by insurer asset size), $P V N e t_{i, t+h}$. Second, we use the logged maturity of new corporate bond purchases (net of the maturity of sold bonds), LogMat ${ }_{i, t+h}$. Third, we look at the share of bonds purchased on the primary market PrimShare ${ }_{i, t+h}$. $X_{t-1}$ contains the usual macro controls, whereas $\mu_{i}$ denotes insurer fixed effects. Figure 11 summarizes the local projections for the model described my Equation 8.

In Panel A, following a 1 s.d. expansionary monetary policy surprise, the par value of net purchases increases on impact by $16 \mathrm{~b} . \mathrm{p}$. (approximately $50 \%$ of the quarterly change in new bond purchases), and so does the maturity of the net purchases (Panel B). Moreover, in line with our theoretical framework in which lower monetary policy rates and higher demand for bonds stimulate bond issuance, the share of bonds purchased in the primary market increases (Panel C).

In addition, we test whether reach-for-yield motives are a driver of the just described adjustments through the following local projection model:

$$
\begin{align*}
y_{i, t+h}= & \beta_{1, h} H Y_{i, y-1}+\beta_{2, h} H Y_{i, y-1} * \Delta M P_{t}+  \tag{9}\\
& +\Gamma_{h} X_{i, t-1}+\mu_{i}+\mu_{t}+u_{i, t+h}
\end{align*}
$$

Our key coefficient of interest is $\beta_{2, h}$, loading the interaction between the policy rate surprise, $\Delta M P_{t}$, and a dummy variable, $H Y_{i, y-1}$, denoting yield-seeking investors (as of year $t-1$ ). In practice, we label as yield-seeking those insurers with

[^18]Figure 11: Insurer Net Purchases after an Expansionary Shock
(a) Par Value
(B) Maturity
(C) Primary Market Share




This figure shows the relative effect of a 1 s.d. expansionary monetary policy surprise ( $\varepsilon_{t}=-10 b . p$.) on US insurers' purchases of corporate bonds, so the coefficients $\beta_{1, h}$ in the regression models described in Equation 8 , in which we proxy monetary policy changes $\Delta M P_{t}$ with the Gürkaynak et al. (2005) monetary policy surprise, $\varepsilon_{t}$. The dependent variable in Panel A is the par value of purchased and acquired corporate bonds net of the par value of sold and disposed-of corporate bonds $P V N e t_{i, t+h}$. The dependent variable in Panel B is the log maturity of purchased and acquired corporate bonds net of the maturity of sold and disposed-of corporate bonds HoldingLogmat $i_{i, t+h}$. Panel C reports the effect on the share of bonds purchased from the primary market PrimShare ${ }_{i, t}$. Throughout we include insurer fixed effects, as well as lagged macro variables (annual GDP growth and inflation rate; a recession dummy; term spread, corporate spread, share of Treasuries with maturity above 20 years) as control variables and cluster standard errors at the insurer level. Displayed are $90 \%$ confidence intervals.
above-median corporate bond portfolio yields in year $y-1 . X_{i, t-1}$ is a vector of lagged insurer level controls - including log asset size and the capital (solvency) ratio - fully interacted with the policy rate shock. As we exploit the cross-sectional response to monetary policy surprises, we introduce time fixed effects $\mu_{t}$ on top of insurer fixed effect, $\mu_{i}$. We cluster the error term, $u_{i, t+h}$, at the insurer level.

Figure 12 reports the estimated IRFs. In Panel A, yield-seeking investors exhibit disproportionately large net purchases. Moreover, in Panel C, they also disproportionately tap the primary market, hence facilitating bond issuance relatively more than other insurers. Both results are in line with our theoretical predictions. Panel B suggests that reach for yield is not associated with a relative lengthening of the maturity of purchased corporate bonds. Indeed, higher demand for long-term bonds following a policy rate cut may be driven by other factors as well, e.g. asset and liability matching (Ozdagli \& Wang 2019). However, crucially, our evidence shows that the bulk of the adjustment (in terms of volume of purchases and presence in the primary market) in insurers' portfolios is carried out by yield-seeking investors,
thereby validating our model predictions.
Figure 12: Insurer Net Purchases after an Expansionary Shock
(a) Par Value
(B) MATURITY
(C) Primary Market SHARE




This figure shows the interaction effect of a $1 \mathrm{~s} . \mathrm{d}$. expansionary monetary policy surprise and the indicator $H Y_{i, y-1}$, which denotes whether insurer $i$ was in the top $50 \%$ of deviations of the par value-weighted yield-to-maturity of their corporate bond portfolio from the average yield of all corporate bonds in year $y-1$, on US insurers' purchases of corporate bonds. The coefficient of interest is $\beta_{2, h}$ in the regression models described in Equation 9, in which we proxy monetary policy changes $\Delta M P_{t}$ with the Gürkaynak et al. (2005) monetary policy surprises, $\varepsilon_{t}$. The dependent variable in Panel A is the par value of purchased and acquired corporate bonds net of the par value of sold and disposed-of corporate bonds. The dependent variable in Panel B is the log maturity of purchased and acquired corporate bonds net of the maturity of sold and disposed-of corporate bonds. Panel C reports the effect on the share of corporate bonds purchased from the primary market. Throughout we include insurer fixed effects, as well as insurer asset size and capital ratios interacted with $\Delta M P_{t}$ as control variables and cluster standard errors at the insurer level. Missing values for insurer capital ratios are imputed by using the quarterly mean value for the capital ratio for the missing insurer-year pair. Displayed are $90 \%$ confidence intervals. These results are robust to defining $H Y_{i, y-1}$ as denoting that an insurer is above the median in terms of value-weighted yield deviation from the average yield within the respective NAIC regulatory categories in $y-1$. The IRF using this definition of $H Y_{i, y-1}$ can be found in the Internet Appendix Figure IA3.

For robustness, we also investigate the reaction to monetary policy surprises of another important class of investors, namely corporate bond mutual funds (CBMF). By doing so, we cover approximately an additional $30 \%$ of total corporate bond holdings in the US. Data on CBMF holdings are aggregated and hence we analyze the response of corporate bond holdings and the average maturity of the debt securities held in the portfolio. ${ }^{27}$ The results are displayed in Figure IA4 of the Internet Appendix, where we also report a detailed description of the model, which mirrors very closely Equation 9. We label high-yield CBMF, identified as in Choi \& Kronlund (2018), as yield-seeking investors. ${ }^{28}$ Panel A describes the relative response of HY funds' corporate bond holdings to a 1 s.d. reduction in $\varepsilon_{t}$. The effect is positive and close to $1 \%$ one quarter after the shock. Next, Panel B confirms that, following

[^19]an interest rate descent, on top of buying corporate bonds, HY funds tilt their portfolio towards debt securities with longer maturity, the effect being more persistent than for the volume of corporate bonds, and peaking around $2 \%$ two years after the monetary policy shock.

## 6. CONCLUSION

We show that monetary policy significantly influences the cyclical dynamics of the maturity structure of corporate debt. At the aggregate level, a 1 s.d. expansionary monetary policy surprise increases the share of long-term debt by $87 \mathrm{~b} . \mathrm{p}$. over a 12-quarter horizon, accounting for about $20 \%$ of the corresponding variation in the share of LT debt. Moreover, examining the cross-section of firms, we show that very large and bond-issuing corporations drive the adjustment.

We rationalize these empirical facts through a simple model combining financial frictions due to firms' moral hazard and yield-oriented investors, who increase the demand for long-term debt when the interest rate goes down. Only large, unconstrained companies can accommodate such an upward shift in demand. We bring the mechanism to the data through several empirical tests on the relation between monetary policy, bond issuance, and insurers' holdings of corporate bonds and find supportive evidence.

Our work shows that monetary policy is a primary and previously overlooked driver of cyclical fluctuations in corporate debt maturity, the implications of which are yet to be fully understood. In particular, corporate debt maturity has been shown to affect firm-level and macroeconomic responses to shocks of different natures, including e.g. financial crises and inflation (Almeida et al. 2009, Gomes et al. 2016, Kalemli-Özcan et al. 2022). Quantifying the extent to which monetary policy — by affecting corporate debt maturity — affects those outcomes is an important task, which we leave for future research.

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## Appendix

TABLES
Table A1: Summary Statistics

| VARIABLES | $\begin{gathered} (1) \\ \text { Scale } \end{gathered}$ | (2) | (3) | $\begin{aligned} & (4) \\ & \mathfrak{p} 25 \end{aligned}$ | $\begin{array}{r} (5) \\ \text { p50 } \\ \hline \end{array}$ | $\begin{aligned} & (6) \\ & \text { p75 } \end{aligned}$ | sd |
| :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: |
| Macro-level Variables |  |  |  |  |  |  |  |
| $\Delta L T-$ Debt $_{t}$ | \% | 112 | 0.154 | -0.360 | 0.0804 | 0.553 | 0.642 |
| $\Delta L T-$ Debt $_{t+1}$ | \% | 111 | 0.315 | -0.467 | 0.0687 | 0.992 | 1.141 |
| $\Delta L T-$ Debt $_{t+2}$ | \% | 110 | 0.480 | -0.590 | 0.232 | 1.558 | 1.602 |
| $\Delta L T-D e b t_{t+3}$ | \% | 109 | 0.651 | -0.794 | 0.123 | 1.834 | 2.026 |
| $\Delta L T-$ Debt $_{t+4}$ | \% | 108 | 0.826 | -0.720 | 0.218 | 2.554 | 2.401 |
| $\triangle E F F R_{t}$ | \% | 112 | -0.0638 | -0.105 | -0.01000 | 0.0900 | 0.445 |
| $\varepsilon_{t}$ | \% | 112 | -0.0340 | -0.0325 | -0.00750 | 0.00500 | 0.0966 |
| $\triangle G D P_{t-1}$ | \% | 112 | 2.458 | 1.650 | 2.600 | 3.650 | 1.712 |
| $\triangle C P I_{t-1}$ | \% | 112 | 2.496 | 1.714 | 2.584 | 3.202 | 1.284 |
| $\mathrm{Rec}_{t-1}$ | 0/1 dummy | 112 | 0.0982 | 0 | 0 | 0 | 0.299 |
| $\Delta L T-$ Treas $_{t-1}$ | \% | 112 | -0.0197 | -0.231 | -0.00926 | 0.254 | 0.357 |
| $\Delta i_{t-1}^{10 y-3 m}$ | \% | 112 | 0.00866 | -0.300 | -0.0550 | 0.275 | 0.485 |
| $\Delta i_{t-1}^{\text {baa-aaa }}$ | \% | 112 | -0.00205 | -0.0900 | -0.01000 | 0.0600 | 0.252 |
| Firm-level Variables |  |  |  |  |  |  |  |
| $\Delta L T-\operatorname{Debt}_{f, t}$ | \% | 327,532 | -0.482 | -2.051 | 0 | 0.800 | 17.03 |
| $\Delta L T-\operatorname{Debt}_{f, t+1}$ | \% | 299,722 | -0.915 | -4.038 | 0 | 1.911 | 21.77 |
| $\Delta L T-\operatorname{Debt}_{f, t+2}$ | \% | 284,670 | -1.238 | -5.662 | 0 | 2.987 | 24.65 |
| $\Delta L T-$ Debt $_{f, t+3}$ | \% | 270,304 | -1.466 | -6.823 | -0.00232 | 3.757 | 26.47 |
| $\Delta L T-$ Debt $_{f, t+4}$ | \% | 259,167 | -1.568 | -7.964 | 0 | 4.795 | 28.08 |
| $\Delta$ Sales $_{f, t-1}$ | \% | 327,532 | 0.75 | -8.14 | 1.17 | 10.4 | 23 |
| Liquid Assets $_{f, t-1}$ | \% | 327,532 | 12.9 | 1.39 | 5.21 | 16.5 | 18.1 |
| Leverage $_{f, t-1}$ | \% | 327,532 | 34.8 | 12.00 | 27.4 | 42.9 | 45.7 |
| Size $_{f, t-1}$ | Log(Mln US\$) | 327,532 | 4.965 | 3.183 | 4.938 | 6.764 | 2.477 |
| $\mathbb{1}$ (BondIssue $_{)_{f, t}}$ | $0 / 1$ dummy | 118,993 | 0.0677 | 0 | 0 | 0 | 0.251 |
| $\mathbb{1}$ (BondIssue) $)_{f, t+1}$ | $0 / 1$ dummy | 110,896 | 0.0745 | 0 | 0 | 0 | 0.263 |
| $\mathbb{1}$ (BondIssue) $)_{f, t+2}$ | 0/1 dummy | 106,453 | 0.0781 | 0 | 0 | 0 | 0.268 |
| $\mathbb{1}$ (BondIssue) $)_{f, t+3}$ | $0 / 1$ dummy | 102,372 | 0.0813 | 0 | 0 | 0 | 0.273 |
| $\mathbb{1}$ (BondIssue) ${ }_{f, t+4}$ | 0/1 dummy | 99,435 | 0.0829 | 0 | 0 | 0 | 0.276 |
| BondSpread $_{f, t}$ | Basis Points | 7,620 | 222.1 | 82.27 | 145 | 324 | 204.0 |
| $B^{\text {BondSpread }}$ f,t+1 | Basis Points | 5,970 | 208.6 | 80 | 136.6 | 293 | 194.7 |
| $B^{\text {BondSpread }}$ f,t+2 | Basis Points | 6,052 | 209.8 | 80 | 137.5 | 295 | 194.9 |
| $B^{\text {BondSpread }}$ f,t+3 | Basis Points | 5,989 | 208.1 | 80 | 137 | 290 | 193.3 |
| BondSpread $_{f, t+4}$ | Basis Points | 5,989 | 206.3 | 80 | 135.8 | 287.5 | 191.8 |
| $\mathbb{1}(\text { Loan Issue) })_{f, t}$ | $0 / 1$ dummy | 235,213 | 0.0929 | 0 | 0 | 0 | 0.29 |
| $\mathbb{1}$ (Loan Issue $)_{f, t+1}$ | $0 / 1$ dummy | 219,401 | 0.0918 | 0 | 0 | 0 | 0.29 |
| $\mathbb{1}$ (Loan Issue $)_{f, t+2}$ | $0 / 1$ dummy | 211,154 | 0.0907 | 0 | 0 | 0 | 0.29 |
| $\mathbb{1}(\text { Loan Issue })_{f, t+3}$ | $0 / 1$ dummy | 202,687 | 0.0909 | 0 | 0 | 0 | 0.29 |
| $\mathbb{1}$ (Loan Issue) ${ }_{f, t+4}$ | 0/1 dummy | 196,487 | 0.0901 | 0 | 0 | 0 | 0.29 |
| AIDSpread $_{f, t}$ | Basis Points | 20,220 | 174.3 | 50.40 | 150 | 255 | 150.9 |
| AIDSpread $_{f, t+1}$ | Basis Points | 18,616 | 173.6 | 50. | 150 | 255 | 150.9 |
| AIDSpread $_{f, t+2}$ | Basis Points | 17,678 | 172.6 | 50 | 150 | 255 | 150.6 |
| AIDSpread $_{f, t+3}$ | Basis Points | 16,999 | 171.6 | 50 | 150 | 255 | 150.1 |
| AIDSpread $_{f, t+4}$ | Basis Points | 16,335 | 170.9 | 50 | 148 | 254 | 150.3 |
| Bank Loan Amount ${ }_{f, t}$ | Mln US\$ | 235,213 | 35.1 | 0 | 0 | 0 | 306.8 |
| Bank Maturity ${ }_{f, t}$ | Months | 21,846 | 41.68 | 17.49 | 37 | 60 | 29.25 |
| StockReturn $_{f, t}$ | \% | 653,090 | 0.201 | -1.53 | 0 | 1.76 | 4.30 |
| Bond $_{f, t}$ | 0/1 dummy | 662,675 | 0.173 | 0 | 0 | 0 | 0.379 |
| Refinancing ${ }_{f, t}$ | $0 / 1$ dummy | 14,919 | 0.24 | 0 | 0 | 0 | 0.43 |
| EarlyRefinancing ${ }_{f, t}$ | 0/1 dummy | 14,919 | 0.23 | 0 | 0 | 0 | 0.42 |


