Long Rates, Life Insurers, and Credit Spreads*

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Abstract

Post-2008, corporate bond credit spreads decline when long-term interest rates increase, particularly for lower-rated bonds. This is true unconditionally but also conditional on monetary policy announcements. In the cross-section, this negative co-movement between long rates and credit spreads is more pronounced for bonds predominantly held by life insurers. I develop a quantitative framework that rationalizes these findings. In the model, life insurers with long-duration liabilities face duration mismatch and therefore realize equity gains when long rates increase. As a result, their effective risk aversion declines, driving down equilibrium credit spreads. The model explains the majority of the empirical credit spread responses to long rates. The model also shows that life insurers' duration mismatch can dampen or even reverse the transmission of unconventional monetary policy to bond yields and issuance.

JEL Codes: G11, G12, G22, E44, E52
Keywords: Corporate Bonds, Credit Spreads, Long-term Interest Rates, Life Insurance Companies, Duration Mismatch, Monetary Policy

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1 Introduction

The US corporate bond market has expanded dramatically in the last three decades, with its total market capitalization reaching 4.4 times GDP in 2023 Q3 from 2.9 times GDP in 1993 Q1 and surpassing the bank loan market as the primary funding source for US corporations. Corporate bond prices are also significant indicators of real investments and aggregate economic activities (Philippon, 2009; Gilchrist and Zakrajšek, 2012). This paper studies the impact of long-term interest rates on the corporate bond market and focuses on credit spreads, a vital component of corporate bond yields. In particular, I find that increases in long-term interest rates have led to declines in credit spreads since the Global Financial Crisis.

The first main contribution of this paper is new empirical evidence on the impact of long rates on credit spreads and the role of life insurers. I document a new *negative* co-movement between the long-term interest rate and corporate bond credit spreads after the Financial Crisis, both unconditionally and conditional on monetary policy announcements. I then attribute this new co-movement to the duration mismatch of life insurers, the largest investor group in the bond market. I provide causal evidence that higher ownership by life insurers leads to more negative co-movement of credit spreads with the long-term interest rate in the cross-section of corporate bonds. I further show that resulting credit spread dynamics have important real implications on bond issuance and investment.

The second main contribution of this paper is a quantitative framework grounded in the new empirical findings. I develop an intermediary asset pricing model featuring life insurers facing duration mismatch. The model quantitatively explains the empirical findings and demonstrates that this duration mismatch channel can dampen or even reverse the effects of unconventional monetary policy on bond yields and issuance.

I begin by documenting novel facts about the co-movement between the long-term interest rate and corporate bond credit spreads. Leveraging detailed microdata on corporate bond prices, I estimate the 10-year Treasury yield pass-through to corporate bonds with different credit ratings. Before the 2007-2008 Financial Crisis, the long-term interest rate and credit spreads exhibited no significant associations. Changes in the long-term interest rate affected all corporate bonds in a similar fashion. However, after the Financial Crisis, there is an economically and statistically significant negative relationship between the long-term interest rate and corporate bond credit spreads. In particular, the yields of low-credit-rating bonds decline relative to high-credit-rating bonds when long-term interest rates increase. In fact, the yields of bonds with the lowest ratings (e.g., single B or lower) even decline in absolute terms when the long-term interest rate increases.

Contrary to conventional wisdom, where high interest rates discourage risk-taking and raise risk premia, increases in the long-term interest rate lead to declines in corporate bond credit spreads, even after controlling for the yield curve and detailed bond characteristics. Moreover, the impact is more significant on bonds with lower credit ratings: when the long-term interest rate declines, the credit spreads of speculative-grade bonds increase significantly more than investment-grade bonds.

High-frequency identifications around monetary policy announcements further demonstrate that increases in long rates reduce corporate bond credit spreads. I construct high-frequency shocks to the 10-year Treasury yield using yield movements around FOMC meetings. Using a local projection method, I find that 10-year Treasury yield shocks did not cause significant changes in bond credit spreads before the Financial Crisis. After the Crisis, positive shocks to long-term Treasury yields induce large negative credit spread responses. For example, a 1% increase in the 10-year Treasury yield around FOMC meetings leads to an almost 1% reduction in the spread between single B corporate bonds and AAA corporate bonds.

Next, I trace this new negative co-movement to the duration mismatch and bond holdings of life insurers, the largest institutional investor group of US corporate bonds, which owns about 30% of all US corporate bonds. The business model of modern life insurers involves investing in fixed-income securities and issuing long-term annuity products. Before the Financial Crisis, life insurers were hedged against interest rate risk. Their equity value was largely shielded from fluctuations in the long-term interest rate, which suggests that their assets and liabilities had matching duration. However, as the economy entered a low-interest-rate environment post-2008, life insurers began to face a severe duration mismatch, and their market equity values became highly sensitive to the long-term interest rate.[1] In this time period, the market equity of the life insurance sector increased by almost 7.2% when the 10-year Treasury yield rose by 1%.

The negative co-movement between credit spreads and the long-term interest rate only exists in bonds with life insurance ownership and is more pronounced in bonds with higher life insurance ownership. Next, I provide causal evidence that higher life insurance ownership generates stronger co-movement between credit spreads and the long-term interest rate.

To sharpen the identification, I exploit a discontinuity in bond ownership structure stemming from mutual funds' investment mandates (e.g., Li and Yu, 2023). Figure 1 visualizes the composition of the US bond market. The bond market is populated by large institutional investors, where the largest investor type is life insurers, followed by mutual funds. Among bond mutual funds, many are "intermediate-term" with fund charters and mandates to invest in bonds whose maturities are less than 10 years, resulting in a discontinuity in investor composition around the 10-year maturity threshold. Bonds with maturities slightly below 10 years are significantly less likely to be held by life insurers than bonds with maturities slightly above 10 years, as the former face higher demand from mutual funds. I then find that the bonds slightly to the left of the cutoff are much less responsive to the 10-year Treasury yield.

The results point to a new channel through which the long-term interest rate affects corporate bond credit spreads — *the duration mismatch channel*. Life insurers' net worth rises following increases in the long-term interest rate. Consequently, their risk-bearing capacity is higher, and they expand their holdings of risky bonds. Furthermore, I show that the bond trading behavior of life insurers supports the duration mismatch channel. Following increases in the 10-year interest rate, life insurers respond by increasing their demand for risky bonds, manifested in the data as more future purchases of risky corporate bonds.

I then show that the credit spread dynamics also generate large real effects on corporate bond issuance and firm investment. Following increases in long-term interest rates, the issuance of investment-grade bonds declines more than the issuance of speculative-

[1] Lower interest rates increase the insurers' liability duration more than their asset duration. Together with other institutional frictions (e.g., regulatory), it can leave insurers in a persistent duration mismatch. See Section 4.1 for detailed discussions on the change in life insurers' hedging behavior.

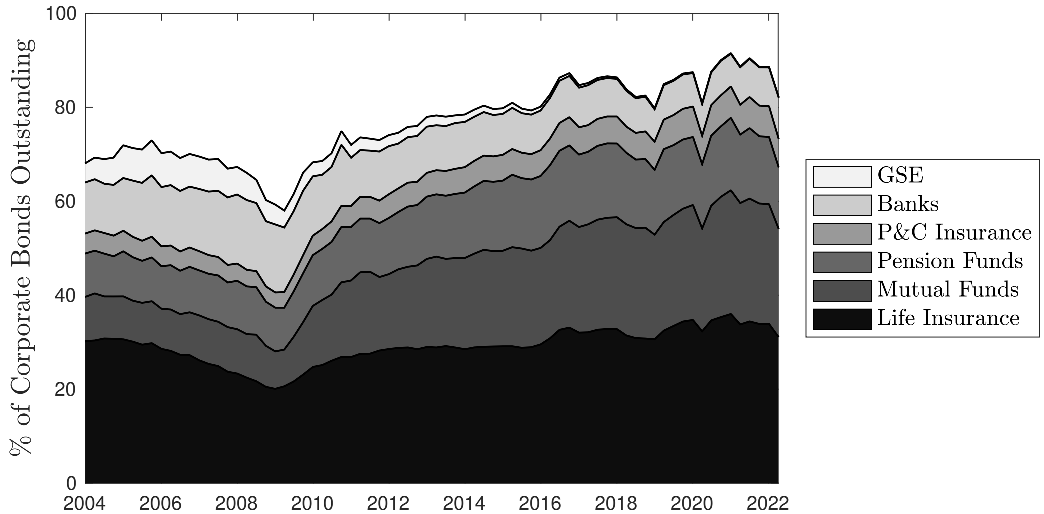

Figure 1. **Institutional Investors in the Bond Market.**

This figure plots the ownership structure of US domestic corporate bonds between 2004 and 2022. From top to bottom, the figure shows the fraction of bonds owned by government-sponsored enterprises (GSE), commercial banks, property-casualty insurance companies, pension funds, mutual funds, and life insurance companies. Foreign investors are excluded. The data is obtained from the Financial Accounts of the United States. The black area indicates that life insurers hold 20-35% of all US corporate bonds throughout the last two decades.

grade bonds. The issuance to outstanding amount ratio by investment-grade firms falls by more than 1.5% relative to speculative-grade firms in an average 6-month period when the 10-year Treasury yield is 1% higher. The finding reveals large credit supply effects from risky firms. Further, I show that life insurers' bond ownership has a significant effect on the transmission of long-term interest rates to firm investment. Firms reduce their capital expenditure when the long-term interest rate increases, but the effect is milder in firms whose bonds are held more by life insurers.

Finally, based on the new empirical evidence, I develop an intermediary asset pricing model centered around the duration mismatch channel to illustrate the impact of duration mismatch facing life insurers and quantify its contribution to the empirical co-movement between long rates and credit spreads. In the model, life insurers hold corporate bonds and Treasuries as assets and issue annuities as liabilities. Life insurers face duration mismatch because the annuities they issue have a much longer duration than their assets. Increases in the Treasury yield cause declines in the price of annuities, which reduce the value of life insurers' liabilities more than their assets and boost their net worth. Life insurers are the main investors of corporate bonds, and their risk-bearing capacity affects equilibrium credit spreads. Life insurers are more averse to bonds with lower credit ratings as they carry greater credit risk and regulatory costs. Therefore, when the long-term Treasury yield rises, life insurers become more willing to hold risky bonds, so the equilibrium credit spreads shrink.

Analytically, I prove that increases in the long-term interest rate depress corporate bond credit spreads in a simplified model. Quantitatively, the model can explain the majority of the observed co-movement between the long-term interest rate and corporate bond credit spreads when matched to the estimated duration mismatch of life insurers. The model can also account for the observations before the Financial Crisis, as the co-movement disappears once life insurers in the model are not subject to duration mismatch.

My findings suggest that policies targeting the long-term interest rate, such as Quantitative Easing and Tightening (QE and QT), have large unintended consequences in the corporate bond market. QE is a policy aimed at boosting the economy by reducing long-term interest rates. However, it may have unintended consequences by heightening credit spreads, especially for companies with higher risk profiles. Conversely, the tightening effects of QT might be dampened as higher long-term interest rates could depress credit spreads and generate favorable financial conditions for bond-issuing firms. In the model, I show that the duration mismatch channel has large real effects. Quantitatively, the credit spread responses reduce the effectiveness of long-term rate hikes on aggregate bond issuance and firm investment by more than half. The channel even reverses the transmission of long rates to speculative-grade firms, resulting in speculative-grade bond booms after positive long-term rate shocks. These results are particularly relevant today, given the increasing prevalence of unconventional monetary policies such as QE and QT.

Related Literature. My results contribute to the extensive literature on corporate bond credit spreads.[2] Following Fama and French (1993), many have modeled corporate bond yields and returns using factor models (e.g., Gebhardt, Hvidkjaer and Swami-

[2]See Huang and Shi (2021) for an overview of the literature on corporate bond returns.

nathan, 2005; Lin, Wang and Wu, 2011; Acharya, Amihud and Bharath, 2013; Jostova et al., 2013; Kelly, Palhares and Pruitt, 2023). The factor approach typically views credit risk and the term structure as unconnected orthogonal factors in determining bond yields. In this paper, I show that the pricing of the two factors is interconnected, as the level of long-term interest rates could affect the pricing of credit risk through the risk-bearing capacity of life insurers. Another strand of literature aims to explain the levels and fluctuations of corporate bond credit spreads (see, e.g., recent works by Eom, Helwege and Huang, 2004; Longstaff, Mithal and Neis, 2005; Schaefer and Strebulaev, 2008; Chen, Collin-Dufresne and Goldstein, 2009; Huang and Huang, 2012; Kuehn and Schmid, 2014; Culp, Nozawa and Veronesi, 2018; Feldhütter and Schaefer, 2018; van Binsbergen, Nozawa and Schwert, 2023). I contribute to this literature by documenting a new co-movement between credit spreads and long-term interest rates and providing an explanation centered around life insurance companies.[3]

This paper also belongs to the burgeoning literature focusing on the role of institutional investors in the corporate bond market. My work is most closely connected to Coppola (2022) and Li and Yu (2022, 2023), who, in different contexts, also show that investor composition matters for bond price dynamics. Using similar empirical frameworks, I show that life insurance ownership induces a negative co-movement between credit spreads and the long-term interest rate after the Financial Crisis. A recent strand of the literature studies the role of mutual funds in liquidity disruptions in crises (e.g., Haddad, Moreira and Muir, 2021; Falato, Goldstein and Hortaçsu, 2021; Jiang et al., 2022; Ma, Xiao and Zeng, 2022). This paper, instead, focuses on life insurers, the largest investor group in the US corporate bond market that currently holds more than 30% of the US corporate market capitalization, to show that their balance sheets are important for bond price dynamics. Recent works such as Bretscher et al. (2022) and Darmouni, Siani and Xiao (2022) use demand system approaches to analyze the equilibrium effects of institutional demand. In this paper, I depart from the logit portfolio choices used in the demand system literature to focus on life insurers' duration mismatch.

[3] Duffee (1998) finds a negative co-movement between short-term interest rates and spreads between corporate bonds and Treasuries, highlighting the role of corporate bond callability. In this paper, I instead study long-term interest rates and spreads in the cross-section of corporate bonds, focusing on the impact of life insurance companies.

The findings of my paper echo recent research on the investing behavior of life insurance companies (e.g., Koijen and Yogo, 2022; 2023; Chodorow-Reich, Ghent and Haddad, 2021; Ellul et al., 2022). Several studies, including Berends et al. (2013), Hartley, Paulson and Rosen (2016), Domanski, Shin and Sushko (2017), Ozdagli and Wang (2019), Koijen and Yogo (2022), and Huber (2022), have also shown that life insurers' interest rate risk exposure changed after the Financial Crisis. I build on this finding and argue that the duration mismatch can significantly influence bond prices since life insurers, on average, hold more than 25% of all US corporate bonds. Prior works by Ellul, Jotikasthira and Lundblad (2011), Nanda, Wu and Zhou (2019), Girardi et al. (2021), Becker, Opp and Saidi (2022), and Murray and Nikolova (2022) have established that the trades of life insurers have potentially large price impacts on corporate bond prices. I confirm that life insurers can considerably influence bond prices in the context of how corporate bond yields respond to long-term interest rate fluctuations. Consistent with Ge and Weisbach (2021) and Bhardwaj, Ge and Mukherjee (2022), I find that insurers tilt their portfolios towards safer investments when they become financially constrained.

Both my empirical and theoretical results add to the recent literature on how the corporate bond market interacts with the corporate sector and the real economy (e.g., Philippon, 2009; Gilchrist, Yankov and Zakrajšek, 2009; Greenwood, Hanson and Stein, 2010; Gilchrist and Zakrajšek, 2012; López-Salido, Stein and Zakrajšek, 2017; Coppola, 2022; Mota, 2023). In this paper, I study how long-term interest rates affect bond issuance and firm investment through their impacts on life insurers and credit spreads.

The approach of this paper connects to the literature on intermediary asset pricing, which emphasizes the role of the financial health of intermediaries on asset prices (e.g., Brunnermeier and Pedersen, 2009; He and Krishnamurthy, 2013, 2018; Brunnermeier and Sannikov, 2014; Adrian, Etula and Muir, 2014; He, Kelly and Manela, 2017; Haddad and Muir, 2021; Baron and Muir, 2022). My paper finds that the duration mismatch of life insurers, the dominant intermediaries in the US corporate bond market, significantly affects bond credit spreads. Unlike previous works where higher interest rates depress intermediaries' net worth (e.g., Gomez et al., 2021; Kekre, Lenel and Mainardi, 2023), I find that life insurers face the opposite duration mismatch and receive equity gains when long rates increase.

Additionally, this paper contributes to the literature on the impact of monetary policy and interest rates on the bond market. Departing from existing works that focus on short-term monetary policy and mutual funds (e.g., Guo, Kontonikas and Maio, 2020; Daniel, Garlappi and Xiao, 2021; Chen and Choi, 2023; Fang, 2023), I instead focus on the long-term interest rate and life insurers. I document evidence that positive shocks to long-term interest rates significantly depress corporate bond credit spreads and encourage bond issuance by risky firms.

Outline. Section 2 describes data sources for the empirical analysis. Section 3 discusses evidence on the co-movement between the long-term interest rate and corporate bond credit spreads. Section 4 investigates the role of life insurers in shaping the co-movement. Section 5 explores the real effects of the long-term interest rate and credit spreads. Section 6 builds an intermediary asset pricing model that accounts for the empirical findings and studies policy implications. Section 7 concludes.

2 Data

In this paper, I combine data from multiple sources to assess the co-movement between the long-term interest rate and corporate bond credit spreads, the significance of life insurers, and their real impacts.

The Long-term Interest Rate. I use the US Treasury yield curve constructed by Liu and Wu (2021) at a daily frequency. In particular, I use the yield on 10-year US Treasury notes as the proxy for the long-term interest rate.

Corporate Bonds. I combine monthly data from the Mergent Fixed Income Securities Database (Mergent FISD), the Trade Reporting and Compliance Engine (TRACE), and the WRDS Bond Returns for corporate bond prices, quantities, and characteristics. The dataset provides comprehensive coverage for US corporate bonds between 2000 and 2019. From Mergent FISD, I obtain information on the bond issuer, maturity, duration, credit ratings, outstanding amount, issuance date, coupons, transaction volume, and default history at a monthly frequency. The WRDS Bond Returns dataset also provides end-of-month transaction prices and yields extracted from TRACE. For any given month, I focus on bonds with at least one observed transaction price.

Mergent FISD reports three "raw" credit ratings from Standard and Poor's (S&P), Moody's Analytics, and the Financial Industry Regulatory Authority (FINRA). For my analysis, I adopt the NAIC system that consolidates the three ratings into one and sorts them into six NAIC categories. The NAIC rating is the most relevant risk metric for insurance companies, as it determines the capital requirement for each bond. Bonds in NAIC 1 and NAIC 2 are investment-grade, while bonds in NAIC 3-6 are speculative-grade. Table 1 replicates Table 2 in Becker and Ivashina (2015), which summarizes the 5-year default rate and capital requirement of each NAIC category. Corporate bonds with lower NAIC ratings have higher default rates and entail more stringent capital requirements.

NAIC Category	Credit Ratings	Investment Grade	5-year Default Rate (1990-2010)	Capital Requirement
NAIC 1 (highest)	AAA, AA, A	✓	0.00%, 0.09%, 0.69%	0.3%
NAIC 2	BBB	✓	2.62%	0.96%
NAIC 3	BB	x	6.76%	3.39%
NAIC 4	B	x	8.99%	7.38%
NAIC 5	CCC	x	34.38%	16.96%
NAIC 6 (lowest)	CC, C, D	x	n.a.	19.50%

Table 1. **The NAIC Rating System.**

This Table summarizes the characteristics of corporate bonds belonging to different NAIC categories. The data on the cumulative 5-year default rates and capital requirements are drawn from Becker and Ivashina (2015).

For part of my empirical analysis, I also use aggregate bond yield indices from the Intercontinental Exchange ("ICE") and bond credit default swaps (CDS) spreads from Markit IHS. In particular, I use the par spreads of credit default swaps with a 5-year tenor, which form the most liquid segment of the CDS market (e.g., Blanco, Brennan and Marsh, 2005). The CDS data has been available since the beginning of 2001.

Life Insurers. I obtain regulatory data on life insurers' end-of-year bond holdings and long-term bond transactions from the National Association of Insurance Commissioners (NAIC). I examine the Schedule D information in life insurers' regulatory reports to NAIC.[4] The NAIC data also contains bond identifiers (CUSIP codes) that allow me to match the bonds held and transacted by life insurers to those in Mergent FISD and WRDS Bond Returns. In addition to the NAIC data, I use data on the aggregate balance sheets of life insurers from the Financial Accounts of the United States and data on life insurers' stock prices from the Center for Research in Security Prices (CRSP).

Investment. To jointly study the life insurers' bond holdings and firm investment, I combine Mergent FISD with Compustat to form a panel of bond-issuing firms. The data include the quarterly capital expenditure and bond ownership structure of all firms in Mergent FISD.

FOMC Meeting Dates. Hillenbrand (2023) summarized the dates of all FOMC meetings since September 1982. In my analysis, I focus on meetings between 1997 and 2022, excluding all unscheduled meetings.

Sample Period. Throughout the paper, I separate my analysis into two time periods: before the Financial Crisis (1997-2007 or 2000-2007) and after the Financial Crisis (2010-2022). The aggregate credit spread data begins in 1997, while the individual bond data in WRDS Bond Returns dates back to January 2000. I remove the Global Financial Crisis (2007-2009), which features large-scale fire sales and liquidity interruptions that potentially confound the mechanism of interest. Similarly, I exclude data from March 2020 in the post-crisis period to avoid capturing the bond market disruptions during the COVID-19 crisis. For analysis involving life insurers' bond ownership or real effects, I shorten the post-crisis sample to 2010-2019 due to data availability.

[4] See https://content.naic.org/sites/default/files/ASB-PCA-22_0.pdf for the structure and content of the dataset.

3 The Long-term Interest Rate and Bond Credit Spreads

In this section, I examine the co-movement between the long-term interest rate and corporate bond credit spreads and contrast the findings before and after the 2008 Financial Crisis.

3.1 Pass-through of the Long-term Interest Rate

I begin by studying the 10-year Treasury yield pass-through in the cross-section of corporate bonds. I run the following regressions to estimate the pass-through

$$y_{it} = \alpha_i + \sum_{k=1}^{6} \beta_k \cdot \mathbf{1}_{\{\text{NAIC } k\}} \cdot y_t^{(10)} + \Gamma \mathbf{X}_{it} + \varepsilon_{it}, \tag{1}$$

where y_{it} is the yield of bond i at time t, $y_t^{(10)}$ is the 10-year Treasury yield, α_i are bond fixed effects, and $\mathbf{X}_{it}$ are additional controls. Here I control for bond characteristics such as bond size (outstanding amount), liquidity (trading volume), maturity, duration, credit ratings, coupon amount, and coupon frequency, as well as the recent default rate of each NAIC category.

The regression coefficient β_k measures the pass-through rate of the 10-year Treasury yield to the yield of NAIC k corporate bonds. Intuitively, β_k measures the average response of the yields of NAIC k bonds when the 10-year Treasury yield increases by one percentage point. Figure 2 shows the estimates of β_k for the sample before the Financial Crisis (2000-2007) and the sample after the Financial Crisis (2010-2022).

After the Financial Crisis, there is a strong relationship between credit ratings and the pass-through of the long-term interest rate. The yields of the safest bonds (i.e., NAIC 1 and 2) move strongly together with the 10-year Treasury yield. However, as the credit rating declines, the sensitivity to the long-term interest rate diminishes and eventually reverses for bonds with the lowest credit ratings (NAIC 5 and 6). The effects are large, especially for bonds in NAIC 5 and 6, whose yields decrease by much more than 1% when the 10-year Treasury yield increases by 1%. Accordingly, the credit spreads between low-rating and high-rating bonds shrink when the 10-year Treasury yield is high.

