Long Rates, Life Insurers, and Credit Spreads*

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April 18, 2025

Abstract

This paper examines how the duration mismatch of life insurance companies, the largest institutional investors in the US corporate bond market, affects credit spread dynamics. Post-GFC, US life insurers face large and negative duration gaps. When long-term interest rates increase, life insurers realize equity gains, which boost their risk-bearing capacity and compress credit spreads. Empirically, post-2008, corporate bond credit spreads decline when long rates rise, which holds both unconditionally and around monetary policy announcements. In the cross-section, I utilize a regression discontinuity design to confirm that this negative co-movement is more pronounced for bonds held more by life insurers.

JEL Codes: G11, G12, G22, E44, E52
Keywords: Credit Spreads, Long-term Interest Rates, Life Insurance Companies, Interest Rate Risk, Monetary Policy

*I am deeply indebted to Markus Brunnermeier, Moritz Lenel, Jonathan Payne, Wei Xiong, and Motohiro Yogo for their guidance and support. I also thank Joseph Abadi, Jennie Bai, Harjoat Bhamra, Adrien d'Avernas, Yuchen Chen, Darrell Duffie, Chuck Fang, Shan Ge, Goutham Gopalakrishna, Brett Green, Nicolas Hommel, Shiyang Huang, John Hund (discussant), Rajkamal Iyer, Mahyar Kargar, Paymon Khorrami, Nobuhiro Kiyotaki, Leonid Kogan, Robert Kosowski, Pete Kyle, Matteo Maggiori, Aytek Malkhozov (discussant), Ian Martin, Alexander Michaelides, Lira Mota, Ricardo Lagos, Jiwon Lee, Seung Joo Lee, Xuewen Liu, Sebastian Merkel, Atif Mian, Don Noh, Jonathan Parker, Neil Pearson, Manju Puri, Tarun Ramadorai, Richard Rogerson, Karthik Sastry, Andreas Schaab, Lukas Schmid (discussant), Lawrence Schmidt, Ran Shi, Kerry Siani (discussant), Alp Simsek, Dimitri Vayanos, Emil Verner, Vish Viswanathan, Michael Woodford, Yucheng Yang, Kathy Yuan, Xiaofei Zhao, Haonan Zhou, seminar participants at Bocconi, CU Boulder, CUHK, CKGSB, Duke, Georgetown, HKU, Imperial College, LSE, Maryland, MIT, MSU, OSU, Princeton, UIUC, UMN, WUSTL, Yale, and conference participants at AFA, Bank of Canada FSRC, Eastern FA, EEA-ESEM, WFA, and the Young Scholars Finance Consortium, and several anonymous referees for constructive feedback.

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1 Introduction

The US corporate bond market has expanded dramatically in the last four decades, with its total outstanding reaching 56% of GDP in 2023 Q4 from 17% of GDP in 1983 Q4 and surpassing the bank loan market as the primary funding source for US firms. Over the past forty years, the US corporate bond market has been dominated by large institutional investors. Among them, life insurance companies have consistently been the largest investor group, holding over 20% of outstanding corporate bonds. Given their dominant size in the bond market, the demand of life insurers can have considerable impacts on bond prices and credit spreads.

How do life insurers affect bond pricing? The existing literature has identified two potential channels. First, insurers face risk-based capital regulations, so negative shocks to their capital or the capital charges of bonds restrict their bond demand, resulting in fire sales and fluctuations in credit spreads (Ellul, Jotikasthira and Lundblad, 2011; Murray and Nikolova, 2022). Second, the liabilities of life insurers mostly consist of life insurance and annuity products, which are inherently long-term and illiquid. Hence, insurers have more stable funding than most other institutional bond investors, and they act as a stabilizing force in times of market distress (Chodorow-Reich, Ghent and Haddad, 2021; Coppola, 2024). In this paper, I propose a novel *duration mismatch channel* where interest rates influence insurers' bond demand and equilibrium credit spreads through life insurers' interest rate risk exposure.

Modern life insurers issue long-duration products and invest the proceeds in fixed income markets. The long-lived liabilities naturally expose life insurers to interest rate risk, as higher interest rates, especially higher long-term interest rates, compress insurers' liabilities more than their assets and generate capital gains. As shown in the earlier literature (e.g., Koijen and Yogo, 2022; Sen, 2023), US life insurers have sustained large negative duration gaps since the 2007-2008 Finance Crisis. Post-GFC, increases in long rates strengthen the capital of life insurers and expand their risk-taking capacity. Consequently, insurers increase their bond demand and rebalance their portfolios towards riskier bonds. Due to the large share of life insurers in the bond market, such demand shifts can significantly reduce credit spreads.

I first develop a tractable model of the corporate bond market featuring duration-mismatched life insurers as key investors to illustrate the mechanism. In the model, life insurers hold corporate bonds and Treasuries as assets and issue long-term annuities as liabilities. Life insurers face duration mismatch because their liabilities have a much longer duration than their assets. Increases in long-term Treasury yields cause larger declines in the value of liabilities than the value of assets, which boosts the insurers' net worth and risk-bearing capacity. Thus, when the long rate rises, life insurers become more willing to hold bonds in lower credit ratings that carry greater credit risk and regulatory costs. The increase in insurers' bond demand then leads to lower equilibrium credit spreads. Therefore, the model implies that negative duration gaps of life insurers cause a negative aggregate co-movement between long rates and credit spreads.

Following the model predictions, I document novel facts about the co-movements between long-term interest rates and corporate bond credit spreads over time. Leveraging detailed microdata on corporate bond prices, I estimate the pass-through of the 10-year Treasury yield to corporate bonds with different credit ratings. Before the 2007-2008 Financial Crisis, when life insurers had no significant interest rate risk exposure, long rates and credit spreads exhibited no significant associations. Changes in the long-term interest rate affected all bonds in a similar fashion. However, after the Financial Crisis, as life insurers face severe negative duration gaps, an economically and statistically significant negative relationship between long rates and credit spreads emerges. 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. In addition, I find that the negative co-movements become more pronounced when life insurers' duration gap is more negative.

Contrary to the conventional wisdom where high interest rates raise risk premia by increasing default risk (as in standard corporate finance models, e.g., Hennessy and Whited, 2007; Gomes and Schmid, 2010) and discouraging risk-taking (e.g., Jiménez et al., 2014; Bauer, Bernanke and Milstein, 2023), increases in the long-term interest rate lead to declines in corporate bond credit spreads, even after controlling for changes in credit risk 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 high-yield bonds increase significantly more than investment-grade bonds. Overall, the evidence suggests that an alternative mechanism, such as the duration mismatch of life insurers, might be at play.

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 the 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 in the US bond market, which owns over 20% of all US corporate bonds. The business model of modern life insurers involves investing in fixed-income securities and issuing long-term life insurance and 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 more than 6% when the 10-year Treasury yield rose by 1%.

I provide causal evidence that higher life insurance ownership generates stronger co-movement between long-term interest rates and credit spreads. To sharpen the identification, I exploit a discontinuity in bond ownership structure stemming from mutual

[1]Lower interest rates increase the insurers' liability duration more than their asset duration. Together with other institutional frictions (e.g., regulations, transaction costs, and market incompleteness), it can leave insurers in a persistent duration mismatch. See Section 5.1 for detailed discussions on the change in life insurers' hedging behavior.

funds' investment mandates (e.g., Li and Yu, 2024; Bretscher, Schmid and Ye, 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 whose time to maturity is slightly below 10 years are much less responsive to the long rate.

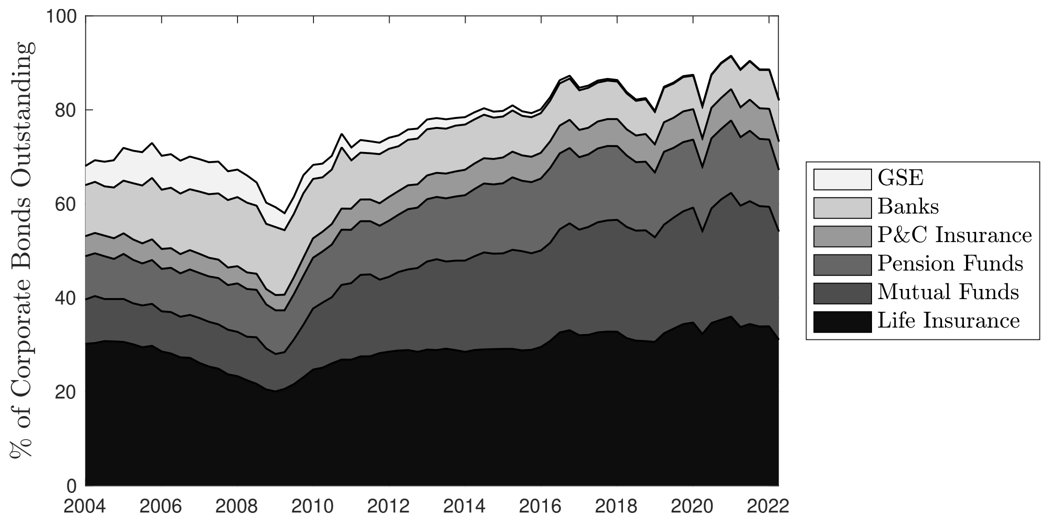

Figure 1. **Institutional Investors in the Bond Market.**

This figure plots the ownership structure of the US corporate and foreign bond market between 2004 and 2022 (excluding holdings by foreign investors). 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. The data is obtained from the Financial Accounts of the United States.

Again, the results point to a new channel through which the long-term interest rate and monetary policy affect the corporate bond market — 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.

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 Swaminathan, 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, 2025). 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 (2024) and Li and Yu (2022, 2024), 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

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

[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. Longstaff and Schwartz (1995) also finds a negative relationship between Treasury yields and the bond-Treasury spreads and emphasizes the role of changes in the risk-neutral default probability. In this paper, I show that the negative relationship between credit spreads and long rates post-2008 is not explained by the market price of default risk, measured by Merton's expected default frequency (EDF).

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 20% 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. (2024) and Darmouni, Siani and Xiao (2025) use demand system approaches to analyze the equilibrium effects of institutional demand. In this paper, I depart from the typical demand functions used in the demand system literature to analyze the role of life insurers' duration mismatch.

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), Huber (2022), Sen (2023), and Kirti and Singh (2024), have also shown that life insurers' interest rate risk exposure changed after the Financial Crisis.[4] I build on this finding and argue that the duration mismatch can significantly influence bond prices since life insurers, on average, hold more than 20% of 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 and propose a new mechanism stemming from their interest rate risk. Consistent with Ge and Weisbach (2021) and Bhardwaj, Ge and Mukherjee (2022), I find that life insurers tilt their portfolios toward safer investments when they become financially constrained.

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

[4] Li and Wenning (2025) study the product market implications of life insurers' 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, 2024; 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.

Another literature on the co-movement between Treasury and stock returns focuses more on inflation and output dynamics. For example, Campbell, Pflueger and Viceira (2020) finds that the Treasury-stock co-movement turned negative around 2001 and that periods of high Treasury returns tend to coincide with periods of low output that hurt stock returns.[5] In this paper, I find that the negative co-movement between long rates and credit spreads emerged much later than the negative Treasury-stock co-movement and is driven mostly by the risk exposure of life insurers instead of default risk.

Outline. Section 2 builds an intermediary asset pricing model to illustrate the mechanism and generate testable implications. Section 3 describes data sources for the empirical analysis. Section 4 discusses evidence on the co-movement between the long-term interest rate and corporate bond credit spreads. Section 5 investigates the role of life insurers in shaping the co-movement. Section 6 concludes.

2 An Intermediary Asset Pricing Model

In this section, I build an intermediary asset pricing model centered around life insurers to illustrate how long rates affect insurers' bond demand and credit spreads. As

[5]See also Baele, Bekaert and Inghelbrecht (2010), David and Veronesi (2013), and Song (2017).

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 (see Section 5.1 for empirical evidence), 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.

2.1 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	Long-duration Annuities
Treasuries	Net Worth

Preferred-habitat investors include all other investors of corporate bonds (e.g., mutual 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

$$\mathrm{d}r_t^T = \frac{(\phi^T - \lambda^T P_t^T)\mathrm{d}t}{P_t^T} + \frac{\mathrm{d}P_t^T}{P_t^T}, \tag{1}$$

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[6]

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

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. I assume that the dynamics of the long-term Treasury yield is driven by (unspecified) monetary policy and is exogenous to the model.

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 $y_t^L := \phi^L / P_t^L - \lambda^L$.

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,

$$\mu_t^{r,T} = \mu_t^{r,L}, \tag{2}$$

where $\mu_t^{r,T} = \mathrm{d}r_t^T/\mathrm{d}t$ and $\mu_t^{r,L} = \mathrm{d}r_t^L/\mathrm{d}t$ 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.

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

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

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

$$\mathrm{d}r_t^n = \frac{\phi^n - \lambda^n P_t^n}{P_t^n}\mathrm{d}t + \frac{\mathrm{d}P_t^n - \nu^n P_t^n \mathrm{d}J_t}{P_t^n}, \tag{3}$$

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.[7] 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.[8]

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,[9] and the insurer chooses its holdings 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

[7] 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.

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

[9] 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).

passing stress tests.[10] 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. The insurer has a mean-variance preference over its growth rate:

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

The portfolio problem can then be written as[11]

$$\max_{\{w_t^{I,n}\}_{n=1}^N,\, w_t^{I,T},\, w_t^{I,L}} \quad \underbrace{w_t^{I,T}\mu_t^T + \sum_{n=1}^N w_t^{I,n}\mu_t^{r,n} - w_t^{I,L}\mu_t^{r,L}}_{\text{expected return}} - \underbrace{\frac{a}{2}\delta\left(\sum_{n=1}^N w_t^{I,n}\sigma_t^{r,n}\right)^2}_{\text{risk aversion}} - \underbrace{\frac{1}{2}\sum_{n=1}^N\left(\zeta^n w_t^{I,n}\right)^2}_{\text{regulatory cost}}, \tag{4}$$

$$\text{s.t.} \quad w_t^{I,T} + \sum_{n=1}^N w_t^{I,n} - w_t^{I,L} = 1, \tag{5}$$

The objective function (4) 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. (5) 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 habitat demand is a reduced-form function of bond prices $D_t^{P,n} = D^{P,n}(P_t^n)$. The preferred-habitat demand is downward-sloping, i.e., $(D^{P,n})'(\cdot) \leq 0$.[12]

[10] 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.

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

[12] The assumption is standard in the literature of institutional asset demand and supported by various empirical estimates (e.g., Bretscher et al., 2024; Darmouni, Siani and Xiao, 2025).

Bond Supply. There exists a firm sector that supplies corporate bonds. For simplicity, I model bond supply as a function of bond prices, $B_t^n = B^n(P_t^n)$. The supply function is naturally upward-sloping (i.e., $(B^n)'(\cdot) \geq 0$), because firms face lower borrowing costs when their bond prices are higher, and they exploit the lower borrowing costs by issuing additional bonds. I provide a micro-foundation and derive an explicit form of the supply function in Appendix D.

Market Clearing. There are three asset markets in this model: Treasury, corporate bond, and annuity. Recall that the Treasury yield is exogenously given. Treasuries are supplied perfectly elastically to clear the market, given the exogenous yield. The market for rating n corporate bonds clears when

$$w_t^{I,n} A_t^I + P_t^n D_t^{P,n} = P_t^n B_t^n. \tag{6}$$

The market clearing condition requires that the sum of the life insurer's demand $w_t^{I,n} A_t^I$ and preferred-habitat investors' demand $P_t^n D_t^{P,n}$ is equal to the total market cap $P_t^n B_t^n$. Finally, the annuity market clears when the insurer's portfolio weight in annuities $w_t^{I,L}$ matches the exogenously fixed annuity demand,

$$w_t^{I,L} A_t^I = P_t^L L. \tag{7}$$

2.2 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 expected to be constant.*

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 (1) and (2), 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,

$$P_t^T = \frac{\phi^T}{y + (1/\tau^T)}, \quad P_t^L = \frac{\phi^L}{y + (1/\tau^L)}. \tag{8}$$

Corresponding to the standard Gordon formula,[13] 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 (8), 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.

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 (8). 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

$$\begin{aligned}\widehat{A} - A &= \overbrace{T\left(\widehat{P}^T - P^T\right)}^{\text{Repricing of Treasuries}} - \overbrace{L\left(\widehat{P}^L - P^L\right)}^{\text{Repricing of Annuities}} \\ &= T\left[\frac{\phi^T}{\widehat{y} + (1/\tau^T)} - \frac{\phi^T}{y + (1/\tau^T)}\right] - L\left[\frac{\phi^L}{\widehat{y} + (1/\tau^L)} - \frac{\phi^L}{y + (1/\tau^L)}\right].\end{aligned} \tag{9}$$

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 ($\widehat{P}^T < P^T, \widehat{P}^L <$

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

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

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 2008 Financial Crisis (see Section 5.1).

Since the corporate bonds are short-term with independent credit risks, their return rates (3) 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 (4)-(5). In this simplified model, the first-order condition for rating n bonds is given by

$$\underbrace{\mu_t^{r,n} - \mu_t^T}_{\text{credit spread}} = \underbrace{a\delta w_t^{I,n}(\nu^n)^2}_{\text{credit risk}} + \underbrace{(\zeta^n)^2 w_t^{I,n}}_{\text{regulatory premium}} = \gamma^n w_t^{I,n}, \tag{10}$$

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 the constant $\gamma^n := a\delta(\nu^n)^2 + (\zeta^n)^2 > 0$ measures the insurer's total cost of holding rating n bonds. We can then write equation (10) as

$$w_t^{I,n} = \frac{\mu_t^{r,n} - \mu_t^T}{\gamma^n}, \tag{11}$$

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.[14]

[14] 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.