| Mutual Fund-level variables |  |  |  |  |  |  |  |
| :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: |
| $\Delta M_{a t u}{ }_{m, t}$ | \% | 72,389 | -0.412 | -3.190 | -0.199 | 2.327 | 15.55 |
| $\Delta$ Matu $_{m, t+1}$ | \% | 72,072 | -0.799 | -5.241 | -0.731 | 3.673 | 19.55 |
| $\triangle M_{\text {atu }}{ }_{\text {m,t+2 }}$ | \% | 71,617 | -1.226 | -6.860 | -1.046 | 4.437 | 22.51 |
| $\Delta M_{\text {atu }}{ }_{m, t+3}$ | \% | 70,940 | -1.614 | -8.246 | -1.409 | 5.167 | 24.40 |
| $\Delta M_{\text {atu }}{ }_{m, t+4}$ | \% | 69,924 | -1.926 | -9.381 | -1.816 | 5.555 | 26.17 |
| $\Delta C B_{m, t}$ | \% | 71,063 | -0.515 | -4.449 | -0.459 | 3.214 | 18.26 |
| $\Delta C B_{m, t+1}$ | \% | 68,201 | -1.042 | -6.847 | -0.974 | 4.485 | 23.23 |
| $\triangle C B_{m, t+2}$ | \% | 65,217 | -1.644 | -8.635 | -1.389 | 5.219 | 26.46 |
| $\triangle C B_{m, t+3}$ | \% | 62,112 | -2.111 | -9.789 | -1.928 | 5.607 | 29.04 |
| $\triangle C B_{m, t+4}$ | \% | 58,889 | -2.810 | -10.99 | -2.422 | 5.823 | 30.79 |
| $\mathrm{HY}_{\text {m }}$ | 0/1 dummy | 72,389 | 0.388 | 0 | 0 | 1 | 0.487 |
| TurnoverRatio ${ }_{m, t-1}$ | \% | 72,389 | 0.763 | 0.440 | 0.720 | 1.490 | 8.429 |
| ExpenseRatio ${ }_{m, t-1}$ | \% | 72,389 | 0.00979 | 0.00570 | 0.00810 | 0.0119 | 0.00507 |
| $N A V_{m, t-1}$ | Log(Mln US\$) | 72,389 | 1.822 | 1.687 | 1.806 | 1.920 | 0.481 |
| Returns $_{\text {m,t-1 }}$ | \% | 72,387 | 0.00960 | -0.000325 | 0.00875 | 0.0215 | 0.0234 |
| Insurer Security Holdings |  |  |  |  |  |  |  |
| $R Y_{b, t}$ | \% | 5,106,789 | 31 | 2.17 | 3.39 | 4.78 | 52588 |
| $I Y_{b, t}$ | \% | 4,973,067 | 21.55 | 1.687 | 4.32 | 5.39 | 1885 |
| Rating $_{6, t}$ | 1-28 | 6,610,114 | 20.72 | 19 | 21 | 23 | 4.93 |
| NAICCat ${ }_{\text {b,t }}$ | 1-9 | 6,610,114 | 1.73 | 1 | 2 | 2 | 0.87 |
| Insurance Firm Variables |  |  |  |  |  |  |  |
| ParValuePur Ac $_{i, t}$ | \% | 89,529 | 0.31 | 0.0003 | 0.002 | 0.006 | 4.26 |
| ParValueSoldDisp $p_{i, t}$ | \% | 89,529 | 0.3 | 0.0003 | 0.001 | 0.46 | 7.76 |
| $\mathrm{PVNet}_{i, t}$ | \% | 89,529 | 0.019 | -0.001 | 0 | 0.002 | 5.7 |
| ${ }_{P V N e t}^{i, t+1}$ | \% | 76,383 | -0.005 | -0.001 | 0 | 0.003 | 5.8 |
| ${ }^{\text {PN }}$ Net ${ }_{\text {i, }, t+2}$ | \% | 73,880 | 0.011 | -0.001 | 0 | 0.002 | 5.9 |
| PVNet ${ }_{\text {, }}$, +3 | \% | 71,627 | 0.016 | -0.001 | 0 | 0.002 | 6.3 |
| PVNet ${ }_{i, t+4}$ | \% | 69,747 | 0.006 | -0.001 | 0 | 0.003 | 6.21 |
| HoldingLogmaturityPur Ac $i_{i, t}$ | Log(Days) | 74,892 | 8.22 | 7.35 | 8.29 | 9.26 | 1.3 |
| HoldingLogmaturitySoldDisp $p_{i, t}$ | Log(Days) | 79,749 | 8.39 | 7.58 | 8.55 | 9.32 | 1.28 |
| Logmat $_{i, t}$ | Log(Days) | 61,782 | -0.12 | -0.7 | -0.09 | 0.43 | 1.01 |
| Logmat $_{\text {, }, t+1}$ | Log(Days) | 58,018 | -0.11 | -0.7 | -0.09 | 0.44 | 1.12 |
| Logmat $_{i, t+2}$ | Log(Days) | 56,363 | -0.1 | -0.7 | -0.08 | 0.44 | 1 |
| Logmat $_{\text {, }, t+3}$ | Log(Days) | 54,857 | -0.1 | -0.7 | -0.08 | 0.44 | 1 |
| Logmat ${ }_{\text {, }, t+4}$ | Log(Days) | 53,410 | -0.1 | -0.7 | -0.08 | 0.44 | 1 |
| PrimShare $_{i, t}$ | \% | 74,893 | 0.15 | 0 | 0.03 | 0.14 | 0.28 |
| PrimShare ${ }_{i, t+1}$ | \% | 66,720 | 0.14 | 0 | 0.03 | 0.13 | 0.26 |
| PrimShare $i_{, t+2}$ | \% | 64,266 | 0.14 | 0 | 0.03 | 0.14 | 0.26 |
| PrimShare $i_{, t+3}$ | \% | 62,156 | 0.14 | 0 | 0.03 | 0.13 | 0.26 |
| PrimShare ${ }_{i, t+4}$ | \% | 60,265 | 0.14 | 0 | 0.03 | 0.13 | 0.26 |
| Insurer AssetSize i,t-1 $^{\text {den }}$ | Log(US\$) | 89,529 | 19.27 | 17.67 | 19.11 | 20.67 | 2.24 |
| InsurerCapitalRatio ${ }_{\text {i,t-1 }}$ | Ratio | 115,474 | 192 | 6.24 | 10.27 | 24.97 | 5932 |
|  | \% | 93,077 | 2.19 | 1.038 | 1.9 | 2.83 | 10.35 |
| YieldSeeking ${ }_{\text {i,y }}$ | \% | 93,077 | -0.25 | -0.87 | -0.43 | -0.02 | 10.28 |
| $H Y_{i, y}$ | \% | 93,023 | 0.5 | 0 | 0 | 1 | 0.5 |

Macro-level Variables. Period: 1990-2017. $\Delta L T-\operatorname{Debt}_{t+h}$ is the change in the aggregate LT debt share (i.e., the fraction of debt with maturity above 1 year) between year-quarter $t-1$ and year-quarter $t+h, h=0,1, \ldots .4 . \Delta E F F R_{t}$ is the quarterly variation in the Effective Funds Rate. $\varepsilon_{t}$ is the 30 -minute surprise in FED-Funds futures around policy announcements from Gürkaynak et al. (2005) (aggregated at the quarterly frequency). $\triangle G D P_{t-1}$ is the lagged annual GDP growth rate. $\triangle C P I$ is the lagged annual inflation rate. $\operatorname{Rec}_{t-1}$ is a lagged recession dummy. $\Delta L T$ - Treas $_{t-1}$ is the lagged quarterly change in the share of Treasuries with maturity above 10 years. $\Delta i_{t-1}^{10 y-3 m}$ is the lagged quarterly variation of the difference between the 10-year and the 3-month yield on benchmark US Treasuries (term spread). $\Delta i_{t-1}^{b a a-a a a}$ is the lagged quarterly variation of the difference between the BAA and the AAA Moody's Seasoned Corporate Bond Yield (corporate spread).

Firm-level Variables. Period: 1990-2017. Sample: Compustat non-financial companies identified as in Ottonello \& Winberry (2020). $\Delta L T$ - $\operatorname{Debt~}_{f, t+h}$ is the change in firm $f^{\prime}$ s LT debt share (i.e., the fraction of debt with maturity above 1 year) between year-quarter $t-1$ and year-quarter $t+h, h=0,1, \ldots .4 . \Delta$ Sales $_{f, t-1}$ is the lagged quarterly change in log sales, expressed in p.p.. Liquid Assets $f_{f, t-1}$ is the lagged share of liquid assets over total assets. Leverage $f_{f, t-1}$ is the lagged ratio between total debt and total assets. Size $f_{f, t-1}$ is the lagged $\log$ assets size. $\mathbb{1}\left(\right.$ BondIssue $\left._{f, t+h}\right)$ is a dummy variable with value 1 if firm $f$ issues bonds with maturity above 1 -year in year-quarter $t+h$ and with value 0 otherwise, $h=0,1, \ldots .4$. BondSpreadf, $t+h$ is the spread at issuance (vis-a-vis US Treasuries yield with comparable maturity) on the bonds issued by firm $f$ in year-quarter $t+h, h=0,1, \ldots .4$. For the bank loan-related variables $\left(\mathbb{1}\left(\right.\right.$ LoanIssue $_{f, t}$, Bank Maturity $_{f, t}$, AIDSpread $_{f, t}$, Bank Loan Amount $\left.{ }_{f, t}\right)$ the sample is the NFC sample of Ottonello \& Winberry (2020), merged with the LPC DealScan database of syndicated loans, as in Chava \& Roberts (2008). $\mathbb{1}$ (LoanIssue $)_{f, t}$ is a dummy variable with value 1 if firm $f$ issues loans in year-quarter $t+h$ and with value 0 otherwise, Bank Maturity $f_{f, t}$ is the value-weighted maturity of new bank loans contracted by a firm in a year-quarter (weighted to account for the cases in which a firm took out several loans with different maturities). AIDSpread $f_{f, t}$ is the all-in-drawn spread over the LIBOR rate, weighted by both value and maturity of the corresponding loans for firm $f$ in year-quarter $t+h, h=0,1, \ldots .4$. Bank Loan Amount $f, t$ is simply the contracted loan amount in $\$$ Million. If reported
in a non-dollar currency, the dollar value is calculated by using the associated dollar exchange rate, provided by DealScan. StockReturn $_{f, t}$ is the $99 \%$-winsorized stock returns of firm $f$ in year-quarter $t$, taken from CRSP, while Bond $d_{f, t}$ is a dummy that denotes whether firm $f$ issued bonds in the Mergent FISD data set prior to year-quarter $t$. The dummies Refinancing $f_{f, t}$ and EarlyRefinancing $f_{f, t}$ denote if firm $f$ has refinanced (early) any of its outstanding bonds in $t$. Refinancing is defined as in Xu (2018): if firm $f$ retires a bond and within a three-month window around the retirement issues a new bond within $+-30 \%$ of the previously retired bond's value, then the retired bond is flagged as refinanced. Early refinancing is defined as occurring more than 6 months before the originally scheduled maturity date. We need sufficient bonds in the sample before and after the date of refinancing, thus the effective sample of the refinancing analysis runs from 1996 to 2013 (out of an original sample of 1990 to 2017).