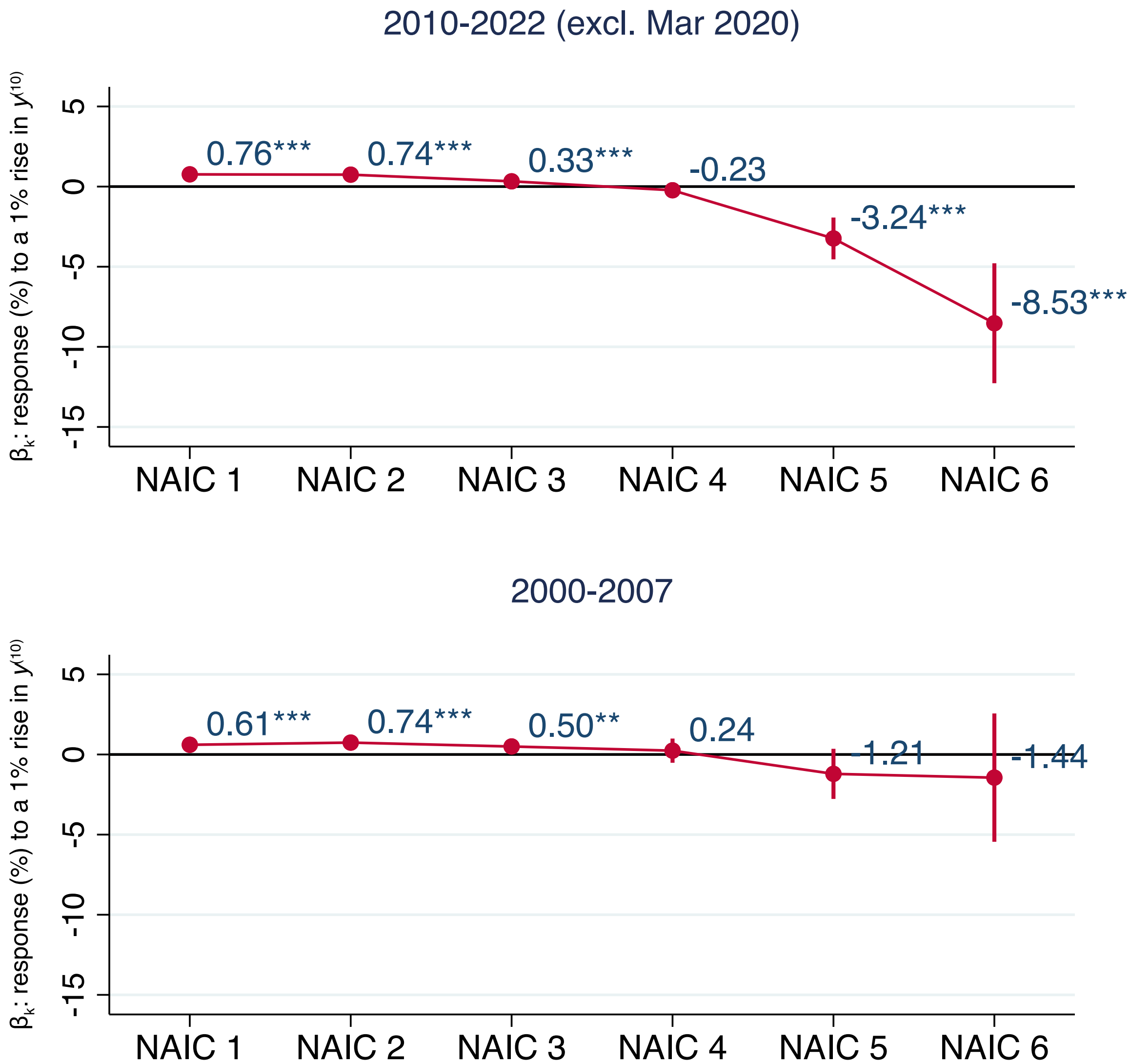

Figure 2. **10-year Treasury yield pass-through.**

This figure plots the coefficients β_k estimated from regression (1), controlling for the trading volume, outstanding amount, maturity, duration, credit ratings, coupon amount, and coupon frequency of each bond, and the recent default rate of each NAIC category. Both corporate bond yields and the Treasury yield are in percentage points. The top panel shows results for the post-crisis sample, while the bottom panel shows results for the pre-crisis sample. The t-statistics shown in brackets are based on standard errors clustered at the issuer and year-month levels. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Before the Crisis, by contrast, the relationship was much weaker before the Financial Crisis. Bonds in NAIC 1 and 2 had a strong positive co-movement with the 10-year Treasury yield. However, the responses of lower-rating bonds were much smaller and statistically insignificant from those in NAIC 1 and 2.

Overall, the results indicate a significant increase in the co-movement between the long-term interest rate and corporate bond yields around the Financial Crisis.

3.2 Credit Spread Responses to the long-term interest rate

Next, I focus on credit spreads in the cross-section of corporate bonds. To this end, I analyze the spreads between bonds with lower ratings (NAIC k, $k \geq 2$) and the highest credit rating (NAIC 1). Relative to the empirical specification in regression (1), I further include maturity-time fixed effects $\alpha_{\tau(i),t}$,

$$y_{it} = \alpha_i + \alpha_{\tau(i),t} + \sum_{k=2}^{6} \beta_k \cdot \mathbf{1}_{\{\text{NAIC } k\}} \cdot y_t^{(10)} + \Gamma \mathbf{X}_{it} + \varepsilon_{it}. \tag{2}$$

To construct the maturity-time fixed effects, I sort bonds into small maturity segments with a size of 1 year and interact these maturity dummies with the time variable. The fixed effects thus control for the maturity differences across each NAIC category[5] and variations in the yield curve over time. Additionally, the fixed effects absorb one NAIC category for any given month, which I normalize as NAIC 1. Hence, the coefficient β_k (for $k \geq 2$) measures the average response that changes in the 10-year Treasury yield induce in the *spreads* between NAIC k bonds and NAIC 1 bonds. A negative coefficient indicates that credit spreads move in opposite directions as the long-term interest rate.

Figure 3 plots the estimated coefficients and delivers one of the main results of this paper. In the post-crisis sample, corporate bond credit spreads fall when the 10-year Treasury yield increases. The result is significant for all NAIC categories and is stronger for lower ratings.

To understand the significance of the results, it is useful to consider the following decomposition of bond yields

$$y_{t,\text{NAIC } k} = y_{t,\text{NAIC } 1} + (\text{Credit Spread})_{t,\text{NAIC } k}.$$

[5] It is important to control for maturity as riskier bonds tend to be of shorter maturity. Figure A.4 and Figure A.5 further confirm that the results also hold when we compare bonds within maturity segments.

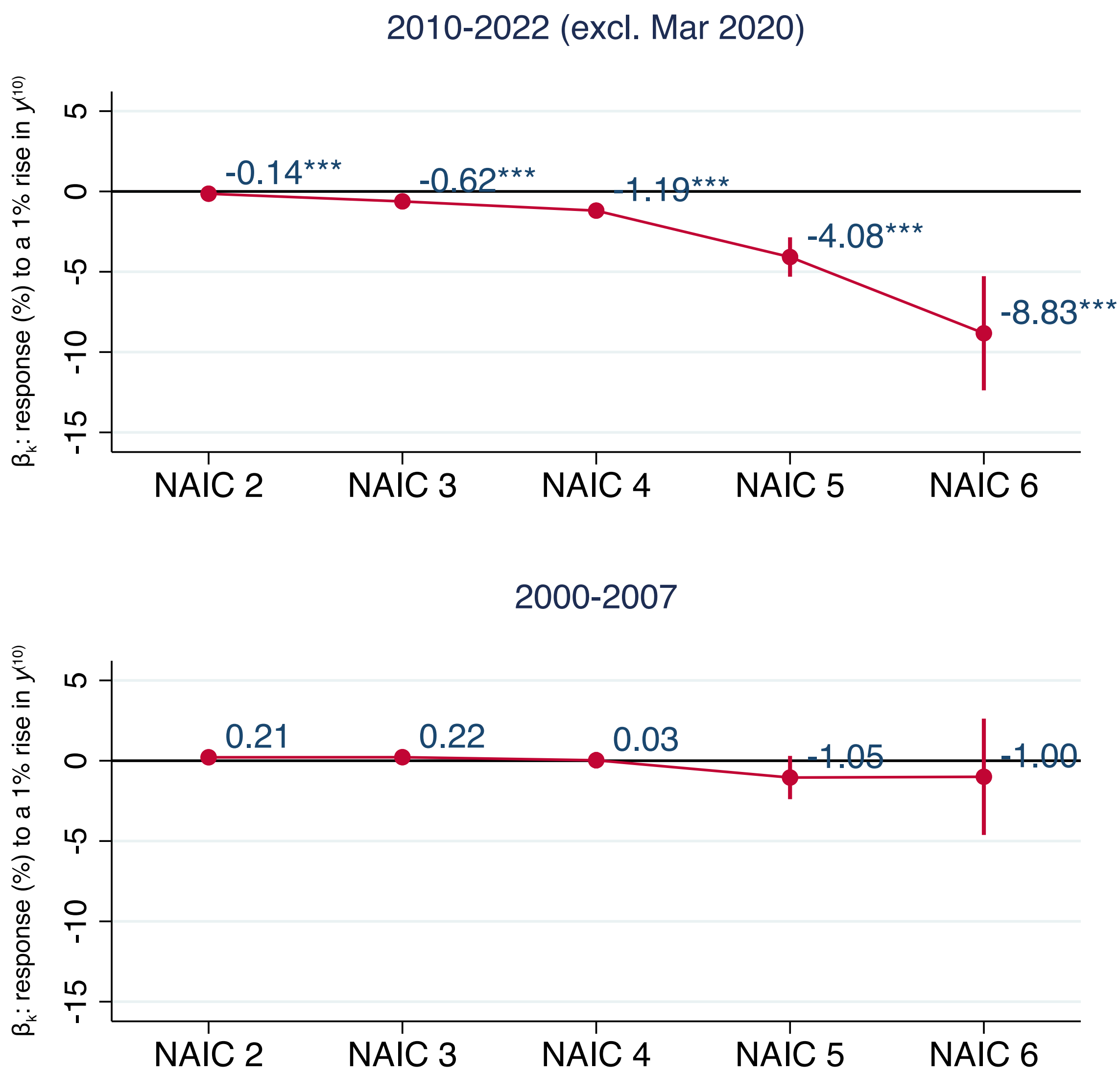

Figure 3. **Credit Spread Responses to the Long-term Interest Rate.**

This figure plots the coefficients β_k estimated from regression (2), controlling for the trading volume, outstanding amount, maturity, duration, credit ratings, coupon amount, and coupon frequency of each bond, and the recent default rate of each NAIC category. Both corporate bond yields and the Treasury yield are in percentage points. The top panel shows results for the post-crisis sample, while the bottom panel shows results for the pre-crisis sample. The t-statistics shown in brackets are based on standard errors clustered at the issuer and year-month levels. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Effects of the level of the long-term interest rate on total bond yields depend on (1) how it affects the safest segment of the bond market (yields of NAIC 1 bonds) and (2) how it affects credit spreads relative to NAIC 1. In Section 3.1, I show that NAIC 1 bond yields co-move positively with the 10-year Treasury yield, with a pass-through coefficient of about 0.76 after the Financial Crisis. For safer bonds (NAIC 2, 3, and 4), the effect on credit spreads partially offsets the changes in NAIC 1 yields, making bond yields less sensitive to the 10-year Treasury yield.

The effects are much larger for the riskiest bonds. For example, when the 10-year Treasury yield increases by 25 basis points, the yields of CCC corporate bonds (NAIC 5) fall by 102 basis points relative to corporate bonds rated A or better (NAIC 1), which is large enough to fully offset the increase in NAIC 1 yields and lower the NAIC 5 yields in absolute terms. As already shown in Figure 2, the yields of NAIC 5 and 6 bonds move in opposite directions as the 10-year Treasury yield. Surprisingly, increases in the long-term interest rate lower the funding costs of the riskiest firms.

It is also worth noting that the effects on credit spreads were absent before the Financial Crisis. In the sample before 2007, credit spreads had very small and statistically insignificant responses to the 10-year Treasury yield. Earlier lack of response implies that the co-movement between the long-term interest rate and credit spreads only emerged after the Crisis.

Figure A.1 and Figure A.2 in the Appendix visualize the time series of the 10-year Treasury yield and various corporate bond credit spreads. The correlation pattern supports the conclusion of this Section. A strong negative correlation exists between corporate bond credit spreads and the 10-year Treasury yield after the Financial Crisis but not before the Financial Crisis.

A potential confounding factor in these findings is credit risk movements. For example, stronger economic growth can lower credit risk, which could further lead to both lower credit spreads and increased interest rates (e.g., Wu and Zhang, 2008). To better control for variations in perceived credit risk, I further control the average CDS spread for each NAIC category, which is a direct and real-time measure of the price of default risk (e.g., Longstaff, Mithal and Neis, 2005).[6] By controlling for CDS spreads,

[6] Absent frictions, CDS spreads should coincide with credit spreads. However, empirically, the CDS-

we can eliminate the impacts of changing credit risk and more clearly understand how long rates affect credit spreads.[7] Figure A.3 shows the coefficients estimated after controlling for CDS spreads, which exhibit the same pattern as in Figure 3 and confirm the validity of the findings.

3.3 High-frequency Evidence: FOMC Announcements

Section 3.1 and Section 3.2 discussed the unconditional co-movements between the long-term interest rate and corporate bond credit spreads. In this section, I strengthen the evidence by showing identified evidence on the impact of 10-year Treasury yields on bond credit spreads using high-frequency shocks around FOMC meetings. A recent study by Hillenbrand (2023) documents that a short window around FOMC meetings explains the majority of the long-run movements in long-term interest rates for the past 30 years. The potential explanation is that FOMC meetings disseminate information concerning the future paths of interest rates. Inspired by Hillenbrand (2023), I construct shocks to the 10-year Treasury yield as the changes in the yield in 2-day windows around FOMC meetings,

$$\Delta y_t^{(10)}\Big|_{\text{FOMC}} = y_{t+1}^{(10)} - y_{t-1}^{(10)},$$

where t is an FOMC announcement day. Using a local projection method (Jordà, 2005), I then estimate the impulse responses of corporate bond credit spreads to the high-frequency 10-year Treasury yield shocks. The regressions are as follows

$$\text{Spread}_{t+h}^k - \text{Spread}_{t-1}^k = \alpha_h + \beta_h \left(\Delta y_t^{(10)}\Big|_{\text{FOMC}} \right) + \varepsilon_{t,h}. \tag{3}$$

Here the left-hand-side variable Spread_t^k is constructed as the difference between the bond yield index for rating k and the bond yield index for the AAA rating. The coefficient β_k measures the cumulative response of Spread_t^k to a 1% positive innovation in the 10-year Treasury yield at a horizon of h trading days.

bond basis is known to be non-zero and dispersed across bonds (e.g., Bai and Collin-Dufresne, 2019). Following Longstaff, Mithal and Neis (2005) and Mota (2023), I view the CDS spread as a better measure of the market price of credit risk, while credit spreads potentially also reflect other factors, including liquidity, convenience yields, and the risk-bearing capacity of bond investors such as life insurers.

[7] Importantly, life insurers are not the major investors in the CDS market, so the CDS spreads are potentially not affected by the balance sheets of life insurers.

Figure 4 and Figure 5 plot the impulse response functions estimated from the local projections (3) for the post-crisis sample and the pre-crisis sample, respectively. After the Financial Crisis, positive shocks to the 10-year Treasury yield led to large and significant declines in the credit spreads between AAA bonds and bonds with lower credit ratings. For example, a 1% increase in the 10-year Treasury yield depresses the spread between single B and AAA bonds around 1% at a 10-day horizon. The negative responses of credit spreads are surprising in light of the literature on short-term monetary policy, which typically finds that higher interest rates raise risk premia and corporate bond spreads (e.g., Gertler and Karadi, 2015).

Nevertheless, similar shocks to the 10-year Treasury yield produced statistically insignificant or even the opposite responses before the Crisis. The results suggest that there was a shift in the transmission of policy shocks around the Financial Crisis — the FOMC windows did not influence credit spreads prior to the Financial Crisis.

The results so far focus on credit spreads relative to AAA bonds. Figure A.7 in the Appendix replaces the spreads relative to AAA bonds with the spreads relative to US Treasury securities and finds similar impulse responses. Figure A.8 instead estimates the local projections directly on the yield indices for different credit ratings. The yield of AAA bonds exhibits a persistently positive dependence on the 10-year Treasury yield, while yields of other investment-grade bonds (AA, A, BBB) are less sensitive. Consistent with the findings of Section 3.1, the yields of speculative-grade bonds (BB and lower) seem to move in the opposite direction as the 10-year Treasury yield.

An alternative explanation might be that the effects are mainly due to short rates that move simultaneously with long rates. As a robustness check, I include changes in the 1-month Treasury yield around FOMC meetings as controls for short-term monetary policy. Further, long-rate movements around FOMC meetings could contain information on the long-run performance of corporate bonds. To distinguish this alternative channel, I include changes in CDS spreads around FOMC windows as controls for credit risk news. Figure A.9-Figure A.12 plot the post-crisis impulses responses estimated after controlling for short rate changes and CDS spreads, which again confirm

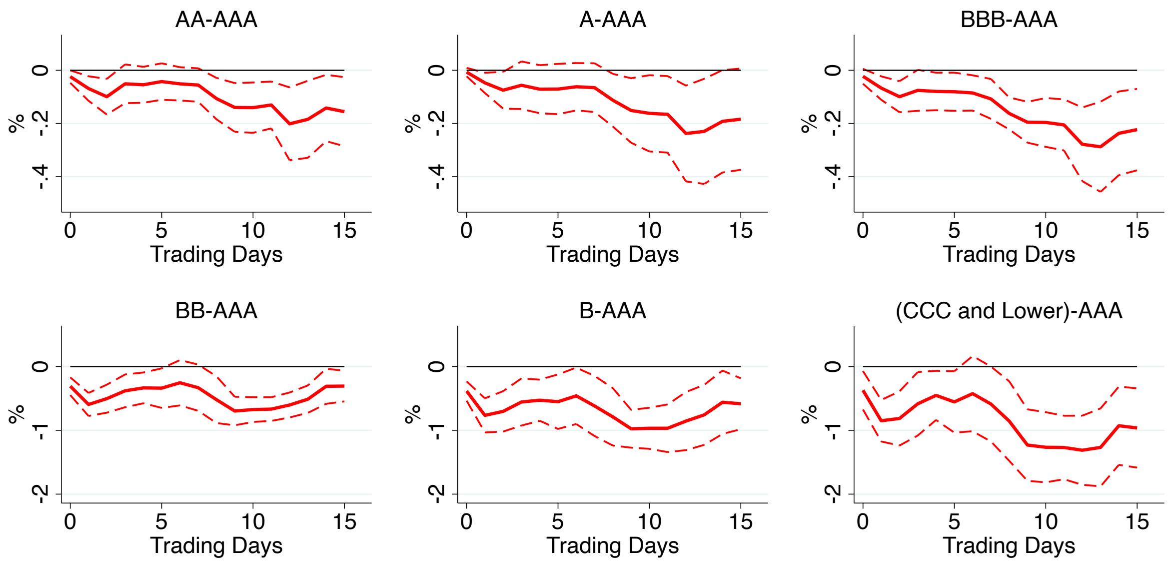

Figure 4. **Impulse Responses of Credit Spreads (2010-2022).**

This figure plots the cumulative responses of credit spreads to a 1% increase in the 10-year Treasury yield and 90% confidence intervals based on Newey-West standard errors with a maximum lag of 30 trading days.

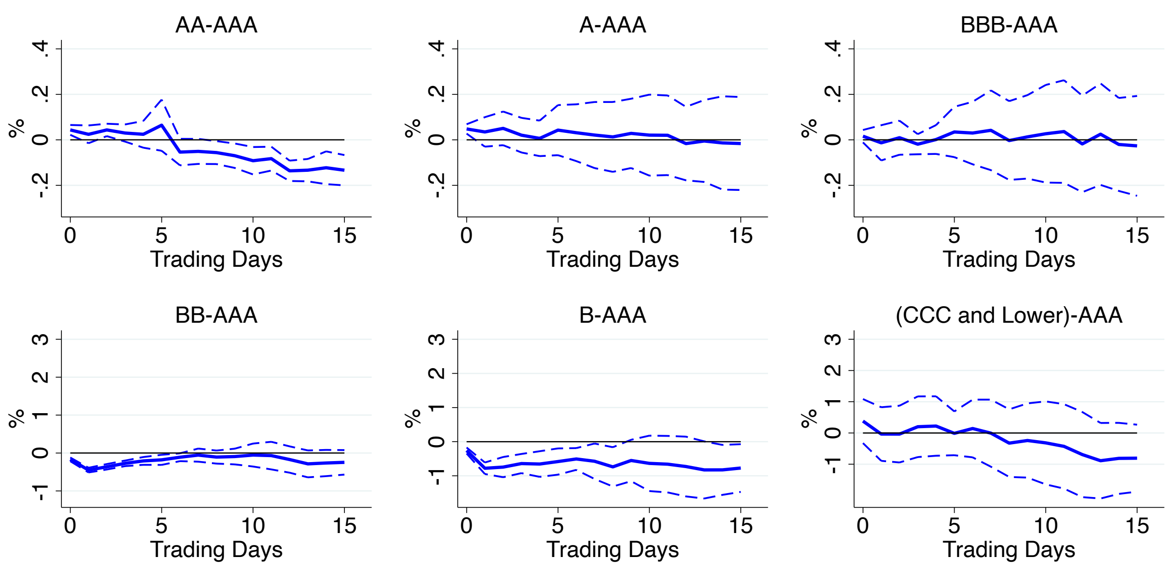

Figure 5. **Impulse Responses of Credit Spreads (1997-2007).**

This figure plots the cumulative responses of credit spreads to a 1% increase in the 10-year Treasury yield and 90% confidence intervals based on Newey-West standard errors with a maximum lag of 30 trading days.

the patterns shown in Figure 4.[8] Figure A.13 further shows that the responses of CDS spreads to long rates are much smaller than the responses of credit spreads.

3.4 Discussion on the Mechanism

In this section, I document a negative co-movement between the long-term interest rate and corporate bond credit spreads after the Financial Crisis — credit spreads rise when the 10-year Treasury yield falls. The results might seem counterintuitive relative to the standard macro-finance literature, where low interest rates lower risk premia by enhancing the balance sheets of financial intermediaries or encouraging reaching-for-yield behaviors. While the standard theory has been proven true in the context of bank lending (e.g., Kashyap and Stein, 2000; Jiménez et al., 2014; Whited, Wu and Xiao, 2021), I show there exists an opposite effect in the corporate bond market.

The favored explanation this paper proposes is the impact of life insurers' duration mismatch. In Section 4.1, I empirically document a shift in the life insurance sector's duration mismatch and investigate their contribution to the relationship between the long-term interest rate and credit spreads. Under the new duration mismatch, life insurers suffer from equity drawdowns when the long-term interest rate declines, which increase their portfolio exposure to corporate bonds. Since life insurers are the largest investor group in the US bond market, elevations in their effective risk aversion lead to higher equilibrium corporate bond credit spreads.

While life insurers are important for this negative co-movement, there exist other potential channels through which long-term interest rates affect risk premia in the bond market. For example, pension funds have a balance sheet structure similar to life insurers, which makes them susceptible to interest rate fluctuations too. Increases in the long-term interest rate could also result from endogenous monetary policy responses to higher credit spreads. Section 6 analyzes a structural model of the corporate bond market to quantify the fraction of the co-movement explained by life insurers. In the model, insurers' duration mismatch can explain a large fraction of the empirical co-movement, while the unexplained fraction is potentially due to other mechanisms discussed above.

[8] Hillenbrand (2023) also argues that the long-rate changes around FOMC meetings $\left(\Delta y_t^{(10)}\Big|_{\text{FOMC}}\right)$ are not driven by short-term economics news or the Fed's policy responses to them.

4 The Role of Life Insurers

Next, I turn to the potential channel through which the long-term interest rate negatively affects corporate bond credit spreads — life insurers' duration mismatch. As shown in Figure 1, life insurers are the largest investor group in corporate bonds and own more than 30% of all corporate bonds in the US as of 2022 (see also Koijen and Yogo, 2023).

4.1 Duration Mismatch

Modern life insurers issue various life insurance and annuity products while investing primarily in fixed-income markets. For example, variable annuities, which are long-term mutual fund products with minimum return guarantees, have now become the largest component of life insurers' liabilities (Koijen and Yogo, 2022). Maturity transformation is a key component of life insurers' operations, as they invest in fixed-income securities (e.g., corporate bonds) and turn them into longer-term liabilities (e.g., variable annuities). Unlike banks that turn long-term loans into short-term deposits, the maturity transformation of life insurers takes the opposite direction. Hence, life insurers' market equity could be hurt when the long-term interest rate is low.

Table 2 summarizes the balance sheet structure of the US life insurance sector. Life insurers' assets are predominantly fixed-income debt securities and mutual fund shares that are ultimately also invested in fixed-income assets. Corporate bonds are an important asset class for life insurers as they make up 80% of all the debt securities they hold. On the liability side, life insurers raise funding mostly by selling life insurance and annuities, which typically have a longer maturity and duration than corporate bonds.

I then estimate the exposure of life insurers' market equity to the 10-year Treasury yield. In particular, I run the following regressions

$$\text{ExcessReturn}_t = \alpha + \beta \Delta y_t^{(10)} + \text{Controls} + \varepsilon_t. \tag{4}$$

Here ExcessReturn_t is the stock return of the life insurance sector in week t in excess of the return on the S&P 500 index, and $\Delta y_t^{(10)}$ is the change in the 10-year Treasury yield in week t. The coefficient β measures the sensitivity of life insurers' market equity to

Financial Assets ($ tn)		Liabilities ($ tn)	
Short-term Assets	0.20	Life Insurance Reserves	2.25
Debt Securities	4.43	Annuity Reserves	3.79
– Corporate Bonds	– 3.56	Other Liabilities	2.34
Loans & Equities	1.54		
Mutual Fund Shares	1.70		
Other Financial Assets	1.55		
Total	9.42	Total	8.38

Table 2. **The Balance Sheets of the Life Insurance Sector.**

This table summarizes the main financial assets and liabilities of life insurers in the US. The data is from the Financial Accounts of the United States at 2020 Q4.

changes in the 10-year Treasury yield. If the duration of life insurers' liabilities exceeds their assets, the coefficient β should be positive, as higher interest rates lower the value of their liabilities more than their assets and thereby boost the value of their equity. Therefore, $(-\beta)$ can also be interpreted as an estimate of the duration of life insurers' market equity, which is defined as $D_E = -\partial E_t / \partial y_t^{(10)}$ (E_t is the market equity of life insurers in week t). A positive estimate of β then indicates that life insurers' equity has a negative duration (e.g., Ozdagli and Wang, 2019; Koijen and Yogo, 2022).

