From Long to Short:
How Interest Rates Shape Life Insurance Markets*

Ziang Li† **Derek Wenning**‡

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Abstract

This paper investigates how interest rate fluctuations shape life insurance markets, focusing on the liability adjustments insurers employ to manage interest rate risk. After the 2008 Financial Crisis, insurers exposed to high interest rate risk – especially those offered variable annuities with minimum return guarantees pre-2008 – shifted their product portfolios toward short-duration policies to hedge against rising duration gaps. Using a combination of theoretical and empirical analysis, we show that this liability rebalancing led to sizable contractions in both the supply of long-duration life insurance products and the aggregate life insurance market. Our findings reveal that interest rate risk can significantly influence financial intermediaries' liability choices, which in turn shape the composition and availability of financial products.

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†Imperial College London. Email: ziang.li@imperial.ac.uk

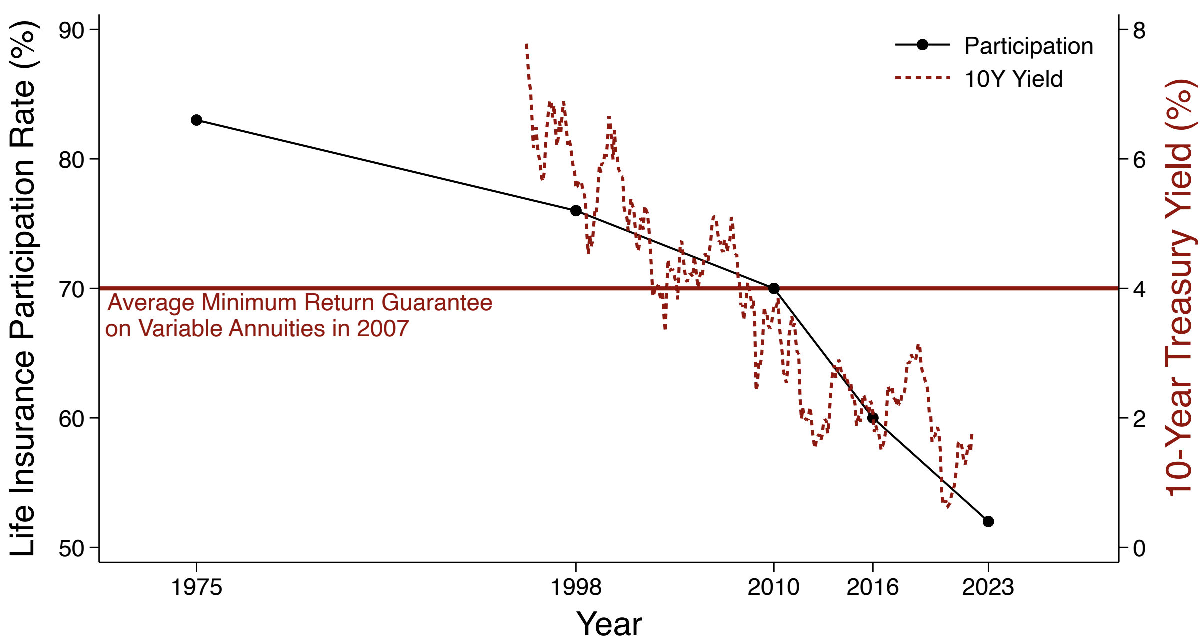
‡Kelley School of Business, Indiana University. Email: dtwennin@iu.edu

1 Introduction

Life insurance participation has steadily declined for the past half century, and at an accelerating pace (see Figure 1). According to a report by the Guardian Life Insurance Company (Guardian, 2023), life insurance participation declined from 83% in 1975 to 70% in 2010, or 0.37 percentage points per year. Participation continued to decline to 60% just six years later — a magnified rate of 1.67 percentage points per year — and today sits at 52%. The sharp drop in participation has important consequences: among households that experience the loss of an income-earner, 84% that did not have life insurance report living paycheck-to-paycheck as opposed to the 36% that did (Guardian, 2023).

It is therefore reasonable to suspect that the sharp decline in participation was driven by forces beyond household demand. In particular, the post-crisis recovery was accompanied by historically low interest rates, as shown in Figure 1. Life insurers — financial institutions with particularly long-lived liabilities — are generally sensitive to the revaluation effects of interest rate changes. Modern life insurance and annuity products are especially exposed due

Figure 1: Life Insurance Participation and Treasury Yields



Note: This figure plots life insurance participation rates (left axis) and 10-year Treasury yields (right axis) over time. Data on life insurance participation come from Guardian (2023), which is itself derived from LIMRA Barometer reports. The size of the gray triangles represents the size of participation declines between measurement years. Monthly 10-year Treasury yields are taken from FRED and cover 1995 to 2022. The average minimum return guarantee is taken from Koijen and Yogo (2022).

to their minimum return guarantees, embedded options whose valuation grows dramatically when interest rates are low. In particular, Koijen and Yogo (2022) highlight that the average minimum return guarantee of variable annuities issued in 2007 sat at 4%, approximately 2 percentage points higher than ensuing Treasury yields just a few years later. As a result, the reserve value of the embedded options grew substantially, leaving life insurers exposed to elevated interest rate risk.

This paper explores a new channel through which life insurers may hedge interest rate risk: *liability rebalancing*. As we discuss in Section 2, life insurance product markets are segmented by maturity, and therefore, degrees of interest rate risk. Ordinary life insurance products (term or whole life) provide long-term coverage, while life insurance accessed through employers (group life) typically provides coverage for a single year. Given limits to duration matching through asset rebalancing (Ozdagli and Wang, 2019; Sen, 2023), insurers may naturally transition from ordinary life to group life issuance to reduce their interest rate risk in a low interest rate environment. However, since group life insurance is only accessible through (large) employers, there could be negative consequences for participation at the market level. Moreover, since group life policies typically provide lower levels of coverage than ordinary life policies (Guardian, 2023), life insurance coverage as a whole may shrink.

We explore these insights formally in Section 3. We provide a model of insurance product markets in which risk-averse insurance companies are exposed to interest rate risk. Insurers care about their operating profits as well as the volatility of their capital returns. Further, they add duration to their capital through new liability issuance, which could amplify the risk of their capital returns when interest rates are uncertain. As a result, insurers trade off current-period profits with future interest rate risk when issuing new policies.