Mutual Fund-level variables. Period: 2010q2-2018q2. Sample: Corporate Bond Mutual Funds, identified as those with CRSP style categories: I, ICQH, ICQM, ICQY, ICDI, ICDS, or IC. $\Delta M a t u_{m, t+h}$ is the change in the log (weighted) average portfolio maturity of fund $m$ between year-quarter $t-1$ and year-quarter $t+h, h=0,1, \ldots .4 . \Delta C B_{m, t+4}$ is the change in the $\log$ corporate bond holdings between year-quarter $t-1$ and year-quarter $t+h, h=0,1, \ldots 4 . H Y_{m}$ is a dummy with value 1 if a fund $m$ is classified as High-Yield, and 0 otherwise. HY funds are those with Lipper style codes: HY, GB, FLX, MSI, or SFI. TurnoverRatio ${ }_{m, t-1}$ is the lagged fund $m^{\prime}$ s turnover ratio, corresponding to the minimum (of aggregated sales or aggregated purchases of securities), divided by the average 12-month Total Net Assets. ExpenseRatio ${ }_{m, t-1}$ is fund m's lagged expense ratio, i.e. the ratio of the total investment that shareholders pay for the fund's operating expenses. $N A V_{m, t-1}$ is the lagged fund net asset value, i.e. the value of assets minus liabilities. Returns ${ }_{m, t-1}$ reflects the lagged fund m's quarterly returns, computed as the growth in net asset value from one year-quarter to the next.

Insurance Firm Security Holding Variables. Period: 2006-2022. Sample: Regulatory insurance holdings data from the NAIC. Insurer security holdings are reported at the end of the year. We match these data with bond ratings data from Mergent FISD and with bond yields obtained from Refinitiv Datastream. We only consider corporate bonds and analyze $5,106,789$ insurer-bond-year entries for which we have both a rating and a bond yield. $I Y_{b, t}$ is the average yearly yield of the bond, as obtained from Datastream, where the current yield is coupon net of income taxes, relative to the bond price, according to the formula: $I Y=g V\left(1-t_{g}\right) /\left(P E-A\left(1-t_{g}\right)+Q\right)$, where $g$ is the coupon, $V$ the nominal value, $t_{g}$ the income tax rate, $P$ the gross bond price, $A$ accrued interest, and $Q$ the amount to still be paid out e.g. for partly paid stock. $R Y_{b, t}$ is the redemption yield (or yield-to-maturity) that equates the current price with the future coupons, according to $P=\sum_{i}^{N} g_{i} V /(1+R Y)^{L_{i}}$, where $L_{i}$ is the time in years up to coupon $g_{i}$ and $N$ is the number of remaining years of the bond. We match the categorical ratings to numerical values from 1-28, following the method in Becker \& Milbourn (2011) to obtain numerical Ratingb,t. Then we summarize the ratings into the NAIC regulatory categories 1 (best) to 6 (worst) following the NAIC guidelines, as characterized in Becker \& Ivashina (2015), Table 2, which we call NAICCat $t_{b, t}$.

Insurance Firm Variables. Period: 2006-2022. Using the end-of-year regulatory reported corporate bond holdings of insurance companies, as provided by the NAIC (Schedule D, Part 1), we construct a quarterly dataset, using the within-year purchases and acquisitions (Schedule D, Part $3 \& 5$ ), as well as sales and disposals (Schedule D, Part $4 \& 5$ ), to track the portfolios of insurers on an insurer-quarter level. We define ParValuePurAc $c_{i, t}$ as the total par value of all bonds purchased and acquired in $t$ as a share of total invested assets, ParValueSoldDisp $i_{i, t}$ as the total par value of all bonds sold and disposed of within $t$ as a share of total invested assets. PVNet $i_{i, t}=$ ParValuePur Ac $c_{i, t}-$ ParValueSoldDisp $_{i, t}$. HoldingLogmaturityPurAc $i_{i, t}$ is the $\log$ of the par-value-weighted mean maturity of purchased and acquired bonds in $t$, HoldingLogmaturitySoldDisp $p_{i, t}$ is the log of the par-value-weighted mean maturity of sold and disposed of bonds in $t$. Logmat $\mathrm{L}_{i, t}$ is the log difference of the par-value-weighted mean maturity of purchased and acquired bonds and the par-value-weighted mean maturity of sold and disposed of bonds in $t$. PrimShare ${ }_{i, t}$ is the share of bonds purchased directly from the primary market. We define a purchase from the primary market as a purchase that occurs immediately at or before the date of the offering, as registered in Mergent FISD. Insurer AssetSize $i_{i, t-1}$ are the logged total assets of the insurer, at $t$. We calculate insurers' capital ratio from insurance company liabilities data provided by the NAIC. The insurer capital ratio InsurerCapital Ratio $i_{i, t-1}$ is calculated by dividing total capital as reported to NAIC by the required regulatory capital. $r_{i, y}$ denotes the value-weighted yield of the corporate bond portfolio of insurer $i$ in year $y$. YieldSeeking ${ }_{i, y}$ is the deviation of $r_{i, y}$ from the non-weighted (to avoid bias towards large insurers) average of all corporate bond yields in $y . H Y_{i, y}$ is a dummy that denotes insurers that are above the median in YieldSeeking ${ }_{i, y}$ in year $y$.
Table A2: Firm-Level Regressions Using $\varepsilon_{t}$

| $\Delta L T-$ Debt $_{\text {f,th }}$ |  |  |  |  |  |  |  |  |  |  |  |  |  |
| :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: |
| $h=$ | (0) | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) | (15) | (20) |
| Large $_{f, t-1}$ | -0.013 | $-0.343^{*}$ | $-0.500^{* *}$ | $-0.507^{*}$ | -0.472 | -0.639* | $-0.885^{* *}$ | $-0.889 * *$ | -0.905* | $-1.359 * * *$ | -1.575*** | 229*** | 219 |
|  | (0.116) | (0.185) | (0.246) | (0.292) | (0.337) | (0.375) | (0.402) | (0.435) | (0.473) | (0.502) | (0.530) | (0.642) | (0.711) |
| $\varepsilon_{t} *$ Large $_{f, t-1}$ | -0.206 | $-2.185^{* *}$ | $-2.521^{* *}$ | $-2.928^{* *}$ | $-3.366^{* *}$ | $-3.970^{* * *}$ | $-5.416^{* * *}$ | $-4.030^{* *}$ | $-5.701^{* * *}$ | $-5.100^{* * *}$ | -4.451** | 0.672 | 119** |
|  | (0.616) | (0.867) | (1.041) | (1.228) | (1.355) | (1.450) | (1.562) | (1.680) | (1.754) | (1.846) | (1.825) | (1.942) | (2.032) |
| $\Delta$ Sales $_{\text {f,t-1 }}$ | 0.013*** | $0.026^{* * *}$ | $0.006^{* *}$ | $0.006{ }^{* *}$ | 0.019*** | $0.022^{2 * *}$ | 0.004 | 0.005 | 0.014*** | 0.019 ** | 0.005 | 0.002 | 0.016*** |
|  | (0.002) | (0.003) | (0.003) | (0.003) | (0.003) | (0.004) | (0.004) | (0.003) | (0.004) | (0.004) | (0.004) | (0.004) | (0.005) |
| $\varepsilon_{t} * \Delta$ Sales $_{\text {f,t-1 }}$ | 0.004 | 0.021 | 0.003 | 0.008 | 0.023 | 0.026 | 0.005 | -0.008 | 0.015 | 0.029 | $0^{0.066 *}$ | 0.022 | -0.018 |
|  | (0.017) | (0.021) | (0.025) | (0.027) | (0.031) | (0.030) | (0.031) | (0.032) | (0.033) | (0.036) | (0.039) | (0.040) | (0.040) |
| Liquid Assets $_{\text {f,t-1 }}$ | $-0.018^{* * *}$ | $-0.039 * * *$ | $-0.054 * * *$ | $-0.068^{* * *}$ | $-0.079^{* * *}$ | $-0.083^{* * *}$ | -0.080 *** | $-0.079^{* * *}$ | $-0.082^{* * *}$ | $-0.094 * *$ | $-0.0988^{* * *}$ | $-0.092^{* * *}$ | $-0.084^{* * *}$ |
|  | (0.004) | (0.006) | (0.009) | (0.010) | (0.012) | (0.013) | (0.014) | (0.015) | (0.016) | (0.018) | (0.018) | (0.023) | (0.026) |
| $\varepsilon_{t} *{\text { Liquid } \text { Assets }_{\text {f,t-1 }}}^{\text {d }}$ | -0.025 | -0.023 | -0.012 | 012 | 0.052 | 0.035 | 0.060 | 0.033 | 0.086 | 0.054 | -0.005 | 0.026 | 0.108 |
|  | (0.024) | (0.033) | (0.043) | (0.056) | (0.054) | (0.063) | (0.067) | (0.069) | (0.073) | (0.078) | (0.080) | (0.080) | (0.084) |
| Leverage $_{\text {f,t-1 }}$ | -0.000 | -0.001 | -0.001 | -0.004 | $-0.007$ | -0.007 | -0.008 | -0.011 | -0.009 | $-0.013$ | -0.016* | $-0.032^{* * *}$ | $-0.047^{* * *}$ |
|  | (0.001) | (0.002) | (0.003) | (0.004) | (0.005) | (0.006) | (0.006) | (0.007) | (0.007) | (0.008) | (0.008) | (0.010) | (0.011) |
|  | -0.010 | -0.012 | ${ }^{-0.016}$ | -0.015 | -0.014 | -0.021 | ${ }^{-0.025}$ | -0.051* | $-0.069 * *$ | $-0.105^{* * *}$ | $-0.088^{* *}$ | $-0.062^{*}$ | $-0.065^{*}$ |
|  | (0.010) | (0.014) | (0.016) | (0.020) | (0.021) | (0.026) | (0.026) | (0.029) | (0.030) | (0.033) | (0.034) | (0.037) | (0.033) |
| Observations | 320,108 | 292,677 | 277,849 | 265,263 | 255,423 | 247,247 | 241,564 | 232,489 | 224,179 | 216,403 | 210,684 | 179,253 | 153,430 |
| R-squared | 0.093 | 0.119 | 0.138 | 0.155 | 0.170 | 0.183 | 0.192 | 0.201 | 0.212 | 0.223 | 0.230 | 0.275 | 0.319 |

In column $h$, the dependent variable is $\Delta L T-$ Debt $_{f, t+h}$, i.e., the change in firm $f^{\prime}$ s share of LT debt between year-quarter $t-1$ and $t+h$. $\varepsilon_{t}$ is the Gürkaynak et al. (2005) high-frequency monetary policy surprise. Large $f_{f, t-1}$ is a dummy with value 1 if firm $f$ is in the top asset-size quartile of the respective ( 3 -digit SIC) industry. $\Delta$ Sales $_{f, t-1}$ is the quarterly variation in firm


Table A3: Firm-Level Regressions using $\Delta E F F R_{t}$

|  |  |  |  |  |  | $\Delta L T$ - D | Debt $f_{f t+h}$ |  |  |  |  |  |  |
| :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: |
| $h=$ | (0) | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) | (15) | (20) |
| Large $_{f, t-1}$ | 0.104 | -0.184 | -0.333 | -0.234 | -0.070 | -0.263 | -0.576 | -1.020 | -1.229* | $-2.048^{* * *}$ | $-2.676 * * *$ | -4.471*** | $-4.189^{* * *}$ |
|  | (0.207) | (0.332) | (0.434) | (0.518) | (0.569) | (0.599) | (0.634) | (0.683) | (0.734) | (0.757) | (0.788) | (0.914) | (1.026) |
| $\triangle E F F R_{t} *$ Large $_{f, t-1}$ | 0.006 | -0.492** | -0.899*** | -0.873** | $-1.376^{* * *}$ | $-1.601^{* * *}$ | $-2.151^{* * *}$ | -1.538*** | -1.799*** | -1.716*** | -1.571*** | -0.006 | -0.018 |
|  | (0.163) | (0.242) | (0.306) | (0.368) | (0.399) | (0.421) | (0.435) | (0.454) | (0.470) | (0.484) | (0.489) | (0.547) | (0.544) |
| $\Delta$ Sales $_{f, t-1}$ | 0.013*** | 0.024** | 0.005* | 0.006** | $0.018^{* * *}$ | 0.021*** | 0.004 | 0.005 | $0.014^{* * *}$ | 0.018*** | 0.003 | 0.001 | 0.016*** |
|  | (0.002) | (0.002) | (0.003) | (0.003) | (0.003) | (0.003) | (0.003) | (0.003) | (0.004) | (0.004) | (0.004) | (0.004) | (0.005) |
| $\Delta E F F R_{t} * \Delta$ Sales $_{f, t-1}$ | -0.000 | -0.000 | -0.004 | 0.003 | 0.003 | 0.006 | 0.003 | -0.001 | 0.004 | 0.001 | 0.005 | 0.004 | -0.006 |
|  | (0.004) | (0.005) | (0.006) | (0.006) | (0.007) | (0.007) | (0.007) | (0.007) | (0.008) | (0.008) | (0.009) | (0.009) | (0.009) |
| ${\text { Liquid } \text { Assets }_{f, t-1}}$ | -0.017*** | $-0.038 * * *$ | $-0.053 * * *$ | $-0.067^{* * *}$ | -0.079*** | $-0.083^{* * *}$ | $-0.080 * * *$ | $-0.079 * * *$ | $-0.084^{* * *}$ | $-0.094^{* * *}$ | $-0.096^{* * *}$ | $-0.091^{* * *}$ | $-0.084^{* * *}$ |
|  | (0.004) | (0.006) | (0.008) | (0.010) | (0.012) | (0.013) | (0.014) | (0.015) | (0.016) | (0.017) | (0.018) | (0.023) | (0.026) |
| $\triangle E F F R_{t} *$ Liquid $^{\text {Assets }}$ f,t-1 | 0.002 | 0.001 | 0.003 | 0.019 | 0.029* | 0.020 | 0.023 | 0.021 | 0.025 | 0.026 | 0.019 | 0.008 | 0.035 |
|  | (0.005) | (0.008) | (0.011) | (0.014) | (0.015) | (0.016) | (0.018) | (0.018) | (0.019) | (0.020) | (0.021) | (0.021) | (0.023) |
| Leverage $_{\text {f,t-1 }}$ | -0.000 | -0.001 | -0.002 | -0.004 | -0.007 | -0.007 | -0.008 | -0.010 | -0.008 | -0.011 | -0.014* | $-0.032^{* * *}$ | $-0.047^{* * *}$ |
|  | (0.001) | (0.002) | (0.003) | (0.004) | (0.005) | (0.006) | (0.006) | (0.007) | (0.007) | (0.008) | (0.008) | (0.010) | (0.011) |
| $\triangle E F F R_{t} *$ Leverage $_{f, t-1}$ | -0.003* | -0.002 | -0.003 | -0.000 | -0.002 | -0.001 | 0.002 | -0.006 | -0.006 | -0.013 | -0.010 | -0.009 | -0.016** |
|  | (0.002) | (0.003) | (0.003) | (0.004) | (0.005) | (0.006) | (0.006) | (0.007) | (0.007) | (0.008) | (0.008) | (0.008) | (0.008) |




 fixed effects. Standard errors are clustered at the Firm and Industry* Year-Quarter level. ${ }^{* * *} \mathrm{p}<0.01,{ }^{* *} \mathrm{p}<0.05,{ }^{*} \mathrm{p}<0.1$.