Table 3 contrasts life insurers' exposure to the 10-year Treasury yield before and after the Financial Crisis. Before the Financial Crisis, the stock returns of the life insurance sector did not seem to be affected by the long-term interest rate. It implies that life insurers were largely hedged against interest rate fluctuations. However, the estimated β is significantly away from zero in the post-crisis period. The estimated coefficient of 7.179 suggests that life insurers' equity value grows by more than 7% when the 10-year rate increases by 1%, so life insurers face severe duration mismatch after the 2007-2008 Financial Crisis.

Figure A.14 in the Appendix shows the evolution of life insurers' duration mismatch over time. In particular, the figure plots the two-year rolling estimates of β. Consistent with existing findings (e.g., Hartley, Paulson and Rosen, 2016; Ozdagli and Wang, 2019; Huber, 2022), the estimated exposure to interest rate risk is small and insignificant from

	Dep. Var.: ExcessReturn_t	
	1997-2007	2010-2020
$\Delta y_t^{(10)}$	-0.845	7.179***
	[-0.788]	[2.782]
S&P 500 Return	✓	✓
$\Delta y_t^{(1m)}$	✓	✓
Observations	520	520

Table 3. **Life Insurers' Duration Mismatch.**

This table shows the coefficients estimated from regression (4), controlling for the market return and changes in the one-month Treasury yield. The first column shows the result for the pre-crisis sample, while the second column shows the result for the post-crisis sample. Both the excess return and the 10-year Treasury yield are in percentage points. The t-statistics shown in brackets are based on robust standard errors. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

zero before the Financial Crisis, whereas large and persistent deviations from zero are found after the Financial Crisis. Overall, the results suggest that life insurers' duration mismatch increased dramatically after the Financial Crisis.

Figure 6 plots the market leverage ratio of the life insurance sector against the 10-year Treasury yield. Consistent with the previous findings, life insurers become less levered when the 10-year Treasury yield is high, as their liabilities become less expensive and their equity declines.

The Source of Duration Mismatch. In this paper, I take a neutral stance on the cause of the shift in life insurers' duration mismatch. Nevertheless, the existing literature typically attributes the new duration mismatch post-2008 to lower interest rates after the Financial Crisis. To illustrate this point, we can write life insurers' equity as $E = A - L$, where A is assets, L is liabilities, and $\ell = L/E$ is the leverage ratio. The duration of equity is then $D_E = (AD_A - LD_L)/E = (1+\ell)D_A - \ell D_L$. Before the Crisis, life insurers were hedged against interest rate risk by choosing $D_A \approx [\ell/(1+\ell)]D_L$ so that $D_E \approx 0$. Meanwhile, the assets and the liabilities both have a positive *convexity* as lower

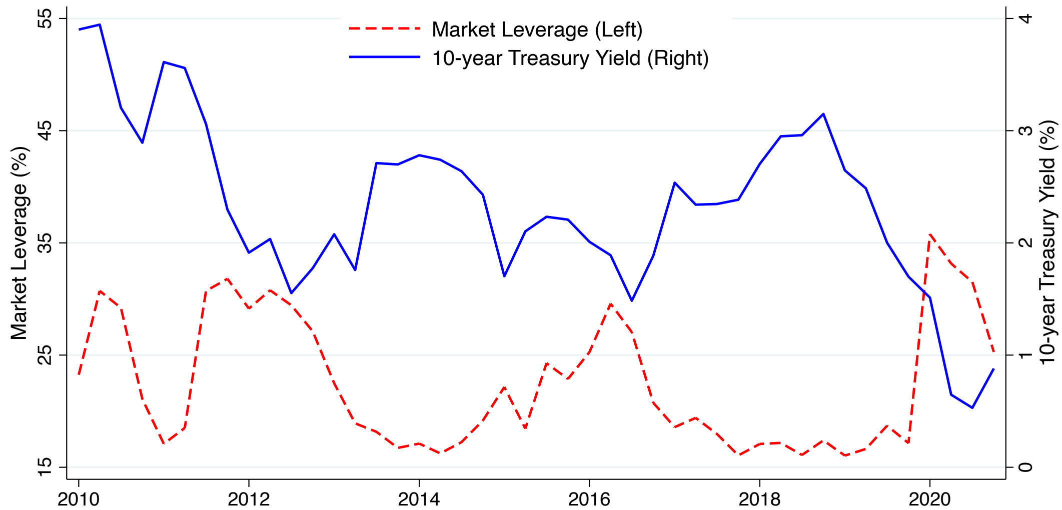

Figure 6. **Life Insurers' Market Leverage.**

This figure plots the market leverage ratio of the US life insurance sector and the 10-year Treasury yield. The market leverage ratio is defined as the "fair" value of their liabilities (from Compustat) divided by their market equity value.

interest rates increase bond duration,

$$C_A := -\frac{\partial D_A}{\partial y_t^{(10)}} > 0, \quad C_L := -\frac{\partial D_L}{\partial y_t^{(10)}} > 0.$$

Crucially, the convexity of life insurers' liabilities is greater than the convexity of their assets ($C_L > C_A$), meaning that D_L rises faster than D_A when interest rates fall.[9] As a result, D_E became negative in the low-interest-rate environment after the Crisis.

Furthermore, there exist other market and institutional frictions that prevent life insurers from increasing their asset duration D_A to close the duration gap. Domanski, Shin and Sushko (2017) and Greenwood and Vissing-Jorgensen (2018) argue that large-scale portfolio rebalancing of life insurers towards longer-term assets puts downward pressure on the long-term interest rate, which could further exacerbate the duration mismatch. Ozdagli and Wang (2019) emphasize the role of portfolio adjustment frictions resulting from the cost of large bond trades. Koijen and Yogo (2021) discussed several other reasons why life insurers do not fully hedge their interest rate exposure,

[9]One likely explanation is that life insurers' liabilities have a longer maturity than their assets, as the convexity of an asset typically increases in its maturity. It could also be because the embedded options in some of life insurers' variable annuity liabilities are less exercised when interest rates are low (e.g., Ozdagli and Wang, 2019; Koijen and Yogo, 2022).

including market incompleteness (i.e., the scarcity of long-term assets and options), risk-shifting motives, and regulatory distortions. In particular, Huber (2022) and Sen (2023) argue that the regulatory framework imposed on life insurers might not properly capture the interest rate risk and thus distorts life insurers' hedging incentives.[10]

4.2 Life Insurers and Long Rate Pass-through

I further show that life insurers play an important role in shaping the co-movement between the long-term interest rate and corporate bond credit spreads. Section 3.1 illustrates that long-term interest rates have heterogeneous impacts on the cross-section of corporate bonds after the Financial Crisis. If life insurers' duration mismatch is the main channel through which the long-term interest rate affects credit spreads, we should expect the pattern to be more pronounced in bonds owned by life insurers, which is indeed confirmed by the following results.

In this Section, I revisit the results of Section 3.1. To examine how life insurers affect the pass-through of the long-term interest rate, I modify regression (1) in Section 3.1 as

$$y_{it} = \alpha_i + \sum_{k=1}^{6} \left(\beta_k^0 + \beta_k^1 \cdot \mathbf{1}_{\{\text{insurance ownership}\}} \right) \cdot \mathbf{1}_{\{\text{NAIC } k\}} \cdot y_t^{(10)} + \Gamma \mathbf{X}_{it} + \varepsilon_{it}. \tag{5}$$

Here $\mathbf{1}_{\{\text{insurance ownership}\}}$ is a dummy variable that indicates whether a bond has positive life insurance ownership at the end of the year. The pass-through coefficients are allowed to differ between bonds with no life insurance ownership and bonds with positive life insurance ownership. β_k^0 measures the pass-through of the 10-year Treasury yield to NAIC k bonds *not* held by life insurers, while β_k^1 measures the pass-through to NAIC k bonds that are held by life insurers.

Figure 7 plots the estimated coefficients from regression (5) for the post-crisis period. Indeed, the relationship between pass-through and credit ratings in Figure 2 is only present in bonds with positive life insurance ownership. Bonds with no life insurance ownership exhibit either no significant pass-through (NAIC 1-4) or the opposite pass-through compared to Figure 2 (NAIC 5-6). Moreover, the pass-through estimates for bonds with life insurance ownership have comparable magnitudes to those shown in

[10] For more examples of regulatory distortions in the insurance sector, see also Lee, Mayers and Smith (1997), Ellul, Jotikasthira and Lundblad (2011), Becker, Opp and Saidi (2022), and Ellul et al. (2022).

Figure 2. Thus, the empirical findings of Section 3.1 are driven primarily by life-insurer-owned bonds.[11]

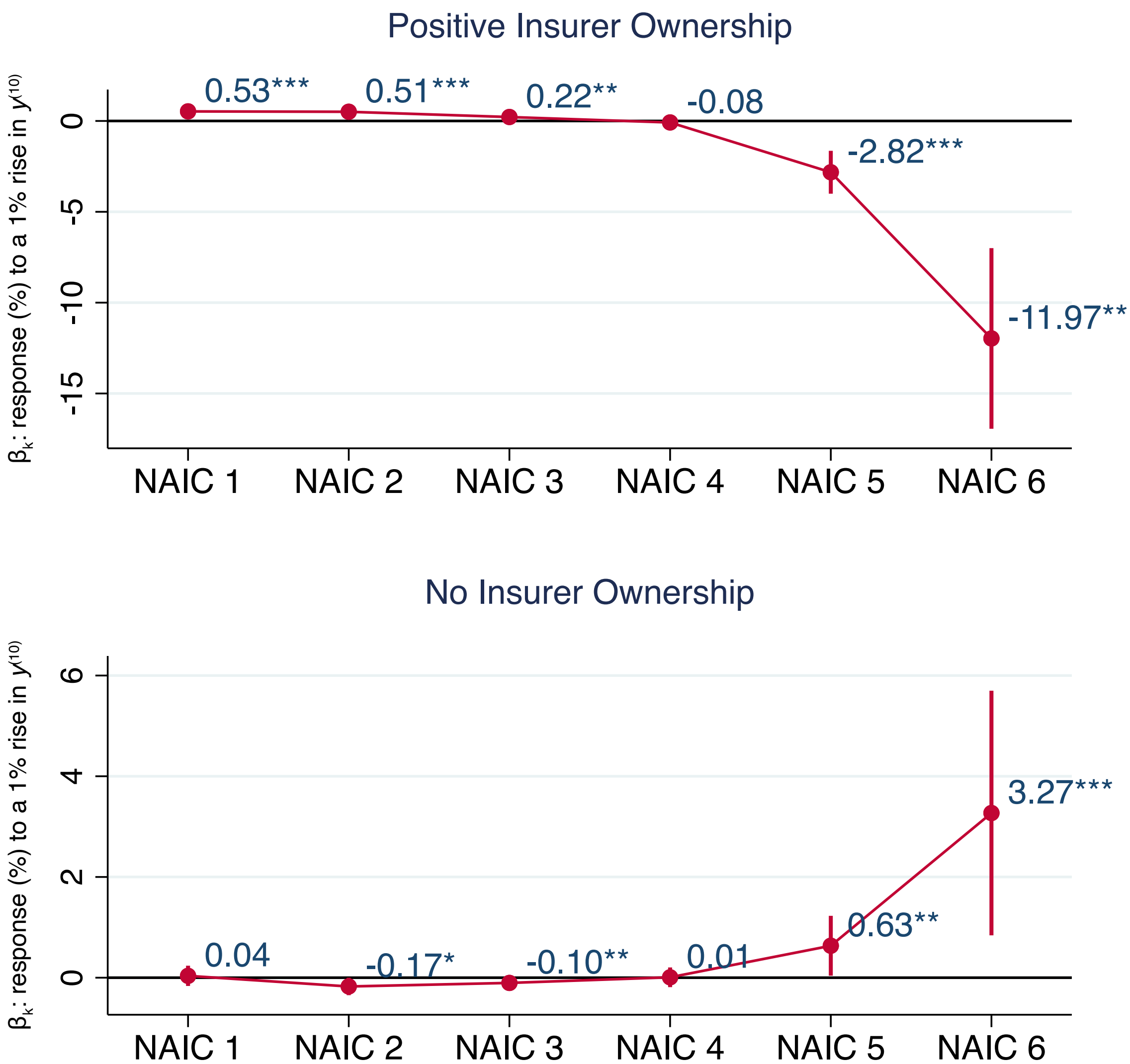

Figure 7. **Pass-through by Life Insurance Ownership.**

This figure plots the coefficients β_k estimated from regression (5) for the period of 2010-2019, controlling for the trading volume, outstanding amount, maturity, duration, credit ratings, coupon amount, and coupon frequency of each bond, and the recent default rate of each NAIC category. Both corporate bond yields and the Treasury yield are in percentage points. The top panel shows estimates for β_k^1, while the bottom panel shows estimates for β_k^0. The t-statistics shown in brackets are based on standard errors clustered at the issuer and year-month levels. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

[11] Figure A.6 shows that, before the Financial Crisis, the path-through pattern is absent in both insurer-owned bonds and non-insurer-owned bonds. This finding is consistent with the fact that life insurers did not face duration mismatch pre-2008.

4.3 Regression Discontinuity Design

Next, I recover the fraction of each bond owned by life insurers from NAIC regulatory reports and study the role of life insurance ownership in shaping the results of Section 3.2.[12]

To further sharpen the identification, I utilize an exogenous discontinuity in investor composition stemming from mutual funds' investment mandates (e.g., Li and Yu, 2023). Bai, Li and Manela (2023) and Li and Yu (2023) document that a large fraction of corporate bond funds are "intermediate-term" and are mandated to only invest in bonds whose maturity is less than 10 years. For a bond with a maturity greater than 10 years at issuance, it will experience a surge in demand from mutual funds once it ages to the point where its maturity drops below 10 years. As a result, the bond's ownership by other investors, such as life insurers, jumps downward.

Figure 8 visualizes the discontinuity in life insurers' ownership share around the maturity threshold of 10 years and a quarter. Indeed, we observe a large discontinuous jump in life insurers' ownership shares at the maturity threshold. Bonds whose maturity falls below the threshold are much less likely to be held by life insurers.[13]

I further test the existence of the discontinuity using both OLS regressions and the robust bias-correction method of Calonico, Cattaneo and Titiunik (2014). In the OLS method, I regress the life insurer ownership share on a dummy variable indicating whether a bond's maturity is above the threshold $\mathbf{1}_{\{\text{maturity}_{it} \geq c\}}$ where $c = 10.25$.

$$\varphi_{it}^{\text{Ins}} = \alpha + \beta \cdot \mathbf{1}_{\{\text{maturity}_{it} > c\}} + \Gamma \mathbf{X}_{it} + \varepsilon_{it}.$$

In the robust bias-correction method, I treat the problem as a sharp regression discontinuity design (RDD), where the treatment status is determined by $\mathbf{1}_{\{\text{maturity}_{it} \geq c\}}$ and the outcome variable is the life insurer ownership share.

[12] In the main regression of this section, I use end-of-year observations as insurance companies' holdings data is only recorded at the end of each year.

[13] The actual maturity cutoff for the discontinuity (10.25 years) is slightly more than 10 years. A likely explanation is mutual funds' window-dressing behavior, meaning that mutual funds only need to comply with their mandates at the end of each quarter when they disclose their bond holdings (Morey and O'Neal, 2006; Agarwal, Gay and Ling, 2014). Therefore, mutual funds could invest in bonds with a maturity between 10 and 10.25 years even if they are mandated to invest in bonds with a maturity of less than 10 years.

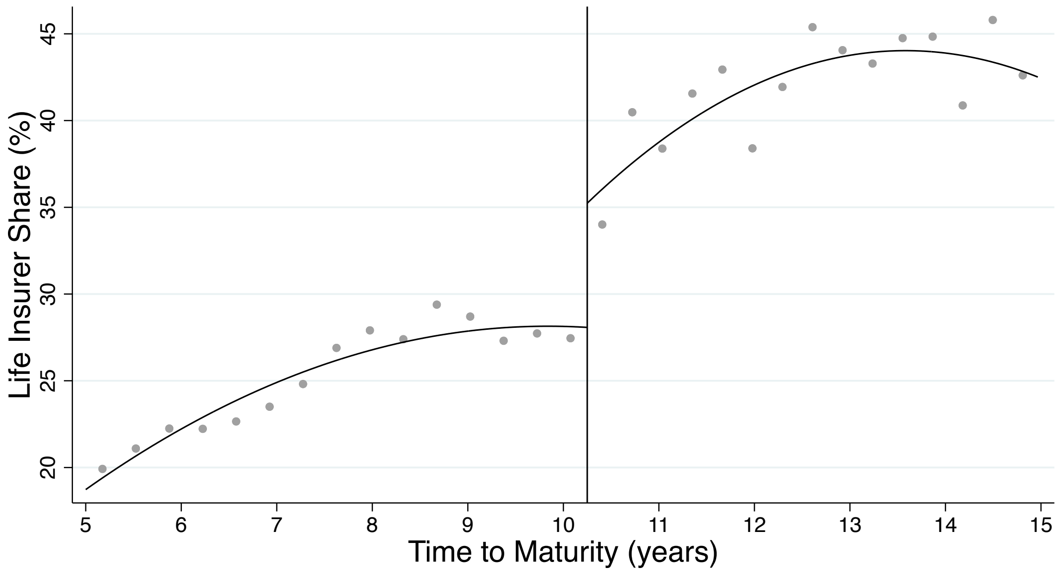

Figure 8. **Discontinuity in Investor Composition.**

This figure shows a bin scatter plot of corporate bonds' life insurance ownership share (%) and their maturity. The vertical line indicates the maturity threshold of 10.25 years. The figure also shows the best-fit second-order polynomials for observations on each side of the threshold.

Table 4 summarizes the effect of the maturity threshold on life insurer shares estimated using different methods. All results imply a strong discontinuity in investor composition at the maturity cutoff. The life insurers' ownership share is around 4-5% higher for bonds whose maturity is slightly above the cutoff.

Method	β	p-value	[95% Conf. Interval]
OLS	4.36	0.000	[3.81, 4.91]
RDD, Conventional	4.41	0.000	[2.45, 6.36]
RDD, Bias-corrected	4.58	0.000	[2.63, 6.54]
RDD, Bias-corrected, Robust	4.58	0.000	[2.31, 6.86]

Table 4. **Testing the Discontinuity.**

This table shows the coefficient on $\mathbf{1}_{\{maturity_{it}>c\}}$, controlling for the trading volume, duration, maturity, and size for each bond, and the recent default rate in each NAIC category. The first row shows the OLS estimate with robust standard errors, while the other rows show the estimates from various RDD methods discussed in Calonico, Cattaneo and Titiunik (2014).

I then examine the effect of life insurance ownership on the response of a bond's credit spread to the long-term interest rate. I follow a fuzzy RDD approach by instrumenting the life insurer share $\varphi_{it}^{\text{Ins}}$ with the dummy variable $\mathbf{1}_{\{\text{maturity}_{it} \geq c\}}$ for observations near the threshold (e.g., Lee and Lemieux, 2010). The identification relies on the discontinuity of the investor base around the threshold. In other words, I am now comparing the co-movement seen in bonds slightly above the threshold to those seen in bonds slightly below the threshold. Specifically, I estimate the following two-stage regression

$$\begin{aligned}\varphi_{it}^{\text{Ins}} &= \alpha + \beta \cdot \mathbf{1}_{\{\text{maturity}_{it} > c\}} + \Gamma \mathbf{X}_{it} + \varepsilon_{it} \\ y_{it} &= \Big(\beta_2 \cdot \mathbf{1}_{\{\text{NAIC 2}\}} + \beta_{3-6} \cdot \mathbf{1}_{\{\text{NAIC 3-6}\}}\Big) y_t^{(10)} + \\ &\quad \Big(\gamma_2 \cdot \mathbf{1}_{\{\text{NAIC 2}\}} + \gamma_{3-6} \cdot \mathbf{1}_{\{\text{NAIC 3-6}\}}\Big) \varphi_{it}^{\text{Ins}} y_t^{(10)} + \alpha_t + \Gamma \mathbf{X}_{it} + \varepsilon_{it}.\end{aligned} \tag{6}$$

The first stage estimates the effect of the maturity cutoff on the life insurer share $\varphi_{it}^{\text{Ins}}$. The second stage estimates the effect of the life insurer share $\varphi_{it}^{\text{Ins}}$ on the co-movement between the long-term interest rate and credit spreads, using $\mathbf{1}_{\{\text{maturity}_{it} > c\}}$ as an instrumental variable.[14] I choose a small bandwidth of 1 year on both sides of the threshold to create a segment that isolates the impacts of the discontinuity.[15]

Table 5 presents the estimates from the RDD regressions above. First, the first-stage F-stats both before and after the Financial Crisis are well above the conventional threshold for weak instruments in Stock and Yogo (2005), again confirming the validity of the discontinuity. Second, after the Financial Crisis, we obtained negative and significant coefficients on $\varphi_{it}^{\text{Ins}} y_t^{(10)}$. The results show that life insurer holdings amplify the negative co-movement between the long-term interest rate and credit spreads.[16] It is worth noting that the special role of life insurers began after the Financial Crisis, as the coefficients on $\varphi_{it}^{\text{Ins}} y_t^{(10)}$ were insignificant and positive before the Financial Crisis.

[14] I instrument $\mathbf{1}_{\{\text{NAIC } k\}} \cdot \varphi_{it}^{\text{Ins}} \cdot y_t^{(10)}$ with $\mathbf{1}_{\{\text{NAIC } k\}} \cdot \mathbf{1}_{\{\text{maturity}_{it} > c\}} \cdot y_t^{(10)}$.

[15] Table B.1 in the Appendix offers additional robustness checks with alternative bandwidth choices.

[16] A potential concern is that the results might mostly be driven by new bonds that have different maturities than existing ones. Table B.2 shows the results still hold when we exclude bonds issued less than 3 months ago. Another threat to identification is that firms might strategically issue bonds with a maturity of less than 10 years to attract mutual fund investors. Table B.3 shows the results still hold when we only include bonds whose maturity at issuance is at least 10 years.

	2000-2007	2010-2019
γ_2	0.141	-0.452***
	[0.662]	[-4.745]
γ_{3-6}	0.488	-0.813**
	[0.834]	[-3.229]
Time FE	✓	✓
R^2	.624	.664
First-stage F	173.5	111.2
# Identifying Bonds	1466	3774

Table 5. **RDD Regressions.**

This table shows the coefficients estimated from regression (6), with the instrument, controlling for the trading volume, duration, maturity, size, coupon amount, and coupon frequency for each bond, and the recent default rate for each NAIC category. The first column shows the result for the pre-crisis sample, while the second column shows the result for the post-crisis sample. The t-statistics shown in brackets are based on standard errors clustered at the issuer and year-month levels. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

4.4 Bond Transactions

So far, I have shown that declines in the long-term interest rate have a larger impact on bonds held by life insurers after the Financial Crisis. The main hypothesis is that life insurers' risk-bearing capacity becomes more restricted when lower long-term interest rates erode their equity (see Section 4.1). To further verify the mechanism, I examine the bond transaction pattern of life insurers following movements in long-term interest rates. Specifically, I expect life insurers to rebalance their bond portfolio towards riskier bonds after the long-term interest rate increases.

I use the following empirical specification to test the hypothesis that life insurers rebalance towards safer bonds after declines in the long-term interest rate,

$$\frac{\text{Net Purchase}_{k,t}^{(m)}}{\text{Market Equity}_t} = \alpha_t + \alpha_k + \sum_{k=2}^{6} \beta_k \cdot \mathbf{1}\{\text{NAIC } k\} \cdot \Delta y_t^{(10)} + \Gamma \mathbf{X}_{kt} + \varepsilon_{kt}. \tag{7}$$

Here Net Purchase$_{k,t}^{(m)}$ is all life insurers' net purchases[17] of NAIC k bonds in a m-month period after month t. The outcome variable is normalized by the aggregate month-end

[17] Net purchases are defined as the total new purchases subtracting the total disposals including bonds sold and matured, using all transactions recorded in NAIC regulatory filings (Schedule D).

market equity of all US life insurers. On the right-hand side, $\Delta y_t^{(10)}$ is the change in the 10-year Treasury yield during month t. After controlling for year-month fixed effects, the coefficients β_k measure the purchases of NAIC k bonds *relative* to NAIC 1 bonds, following a 1% increase in the 10-year Treasury yield. I also include NAIC fixed effects to control market size differences across credit ratings.

Table 6 shows the estimated coefficients from regression (7) for the post-crisis period. After an increase in the 10-year Treasury yield, life insurers purchase more bonds from NAIC 2-5 relative to NAIC 1, tilting their bond portfolio towards riskier segments of the corporate bond market. The effects are large and significant except for NAIC 6.[18] After a 1% increase in the 10-year Treasury yield, life insurers purchase more lower-rating bonds than NAIC 1 bonds in the following 6 months, and the difference amounts to more than 0.7% of their market equity for bonds in NAIC 2 and more than 2% for bonds in NAIC 3, 4, and 5.