We first show formally that insurers hedge interest rate risk through product markets, consistent with our concept of liability rebalancing. In particular, when interest rate uncertainty rises, insurers issue fewer long-duration policies but increase their issuance of short-duration policies. This effect is especially pronounced for insurers with more negative duration gaps and larger capital convexity: because their capital returns respond more to declines in interest rates, they rebalance toward short-duration policies more greater intensity.

We then cast the model in general equilibrium to study how liability rebalancing affects product markets. In contrast to the partial equilibrium setting, we show that less exposed (but not unexposed) insurers may increase their long-duration product issuance due to the decline in competition. In this sense, less exposed insurers try to fill the gap left by more exposed insurers. However, due to decreasing returns to scale, the substitution across insurers is not enough to stabilize the market and total issuance of long-duration policies declines.

With these predictions in hand, we next turn to our empirical analysis in Section 4. Our data are taken from life insurers' annual statutory filings. For each insurer, we have access to both new issuance and insurance in force for their term life, whole life, and group life businesses from 2005 to 2023. We also collect data on monthly term life prices from Compulife, a quotation software used by life insurance agents. We use information on the account value of insurers' variable annuities in the pre-crisis period to classify them into exposed and non-exposed insurance groups. The exposed insurers are relatively large in terms of assets and capital, but they are 2-3 times more levered. Beyond their exposure to variable annuity guarantees, they also hold a higher share of interest-sensitive life insurance reserves.

We begin by revisiting duration gap estimates for our two insurer classifications. We first replicate the finding in the literature that at the industry level, duration gaps became negative after the financial crisis (Berends et al., 2013; Koijen and Yogo, 2022; Kirti and Singh, 2024). We take one step further and utilize liability duration estimates from Huber (2022) to examine the differences in duration gaps between exposed and non-exposed insurers following the financial crisis. Consistent with our narrative, we find that the duration gaps of exposed insurers became relatively more negative during the low-interest-rate period. Our results are robust to controlling for year and insurer fixed effects.

According to our theory, product prices should reflect widening duration gaps. We use our data on term life insurance prices to test this hypothesis. We first show that relative maturity markups — the difference between long- vs. short-maturity policy prices sold by exposed relative to non-exposed insurers — correlates negatively with 10-year Treasury yields at the monthly level.[1] This finding is consistent with our theory: during low interest rate periods, exposed insurers face relatively wider (negative) duration gaps, which increases the spread between the prices of their long- and short-duration products. We then show that the result holds more generally in a regression format, where we can control for a variety of granular fixed effects. Similar to low rates, interest rate uncertainty also amplifies insurers' exposure to interest rate risk. Accordingly, we find that relative maturity markups also rise when interest rate uncertainty is heightened. The results indicate that insurers pass through interest rate risk to their products on the maturity dimension, which we interpret as indirect evidence of liability rebalancing.

We then turn to a more direct test of liability rebalancing using our data on the issuance

[1]We show this for 20 year relative to 10 year term life products.

of insurance coverage and policies. Since we cannot observe issuance at the individual product level, we instead explore how insurers rebalance their issuance between generally long-duration products (term and whole life insurance) and short-duration group life insurance. We demonstrate this in both the raw data and through a regression specification with granular fixed effects, showing that exposed insurers rebalance their liabilities toward group life insurance relative to non-exposed insurers throughout the low-interest-rate period. We further connect this directly to the level of interest rate risk exposure that we estimated before. We interpret these results as more direct evidence of liability rebalancing.

We then aggregate policy issuance across insurers to reinterpret product market trends as consequences of insurers' risk management strategies. We document that yearly aggregate issuance of ordinary life insurance as a percentage of GDP declined by 48% between 2005 and 2023, with two thirds of the decline stemming from exposed insurers.[2] Group life insurance issuance also declined as a percentage of GDP, but to a much lesser extent. The results suggest that both supply-side and demand-side effects contributed to the reduction in life insurance issuance, but that the supply-side effects, primarily driven by exposed insurers, exacerbated the demand-side effects in ordinary life markets.

Despite the decline in issuance rates, the size of the insurance market could remain stable if new issuance exceeded claims and lapsation. We show that this was not the case: from peak to trough, ordinary life insurance in force as a percentage of GDP declined from 150.4% to 107%, three quarters of which stemmed from exposed insurance groups. Group life insurance in force only declined from 64% to 53% of GDP, the bulk of which came after the COVID-19 crisis. Due to the relative market sizes, total insurance in force declined from 213% to 160% of GDP, a contraction of nearly a quarter.

Our results suggest that the interest rate risk exposure of variable annuity issuers had severe consequences for life insurance markets. While we cannot yet disentangle the demand-side effects from the supply-side consequences, our work highlights a growing need for regulation to address insurers' risk exposures and mitigate the resulting cross-product spillovers.

Related Literature

Our work relates most closely to the literature on variable annuities and insurers' hedging behavior. In terms of market risk, Barbu (2023) shows that insurers reduce their exposure to variable annuities by having customers exchange them into less generous products. Further,

[2] This does not imply that nominal issuance of non-exposed insurers declined; in fact, total non-exposed insurer issuance increased by 34% by the end of the sample period. The discrepancy is due to relative changes in real GDP. See Section 4.5 for more details.

Barbu and Sen (2024) document that insurers have begun selling long-dated short put products in an attempt to reduce exposure to downside market risk. Ellul et al. (2022) shows both theoretically and empirically that insurers only partially hedge their variable annuity guarantee exposures by rebalancing their bond portfolios, but in doing so, exacerbate systemic risk. In terms of interest rate risk, Ozdagli and Wang (2019) show that transaction costs make it difficult for insurers to fully hedge their duration gap. Sen (2023) conducts a detailed study of variable annuity hedging and finds that differences in accounting methods used for assets and liabilities can further lead to imperfect hedging. We offer a new channel — liability rebalancing — through which insurers can reduce their exposure to variable annuities and interest rate risk by issuing shorter-duration products such as group life insurance and reducing their issuance of long-duration products.