## Table A4: Monetary Policy Shocks and Bond Issuance by Large-vs-Small Firms

|  | $\mathbb{1}\left(\right.$ BondIssue $_{f, t}$ |  |  | $\mathbb{1}^{\left(\text {BondIssue }_{f, t+1}\right.}$ |  |  |
| :--- | :---: | :---: | :---: | :---: | :---: | :---: |
|  | $(1)$ | $(2)$ | $(3)$ | $(4)$ | $(5)$ | $(6)$ |
| $\varepsilon_{t}$ | $-0.039^{* *}$ |  |  | $-0.058^{* * *}$ |  |  |
| Large $_{f, t-1}$ | $(0.017)$ |  |  | $(0.017)$ |  |  |
|  | $0.019^{* * *}$ | $0.019^{* * *}$ | $0.017^{* * *}$ | $0.020^{* * *}$ | $0.019^{* * *}$ | $0.017^{* * *}$ |
| $\varepsilon_{t} *$ Large $_{f, t-1}$ | $(0.003)$ | $(0.003)$ | $(0.003)$ | $(0.003)$ | $(0.003)$ | $(0.004)$ |
|  | $-0.026^{*}$ | -0.025 | $-0.038^{* *}$ | $-0.038^{* *}$ | $-0.037^{* *}$ | $-0.052^{* * *}$ |
|  | $(0.016)$ | $(0.015)$ | $(0.017)$ | $(0.017)$ | $(0.017)$ | $(0.019)$ |
| Observations $_{\text {R-squared }}$ | 114,285 | 114,285 | 108,311 | 107,121 | 107,121 | 100,978 |
| Firm Controls* $\varepsilon_{t}$ | 0.088 | 0.094 | 0.209 | 0.091 | 0.096 | 0.218 |
| Firm FE | Yes | Yes | Yes | Yes | Yes | Yes |
| Year:Quarter FE | Yes | Yes | Yes | Yes | Yes | Yes |
| Industry* Year-Quarter FE | No | Yo | No | - | No | Yes |

In columns (1)-(3), $\mathbb{1}$ (BondIssue $)_{f, t}$ is a dummy variable with value if firm $f$ issues LT bonds in year-quarter $t$. In columns (3)-(6), we use the same dependent variable, though measured as of year-quarter $t+1 . \varepsilon_{t}$ is the Gürkaynak et al. (2005) high-frequency monetary policy surprise. Large $f_{, t-1}$ is a dummy with value 1 if firm $f$ is in the top asset-size quartile of the respective (3-digit SIC) industry. Firm controls include lagged sales growth, leverage and share of liquid assets. Standard errors are clustered at the Firm and Industry*Year-Quarter level. ${ }^{* * *} \mathrm{p}<0.01,{ }^{* *} \mathrm{p}<0.05,{ }^{*} \mathrm{p}<0.1$.

Table A5: Monetary policy and financing costs through LT bonds

|  | BondSpread $_{f, t}$ |  | BondSpread $_{f, t+1}$ |  |
| :---: | :---: | :---: | :---: | :---: |
|  | (1) | (2) | (3) | (4) |
| $\Xi_{t}$ | -27.928* |  | -16.857 |  |
|  | (16.038) |  | (14.720) |  |
| Large $_{f, t-1}$ | 45.177 | $73.669^{* * *}$ | 27.743 | 54.199** |
|  | (29.944) | (26.941) | (25.345) | (24.707) |
| $\Xi_{t} * \operatorname{Large}_{f, t-1}$ | 15.582* | 24.948*** | 9.029 | $21.324^{* * *}$ |
|  | (9.440) | (8.187) | (8.346) | (7.709) |
| Observations | 3,760 | 1,876 | 3,983 | 1,912 |
| R -squared | 0.649 | 0.894 | 0.626 | 0.881 |
| Firm Controls*EFFR | Yes | Yes | Yes | Yes |
| Firm FE | Yes | Yes | Yes | Yes |
| Industry*Year:Quarter FE | No | Yes | No | Yes |

In columns (1)-(2), the dependent variable BondSpread $_{f, t}$ is the bond sprad of firm $f^{\prime}$ s newly issued LT bonds in year-quarter $t$. In columns (3)-(4), the left-hand side variable is the same, though measured in year-quarter $t+1$. $\Xi_{t}$ are the aggregate monetary policy shocks, as in Coibion (2012), a proxy for the current level of the policy rate. Large $f_{f, t-1}$ is a dummy with value 1 if firm $f$ is in the top asset-size quartile of the respective (3-digit SIC) industry. Firm controls include lagged sales growth, leverage, and share of liquid assets. Macro controls are given by annual GDP growth and inflation rate; a recession dummy; term spread, corporate spread, and share of Treasuries with maturity above 20 years. Standard errors are double-clustered at the firm and industry*year-quarter level. ${ }^{* * *} \mathrm{p}<0.01,{ }^{* *} \mathrm{p}<0.05,{ }^{*} \mathrm{p}<0.1$.

Table A6: Monetary policy and financing costs through LT bonds Robustness

| Size Distribution | All Firms |  |  |  | Sectoral |  |  |  | 2-Digit SIC |  |  |  |
| :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: |
| Dependent Variable | BondSp <br> (1) | $\operatorname{rread}_{f, t}$ (2) | $\text { BondSpread }_{f, t+1}$ |  | $\text { BondSpread }_{f, t}$ |  | BondSpread $_{f, t+1}$ |  | BondSpread $_{f, t}$ |  | BondSp <br> (11) | $\operatorname{read}_{f, t+1}$ <br> (12) |
| $\Xi_{t}$ | $\begin{gathered} -174.248^{* * *} \\ (41.079) \end{gathered}$ |  | $-178.784^{* * *}$ (44.875) |  | $\begin{aligned} & -31.489 \\ & (23.276) \end{aligned}$ |  | $\begin{aligned} & -22.986 \\ & (19.509) \end{aligned}$ |  | $\begin{aligned} & -28.898 \\ & (20.123) \end{aligned}$ |  | $\begin{aligned} & -20.703 \\ & (16.224) \end{aligned}$ |  |
| Large $_{f, t-1}$ | $\begin{gathered} 428.053^{* * *} \\ (107.949) \end{gathered}$ | $401.780^{* * *}$ (102.470) | $\begin{gathered} 476.184^{* * *} \\ (119.350) \end{gathered}$ | $447.648^{* * *}$ (113.381) | $\begin{aligned} & 49.186 \\ & (39.607) \end{aligned}$ | $\begin{gathered} 112.868^{* * *} \\ (33.261) \end{gathered}$ | $\begin{aligned} & 28.919 \\ & (35.072) \end{aligned}$ | $92.057^{* * *}$ (28.104) | $\begin{aligned} & 60.288^{*} \\ & (33.449) \end{aligned}$ | $124.414^{* * *}$ <br> (27.716) | $\begin{gathered} 36.743 \\ (29.123) \end{gathered}$ | $98.849^{* * *}$ (24.683) |
| $\Xi_{t} *$ Large $_{f, t-1}$ | $155.452^{* * *}$ (33.349) | $141.375^{* * *}$ (31.431) | $164.969^{* * *}$ (37.056) | $152.888^{* * *}$ (34.934) | $49.186$ (12.700) | $112.868^{* * *}$ (10.482) | $\begin{aligned} & 28.919 \\ & (10.858) \end{aligned}$ | $92.057^{* * *}$ (8.919) | $\begin{aligned} & 60.288^{*} \\ & (33.449) \end{aligned}$ | $124.414^{* * *}$ <br> (27.716) | $\begin{aligned} & 36.743 \\ & (29.123) \end{aligned}$ | $\begin{gathered} 98.849^{* * *} \\ (24.683) \end{gathered}$ |
| Observations | 3,557 | 3,557 | 3,733 | 3,733 | 3,557 | 3,391 | 3,733 | 3,553 | 3,557 | 2,363 | 3,733 | 2,473 |
| R-squared | 0.650 | 0.782 | 0.626 | 0.774 | 0.644 | 0.817 | 0.618 | 0.813 | 0.643 | 0.861 | 0.618 | 0.861 |
| Firm Controls* ${ }^{*} P_{t}$ | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Firm FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Year:Quarter FE | No | Yes | No | Yes | No | - | No | - | No | - | No | - |
| Sector*Year:Quarter FE | No | No | No | No | No | Yes | No | Yes | - | - | - | - |
| Sic2*Year:Quarter FE | No | No | No | No | No | No | No | No | No | Yes | No | Yes |

In columns (1)-(2), (5)-(6) and (9)-(10), the dependent variable is the bond spread on firm $f^{\prime}$ 's newly issued LT bonds in yearquarter $t$. In columns (3)-(4), (7)-(8), and (11)-(12), the left-hand side variable is the same, though measured in year-quarter $t+1$. $\Xi_{t}$ are the aggregate monetary policy surprises, as in Coibion (2012), a proxy for the current level of the policy rate. Large $_{f, t-1}$ is a dummy with value 1 if firm $f$ is in the top quartile of the asset-size distribution. As indicated in the top row "Size Distribution", in columns (1)-(4) the relevant distribution includes all firms; in columns (5)-(8), the large dummy is computed within sectors (i.e., 1-digit SIC code); finally, in columns (9)-(12) within 2-digit SIC code industries. Firm controls include lagged sales growth, leverage, and share of liquid assets. Standard errors are double-clustered at the: firm and yearquarter level in columns (1)-(4); firm and sector level in columns (5)-(8); firm and (2-digit SIC)-industry*year-quarter level in columns (9)-(12). ${ }^{* * *} \mathrm{p}<0.01,{ }^{* *} \mathrm{p}<0.05,{ }^{*} \mathrm{p}<0.1$.

## Figures

Figure A1: Measures of Changes in the Policy Rate


The gray bars show the quarterly variation of the Effective FED Funds Rate. The black line reports the Gürkaynak et al. 2005 monetary policy surprises.

Figure A2: Monetary Policy and Debt Maturity Structure: Aggregate Response - Pre-Crisis Period
(a) Quarterly Variation in Effective Federal Funds Rate: $\Delta E F F R_{t}$

(B) GÜRKAYNAK ET AL. (2005) SURPRISES: $\varepsilon_{t}$


This figure depicts the IRF of the aggregate-level share of LT debt to a change in monetary policy. For this robustness check the sample is restricted to the pre-crisis period (the baseline sample prior to Q4 2008). Both Panel A and Panel B show the coefficients $\beta_{1, h}$ from the estimation of Equation $1, h=0,1, \ldots, 20$, with the proxy of monetary policy change (i.e., $\Delta M P_{t}$ ) given, respectively, by the quarterly variation in the Effective Federal Funds Rate, $\triangle E F F R_{t}$ and the Gürkaynak et al. (2005) monetary policy surprises, $\varepsilon_{t}$. In Panel A, the IRF is calibrated to a $\triangle E F F R_{t}=-25$ b.p. change. In Panel B, the IRF is calibrated to a 1 s.d. expansionary monetary policy surprise, i.e. $\varepsilon_{t}=-10$ b.p.. The solid line reports the point estimates for $\beta_{1, h}$; the dashed lines the $90 \%$ confidence intervals. In both panels, we apply robust standard errors.

Figure A3: Monetary Policy and Debt Maturity Structure - Relative Response of Large Companies: Pre-Crisis
(a) Quarterly Variation in Effective Federal Funds Rate: $\Delta E F F R_{t}$

(B) GÜRKAYNAK ET AL. (2005) SURPRISES: $\varepsilon_{t}$


This figure depicts the relative response of large firms to a monetary policy shock (as compared to smaller firms). For this robustness check the sample is restricted to the pre-crisis period (the baseline sample prior to Q4 2008). Both Panel A and Panel B show the coefficients $\beta_{2, h}$ from the estimation of Equation $2, h=0,1, \ldots, 20$, with the proxy of monetary policy change (i.e., $\Delta M P_{t}$ ) given, respectively, by the quarterly variation in the Effective Federal Funds Rate, $\triangle E F F R_{t}$ and the Gürkaynak et al. (2005) monetary policy surprises, $\varepsilon_{t}$. In Panel A, the IRF is calibrated to a $\Delta E F F R_{t}=-25$ b.p. change. In Panel B, the IRF is calibrated to a 1 s.d. expansionary monetary policy surprise, i.e. $\varepsilon_{t}=-10$ b.p.. In both panels, we double-cluster standard errors at the firm and industry* year-quarter level.