Equation (11) 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 relative to the 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.

Proposition 1 *When the Treasury yield increases from y to $\widehat{y}$, if the life insurer faces a negative duration gap (i.e., $\frac{\partial \widehat{A}}{\partial \widehat{y}} > 0$), then the following predictions hold under Assumptions 1-2.*

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 $n > 0$.*

The magnitudes of the predictions increase in the severity of duration mismatch (i.e., $\frac{\partial \widehat{A}}{\partial \widehat{y}}$).

The proof is relegated to Appendix C.2. Proposition 1 summarizes the main analytical results from this model. It 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.

2.3 Model Extensions

Illustration of Quantitative Relevance. While the model presented so far is primarily analytical, I quantify the model in Appendix D and use a simple calibration to illustrate the quantitative importance of the channel. In the extension, I assume explicit functional forms for the bond supply function and the habitat demand. I then match key model parameters, including the life insurer's preferences, the bond market share of life insurers and habitat investors, and the habitat investors' demand elasticity, to existing estimates and empirical moments. The results show that the duration mismatch channel can potentially lead to sizable fluctuations in credit spreads, bond supply, and bond market composition following changes in the long-term interest rate.

Duration Management Motives. The existing literature has also pointed to a *duration management channel* where changes in interest rates affect insurers' demand for duration (e.g., Ozdagli and Wang, 2019; Yu, 2020). The duration management channel suggests that insurers tilt their bond portfolios toward longer-duration bonds when low interest rates worsen their duration gaps.

In Appendix E, I present an extension of the model that includes the duration management motives of life insurers to examine whether and how the duration mismatch channel proposed in this paper interacts with the duration management channel. In the extended model, life insurers first choose between bonds with different durations ("long-term vs. short term") and then choose between bonds with different credit ratings ("risky vs. safe") within each duration.

Analytically, I show that the duration management channel governs the first portfolio choice problem, while the duration mismatch channel governs the second portfolio choice problem. Indeed, the life insurer rebalances its portfolio towards longer-duration bonds following interest rate decreases. However, as shown by Domanski, Shin and Sushko (2017) and Ozdagli and Wang (2019), the reaching-for-duration behavior cannot entirely close the duration gap. Hence, the insurer still faces a negative duration gap in equilibrium, and declines in interest rates still erode the insurer's net worth. Accordingly, in the second portfolio choice problem, the insurer further rebalances its portfolio towards safer bonds, within each duration.

In summary, through this model extension, I show that the effects of the duration management channel and the duration mismatch channel are *separate*: the former determines the insurer's asset allocations *across* durations, while the latter determines the credit spread *within* each duration. If we focus on bonds of the same duration, the duration mismatch channel predicts that higher interest rates result in lower credit spreads, for any given duration.

3 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 at a daily frequency constructed by Liu and Wu (2021). 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, yields, outstanding amounts, issuance, and other characteristics. The dataset provides comprehensive coverage for publicly traded US corporate bonds between 2002 and 2022. 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, which consolidates the three ratings into one and sorts them into six NAIC categories.[15] 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 high-yield. Table 1 replicates Table 2 in Becker and Ivashina (2015), which summarizes the

[15]The rating scheme only applies to publicly traded bonds. For private placement bonds not rated by credit rating agencies, the ratings are assigned by NAIC directly.

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).

Bond Indices and CDS. For part of my empirical analysis, I use aggregate bond yield indices from the Intercontinental Exchange ("ICE"), which are accessed through FRED and began in 1997. I also use 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) between 2002 and 2019. I examine the Schedule D information in life insurers' regulatory reports to NAIC. 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. 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).

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 (pre-2007) and after the Financial Crisis (post-2009). 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.

4 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.

4.1 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

$$\Delta y_{it} = \alpha_i + \sum_{k=1}^{6} \beta_k \cdot \mathbf{1}_{\{\text{NAIC } k\}} \cdot \Delta y_t^{(10)} + \Gamma \mathbf{X}_{it} + \varepsilon_{it}, \tag{12}$$

where Δy_{it} is the change in bond i's yield from month t to $t+3$, $\Delta y_t^{(10)}$ is the change in the 10-year Treasury yield over the same period, 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 (pre-2007) and the sample after the Financial Crisis (post-2009).

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.

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.

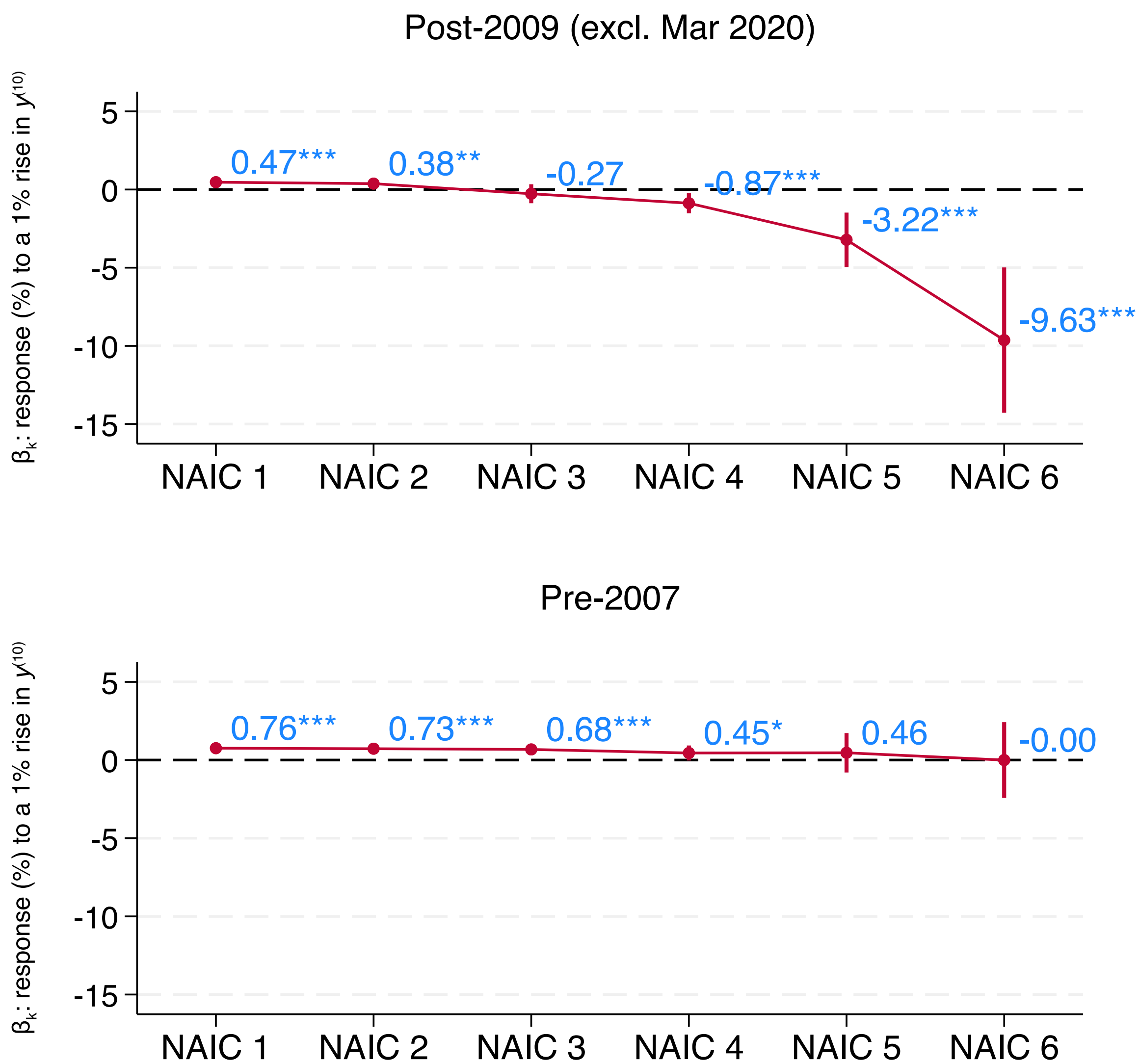

Figure 2. **10-year Treasury yield pass-through.**

This figure plots the coefficients β_k estimated from regression (12), controlling for the trading volume, outstanding amount, time to maturity, duration, credit ratings, coupon rate, 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 p-values are based on standard errors clustered at the issuer and year-month levels. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

4.2 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 (12), I further include duration-time fixed effects $\alpha_{D(i),t}$,

$$\Delta y_{it} = \alpha_i + \alpha_{D(i),t} + \sum_{k=2}^{6} \beta_k \cdot \mathbf{1}_{\{\text{NAIC } k\}} \cdot \Delta y_t^{(10)} + \Gamma \mathbf{X}_{it} + \varepsilon_{it}. \tag{13}$$

To create the duration-time fixed effects, I group bonds into small segments based on their Macaulay duration, each with a size of 1 year. I then interact these duration dummies with the observation date. The time 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. The interaction of time fixed effects with duration fixed effects further ensures that we construct credit spreads using bonds of similar durations.[16] 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

$$y_t^{\text{NAIC } k} = y_t^{\text{NAIC } 1} + (\text{Credit Spread})_t^{\text{NAIC } k}.$$

The effects of changes in the long-term Treasury yield 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 4.1, I have shown that

[16] As noted by van Binsbergen, Nozawa and Schwert (2025), it is important to control for the duration when studying the pricing of corporate bonds. Also, riskier bonds tend to be of shorter duration, which makes their prices less sensitive to Treasury yields. Figure A.5 further confirms that the results also hold when we compare bonds within maturity segments.

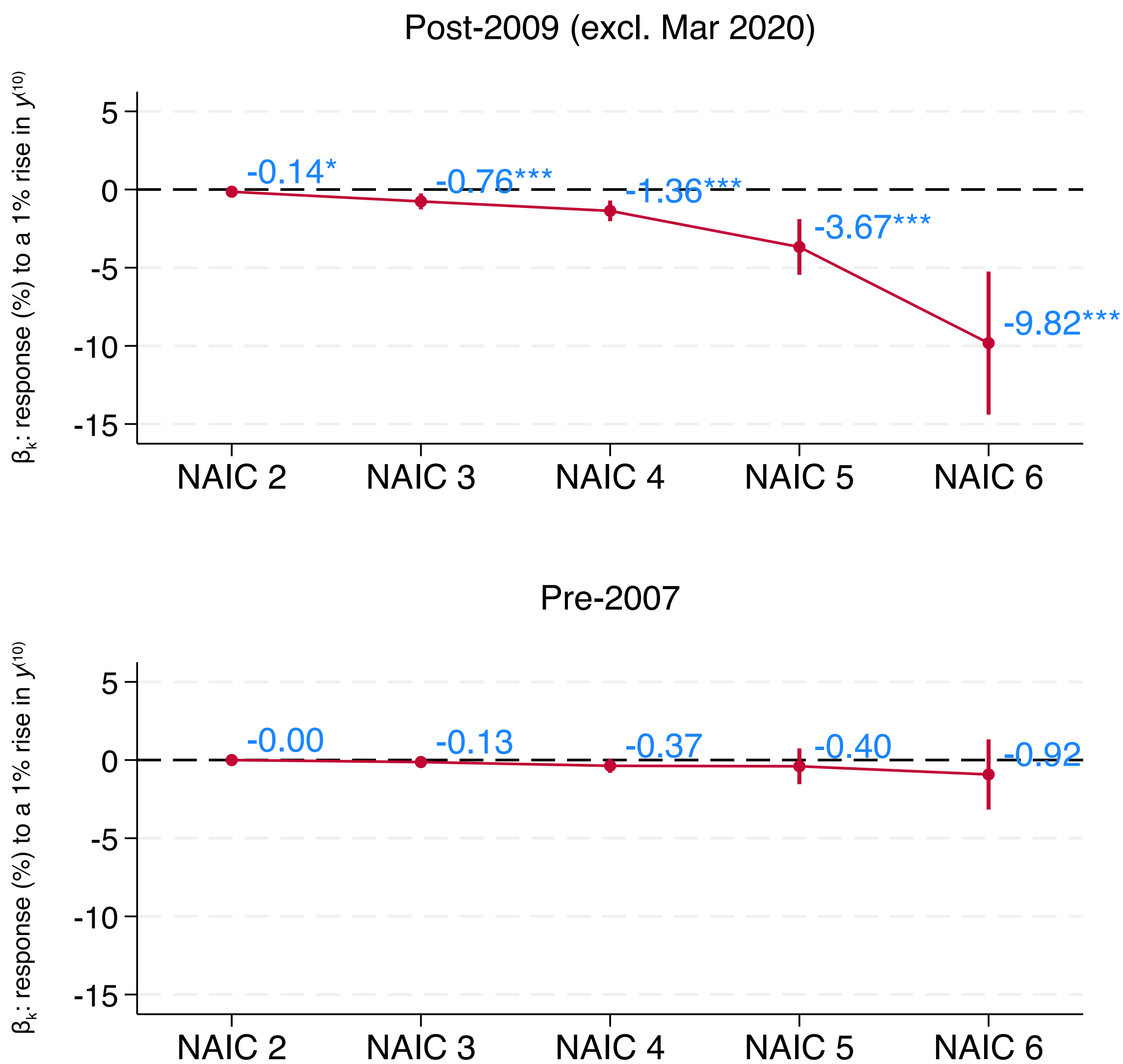

Figure 3. **Credit Spread Responses to the Long-term Interest Rate.**

This figure plots the coefficients β_k estimated from regression (13), controlling for the trading volume, outstanding amount, time to maturity, duration, credit ratings, coupon rate, 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 p-values are based on standard errors clustered at the issuer and year-month levels. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

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 92 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.

It is interesting to further examine whether the credit spread dynamics are driven by credit risk or other factors (such as insurers' risk-bearing capacity), as stronger economic growth can reduce firms' credit risk, which could further lead to both lower credit spreads and increased interest rates (e.g., Wu and Zhang, 2008). To account for variations in credit risk, I control for changes in bond issuers' expected default frequency (EDF) based on the Merton (1974) distance to default model following Bharath and Shumway (2008).[17] Figure A.4 shows the coefficients estimated in the sample of publicly traded firms after controlling for EDF, which exhibit the same pattern as in Fig-

[17] Bharath and Shumway (2008) propose a way of constructing Merton EDF for public firms using Compustat data. Using their method and the implementation by Gao (2022), I am able to compute the Merton EDF for 75.64% of observations in my dataset.

ure 3 and confirm that credit risk itself has little impact on the negative co-movements between long rates and credit spreads.

Another important potential confounding factor is the value of call options. As has been noted in Ma, Streitz and Tourre (2023), most corporate bonds issued recently have embedded call options. Typically, call options become more valuable when increases in interest rates lower bond prices. Hence, call options can potentially also contribute to the negative co-movements between long rates and credit spreads, if bonds with different credit ratings differ systemically in the moneyness of their call options (Longstaff and Schwartz, 1995; Duffee, 1998). To understand whether the findings in Figure 3 are mostly driven by call options, I collect the call prices and moneyness ratio (bond price / call price) of a large fraction of bonds in my sample.[18] Figure A.3 shows that the results remain largely the same after controlling for callability dummies and the interaction of moneyness and $\Delta y_t^{(10)}$ at the bond level. The results show that the value of call options can, at most, explain a small fraction of the credit spread dynamics of interest.

4.3 High-frequency Evidence: FOMC Announcements

Section 4.1 and Section 4.2 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-

[18] I am able to recover the moneyness for 66.44% of observations in my dataset. The procedure of constructing call prices is detailed in Appendix B.

frequency 10-year Treasury yield shocks. The regressions are as follows

$$\text{Spread}_{t+h}^{k} - \text{Spread}_{t-1}^{k} = \alpha_h + \beta_h \left(\Delta y_t^{(10)} \Big|_{\text{FOMC}} \right) + \varepsilon_{t,h}. \tag{14}$$

Here, the left-hand-side variable Spread_t^k is constructed as the difference between the option-adjusted spread (OAS) indices for rating k and the AAA rating,

$$\text{Spread}_t^k = \text{OAS}_t^k - \text{OAS}_t^{AAA}.$$

The index OAS_t^k is constructed as the weighted OASs of all bonds in rating k measured against the spot Treasury yield curve. The coefficient β_h then 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.

Figure 4 and Figure 5 plot the impulse response functions estimated from the local projections (14) 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.