We also contribute to the literature on how the financial health of life insurance companies spills over into their product markets.[3] Koijen and Yogo (2015) document that the wedge between actuarial and statutory reserve valuation methods affects pricing behavior. Barbu et al. (2024) show that the introduction of risk-based capital accounting affects both the prices of, the supply of, and the demand for life insurance. Ge (2022) further shows that for insurance groups with both P&C and life divisions, P&C losses worsen their financial health, spilling over to their life insurance division and leading to nuanced pricing behavior. Knox and Sørensen (2024) show that insurance prices reflect the gains and losses stemming from insurers' asset performance. Verani and Yu (2024) show that interest rate risk management plays a key role in annuity pricing, and the cost of interest rate risk hedging has pushed up annuity premia post-2008. Ellis et al. (2025) show that the roll-out of enterprise risk management mandates prompted insurers to shift from variable annuities with risky guarantees to index-linked annuities. Our contribution highlights the long-term effects on product markets when interest rate risk cannot be perfectly hedged.

Our paper builds on earlier works on the interest rate risk of US life insurance companies. Since Berends et al. (2013), a large literature has established that the duration gap of US life insurers switched from positive to negative after the 2007-2008 Financial Crisis (e.g., Hartley et al., 2016; Ozdagli and Wang, 2019; Koijen and Yogo, 2021, 2022; Huber, 2022; Sen, 2023; Kirti and Singh, 2024; Li, 2024). The majority of existing works identify the insurers' duration gap by using two-factor regression models to estimate the sensitivity of US life insurers' excess stock returns with respect to 10-year Treasury yields. In addition,

[3] In the context of P&C insurance, Gron (1994), Froot (2001), and Zanjani (2002) show that insurers' capital constraints can affect their insurance supply and demand. Damast et al. (2025) study how monetary policy affects homeowner insurance products through P&C insurers' interest rate risk exposure.

Huber (2022) and Sen (2023) provide more direct evidence from insurers' balance sheets showing that insurers hedge their long-term liabilities imperfectly post-2008. We relegate a more detailed discussion on insurers' duration gaps to Section 2.3. Domanski et al. (2017), Kirti and Singh (2024), and Li (2024) further show that interest rate risk has important asset pricing implications through insurers' asset demand.

Last, our paper connects to the literature on the industrial organization of insurance markets, particularly life insurance. Koijen and Yogo (2015, 2022) estimate demand for life insurance and variable annuities, respectively, and use their framework to understand how regulation affects product markets. Tang (2022) uses a structural model to evaluate the effects of regulatory competition across US state insurance regulators and the establishment of captive reinsurance. Wenning (2024) estimates a model of life insurance agent distribution across a rich geography to explore the consequences of national price-setting behavior. While we have not yet done so, our model is amenable to estimation and will be used to carry out counterfactual analyses in future work.

2 Institutional Setting

We begin with a broad description of insurance product markets. We then discuss the interaction between insurance reserves and interest rates. We end with a discussion of the regulatory and economic motives for life insurance companies to hedge interest rate risk and highlight why product markets are a feasible outlet for hedging.

2.1 An Overview of Life Insurance Products

Life insurance markets have evolved considerably since their inception. The earliest forms of life insurance were short-duration policies with minor payouts. Prominent insurers by today's standards often began with such policies (Knight, 1920): for example, at its inception in 1875, Prudential Financial, which today has over $1.4 trillion in assets under management, primarily sold industrial life insurance — small policies with maturities of about a week that targeted laborers in poor urban neighborhoods (Carr, 1975).

Since then, life insurance products have evolved considerably. The closest category to the traditional industrial life policy is what is known as group life insurance.[4] Insurers write group contracts with firms rather than individuals, and the firm itself issues insurance certificates to their workers. These certificates function primarily as yearly-renewable policies

[4]Industrial life insurance still exists today but the market is minuscule: as of 2023, it only accounts for 0.016% of gross life insurance coverage in force.

with premium rates that are renegotiated at renewal. Employer-sponsored group policies are especially small, typically covering only one to two years of an employee's salary (Guardian, 2023), and are less accessible, since not all employers offer group life insurance as a benefit. Group life coverage totaled about 55% of GDP in 2005.

Ordinary life insurance departs from group life insurance in both the coverage and the time dimension. On the coverage dimension, policyholders are free to choose their desired level of coverage rather than being fixed at one year's wage.[5] On the time dimension, products can be split into two broad categories: term life policies and whole or permanent policies. Term life policies pay out a pre-specified benefit upon the death of the insured, conditional on the death happening during a set number of years.[6] For example, a 10-year term life policy pays out if the insured dies between the time of issuance and 10 years. Whole life policies, on the other hand, do not expire unless premiums are not paid.[7]

Whole life policies are notable due to their embedded savings components. These policies typically have lower coverage but redirect a fraction of the premium revenue toward a savings vehicle that accrues interest. This is referred to as the cash value of the policy. Traditionally, the cash value is invested in fixed income assets whose investment returns are fairly stable. New innovations in whole life policies have emerged over time, such as variable, indexed, and universal life products, that invest the cash value in a variety of non-fixed-income assets and may come with additional embedded options, such as minimum return guarantees.

Life insurers also issue annuities, products that insure longevity as opposed to mortality. Standard annuity products are paid for upfront and provide a fixed stream of payments until the death of the insured. The payments can either start alongside the initial payment (immediate annuities) or after a set number of years (deferred annuities). Similar to whole life insurance, insurers have innovated on annuity products by allowing the payments to fluctuate with an underlying mutual fund. These are known as variable annuities. A key similarity with variable life insurance is that the returns often come with a minimum return guarantee. For example, if the return guarantee is 4% per year and the mutual fund only returns 2% in a given year, the insurance company must pay the remaining 2% out of pocket.

[5] In our data that we discuss in Section 4, the average ordinary life insurance policy covers $144,281 in 2023, while the average group life policy covers $67,185.

[6] Some policies allow for yearly renewals with adjusted premiums, but policyholders are not permitted to renew for the full term. Other provisions may allow the policies to be converted to permanent contracts.

[7] These policies technically expire at a very old age, such as 100 or 121. Since most individuals do not live this long or lapse well before this, the restriction is typically not binding.

2.2 Insurance Reserves and Interest Rate Risk

Insurance companies must hold reserves to ensure available payment for policyholders. The value of the reserves for traditional policies directly accounts for mortality risk conditional on the age, gender, and health status of the policyholder. Since many policies have a time component, the value of a given policy's reserves may change over time due to a higher loading on a higher mortality risk or due to changes in the discounted value of future payouts (Koijen and Yogo, 2015; Huber, 2022). As such, these policies, especially whole life and long maturity term life, carry implicit interest rate risk. Group policies, which are often yearly renewable, typically have a low reserve requirement and are not sensitive to interest rate risk due to their short maturities.