Figure A4: Monetary Policy and Bond Issuance - Relative Response of Large Companies: Pre-Crisis

## (A) Likelihood of Issuing Bonds


(B) Yield Spread


This figure depicts the relative response in the likelihood of issuing bonds and in the related yield spread to an expansionary monetary policy shock for large firms (as compared to smaller firms). In this robustness check the sample is restricted to the pre-crisis period (before Q4 2018). Panel A shows the coefficients $\beta_{2, h}$ from the estimation of Equation 2, $h=0,1, \ldots, 20$, with $\mathbb{1}$ (BondIssue $)_{f, t+h}$ as dependent variable, i.e. a dummy variable with value 1 if firm $f$ issues LT bonds in year-quarter $t+h$ and with value 0 if it does not. The IRF is calibrated to a 1 s.d. expansionary Gürkaynak et al. (2005) monetary policy surprise ( $\varepsilon_{t}=-10 b$.p.). Panel B shows the coefficients $\beta_{2, h}$ from the estimation of Equation $6, h=0,1, \ldots, 20$, with $\mathbb{1}(\text { BondSpread })_{f, t+h}$ as dependent variable. The IRF is calibrated to a $1 \mathrm{~s} . \mathrm{d}$. variation in the cumulative Gürkaynak et al. (2005) monetary policy surprises, $\Xi_{t}=-90 \mathrm{~b} . \mathrm{p}$. . In both panels, the solid line represents the point estimates, and the dashed lines the $90 \%$ confidence interval. Standard errors are double-clustered at the firm and industry*year-quarter level.

## IA. Internet Appendix

## IA.1. Additional Empirical Results

## Figure IA1: Monetary Policy and the Term Spread



This figure depicts the response of the term spread to a change in the monetary policy rate. We employ several definitions of the term spread, all based on the benchmark Treasury yields at constant maturity. The 1 -year $/ 3$-month spread is in the top-left sub-plot; the 5 -year/3-month spread is in the top-right sub-plot; the 10 -year/3-month spread is in the bottom-left sub-plot, whereas the 20 -year/3-month spread is displayed on the bottom-right. Formally, we shows the coefficients $\beta_{1, h}$ from the estimation of the following local projection model:

$$
\Delta_{h} y_{t+h}=\beta_{1, h} \varepsilon_{t}+\text { MacroControls }_{t-1}+u_{t, h}
$$

The dependent variable, $\Delta_{h} y_{t+h}$, represents the growth of the term spread (expressed in p.p.) from year-quarter $t-1$ to yearquarter $t+h . \varepsilon_{t}$ is the Gürkaynak et al. (2005) monetary policy surprise at time $t$. MacroControls $s_{t-1}$ is a vector of lagged macroeconomic controls, including annual GDP growth and inflation rate, a dummy for recessions, the quarterly variation in the share of LT Treasuries and in the corporate spread. $u_{t, h}$ is a robust error term. The $x$-axis is measured in terms of quarters after the shock. The solid lines report the point estimates for $\beta_{1, h}$; the dashed lines the $90 \%$ confidence intervals.

Figure IA2: Monetary Policy and the Term Premium


This figure depicts the response of the term premium to a 1 s.d. reduction in the monetary policy shock. We employ estimates of the term premium from Adrian et al. (2013). The 1-year term premium is in the top-left sub-plot. The 2-year term premium is in the top-right sub-plot. The 5 -year term premium is in the bottom-left sub-plot, whereas the 10 -year term premium is displayed on the bottom-right. Formally, it shows the coefficients $\beta_{1, h}$ from the estimation of the following local projection model:

$$
\Delta_{h} y_{t+h}=\beta_{1, h} \varepsilon_{t}+\text { MacroControls }_{t-1}+u_{t, h}
$$

The dependent variable, $\Delta_{h} y_{t+h}$, represents the growth of the term premium (expressed in p.p.) from year-quarter $t-1$ to year-quarter $t+h . \varepsilon_{t}$ is the Gürkaynak et al. (2005) monetary policy surprise at time $t$. MacroControls $s_{t-1}$ is a vector of lagged macroeconomic controls, including annual GDP growth and inflation rate, a dummy for recessions, the quarterly variation in the share of LT Treasuries, and in the corporate spread. Moreover, to account for the fact that the term premium may spuriously reflect changes in (the expectation-hypothesis component of) the term spread, we also include the lagged $h$-quarter growth in the term spread. $u_{t, h}$ is a robust error term. The $x$-axis is measured in terms of quarters after the shock.

## Figure IA3: Insurer Net Purchases after an Expansionary Shock Reach for Yield by NAIC category.

## (A) Par Value


(в) MATURITY

(C) Primary Market

SHARE


This figure shows the interaction effect of a 1 s.d. expansionary monetary policy surprise and the indicator $H Y_{i, y-1}$, which denotes whether insurer $i$ was in the top $50 \%$ of deviations of the par value-weighted yield-to-maturity of their corporate bond portfolio from the average yield of all corporate bonds, within NAIC-regulatory category, in year $y-1$, on US insurers' purchases of corporate bonds. The coefficient of interest is $\beta_{2, h}$ in the regression models described in Equation 9 , in which we proxy monetary policy changes $\Delta M P_{t}$ with the Gürkaynak et al. (2005) monetary policy surprises, $\varepsilon_{t}$. The dependent variable in Panel A is the par value of purchased and acquired corporate bonds net of the par value of sold and disposed-of corporate bonds. The dependent variable in Panel B is the log maturity of purchased and acquired corporate bonds net of the maturity of sold and disposed-of corporate bonds. Panel C reports the effect on the share of corporate bonds purchased from the primary market. Throughout we include insurer fixed effects, as well as insurer asset size and capital ratios interacted with $\Delta M P_{t}$ as control variables and cluster standard errors at the insurer level. Missing values for insurer capital ratios are imputed by using the quarterly mean value for the capital ratio for the missing insurer-year pair. Displayed are $90 \%$ confidence intervals.

## IA.2. Corporate Bond Holdings by Corporate Bond Mutual Funds

## IA.2.1. DATA

We retrieve data on corporate bond mutual funds (CBMF) holdings from the CRSP Survivor Bias-Free dataset, including information on both active and inactive funds. We split funds into High Yield (HY) and Investment Grade (IG) funds based on standard Lipper style codes. ${ }^{29}$ We label HY funds as yield-seeking: as shown by Choi \& Kronlund (2018), they invest relatively more in longer and riskier debt securities, i.e. they reach for yield relatively more. Overall, we analyze 3,487 funds ( 2,034 are IG and 1,453 are HY) over the data period 2010q2-2018q2. Appendix Table A1 describes the related summary statistics. A first outcome variable of interest is the cumulative growth rate of the volume of corporate bond holdings through time, $\Delta C B_{m, t+h}$, which displays a large extent of heterogeneity across funds. Second, we retain data on the changes in the fund's average (weighted) portfolio maturity over time, $\Delta M a t u_{m, t+h}$, equally showing significant differences in the cross-section of funds. We gather additional information on fund characteristics, used as controls in our models, including the fund's turnover and expenses ratio, the net asset value, and returns.

## IA.2.2. Empirical Analysis

We employ the following model:

$$
\begin{equation*}
\Delta y_{m, t+h}=\beta_{1, h} \Delta M P_{t} * H Y_{m}+\Gamma_{h} X_{m, t-1}+\mu_{m}+\mu_{y q}+e_{m, t+h} \tag{10}
\end{equation*}
$$

The dependent variable, $\Delta y_{m, t+h}$, is the growth between year-quarter $t-1$ and $t+h$ of fund $m$ 's $\log$ volume of corporate bond holdings (or their log average weighted maturity). Our coefficient of interest is $\beta_{1, h}$, loading the interaction between the monetary policy rate variation, $\Delta M P_{t}$, and a dummy, $H Y_{m}$, with value 1 if fund $m$ is high-yield, our proxy for yield-seeking mutual funds. Importantly, $\beta_{1, h}$ captures the relative response of high-yield

[^20]mutual funds' portfolios (as compared to investment-grade ones) to modifications of the interest rate, therefore sizing the impact of reach-for-yield motives on such relation. $X_{m, t-1}$ is a vector of time-varying fund-level controls, including the interaction of: i) $H Y_{m}$ with macro-level controls; ii) other lagged fund characteristics (turnover ratio, expense ratio, log asset size and returns) with $\Delta M P_{t}$. Moreover, we augment the model with fund and yearquarter fixed effects ( $\mu_{m}$ and $\mu_{y q}$ ), respectively controlling for time-invariant heterogeneity at the level of the fund and for common shocks across all funds in a given year-quarter. $e_{m, t+h}$ is an error term, double-clustered at the fund and year-quarter level.

## IA.2.3. Results

The following Figure IA4 reports the estimated IRFs.

## Figure IA4: Monetary policy and CBMFs' Debt Securities Holdings HOLDINGS

(a) Corporate Bonds Holdings

(b) Average Maturity of Debt Securities Holdings


This figure shows the relative response of High-Yield (HY) corporate bonds mutual funds to a 1 s.d. expansionary monetary policy shock (as compared to Investment-Grade funds). Formally, it shows the coefficients $\beta_{1, h}$ from the estimation of local projection model 10, in which we proxy monetary policy changes $\Delta M P_{t}$ with the Gürkaynak et al. (2005) monetary policy surprises, $\varepsilon_{t}$. In Panel A, the dependent variable is given by the growth between year-quarter $t-1$ and $t+h$ of fund $m^{\prime}$ s log volume of corporate bond holdings, expressed in p.p.. In Panel B, the dependent variable is given by the growth between year-quarter $t-1$ and $t+h$ of fund $m$ weighted average maturity. The solid lines report the point estimates for $\beta_{1, h}$; the dashed line the $90 \%$ confidence intervals, based on s.e. double-clustered at the fund and year-quarter level.

## IB. Theoretical Model: Discussion of the Setup, Proofs and Extensions

## IB.1. Discussion of the Setup

The model by Holmström \& Tirole $(1998,2000)$, provides a tractable framework to study how the maturity structure of corporate debt interacts with financial constraints and investors demand. In this model, the defining difference between short-term and long-term financing is credit risk that affects only debt of longer maturity. As do Holmström \& Tirole $(1998,2000)$, we abstract from other sources of risk, such as duration or rollover risk. In Section IB.4, we show the robustness of our results to explicitely adding rollover risk in our framework.

Our goal is to have a simple model that endogenizes the corporate sector's maturity structure. Since our investors are risk-averse and thus more complex than those in Holmström \& Tirole $(1998,2000)$, we make some simplifying assumptions, namely that there is no storage technology, and that short-term debt is riskless. The latter can be rationalized when the intermediate cash flow $r$ is large enough. In other words, the firm must be sufficiently "cash-rich". The absence of rollovers in the baseline model allows us to solve for a well-defined maturity structure and is without loss of generality: the contract specifies all contingencies, which are known by both firms and investors in $t=0$. Thus, there will be no incentive to refinance once the liquidity shock has been realized.

There has long been an established theoretical literature on firms' optimal debt maturity choice such as Flannery (1986, 1994), Diamond (1991), Diamond \& He (2014), He \& Milbradt (2016). The aforementioned papers do not focus on the role of firms' financial constraints for corporate debt maturity, which our empirical evidence suggests is the central determinant of the effect of monetary policy on debt maturity. Other papers consider the effect that a given debt maturity has on financial outcomes, such as rollover risk and credit risk He \& Xiong (2012b), He \& Xiong (2012a). Relative to these papers, we are exploring the relationship in the opposite direction in that we are trying to understand how changes in financing conditions affect maturity choices.

Moreover, as we show in section IB.4, our model without yield-seeking investors does not match our empirical facts. This is our motivation to extend the model such that investor demand is sensitive to interest rate changes by introducing risk aversion and reach-for-yield behavior. ${ }^{30}$

## IB.2. Equilibrium and First Order Conditions

First, consider the firms' maximization problem

$$
\begin{equation*}
\max _{\rho^{*}(A), d_{s}(A), d_{l}(A)} \frac{r}{1+i_{1}}-\frac{\int_{0}^{\rho^{*}(A)} \rho f(\rho) d \rho}{1+i_{1}}+\frac{\delta F\left(\rho^{*}(A)\right)}{\left(1+i_{1}\right)\left(1+i_{2}\right)} R+\left(P_{l}(A)-\frac{\delta F\left(\rho^{*}(A)\right)}{\left(1+i_{1}\right)\left(1+i_{2}\right)}\right) d_{l}(A) \tag{11}
\end{equation*}
$$

subject to the LL, IC, IR constraints.
The objective function represents expected firm profits as of $t=0$. The first term is the risk-free revenue at $t=1$ and the second subtracts the expected $t=1$ payments due to the liquidity shock. The third term is the expected $t=2$ revenue, influenced by both idiosyncratic liquidity shock and aggregate risk. The last term of the equation collects the net proceeds from the issuance of LT bonds, i.e., their price minus total expected repayments. Moreover, recognize that the explicit mean and variance of investor wealth at $t=2$ are:

$$
\begin{aligned}
& E\left[w^{j}\right]=\delta \int_{0}^{I} F\left(\rho^{*}(A)\right) d_{l}^{j}(A) d A-\iota^{j}\left(i_{1}, i_{2}\right) \int_{0}^{I} P_{l}(A) d_{l}^{j}(A) d A \\
& \operatorname{Var}\left[w^{j}\right]=\operatorname{Var}\left[d_{l}^{* j}\right]=\left(\int_{0}^{I} F\left(\rho^{*}(A)\right) d_{l}^{j}(A) d A\right)^{2} \delta(1-\delta) .
\end{aligned}
$$

In the expression for expected wealth, the expected revenues reflect the fact that, for a firm with endowment $A$, the likelihood of repayment equals $\delta F\left(\rho^{*}(A)\right)$. Next, in the variance term, $\int_{A} F\left(\rho^{*}(A)\right) d_{l}^{j}(A)$ is treated like a constant due to full diversification of firms' idiosyncratic risk; in other terms, investors' risk only depends on aggregate shocks. By plugging in