	1 month	3 months	6 months
β_2	0.331***	0.0647	0.723***
	[526.5]	[1.687]	[36.11]
β_3	0.423***	0.988***	2.256***
	[20.17]	[20.44]	[14.92]
β_4	0.522***	1.120***	2.297***
	[25.04]	[27.38]	[14.63]
β_5	0.553***	1.217***	2.433***
	[21.34]	[20.22]	[13.13]
β_6	0.224	-0.225	0.327
	[1.323]	[-0.302]	[0.288]
NAIC FE	✓	✓	✓
Time FE	✓	✓	✓
R^2	.851	.894	.908

Table 6. **Insurer Bond Transactions.**

This table shows the coefficients estimated from regression (7), controlling for the recent default rate for each NAIC category. The three columns show the results for purchases in different time horizons. The t-statistics shown in brackets are based on standard errors clustered at the NAIC category level. $^*p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

[18] NAIC 6 (bonds rated CC and lower) is the smallest segment of the corporate bond market where trading is relatively infrequent. Life insurers also own relatively fewer NAIC 6 bonds.

5 Real Effects of the Long-term Interest Rate

A natural question is whether the nexus between the long-term interest rate and credit spreads affects bond-issuing firms' borrowing and investment. In this section, I show that the effects of the long-term interest rate on credit spreads indeed have considerable real consequences on bond issuance and firm investment.

5.1 The Long-term Interest Rate and Bond Issuance

In Section 3, I documented that a lower long-term interest rate increases credit spreads, especially for speculative-grade bonds. Credit spreads are important indicators of firm borrowing costs — existing studies show that primary market bond prices (i.e., new bond issuance costs) are highly correlated with secondary market conditions (e.g., Coppola, 2022; Flanagan, Kedia and Zhou, 2019). Therefore, the co-movement between the long-term interest rate and corporate bond credit spreads likely affects firms' incentives to issue new bonds.

To investigate the co-movement between the long-term interest rate and corporate bond issuance, I aggregate new bond issuance for each NAIC category in Mergent FISD. I then estimate the following empirical specification

$$\frac{\text{Issuance}_{k,t}^{(6m)}}{\text{Outstanding}_{k,t}} = \alpha_k + \alpha_t + \sum_k \beta_k \cdot \mathbf{1}_{\{\text{NAIC } k\}} \cdot y_t^{(10)} + \Gamma \mathbf{X}_{kt} + \varepsilon_{kt}, \tag{8}$$

where $\text{Issuance}_{k,t}^{(6m)}$ is the total NAIC k issuance in the 6-month period between month $t+1$ and month $t+6$, $\text{Outstanding}_{k,t}$ is the total outstanding amount of NAIC k bonds at the end of the month t, $y_t^{(10)}$ is the 10-year yield at the end of the month t, and the controls $\mathbf{X}_{kt}$ include the average maturity of new issues, the average maturity of new issues interacted with the 10-year Treasury yield, and the recent default rates of each NAIC category. The rating fixed effects α_k capture the steady state differences in the growth rates of individual NAIC categories. The time fixed effects α_t capture aggregate fluctuations in bond issuance.

I consider two specifications — one with the time fixed effects α_t and one without. When α_t is not imposed, the coefficient β_k captures the effect of the 10-year Treasury yield on the *amount* of NAIC k bond issuance. The issuance of NAIC k bonds increases by a β_k fraction of the current outstanding in the next 6 months when the 10-year Treasury yield increases by 1%. When α_t is imposed, the fixed effects absorb the effect on NAIC 1 issuance, so the coefficient β_k measures how the 10-year Treasury yield affects the *difference* in issuance between NAIC k and NAIC 1. The issuance-to-outstanding ratio of NAIC k increases by β_k relative to NAIC 1 when the 10-year Treasury yield increases by 1%.[19]

Table 7 contains the estimated coefficients from regression (8). The second column reports the results after the Financial Crisis. It suggests that increases in the 10-year Treasury yield boost the issuance of NAIC 2-6 bonds relative to NAIC 1 bonds. Following a 1% increase in the long-term interest rate, the issuance of speculative-grade (NAIC 3-6) bonds grows by more than 1.5% relative to the issuance of NAIC 1 bonds. The results suggest that the long-term interest rate potentially alters the composition of the corporate bond market. Investment-grade bonds make up a larger share of total new issuance when the long-term interest rate is low. After the Financial Crisis, the US economy entered a sustained period of low interest rates accompanied by a large bond market expansion. Consistent with my findings, the post-crisis bond market expansion was concentrated in investment-grade firms (e.g., Mota, 2023).

In contrast, the issuance differences are not correlated with the long-term interest rate before the Financial Crisis. The lack of correlation is expected since the long-term interest rate affected all corporate bonds equally during that period, as demonstrated in Section 3.1.

5.2 The Long-term Interest Rate and Firm Investment

As demonstrated in the previous section, credit spread responses to long-term interest rates are crucial for long-term interest rates' transmission to bond issuance. Next, I examine whether such credit spread responses affect firms' real investment. I adopt the following empirical specification

$$\left(\frac{\text{CapEx}}{\text{Asset}}\right)_{jt} = \alpha_j + \sum_k \beta_k \cdot \mathbf{1}_{\{\text{NAIC } k\}} \cdot y_t^{(10)} + \sum_k \gamma_k \cdot \mathbf{1}_{\{\text{NAIC } k\}} \cdot \varphi_{jt}^{\text{Ins}} y_t^{(10)} + \Gamma \mathbf{X}_{jt} + \varepsilon_{jt}. \tag{9}$$

[19] The 10-year Treasury yield is highly persistent (e.g., Caporale, Gil-Alana and Yaya, 2022). The current 10-year Treasury yield has strong predictive power for the future path of the 10-year Treasury yield.

	2000-2007	2010-2019
β_2	1.418	0.782***
	[1.585]	[2.900]
β_{3-6}	0.606	1.550***
	[0.864]	[4.712]
NAIC FE	✓	✓
Time FE	✓	✓
R^2	.858	.781

Table 7. **Bond Issuance Responses to the Long-term Interest Rate.**

This table shows the coefficients estimated from regression (8), controlling for the average maturity of new issues, the average maturity of new issues interacted with the 10-year Treasury yield, and the recent default rates of each NAIC category. The first two columns show results for the post-crisis sample, while the last two columns show results for the pre-crisis sample. Both the responsible variable and the 10-year Treasury yield are in percentage points. The t-statistics shown in brackets are based on standard errors clustered at the NAIC category by year-month levels. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

The response variable is the investment rate of firm j in quarter t, measured as the firm's capital expenditure to assets ratio. I include the 10-year Treasury yield $y_t^{(10)}$ and the interaction term $\varphi_{jt}^{\text{Ins}} y_t^{(10)}$ as explanatory variables. $\varphi_{jt}^{\text{Ins}}$ is the average life insurance ownership across all bonds issued by firm j measured at the end of the previous year.[20] The coefficient β_k measures how much an NAIC k firm's investment rate reacts to the 10-year Treasury yield if the firm's bonds are not owned by life insurers at all. γ_k, in turn, measures how much life insurance ownership affects the relationship between the 10-year Treasury yield and a firm's investment rate. I also include firm characteristics and firm fixed effects as controls.

Table 8 summarizes the estimates from regression (9). In the pre-crisis sample, neither the 10-year Treasury yield nor the bond ownership by life insurers had a significant impact on firm investment. Nevertheless, the results indicate that $\beta_k < 0$ and $\gamma_k > 0$ after the Financial Crisis, suggesting that higher long-term interest rates depress investment, but life insurance ownership dampens or reverses this relationship. This is consistent with the results of Section 4 where life insurance ownership induces nega-

[20] Figure A.15 visualizes the distribution of φ_j^{Ins} across firms.

tive credit spread responses, which offset the effects of elevated long rates. Notably, I find $\gamma_{3-6} > \gamma_2 > \gamma_1$, meaning that the effect of bond ownership by life insurers on investment is the strongest for speculative-grade firms. The result is consistent with the findings of Section 3.2, which shows speculative-grade firms exhibit the largest credit spread responses to the long-term interest rate. Since $\beta_{3-6} = -0.653, \gamma_{3-6} = 0.903$, the investment rate of a speculative-grade firm could *increase* with the 10-year Treasury yield if life insurers own more than 72% of its bonds.

	2000-2007	2010-2019
β_1	-0.0427	-0.440***
	[-0.618]	[-9.625]
β_2	0.0381	-0.623***
	[0.353]	[-12.32]
β_{3-6}	0.206*	-0.653***
	[1.925]	[-8.363]
γ_1	0.0372	0.195**
	[0.350]	[2.524]
γ_2	0.185	0.340***
	[1.635]	[4.601]
γ_{3-6}	0.0566	0.903***
	[0.383]	[2.727]
Firm FE	✓	✓
Controls	✓	✓
R^2	.657	.644

Table 8. **Firm Investment Responses to the Long-term Interest Rate.**

This table shows the coefficients estimated from regression (9), controlling for the average maturity of firms' existing bonds and fixed effects for firms, size deciles, leverage deciles, and NAIC ratings. The first column shows results for the pre-crisis sample, while the second column shows results for the post-crisis sample. Both the responsible variable and the 10-year Treasury yield are in percentage points. The t-statistics shown in brackets are based on standard errors clustered at the firm level. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

6 An Intermediary Asset Pricing Model

In the previous sections, I have shown that a negative co-movement between long rates and credit spreads emerged post-crisis and that life insurers' bond holdings matter for this co-movement. In this section, I build an intermediary asset pricing model centered around life insurers to explain the observations. The model illustrates how long rates affect credit spreads through life insurers' duration mismatch and quantifies the strength of this new channel.

As in He and Krishnamurthy (2013), the risk-bearing capacity of intermediaries (life insurers) is key to equilibrium credit spreads. As in Koijen and Yogo (2023), life insurers take on leverage by issuing annuity liabilities and investing in corporate bonds. Importantly, life insurers' portfolios expose them to duration mismatch as found in Section 4.1, and their net worth increases with the long-term interest rate. The equity gains reduce life insurers' effective risk aversion and increase their risk-bearing capacity, which leads to lower credit spreads in equilibrium.

6.1 Model Setup

Investors. There are two types of corporate bond investors in the model — life insurers and preferred-habitat investors.

The *life insurer* invests its portfolio in corporate bonds and Treasuries while issuing long-term annuities to households. The duration mismatch between the insurer's assets (corporate bonds and Treasuries) and liabilities (annuities) exposes it to interest rate risk. When the interest rate on long-term Treasuries rises, the insurer's liabilities decline more than its assets, so its net worth increases. The insurer's balance sheets are as follows.

Assets	*Liabilities*
Corporate Bonds	Annuities
Treasuries	Net Worth

Preferred-habitat investors include all other investors of corporate bonds (e.g., mutual funds and pension funds). For simplicity, their demand for corporate bonds is assumed to be reduced-form functions of bond prices, in the spirit of Vayanos and Vila (2021).

Treasuries and Annuities. Time is continuous. *Treasuries* are long-term consol bonds with a geometric maturity structure. The Treasuries have a price P_t^T, coupon rate ϕ^T, and a geometric decaying rate λ^T. Specifically, each bond pays a stream of coupon payments $\phi^T \mathrm{d}t$ and has a face value that decays over time at a constant rate of λ^T. The return rate on Treasuries is

$$\mathrm{d}r_t^T = \frac{(\phi^T - \lambda^T P_t^T)\mathrm{d}t}{P_t^T} + \frac{\mathrm{d}P_t^T}{P_t^T}, \tag{10}$$

where $\phi^T \mathrm{d}t$ captures the coupon payments, $\lambda^T P_t^T \mathrm{d}t$ captures the depreciated face value, and $\mathrm{d}P_t^T$ is the capital gain from price fluctuations. Based on the geometric decaying maturity structure, the average maturity of this bond is $\tau^T = 1/\lambda^T$. Each unit of the face value is expected to last a period of length τ^T. I define the Treasury yield as[21]

$$y_t^T := \frac{\phi^T}{P_t^T} - \lambda^T.$$

The Treasury yield has an exogenous law of motion and always reverts to the "natural" level $\overline{y}^T$ at a speed of α_y. Its dynamics is given by

$$\mathrm{d}y_t^T = \alpha_y \left(y_t^T - \overline{y}^T \right) \mathrm{d}t. \tag{11}$$

Annuities are similar consol bonds with a price P_t^L, coupon rate ϕ^L, and a geometric decaying rate λ^L. Annuities have a maturity of $\tau^L = 1/\lambda^L$ and a return rate of

$$\mathrm{d}r_t^L = \frac{(\phi^L - \lambda^L P_t^L)\mathrm{d}t}{P_t^L} + \frac{\mathrm{d}P_t^L}{P_t^L}.$$

Similar to Treasuries, I define the annuity yield as $y_t^L := \phi^L / P_t^L - \lambda^L$.

[21] The yield to maturity is defined as the rate at which future payoffs are discounted and aggregated into the current bond price:

$$P_t^T = \int_t^\infty e^{-(y_t^T + \lambda^T)s} \phi^T \mathrm{d}s.$$

No-arbitrage Pricing of Annuities. In this model, both Treasuries and annuities are risk-free assets with deterministic returns, and the life insurer holds both in equilibrium. Therefore, the returns of Treasuries and annuities must be equalized,

$$\mu_t^{r,T} = \mu_t^{r,L}, \tag{12}$$

where $\mu_t^{r,T} = \mathrm{d}r_t^T/\mathrm{d}t, \mu_t^{r,L} = \mathrm{d}r_t^L/\mathrm{d}t$ are the return rates of Treasuries and annuities, respectively. This no-arbitrage condition allows us to solve for the annuity price P_t^L and yield y_t^L as functions of the Treasury yield y_t^T.

Corporate Bonds. Corporate bonds are long-term bonds with credit risk. There are N types of corporate bonds corresponding to N credit ratings. The exposure to credit risk varies across different ratings. Investors hold diversified portfolios within each rating. The return rate on rating $n \in \{1, \ldots, N\}$ bonds is

$$\mathrm{d}r_t^n = \frac{\phi^n - \lambda^n P_t^n}{P_t^n}\mathrm{d}t + \frac{\mathrm{d}P_t^n - \nu^n P_t^n \mathrm{d}J_t}{P_t^n}, \tag{13}$$

The coupon rate is denoted by ϕ^n, while the rate of decay is represented by λ^n. $\mathrm{d}J_t$ is a Poisson jump process with intensity δ that captures the bonds' credit risk.[22] When a jump is realized, the bond market enters a disrupted period, and a fraction ν^n of rating n bonds default and lose their value. As before, the corporate bond yields are defined as $y^n := \phi^n/P_t^n - \lambda^n$. Further, I define

$$\mathrm{d}r_t^n = \mu_t^{r,n}\mathrm{d}t + \sigma_t^{r,n}(\mathrm{d}J_t - \delta\mathrm{d}t),$$

so that $\mu_t^{r,n}$ is the expected return on bonds n and $\sigma_t^{r,n}$ is the overall risk exposure of bonds n.[23]

Life Insurer's Portfolio Problem. There is one representative life insurer that holds corporate bonds and Treasuries while issuing annuities to households. The quantity of annuities is assumed to be exogenous at L,[24] and the insurer chooses its holdings

[22] The credit risk is correlated across different bonds. This is consistent with the findings of Das et al. (2007) and Duffie et al. (2009), who show that corporate bond defaults in the US are highly correlated.

[23] $(\mathrm{d}J_t - \delta\mathrm{d}t)$ is a martingale since the process $\mathrm{d}J_t$ drifts upwards in expectation ($\mathbb{E}_t[\mathrm{d}J_t] = \delta\mathrm{d}t$).

[24] In practice, the quantity of life insurers' liabilities is stable as they are typically long-term products with few early withdrawals (e.g., Chodorow-Reich, Ghent and Haddad, 2021).

of corporate bonds and Treasuries. Its portfolio weight in Treasuries, annuities, and corporate bonds are denoted as $w_t^{I,T}, w_t^{I,L}$ and $\left(w_t^{I,n}\right)_{n=1}^{N}$, respectively. The insurer's net worth A_t^I evolves according to

$$\frac{\mathrm{d}A_t^I}{A_t^I} = \left[w_t^{I,T}\mu_t^{r,T} + \sum_{n=1}^{N} w_t^{I,n}\mu_t^{r,n} - w_t^{I,L}\mu_t^{r,L} - \sum_{n=1}^{N}\frac{1}{2}\left(\zeta^n w_t^{I,n}\right)^2 + \psi_t\right]\mathrm{d}t + \sum_{n=1}^{N} w_t^{I,n}\sigma_t^{r,n}\mathrm{d}J_t.$$

$w_t^{I,L}\mu_t^{r,L}$ enters the net worth evolution negatively because annuities are liabilities. The insurer incurs regulatory cost $\frac{1}{2}\sum_{n=1}^{N}\left(\zeta^n w_t^{I,n}\right)^2 A_t^I$ for holding corporate bonds, which potentially includes the costs of complying with risk-weighted leverage constraints and passing stress tests.[25] Following Koijen and Yogo (2023), the regulatory cost is assumed to be quadratic in the bond holdings, while the parameter ζ^n governs the marginal regulatory cost of holding bonds in rating n. The process $\psi_t \mathrm{d}t$ captures an exogenous process of equity injection and dividend payout, which ensures that net worth A_t^I is stationary. In particular, I assume

$$\psi_t = \psi(A_t^I - \overline{A}^I), \quad \psi < 0,$$

so the insurer pays out dividends (raises equity) at the rate of ψ when its net worth is greater (less) than the reference level $\overline{A}^I$.

The insurer has a mean-variance preference over the return on its net worth:

$$\mathbb{E}_t\left[\frac{\mathrm{d}A_t^I}{A_t^I}\right] - \frac{a}{2}\mathrm{Var}_t\left[\frac{\mathrm{d}A_t^I}{A_t^I}\right].$$

The portfolio problem can then be written as[26]

$$\max_{\{w_t^{I,n}\}_{n=1}^{N},\, w_t^{I,T},\, w_t^{I,L}} \quad \underbrace{w_t^{I,T}\mu_t^{T} + \sum_{n=1}^{N} w_t^{I,n}\mu_t^{r,n} - w_t^{I,L}\mu_t^{r,L}}_{\text{expected return}} - \underbrace{\frac{a}{2}\delta\left(\sum_{n=1}^{N} w_t^{I,n}\sigma_t^{r,n}\right)^2}_{\text{risk aversion}} - \underbrace{\frac{1}{2}\sum_{n=1}^{N}\left(\zeta^n w_t^{I,n}\right)^2}_{\text{regulatory cost}}, \tag{14}$$

$$\text{s.t.} \quad w_t^{I,T} + \sum_{n=1}^{N} w_t^{I,n} - w_t^{I,L} = 1, \tag{15}$$

[25] Life insurers also face other potential costs, such as transaction and informational costs, which are conceptually similar to the regulatory cost from the modeling perspective.

[26] It follows from the fact the $\mathrm{Var}_t(\mathrm{d}J_t) = \delta \mathrm{d}t$.

The objective function (14) can be decomposed into three components: expected portfolio return, exposure to credit risk, and additional regulatory costs. The parameter a symbolizes the degree of risk aversion. (15) is the insurer's balance sheet constraint, which simply states that the portfolio weight in assets (Treasuries and corporate bonds) minus the portfolio weight in liabilities (annuities) equals one, the portfolio weight in net worth.

Preferred-Habitat Investors. The preferred-habitat investors also participate in the corporate bond market. I denote the shares of rating n bonds held by the preferred-habitat investors as $D_t^{P,n}$. I assume that the preferred-habitat investors have the following demand functions

$$\log D_t^{P,n} = \alpha^n - \beta \log P_t^n. \tag{16}$$

In this specification, β is the price elasticity of demand, and the intercept α^n captures the average propensity to hold bonds of rating n by the preferred-habitat investors.

Firms. There are N sectors. Each sector consists of a continuum of ex-ante identical firms with a mass of one. Let K_t^n denote both the total and average capital stock of sector n firms. Each firm in sector n produces the following stream of output

$$Y_t^n \mathrm{d}t = \frac{(K_t^n)^{1-\theta}}{1-\theta} \mathrm{d}t.$$

The production function features decreasing returns to scale. Firms issue corporate bonds to finance their capital. Capital is elastically supplied at a price of one. For simplicity, I assume that the firms can freely adjust capital stock and debt quantity but are not allowed to accumulate capital. As a result, their balance sheet constraint is simply $K_t^n = P_t^n B_t^n$, i.e., the value of their assets K_t equals the value of their debt $P_t^n B_t^n$.

In normal times (i.e., when jumps $\mathrm{d}J_t$ do not realize), all firms operate normally, and no bond defaults. However, when the bond market is disrupted (i.e., when a jump arrives), a fraction ν^n of sector n firms are destroyed. The affected firms lose all their capital, default on the bonds, exit the economy, and get replaced by new firms after the market disruption is over. Therefore, as in equation (13), a diversified portfolio in rating n bonds loses a ν^n fraction of its value when a jump materializes.

Myopic firms solve a static profit maximization problem where they choose capital stock and bond supply to maximize their expected profits subject to the balance sheet constraint.

$$\max_{K_t^n, B_t^n} \quad \mathbb{E}_t \left[\mathbf{1}_{\{\text{survive}_t\}} \left(\frac{(K_t^n)^{1-\theta}}{1-\theta} - \phi^n B_t^n \right) \right] \quad \text{s.t.} \quad K_t^n = P_t^n B_t^n.$$

The firms' profits equal their outputs subtracting the coupon payments on bonds, conditional on survival. The firm problem leads to a tractable bond supply function

$$B_t^n = \left[\frac{(P_t^n)^{1-\theta}}{\phi^n} \right]^{\frac{1}{\theta}}. \tag{17}$$

The bond supply B_t^n is increasing in the bond price P_t^n, implying that firms borrow more when their debt is more valuable.

Market Clearing. Recall that the Treasury yield y_t^T follows the exogenous law of motion (11). Treasuries are supplied perfectly elastically to clear the market, given the exogenous yield.

The supplies of corporate bonds are B_t^n. The market for rating n bonds clears when

$$w_t^{I,n} A_t^I + P_t^n D_t^{P,n} = P_t^n B_t^n. \tag{18}$$

Market clearing requires that the demand from the life insurer $w_t^{I,n} A_t^I$ and the demand from preferred-habitat investors $P_t^n D_t^{P,n}$ add to the total market cap $P_t^n B_t^n$. The annuity market clears when the insurer's portfolio weight in annuities equals the exogenously fixed annuity demand.

$$w_t^{I,L} A_t^I = P_t^L L. \tag{19}$$

Equilibrium. An *equilibrium* is a mapping from the Treasury yield and insurer net worth $\{y_t^T, A_t^I\}$ to insurer portfolio weights $\{w_t^{I,n}\}$ and asset prices $\{P_t^L, P_t^n\}$ that satisfy the insurer's portfolio choice problem (14)-(15), the habitat demand function (16), the bond supply function (17), and the market clearing conditions (18)-(19).

6.2 Analytical Insights

In this section, I analyze the effects of an unexpected shock on the long-term Treasury yield on credit spreads. I focus on a simplified version of the model that permits analytical solutions and illustrates the mechanism. To this end, I make two simplifying assumptions.

Assumption 1 *The Treasury yield is constant $y_t^T \equiv y$.*

Assumption 2 *Corporate bonds are short-term bonds with independent default risks.*

Since the Treasury yield is constant, the prices of Treasuries and annuities are constant (so $\mathrm{d}P_t^T = \mathrm{d}P_t^L = 0$). From equations (10) and (12), the return rates on Treasuries and annuities are simply

$$\mu_t^{r,T} = \mu_t^{r,L} = y.$$

The prices of Treasuries and annuities are given by the following Gordon growth formulas,

$$P_t^T = \frac{\phi^T}{y + (1/\tau^T)}, \quad P_t^L = \frac{\phi^L}{y + (1/\tau^L)}. \tag{20}$$

Corresponding to the standard Gordon formula,[27] the current dividend rate of the Treasuries is ϕ^T, the required return rate is y, and the dividend growth rate is minus the depreciation rate $-\lambda^T = -1/\tau^T$.

When the government (unexpectedly) raises the Treasury yield, it also raises the annuity yield of the same magnitude. According to (20), the increase in y lowers the price of annuities P_t^L, making life insurers' liabilities less expensive. The results are intuitive — for the same stream of coupon payments, the price of the annuities must decline to be consistent with a higher equilibrium yield.

[27]The Gordon growth formula gives the price of an asset whose dividend growth rate and required return rate are constant:

$$\text{Price} = \frac{D}{r - g} = \frac{\text{Dividend Rate}}{\text{Required Return Rate} - \text{Dividend Growth Rate}}.$$

Next, I consider a scenario where the government unexpectedly increases the Treasury yield from y to $\widehat{y} > y$, focusing on how the Treasury yield shock affects the life insurer's net worth and bond prices.