Figure A.8 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 4.1, the yields of high-yield bonds (BB and lower) 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

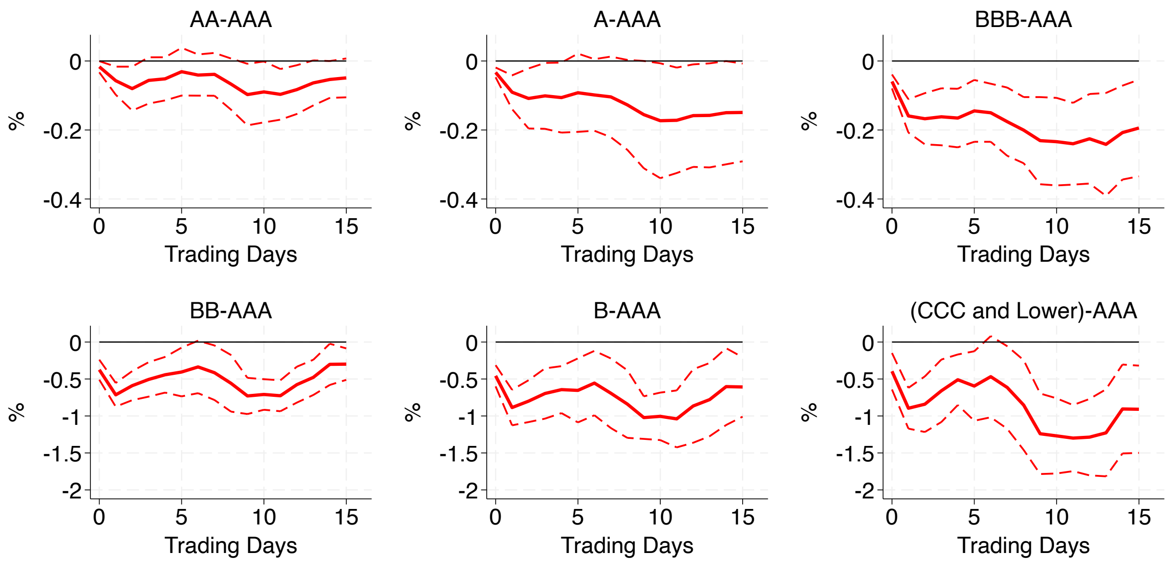

Figure 4. **Impulse Responses of Credit Spreads (2010-2022).**

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.

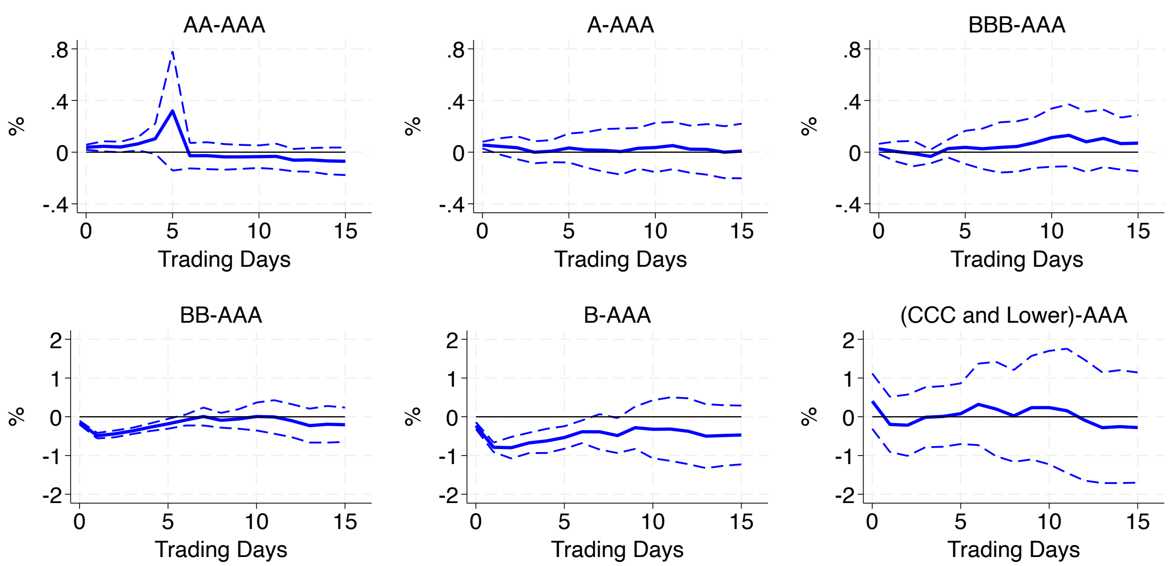

Figure 5. **Impulse Responses of Credit Spreads (1997-2007).**

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 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 credit risk and the long-run performance of corporate bonds. To control for shocks to credit risk, I further control the changes in the average CDS spread for each NAIC category against AAA bonds around FOMC meetings, which is a direct and real-time measure of the price of default risk (e.g., Longstaff, Mithal and Neis, 2005).[19] By controlling for CDS spreads, we can eliminate the impacts of changing credit risk and more clearly understand how long rates affect credit spreads.[20] Figure A.7 plots the post-crisis impulses responses estimated after controlling for short rate changes and CDS spreads, which again confirm the patterns shown in Figure 4.

Another potential confounding mechanism is the "Fed response to news" channel, where long rate increases around FOMC meetings reflect policy responses to improved economic conditions. However, while the "Fed response to news" channel is important for understanding short rate surprises around FOMC meetings (Bauer and Swanson, 2023), there is no evidence that it plays a significant role in the dynamics of long rates. Hillenbrand (2023) finds that changes in economic conditions before FOMC meetings do not predict changes in the 10-year Treasury yield around FOMC meetings. Gürkaynak, Sack and Swanson (2005) shows that long rates actually tend to fall in response to FOMC tightenings, which are more likely to occur after improved economic conditions. Indeed, Bauer and Swanson (2023) also argued that "an information effect is not needed to explain the response of long-term Treasury yields to FOMC announcements."

4.4 The Long-term Interest Rate and Bond Issuance

In previous sections, I documented that decreases in long-term interest rates are accompanied by increases in credit spreads, especially for high-yield bonds. Credit spreads are important indicators of firm borrowing costs — existing studies show that

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

[20] 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.

primary market bond prices (i.e., new bond issuance costs) are highly correlated with secondary market conditions (e.g., Coppola, 2024; 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 investment-grade (IG, NAIC 1-2) and high-yield (HY, NAIC 3-6) bonds in WRDS Bond Returns. I then estimate the following empirical specification

$$\Delta\text{Issuance Rate}_{k,t} = \alpha_t + \alpha_k + \beta \cdot \mathbf{1}_{\{k=\text{HY}\}} \cdot \Delta y_t^{(10)} + \Gamma \mathbf{X}_{kt} + \varepsilon_{kt}, \tag{15}$$

where $k \in \{\text{IG}, \text{HY}\}$, $\text{Issuance Rate}_{k,t}$ is defined as the total IG/HY issuance in the 6-month period after month t divided by the current total outstanding amount of IG/HY bonds at the end of the month t, $\Delta y_t^{(10)}$ is the change in the 10-year Treasury yield from month t to $t+3$, $\Delta\text{Issuance Rate}_{k,t}$ is the change in the issuance rate over the same period, and the controls $\mathbf{X}_{kt}$ include the recent default rates of IG and HY bonds. The rating fixed effects α_k capture the steady state difference in the growth rates of IG and HY markets. The time fixed effects α_t capture aggregate fluctuations in bond issuance. The coefficient β measures how the 10-year Treasury yield affects the *difference* in issuance between IG and HY bonds. The issuance rate of HY bonds increases by β relative to IG bonds when the 10-year Treasury yield increases by 1%.

Table 2 contains the estimated coefficients from regression (15). The second column reports the results after the Financial Crisis. It suggests that increases in the 10-year Treasury yield boost the issuance of high-yield bonds relative to investment-grade bonds. Following a 1% increase in the long-term interest rate, the issuance of high-yield (NAIC 3-6) bonds grows by more than 0.9% relative to the issuance of investment-grade (NAIC 1-2) 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. 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 4.1.

	Pre-2007	Post-2009
β	-0.136	0.939**
	[0.712]	[0.047]
NAIC FE	✓	✓
Time FE	✓	✓
R^2	.589	.725

Table 2. **Bond Issuance Responses to the Long-term Interest Rate.**

This table shows the coefficients estimated from regression (15), controlling for the recent default rates of each NAIC category and the average issuance maturity interacted with $\Delta y_t^{(10)}$. 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 p-values shown in brackets are based on standard errors clustered at the IG/HY category by year-month levels. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

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, Figure 6 shows that the corporate bond market has shifted towards the investment-grade segment during the low-interest-rate environment post-2008 (see also Mota, 2023).

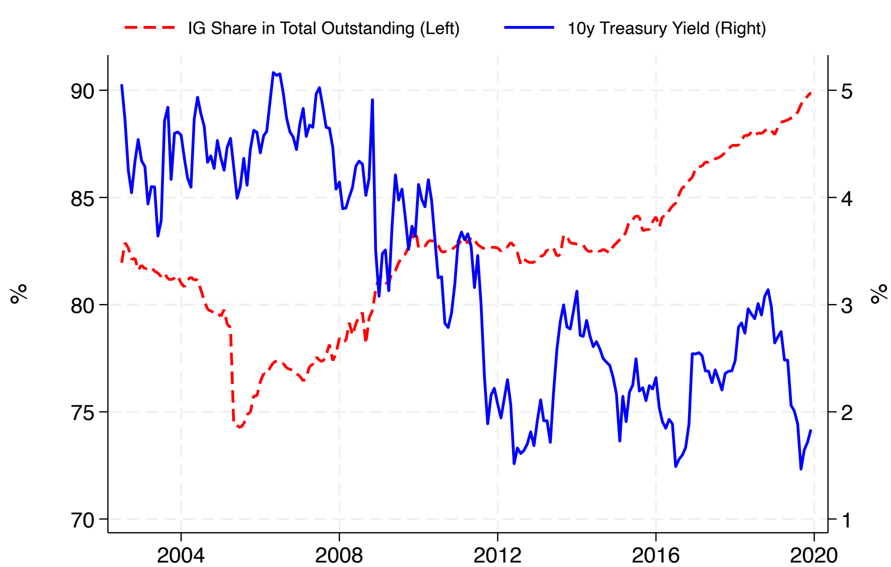

Figure 6. **Share of Investment-grade Bonds in WRDS Bond Returns.**

5 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 20% of all corporate bonds in recent years (Koijen and Yogo, 2023).

5.1 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 3 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 70% 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 estimate the following regression

$$\text{InsurerReturn}_t = \alpha + \beta \Delta y_t^{(10)} + \gamma \text{MarketReturn}_t + \text{Controls} + \varepsilon_t. \tag{16}$$

InsurerReturn_t is the stock return of the life insurance sector in week t, MarketReturn_t is the S&P 500 return in week t, 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 changes in the 10-year Treasury yield. If the duration of life insurers' liabilities exceeds

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 3. **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.

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 4 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 6.008 suggests that life insurers' equity value grows by more than 6% when the 10-year rate increases by 1%, so life insurers face severe duration mismatch after the 2007-2008 Financial Crisis.

Figure 7 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 long rate is high, as their liabilities become less expensive and their equity increases.

	Pre-2007	Post-2019
$\Delta y_t^{(10)}$	-0.0723	6.008***
	[0.947]	[0.000]
S&P 500 Return	✓	✓
$\Delta y_t^{(1m)}$	✓	✓
Observations	260	663

Table 4. **Life Insurers' Duration Mismatch.**

This table shows the coefficients estimated from regression (16) between 2002 and 2022, 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 p-values shown in brackets are based on robust standard errors. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

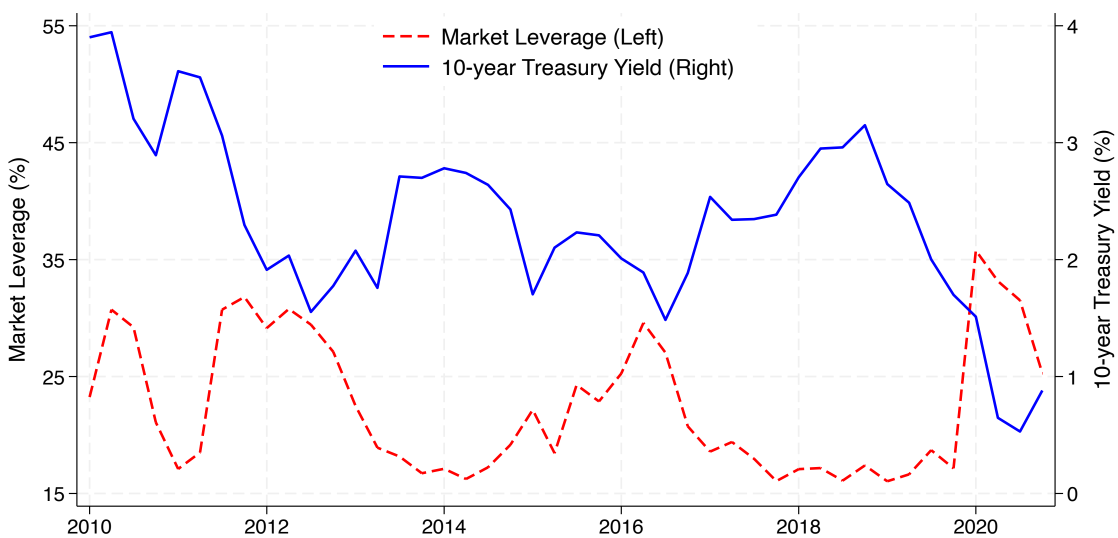

Figure 7. **Life Insurers' Market Leverage.**

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.

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 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.[21] 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, 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.[22]

5.2 Life Insurers and Credit Spread Dynamics

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 4.2 il-

[21]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).

[22]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).

lustrates 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.

First, I provide *suggestive* evidence that life insurers matter for credit spread dynamics by estimating regression (13) in Section 4.2 separately for two sub-samples: bonds that life insurance companies hold and bonds that life insurance companies do not hold. Figure 8 plots the estimated coefficients from the two separate regressions for the post-crisis period. Indeed, the negative relationship between credit spreads and long rates in Figure 3 is only present in bonds with positive life insurance ownership. Bonds with no life insurance ownership mostly exhibit no significant responses (NAIC 2, 4, and 5) or positive responses (NAIC 3). Among bonds not held by life insurers, only those in NAIC 6 (less than 0.5% of the bond market) show negative responses (yet the magnitude is much smaller compared to Figure 3).[23]

To further establish the link between life insurers' duration mismatch and credit spread dynamics, I augment regression (13) with an interaction term between insurers' risk exposure and $\Delta y_t^{(10)}$,

$$\Delta y_{it} = \alpha_i + \alpha_{D(i),t} + \sum_{k=2}^{6} \left(\beta_k + \gamma_k \text{Exposure}_t\right) \cdot \mathbf{1}_{\{\text{NAIC } k\}} \cdot \Delta y_t^{(10)} + \Gamma \mathbf{X}_{it} + \varepsilon_{it}. \tag{17}$$

where Exposure_t is the insurers' exposure to long rates estimated from (16) using data one year prior to date t. If estimated $\gamma_k < 0$, we know that the co-movements between long rates and credit spreads become more negative when insurers have more negative duration gaps. Figure A.12 summaries the γ coefficients estimated from regression (17), which again confirms that the co-movements between long rates and credit spreads are negatively linked to the size of the life insurance sector's duration gap. Moreover, the negative association is stronger for lower credit ratings, which is consistent with the findings of Section 4.2.

[23]Figure A.6 further shows that, before the Financial Crisis, the co-movement between credit spreads and long rates was 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-crisis.

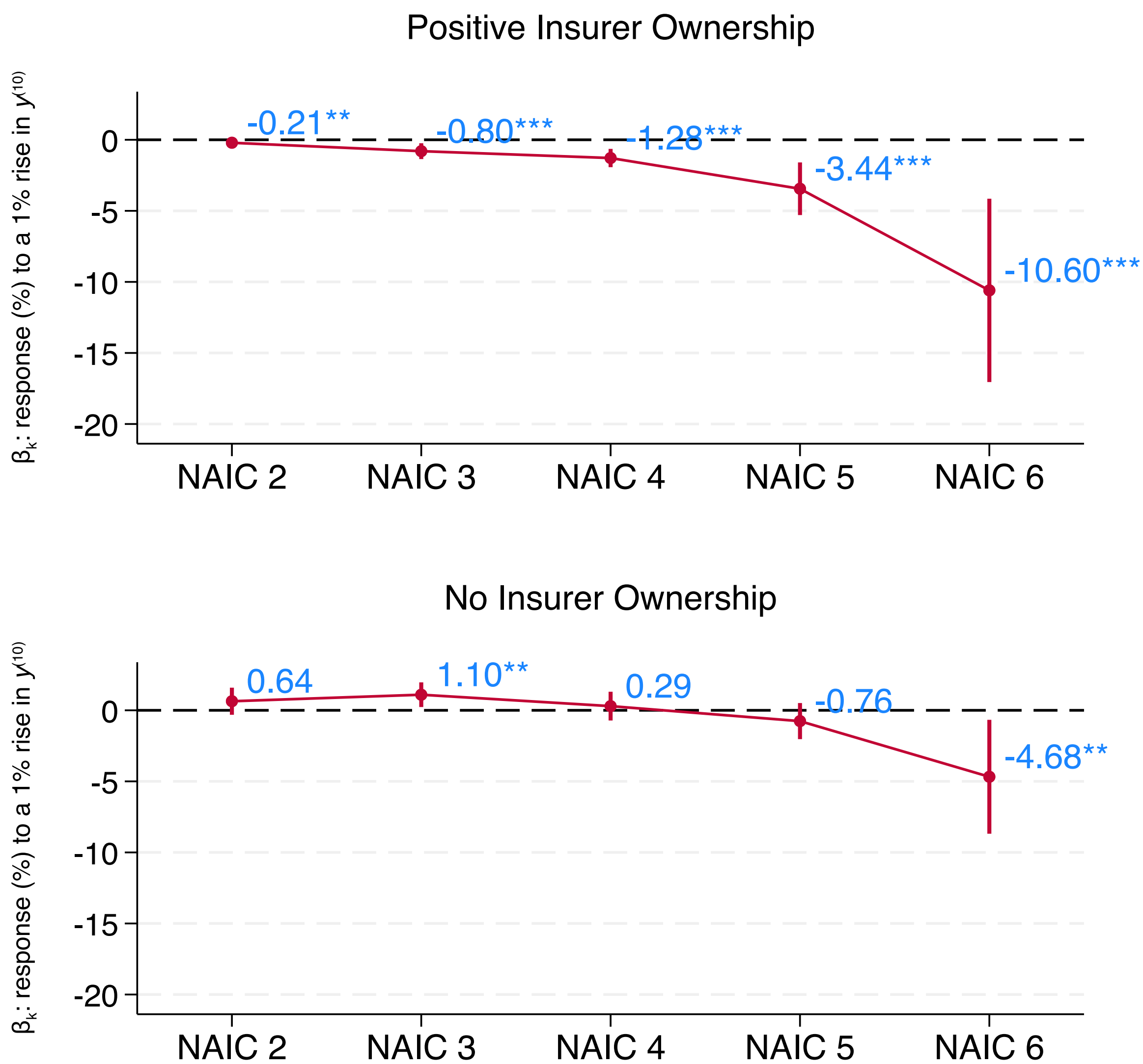

Figure 8. **Credit Spread Responses by Life Insurance Ownership.**

This figure plots the coefficients β_k estimated from regression (13) for the period of 2010-2019, controlling for the trading volume, outstanding amount, time to maturity, duration, credit ratings, coupon rate, 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 the sub-sample of bonds with life insurance ownership. The bottom panel shows estimates for the sub-sample of bonds without life insurance ownership. The p-values are based on standard errors clustered at the issuer and year-month levels. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

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 4.2.[24] To further sharpen the identification, I utilize an exogenous discontinuity in investor composition stemming from mutual funds' investment mandates (e.g., Li and Yu, 2024; Bretscher, Schmid and Ye, 2023). Bai, Li and Manela (2023) and Li and Yu (2024) 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 insurance companies, jumps downward.