[^21]these expressions into the objective function (4), we arrive at the following problem
\[

$$
\begin{equation*}
\max _{d_{l}^{j}(A)} \delta \int_{0}^{I} F\left(\rho^{*}(A)\right) d_{l}^{j}(A) d A-\iota^{j}\left(i_{1}, i_{2}\right) \int_{0}^{I} P_{l}(A) d_{l}^{j}(A) d A-\frac{\gamma}{2}\left(\int_{0}^{I} F\left(\rho^{*}(A)\right) d_{l}^{j}(A) d A\right)^{2} \delta(1-\delta) . \tag{12}
\end{equation*}
$$

\]

Definition 1. A competitive equilibrium is a set of quantities $\left\{d_{s}(A), d_{l}(A), d_{l}^{R}(A), d_{l}^{Y}(A)\right\}_{A \in[0, I]}$, cut-off rules $\left\{\rho^{*}(A)\right\}_{A \in[0, I]}$ and prices $\left\{P_{l}(A)\right\}_{A \in[0, I]}$ such that:

1. $\left\{d_{s}(A), d_{l}(A), \rho^{*}(A)\right\}_{A \in[0, I]}$ solve firms' optimization problem (11), given $\left\{P_{l}(A)\right\}_{A \in[0, I]}$.
2. $\left\{d_{l}^{R}(A), d_{l}^{Y}(A)\right\}_{A \in[0, I]}$ solve rational and yield-seeking investors' respective maximization problems (12).
3. The LT bond market clears:

$$
\begin{equation*}
d_{l}(A)=\alpha d_{l}^{R}(A)+(1-\alpha) d_{l}^{Y}(A)+\frac{g}{\int_{0}^{I} F\left(\rho^{*}(A)\right) d A} \cdot{ }^{31} \tag{13}
\end{equation*}
$$

The FOCs for rational and yield-seeking investors are, respectively:

$$
\begin{equation*}
P_{l}(A)=\frac{\delta F\left(\rho^{*}(A)\right)-\gamma F\left(\rho^{*}(A)\right) \delta(1-\delta) \int_{0}^{I} F\left(\rho^{*}(A)\right) d_{l}^{j}(A) d A}{\dot{\nu}\left(i_{1}, i_{2}\right)} \quad \forall j \in\{R, \gamma\} \tag{14}
\end{equation*}
$$

where $d_{l}^{j}(A)$ is the demand for firm $A^{\prime}$ 's bonds by investor type $j$. Rearranging equations (14) and plugging them into the market clearing condition (13) yields the inverse demand for firm $A$ 's LT debt (5). Taking first order conditions of the firms' problem, we get
$\lambda_{3}(A)=\lambda_{1}(A)$
$P_{l}(A)=\frac{1+\lambda_{2}(A)}{1+\lambda_{3}(A)} \frac{\delta F\left(\rho^{*}(A)\right)}{\left(1+i_{1}\right)\left(1+i_{2}\right)}$
$f\left(\rho^{*}(A)\right)\left[\frac{\delta R}{1+i_{2}}-\rho^{*}(A)\right]+\left(\left(1+i_{1}\right)\left(1+\lambda_{3}(A)\right) \frac{\partial P_{l}(A)}{\partial \rho^{*}(A)}-\frac{\delta f\left(\rho^{*}(A)\right)}{1+i_{2}}\right) d_{l}(A)-\lambda_{1}(A)=0$
where $\lambda_{1}, \lambda_{2}, \lambda_{3}$ are the multipliers linked to the three constraints LL, IC and IR, respectively. Condition (15) signals that LL binds if and only if IR does. In Equation (16), the firm val-

[^22]uation of one unit of LT debt equals the NPV of the risky project times a factor positively (negatively) related to the tightness of the IC (IR) constraint. Finally, in Condition (17), the optimal liquidation cutoff decreases when the LL constraint tightens.

## IB.3. Proofs

For the proof of Proposition 1 we will find it useful to first introduce three Lemmas.
LEMMA 1. The price for LT debt of a firm with endowment $A$ is unaffected by changes in the supply of $L T$ debt $d_{l}(A)$ and increases in $\rho^{*}(A)$.

Proof of Lemma 1. Firms' endowments are distributed uniformly and continuously on $[0, I]$. This implies that each firm is atomistic relative to the set of all firms. The expected return of the market portfolio thus does not respond to changes in a single firm's $d_{l}(A)$ and $\rho^{*}(A)$ :

$$
\frac{\partial \int_{0}^{I} P_{l}(A) d_{l}(A) d A}{\partial d_{l}(A)}=0 \quad \frac{\int_{0}^{I} P_{l}(A) d_{l}(A) d A}{\partial \rho^{*}(A)}=0 .
$$

Taking derivatives of (5) yields

$$
\frac{\partial P_{l}(A)}{\partial d_{l}(A)}=0 \quad \frac{\partial P_{l}(A)}{\partial \rho^{*}(A)}=\delta f\left(\rho^{*}(A)\right) \frac{1-\gamma(1-\delta)\left(\int_{0}^{I} F\left(\rho^{*}(A)\right) d_{l}(A) d A-g\right)}{\left(1+i_{1}\right)\left[\alpha\left(1+i_{1}\right)+(1-\alpha)\left(1+i_{2}\right)\right]} .
$$

Note that $\partial P_{l}(A) / \partial \rho^{*}(A)>0$ iff $P_{l}(A)>0$, which we know to be true from the first order conditions of the firm, specifically Equation (16).

Lemma 2. The three constraints of the firm problem only bind concurrently. Unconstrained and constrained firms coexist in equilibrium if:

$$
\bar{A} \in(0, I)
$$

where:

$$
\bar{A}=I-\frac{r-\frac{\delta R}{1+i_{2}}}{1+i_{1}}-\frac{F\left(\frac{\delta R}{1+i_{2}}\right)}{\left(1+i_{1}\right)\left(1+i_{2}\right)} \delta\left(R-\frac{B}{\Delta P}\right)
$$

Proof of Lemma 2. The three constraints of the firms' problem are:

$$
\begin{equation*}
r-\rho^{*}(A) \geq d_{s}(A) \tag{LL}
\end{equation*}
$$

$$
\begin{gather*}
R-\frac{B}{\Delta P} \geq d_{l}(A)  \tag{IC}\\
\frac{d_{s}(A)}{1+i_{1}}+P_{l}(A) d_{l}(A) \geq I-A \tag{IR}
\end{gather*}
$$

First, we will show that these three constraints will only bind simultaneously, such that, notationally, we only need one Lagrange-multiplier $\lambda(A)$. If there exists any one firm that is unconstrained in equilibrium, i.e., if there exists $A^{\prime}$ such that $0 \leq A^{\prime} \leq I$ and $\lambda_{2}\left(A^{\prime}\right)=\lambda_{3}\left(A^{\prime}\right)=0$, then (16) implies that

$$
P_{l}\left(A^{\prime}\right)=\frac{1+\lambda_{2}\left(A^{\prime}\right)}{1+\lambda_{3}\left(A^{\prime}\right)} \frac{\delta F\left(\rho^{*}\left(A^{\prime}\right)\right)}{\left(1+i_{1}\right)\left(1+i_{2}\right)}=\frac{\delta F\left(\rho^{*}\left(A^{\prime}\right)\right)}{\left(1+i_{1}\right)\left(1+i_{2}\right)} .
$$

(16) must hold for all possible endowments $A$ and thus

$$
\frac{P_{l}(A)}{\delta F\left(\rho^{*}(A)\right)}=\frac{1+\lambda_{2}(A)}{1+\lambda_{3}(A)} \frac{1}{\left(1+i_{1}\right)\left(1+i_{2}\right)}
$$

and since the LHS is identical for all $A$, it follows that

$$
\frac{1+\lambda_{2}(A)}{1+\lambda_{3}(A)}=1 \quad \Longleftrightarrow \quad \lambda_{2}(A)=\lambda_{3}(A) \quad \forall A
$$

and from (15) follows that

$$
\lambda(A) \equiv \lambda_{1}(A)=\lambda_{2}(A)=\lambda_{3}(A) \forall A \quad \forall A
$$

Now, we can find the threshold endowment $\bar{A}$, which is is the lowest endowment at which $\lambda(\bar{A})=0$. At the critical value $\bar{A}$ the constraints bind exactly, for a firm that chooses the optimal cutoff value $\rho^{*}=\delta R /\left(1+i_{2}\right)$. We find it by plugging the LL and the IC constraints into the IR constraint. From (16), the price of LT debt takes the form

$$
P_{l}(A)=\frac{\delta F\left(\frac{\delta R}{1+i_{2}}\right)}{\left(1+i_{1}\right)\left(1+i_{2}\right)}
$$

and we get that the threshold endowment must be:

$$
\bar{A}=I-\frac{r-\frac{\delta R}{1+i_{2}}}{1+i_{1}}-\frac{\delta F\left(\frac{\delta R}{1+i_{2}}\right)}{\left(1+i_{1}\right)\left(1+i_{2}\right)}\left(R-\frac{B}{\Delta P}\right) .
$$

Both types, unconstrained and constrained firms, exist concurrently if

$$
\bar{A} \in(0, I) .
$$

Lemma 3. Assuming $\frac{\chi \delta}{1+i_{2}}\left(R-\frac{B}{\Delta p}\right)<1$, a constrained firm chooses a continuation value $\rho^{*}(A) \in\left[0, \frac{\delta R}{1+i_{2}}\right)$. It also follows that:

1. $\frac{\partial \rho^{*}(A)}{\partial A}>0$ if $A \leq \bar{A}$ and $\frac{\partial \rho^{*}(A)}{\partial A}=0$ if $A>\bar{A}$
2. $\frac{\partial \rho^{*}(A)}{\partial i_{1}}<0$ if $A \leq \bar{A}$ and $\frac{\partial \rho^{*}(A)}{\partial i_{1}}=0$ if $A>\bar{A}$

Proof of Lemma 3. Combining (16) and (17) we see that:

$$
\rho^{*}(A)=\frac{\delta R}{1+i_{2}}-\lambda(A)\left[\frac{1}{f\left(\rho^{*}\right)}-\frac{\delta d_{l}(A)}{1+i_{2}}\right] .
$$

An increase in $\rho^{*}(A)$ has two effects for constrained firms. On the one hand, it tightens the LL constraint, such that less short-term debt can be issued. On the other hand, a higher probability of continuation yields a higher price. We will show that our assumption $\chi \delta(1+$ $\left.i_{2}\right)^{-1}(R-B / \Delta p)<1$ assures that the former effect always dominates the latter and that all statements in this lemma follow from this fact. Note that for constrained firms, for which $\lambda(A)>0$, it is true that:

$$
\rho^{*}(A)<\frac{\delta R}{1+i_{2}} \quad \Longleftrightarrow \quad f\left(\rho^{*}\right) \frac{\delta d_{l}(A)}{1+i_{2}}<1 .
$$

The inequality $f\left(\rho^{*}\right) \frac{\delta d_{l}(A)}{1+i_{2}}<1$ holds due to the assumption that $\frac{\chi \delta}{1+i_{2}}\left(R-\frac{B}{\Delta p}\right)<1$, as

$$
f\left(\rho^{*}\right) \frac{\delta d_{l}(A)}{1+i_{2}}<f(0) \frac{\delta\left(R-\frac{B}{\Delta p}\right)}{1+i_{2}}<\frac{\chi \delta}{1+i_{2}}\left(R-\frac{B}{\Delta p}\right)<1
$$

Therefore,

$$
\rho^{*}(A)<\frac{\delta R}{1+i_{2}} .
$$

Statements 1. and 2. in the Lemma follow immediately from the IR constraint. As the unconstrained firms, those with $A>\bar{A}$, always choose $\rho^{*}(A)=\delta R /\left(1+i_{2}\right)$, there will be no effect of a change in either $A$ nor $i_{1}$ on their choice of $\rho^{*}(A)$. For the constrained firms we can derive the change in $\rho^{*}(A)$ from the constraints.