Insurers hold non-traditional policy reserves in their separate accounts rather than their general accounts (Koijen and Yogo, 2022). This is due to the fluctuating nature of the savings components. However, when these policies are bundled with minimum return guarantees, the value of the separate account does not cover the residual returns between the underlying mutual fund and the minimum return guarantee when the guarantee is in the money. Insurers therefore hold reserves in their general accounts to account for these options.

Variable annuity and life insurance reserves are therefore convex. When interest rates and stock market returns are high, the likelihood that the minimum return guarantee will be exercised is low. Reserve positions are therefore small since insurers are less likely to have to cover the gap in returns. However, when rates and stock returns are low and declining, the guarantees are more likely to be exercised, and the reserve valuations increase substantially. For example, as discussed in Huber (2022), Metlife's "5 Year Ratchet & ROP-d, GMIB w/ 10y, 7 to 8" variable annuity had a reserve value that increased 4-fold between 2009 and 2011. In general, Sen (2023) estimates that the duration of minimum return guarantees is between 9 and 17 years. The high duration and convexity of minimum return guarantees have also raised concerns among insurance practitioners. For example, a report by AM Best (Panko, 2012) also argues that large blocks of legacy annuities with minimum return guarantees created severe pressures on insurers' balance sheets post-2008.

2.3 Duration Matching Motives in the Life Insurance Industry

Insurers that specialize in ordinary life insurance hold reserves with long maturities, often spanning more than 30 years. Minimum return guarantees on their variable liabilities add both duration and convexity to their total reserve positions. Given the sensitivity of their reserves to interest rates, a natural interest rate risk management strategy is to hold assets that match the duration of their reserves.

However, duration matching is not always a successful or even feasible strategy. Market incompleteness may prevent insurers from perfectly matching the duration between their assets and liabilities. Corporate bonds, which account for the majority of insurers' asset portfolios (Koijen and Yogo, 2023), have an average duration of only around 7-8 years. While Treasury bonds can have a longer duration, their maturities are also capped at 30 years, and insurers in general dislike Treasuries for their relatively low returns.

Beyond market incompleteness, insurers also face a variety of other frictions that push against duration-matching motives. First, insurance regulations might inadvertently distort insurers' hedging motives. Sen (2023) argues that the mismatch in the accounting methods used for assets and liabilities discourages insurers from using interest rate derivatives to hedge variable annuities. Second, Ozdagli and Wang (2019) finds that illiquidity and transaction costs in the corporate bond market are potentially important factors preventing insurers from closing their duration gaps, as doing so requires insurers to turn over large fractions of their bond holdings, which could be prohibitively expensive.[8] This is consistent with the evidence in Huber (2022), which shows that the asset duration of individual life insurers did increase somewhat after the financial crisis, but not substantially.

Consequently, life insurers' duration gaps became negative after the financial crisis. Several existing studies (e.g., Berends et al., 2013; Hartley et al., 2016; Ozdagli and Wang, 2019; Koijen and Yogo, 2021, 2022; Kirti and Singh, 2024; Li, 2024) arrived at this conclusion by examining how insurers' stock returns co-moves with interest rates. After carefully studying insurers' balance sheets, Sen (2023) finds direct evidence that many insurers failed to hedge a significant proportion of their variable annuity liabilities. By calculating the insurance companies' asset and liability durations directly, Huber (2022) finds that the aggregate gap switched from positive to negative after 2010.[9] Additionally, Li (2024) shows that after the financial crisis, the market leverage of life insurance companies co-moved negatively with long-term Treasury yields. The negative impacts of low interest rates on the life insurance sector have also been voiced frequently in practitioner publications (e.g., Panko, 2012; Dobbyn, 2015; Scism, 2023)

[8] Furthermore, Domanski et al. (2017) and Greenwood and Vissing-Jorgensen (2018) suggest that, due to their large scale, the reach-for-duration by insurers could lead to a substantial increase in the total demand for long-term assets, which could further push down long-term interest rates, resulting in a vicious cycle.

[9] Note that despite the duration estimation of minimum return guarantees by Koijen and Yogo (2022) and Sen (2023), Huber (2022) sets the duration of the minimum return guarantees for variable annuities and life insurance policies to zero. Incorporating these liabilities would likely lead to an even stronger decline in duration gaps.

Given the limits to duration matching through asset rebalancing, we explore an alternative channel: *liability rebalancing*. Insurers can reduce the duration of their liabilities by allowing their legacy reserves to expire and shifting new issuance toward shorter-duration policies. In the following section, we present a model of insurance product markets in the presence of interest rate risk to explore how liability rebalancing can be used as a risk management strategy.

3 A Model of Product Markets and Interest Rate Risk

We first present a simple model of duration matching to organize the empirical exploration. We discuss the structure of the model in Section 3.1. We then explore how duration mismatch affects product pricing and liability rebalancing in Section 3.2. We end with a discussion on the cross-market equilibrium outcomes in Section 3.3.

3.1 Setup

Time is discrete, $t \in \mathbb{N}$. There are a large number of insurance companies, $j \in \mathcal{J}$, that sell a variety of insurance and annuity products, $i \in \mathcal{I}$, to a unit measure of households. Insurers have two functions. First, they sell insurance to households, strategically setting prices and the extent of their market penetration for each product. Second, they manage a portfolio of assets with exogenous insurer-specific returns. These two activities shape the behavior of insurers' capital.

We denote insurer j's portfolio's return between periods $t-1$ and t as R^A_{jt}, which the insurer takes as given.[10] Insurers can expand their balance sheets and increase their assets by selling new insurance policies. When selling new products, insurers can attract more demand by setting lower prices, P_{ijt}, or by hiring more agents to market their products, T_{ijt}. We assume demand for each product-insurer pair takes the form

$$Q_{ijt} \equiv \overline{Q}_{ijt}\kappa(T_{ijt})P_{ijt}^{-\varepsilon_{it}}, \tag{1}$$

where $\overline{Q}_{ijt}$ is a insurer-product-specific component that we elaborate on in Section 3.3, $\kappa(T_{ijt})$ is an increasing function of T_{ijt} that varies between 0 and 1, and ε_{it} is the demand elasticity for policy i at time t. We assume for simplicity that the total number of agents attracted to

[10] In practice, insurers hold 60-70% of their asset portfolios in corporate bonds and, therefore, have asset returns close to the average return of the bond market (Koijen and Yogo, 2023). This assumption can in principle be relaxed to allow for reaching-for-duration by insurers (Ozdagli and Wang, 2019).

sell the insurer's products is linear in the commissions paid, $T_{ijt} = \eta_{it}^{-1} F_{ijt}$, for some constant $\eta_{it} > 0$.