Figure 9 visualizes the discontinuity in life insurers' ownership share around the maturity threshold of 10 years and a quarter.[25] 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. Figure A.10 and Figure A.11 show that the discontinuity disappears if we use other cutoff values such as 9 and 11 years.

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 $c = 10.25$,

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

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.

[24] 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.

[25] 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.

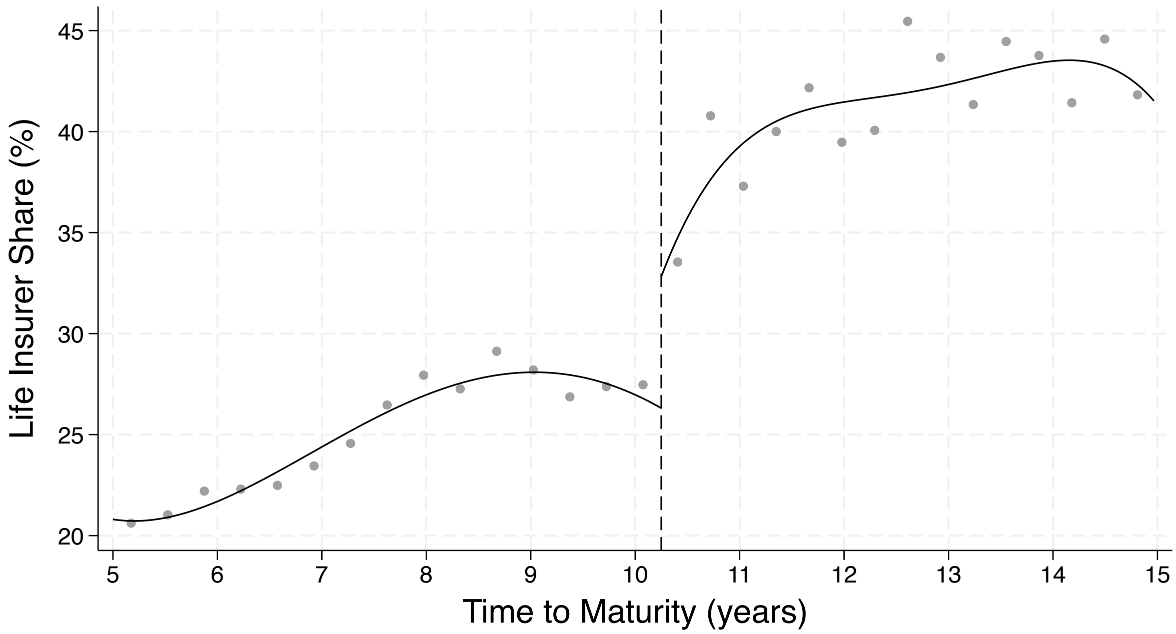

Figure 9. **Discontinuity in Investor Composition.**

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

Table 5 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.

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 use the 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 compare the co-movement seen in bonds slightly above the threshold to that seen in bonds slightly below the threshold. Specifically, I estimate the following two-stage regression

$$\begin{aligned} \varphi_{it}^{\text{Ins}} &= \alpha + \delta \cdot \mathbf{1}_{\{\text{maturity}_{it} > c\}} + \Gamma \mathbf{X}_{it} + \varepsilon_{it} \\ \Delta y_{it} &= \alpha_t + \left(\beta + \gamma \varphi_{it}^{\text{Ins}}\right) \mathbf{1}_{\{\text{NAIC } 3-6\}} \Delta y_t^{(10)} + \Gamma \mathbf{X}_{it} + \varepsilon_{it}. \end{aligned} \tag{18}$$

Method	δ	p-value	[95% Conf. Interval]
OLS	4.73	0.000	[4.15, 5.32]
RDD, Conventional	4.43	0.000	[2.46, 6.39]
RDD, Bias-corrected	4.52	0.000	[2.55, 6.48]
RDD, Bias-corrected, Robust	4.52	0.000	[2.22, 6.81]

Table 5. **Testing the Discontinuity.**

This table shows the coefficient on $\mathbf{1}_{\{maturity_{it}>c\}}$, controlling for the trading volume, duration, maturity, size, coupon rate, and coupon frequency 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).

The first stage estimates the effect of the maturity cutoff on the life insurer share $\varphi_{it}^{\text{Ins}}$. I then use the optimal bandwidth from the first stage (2.446 years) following Calonico, Cattaneo and Titiunik (2014) in the second stage.[26]

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.[27] In the second stage, I focus on the credit spread between investment-grade bonds (NAIC 1-2) and high-yield bonds (NAIC 3-6). The time fixed effects α_t absorb the average responses of IG bond yields to the 10-year Treasury yield. The term $(\beta + \gamma\varphi_{it}^{\text{Ins}})\mathbf{1}_{\{\text{NAIC 3}-6\}}\Delta y_t^{(10)}$ captures the responses of IG-HY credit spreads to the 10-year Treasury yield. Further, we can decompose the credit spread responses into two components: the first term $\beta \cdot \mathbf{1}_{\{\text{NAIC 3}-6\}}\Delta y_t^{(10)}$ captures the responses of credit spreads of bonds not held by life insurance companies, and the interaction term $\gamma \cdot \varphi_{it}^{\text{Ins}}\mathbf{1}_{\{\text{NAIC 3}-6\}}\Delta y_t^{(10)}$ captures the degree to which the credit spread responses depend on the bonds' life insurance ownership.

Table 6 presents the estimates from the RDD regressions above. The first-stage F-stats both before and after the Financial Crisis are well above the conventional threshold for strong instruments in Stock and Yogo (2005), again confirming the validity of the discontinuity in bond ownership structure.

[26]Figure A.9 shows the estimates using alternative bandwidths around the optimal bandwidth.

[27]Specifically, I instrument $\varphi_{it}^{\text{Ins}}\mathbf{1}_{\{\text{NAIC 3}-6\}}\Delta y_t^{(10)}$ with $\mathbf{1}_{\{\text{maturity}_{it}>c\}}\mathbf{1}_{\{\text{NAIC 3}-6\}}\Delta y_t^{(10)}$.

	Pre-2007	Post-2009
γ	-1.529	-13.81***
	[0.593]	[0.001]
Controls	✓	✓
Time FE	✓	✓
Kleibergen-Paap F-stat	131.927	79.925
Observations	4447	10795

Table 6. **RDD Regressions.**

This table shows the coefficients estimated from regression (18), with the instrument, controlling for the trading volume, time to maturity, duration, maturity, size, coupon rate and frequency, credit rating, and insurer ownership share 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 p-values shown in brackets are based on standard errors clustered at the issuer and year-month levels. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

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. Consider two high-yield bonds whose life insurance ownership is 3% and 10% (a one-standard-deviation difference), respectively. The estimates suggest that the credit spread of the second bond would decline by $13.81 \times (10\% - 3\%) = 0.97\%$ more than the first when the 10-year rate increases by 1%. It is worth noting that the special role of life insurers began after the Financial Crisis, as the coefficient on $\varphi_{it}^{\text{Ins}} y_t^{(10)}$ is statistically insignificant before the GFC.

A threat to identification is that bond issuers might try to cater to mutual fund investors by issuing bonds just below 10 years, so the bonds on the two sides of the cutoff are systematically different. To alleviate such concerns, I conduct two additional robustness checks. The first exercise excludes recently issued bonds, and the second focuses on bonds whose maturity is above the maturity cutoff at issuance. Table A.2 shows that the results survive these robustness checks.[28]

[28] Table A.3 further shows that the results remain the same after controlling for bonds' callability and moneyess as in Section 4.2.

5.3 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 5.1). 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.

To study changes in insurers' portfolio demand, I first construct the share of their bond purchases in the high-yield market,

$$\text{(Purchase Share)}_t^{\text{HY}} = \frac{\text{Net Purchase}_t^{\text{HY}}}{\text{Net Purchase}_t^{\text{Total}}}.$$

Here $\text{Net Purchase}_t^{\text{HY}}$ is life insurers' net purchases of high-yield (NAIC 3-6) bonds in the 3-month period after month t,[29] $\text{Net Purchase}_t^{\text{Total}}$ is life insurers' net purchases of all bonds during the same period, and $\text{(Purchase Share)}_t^{\text{HY}}$ measures the fraction of insurers' bond purchases in HY segments.

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

$$\Delta\text{(Purchase Share)}_t^{\text{HY}} = \alpha + \beta \cdot \Delta y_t^{(10)} + \Gamma \mathbf{X}_t + \varepsilon_t. \tag{19}$$

In this specification, $\Delta y_t^{(10)}$ is the change in the 10-year Treasury yield from month t to month $t+3$. The outcome variable is the change in the HY purchase share from month t to $t+3$ or $t+6$. The coefficient β measures how much life insurers shift their bond purchases toward HY bonds following a 1% rise in the 10-year Treasury yield. Table 7 shows the estimated coefficients from regression (19). After an increase in the 10-year yield, life insurers tilt their bond portfolio towards riskier bonds after the GFC, but not before, again confirming the predictions of the duration mismatch channel.

Next, I investigate whether the portfolio rebalancing pattern also holds in the cross-section of life insurers, i.e., whether insurers with greater exposure to interest rates

[29] Net purchases are defined as the total bond purchases net of the total bond disposals for any reasons, using all transactions recorded in NAIC regulatory filings. After constructing the bond transaction records, I connect them to WRDS Bond Returns for credit rating information.

	Pre-2007		Post-2009	
	$h = 3m$	$h = 6m$	$h = 3m$	$h = 6m$
β	-0.269	0.562	0.750*	2.346***
	[0.571]	[0.280]	[0.071]	[0.000]
R^2	.537	.723	.305	.387
Observations	54	54	114	111

Table 7. **Long Rates and the Composition of Insurers' Bond Purchases.**

This table shows the coefficients estimated from regression (19), controlling for changes in the one-month Treasury yield and month and year fixed effects. h denotes the horizon of change for Δ*(Purchase Share)*$_t^{HY}$*. The p-values shown in brackets are based on standard errors clustered at the month level.* $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

adjust their bond purchases more strongly. As revealed by the previous literature (e.g., Koijen and Yogo, 2022; Ellul et al., 2022; Sen, 2023), a particular type of product sold by life insurers — variable annuities (VAs) — has extremely high duration and convexity due to its minimum return guarantees and embedded options. Consequently, insurers that issued large quantities of VAs pre-GFC ended up facing more severe duration mismatch post-GFC. In light of the link between VA liabilities and insurers' risk exposure, I re-run the previous regression at the individual insurer level, further adding an interaction term between $\Delta y_t^{(10)}$ and insurers' VA liability shares in 2009, the beginning of the post-GFC sample,[30]

$$\Delta(\text{Purchase Share})_t^{\text{HY}} = \alpha_j + \alpha_t + \beta \cdot (\text{VA Share})_{j,2009} \cdot \Delta y_t^{(10)} + \Gamma \mathbf{X}_{jt} + \varepsilon_{jt}. \tag{20}$$

Table 8 shows the results from regression (20). I have added insurer and time fixed effects to further control for any factors that are insurer-specific or insurer-specific. The results show that the insurers with stronger exposures to interest rates (i.e., those that issued more VAs) exhibited larger portfolio rebalancing towards HY bonds following long rate increases in the post-GFC period. The results verify our hypothesis that the duration mismatch of life insurers drives credit spread dynamics through its impact on insurers' bond demand.

In order to assess the magnitude and quantitative significance of insurers' portfolio

[30] I adopt the same definition of the VA liability share as in Koijen and Yogo (2022).

	$h = 3m$		$h = 6m$	
$(\text{VA Share})_{j,2009} \cdot \Delta y_t^{(10)}$	0.133^{**}	0.152^{**}	0.926^{***}	0.963^{***}
	[0.016]	[0.012]	[0.000]	[0.000]
Insurer FE	✓	✓	✓	✓
Time FE		✓		✓
R^2	.009	.021	.019	.034
Observations	27518	27518	23755	23755

Table 8. **Duration Mismatch and Insurers' Bond Purchases.**

This table shows the coefficients estimated from regression (20). The first column controls for insurer fixed effects, changes in 10-year and one-month Treasury yields, and changes in the one-month Treasury yield interacted with (VA Share)$_{j,2009}$. The second column controls for insurer fixed effects, time fixed effects, and changes in the one-month Treasury yield interacted with (VA Share)$_{j,2009}$. The variable (VA Share)$_{j,2009}$ is standardized. h denotes the horizon of change for Δ(Purchase Share)$_t^{HY}$. The p-values shown in brackets are based on standard errors clustered at the insurer level. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

rebalancing, I adopt the following specification to estimate the *difference* in the quantities of bonds purchased across credit ratings

$$\Delta \text{Holdings}_t^{\text{NAIC } k} = \alpha_t + \sum_k \beta_k \cdot \mathbf{1}_{\{\text{NAIC } k\}} \cdot \Delta y_t^{(10)} + \Gamma \mathbf{X}_t + \varepsilon_t. \tag{21}$$

The dependent variable $\Delta \text{Holdings}_t^{\text{NAIC } k}$ is the change in insurers' holdings of NAIC k bonds over the 3 months after month t, which by definition equals Net $\text{Purchase}_t^{\text{NAIC } k}$, life insurers' net purchases in NAIC k bonds. The time fixed effect α_t absorbs the average purchases in IG bonds, and the coefficient β_k measures the amount of insurers' net purchases in NAIC k that is *in excess of* their purchases of IG bonds, following a 1% positive increase in the 10-year rate. The results are presented in Table 9, which again verify that insurers increase their purchases of HY bonds relative to IG bonds following increases in the long rate.

In the following, I perform simple back-of-the-envelope computations based on Table 9 to gauge the impact of insurers' bond demand on credit spreads. As implied by Table 9, after a 1% rise in the long rate, insurers purchase $6.112 + 6.900 + 7.268 = \$20.28$ bn of high-yield (NAIC 3-6) bonds in excess of their purchases of investment-grade bonds over the next three months. The average total outstanding amount of high-yield

	Pre-2007	Post-2009
β_3	0.712	6.112**
	[0.855]	[0.017]
β_4	0.549	6.900**
	[0.892]	[0.014]
β_{5-6}	0.659	7.268**
	[0.878]	[0.013]
Time FE	✓	✓
R^2	.162	.108
Observations	270	582

Table 9. **Duration Mismatch and Insurers' Bond Purchases.**

This table shows the coefficients estimated from regression (21), controlling for changes in the one-month Treasury yield interacted with NAIC ratings. NAIC groups 5 and 6 are combined as the market size for NAIC 6 is too small to be estimated individually. The p-values shown in brackets are based on robust standard errors. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

bonds in WRDS Bond Returns is $727.6 bn in my sample (Table A.1). Therefore, the excess purchases of HY bonds amount to 2.8% of the aggregate HY market. Darmouni, Siani and Xiao (2025) estimate the bond demand elasticity of active mutual funds to be 0.75. Suppose that life insurers mainly trade against mutual funds in the HY market. Then, the price impact of the excess purchases would be about 2.8%/0.75 = 3.73%. Because the average duration of HY corporate bonds in WRDS Bond Returns is 4.45 years, a 3.73% price impact could translate to a 3.73%/4.45 = 0.84% change in the credit spread between IG and HY bonds.