Recall that for the constrained firm the following must hold:

$$
\frac{r-\rho^{*}(A)}{1+i_{1}}+\delta F\left(\rho^{*}(A)\right) \frac{1-\gamma(1-\delta)\left(\int_{0}^{I} F\left(\rho^{*}(A)\right) d_{l}(A) d A-g\right)}{\left(1+i_{1}\right)\left[\alpha\left(1+i_{1}\right)+(1-\alpha)\left(1+i_{2}\right)\right]}\left(R-\frac{B}{\Delta P}\right)=I-A .
$$

By plugging in the equilibrium price

$$
P_{l}(A)=\frac{\delta F\left(\rho^{*}(A)\right)}{\left(1+i_{1}\right)\left(1+i_{2}\right)}
$$

this becomes:

$$
\frac{r-\rho^{*}(A)}{1+i_{1}}+\frac{\delta F\left(\rho^{*}(A)\right)}{\left(1+i_{1}\right)\left(1+i_{2}\right)}\left(R-\frac{B}{\Delta P}\right)=I-A
$$

We apply the Implicit Function Theorem to the equality

$$
l: \frac{r-\rho^{*}(A)}{1+i_{1}}+\frac{\delta F\left(\rho^{*}(A)\right)}{\left(1+i_{1}\right)\left(1+i_{2}\right)}\left(R-\frac{B}{\Delta P}\right)-I+A=0 .
$$

For $A<\bar{A}$ we get

$$
\frac{\partial \rho^{*}(A)}{\partial i_{1}}=-\frac{\frac{\partial l}{\partial i_{1}}}{\frac{\partial l}{\partial \rho^{*}(A)}}<0
$$

since

$$
\frac{\partial l}{\partial i_{1}}<0
$$

and

$$
\frac{\partial l}{\partial \rho^{*}(A)}=-\frac{1}{1+i_{1}}+\frac{\delta f\left(\rho^{*}(A)\right)}{\left(1+i_{1}\right)\left(1+i_{2}\right)}\left(R-\frac{B}{\Delta P}\right)<0
$$

as

$$
f\left(\rho^{*}\right) \frac{\delta\left(R-\frac{B}{\Delta P}\right)}{1+i_{2}}<1
$$

Moreover,

$$
\frac{\partial \rho^{*}(A)}{\partial A}=-\frac{1}{\frac{\partial l}{\partial \rho^{*}(A)}}>0 .
$$

Proof of Proposition 1. We start from the first result. As the unconstrained firms, those with $A>\bar{A}$, always choose $\rho^{*}(A)=\delta R /\left(1+i_{2}\right)$, there will be no effect of a change in $i_{1}$ on their choice. For the second part, it amounts to show that $\exists \phi>0$ such that if $\varkappa>\phi$,
unconstrained firms increase their LT debt issuance, i.e.:

$$
\frac{\partial d_{l}(A)}{\partial i_{1}}<0 \text { for } \mathrm{A} \in(\bar{A}, \mathrm{I}] .
$$

Recall that the total expected revenue from long-term $\operatorname{debt} \int_{0}^{I} F\left(\rho^{*}(A)\right) d_{l}(A) d A$ is pinned down by the market clearing equation:

$$
\frac{1-\gamma(1-\delta)\left(\int_{0}^{I} F\left(\rho^{*}(A)\right) d_{l}(A) d A-g\right)}{\left[\alpha\left(1+i_{1}\right)+(1-\alpha)\left(1+i_{2}\right)\right]}=\frac{1}{1+i_{2}}
$$

We can reformulate this equation as

$$
\int_{0}^{I} F\left(\rho^{*}(A)\right) d_{l}(A) d A=\frac{\varkappa\left(i_{2}-i_{1}\right)}{\left(1+i_{2}\right)(1-\delta)}+g .
$$

We can then split the left-hand side into the unconstrained and constrained components of LT debt

$$
\int_{0}^{\bar{A}} F\left(\rho^{*}(A)\right) d_{l}(A) d A+\int_{\bar{A}}^{I} F\left(\rho^{*}(A)\right) d_{l}(A) d A=\frac{\varkappa\left(i_{2}-i_{1}\right)}{\left(1+i_{2}\right)(1-\delta)}+g
$$

and get

$$
\int_{\bar{A}}^{I} d_{l}(A) d A=\frac{1}{F\left(\frac{\delta R}{1+i_{2}}\right)}\left[g+\frac{\varkappa\left(i_{2}-i_{1}\right)}{\left(1+i_{2}\right)(1-\delta)}-\left(R-\frac{B}{\Delta P}\right) \int_{0}^{\bar{A}} F\left(\rho^{*}(A)\right) d A\right] .
$$

Thus,

$$
\frac{\partial \int_{\bar{A}}^{I} d_{l}(A) d A}{\partial i_{1}}=\frac{1}{F\left(\frac{\delta R}{1+i_{2}}\right)}\left[\frac{-\varkappa}{\left(1+i_{2}\right)(1-\delta)}-\left(R-\frac{B}{\Delta P}\right) \frac{\partial \int_{0}^{\bar{A}} F\left(\rho^{*}(A)\right) d A}{\partial i_{1}}\right]
$$

which is negative, as the data suggests, if

$$
\frac{-\varkappa}{\left(1+i_{2}\right)(1-\delta)}-\left(R-\frac{B}{\Delta P}\right) \frac{\partial \int_{0}^{\bar{A}} F\left(\rho^{*}(A)\right) d A}{\partial i_{1}}<0
$$

Plugging in the explicit expressions into this inequality, we have

$$
\frac{\varkappa}{\left(1+i_{2}\right)(1-\delta)}>\left(R-\frac{B}{\Delta P}\right) \int_{0}^{\bar{A}} f\left(\rho^{*}(A)\right) \frac{\frac{r-\rho^{*}(A)}{1+i_{1}}+\frac{\delta F\left(\rho^{*}(A)\right)}{\left(1+i_{1}\right)\left(1+i_{2}\right)}\left(R-\frac{B}{\Delta P}\right)}{1-\frac{\delta f\left(\rho^{*}(A)\right)}{1+i_{2}}\left(R-\frac{B}{\Delta P}\right)} d A .
$$

The right-hand side of this inequality is maximized by setting $\rho^{*}(A)=0$, so as low as possible, in this case $f(0)=\chi$ and $F(0)=0$. Thus, in this way, the right-hand side simplifies to

$$
\frac{\varkappa}{\left(1+i_{2}\right)(1-\delta)}>\bar{A}\left(R-\frac{B}{\Delta P}\right) \chi \frac{\frac{r}{1+i_{1}}}{1-\frac{\delta \chi}{1+i_{2}}\left(R-\frac{B}{\Delta P}\right)} .
$$

Inspecting this result, we can see that if $\varkappa$ is large enough, namely larger than $\phi$

$$
\phi=\frac{\left(1+i_{2}\right) \chi(1-\delta) \bar{A}\left(R-\frac{B}{\Delta P}\right) \frac{r}{1+i_{1}}}{1-\frac{\delta \chi}{\left(1+i_{2}\right)}\left(R-\frac{B}{\Delta P}\right)}
$$

then we have a sufficient condition for

$$
\frac{\partial \int_{A}^{I} d_{l}(A) d A}{\partial i_{1}}<0
$$

The second part of the proposition requires us to show that the aggregate change of LT debt of the unconstrained firms generalizes to individual firm behaviour, namely

$$
\frac{\partial \int_{A}^{I} d_{l}(A) d A}{\partial i_{1}}<0 \Longrightarrow \frac{\partial d_{l}(A)}{\partial i_{1}}<0 \text { for } \mathrm{A} \in(\bar{A}, \mathrm{I}]
$$

Recall the assumption that unconstrained firms choose at least the minimum LT debt that allows them to equalize the highest LT debt share of the constrained firms, which is the LT debt share of the firm with endowment $\bar{A}$. This minimum component can be expressed as:

$$
d_{l}^{\min }(A)=\frac{\left(1+i_{1}\right)(I-A)}{\frac{1-\kappa}{\kappa}+\frac{\delta F\left(\frac{\delta R}{1+i_{2}}\right)}{1+i_{2}}}
$$

Here,

$$
\kappa=\frac{R-\frac{B}{\Delta p}}{R-\frac{B}{\Delta p}+r-\frac{\delta R}{1+i_{2}}}
$$

is the LT debt share of the firm with endowment $\bar{A}$. Additionally, we impose that for any unconstrained firm $d_{l}^{\min }(A)<d_{l}(A)<(R-B / \Delta p)$, and that any change in the aggregate LT debt of unconstrained firms is distributed among all unconstrained firms as a change in their LT debt, and that this change, relative to the aggregate change, for any subset of
unconstrained firms that has non-zero measure, is larger than zero.
Now, consider an infinitesimal increase in $i_{1}$ : it will increase the minimum amount of LT debt needed to match the highest LT debt ratio of constrained firms. Concretely:

$$
\frac{\partial d_{l}^{\min }(A)}{\partial i_{1}}=\frac{I-A}{\frac{1-\kappa}{\kappa}+\frac{\delta F\left(\frac{\delta R}{1+i_{2}}\right)}{1+i_{2}}}
$$

However, as for all $A \in(\bar{A}, I]$ the choice of LT debt before the increase was strictly higher than the minimum LT debt due to our assumption, the infinitesimal change in the minimum LT debt will not make it surpass the previous amount. Instead, as the aggregate LT debt for the unconstrained firms must decrease, and this decrease is distributed among all unconstrained firms, we find that $\partial d_{l}(A) / \partial i_{1}<0$ for $A \in(\bar{A}, I]$. For an infinitesimal decrease in $i_{1}$, the aggregate long-term debt increases while $d_{l}^{\min }(A)$ decreases, thus $d_{l}(A)$ must increase for all firms with $A \in(\bar{A}, I]$.
Moving on, from

$$
\frac{\partial \int_{A}^{I} d_{l}(A) d A}{\partial i_{1}}=\frac{1}{F\left(\frac{R}{1+i_{2}}\right)}\left[\frac{-\varkappa}{\left(1+i_{2}\right)(1-\delta)}-\left(R-\frac{B}{\Delta P}\right) \frac{\partial \int_{0}^{\bar{A}} F\left(\rho^{*}(A)\right) d A}{\partial i_{1}}\right]
$$

we can see that if $\varkappa \rightarrow \infty$ then

$$
\frac{\partial \int_{A}^{I} d_{l}(A) d A}{\partial i_{1}} \rightarrow-\infty .
$$

Due to the assumptions that a change in the aggregate must be shared across firms and furthermore each share cannot be trivially small, by the same logic as above

$$
\frac{\partial \int_{\bar{A}}^{I} d_{l}(A) d A}{\partial i_{1}} \rightarrow-\infty \Longrightarrow \frac{\partial d_{l}(A)}{\partial i_{1}} \rightarrow-\infty \text { for } \mathrm{A} \in(\bar{A}, \mathrm{I}] .
$$

In order to prove the second part, namely that

$$
\frac{\partial \frac{\partial d_{s}(A)}{\partial i_{1}}}{\partial \varkappa}=0 \quad \forall A
$$

recall that for the constrained firms, $\rho^{*}(A)$ is pinned down by the constraints, in which $\varkappa$ plays no role, as it does not affect the price. A change in $\varkappa$ thus has no effect on the choice of $\rho^{*}(A)$ for firms with $A<\bar{A}$. This concludes the proof of result 2 .

As shown in Lemma 3, constrained firms, those with $A \leq \bar{A}$, increase $\rho^{*}(A)$ in $i_{1}$. Their

LT debt is determined by the IC constraint to be $d_{l}(A)=(R-B / \Delta p)$. This proves result 3.

## IB.4. Extensions

Effect of a Policy Rate Change with Risk-neutral and Rational Investors
To showcase the consequences of our departure from a model with only rational and riskneutral investors, we analyze a baseline specification, in which we assume that investors are risk-neutral and rational: $\gamma=\alpha=0$. In this case, investor demand for bonds is horizontal, which means that they are willing to hold any amount of LT debt at price

$$
P_{l}(A)=\frac{\delta F\left(\rho^{*}(A)\right)}{\left(1+i_{1}\right)\left(1+i_{2}\right)} .
$$

This is also the price at which a firm with endowment $A$ would inelastically sell LT debt. Thus, the total amount of LT debt is not pinned down by market clearing. The requirements for constrained and unconstrained firms to exists are the same as derived above, and it is still true in this benchmark case that for constrained firms

$$
\frac{\partial \rho^{*}(A)}{\partial i_{1}}<0
$$

However, because $\int_{0}^{I} P_{l}(A) d_{l}(A) d A$ is not pinned down by a downward sloping investor demand curve, we cannot say whether or how unconstrained firms adjust their maturity structure. Thus, we have that constrained firms take on less short-term debt when the the monetary authority eases, while unconstrained firms have no incentive to change their maturity structure. This is counterfactual, as we see in the data that unconstrained firms should lengthen their maturity structure and do so more than constrained firms.

## Effect of a Policy Rate Change without Yield-Seeking Investors

The absence of reach-for-yield motives represents a special case of the model, in which $\alpha=$ 0. In such a setting, rational investors have no incentive to hold LT debt, because it is risky
and firms are just willing to price debt in a risk-neutral fashion. ${ }^{32}$ In practical terms, the inverse demand function for LT bonds equals:

$$
P_{l}(A)=\delta F\left(\rho^{*}(A)\right) \frac{1-\gamma(1-\delta)\left(\int_{0}^{I} F\left(\rho^{*}(A)\right) d_{l}(A) d A-g\right)}{\left(1+i_{1}\right)\left(1+i_{2}\right)} .
$$

Market clearing implies:

$$
\frac{1-\gamma(1-\delta)\left(\int_{0}^{I} F\left(\rho^{*}(A)\right) d_{l}(A) d A-g\right)}{\left(1+i_{1}\right)\left(1+i_{2}\right)}=\frac{1}{\left(1+i_{1}\right)\left(1+i_{2}\right)},
$$

or

$$
\begin{equation*}
\int_{0}^{I} P_{l}(A) d_{l}(A) d A=\int_{0}^{\bar{A}} F\left(\rho^{*}(A)\right) d_{l}(A) d A+\int_{\bar{A}}^{I} F\left(\rho^{*}(A)\right) d_{l}(A) d A=g . \tag{18}
\end{equation*}
$$

That is, the LT bonds issued by constrained and unconstrained firms have to net out. In this case, in reaction to a descent in the interest rate $i_{1}$, by Lemma 3, constrained firms increase their continuation value $\rho^{*}$. The right hand side of (18) is constant, and thus unconstrained firms must decrease their LT debt, therefore shortening debt maturity. This result is in contrast with the cross-sectional empirical evidence.

## Rollover Risk

In this extension, we allow firms to choose between financing investment with short-term debt from $t=0$ to $t=1$, LT debt, and rolling over short-term debt at at $t=1$. The following timing is applied in $t=1$ : first, a liquidity shock is realized, then a potential rollover crisis is realized and lastly short-term payout $r$ is paid out. This implies that, as in the base model, there exists a maximum $\rho^{*}$ that a firm will be able to cover with available liquidity. In the following, we show that for constrained firms, $\rho^{*}$ is larger if they roll over some of their short-term debt. Debt rollover has the advantage that it is conditional on the liquidity shock realizing and thus cheaper ex-ante, however, there is a risk that the market will not supply funding at $t=1$. Let the probability of such an event, a "rollover crisis", be $1-\varepsilon$.

If a firm is in the range in which a rollover crisis would lead to default, despite $\rho<\rho^{*}$,
${ }^{32}$ This can be inferred from Equation (16). For unconstrained companies, $\lambda_{2}=\lambda_{3}=0$, such that the resulting price just discounts aggregate and idiosyncratic risk, without offering any risk premium.
then the firm defaults with probability $1-\varepsilon$. As before, we assume that default is costly and thus firm value is reduced to zero in case of default.