Let A_{jt} denote insurer j's assets at the *beginning* of period t — i.e., the level of assets inherited from period $t-1$ and before issuing new products at time t. New policies issued during period t contribute to the insurer j's assets at the beginning of the next period, $t+1$. Hence, insurers j's assets evolve according to the law of motion

$$A_{jt+1} = R^A_{jt+1}\left[A_{jt} + \sum_{i\in\mathcal{I}} \left(P_{ijt}Q_{ijt} - F_{ijt}\right)\right]. \tag{2}$$

When issuing new policies, insurers add to their existing liabilities, L_{jt}, through the creation of reserves. We refer to V_{it} as product i's reserve value. The total reserves created through the issuance of policy i at time t is then $V_{it}Q_{ijt}$.[11] We denote the return on an insurer's stock of existing reserves as R^L_{jt}. We then refer to the return on a particular product's reserves as R_{it}, which we assume is fixed constant across insurers.[12] Similar, we use L_{jt} to denote insurer j's liabilities at the beginning of period t, before the issuance of new products. Insurer j's liabilities therefore evolve according to

$$L_{jt+1} = R^L_{jt+1}L_{jt} + \sum_{i\in\mathcal{I}} R_{it+1}V_{it}Q_{ijt}. \tag{3}$$

Combining (2) and (3) therefore gives us the evolution of insurers' capital:

$$\begin{aligned} K_{jt+1} &= A_{jt+1} - L_{jt+1} \\ &= \underbrace{R^A_{jt+1}A_{jt} - R^L_{jt+1}L_{jt}}_{\text{Legacy Capital, } \equiv \widetilde{K}_{jt+1}} + \underbrace{\sum_{i\in\mathcal{I}} \left[\left(R^A_{jt+1}P_{ijt} - R_{it+1}V_{it}\right)Q_{ijt} - R^A_{jt+1}F_{ijt}\right]}_{\text{Return on New Policy Issuance in Period } t} \end{aligned} \tag{4}$$

where insurer j's legacy capital, $\widetilde{K}_{jt+1}$, is their level of capital at $t+1$ if they did not issue any new policies in period t. Insurer j's capital evolution therefore depends on the financial returns from their legacy capital and the return on new policy issuance.

Asset and reserve returns have two components: a guaranteed component (e.g., coupon payments, bonds maturing, policy claims, and lapsation) and a revaluation component due

[11]Statutory values for insurance policies are typically more conservative than their actuarial value, which can also affect pricing (Koijen and Yogo, 2015). For our purposes, this distinction is not necessary.

[12]This implies that R^L_{jt} is determined through the composition of insurer j's outstanding insurance policies.

to changes in market interest rates, $\Delta R_{t+1} = R_{t+1} - R_t$.[13] We assume returns take the form

$$R^A_{jt+1} = \overline{R}^A_{jt+1} - D^A_{jt}\Delta R_{t+1}, \quad R^L_{jt+1} = \overline{R}^L_{jt+1} - D^L_{jt}\Delta R_{t+1}, \quad R_{it+1} = \overline{R}_{it+1} - D_{it}\Delta R_{t+1},$$

where the guaranteed components of returns $\overline{R}^A_{jt+1}, \overline{R}^L_{jt+1}$ and $\overline{R}_{it+1}$ are assumed to be exogenous, reflecting the characteristics of the underlying securities. We refer to D^A_{jt} as insurer j's asset duration, D^L_{jt} as insurer j's liability duration, and D_{it} as policy i's duration, as they measure the sensitivities of the returns to the market interest rate.

Insurers have two objectives. First, they maximize their operating profits. Second, they minimize the volatility of their growth rate. Let R^K_{jt+1} denote insurer j's capital growth rate from period t to $t+1$,

$$R^K_{jt+1} = \frac{K_{jt+1}}{K_{jt}} = \frac{\widetilde{K}_{jt+1}}{K_{jt}} + \sum_{i\in\mathcal{I}} \frac{\left(R^A_{jt+1}P_{ijt} - R_{it+1}V_{it}\right)Q_{ijt} - R^A_{jt+1}F_{ijt}}{K_{jt}}.$$

We assume insurers are risk averse, and capture their risk management motives through an decreasing and concave function $\Lambda_j(R^K_{jt+1} - \mathbb{E}_t[R^K_{jt+1}])$, where the expectation $\mathbb{E}_t[\cdot]$ is taken over the distribution of policy rate innovations, ΔR_{t+1}. Therefore, their objective function can be summarized as

$$\max_{\{P_{ijt},F_{ijt}\}} \underbrace{\sum_{i\in\mathcal{I}} \Big[(P_{ijt} - V_{it})\, Q_{ijt} - F_{ijt} \Big]}_{\text{Operating Profits}} + \underbrace{\mathbb{E}_t\left[\Lambda_j\Big(R^K_{jt+1} - \mathbb{E}_t[R^K_{jt+1}]\Big)\right]}_{\text{Risk Management}}.$$

One can interpret $\Lambda_j(\cdot)$ as insurer j's disutility for the volatility of its capital growth rate. For example, if $\Lambda_j(x) = -\gamma_j x^2$, the risk management component of the objective function becomes $-\gamma_j \mathrm{Var}_t(R^K_{jt+1})$, and we can interpret the objective function as a mean-variance preference that trades off the operating profits and the variance of capital growth.