Next, I compare the 0.84% change in the HY-IG spread to its empirical counterpart. According to Table A.1, in my sample, the share of NAIC 3, 4, 5, and 6 bonds in the HY market is 54.5%, 35.2%, 9.7%, and 0.95%, respectively, and the share of NAIC 2 among IG bonds is 44.3%. The estimates shown in Figure 3 then imply that the response of the average HY-IG spread to a 1% increase in the 10-year rate is

$$\begin{aligned}&\Delta(\text{HY-IG Spread})\\ &= \Delta(\text{Average HY Yield}) - \Delta(\text{Average IG Yield})\end{aligned}$$

$$= \sum_{k=3}^{6} (\text{NAIC } k \text{ Share in HY}) \Delta y^{\text{NAIC } k} - \sum_{k=1}^{2} (\text{NAIC } k \text{ Share in IG}) \Delta y^{\text{NAIC } k}$$

$$= \sum_{k=3}^{6} (\text{NAIC } k \text{ Share in HY}) \Delta (y^{\text{NAIC } k} - y^{\text{NAIC } 1}) - (\text{NAIC 2 Share in IG}) \Delta (y^{\text{NAIC } 2} - y^{\text{NAIC } 1})$$

$$= 54.5\%(-0.76\%) + 35.2\%(-1.36\%) + 9.7\%(-3.67\%) + 0.95\%(-9.82\%) - 44.3\%(-0.14\%)$$

$$= -1.27\%.$$

Therefore, the 0.84% decline in the HY-IG spread derived from the back-of-the-envelope calculation accounts for the majority (around 2/3) of the empirical co-movement shown in Figure 3.[31]

The back-of-the-envelope calculations rely on two simplifying assumptions: (1) the net purchases of investment-grade bonds are zero,[32] and (2) the average demand elasticity of the corporate bond market is the same as the elasticity of active mutual funds. While the precision can potentially be improved, these results provide approximations of the plausible magnitudes of how much life insurers' bond purchases can move bond yields, and the results are very much in line with the main empirical estimates.

6 Conclusion

Following the 2007-2008 Financial Crisis, US life insurers faced severe duration mismatch, and their equity value declined sharply with reductions in long-term interest rates. In this paper, I propose a *duration mismatch channel* where declines in long rates squeeze the balance sheet equity of life insurers and thereby increase their effective risk aversion, resulting in higher equilibrium credit spreads. To provide empirical support for the duration mismatch channel, 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

[31] It is likely that the duration mismatch channel of life insurers does not explain the entirety of the co-movement between long rates and credit spreads. The unexplained variations in credit spread dynamics might be driven by other factors, such as the duration mismatch of other investors (e.g., pension funds).

[32] This assumption likely biases the inferred yield impacts downwards. Since investment-grade bonds have larger durations and amounts outstanding than high-yield bonds, the same purchases will have smaller yield impacts on them. Hence, considering purchases in NAIC 1 and 2 would likely increase the magnitude of the response in the HY-IG spread.

Crisis, when life insurers started to face negative duration gaps. I further present causal evidence that this co-movement is more pronounced in bonds held by life insurers using an exogenous discontinuity in bonds' ownership structure. The duration mismatch channel can have significant potential impacts on the bond market and monetary policy transmission, given the long-run trend of falling interest rates and the increasing prevalence of unconventional monetary policies (e.g., central bank asset purchases and forward guidance) that specifically target long-term interest rates.

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Appendix

A Supplementary Figures

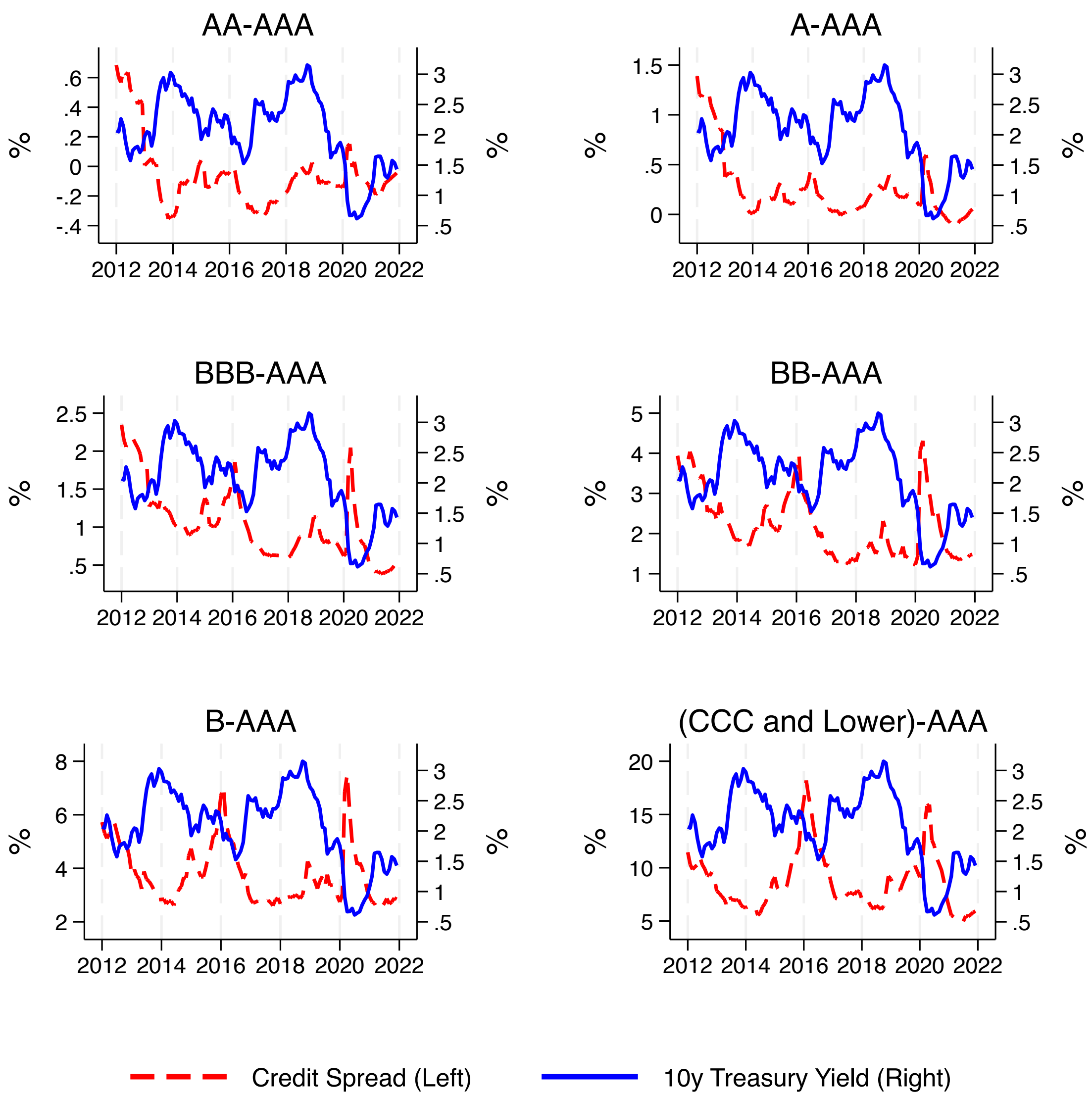

Figure A.1. **The 10-year Treasury Yield and Bond Credit Spreads (2010-2019).**
This figure plots the average monthly 10-year Treasury yield and the spreads of corporate bond indices relative to the index of AAA bonds. The 10-year Treasury yield is shown as the solid blue line and corresponds to the 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.

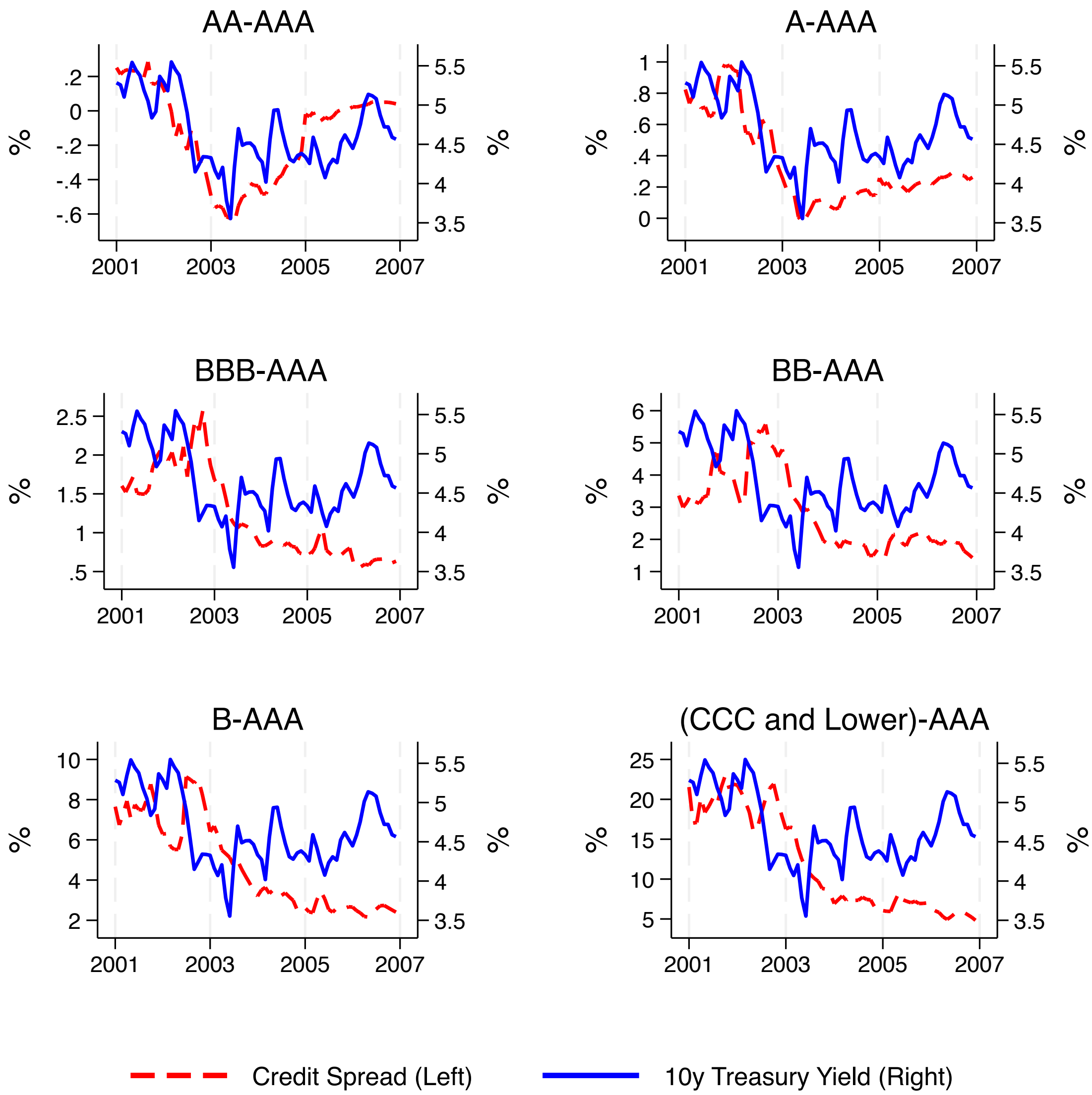

Figure A.2. **The 10-year Treasury Yield and Bond Credit Spreads (2001-2007).** *This figure plots the average monthly 10-year Treasury yield and the spreads of corporate bond indices relative to the index of AAA bonds. The 10-year Treasury yield is shown as the solid blue line and corresponds to the 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.*

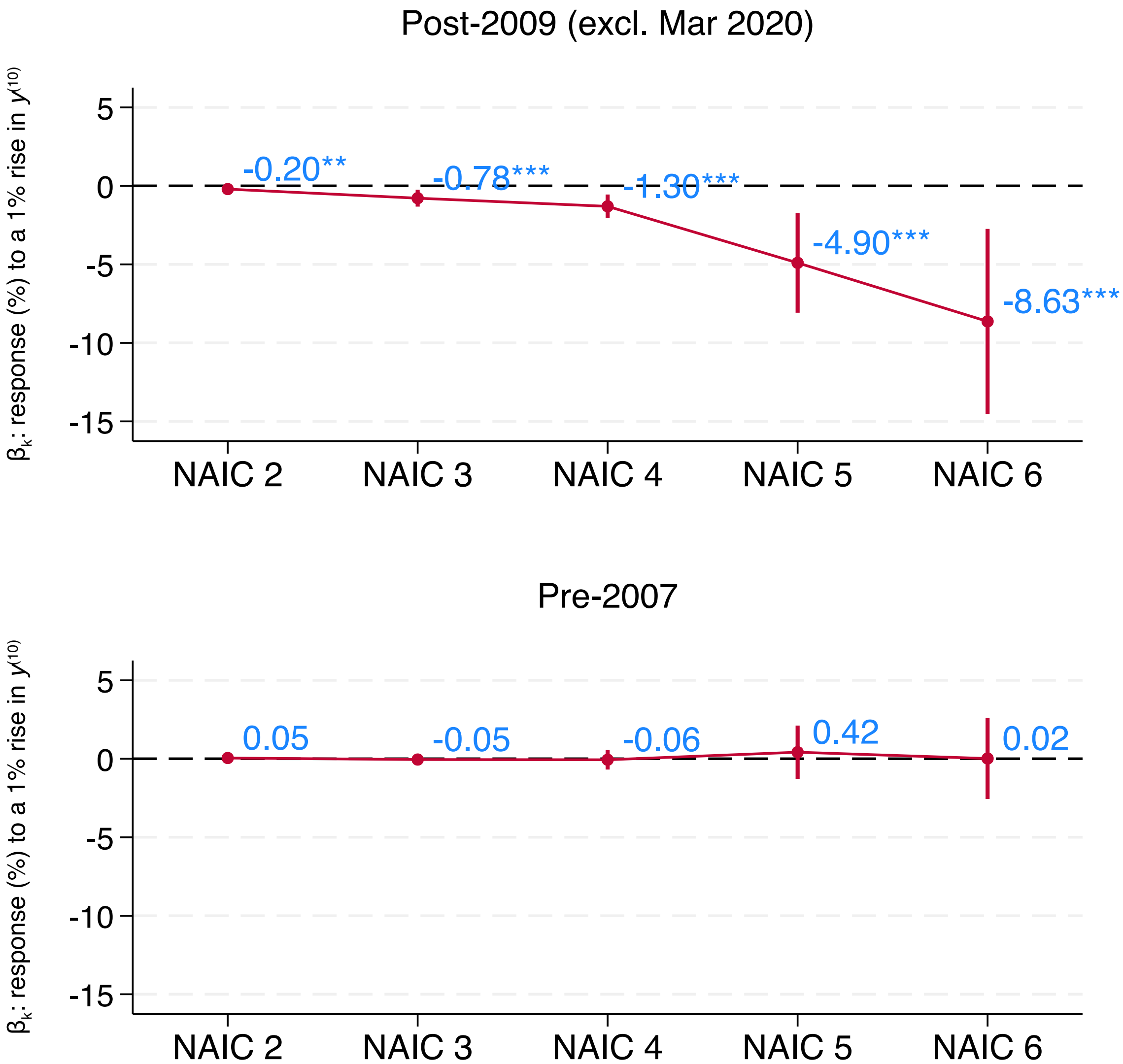

Figure A.3. **Credit Spread Responses (Controlling for Callability).**

This figure plots the coefficients β_k estimated from regression (13), controlling for the trading volume, outstanding amount, time to maturity, duration, credit ratings, coupon rate, and coupon frequency of each bond, the recent default rate of each NAIC category, ***callability dummies****, and* ***the interaction between the bonds' moneyness ratio and*** $\Delta y_t^{(10)}$. *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 p-values are based on standard errors clustered at the issuer and year-month levels.* $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