When the Treasury yield changes, long-term assets, such as the Treasuries and annuities, are repriced according to the Gordon formula (20). I denote the price of Treasuries before and after the shock as P^T and $\widehat{P}^T$, respectively. Similarly, the price of annuities before and after the shock is written as P^L and $\widehat{P}^L$. As a result, the life insurer's net worth could also change. Denote the insurers' net worth before and after the shock as A and $\widehat{A}$, respectively. The change in net worth follows

$$\begin{aligned}\widehat{A} - A &= \overbrace{T\left(\widehat{P}^T - P^T\right)}^{\text{Repricing of Treasuries}} - \overbrace{L\left(\widehat{P}^L - P^L\right)}^{\text{Repricing of Annuities}} \\ &= T\left[\frac{\phi^T}{\widehat{y} + (1/\tau^T)} - \frac{\phi^T}{y + (1/\tau^T)}\right] - L\left[\frac{\phi^L}{\widehat{y} + (1/\tau^L)} - \frac{\phi^L}{y + (1/\tau^L)}\right]. \qquad (21)\end{aligned}$$

Here $T := Aw^{I,T}/P^T$ is the quantity of Treasuries held by the insurer before the shock. A higher Treasury yield depresses the values of Treasuries and annuities ($\widehat{P}^T < P^T, \widehat{P}^L < P^L$). The change in the insurer's net worth, $\widehat{A} - A$, is determined by the extent to which Treasuries and annuities are repriced. Consider a simple case where the Treasury and the annuity have the same coupon rate ($\phi^T = \phi^L = \phi$), where

$$\widehat{A} - A = \phi(\widehat{y} - y)\left[\frac{L}{(\widehat{y} + (1/\tau^L))(y + (1/\tau^L))} - \frac{T}{(\widehat{y} + (1/\tau^T))(y + (1/\tau^T))}\right].$$

Since, in the scenario of interest, annuities have a longer maturity than Treasuries ($\tau^L > \tau^L$) and the insurer issues more annuities than the Treasuries it holds ($L > T$), then the insurer's net worth increases with the Treasury yield (i.e., $\widehat{A} > A$), which corresponds to the situation after the Financial Crisis (Section 4.1).

Since the corporate bonds are short-term with independent credit risks, their return rates (13) simply become

$$\mathrm{d}r_t^n = \frac{\phi^n}{P_t^n}\mathrm{d}t - \nu^n \mathrm{d}J_t^n,$$

where $\{\mathrm{d}J_t^n\}$ are independent from each other. To understand how the insurer's net worth affects bond credit spreads, it is useful to characterize the solution to the portfolio problem (14)-(15). In this simplified model, the first-order condition for rating n bonds is given by

$$\underbrace{\mu_t^{r,n} - \mu_t^T}_{\text{credit spread}} = \underbrace{a\delta w_t^{I,n} \nu^n}_{\text{risk premium}} + \underbrace{\zeta^n w_t^{I,n}}_{\text{regulation premium}} = \gamma^n w_t^{I,n}, \tag{22}$$

where $\mu_t^{r,n} := \mathbb{E}_t[\mathrm{d}r_t^n]/\mathrm{d}t = \phi^n / P_t^n - \nu^n \delta$ is the expected return on rating n bonds and $\gamma^n := a\delta\nu^n + \zeta^n > 0$ measures the total cost of holding rating n bonds for the insurer. We can write equation (22) as

$$w_t^{I,n} = \frac{\mu_t^{r,n} - \mu_t^T}{\gamma^n}, \tag{23}$$

which is the standard portfolio choice condition under mean-variance preferences. $w_t^{I,n}$, the insurer's portfolio weight in rating n bonds, equals the risk premium $\mu_t^{r,n} - \mu_t^T$ divided by the cost parameter γ^n.[28]

Equation (23) allows us to express $A_t^I w_t^{I,n}$, the (dollar) amount of corporate rating n bonds demanded by the life insurer, as

$$A_t^I w_t^{I,n} = \frac{\mu_t^{r,n} - \mu_t^T}{\gamma^n / A_t^I}.$$

The demand is given by the credit spread divided by the insurer's *effective risk aversion*, which I define as γ^n / A_t^I, the total holding cost γ^n divided by net worth A_t^I. Under duration mismatch, the increase in the Treasury yield boosts the insurer's net worth, lowering the insurer's effective risk aversion. As a result, the insurer increases its demand for risky bonds, putting downward pressure on equilibrium credit spreads.

I make two further assumptions before presenting the analytical results. The first assumption states that the life insurer is subject to a duration mismatch of the same kind found in Section 4.1. In Appendix D.1, I provide a sufficient condition for this assumption to hold in equilibrium. The second assumption is that the preferred-habitat demand is downward-sloping. The assumption is standard in the literature of institutional bond demand and supported by various empirical estimates (Bretscher et al., 2022; Darmouni, Siani and Xiao, 2022).

[28] In the absence of the regulatory cost ζ^n, the parameter $\gamma^n = a\delta\nu^n$ equals the risk aversion coefficient a times the default risk $\delta\nu^n$, which is exactly the solution of a standard mean-variance portfolio problem.

Assumption 3 *The life insurer is subject to duration mismatch (i.e.,* $\frac{\partial \widehat{A}}{\partial \widehat{y}} > 0$ *for* $\widehat{y} \geq 0$*).*

Assumption 4 *The preferred-habitat demand functions are downward-sloping (i.e.,* $\beta \leq 0$*).*

The next Proposition summarizes the main analytical results from this model.

Proposition 1 *Under assumptions 1-4, the following predictions hold when the Treasury yield increases from* y *to* $\widehat{y} > y$*:*

1. *the insurer's net worth increases*
2. *the insurer's exposure to credit risk* $w_t^{I,n}$ *declines for all* $n > 1$
3. *the credit spread* $\mu_t^{r,n} - \mu_t^{r,T}$ *declines for all* $n > 0$

In addition, the magnitudes of predictions 2-3 increase in the duration mismatch (i.e., $\frac{\partial \widehat{A}}{\partial \widehat{y}}$*).*

Proof. *See Appendix D.2.*

Proposition 1 shows that when the yield on long-term Treasuries rises, the insurer's balance sheet net worth rises as it faces duration mismatch. The insurer's portfolio becomes less concentrated in risky bonds, and the insurer is less exposed to the credit risks and regulatory burdens of corporate bonds. Consequently, the equilibrium credit spreads decline. Moreover, the impact of a Treasury yield shock is larger when the life insurer's duration mismatch is more severe, and its net worth is more sensitive to the Treasury yield.

6.3 Quantitative Model

In the following sections, I match the model to empirical estimates and key moments in data in order to quantify the contribution of life insurers' duration mismatch to the observed empirical patterns and the transmission of unconventional policy.

Calibrated Parameters. Table 9 discusses model calibration. I consider two corporate bond ratings ($N = 2$) where $n = 1$ represents investment-grade bonds (NAIC 1-2), and $n = 2$ represents speculative-grade bonds (NAIC 3-6). Both types of corporate bonds have a maturity of 8.55 years, which is the average time to maturity of all corporate bonds in Mergent FISD. The parameter of firm production function θ is set to 0.34, matching the usual Cobb-Douglas capital share in the literature.[29]

I normalize all the coupon rates and the life insurer's reference net worth as 1 ($\phi^T = \phi^L = \phi^n = 1$). I set $\tau^T = 10$ and $\tau^L = 20$, so the maturity of Treasuries is 10 years, and the maturity of annuities is 20 years. I let the steady-state value of the 10-year Treasury yield be 2.5%, which is roughly the average observed 10-year US Treasury yield between 2010 to 2020. I set the speed of mean-reversion as $\alpha_y = \psi = 2$, in which case the half-life of Treasury yield shocks is about 3 quarters. I normalize the insurer's reference net worth to $\overline{A}^I = 1$. I use a standard value of risk aversion $a = 2$.

Parameter	Target
Treasuries and Annuities	
$\tau^T = 10$, $\tau^L = 20$	Treasury maturity = 10 yrs, Annuity maturity = 20 yrs
$\phi^T = \phi^L = 1$	Normalization
$\overline{y}^T = 2.5\%$	Average 10-year Treasury yield (2010-2020)
$\alpha_y = \psi = -2$	Half-life of shocks $\approx$ 3 qtrs
Corporate Bonds	
$\tau^1 = \tau^2 = 8.55$	Corporate bond maturity = 8.55 yrs
$\phi^1 = \phi^2 = 1$	Normalization
$\delta = 1.635$	Variance of speculative-grade default rates
$\theta = 0.34$	Standard Cobb-Douglas capital share ($1 - \theta = 0.66$)
Life Insurer	
$\overline{A}^I = 1$	Normalization
$a = 2$	Standard

Table 9. **Calibrated Parameters.**

[29] To map the production function into a Cobb-Douglas form, we can assume that every firm has a single unit of labor supply $L_t^n = 1$ and the production function is $Y_t^n = \frac{1}{1-\theta}(K_t^n)^{1-\theta}(L_t^n)^{\theta}$.

Estimated Parameters. I estimate a few other key parameters using empirical data. The estimated parameters are summarized in Table 10.

Parameter	Estimation Strategy
Corporate Bonds	
$\nu^1 = 0.001,\ \nu^2 = 0.017$	Average default rates
$\delta = 1.635$	Variance of speculative-grade default rates
Life Insurer	
$\zeta^1 = 2.85$	Relative portfolio share $(w^{I,2}/w^{I,1})^{ss} = 0.059$
$\zeta^2 = 7.21$	Relative bond supply $(P^2B^2/P^1B^1)^{ss} = 0.195$
$L = 8.79$	Empirical duration mismatch
Habitat Investor	
$\beta = 1.106$	Mutual fund demand elasticity (Darmouni et al., 2023)
$\alpha^1 = 0.47,\ \alpha^2 = 0.12$	Life insurers' share in each category (35.7%, 10.3%)

Table 10. **Estimated Parameters.**

I estimate the loadings on credit risk ν^1, ν^2 and the intensity of the credit risk process δ from the average annual default rates of both investment-grade and speculative bonds and the variance of speculative-grade default rate. The average one-year default rate between 2003 and 2019 is 0.156% for investment-grade bonds and 2.829% for speculative-grade bonds. The standard deviation of the one-year investment-grade bond default rate is 0.004. In the model, these three moments are given by $\delta\nu^1, \delta\nu^2$ and $\sqrt{\delta}\nu^1$. I estimate values of ν^1, ν^2, δ from the data by equating the model moments to the empirical counterparts.

I estimate regulatory cost parameters ζ^1, ζ^2 using two moments: the insurer's relative portfolio share investment-grade and speculative-grade bonds and the relative market cap of the two types of bonds (see Table B.4). I choose the values of ζ^1, ζ^2 so the model replicates the two empirical moments in the steady state.

The annuities supply L is obtained from the empirically estimated duration mismatch (see Section 4.1). Under my parametrization, the insurer's duration mismatch is more severe when they are more levered (i.e., larger L). The L parameter is chosen such that the insurer's net worth increases by 7.18% in response to a 1% positive Treasury yield shock starting from the steady state.

For the habitat investors' demand elasticity β, I adopt the demand elasticity of mutual funds estimated by Darmouni, Siani and Xiao (2022). I then estimate the demand intercepts α^1, α^2 using the market share of life insurers in each risk category at the end of 2010.[30] In the steady state, the insurer owns 35.7% of investment-grade bonds and 10.3% speculative-grade bonds, which match the empirical observation.

Model Performance. The model is successful at matching a number of important untargeted empirical responses. I consider an unexpected 1% positive shock to the 10-year Treasury yield from the steady state. Table 11 summarizes both the model-implied and empirical responses of three key variables relating to credit spreads, the insurer's portfolio adjustments, and bond issuance.[31] The model generates a 0.92% reduction in the credit spread between investment-grade bonds and speculative-grade on impact, which amounts to 86% of the empirical estimate. The model also implies that the insurer increases its portfolio weight in speculative-grade bonds by 1.06% relative to investment-grade bonds in a quarter after the shock, matching the 1.00% empirical counterpart. In the model, the one-quarter issuance of speculative-grade bonds outpaces that of investment-grade bonds by 0.92%, which is reasonably close to the empirical estimate of 0.63%.

Definition	Description	Model	Empirical
$\Delta(y_t^2 - y_t^1)\|_{t=0}$	Immediate Credit Spread Response	-0.92%	-1.06%
$\left(\Delta w_t^{I,2} - \Delta w_t^{I,1}\right)\|_{t=0.25}$	1-qtr Portfolio Adjustment	1.06%	1.00%
$\left(\frac{\Delta B_t^2}{B^{2,ss}} - \frac{\Delta B_t^1}{B^{1,ss}}\right)\|_{t=0.25}$	1-qtr Bond Issuance Response	0.92%	0.63%

Table 11. **Untargeted Responses.**

[30] In Section 4.2, I show that the co-movement of interest only exists in bonds held by life insurers. Therefore, I calibrate α^n to those bonds with life insurance ownership (see column 2 of Table B.4).

[31] See Table B.5 for details on the empirical estimates.

To understand these responses intuitively, it is useful to examine the insurer's trading and the induced bond price changes. In response to the 1% positive Treasury yield shock, the insurer lowers its portfolio weight in investment-grade bonds by 1.06% and leaves its portfolio weight in speculative-grade bonds largely unchanged by $t = 0.25$. Again, the difference in portfolio adjustments is because higher interest rates boost the insurer's net worth and its demand for risky bonds relative to safe bonds. The insurer sells investment-grade bonds that amount to 0.48% of the total steady-state outstanding. In contrast, it purchases speculative-grade bonds that are 0.84% of the total steady-state outstanding. The trading then lowers the price of investment-grade bonds by about 0.2% and increases the price of speculative-grade bonds by about 0.69%. The price changes trigger further supply responses, inducing less investment-grade bond supply and greater speculative-grade bond supply. Finally, the price movements translate to a 0.25% increase in the investment-grade yield and a 0.66% decline in the speculative-grade yield, thus generating a 0.91% decline in the credit spread.[32]

6.4 The Duration Mismatch Channel and Policy Implications

The model has several new implications for unconventional monetary policy. In recent years, the Federal Reserve has adopted policies aiming to control long-term interest rates (e.g., Quantitative Easing and Tightening). For example, since 2022, the Federal Reserve has conducted Quantitative Tightening (QT) to shrink its balance sheets and control inflation, which increases the long-term interest rate.

Next, I investigate the model implications of a positive shock to the 10-year Treasury yield, which I view as a result of a QT policy.[33] The goal of this section is to quantitatively examine the policy's impacts on corporate bond yields, spreads, and issuance, as

[32]Under my parametrization, a 1% corporate bond price change generates yield movements between 0.75% and 1.25%. The result suggests that the bonds' have a duration close to 1, which is consistent with the empirical price-yield sensitivity from observed market transactions (shown in Appendix C), albeit less than the usual measure of average bond duration. This is because I target the duration of shorter-term bonds, which are traded more frequently and have larger weights in the bond price dataset and my empirical results.

[33]In this paper, I view the unexpected Treasury yield shock as the end product of a QT policy and do not model the details regarding the implementation of the QT policy. See, for example, Krishnamurthy and Vissing-Jorgensen (2011), D'Amico and King (2013) and Vayanos and Vila (2021) for mechanisms of how unconventional monetary policies influence long-term interest rates.

well as the role of the duration mismatch channel in the transmission.

Specifically, I analyze the responses of key model variables to an unexpected shock that moves the 10-year Treasury yield from $y_t^T = 2.5\%$ to $y_t^T = 3.5\%$. In the following analysis, I study a situation where the default shocks J_t are not realized, so the yield change is the only shock to the system.

In particular, I consider two scenarios: (1) the full model and (2) the model with no duration mismatch. In the full model, the insurer's duration mismatch is calibrated to the empirical estimate of Section 4.1, meaning that the insurer's net worth increases by 7.18% as the Treasury yield increases by 1%. In the model with no duration mismatch, I assume that the equity injection and dividend payout process $\psi_t \mathrm{d}t$ is chosen such that the insurer's net worth A_t^I always stays constant. In this case, the net worth does not respond to the Treasury yield shock. In the broad context, we can use the full model to represent the post-crisis scenario and the model without duration mismatch to represent the pre-crisis scenario. By contrasting these two scenarios, I aim to quantitatively assess the consequences of the duration mismatch channel on the transmission of long-term interest rates.

Figure 9 visualizes the shock and the response of insurer net worth A_t^I. The left panel plots the path of the Treasury yield, which jumps from 2.5% to 3.5% at $t = 0$ and gradually reverts back to the steady state level 2.5%. The right panel shows the change in the insurer's net worth as a percentage of its steady-state value for the two different scenarios. In the full model, the insurer's net worth rises on impact and slowly returns to the steady state value. In the model without duration mismatch, the insurer's net worth stays constant.

Figure 10 plots the responses of corporate bond yields relative to the steady state. In the model without duration mismatch, yields of both bonds 1 and bonds 2 increase with the Treasury yield. This is due to the standard portfolio rebalancing channel — the insurer reduces demand for corporate bonds as Treasuries become more attractive. In addition, higher interest rates raise risk premia, as the credit spread between bonds 1 and bonds 2 widens as the Treasury yield increases. This is consistent with the empirical results of Figure 3, where higher long-term interest rates weakly increased the credit spreads for NAIC 2 and NAIC 3 bonds before the Crisis.

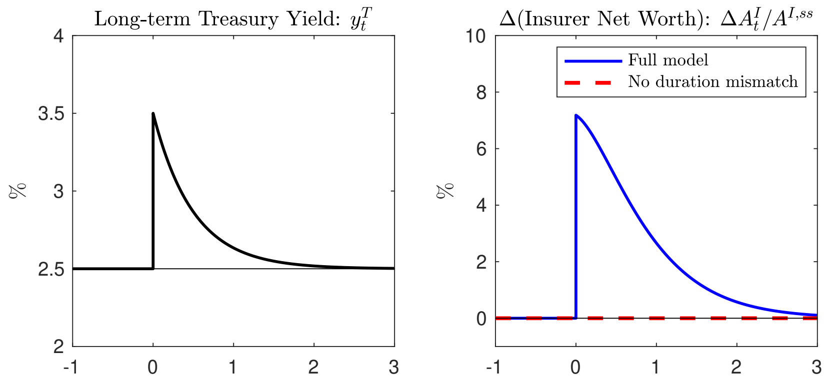

Figure 9. **Treasury Yield Shock and the Insurer's Net Worth Responses.**

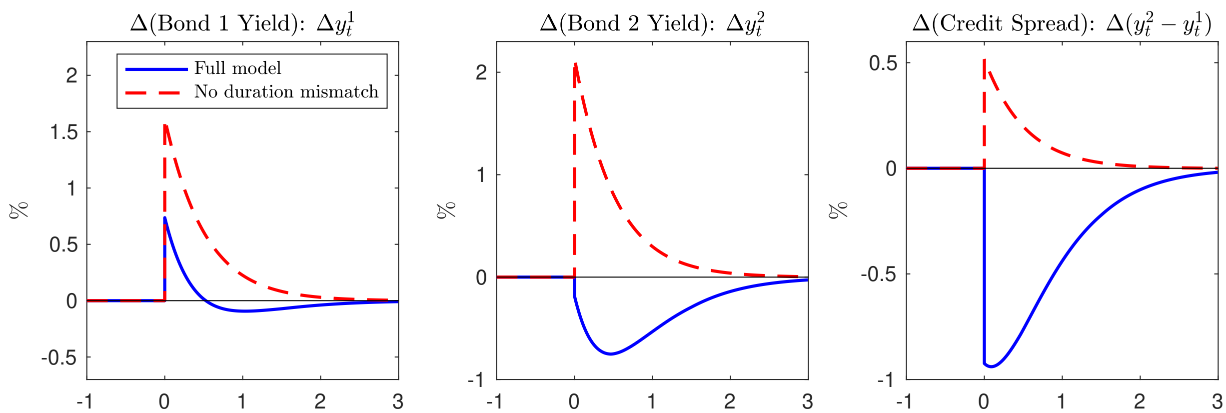

Figure 10. **Responses of Bond Yields and Credit Spreads.**

In the full model, there is an additional duration mismatch channel, where a higher Treasury yield increases the insurer's net worth, thereby boosting its demand for corporate bonds. Quantitatively, on impact, the duration mismatch channel dampens the investment-grade bond yield response by half while reversing the speculative-grade bond yield response slightly. Over time, the dampening and reversing effects become stronger before the yields revert to the steady state. Further, the duration mismatch channel has a larger effect on the speculative-grade yield, and the credit spread between the two bonds falls sharply as the Treasury yield rises. Thus, the QT policy unintentionally tightens the credit spread of corporate bonds. Notably, we can only generate large negative credit spread responses that are in line with Section 3.2 when the insurer faces duration mismatch.

I then study whether such credit spread responses can dampen or reverse the transmission of QT to real outcomes, including bond issuance and firm investment.

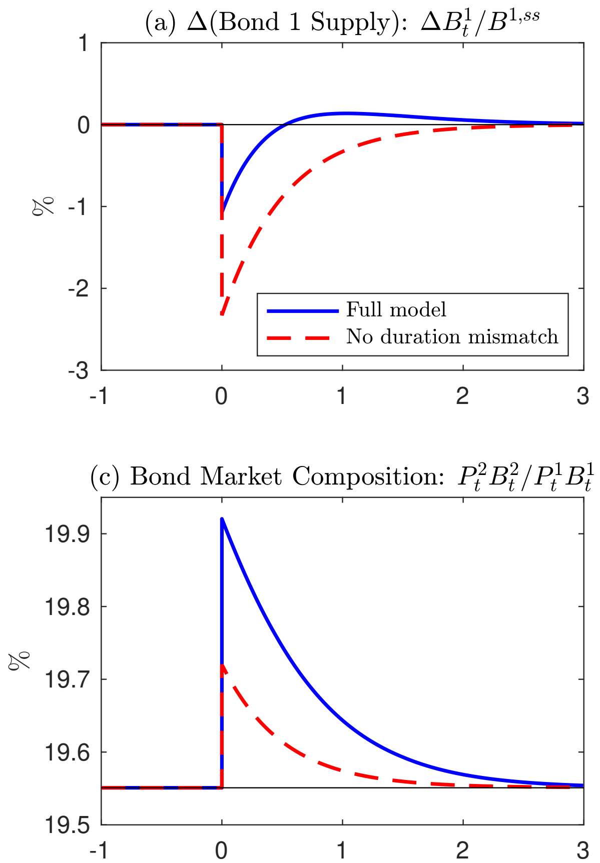

Figure 11. **Responses of Bond Supply and Investment.**

Panels (a) and (b) of Figure 11 plot the responses of the supply of bonds 1 and 2. In the model without duration mismatch, the surge in the Treasury yield drives up corporate bond yields. In response, firms borrow less, showing that the policy is effective at cooling the bond market, at least when the life insurer is not subject to duration mismatch. In the full model, the duration mismatch channel counteracts the increase in corporate bond yields. Quantitatively, the results demonstrate that the duration mismatch channel can offset half of the contraction in investment-grade bond supply and even generate an unintended expansionary effect on speculative-grade bond supply.

Panel (c) of Figure 11 shows the relative market cap of bond 2 relative to bond 1. The QT policy tilts the market towards the speculative-grade segment, even more so in the full model, where credit spreads fall in response to the positive long-term interest rate shock. The results demonstrate that QT potentially alters the composition of the bond market, favoring risky issuers over safe ones.

Panel (d) of Figure 11 displays the aggregate investment response. Absent the duration mismatch channel, firms disinvest 3% of their capital following the Treasury yield increase. The investment response is also heavily muted in the full model, at about 1/3 of that in the model without duration mismatch.

Quantitative Easing (QE) is the opposite of QT, where the Fed purchases long-term Treasuries and lowers long-term interest rates. Empirical evidence shows that QE and central bank asset purchases can effectively lower long rates (e.g., Krishnamurthy and Vissing-Jorgensen, 2011; D'Amico and King, 2013; Vayanos and Vila, 2021), which potentially boosts economic activities. However, my results show that the effects of QE are achieved at the cost of increased corporate bond credit spreads, which offsets some of the postulated economic benefits of QE. Similar to QT, the duration mismatch channel could dampen the transmission of QE to the yields and issuance of investment-grade bonds and even reverse the effects of QE on the yields and issuance of speculative-grade bonds. Overall, the results of this section suggest that unconventional monetary policies that target the long-term interest rate could have large unintended effects in the corporate bond market due to life insurers' duration mismatch.

7 Conclusion

In this paper, I document a shift in the co-movement between the long-term interest rate and corporate bond credit spreads. In particular, declines in the long-term interest rate led to large increases in credit spreads only after the Financial Crisis. This is important given the long-run trend of the falling long-term interest rate and unconventional monetary policies that specifically target the long-term interest rate. I further establish that the new co-movement is caused by life insurers' bond holdings. Life insurers, the largest institutional investor group in the corporate bond market, comprise 20-35% of the bond market on average. After the Financial Crisis, they faced a large duration mismatch, and their equity value declined sharply with the long-term interest rate. I present causal evidence that this co-movement is more pronounced in bonds held by life insurers. This mechanism also has significant real effects on bond issuance and investment in the cross-section of firms. I build an intermediary asset pricing model with life insurers to account for the empirical findings. In the model, declines in the long-term interest rate squeeze life insurers' balance sheet equity and increase their effective risk aversion, resulting in higher equilibrium credit spreads. The model explains more than 80% of the observed co-movement between the long-term interest rate and corporate bond credit spreads as a result of the *duration mismatch channel*. The model implies large real effects and shows that life insurers' balance sheets can play an important role in the transmission of unconventional monetary policy to bond yields, issuance, and investment.