The problem of the firm looks similar to before, except that, if the firm relies on rollover debt the prices $P_{s}$ and $P_{l}$, as well as the firm value, will be scaled down by $\varepsilon$ (the probability of not facing a rollover crisis). Thus, rolling over will be avoided by unconstrained firms, but it will be used by some constrained firms to extend their financial capacity. Define

$$
\bar{\varepsilon}=\left\{\begin{array}{ll}
\varepsilon, & \text { for } \rho^{*}(A)+d_{s}(A) \geq r \\
1, & \text { for } \rho^{*}(A)+d_{s}(A)<r
\end{array}\right\}
$$

as the threshold such that firms decide to roll over or not, and $d_{s}^{\prime}$ the amount of debt that the firm schedules for rollover in $t=1$. The problem of firm $A$ at time 0 now reads:
$\max _{\rho^{*}(A), d_{s}(A), d_{l}(A), d_{s}^{\prime}(A)} \bar{\varepsilon}\left(\frac{r}{1+i_{1}}-\frac{\int_{0}^{\rho^{*}(A)} \rho f(\rho) d \rho}{1+i_{1}}+\frac{\delta F\left(\rho^{*}(A)\right)}{\left(1+i_{1}\right)\left(1+i_{2}\right)} R+\left(P_{l}(A)-\frac{\delta F\left(\rho^{*}(A)\right)}{\left(1+i_{1}\right)\left(1+i_{2}\right)}\right) d_{l}(A)\right)$
subject to

$$
\begin{aligned}
d_{s}^{\prime}+d_{l} & \leq R-\frac{B}{\Delta p^{\prime}} \\
\rho^{*}(A)+d_{s} & \leq r+\frac{d_{s}^{\prime}}{1+i_{2}}, \\
\frac{\bar{\varepsilon} d_{s}(A)}{1+i_{1}}+P_{l}(A) d_{l}(A) & \geq I-A,
\end{aligned}
$$

and

$$
P_{l}(A)=\frac{\bar{\varepsilon} \delta F\left(\rho^{*}(A)\right)}{\left(1+i_{1}\right)\left(1+i_{2}\right)}
$$

First note that an unconstrained firm chooses $d_{s}^{\prime}=0$, as otherwise it exposes itself to rollover risk and it already attained the optimum. These firms use their financial capacity to avoid rollover risk. This implies that the previously derived $\bar{A}$ threshold still governs the distinction between constrained and unconstrained firms.

Within the set of constrained firms, firms have to choose whether they want to roll over debt or not. The trade-off is that rollover debt has a positive effect on financing capacity, as
exemplified if we combine constraints and objective function to find

$$
\bar{\varepsilon}\left\{\frac{r-\rho^{*}(A)}{1+i_{1}}+\frac{\delta F\left(\rho^{*}(A)\right)}{\left(1+i_{1}\right)\left(1+i_{2}\right)}\left(R-\frac{B}{\Delta p}\right)+\left(1-\delta F\left(\rho^{*}(A)\right)\right) \frac{d_{s}^{\prime}(A)}{\left(1+i_{1}\right)\left(1+i_{2}\right)}\right\}=I-A .
$$

Conditional on $\bar{\varepsilon}$, rolling over increases the financing capacity and thus lets firms choose a higher $\rho^{*}$ (given that the LHS is decreasing in $\rho^{*}$, which is true under our assumption on the distribution of $\rho$ ). Therefore, if a firm chooses to roll over debt, it will choose $d_{l}=0$ and $d_{s}^{\prime}=\left(R-\frac{B}{\Delta p}\right)$.

A constrained firm faces the trade-off between rolling over and taking on LT debt on the external margin as well. Note that firms that are close to the optimal $\rho^{*}$ would opt for LT debt, as they would trade off a close to optimal $\rho^{*}$ by exposing themselves to rollover risk, which induces a discrete increase in the probability of a rollover crisis. In order to see this, we can compare, which of the two cases maximizes the initial financing and the objective function while holding $\rho^{*}$ constant. Define the financing capacity with rollover as

$$
\mathcal{D}^{\text {rollover }} \equiv \varepsilon\left\{\frac{r-\rho^{*}(A)}{1+i_{1}}+\frac{1}{\left(1+i_{1}\right)\left(1+i_{2}\right)}\left(R-\frac{B}{\Delta p}\right)\right\}
$$

and without rollover as

$$
\mathcal{D}^{\text {no-rollover }} \equiv\left\{\frac{r-\rho^{*}(A)}{1+i_{1}}+\frac{\delta F\left(\rho^{*}(A)\right)}{\left(1+i_{1}\right)\left(1+i_{2}\right)}\left(R-\frac{B}{\Delta p}\right)\right\} .
$$

The price of long-term bonds is higher than for rolled-over short-term bonds, as short-term debt includes the option of not getting renewed. For LT debt the price is $\delta F\left(\rho^{*}(A)\right) /[(1+$ $\left.\left.i_{1}\right)\left(1+i_{2}\right)\right]$, for rolled-over short-term debt it is $1 /\left[\left(1+i_{1}\right)\left(1+i_{2}\right)\right]$. This implies that rolledover short-term debt is cheaper for the issuing firm and thus increases financial capacity. Taking the difference we get

$$
\Delta \mathcal{D} \equiv \mathcal{D}^{\text {no-rollover }}-\mathcal{D}^{\text {rollover }} \equiv(1-\varepsilon) \frac{r-\rho^{*}(A)}{1+i_{1}}+\frac{\delta F\left(\rho^{*}(A)\right)-\varepsilon}{\left(1+i_{1}\right)\left(1+i_{2}\right)}\left(R-\frac{B}{\Delta p}\right) .
$$

Thus, choosing to roll over increases financing capacity if: $\varepsilon$ is large (specifically, at least as large as $\delta F\left(\rho^{*}(A)\right)$ ) and if $\rho^{*}$ is small (which is the case for low $A$ firms). We now get three types of firms, instead of two. The smallest firms increase their financial capacity by
rolling over debt, while larger constrained firms and unconstrained firms do not. Even these smallest constrained firms will only rely on rollover debt if the increase in $\rho^{*}$ is worth their decreased firm value, due to rollover risk.

For our general prediction that large firms issue more LT debt if interest rates are low, the explicit introduction of rollover risk into the model does not alter the key insights, as these rely on the term-premium channel. Very small firms now do not hold any LT debt, while small firms remain constrained in their LT debt due to moral hazard and can only issue more short-term debt if rates are low. Thus only large firms are reacting to excess demand of LT debt by issuing more LT debt and arbitraging the term premium in response to a monetary easing.

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    ${ }^{\dagger}$ Banca d’Italia, andrea.fabiani@bancaditalia.it
    $\ddagger$ University of Pennsylvania, luigif@sas.upenn.edu
    §University of Bonn, janko.heineken@uni-bonn.de

[^1]:    ${ }^{1}$ Throughout the paper, in line with Figure 1, we focus on a sample period starting in 1990, due to the unavailability of high-frequency monetary policy surprises based on Fed Funds Futures before 1990 (Gürkaynak et al. 2005).

[^2]:    ${ }^{2}$ For robustness purposes, we also gather data on holdings of debt securities by corporate bond mutual funds, representing the second largest class of investors in the corporate bond market. In particular, we exploit CRSP Survivor Bias Free mutual fund data.

[^3]:    ${ }^{3}$ Recent papers, such as Ottonello \& Winberry (2020), Kroen et al. (2021) and Ozdagli (2018), provide empirical evidence and mechanisms which, as in our case, are consistent with larger benefits of low monetary policy rates for unconstrained firms. However, those studies do not explicitly model nor explore empirically the implications for debt maturity.

[^4]:    ${ }^{4}$ We show that in our sample monetary policy surprises trigger consistent variations in both the term spread and the term premium.

[^5]:    ${ }^{5}$ Consistently, in unreported regressions, we do not find evidence of larger investment by large firms in reaction to an expansionary monetary policy surprise.

[^6]:    ${ }^{6}$ Andreolli (2021) conditions monetary policy transmission to the macro-economy on sovereign debt maturity.

[^7]:    ${ }^{7}$ Hong et al. (2017) study the reaction of the demand for long-vs-short bonds to inflation shocks and discuss the contribution of reach-for-yield motives. Chen \& Choi (2023) investigate how reach-foryield motives influence the cross-section of bond returns.

[^8]:    ${ }^{8}$ We match firm-level balance sheets from Compustat with bond-issuance data from Mergent FISD through the 6-digit CUSIP. Since CUSIPs generally vary over time but Compustat only retains the latest one, we follow a 2-step procedure (Jungherr et al. 2022). We first link Compustat with CRSP (via a common unique identifier). CRSP returns both the current and historical CUSIPs of a given firm, that we use for linking Compustat and Mergent FISD. In 25 cases, multiple companies in Compustat share the same 6-digit CUSIP, referring to different subsidiaries of the same group. In these cases, we retain the largest company in Compustat among the ones with the same 6-digit CUSIP in an effort to identify the mother company. Excluding all such companies would not affect the results.
    ${ }^{9}$ We match Compustat with syndicated loan data from Dealscan following Chava \& Roberts (2008).

[^9]:    ${ }^{10}$ For both loans and bonds, whenever there are multiple issuances by a firm in a given year-quarter (which is quite rare), we compute the firm-level spread as the weighted average across such issuances, with weights given the bond or loan volume.
    ${ }^{11}$ For a similar approach, see Becker \& Ivashina (2015).

[^10]:    ${ }^{12}$ Results are robust using the simple quarterly variation of the EFFR, but our focus rests on exogenous monetary policy interest rate surprises.
    ${ }^{13}$ From a formal perspective, Figure 6 pins down the absolute variation in the LT debt share. Nonetheless, we also refer to Figure 5 as it estimates precisely the relative adjustment of large companies, and the baseline effect on smaller firms can be placed at 0 .

[^11]:    ${ }^{14}$ Section IB. 4 of the Internet Appendix also shows the robustness of our results to the introduction of rollover risk as an additional friction.

[^12]:    ${ }^{15}$ From now on we will index the choice variables $d_{s}, d_{l}$, and $\rho^{*}$ by the endowment $A$.

[^13]:    ${ }^{16}$ We assume that the mass of firms is $I$, such that the density of each firm type $A$ is 1 and can be omitted from notations.
    ${ }^{17}$ The fact that these investors are discounting using an incorrect rate generates the yield-seeking behavior. This modeling choice can be justified by agency or accounting considerations that lead investors to worry about short-term measures of reported performance.
    ${ }^{18}$ Greenwood et al. (2010) make a similar assumption and describe such investors as pension funds, life insurance companies or any institution with an inelastic demand for long-term assets.

[^14]:    ${ }^{19}$ For a formal definition of competitive equilibrium in the context of our model and the detailed derivation of all provided results, we refer the reader to Internet Appendix section IB.2.
    ${ }^{20}$ We refer to Lemma 2 in the Internet Appendix section IB. 3 for the detailed conditions regarding the co-existence of both firms in equilibrium.
    ${ }^{21}$ We assume that unconstrained firms choose a combination of short-term and LT debt consistent with the highest LT debt ratio among constrained firms; by doing so, we match the stylized empirical fact in Panel A of Figure 3 that large companies issue relatively more LT debt. Detailed assumptions can be found in the Internet Appendix section IB.3, in the proof of Proposition 1. This assumption is possible thanks to an appropriately large preferred-habitat investor demand, $g$, which generates a relatively large excess demand for LT debt, filled by unconstrained companies.

[^15]:    ${ }^{22}$ To show the effect of adding yield-seeking investors to the model, Appendix IB. 4 describes the effects of a rate change in two benchmark models, which both exclude yield-seeking investors (i.e., reach-for-yield motives) and one excluding risk-aversion. We show that, without our key assumptions, these two models yield predictions counterfactual to the data.

[^16]:    ${ }^{23}$ Put differently, plotting the estimated coefficients $\hat{\beta}_{2, h}$ does not return a standard IRF, which would require, at horizon $H$, summing up coefficients $\sum_{h=0,1, . . H} \hat{\beta}_{3, h}$.
    ${ }^{24}$ Relative to previous models, we report some macro controls (term spread, corporate spread and share of Treasuries with maturity above 20 years) in levels rather than in first differences.

[^17]:    ${ }^{25}$ This does not exclude real effects for large firms, which could be indirectly transmitted through the change in their liability structure.

[^18]:    ${ }^{26}$ See Table A1 for the detailed definitions of variables used in Equations 8 and 9 .

[^19]:    ${ }^{27}$ Ideally, we would look at the maturity of corporate bonds, rather than at the maturity of all debt securities held in the portfolio of CBMF. However, unfortunately, this information is not available in CRSP Survivor Bias Free data.
    ${ }^{28}$ We provide details about the classification in Section IA.2.1 of the Internet Appendix.

[^20]:    ${ }^{29}$ For selecting our sample of funds and classifying them as HY or IG, we follow Choi \& Kronlund (2018). In practical terms, we consider the sample of CBMF by limiting our focus to funds falling in the following CRSP style categories: I, ICQH, ICQM, ICQY, ICDI, ICDS, IC. Moreover, we label Investment Grade (IG) funds as those with a Lipper style code of either A, BBB, IID, SII, SID, or USO; High Yield (HY) funds are those coded HY, GB, FLX, MSI, or SFI.

[^21]:    ${ }^{30}$ There are various approaches to model reach-for-yield behavior, such as Acharya \& Naqvi (2019), Lu et al. (2019), Campbell \& Sigalov (2022). We follow the approach of Hanson \& Stein (2015) who model reach for yield as a subset of agents using the current interest rate to discount future income, instead of the path of expected future interest rates. We take this path as their modeling approach is one that considers yield-seeking explicitly with regard to the asset's maturity, which matches our main empirical results.

[^22]:    ${ }^{31}$ The inclusion of a large enough $g$ ensures that, under any circumstances, all firms borrow a positive amount of LT debt.