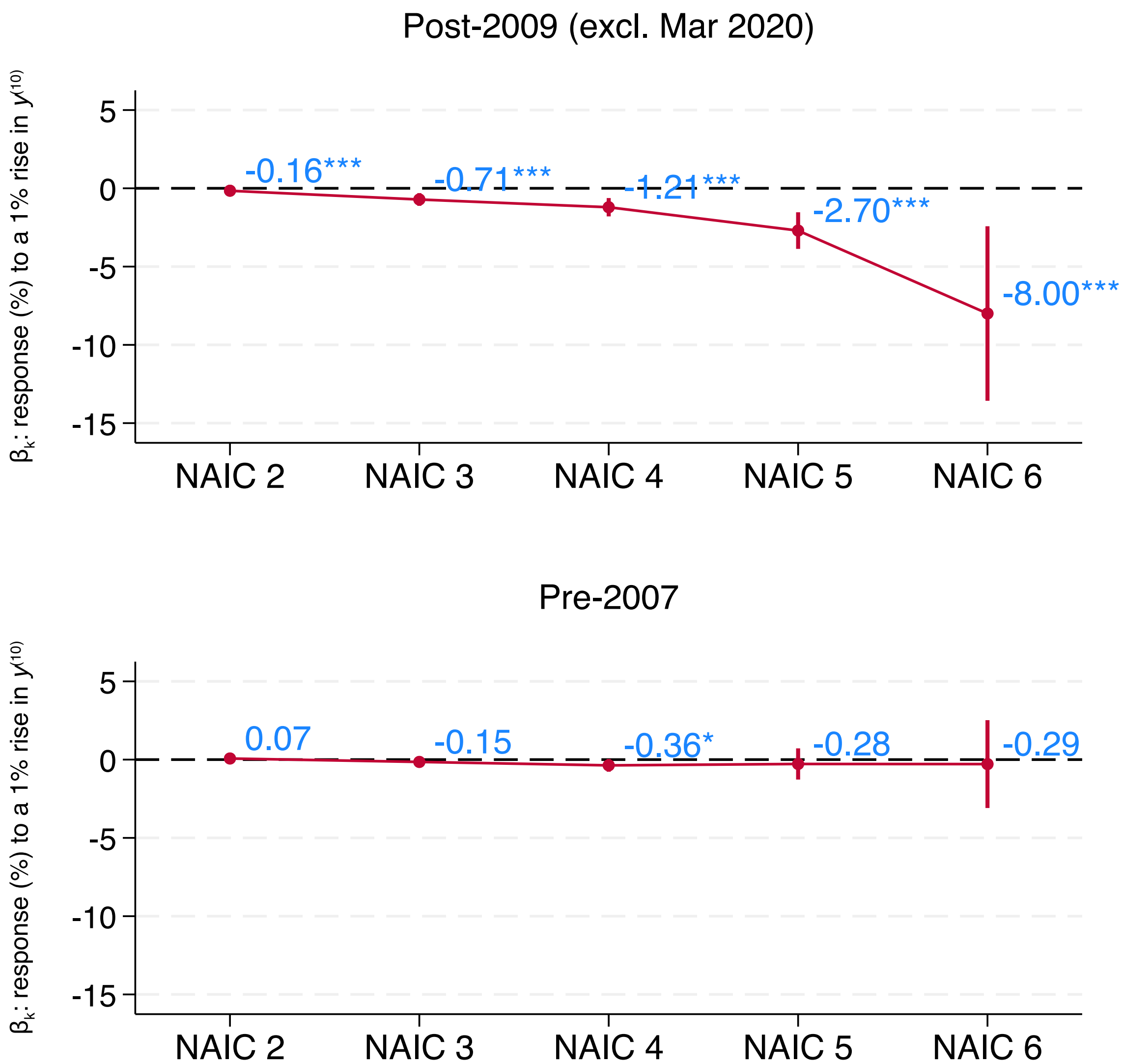

Figure A.4. **Credit Spread Responses (Controlling for Merton EDF).**

This figure plots the coefficients β_k estimated from regression (13), controlling for the trading volume, outstanding amount, time to maturity, duration, credit ratings, coupon rate, and coupon frequency of each bond, the recent default rate of each NAIC category, and ***changes in the expected default frequency (EDF)*** *for each issuer. 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 p-values shown in brackets are based on standard errors clustered at the issuer and year-month levels. The p-values shown in brackets are based on standard errors clustered at the issuer and year-month levels. The p-values are based on standard errors clustered at the issuer and year-month levels.* $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

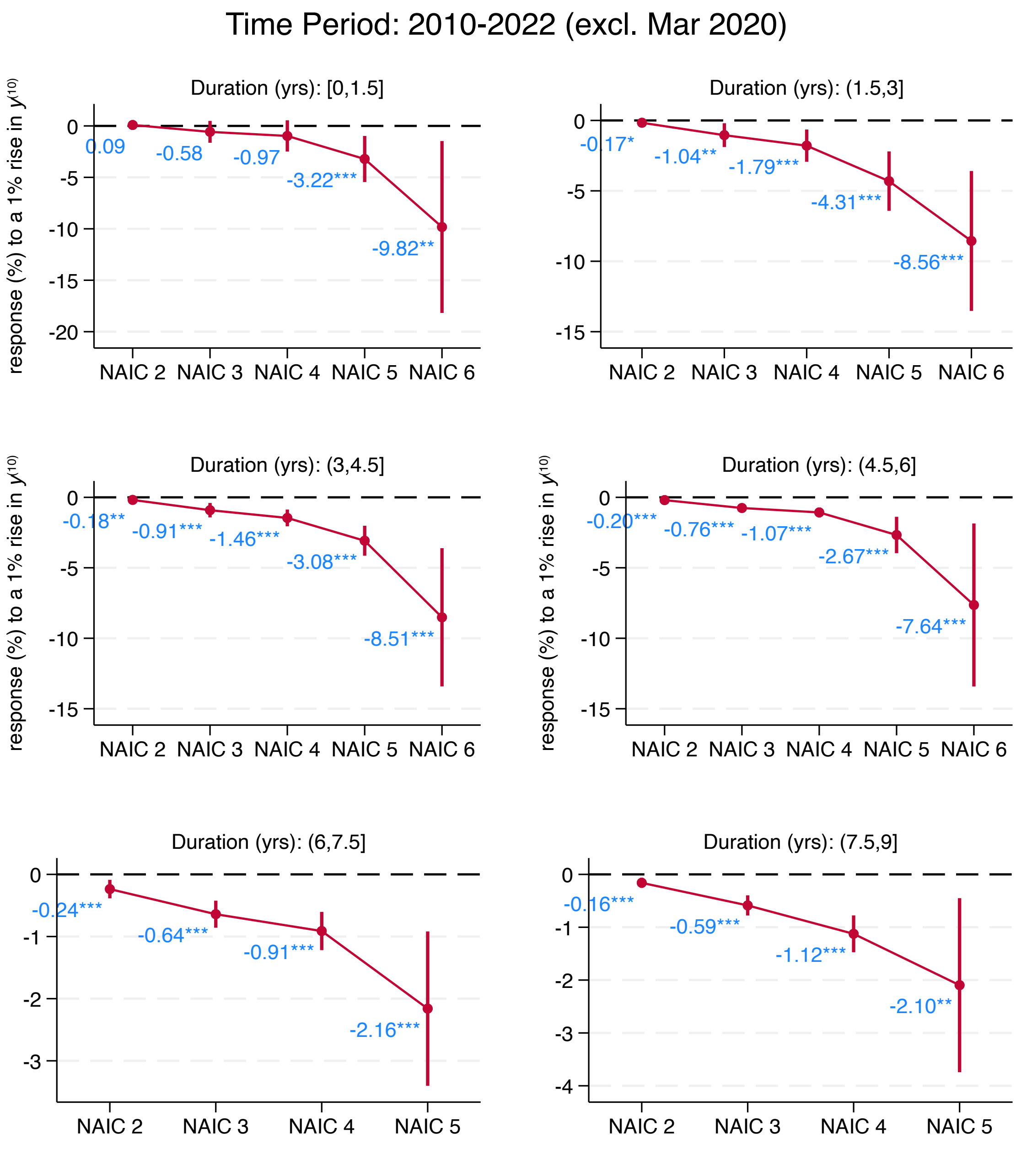

Figure A.5. **Credit Spread Responses in Different Duration Groups (post-2009).**

This figure plots the coefficients β_k estimated from regression (13), controlling for the trading volume, outstanding amount, time to maturity, duration, credit ratings, coupon rate, 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 different panels estimate coefficients for bonds in different duration segments. The NAIC 6 group is omitted in the last two panels because there are too few NAIC 6 bonds with a sufficiently long duration. The p-values are based on standard errors clustered at the issuer and year-month levels. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

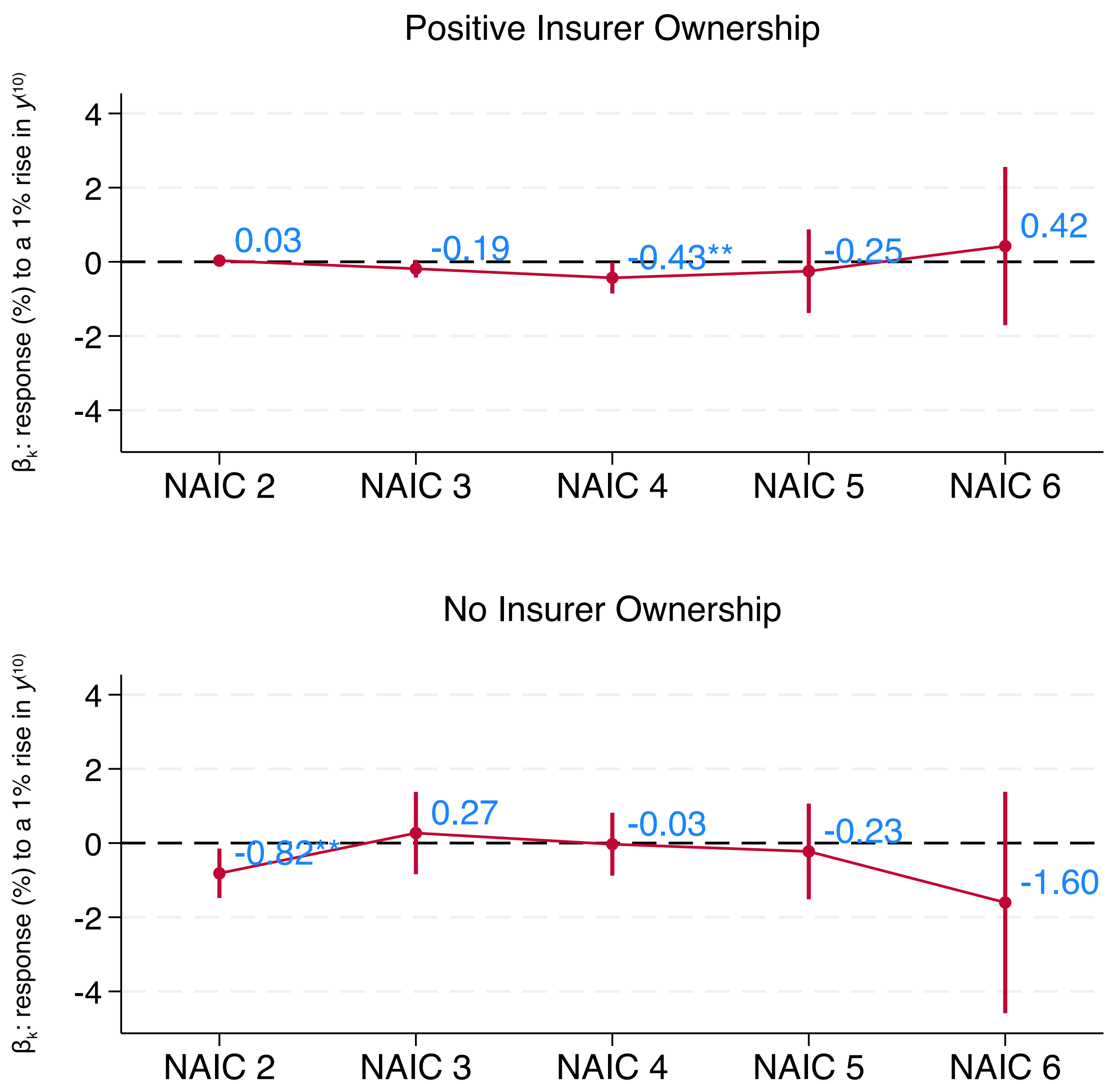

Figure A.6. **Credit Spread Responses by Life Insurance Ownership (pre-2007).**

This figure plots the coefficients β_k estimated from regression (13) for pre-GFC period, controlling for the trading volume, outstanding amount, time to maturity, duration, credit ratings, coupon rate, 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 the sub-sample of bonds with life insurance ownership. The bottom panel shows estimates for the sub-sample of bonds without life insurance ownership. The p-values are based on standard errors clustered at the issuer and year-month levels. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

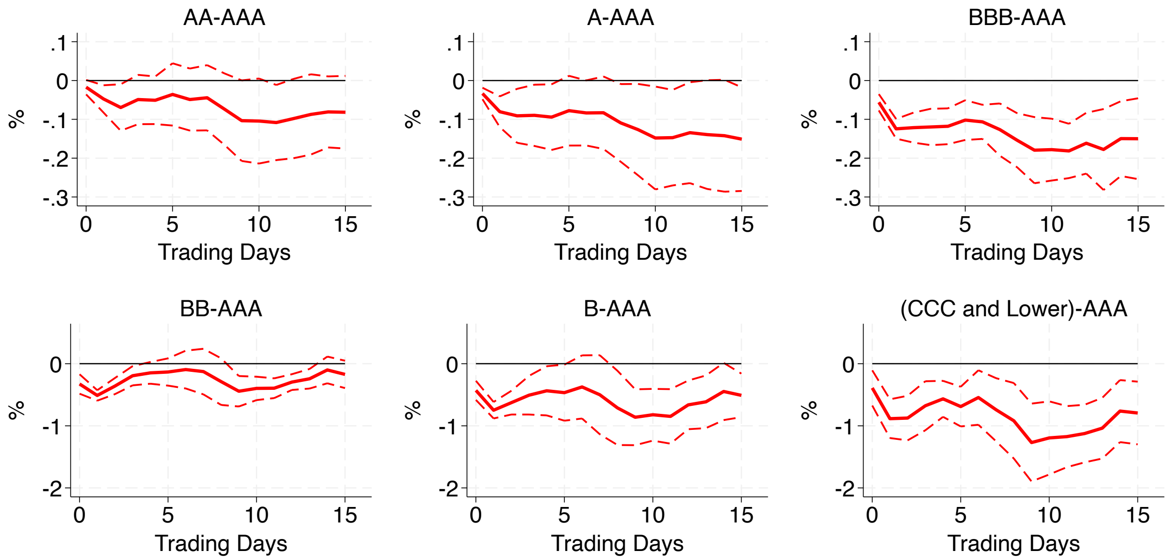

Figure A.7. **Impulse Responses of Credit Spreads (2010-2022).**

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

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

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

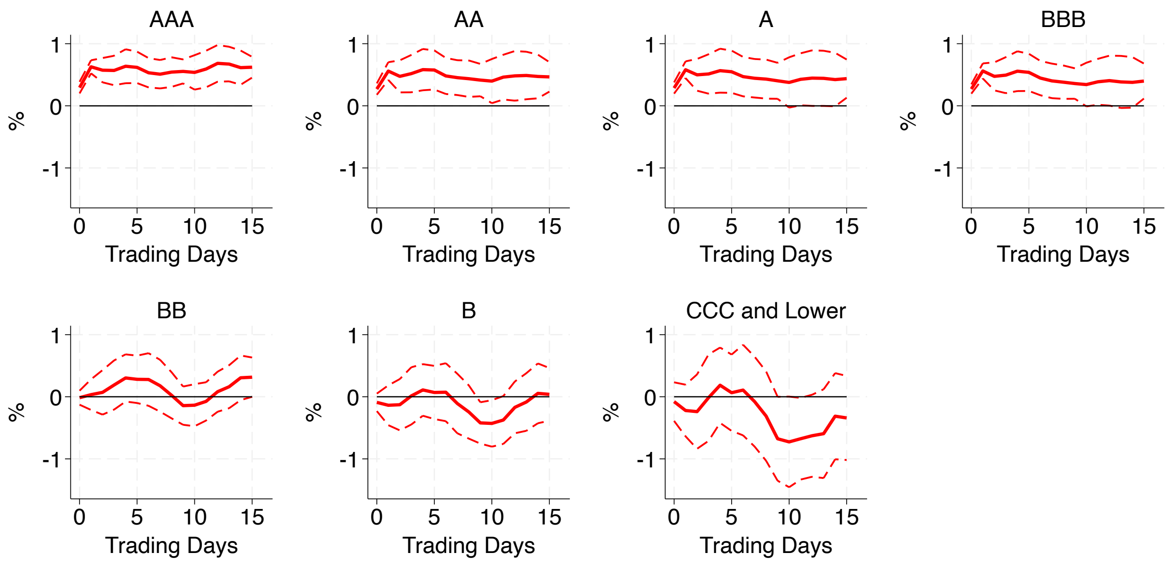

Figure A.8. **Impulse Responses of Yield Indices (2010-2022).**

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.

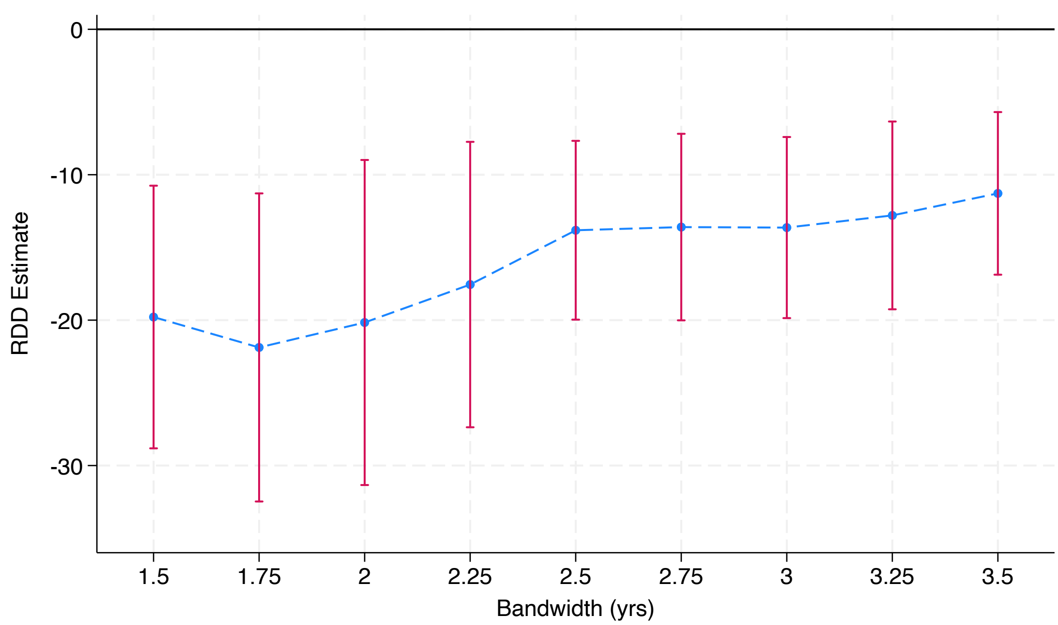

Figure A.9. **Alternative RDD Bandwidth Choice.**

This figure plots the estimated γ and the associated 95% confidence intervals from (18) under different choices of bandwidth.