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Online Appendix

A Supplementary Figures

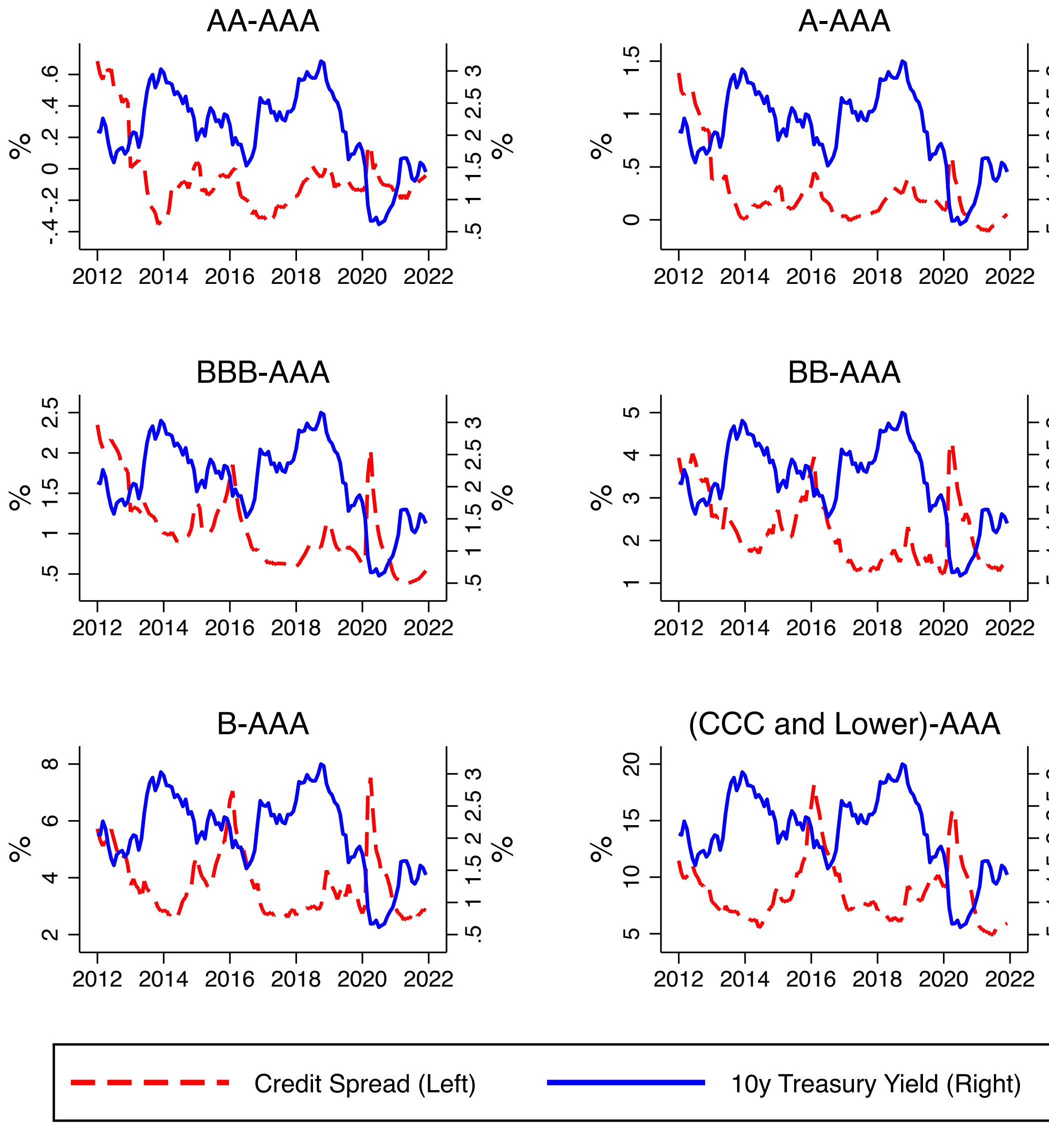

Figure A.1. **The 10-year Treasury Yield and Bond Credit Spreads (2010-2019).**
This figure plots the 10-year Treasury yield and the spreads of corporate bond indices relative to the index of AAA bonds. The 10-year Treasury yield is shown as the solid blue line and corresponds to the y-axis on the right-hand side. The credit spreads are shown as dashed red lines and correspond to the y-axis on the left-hand side.

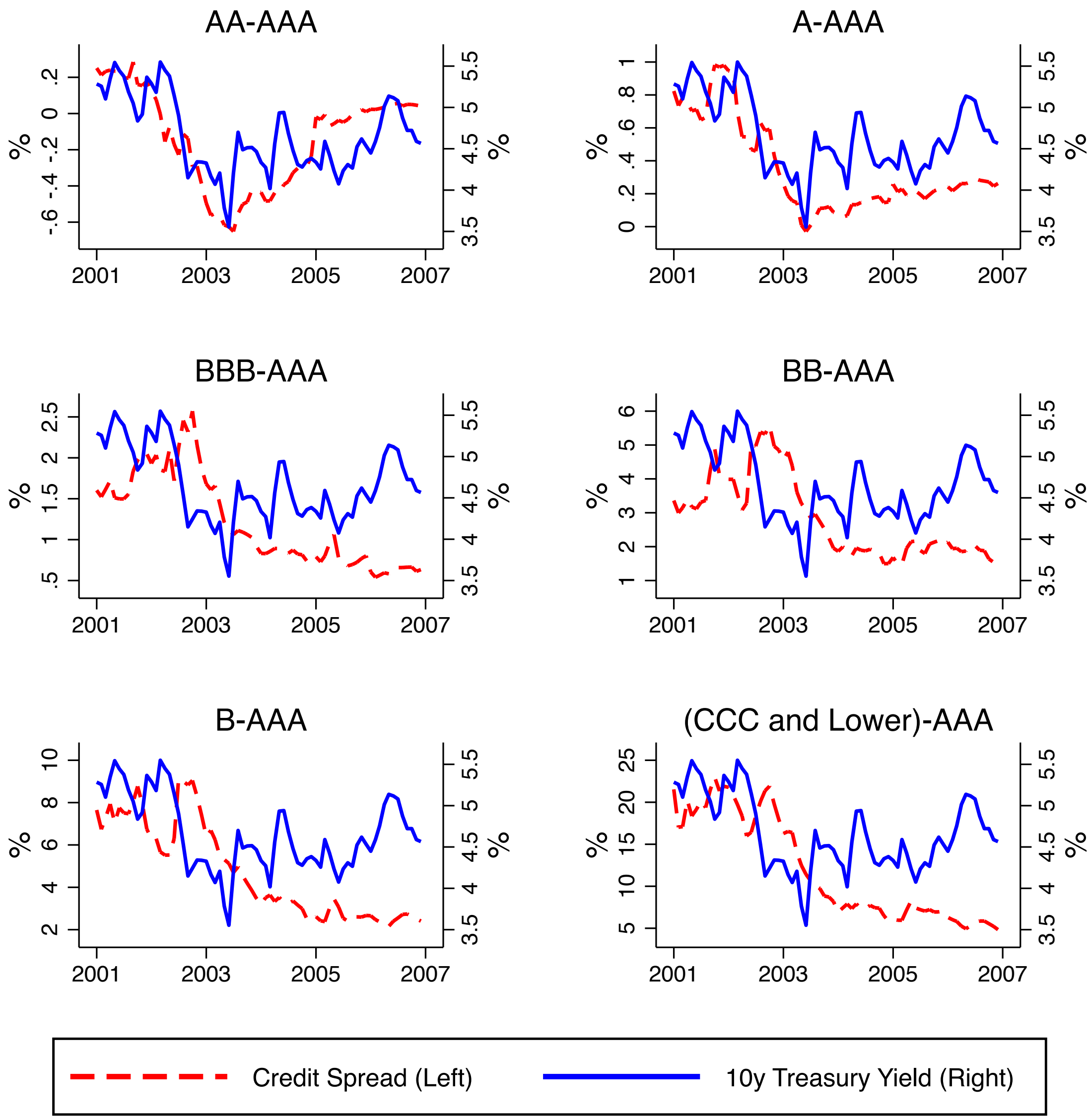

Figure A.2. **The 10-year Treasury Yield and Bond Credit Spreads (2001-2007).**

This figure plots the 10-year Treasury yield and the spreads of corporate bond indices relative to the index of AAA bonds. The 10-year Treasury yield is shown as the solid blue line and corresponds to the y-axis on the right-hand side. The credit spreads are shown as dashed red lines and correspond to the y-axis on the left-hand side.

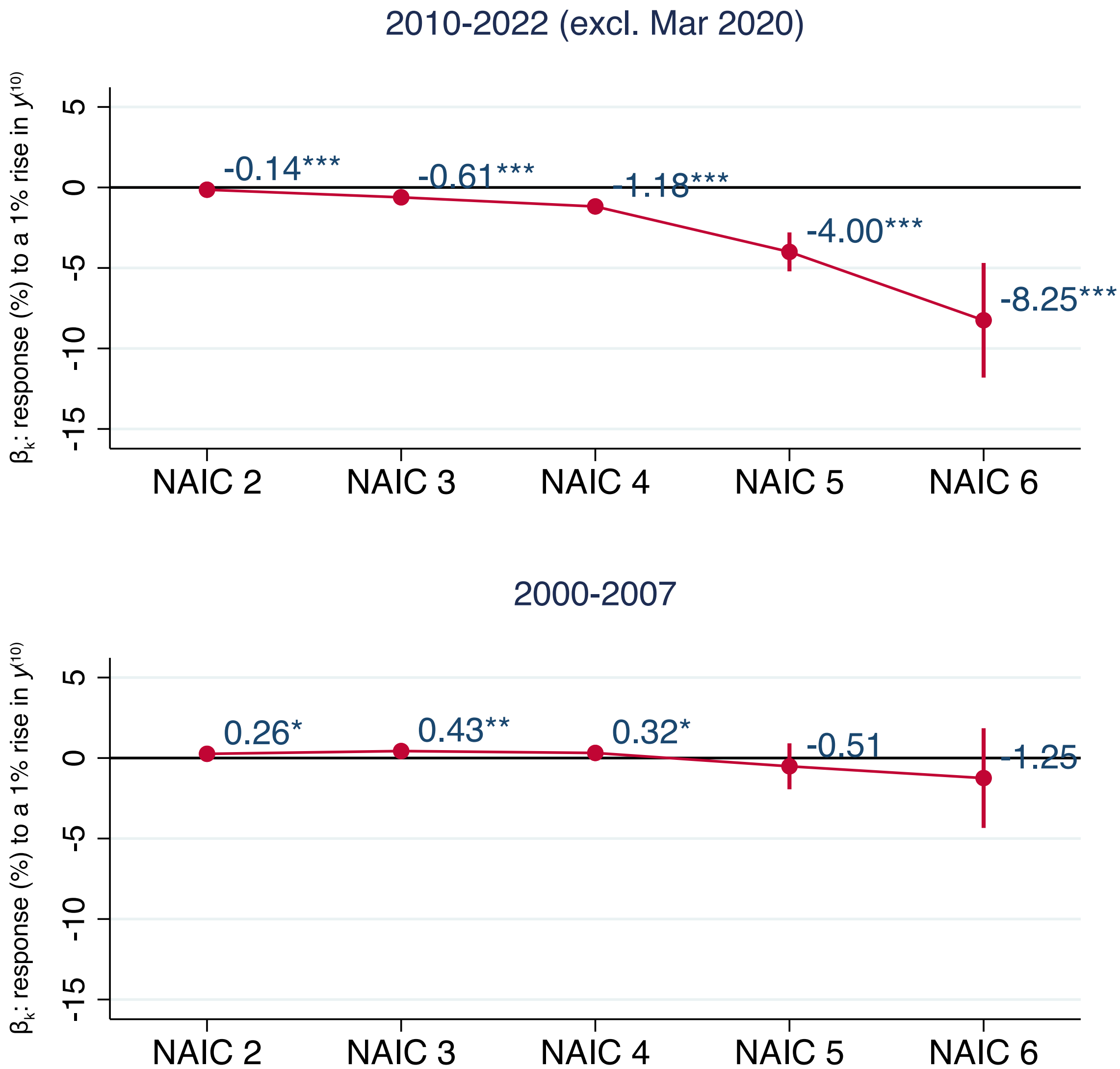

Figure A.3. **Credit Spread Responses (Controlling for CDS Spreads).**

This figure plots the coefficients β_k estimated from regression (2), controlling for the trading volume, outstanding amount, maturity, duration, credit ratings, coupon amount, and coupon frequency of each bond, and the recent default rate and average CDS spreads of each NAIC category. Both corporate bond yields and the Treasury yield are in percentage points. The top panel shows results for the post-crisis sample, while the bottom panel shows results for the pre-crisis sample. The t-statistics shown in brackets are based on standard errors clustered at the issuer and year-month levels. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

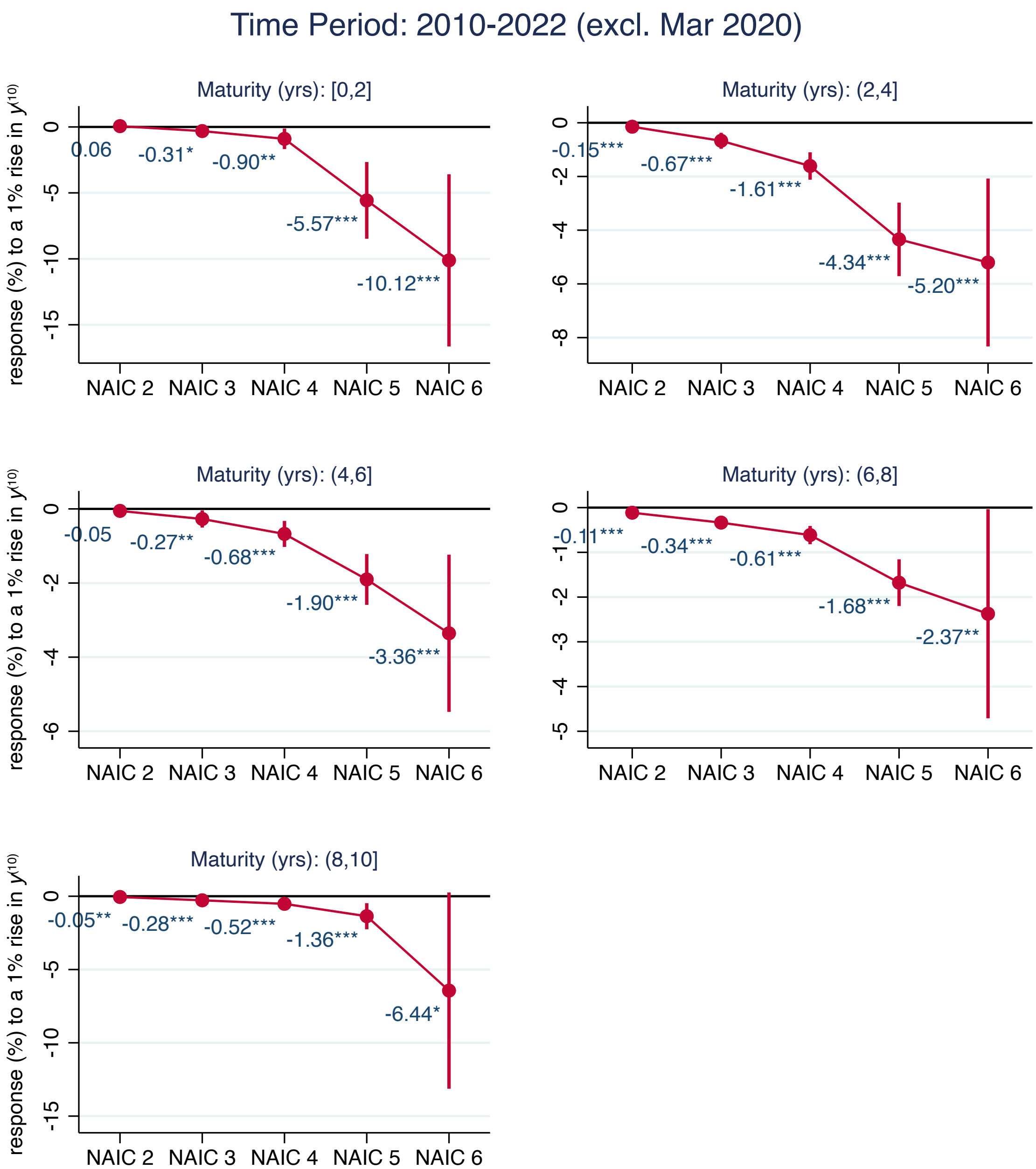

Figure A.4. **Credit Spread Responses in Different Maturity Groups (post-2008).**

This figure plots the coefficients β_k estimated from regression (2), controlling for the trading volume, outstanding amount, maturity, duration, credit ratings, coupon amount, and coupon frequency of each bond, and the recent default rate and average CDS spreads of each NAIC category. Both corporate bond yields and the Treasury yield are in percentage points. The different panels estimate coefficients for bonds in different maturity segments. The t-statistics shown in brackets are based on standard errors clustered at the issuer and year-month levels. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

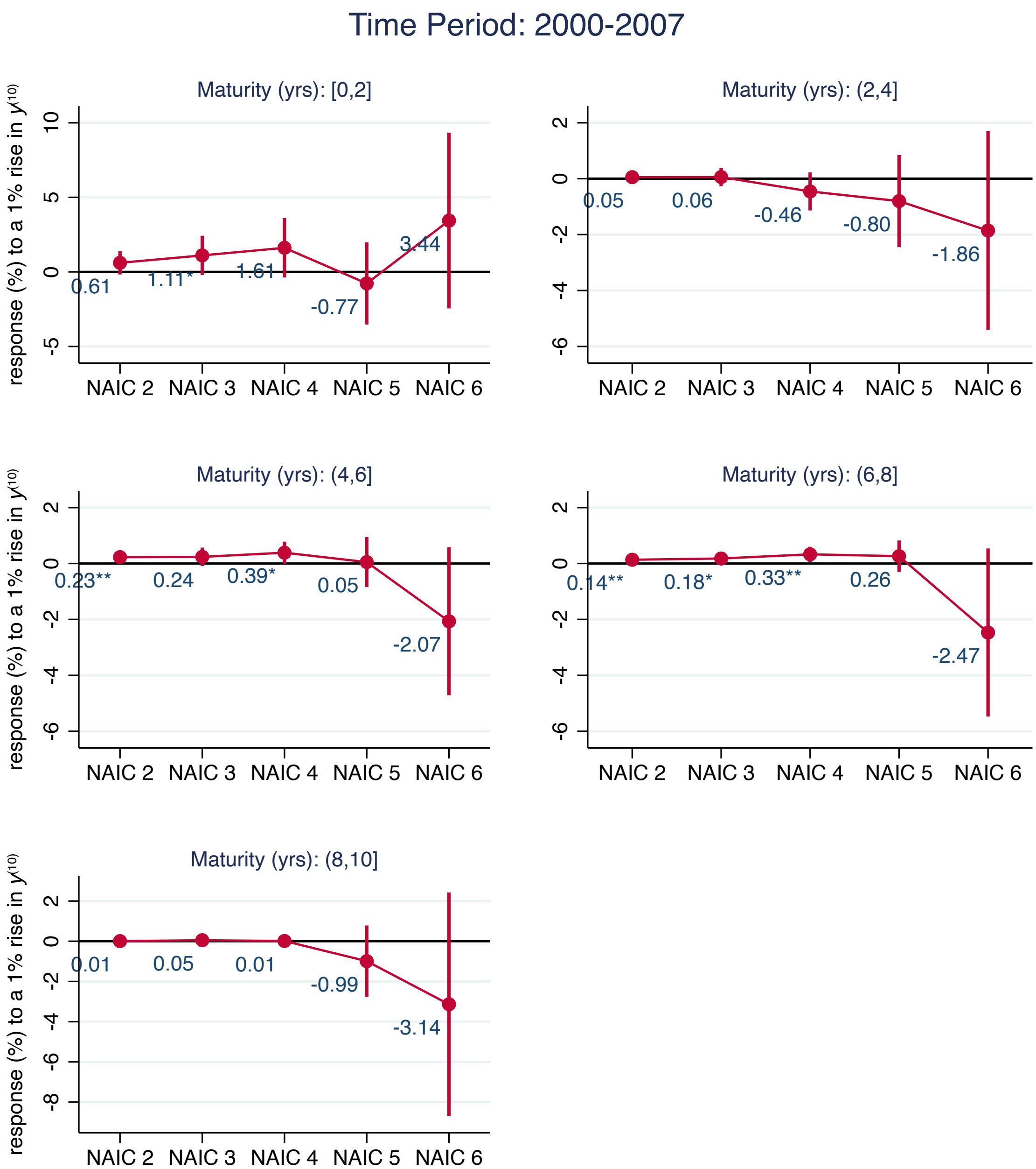

Figure A.5. **Credit Spread Responses in Different Maturity Groups (pre-2008).**

This figure plots the coefficients β_k estimated from regression (2), controlling for the trading volume, outstanding amount, maturity, duration, credit ratings, coupon amount, and coupon frequency of each bond, and the recent default rate and average CDS spreads of each NAIC category. Both corporate bond yields and the Treasury yield are in percentage points. The different panels estimate coefficients for bonds in different maturity segments. The t-statistics shown in brackets are based on standard errors clustered at the issuer and year-month levels. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

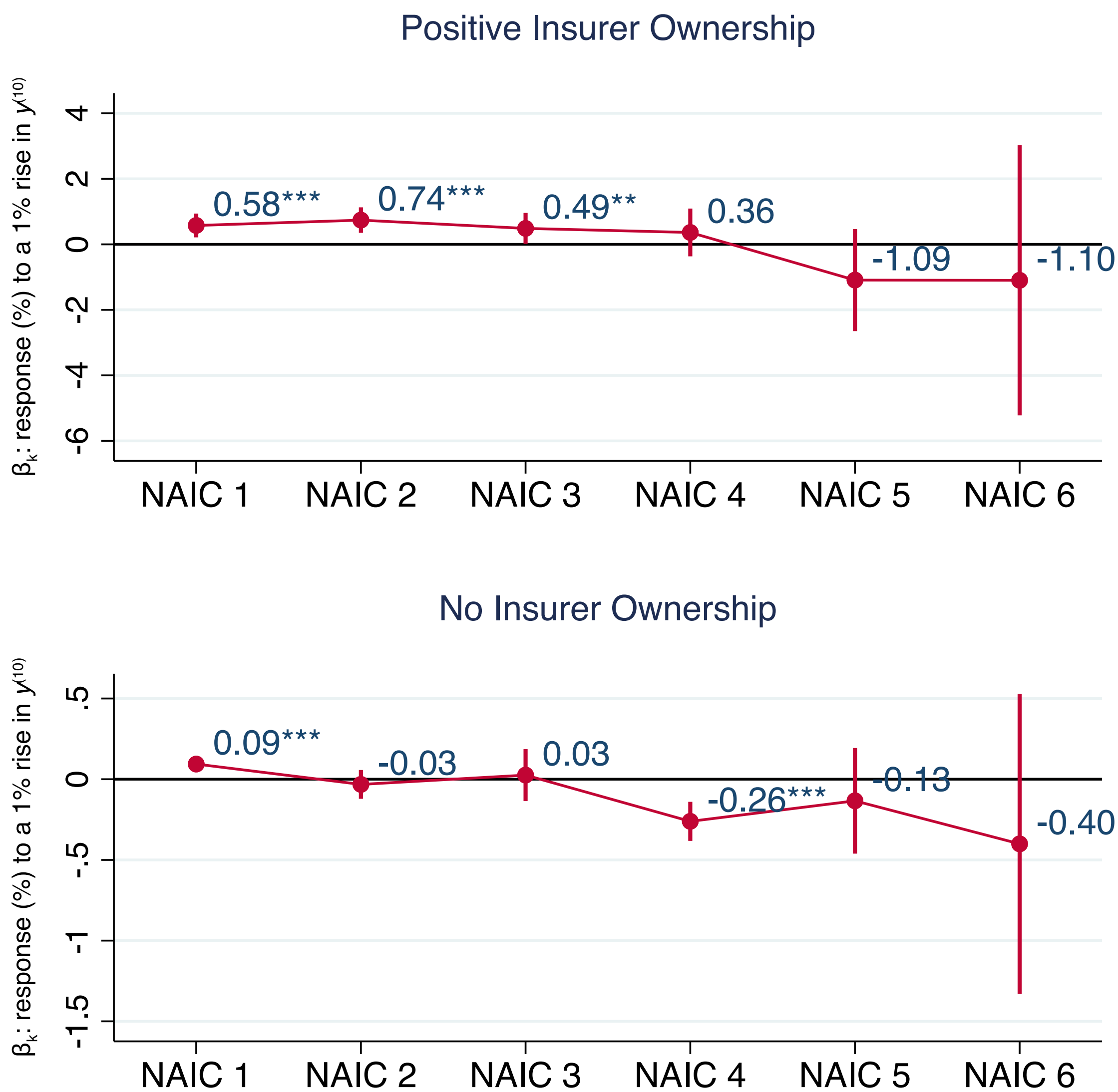

Figure A.6. **Pass-through by Life Insurance Ownership (pre-2008).**

This figure plots the coefficients β_k estimated from regression (5) for the period of 2000-2007, controlling for the trading volume, outstanding amount, maturity, duration, credit ratings, coupon amount, and coupon frequency of each bond, and the recent default rate of each NAIC category. Both corporate bond yields and the Treasury yield are in percentage points. The top panel shows estimates for β_k^1, while the bottom panel shows estimates for β_k^0. The t-statistics shown in brackets are based on standard errors clustered at the issuer and year-month levels. $^*p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

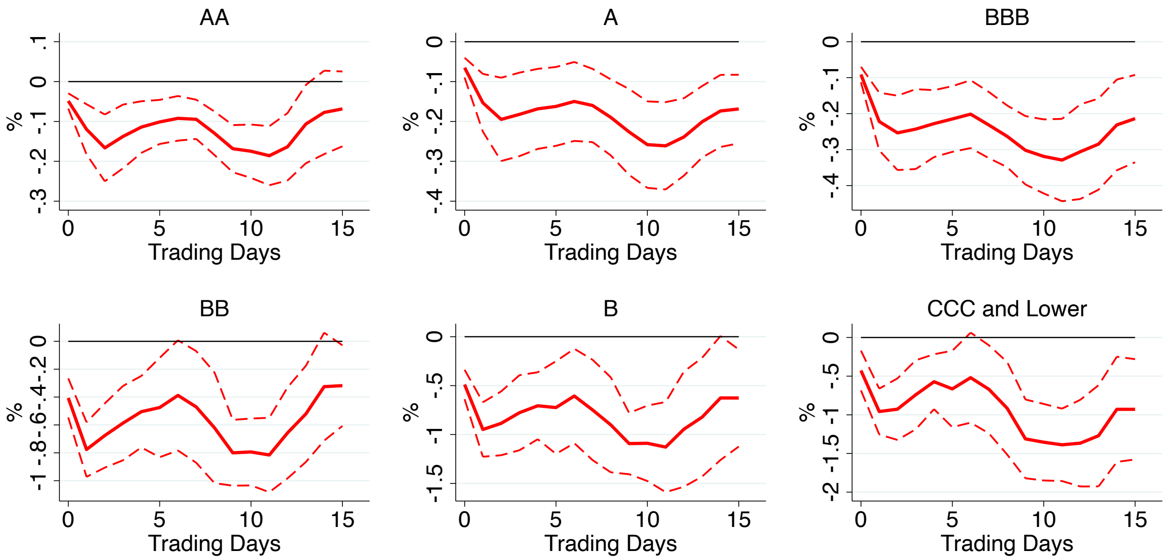

Figure A.7. **Impulse Responses of Spreads Relative to Treasuries (2010-2022).**

This figure plots the cumulative responses of credit spreads relative to Treasury securities to a 1% increase in the 10-year Treasury yield and 90% confidence intervals based on Newey-West standard errors with a maximum lag of 30 trading days.