In what follows, we will use a first-order approximation of $\Lambda_j(\cdot)$ around legacy returns, $R^K_{jt+1} \approx \widetilde{R}^K_{jt+1}$, where the legacy return $\widetilde{R}^K_{jt+1}$ is the return on capital from t to $t+1$ absent any new policy issuance during period t:

$$\widetilde{R}^K_{jt+1} \equiv \frac{\widetilde{K}_{jt+1}}{K_{jt}} = \frac{R^A_{jt+1}A_{jt} - R^L_{jt+1}L_{jt}}{K_{jt}}.$$

[13] One could argue that claims and lapsation rates themselves are both inherently random (e.g., Gottlieb and Smetters, 2021; Koijen et al., 2024; Kubitza et al., 2025). Since our framework considers atomistic households, after aggregating, we treat the idiosyncratic components of such risks as diversified. We also let the returns be time-dependent, which allows for aggregate claim and lapsation risks.

Then, the approximation of $\Lambda_j(\cdot)$ can be written as follows:

$$\Lambda_j\Big(R^K_{jt+1} - \mathbb{E}_t[R^K_{jt+1}]\Big) \approx \Lambda_j\Big(\widetilde{R}^K_{jt+1} - \mathbb{E}_t[\widetilde{R}^K_{jt+1}]\Big) + \frac{\Lambda'_j\Big(\widetilde{R}^K_{jt+1} - \mathbb{E}_t[\widetilde{R}^K_{jt+1}]\Big)}{K_{jt}} \times$$
$$\sum_{i\in\mathcal{I}}\Bigg[(R^A_{jt+1} - \overline{R}^A_{jt+1})(P_{ijt}Q_{ijt} - F_{ijt}) - (R_{it+1} - \overline{R}_{it+1})V_{it}Q_{ijt}\Bigg]. \tag{5}$$

The first term is independent of the product issuance decisions made by the insurer, and therefore is taken as given. The second term captures the marginal value of risk management from the issuance of new products, and is the relevant piece of our model. For notational convenience, we denote $\lambda_{jt+1} \equiv \Lambda'_j\Big(\widetilde{R}^K_{jt+1} - \mathbb{E}_t[\widetilde{R}^K_{jt+1}]\Big)/K_{jt}$, which depends on the realization of interest rates ΔR_{t+1} but not on products issued in period t. Formally, insurer j solves

$$\max_{\{P_{ijt},F_{ijt}\}} \overbrace{\sum_{i\in\mathcal{I}}\Big[(P_{ijt} - V_{it})Q_{ijt} - F_{ijt}\Big]}^{\text{Operating Profits}}$$
$$+ \underbrace{\mathbb{E}_t\Bigg[\lambda_{jt+1}\sum_{i\in\mathcal{I}}\Big((R^A_{jt+1} - \overline{R}^A_{jt+1})(P_{ijt}Q_{ijt} - F_{ijt}) - (R_{it+1} - \overline{R}_{it+1})V_{it}Q_{ijt}\Big)\Bigg]}_{\text{Expected Value of Risk Management}}. \tag{6}$$

The insurer trades off its immediate profits with its expected return on its capital in the next period. The expectation is taken over the distribution of market rate innovations, ΔR_{t+1}. The choice of product prices and agent distribution in the current period will therefore depend on the insurer's *interest rate risk* and, in particular, the strength of the insurer's risk management motive, λ_{jt+1}.

3.2 Duration Gaps and Liability Rebalancing

Given the trade-off between profits and return risk, how should an insurer design its product portfolio? To study this question, we first need to understand the determinants of pricing and agent distribution and, therefore, their product issuance. We begin by characterizing the optimal decisions of a given insurer in the following lemma.

Lemma 1: Optimal Issuance Decisions

Insurer j's optimal price for product i and the optimal number of agents hired to sell product i satisfy

$$\frac{P_{ijt}}{V_{it}} = \left(\frac{\varepsilon_{it}}{\varepsilon_{it}-1}\right)\mathcal{M}_{ijt}, \qquad T_{ijt} = \max\left\{(\kappa')^{-1}\left(\frac{\eta_{it}}{\mathcal{E}_{it}\overline{Q}_{ijt}\mathcal{M}_{ijt}^{1-\varepsilon_{it}}}\right),\ 0\right\},$$

where $\mathcal{E}_{it} \equiv \varepsilon_{it}^{-\varepsilon_{it}}(\varepsilon_{it}-1)^{\varepsilon_{it}-1}$ and the risk management markup, $\mathcal{M}_{ijt}$, satisfies

$$\mathcal{M}_{ijt} = \frac{1+\mathbb{E}_t\Big[\lambda_{jt+1}\big(R_{it+1}-\overline{R}_{it+1}\big)\Big]}{1+\mathbb{E}_t\Big[\lambda_{jt+1}\big(R^A_{jt+1}-\overline{R}^A_{jt+1}\big)\Big]}.$$

Proof: See Appendix A.1.

For a given product, both prices and agent distribution depend explicitly on the returns to that product's reserve value as well as to its interaction with the insurer's marginal value of risk management. Risk management markups, $\mathcal{M}_{ijt}$, are increasing in $\mathbb{E}_t[\lambda_{jt+1}(R_{it+1} - \overline{R}_{it+1})]$. In other words, the insurer charges higher markups on liabilities that grow faster than expectation ($R_{it+1} > \overline{R}_{it+1}$) when the marginal benefit of risk management (λ_{jt+1}) is high. To examine this case, we consider an approximation of λ_{jt+1} around $\Delta R_{t+1} = 0$:

$$\lambda_{jt+1} \approx \underbrace{\frac{\Lambda_j'(0)}{K_{jt}}}_{\equiv \bar{\lambda}_{jt+1}} - \underbrace{\frac{\Lambda_j''(0)}{K_{jt}}}_{\equiv \bar{\lambda}'_{jt+1} < 0} D^K_{jt}\Delta R_{t+1}. \tag{7}$$

where $D^K_{jt} \equiv (D^A_{jt}A_{jt} - D^L_{jt}L_{jt})/K_{jt}$ is insurer j's duration gap. Since the function capturing the risk management motive $\Lambda_j(\cdot)$ is concave, $\bar{\lambda}'_{jt+1} \equiv \Lambda_j''(0)/K_{jt} < 0$. As highlighted in Section 2.3, many life insurers faced a negative duration gap after the financial crisis. This fact is of first order when analyzing insurers' pricing and issuance patterns. To do so, we use the following lemma to understand how a product's duration affects its pricing.