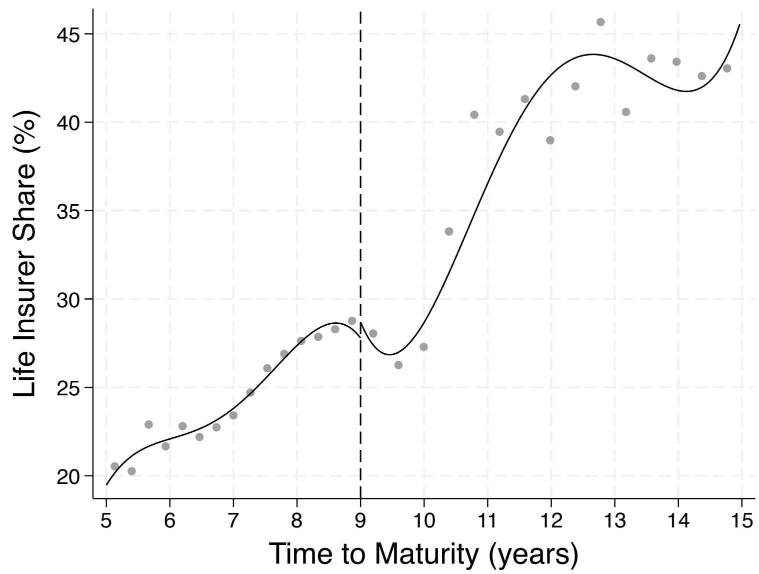

Figure A.10. **Discontinuity in Investor Composition.**

This figure shows a bin scatter plot of corporate bonds' life insurance ownership share (%) and their time to maturity. The vertical line indicates the maturity threshold of 9 years. The figure also shows the best-fit fourth-order polynomials for observations on each side of the threshold.

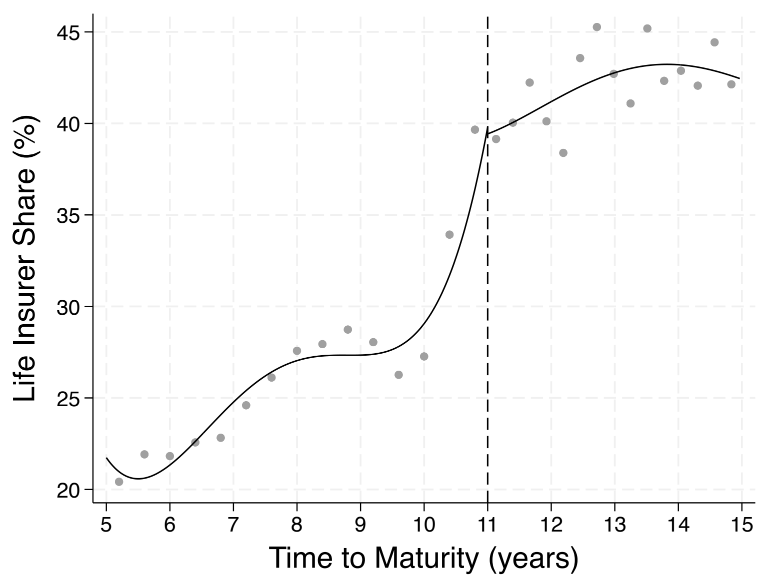

Figure A.11. **Discontinuity in Investor Composition.**

This figure shows a bin scatter plot of corporate bonds' life insurance ownership share (%) and their time to maturity. The vertical line indicates the maturity threshold of 11 years. The figure also shows the best-fit fourth-order polynomials for observations on each side of the threshold.

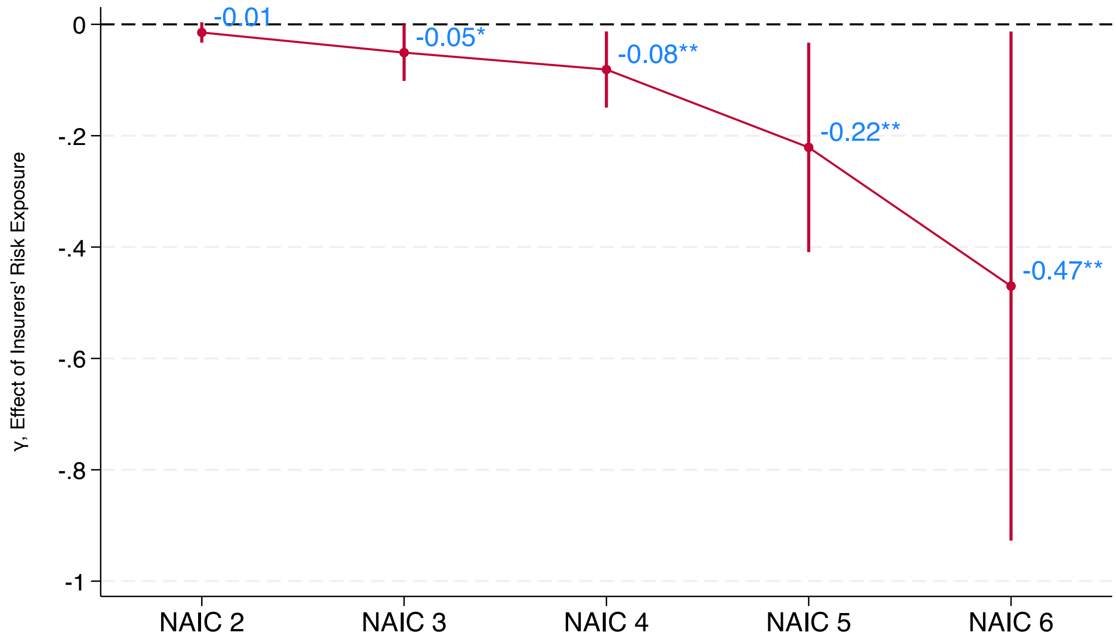

Figure A.12. **Insurers' Risk Exposure and Credit Spread Dynamics.**

This figure plots the coefficients γ_k estimated from regression (17), controlling for the trading volume, outstanding amount, time to maturity, duration, credit ratings, coupon rate, 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 regression is estimated for the period between 2002 and 2020, excluding the GFC (2007-2009). The p-values are based on standard errors clustered at the issuer and year-month levels. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

	$\mathbf{1}\{\varphi_{it}^{\text{Ins}} > 0\}$	$\mathbb{E}_t[\varphi_{it}^{\text{Ins}} \mid \varphi_{it}^{\text{Ins}} > 0]$	$\max \varphi_{it}^{\text{Ins}}$	Amount Outstanding ($ bn)
NAIC 1	88.7%	29.4%	100%	2268.3
NAIC 2	93.6%	31.7%	100%	1821.2
NAIC 3	89.3%	13.0%	98.0%	381.3
NAIC 4	79.5%	5.6%	90.9%	254.0
NAIC 5	58.8%	3.4%	75.7%	80.2
NAIC 6	31.5%	2.4%	64.9%	12.1
NAIC 1-2	91.0%	30.5%	100%	4089.5
NAIC 3-6	80.9%	9.3%	98.0%	727.6

Table A.1. **Summary Statistics.**

This table summarizes life insurers' bond ownership in WRDS Bond Returns between 2010 and 2019. The first column shows the average 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 observed in each category. The last column shows the average total outstanding amount of each category.

	Bond Age $> 1m$	Issuance Maturity > 10.25
γ	-13.78***	-12.02***
	[0.001]	[0.000]
Controls	✓	✓
Time FE	✓	✓
Kleibergen-Paap F-stat	79.859	75.024
Observations	10680	3427

Table A.2. **RDD Regressions: Robustness Checks.**

This table shows the coefficients estimated from regression (18), with the instrument, controlling for the trading volume, time to maturity, duration, maturity, size, coupon rate and frequency, credit rating, and insurer ownership share for each bond, and the recent default rate for each NAIC category. The first column shows the result excluding bonds issued less than a month ago. The second column shows the result for bonds whose maturity is above 10.25 years at issuance. The p-values shown in brackets are based on standard errors clustered at the issuer and year-month levels. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

	Pre-2007	Post-2009
γ	0.800	-15.12***
	[0.891]	[0.000]
Controls	✓	✓
Time FE	✓	✓
Kleibergen-Paap F-stat	23.823	73.663
Observations	2443	7855

Table A.3. **RDD Regressions: Controlling for Callability.**

This table shows the coefficients estimated from regression (18), with the instrument, controlling for the trading volume, time to maturity, duration, maturity, size, coupon rate and frequency, credit rating, insurer ownership share, ***callability dummies****, and the* ***interaction of the moneyness ratio and*** $\Delta y_t^{(10)}$ *for each bond, and the recent default rate for each NAIC category. The p-values shown in brackets are based on standard errors clustered at the issuer and year-month levels.* $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

B Constructing Bond Call Prices

In this section, I describe the method I use to construct the callability and strike prices of a large fraction of corporate bonds in my sample. As in Ma, Streitz and Tourre (2023), we can categorize bonds into four types based on the type of their call options:

1. Non-callable bonds: bonds without any call option.
2. Fixed-price callable bonds: bonds with fixed call prices (which can potentially vary over time).
3. Make-whole callable bonds: bonds for which the call price is equal to the present value of remaining cash flows, discounted at the US Treasury yield curve plus a make-whole spread.
4. Hybrid bonds: bonds that are initially make-whole but become fixed-price after a certain amount of time.

Due to two data limitations, I can only obtain call prices for a subsample of my data. First, the main source of data on bond callability used in the literature is the 'redemption' dataset from Mergent FISD, which does not cover the entire sample of WRDS Bond Returns or TRACE. Second, some fixed-price callable bonds have complicated call price schedules. Without the sub-dataset 'call schedule' of Mergent FISD, I do not observe how the call prices of fixed-price bonds vary over time.

Instead, I obtain the text descriptions of bonds' call options from the 'redemption' dataset of Mergent FISD, which indicates if fixed-price bonds are callable at "par" (call prices are equal to the par value) or "prices" (call prices change over time, but the exact prices are not reported). I then keep the following types of bonds in my new sample: (1) non-callable bonds, (2) fixed-price callable bonds whose call price is equal to par, (3) pure make-whole bonds, (4) hybrid bonds whose fixed-price part has a call price equal to par. In other words, I discard bonds with fixed call prices that vary over time due to the lack of data. After this procedure, the new sample includes 66.44% of the observations of my main sample. To my knowledge, Ma, Streitz and Tourre (2023) is the only recent study that constructs bond-level call prices for the US corporate bond market. In comparison, the final dataset of Ma, Streitz and Tourre (2023) includes 21,508 bonds out of 602,446 bonds issued between 2000 and 2022 (3.57% of the US corporate bond universe).

For make-whole bonds, I use the following formula to compute call prices,

$$(\text{Call Price})_t = \sum_{n=1}^{N} \frac{\text{CF}_{t+n}}{\left[1 + y_t^{(n)} + (\text{make-whole spread})\right]^n}.$$

Here, CF_{t+n} is the bond's cash flow (e.g., coupons and face value) at time $t+n$, $y_t^{(n)}$ is the Treasury yield of maturity n at time t from Liu and Wu (2021),[33] and the make-whole spread is obtained from Mergent FISD. Following Ma, Streitz and Tourre (2023), if the make-whole spread is missing in Mergent FISD, I assume that the spread is 25 bps for investment-grade bonds and 50 bps for high-yield bonds. As in Ma, Streitz and Tourre (2023), for hybrid bonds, the cash flows include future coupons between now and the first fixed-price call date plus the first fixed-price call price.

Following the literature on derivative pricing, I measure the bonds' moneyness using the following ratio,

$$(\text{Moneyness Ratio})_t = \frac{(\text{Bond Price})_t}{(\text{Call Price})_t},$$

which compares the current bond price to the call strike price. Intuitively, the moneyness of an option increases if the market price increases relative to the call price. I have also included non-callable bonds in this new sample, for which I assume the strike price is infinity and the moneyness is zero.

[33] I replace $y_t^{(n)}$ with the Treasury yield of the longest available maturity (typically 30 years) if n exceeds the maximum available Treasury yield maturity.

C Analytical Model: Derivation

C.1 A Sufficient Condition for Duration Mismatch

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

First, differentiating equation (9),

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

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

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

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

Condition (C.2) can be written as

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

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 (C.3) holds for $\gamma^n = 0$ if

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

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

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

In Section C.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 (C.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:

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

C.2 Proof of Proposition 1

The market clearing condition (6) can be written as

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

Since the bond supply is upward-sloping and the habitat demand is downward-sloping, 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

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

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

Differentiating (C.6) yields

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

As the insurer faces a negative duration gap, $\partial A_t^I/\partial y_t^T > 0$, 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 (C.5), we see that the credit spreads must also increase. Equation (C.7) also tells us the magnitude of the effect $\left|\frac{\partial w_t^{I,n}}{\partial y_t^T}\right|$ increases in the severity of duration mismatch $\left|\frac{\partial A_t^I}{\partial y_t^T}\right|$.

Online Appendix

D 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.

Additional Assumption for the Quantitative Model. We need a few additional assumptions to solve the model quantitatively.

First, I assume that the dynamics of the Treasury yield is given by

$$\mathrm{d}y_t^T = \alpha_y \left(y_t^T - \overline{y}^T\right)\mathrm{d}t. \tag{D.1}$$

Second, I assume that the equity injection process follows,

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

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

Third, I assume that the preferred-habitat investors have the following demand functions

$$\log D_t^{P,n} = \alpha^n - \beta \log P_t^n. \tag{D.2}$$

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.

Fourth, I micro-found and derive an explicit bond supply function. There are N sectors of firms. 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 (3), 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.

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

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

$$B_t^n = \left[\frac{(P_t^n)^{1-\theta}}{\phi^n} \right]^{\frac{1}{\theta}}. \tag{D.3}$$

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.

Calibrated Parameters. Table A.4 discusses model calibration.

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 high-yield 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 A.4. **Calibrated Parameters.**

I consider two corporate bond ratings ($N = 2$) where $n = 1$ represents investment-grade bonds (NAIC 1-2), and $n = 2$ represents high-yield 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.[34]

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 and 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 $a = 2$.

[34]To map the production function into the standard Cobb-Douglas form, we can assume that every firm has a single unit of labor supply $L^n_t = 1$ and the production function is $Y^n_t = \frac{1}{1-\theta}(K^n_t)^{1-\theta}(L^n_t)^\theta$.

Estimated Parameters. I estimate a few other key parameters using empirical data. The estimated parameters are summarized in Table A.5.

Parameter	Estimation Strategy / Source
Corporate Bonds	
$\nu^1 = 0.001,\ \nu^2 = 0.017$	Average default rates
$\delta = 1.635$	Variance of high-yield 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 $(P^2B^2/P^1B^1)^{ss} = 0.195$
$L = 8.05$	Empirical duration mismatch
Habitat Investor	
$\beta = 1.1$	Darmouni, Siani and Xiao (2025)
$\alpha^1 = 0.47,\ \alpha^2 = 0.12$	Life insurers' share in each category (35.7%, 10.3%)

Table A.5. **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 high-yield default rate. The average one-year default rate between 2003 and 2019 is 0.156% for investment-grade bonds and 2.829% for high-yield 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 high-yield bonds and the relative market cap of the two types of bonds (see Table A.1). 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 5.1). 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 6% 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 (2025). 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.[35] In the steady state, the insurer owns 35.7% of investment-grade bonds and 10.3% high-yield bonds, which match the empirical observation.

D.1 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.[36] The goal of this section is to quantitatively examine the policy's impacts on corporate bond yields, spreads, and issuance, as 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 5.1, 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

[35] In Section 5.2, 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 A.1).

[36] 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.

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 A.13 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.

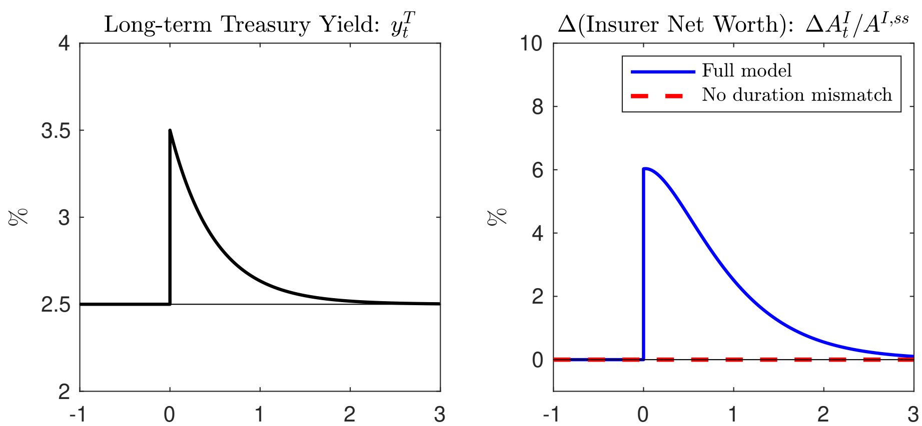

Figure A.13. **Treasury Yield Shock and the Insurer's Net Worth Responses.**

Figure A.14 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.