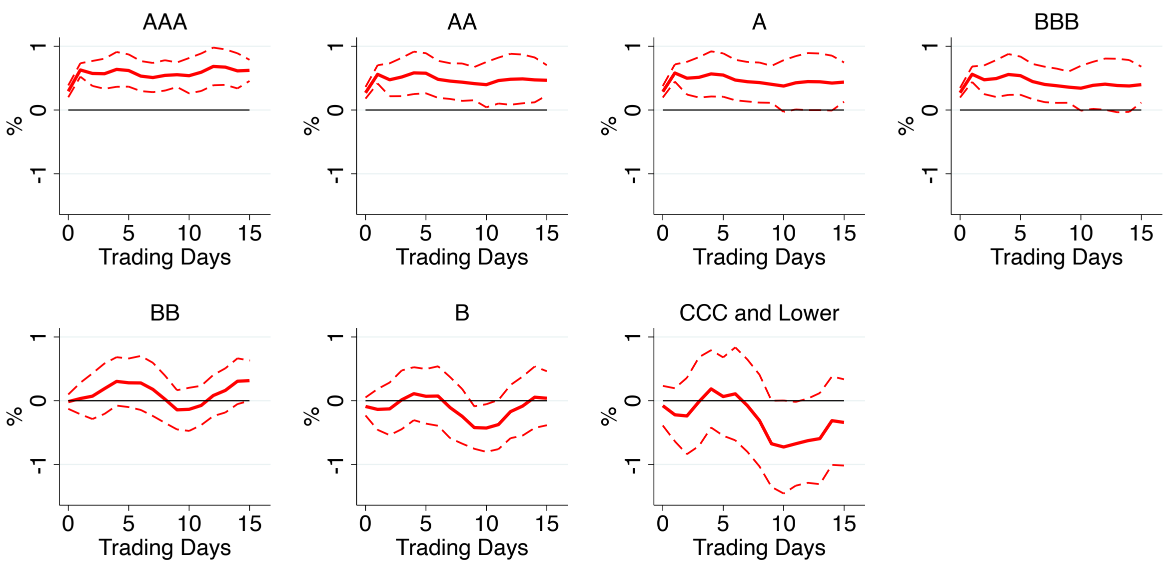

Figure A.8. **Impulse Responses of Yield Indices (2010-2022).**

This figure plots the cumulative responses of yield indices to a 1% increase in the 10-year Treasury yield and 90% confidence intervals based on Newey-West standard errors with a maximum lag of 30 trading days.

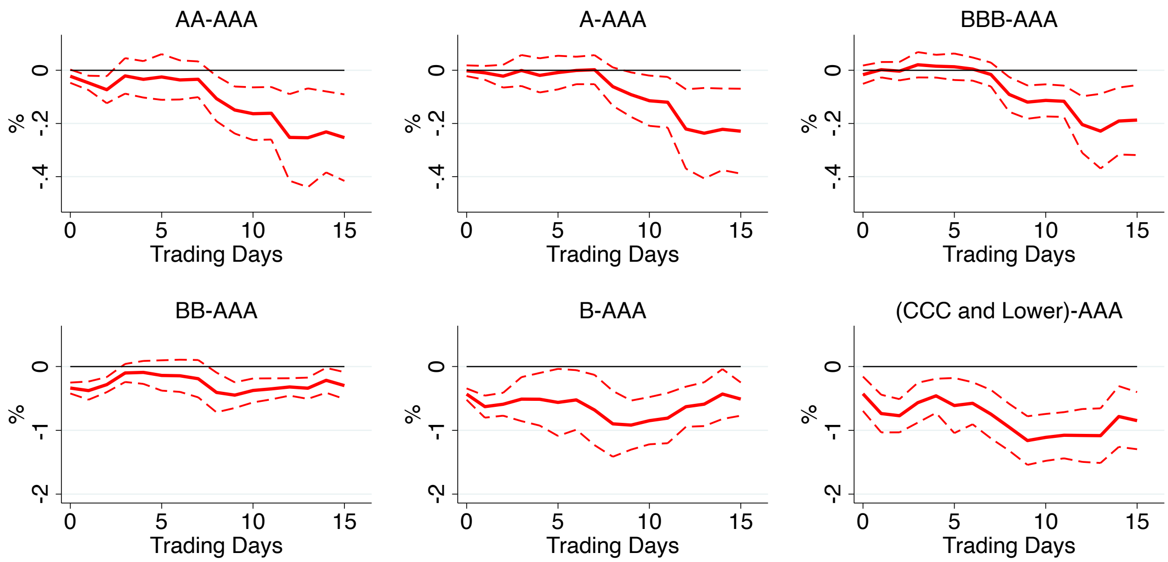

Figure A.9. **Impulse Responses of Credit Spreads (2010-2022).**

This figure plots the cumulative responses of credit spreads to a 1% increase in the 10-year Treasury yield and 90% confidence intervals based on Newey-West standard errors with a maximum lag of 30 trading days. ***I now include changes in the 1-month Treasury yield and average CDS spreads around 2-day FOMC windows as controls.*** *In particular, the regression specification is*

$$\begin{aligned} Spread^k_{t+h} - Spread^k_{t-1} = \alpha_h &+ \beta_h \left(\Delta y^{(10)}_t \Big|_{FOMC} \right) \\ &+ \gamma_h \left(\Delta y^{(1m)}_t \Big|_{FOMC} \right) \\ &+ \delta_h \left(\Delta \left(CDS\ Spread \right)^{AAA}_t \Big|_{FOMC} \right) \\ &+ \zeta_h \left(\Delta \left(CDS\ Spread \right)^{Rating\ k}_t \Big|_{FOMC} \right) + \varepsilon_{t,h}. \end{aligned}$$

Here $\left(\Delta y^{(1m)}_t \Big|_{FOMC} \right)$ *are the changes in the 1-month Treasury yield around 2-day FOMC windows.* $\left(\Delta \left(CDS\ Spread \right)^{AAA}_t \Big|_{FOMC} \right)$ *are the changes in the average CDS spread of AAA bonds around 2-day FOMC windows.* $\left(\Delta \left(CDS\ Spread \right)^{Rating\ k}_t \Big|_{FOMC} \right)$ *are the changes in the average CDS spread of rating k bonds around 2-day FOMC windows.*

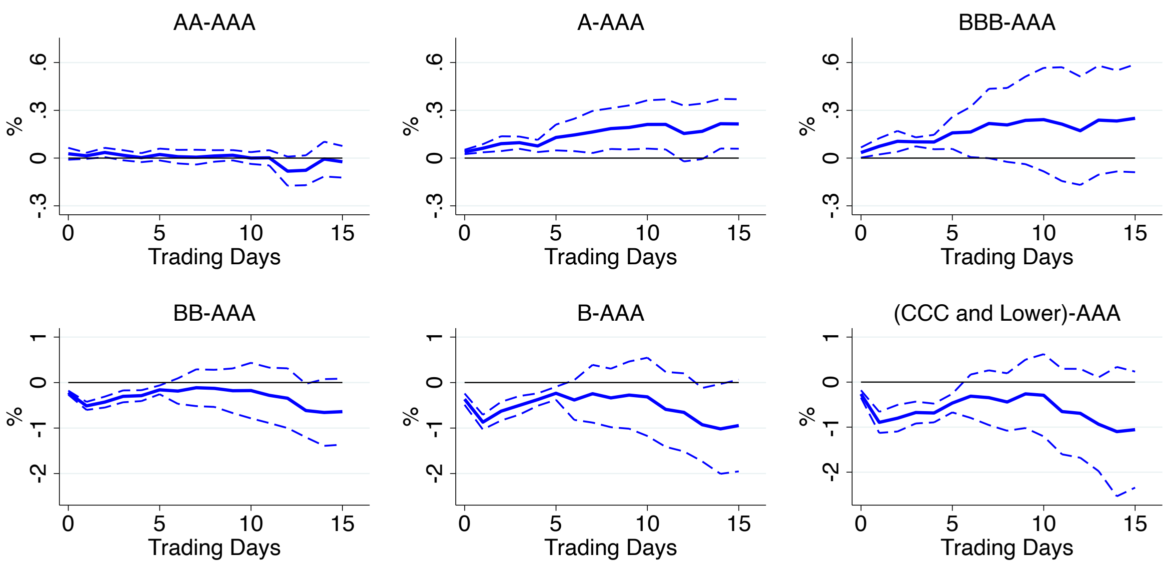

Figure A.10. **Impulse Responses of Credit Spreads (1997-2007).**

This figure plots the cumulative responses of credit spreads to a 1% increase in the 10-year Treasury yield and 90% confidence intervals based on Newey-West standard errors with a maximum lag of 30 trading days. ***I now include changes in the 1-month Treasury yield and average CDS spreads around 2-day FOMC windows as controls.*** *In particular, the regression specification is*

$$\begin{aligned} Spread^k_{t+h} - Spread^k_{t-1} = \alpha_h + \beta_h \left(\Delta y^{(10)}_t \Big|_{FOMC} \right) \\ + \gamma_h \left(\Delta y^{(1m)}_t \Big|_{FOMC} \right) \\ + \delta_h \left(\Delta \left(CDS\ Spread \right)^{AAA}_t \Big|_{FOMC} \right) \\ + \zeta_h \left(\Delta \left(CDS\ Spread \right)^{Rating\ k}_t \Big|_{FOMC} \right) + \varepsilon_{t,h}. \end{aligned}$$

Here $\left(\Delta y^{(1m)}_t \Big|_{FOMC} \right)$ *are the changes in the 1-month Treasury yield around 2-day FOMC windows.* $\left(\Delta \left(CDS\ Spread \right)^{AAA}_t \Big|_{FOMC} \right)$ *are the changes in the average CDS spread of AAA bonds around 2-day FOMC windows.* $\left(\Delta \left(CDS\ Spread \right)^{Rating\ k}_t \Big|_{FOMC} \right)$ *are the changes in the average CDS spread of rating k bonds around 2-day FOMC windows.*

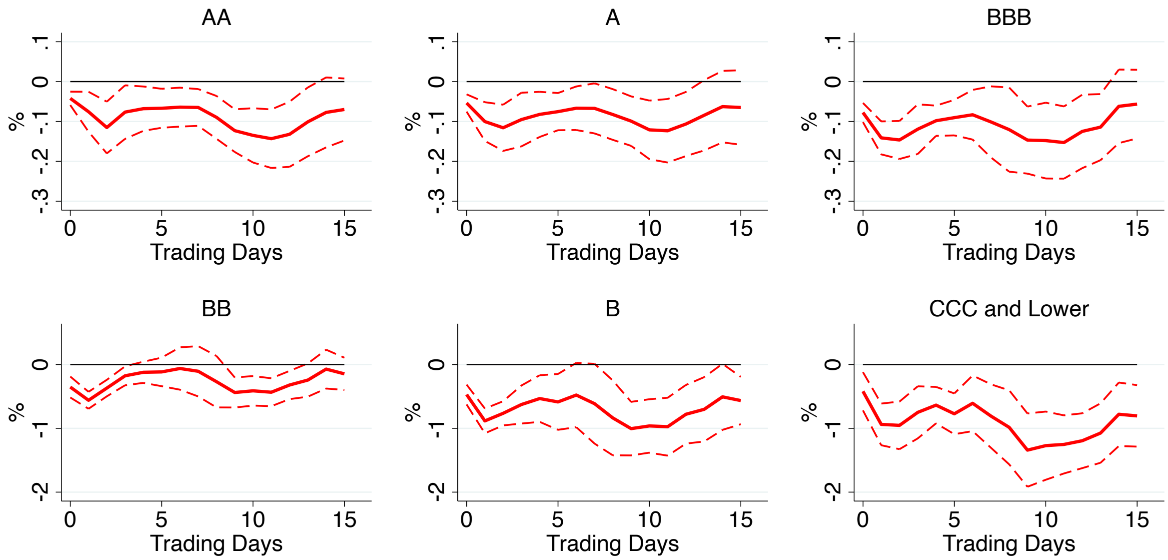

Figure A.11. **Impulse Responses of Spreads Relative to Treasuries (2010-2022).**

This figure plots the cumulative responses of credit spreads relative to Treasury securities to a 1% increase in the 10-year Treasury yield and 90% confidence intervals based on Newey-West standard errors with a maximum lag of 30 trading days. ***I now include changes in the 1-month Treasury yield and average CDS spreads around 2-day FOMC windows as controls.*** *In particular, the regression specification is*

$$\begin{aligned} Spread^k_{t+h} - Spread^k_{t-1} = \alpha_h + \beta_h \left(\Delta y^{(10)}_t \Big|_{FOMC} \right) \\ + \gamma_h \left(\Delta y^{(1m)}_t \Big|_{FOMC} \right) \\ + \zeta_h \left(\Delta \left(CDS\ Spread \right)^{Rating\ k}_t \Big|_{FOMC} \right) + \varepsilon_{t,h}. \end{aligned}$$

Here $\left(\Delta y^{(1m)}_t \Big|_{FOMC} \right)$ *are the changes in the 1-month Treasury yield around 2-day FOMC windows.* $\left(\Delta \left(CDS\ Spread \right)^{Rating\ k}_t \Big|_{FOMC} \right)$ *are the changes in the average CDS spread of rating k bonds around 2-day FOMC windows.*

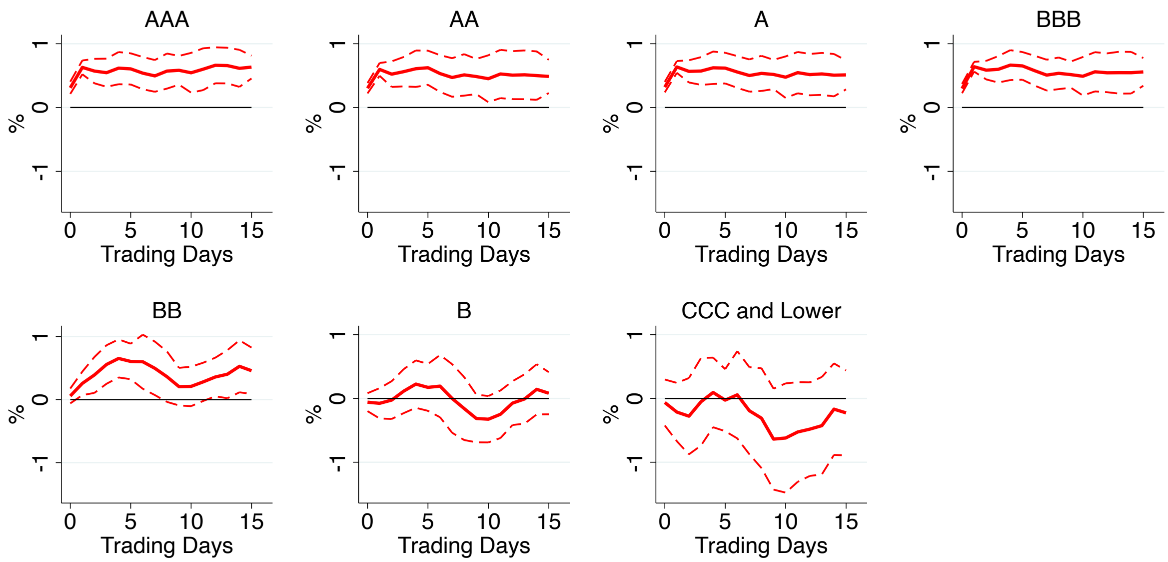

Figure A.12. **Impulse Responses of Yield Indices (2010-2022).**

This figure plots the cumulative responses of yield indices to a 1% increase in the 10-year Treasury yield and 90% confidence intervals based on Newey-West standard errors with a maximum lag of 30 trading days. ***I now include changes in the 1-month Treasury yield and average CDS spreads around 2-day FOMC windows as controls.*** *In particular, the regression specification is*

$$\begin{aligned} Yield^k_{t+h} - Yield^k_{t-1} = \alpha_h + \beta_h \left(\Delta y^{(10)}_t \Big|_{FOMC} \right) \\ + \gamma_h \left(\Delta y^{(1m)}_t \Big|_{FOMC} \right) \\ + \zeta_h \left(\Delta \left(CDS\ Spread \right)^{Rating\ k}_t \Big|_{FOMC} \right) + \varepsilon_{t,h}. \end{aligned}$$

Here $\left(\Delta y^{(1m)}_t \Big|_{FOMC} \right)$ *are the changes in the 1-month Treasury yield around 2-day FOMC windows.* $\left(\Delta \left(CDS\ Spread \right)^{Rating\ k}_t \Big|_{FOMC} \right)$ *are the changes in the average CDS spread of rating k bonds around 2-day FOMC windows.*

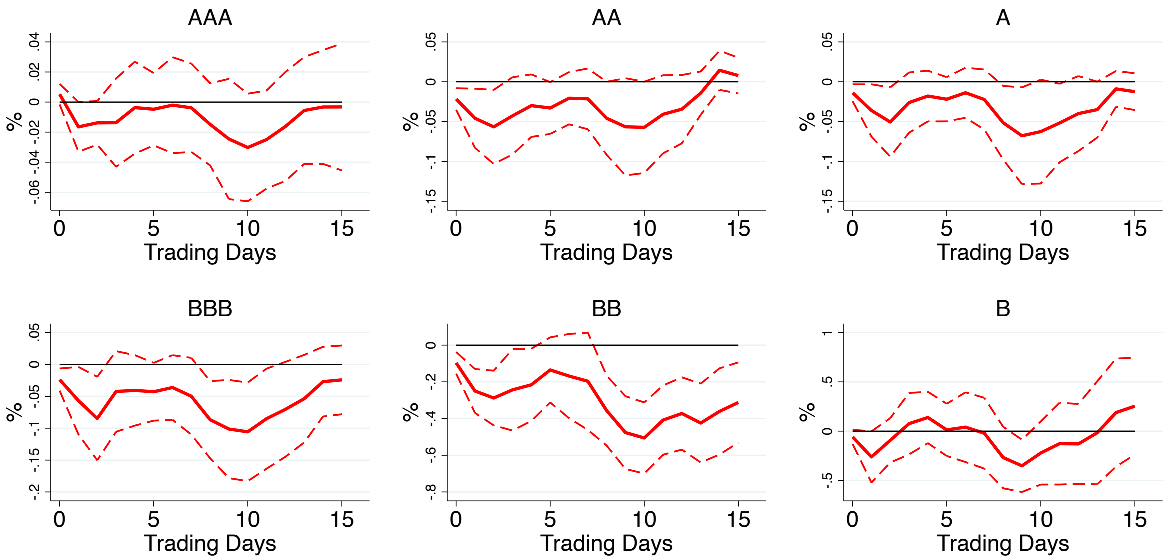

Figure A.13. **Impulse Responses of CDS Spreads (2010-2022).**

This figure plots the cumulative responses of the average CDS spreads of bonds in different risk categories to a 1% increase in the 10-year Treasury yield and 90% confidence intervals based on Newey-West standard errors with a maximum lag of 30 trading days.

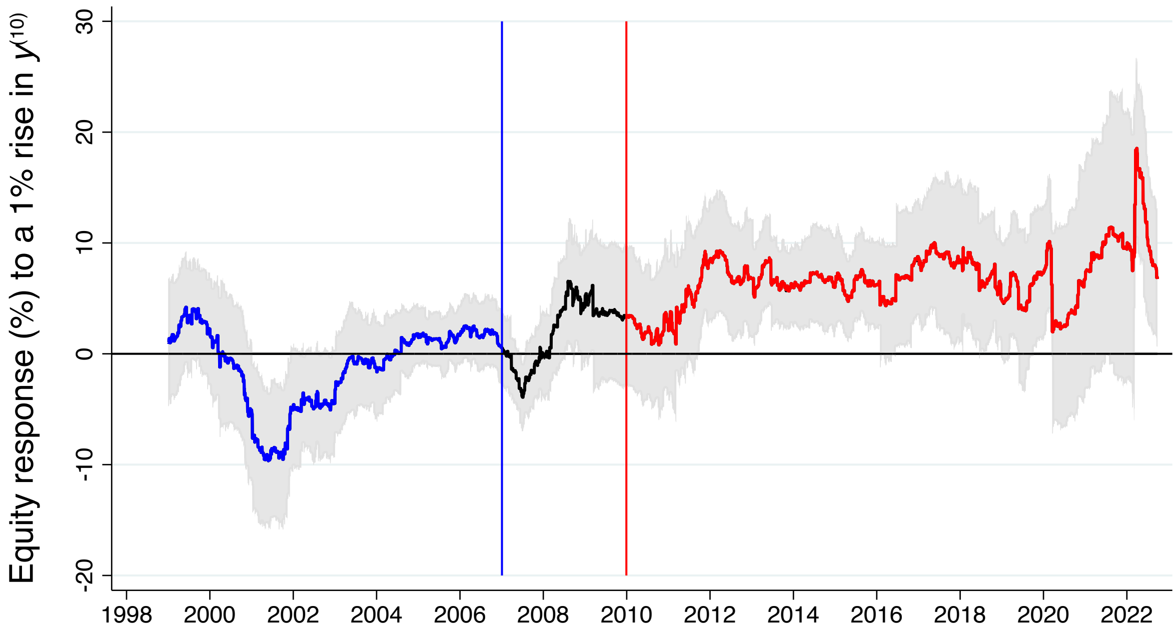

Figure A.14. **Life Insurers' Interest Rate Risk Exposure Over Time.**

This figure plots the two-year rolling estimates of β from regression (4), controlling for the market return and changes in the one-month Treasury yield. The blue segment indicates the pre-Crisis sample, and the red segment indicates the post-Crisis sample. The stock returns are winsorized at 2.5% and 97.5% thresholds. The shaded area indicates 95% confidence intervals using robust standard errors.