Lemma 2: Approximate Risk Management Markups

Suppose that the market interest rate R_t follows a martingale process with variance σ_t^2. Then under the approximation (7), risk management markups $\mathcal{M}_{ijt}$ can be written as

$$\mathcal{M}_{ijt} = \frac{1 + (\bar{\lambda}'_{jt+1} D^K_{jt} \sigma^2_{t+1}) D_{it}}{1 + (\bar{\lambda}'_{jt+1} D^K_{jt} \sigma^2_{t+1}) D^A_{jt}}. \tag{8}$$

Proof: See Appendix A.2.

The lemma highlights an important result: if insurers face a negative duration gap, $D^K_{jt} < 0$, then long duration policies have higher markups, all else equal. Since insurers are risk-averse over capital returns, they put a higher weight on capital losses than they do capital gains. Therefore, when they have a negative duration gap, their value of capital losses due to interest rate declines outweighs their value of capital gains due to interest rate hikes. They therefore set a higher price on long-duration policies when this gap is larger to justify the higher potential losses. Higher prices further translate into reduced agent distribution and commissions as they lower the profitability of long-duration policies. Equipped with this insight, we present our first result.

Proposition 1: Interest Rate Risk and Product Issuance

Consider two interest rate environments, 1 and 2. The interest rate uncertainty in the second environment is higher, $\sigma^2_{2,t+1} > \sigma^2_{1,t+1}$. Then for any insurer j such that $D^K_{jt} < 0$,

$$Q^2_{ijt} > Q^1_{ijt} \qquad \text{if } D_{it} < D^A_{jt}$$

$$Q^2_{ijt} < Q^1_{ijt} \qquad \text{if } D_{it} > D^A_{jt}$$

Proof: See Appendix A.3.

Proposition 1 says that if interest rate uncertainty increases, then relative to their asset duration, insurers with a negative duration gap decrease the issuance of long-duration products and increase the issuance of short-duration products. Since their duration gap is negative, their capital is already exposed to interest rate risk. Therefore, they optimally move away from long-duration products that exacerbate their duration gap in an attempt to hedge additional interest rate risk.

We next explore how this result changes in the cross-section of insurers in different interest rate environments. In particular, we are interested in the role of capital *convexity*. If some insurers have especially convex liabilities — such as insurers that previously issued variable life insurance or annuities with generous minimum return guarantees (Koijen and Yogo, 2022; Sen, 2023) — then in a low rate environment, their duration gap should increase. This makes them especially susceptible to interest rate risk, even if the volatility of interest rates remains unchanged.

Denote the convexity of an insurer's capital as $\gamma^K_{jt} = -\partial D^K_{jt+1}/\partial R_{t+1} < 0$. The following proposition considers how two insurers with different capital convexity respond to a decline in interest rates, holding fixed interest rate volatility.

Proposition 2: Capital Convexity and Product Issuance

Consider two interest rate environments, 1 and 2, that are identical except that interest rates are lower in the second environment, $R^2_t < R^1_t$. In addition, consider two insurers, j and j', that are identical except that insurer j' has more convex capital, $|\gamma^K_{j't}| > |\gamma^K_{jt}|$. Then,

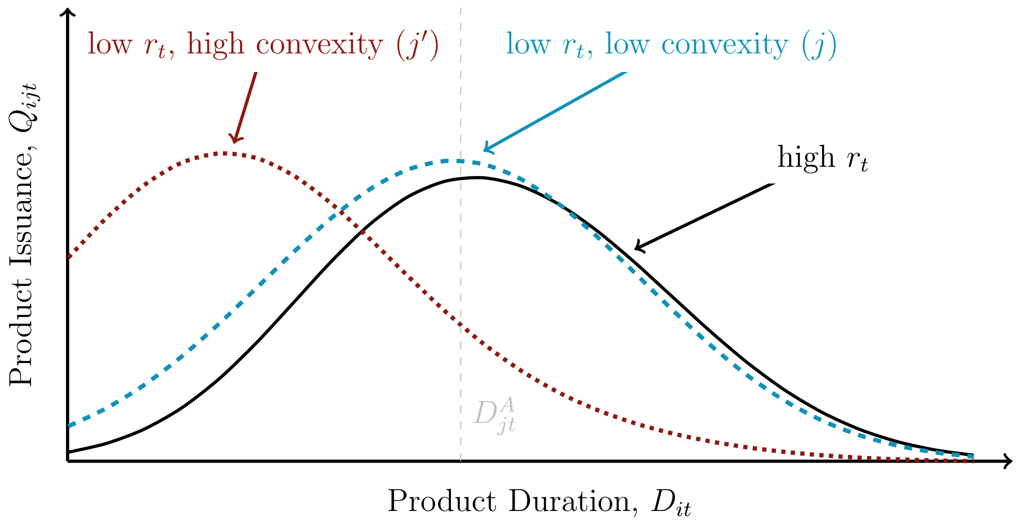
$$\frac{Q^2_{ij't}}{Q^1_{ij't}} > \frac{Q^2_{ijt}}{Q^1_{ijt}} > 1 \qquad \textit{if } D_{it} < D^A_{jt}$$

$$\frac{Q^2_{ij't}}{Q^1_{ij't}} < \frac{Q^2_{ijt}}{Q^1_{ijt}} < 1 \qquad \textit{if } D_{it} > D^A_{jt}$$

Proof: See Appendix A.4.

We summarize Proposition 2 in Figure 2. The Figure plots the product issuance distribution of the two insurers, j and j'. Initially, in a high-interest-rate environment, the two insurers have the same duration gaps and issue products with the same intensity. In the meantime, insurer j' has a higher convexity of capital than insurer j, $|\gamma^K_{j't}| > |\gamma^K_{jt}|$, for example due to previously issuing variable annuities with generous guarantees. Hence, as they transition into an environment with lower rates, the duration gap of j' becomes more negative than the duration gap of j. Both insurers respond to lower rates by shifting their issuance to-

FIGURE 2: INTEREST RATE RISK, CAPITAL CONVEXITY, AND PRODUCT ISSUANCE



Note: This figure presents hypothetical product issuance curves as a function of product duration. The black curve reflects the decisions of two insurers with identical duration gaps in a high interest rate environment. The red dotted and blue dashed line respectively reflect decisions of the more convex and less convex insurer when interest rates decline. The faint dashed gray line represents their shared asset duration.

ward low-duration policies, but since insurer j' is especially sensitive, their response is more pronounced.