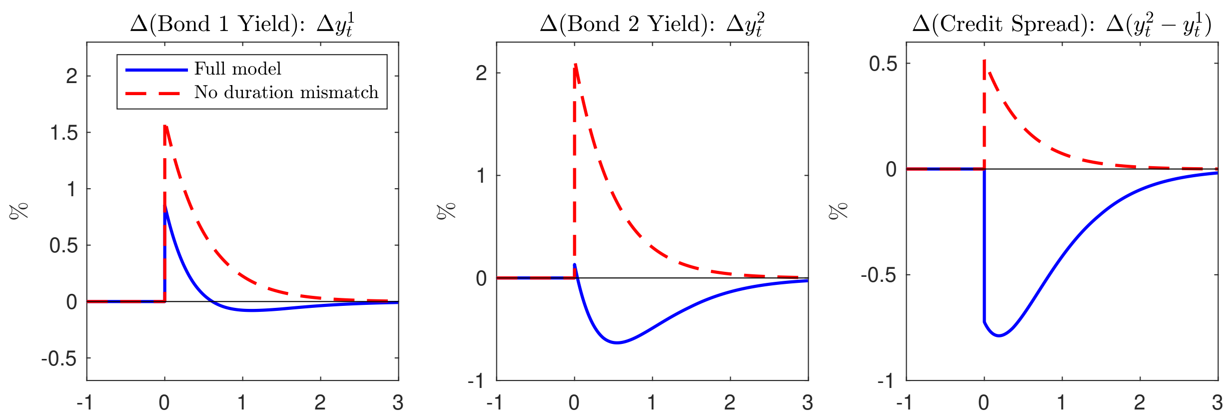

Figure A.14. **Responses of Bond Yields and Credit Spreads.**

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 high-yield 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 high-yield yield, and the credit spread between the two bonds falls by almost 0.8% 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 4.2 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.

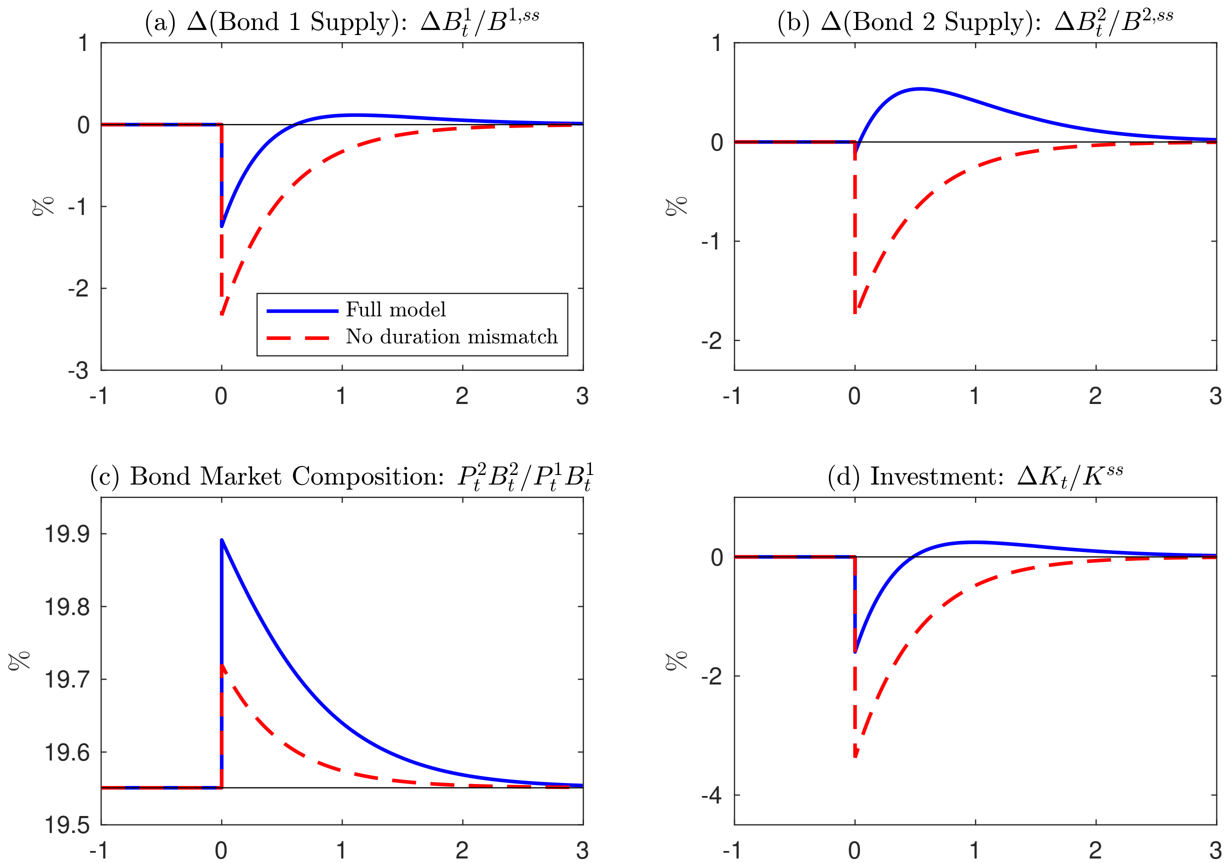

Figure A.15. **Responses of Bond Supply and Investment.**

Panels (a) and (b) of Figure A.15 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 high-yield bond supply.

Panel (c) of Figure A.15 shows the relative market cap of bond 2 relative to bond 1. The QT policy tilts the market towards the high-yield 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 A.15 displays the aggregate investment response. Absent the duration mismatch channel, firms disinvest more than 3% of their capital following the Treasury yield increase. The investment response is also heavily muted in the full model, at about 40% 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 high-yield 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.

D.2 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

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

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:

$$\begin{aligned}
\frac{\partial P_t^n}{\partial t} = &- \alpha_y \left(y_t^T - \bar{y}^T\right) \frac{\partial P_t^n}{\partial y_t^T} - \left(A_t^I \mu_t^{A,I}\right) \frac{\partial P_t^n}{\partial A_t^I} \\
&+ \left[y_t^T - \frac{1}{y_t^T + \lambda^T} \alpha_y \left(y_t^T - \bar{y}^T\right) + \lambda^n + (a \sigma_t^{A,I} - 1) \delta \sigma_t^{r,n} + (\zeta^n)^2 w_t^{I,n} \right] P_t^n - \phi^n,
\end{aligned} \tag{D.4}$$

where

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

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

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

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

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 (D.4) 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.

E The Duration Management Channel

In this section, I discuss an extension model with a continuum of bonds of different durations where the life insurer has a duration management motive.

Assumptions. As in Section 2.2, I make two simplifying assumptions to make the model tractable. The first assumption is the same as before (i.e., Assumption 1).

Assumption 3 *Treasury yields are expected to be constant.*

However, as we will next introduce bonds with different durations, it is no longer appropriate to assume that corporate bonds are short-term as in Assumption 3. Instead, I assume that the credit default risk is small and its impact on credit spreads is negligible (as confirmed in Section 4.2). As a result, credit spreads in this model will almost entirely be driven by regulatory costs.

Assumption 4 *Default risk is negligible compared to regulatory costs, i.e.,* $\zeta^n \gg \nu^n \approx 0$*.*

Assets. In the baseline model of Section 2, the insurer's asset portfolio consists of a single long-term Treasury bond and N corporate bonds. In the extended model, I instead assume that there exists a continuum of Treasuries. As in the baseline model, under the simplifying assumptions, the yield of a Treasury bond is given by

$$y^T = \frac{\phi^T}{P^T} - \lambda^T.$$

The duration of the Treasury bond is then

$$D^T = -\frac{1}{P^T}\frac{\partial P^T}{\partial y^T} = \frac{1}{y^T + \lambda^T} = \frac{P^T}{\phi^T}.$$

Instead of assuming that the decaying rate λ^T is constant, there exists a continuum of Treasuries with different decaying rates λ^T, which takes value within the range $[\underline{\lambda}^T, \overline{\lambda}^T]$. The Treasury yields are still exogenously given. I denote the resulting range of Treasury durations as $[\underline{D}^T, \overline{D}^T]$.

Similarly, for each credit rating n, I assume that there exists a continuum of corporate bonds with different decaying rates. Like Treasuries, we can write the duration of a corporate bond as

$$D^n = \frac{1}{y^n + \lambda^n} = \frac{P^n}{\phi^n}.$$

Ceteris paribus, it can be shown that bonds with a higher decaying rate have lower prices. Hence, there exists a one-to-one and downward-sloping mapping from λ^n to D^n in equilibrium. Suppose that the decaying rate of rating n corporate bonds λ^n takes values within the range $[\underline{\lambda}^n, \overline{\lambda}^n]$. We can then denote the range of durations of rating n corporate bonds as $[\underline{D}^n, \overline{D}^n]$.

As we are primarily interested in credit spreads between bonds with matching durations, I assume that the available durations of Treasuries coincide with those of the corporate bonds. In other words, the distributions of λ^T and λ^n are chosen such that

$$[\underline{D}^T, \overline{D}^T] = [\underline{D}^n, \overline{D}^n] = [\underline{D}, \overline{D}]. \tag{E.1}$$

I further assume that the range of durations is time-invariant. In other words, following a surprise change in interest rates, the distributions of λ^T and λ^n will adjust accordingly to ensure that the durations of Treasuries and corporate bonds still fall within $[\underline{D}, \overline{D}]$.

Under Assumptions 3 and 4, the return rate of a Treasury equals its yield,

$$\frac{\mathrm{d}r^{T,D}}{\mathrm{d}t} = y^{T,D}.$$

where $y^{T,D}$ is the yield of a Treasury with duration D. For a corporate bond of rating n, the return process is given by equation (3),

$$\mathrm{d}r^n = \frac{\phi^n - \lambda^n P^n}{P^n}\mathrm{d}t + \frac{\mathrm{d}P^n - \nu^n P^n \mathrm{d}J_t}{P_t}. \tag{E.2}$$

Since credit risk is small ($\nu^n \approx 0$) and the Treasury yields are expected to be constant ($\mathrm{d}P^n = 0$), we can similarly simplify the return of corporate bonds to

$$\frac{\mathrm{d}r^{n,D}}{\mathrm{d}t} = y^{n,D} = \frac{\phi^n}{P^{n,D}} - \lambda^{n,D}.$$

where $y^{T,D}, P^{n,D}, \lambda^{n,D}$ are the yield, price, and decaying rate of a rating n corporate bond with duration D.

In addition, there exists a zero-duration asset ("cash"), that is in unlimited supply and has a return rate of y^0, $\mathrm{d}r^0 = y^0\mathrm{d}t$.

The Insurer's Portfolio Problem. The insurer has a net worth of A^I, and it first chooses how to allocate its net worth across durations. I denote the insurer's portfolio weight in duration-D assets as $w^{I,D}$. It invests the remaining wealth, in the zero-duration assets, so the portfolio weight in cash is $w^{I,0} = 1 - \int_{\underline{D}}^{\overline{D}} w^{I,D}\mathrm{d}D + w^{I,L}$.

Second, within each duration D, the insurer chooses how to allocate its holdings between Treasuries and corporate bonds. I use $\theta^{n,D}$ to denote the insurer's holdings of rating n bonds of duration D as a fraction of its total assets of duration D. Similarly, its holdings of duration-D Treasuries as a fraction of its total duration-D assets is then $\theta^{T,D} = 1 - \sum_n \theta^{n,D}$.

Under Assumption 4, the portfolio problem can then be written as[37]

$$\max_{w^{I,D},\,\theta^{T,D},\,\theta^{n,D},\,w^{I,0}} \overbrace{\int_{\underline{D}}^{\overline{D}} w^{I,D}\left[\theta^{T,D}y^{T,D} + \sum_{n=1}^{N}\theta^{n,D}y^{n,D}\right]\mathrm{d}D + w^{I,0}y^{0}}^{\text{expected asset return}} \tag{E.3}$$
$$- \underbrace{\frac{1}{2}\sum_{n=1}^{N}\int_{\underline{D}}^{\overline{D}}\left(\zeta^{n}w^{I,D}\theta^{n,D}\right)^{2}\mathrm{d}D}_{\text{regulatory cost}} - \underbrace{\frac{\psi}{2}(\text{Duration Gap})^{2}}_{\text{duration management}},$$
$$\text{s.t.}\quad w^{I,0} = 1 - \int_{\underline{D}}^{\overline{D}} w^{I,D}\mathrm{d}D + w^{I,L}, \quad \theta^{T,D} = 1 - \sum_{n=1}^{N}\theta^{n,D}.$$

The new term $\frac{\psi}{2}(\text{Duration Gap})^2$ captures the insurer's duration management motive, and the duration gap is defined as:

$$\text{Duration Gap} := \frac{1}{A^I}\left(A^I\int_{\underline{D}}^{\overline{D}} w^{I,D}D\mathrm{d}D - LD_L\right).$$

Here, D_L is the duration of annuities, LD_L is the total duration of the insurer's liabilities, and $A^I\int_D w^{I,D}D\mathrm{d}D$ is the total duration of the insurer's assets. The insurer incurs disutility from non-zero duration gaps, and the constant $\psi > 0$ measures the strength of the duration management motive. As I am mainly interested in the situation where the insurer has a negative duration gap, I assume that Duration Gap < 0 in equilibrium.

Habitat Demand and Bond Supply. In the spirit of Vayanos and Vila (2021), I assume that both the habitat demand and the supply of corporate bonds are segmented by duration. In particular, corporate bonds of rating n and duration D have a reduced-form habitat demand function $HD^{n,D}(P^{n,D})$ that is decreasing in the price $P^{n,D}$. Similarly, the supply of corporate bonds of rating n and duration D is given by an increasing function $B^{n,D}(P^{n,D})$. The market for corporate bonds of rating n and duration D clears when the insurer's demand and the habitat demand sum to the total supply,

$$w^{I,D}\theta^{n,D}A^I + P^{n,D}HD^{n,D} = P^{n,D}B^{n,D}. \tag{E.4}$$

[37]Here, I omit the terms related to annuity returns since the portfolio weight in annuities is exogenously given in equilibrium.

Credit Spread Dynamics. Solving the problem (E.3), the first-order conditions with respect to the portfolio weights within durations $\theta^{n,D}$ are

$$y^{n,D} - y^{T,D} = (\zeta^n)^2 w^{I,D} \theta^{n,D}.$$

We can therefore write the insurer's demand for bonds of rating n and duration D as

$$w^{I,D}\theta^{n,D}A^I = \frac{y^{n,D} - y^{T,D}}{(\zeta^n)^2/A^I}, \tag{E.5}$$

Note that the insurer's demand (E.5) and the market clearing condition (E.4) are exactly analogous to those in the baseline model, i.e., (11) and (6). It then follows from the proof of Proposition 1 that, if the insurer has a negative duration gap, an increase in interest rates (e.g., a uniform increase in Treasury yields across all durations) reduces the insurer's effective risk aversion $(\zeta^n)^2/A^I$ and lowers the credit spreads $y^{n,D} - y^{T,D}$ for all rating n and duration D.

The Duration Management Channel. Next, we analyze how the insurer allocates its portfolio across durations. The first-order conditions with respect to $w^{I,D}$ are

$$y^D - y^0 = \sum_{n=1}^{N} \left(\zeta^n \theta^{n,D}\right)^2 w^{I,D} + \psi(\text{Duration Gap})D.$$

Here y^D is the average yield of duration-D bonds in the insurer's portfolio,

$$y^D = y^{T,D} + \sum_{n=1}^{N} \theta^{n,D}(y^{n,D} - y^{T,D}).$$

When Duration Gap < 0, we can further rewrite the first order condition as

$$w^{I,D} = \frac{(y^D - y^0) + \psi D|\text{Duration Gap}|}{\sum_{n=1}^{N} \left(\zeta^n \theta^{n,D}\right)^2}.$$

The first term in the numerator, $(y^D - y^0)$, captures the effect of term spreads on the insurer's portfolio decisions. Ceteris paribus, the insurer holds more long-term assets when the term spread is higher.

The duration management channel influences the insurer's portfolio allocations via the second term, $\psi D|\text{Duration Gap}|$. When interest rates increase, the insurer's duration gap shrinks (|Duration Gap| $\downarrow$),[38] which can reduce the insurer's demand for long-duration assets ($w^{I,D}$ $\downarrow$). Moreover, the effect of the duration management channel is stronger in bonds of higher durations.

In richer models with a term structure of interest rates (e.g., Vayanos and Vila, 2021), such demand shifts across durations from the insurance sector can lead to significant impacts on the yield curve (Domanski, Shin and Sushko, 2017; Greenwood and Vissing-Jorgensen, 2018). However, since the focus of the current paper is on the credit spreads between bonds with matching durations, I leave a more thorough analysis of the impacts of the duration management channel on the term structure of corporate bonds for future research.

[38] In this model, the duration gap shrinks because the duration of the insurer's liabilities D^L becomes shorter when Treasury yields increase.