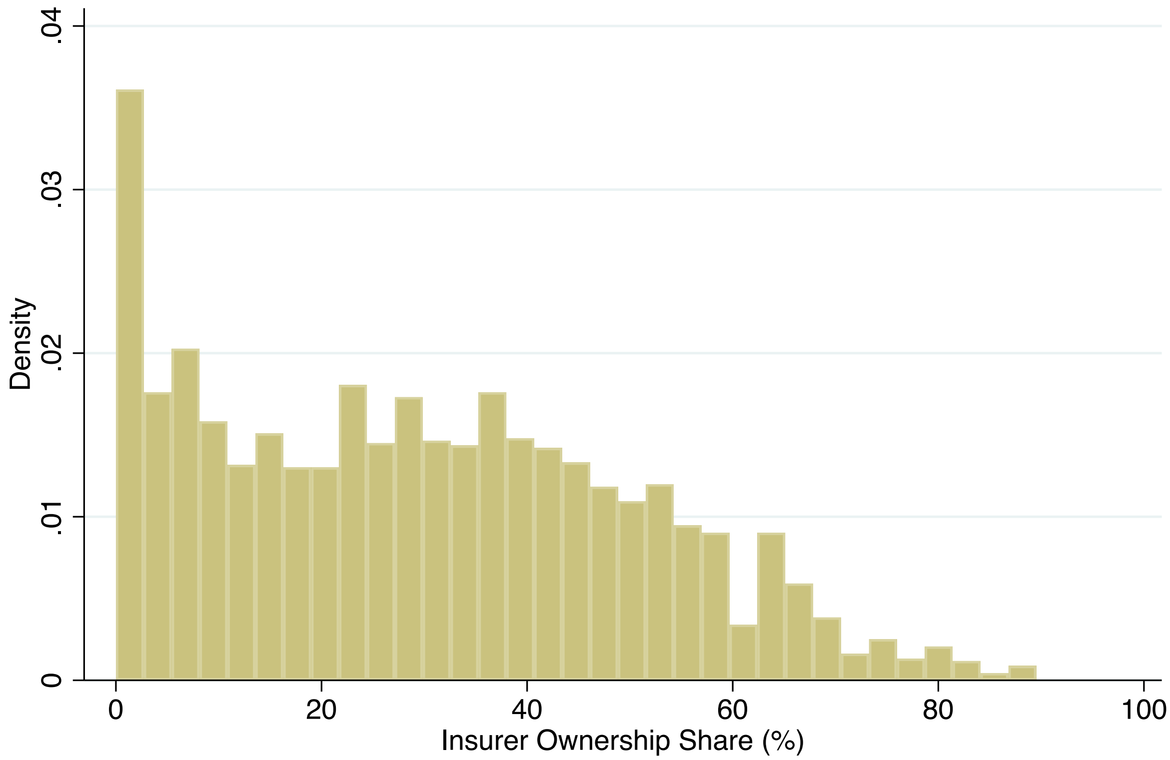

Figure A.15. **Insurer Ownership in the Cross-section of Firms.**

This histogram shows the average bond ownership by insurers (φ_j^{Ins}) in the cross-section of Compustat firms at the beginning of 2010.

B Supplementary Tables

	Bandwidth				
	1yr	2yrs	4yrs	6yrs	8yrs
γ_2	-0.416***	-0.452***	-0.423***	-0.403***	-0.395***
	[-4.903]	[-4.745]	[-4.131]	[-4.485]	[-4.237]
γ_{3-6}	-0.402	-0.813**	-0.712**	-0.962***	-1.392***
	[-0.935]	[-3.229]	[-2.574]	[-4.449]	[-5.173]
Time FE	✓	✓	✓	✓	✓
R^2	.696	.664	.667	.632	.581
First-stage F	112.7	111.2	92.1	55	60.8
# Identifying Bonds	2011	3774	4894	5823	7340

Table B.1. **Robustness: Bandwidth Choice.**

This table shows the coefficients estimated from regression (6) in the post-Crisis sample for different choices of bandwidth. The controls are the same as in Table 5. The t-statistics shown in brackets are based on standard errors clustered at the issuer and year-month levels. $^*p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

	2000-2007	2010-2019
γ_2	0.197	-0.465***
	[0.805]	[-4.862]
γ_{3-6}	0.473	-0.824***
	[0.786]	[-3.383]
Time FE	✓	✓
R^2	.599	.67
First-stage F	310	1950.4
# Identifying Bonds	1092	2802

Table B.2. **Robustness: Excluding New Bonds.**

This table shows the coefficients estimated from regression (6) excluding bonds issued less than 3 months ago. The controls are the same as in Table 5. The t-statistics shown in brackets are based on standard errors clustered at the issuer and year-month levels. $^*p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

	2000-2007	2010-2019
γ_2	0.131	-0.450***
	[0.599]	[-4.713]
γ_{3-6}	0.442	-0.895***
	[0.786]	[-3.375]
Time FE	✓	✓
R^2	.591	.659
First-stage F	3.6	122
# Identifying Bonds	1209	3617

Table B.3. **Robustness: Long Maturity at Issuance.**

This table shows the coefficients estimated from regression (6), including only bonds whose maturities are longer than 10 years when issued. The controls are the same as in Table 5. The t-statistics shown in brackets are based on standard errors clustered at the issuer and year-month levels. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

	$\mathbf{1}\{\varphi_{it}^{\text{Ins}} > 0\}$	$\mathbb{E}_t[\varphi_{it}^{\text{Ins}} \mid \varphi_{it}^{\text{Ins}} > 0]$	$\max\ \varphi_{it}^{\text{Ins}}$	Market Cap ($ bn)
NAIC 1	84.6%	32.5%	100%	1990.6
NAIC 2	98.5%	39.2%	98.6%	1102.6
NAIC 3	94.3%	16.0%	95.9%	255.8
NAIC 4	84.6%	6.0%	50.1%	260.5
NAIC 5	66.3%	3.9%	60.4%	81.6
NAIC 6	52.4%	3.7%	10.1%	6.2
NAIC 1-2	90.9%	35.7%	100%	3093.2
NAIC 3-6	85.1%	10.3%	95.9%	604.1

Table B.4. **Life Insurance Ownership (end of 2010).**

This table summarizes life insurers' bond ownership at the end of 2010. The first column shows the fraction of bonds with positive insurer holdings. The second column shows the average life insurance ownership share in bonds with life insurance ownership. The third column shows the maximum life insurer share for each category. The last column shows the total market cap of each category.

	Dependent Variable		
	Bond Yields (%)	3-month Net Purchases (%)	3-month Issuance (%)
β_{3-6}	-1.064***	1.002***	0.633***
	[-7.096]	[8.340]	[2.999]
Maturity-Time FE	✓		
Bond FE	✓		
Time FE		✓	✓
NAIC FE		✓	✓
Controls	✓	✓	✓
R^2	.753	.893	.65

Table B.5. **Untargeted Responses: The Empirical Estimates.**

This table shows the coefficients estimated from variations of regressions (2), (7), and (8) where $y_t^{(10)}$ interacts only with $\mathbf{1}_{\{NAIC\ 3\text{-}6\}}$. The controls are the same as in the corresponding regressions in the main context. The t-statistics shown in brackets are based on clustered standard errors. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

C Price-Yield Sensitivity

I estimate the sensitivity of bond prices in different NAIC categories with respect to their own yields as follows:

$$\log P_{it} = \alpha - \sum_k \widehat{D}_k \cdot \mathbf{1}_{\{\text{NAIC}\,k\}} \cdot y_{it} + \varepsilon_{it}. \tag{C.1}$$

Here $\log P_{it}$ is the natural log of the recorded market price of the bond i at time t. y_{it} is the bond's yield. The regression coefficients $\widehat{D}_k$ measures the percentage decrease in the price of NAIC k bonds when their yields increase by 1%. We can thus view $\widehat{D}_k$ as an estimate of the average duration of NAIC k bonds in my sample.

Table C.1 shows the estimates. Column (1) shows that, for the average transacted bond in my dataset, the bond price moves by 1.29% when its yield increases by 1%. Column (2) further shows that the duration of investment-grade bonds is between 0.5 and 1, while the duration of speculative-grade bonds is greater than 1.

	Traded Bonds (All)		Traded Bonds (Long-term)	
	(1)	(2)	(3)	(4)
$\widehat{D}_{1-6}$	1.228***		8.603***	
	[11.77]		[55.94]	
$\widehat{D}_1$		0.488***		9.232***
		[5.717]		[49.32]
$\widehat{D}_2$		0.929***		8.796***
		[6.648]		[62.98]
$\widehat{D}_{3-6}$		1.909***		7.988***
		[23.71]		[77.85]
# Bonds	23602	23602	13228	13228

Table C.1. **Price-Yield Sensitivity: Estimates.**

This table shows the coefficients $\widehat{D}_k$ estimated from regression (C.1). The sample for the first two columns includes the full sample of WRDS Bond Returns between 2010 and 2022. The sample for the last two columns uses corporate bonds with a maturity of over 5 years. The t-statistics shown in brackets are based on standard errors clustered at the issuer and year-month levels. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

The estimated duration is lower than the average duration of all corporate bonds and biased towards shorter-term bonds because they are traded more frequently and make up a larger fraction of the bond price data.

However, the duration of these shorter-term bonds is, in fact, the appropriate target for the quantitative model because the empirical results in Section 3 and Section 4 are also based on the same price data that use more observations from shorter-term bonds. Figure C.1 shows that the key results of Section 3.2 hold both in the full sample and for long-term bonds whose maturity is greater than 5 years, but the magnitudes obtained from long-term bonds are smaller (see also Figure A.4). The figure shows that short-term bonds play a more important role in the magnitude of the aggregate patterns found in Section 3.

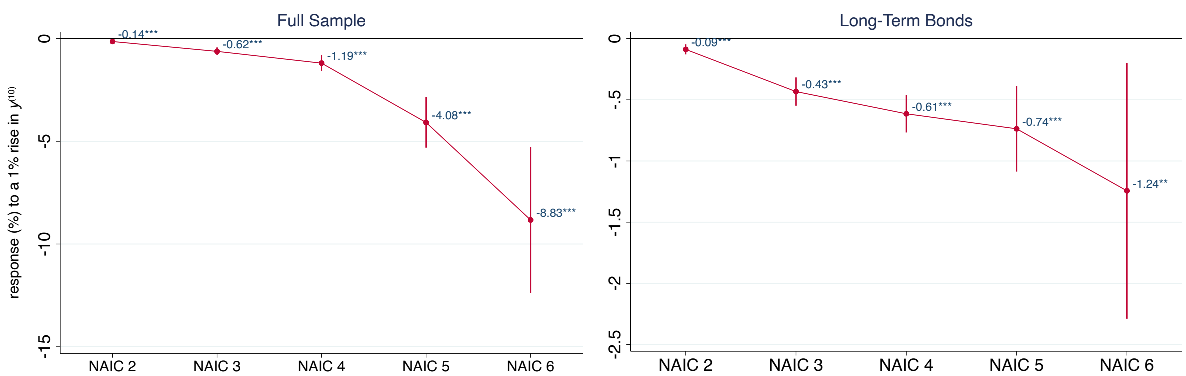

Figure C.1. **Credit Spread Responses: Long-term Bonds.**

This figure plots the coefficients β_k estimated from regression (2), controlling for the trading volume, outstanding amount, maturity, duration, credit ratings, coupon amount, and coupon frequency of each bond, and the recent default rate of each NAIC category. Both corporate bond yields and the Treasury yield are in percentage points. The left panel shows results for the full sample, while the right panel shows results for bonds with a maturity over 5 years. The data is from 2010 to 2022, excluding March 2020. The t-statistics shown in brackets are based on standard errors clustered at the issuer and year-month levels. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

To further validate regression (C.1), columns (3) and (4) of Table C.1 show that if we estimate duration only for long-term bonds whose maturity is greater than 5 years, we obtain a larger duration that is similar to other conventional measures.

D Analytical Model: Derivation

D.1 A Sufficient Condition for Duration Mismatch

Now we show that a set of sufficient conditions for Assumption 3 is that (1) $\tau^T < \tau^L$ and (2) the annuity demand L is sufficiently large.

First, differentiating equation (21),

$$\frac{\partial \widehat{A}}{\partial \widehat{y}} = \frac{L\phi^L}{(\widehat{y} + (1/\tau^L))^2} - \frac{T\phi^T}{(\widehat{y} + (1/\tau^T))^2}.$$

A set of sufficient conditions for $\partial \widehat{A}/\partial \widehat{y} > 0, \forall \widehat{y} \geq 0$ is that

$$\tau^T < \tau^L \tag{D.1}$$

$$T\phi^T < L\phi^L \tag{D.2}$$

Condition (D.2) can be written as

$$\begin{aligned}
\frac{w^{I,0}A}{P^T}\phi^T &< L\phi^L \\
w^{I,0}A(y+\lambda^T) &< L\phi^L \\
w^{I,0} &< \frac{L\phi^L}{A(y+\lambda^T)} \\
1 + \frac{L\phi^L}{A(y+\lambda^L)} - \sum_{n=1}^{N} w^{I,n} &< \frac{L\phi^L}{A(y+\lambda^T)} \\
1 + \frac{L\phi^L(\lambda^T - \lambda^L)}{A(y+\lambda^T)(y+\lambda^L)} &< \sum_{n=1}^{N} w^{I,n} \qquad \text{(D.3)}
\end{aligned}$$

Consider the case with $\gamma^n = 0$. Before the yield shock, the first-order condition implies

$$\frac{\phi^n}{P^n} - \nu^n\delta = y \qquad \Longrightarrow \qquad P^n = \frac{\phi^n}{y + \nu^n\delta}.$$

Then the market clearing conditions imply

$$\begin{aligned}
w^{I,n} &= \frac{P^n\left[B^n(P^n) - D^{P,n}(P^n)\right]}{A} \\
&= \frac{\phi^n}{(y+\nu^n\delta)A}\left[B^n\left(\frac{\phi^n}{y+\nu^n\delta}\right) - D^{P,n}\left(\frac{\phi^n}{y+\nu^n\delta}\right)\right].
\end{aligned}$$

Thus, condition (D.3) holds for $\gamma^n = 0$ if

$$\sum_{n=1}^{N} \phi^n \left[B^n \left(\frac{\phi^n}{y + \nu^n \delta} \right) - D^{P,n} \left(\frac{\phi^n}{y + \nu^n \delta} \right) \right] > (y + \nu^n \delta) A \left[1 + \frac{L\phi^L (\lambda^T - \lambda^L)}{(y + \lambda^T)(y + \lambda^L)} \right]$$

$$N(y + \nu^n \delta)^{\frac{\theta - 1}{\theta}} - \sum_{n=1}^{N} \alpha^n \left(\frac{\phi^n}{y + \nu^n \delta} \right)^{-\beta} > (y + \nu^n \delta) A \left[1 + \frac{L\phi^L (\lambda^T - \lambda^L)}{(y + \lambda^T)(y + \lambda^L)} \right]. \tag{D.4}$$

The condition (D.4) holds when N and ϕ^n are large enough.

In Section D.2, I will show that the portfolio weight $w^{I,n}$ decreases in γ^n. Since $w_t^{I,n}$ is a continuous function in γ^n, there exists some positive constant condition $\overline{\gamma} > 0$ such that (D.3) holds for $\gamma^n \in (0, \overline{\gamma})$. To summarize, we have now found a set of conditions under which $\partial \widehat{A} / \partial \widehat{y} > 0$ holds:

$$\gamma^n \in (0, \overline{\gamma}), (D.1), \text{ and, } (D.4).$$

D.2 Proof

Proof of Proposition 1. The market clearing condition (18) can be written as

$$w_t^{I,n} A_t^I = \underbrace{P_t^n \left(B^n(P_t^n) - D^{P,n}(P_t^n) \right)}_{\uparrow \text{ in } P_t^n}$$

Under Assumption 4, the right-hand side is increasing in P_t^n. The equation implies that P_t^n is a function of $w_t^{I,n}$ and A_t^I. Denote $P_t^n = P^n(w_t^{I,n}, A_t^I)$. It is easy to show that

$$\frac{\partial P_t^n}{\partial w_t^{I,n}} > 0, \quad \frac{\partial P_t^n}{\partial A_t^I} > 0.$$

In the simplified model, the return rate on Treasuries is simply $\mu_t^T = y_t^T$. Consider the insurer's first-order condition

$$\mu_t^{r,n} - y_t^T = \gamma^n w_t^{I,n} \tag{D.5}$$

$$\frac{\phi^n}{P^n(w_t^{I,n}, A_t^I)} - \nu^n \delta - y_t^T = \gamma^n w_t^{I,n}. \tag{D.6}$$

Differentiating (D.6) yields

$$\frac{\partial w_t^{I,n}}{\partial y_t^T} = - \left[\gamma^n + \frac{\phi^n}{(P_t^n)^2} \frac{\partial P_t^n}{\partial w_t^{I,n}} \right]^{-1} \left[1 + \frac{\phi^n}{(P_t^n)^2} \frac{\partial P_t^n}{\partial A_t^I} \frac{\partial A_t^I}{\partial \mu_t^T} \right]. \tag{D.7}$$

Under Assumption 3, $\partial A_t^I / \partial y_t^T > 0$, so $\partial w_t^{I,n} / \partial y_t^T < 0$. That is, the portfolio weight $w_t^{I,n}$ increases when the Treasury yield μ_t^T decreases. From (D.5), we see that the credit spreads must also increase. Equation (D.7) also tells us the magnitude of the effect (i.e., $\left|\frac{\partial w_t^{I,n}}{\partial y_t^T}\right|$) is increasing in the severity of duration mismatch (i.e., $\left|\frac{\partial A_t^I}{\partial y_t^T}\right|$).

We can also show that the portfolio weight $w_t^{I,n}$ decreases in the holding cost γ^n by differentiating (D.6),

$$\frac{\partial w_t^{I,n}}{\partial \gamma^n} = -\left[\gamma^n + \frac{\phi^n}{(P_t^n)^2}\frac{\partial P_t^n}{\partial w_t^{I,n}}\right]^{-1} w_t^{I,n} < 0.$$

E Quantitative Model: Numerical Solution Method

Consider the insurer's net worth dynamics

$$\begin{aligned}
\frac{\mathrm{d}A_t^I}{A_t^I} &= \left[w_t^{I,0}\mu_t^{r,T} + \sum_{n=1}^N w_t^{I,n}\mu_t^{r,n} - w_t^{I,L}\mu_t^{r,L} - \sum_{n=1}^N \frac{1}{2}(\zeta^n w_t^{I,n})^2 + \psi_t\right]\mathrm{d}t + \sum_{n=1}^N w_t^{I,n}\sigma_t^{r,n}(\mathrm{d}J_t - \delta\mathrm{d}t) \\
&= \left[w_t^{I,0}\mu_t^{r,T} + \sum_{n=1}^N w_t^{I,n}(\mu_t^{r,n} - \delta\sigma_t^{r,n}) - w_t^{I,L}\mu_t^{r,L} - \sum_{n=1}^N \frac{1}{2}(\zeta^n w_t^{I,n})^2 + \psi_t\right]\mathrm{d}t + \sum_{n=1}^N w_t^{I,n}\sigma_t^{r,n}\mathrm{d}J_t \\
&:= \mu_t^{A,I}\mathrm{d}t + \sigma_t^{A,I}\mathrm{d}J_t.
\end{aligned}$$

Suppose $P_t^n = P^n(y_t^T, A_t^I)$. By Ito's Lemma,

$$\begin{aligned}
\mathrm{d}r_t^n &= \frac{\phi^n - \lambda^n P_t^n}{P_t^n}\mathrm{d}t + \frac{\mathrm{d}P_t^n}{P_t^n} - \nu^n \mathrm{d}J_t \\
&= \underbrace{\frac{1}{P_t^n}\left[(\phi^n - \lambda^n P_t^n) + \frac{\partial P_t^n}{\partial y_t^T}\alpha_y\left(y_t^T - \bar{y}^T\right) + \frac{\partial P_t^n}{\partial A_t^I}A_t^I\mu_t^{A,I}\right]}_{:=\mu_t^{r,n} - \delta\sigma_t^{r,n}}\mathrm{d}t + \underbrace{\left[1 - \frac{P^n(y_t^T, A_t^I - \sigma_t^{A,I})}{P^n(y_t^T, A_t^I)} - \nu^n\right]}_{:=\sigma_t^{r,n}}\mathrm{d}J_t
\end{aligned}$$

Plugging $\mu_t^{r,n}$ into $\mu_t^{A,I}$, we can solve for $\mu_t^{A,I}$ as

$$\begin{aligned}
\mu_t^{A,I} = &\left[1 - \sum_n w_t^{I,n}\frac{A_t^I}{P_t^n}\frac{\partial P_t^n}{\partial A_t^I}\right]^{-1} \cdot \\
&\left\{w_t^{I,0}\mu_t^{r,T} - w_t^{I,L}\mu_t^{r,L} + \sum_n w_t^{I,n}\frac{1}{P_t^n}\left[\phi^n - \lambda^n P_t^n + \frac{\partial P_t^n}{\partial y_t^T}\alpha_y\left(y_t^T - \bar{y}^T\right)\right] - \sum_n \frac{1}{2}(\zeta^n w_t^{I,n})^2 + \psi_t\right\}.
\end{aligned}$$

To obtain a solvable partial differential equation, we now add the time dimension and postulate $P_t^n = P^n(t, y_t^T, A_t^I)$. By Ito's Lemma,

$$\begin{aligned} \mathrm{d}r_t^n &= \frac{\phi^n - \lambda^n P_t^n}{P_t^n}\mathrm{d}t + \frac{\mathrm{d}P_t^n}{P_t^n} - \nu^n \mathrm{d}J_t \\ &= \underbrace{\frac{1}{P_t^n}\left[(\phi^n - \lambda^n P_t^n) + \frac{\partial P_t^n}{\partial t} + \frac{\partial P_t^n}{\partial y_t^T}\alpha_y\left(y_t^T - \bar{y}^T\right) + \frac{\partial P_t^n}{\partial A_t^I}A_t^I \mu_t^{A,I}\right]}_{:=\mu_t^{r,n} - \delta\sigma_t^{r,n}}\mathrm{d}t \\ &\quad + \underbrace{\left[1 - \frac{P^n(y_t^T, A_t^I - \sigma_t^{A,I})}{P^n(y_t^T, A_t^I)} - \nu^n\right]}_{:=\sigma_t^{r,n}}\mathrm{d}J_t. \end{aligned}$$

The first-order condition of the portfolio choice problem is

$$\mu_t^{r,n} - \mu_t^{r,T} = a\delta\sigma_t^{A,I}\sigma_t^{r,n} + (\zeta^n)^2 w_t^{I,n}.$$

Plugging in the expression for $\mu_t^{r,n}$ and μ_t^T, we get

$$\begin{aligned} &\frac{1}{P_t^n}\left[(\phi^n - \lambda^n P_t^n) + \frac{\partial P_t^n}{\partial t} + \frac{\partial P_t^n}{\partial y_t^T}\alpha_y\left(y_t^T - \bar{y}^T\right) + \frac{\partial P_t^n}{\partial A_t^I}A_t^I\mu_t^{A,I}\right] \\ &\quad - \left[\frac{(\phi^T - \lambda^T P_t^T)}{P_t^T} - \frac{1}{y_t^T + \lambda^T}\alpha_y\left(y_t^T - \bar{y}^T\right)\right] + \delta\sigma_t^{r,n} = a\delta\sigma_t^{A,I}\sigma_t^{r,n} + (\zeta^n)^2 w_t^{I,n}. \end{aligned}$$

Therefore, we get the following system of partial differential equations:

$$\begin{aligned} \frac{\partial P_t^n}{\partial t} =& -\alpha_y\left(y_t^T - \bar{y}^T\right)\frac{\partial P_t^n}{\partial y_t^T} - \left(A_t^I\mu_t^{A,I}\right)\frac{\partial P_t^n}{\partial A_t^I} \\ &+ \left[y_t^T - \frac{1}{y_t^T + \lambda^T}\alpha_y\left(y_t^T - \bar{y}^T\right) + \lambda^n + (a\sigma_t^{A,I} - 1)\delta\sigma_t^{r,n} + (\zeta^n)^2 w_t^{I,n}\right]P_t^n - \phi^n, \end{aligned} \tag{E.1}$$

where

$$B_t^n = \left(1/\phi^n\right)^{\frac{1}{\theta}}\left(P_t^n\right)^{\frac{1-\theta}{\theta}}, \qquad D_t^n = \alpha^n (P_t^n)^{-\beta}, \qquad w_t^{I,n} = P_t^n(B_t^n - D_t^n)/A_t^I$$

$$\mu_t^{A,I} = \left[1 - \sum_n w_t^{I,n}\frac{A_t^I}{P_t^n}\frac{\partial P_t^n}{\partial A_t^I}\right]^{-1}.$$

$$\left\{w_t^{I,0}\mu_t^{r,T} - w_t^{I,L}\mu_t^{r,L} + \sum_n w_t^{I,n}\frac{1}{P_t^n}\left[\phi^n - \lambda^n P_t^n + \frac{\partial P_t^n}{\partial y_t^T}\alpha_y\left(y_t^T - \bar{y}^T\right)\right] - \sum_n \frac{1}{2}(\zeta^n w_t^{I,n})^2 + \psi_t\right\}$$

$$\sigma_t^{r,n} = 1 - \frac{P^n(y_t^T, A_t^I - \sigma_t^{A,I})}{P^n(y_t^T, A_t^I)} - \nu^n$$

Finally, $\sigma_t^{A,I}$ is obtained by solving the following system of equations

$$\begin{aligned}
\sigma_t^{A,I} &= \sum_{n=1}^{N} w_t^{I,n} \sigma_t^{r,n} \\
&= \sum_{n=1}^{N} w_t^{I,n} \left[1 - \frac{P^n(y_t^T, A_t^I - \sigma_t^{A,I})}{P^n(y_t^T, A_t^I)} - \nu^n \right] \\
&= \sum_{n=1}^{N} w_t^{I,n}(1 - \nu^n) - \sum_{n=1}^{N} w_t^{I,n} \frac{P^n(y_t^T, A_t^I - \sigma_t^{A,I})}{P^n(y_t^T, A_t^I)},
\end{aligned}$$

which can be simplified to

$$\sigma_t^{A,I} - \sum_{n=1}^{N} w_t^{I,n}(1 - \nu^n) + \sum_{n=1}^{N} w_t^{I,n} \frac{P^n(y_t^T, A_t^I - \sigma_t^{A,I})}{P^n(y_t^T, A_t^I)} = 0.$$

I solve the PDE system (E.1) using a finite difference method. I start with a guess for $P^n(0, y_t^T, A_t^I)$ and iterate backward through time until the system converges.