It is important to note that the results of this section are partial equilibrium results. If a large insurer such as Metlife responds to a decline in rates by no longer selling long-duration policies, less exposed insurers may step in to fill the gap in demand despite also having some exposure to the decline in rates. We therefore turn to an analysis of product market equilibrium to study the market level effects of interest rate risk and duration gaps.

3.3 Duration Mismatch and the Size of Insurance Markets

We begin by zooming in on household purchasing behavior. For simplicity, we assume that households may hold multiple life insurance policies and treat each product market in isolation.[14] We assume households have identical preferences within a product class, but that their preferences may differ across product classes. Household h's indirect utility from purchasing product i sold from insurer j is

$$u^h_{ijt} = \log \alpha_j + \log \kappa_{ijt}(T_{ijt}) - (\varepsilon_{it} - 1) \log \left(\frac{P_{ijt}}{V_{it}} \right) + \nu^h_{ijt}$$

[14]This is not an unrealistic assumption: according to data from the 2018 Health and Retirement Survey, of the 54% of households that hold a life insurance policy, 38% of households hold more than one policy.

where α_j is an insurer-specific characteristic ("quality") and ν^h_{ijt} is an idiosyncratic taste shock distributed according to an extreme value type I distribution with unit variance.[15] Household h spends a constant amount, Y^h_{it}, on coverage through product i. They therefore purchase $Q^h_{ijt} = Y^h_{it}/P_{ijt}$ units of coverage conditional on buying from insurer j. Households may also choose an outside option 0 (e.g., cash) with preferences satisfying $u^h_{i0t} = \log \alpha^0_{it}$. We normalize the price of the outside option to 1. With these assumptions, insurer j faces the following demand curve

$$Q_{ijt}(P_{ijt}, T_{ijt}) = \kappa(T_{ijt}) \frac{Y_{it}}{P_{ijt}} \left(\frac{P_{ijt}/V_{it}}{\mathcal{P}_{it}} \right)^{1-\varepsilon_{it}},$$

where aggregate expenditures, Y_{it}, and the product market price index, $\mathcal{P}_{it}$, respectively satisfy

$$Y_{it} \equiv \int_0^1 Y^h_{it} dh, \qquad \mathcal{P}^{1-\varepsilon_{it}}_{it} \equiv \alpha^0_{it} + \sum_{j \in \mathcal{J}} \alpha_j \kappa(T_{ijt}) \left(\frac{P_{ijt}}{V_{it}} \right)^{1-\varepsilon_{it}}.$$

We also introduce a functional form for the market penetration function,

$$\kappa(T_{ijt}) = 1 - \exp(-T_{ijt}).$$

This functional form allows us to solve for $\mathcal{P}_{it}$ in closed form, which greatly simplifies the analysis.

We begin by addressing the question asked at the end of Section 3.2: in response to a decline in interest rates, how do insurers adjust within a product market when we account for cross-sectional differences in capital convexity? As highlighted by Huber (2022), some insurers did not see a large decline in their duration gaps post-2008 and should therefore respond differently than insurers whose duration gaps widened. The following result highlights a condition that determines whether or not the competitive effects of reduced issuance by highly exposed firms outweigh the direct effects of additional exposure by other firms.

[15] We include market penetration explicitly in indirect utility for simplicity. The interpretation is that if insurer j has more agents, they are more accessible, which reduces the cost of search or travel for households. One could alternatively model market penetration as the share of households reached by the insurer, but the resulting price index would only be an approximation when there are a finite number of insurers.

Proposition 3: Insurer Substitution

Consider two interest rate environments, 1 and 2, that are equivalent except that all insurers face a more severe duration gap in the second environment. Formally, $|D_{jt}^{K,2}| \geq |D_{jt}^{K,1}|$ for all j with a strict inequality for at least one j. Let $\psi_{ijt} = \mathcal{M}_{ijt}^2 / \mathcal{M}_{ijt}^1$ be the ratio of risk management markups in the two environments.

If $D_{it} > D_{jt}^A$ for all j, then there exists a threshold $\overline{\psi}_{it}$ such that

$$Q_{ijt}^2 < Q_{ijt}^1 \qquad \text{if } \psi_{ijt} > \overline{\psi}_{it}$$

$$Q_{ijt}^2 > Q_{ijt}^1 \qquad \text{if } \psi_{ijt} < \overline{\psi}_{it}$$

The reverse inequalities are true if $D_{it} < D_{jt}^A$ for all j.

Proof: See Appendix A.5.

The partial equilibrium setting of Section 3.2 suggested that even lightly more exposed insurers alter their behavior, and that insurers whose interest rate risk exposure does not change ($D_{jt}^K = 0$) do not adjust their issuance. Instead, in equilibrium, Proposition 3 says that the retreat of the exposed insurers opens up demand for the unexposed insurers, leading them to increase their issuance. This occurs both due to an increase in the number of agents and, therefore, the share of households that they reach, as well as cross-insurer substitution by market participants. We will see in the following section that this pattern holds in the data.

Nevertheless, it is unclear whether unexposed insurers can fully pick up the slack left by the exposed insurers. For example, if lower-quality insurers are the ones with higher exposure, we might expect the higher-quality insurers to easily buy up the policies that they left on the table. However, this may not be sufficient if households' preferences are sufficiently dispersed or if the decreasing returns to scale implied by their market penetration is too strong.

To study this trade-off, note that we can write the share of expenditures that accrue to the outside option as

$$\frac{Q_{it}^0}{Y_{it}} = \alpha_{it}^0 \mathcal{P}_{it}^{\varepsilon_{it}-1} = \frac{\alpha_{it}^0}{\alpha_{it}^0 + \sum_{j \in \mathcal{J}} \alpha_j \kappa_{ijt} (P_{ijt}/V_{it})^{1-\varepsilon_{it}}}. \tag{9}$$

Holding fixed the outside option value α_{it}^0, a ubiquitous increase in prices at the market level points to an increase in the outside option share, and therefore, a decline in the expenditures spent on insurance. Since prices are increasing while expenditures are falling, this would immediately imply that total new coverage issued should decline as well. The following result confirms this finding conditional on insurers having the same initial exposure.

Proposition 4: Product Market Issuance Dynamics

Consider two interest rate environments, 1 and 2, that are equivalent except that all insurers face a more severe duration gap in the second environment, i.e., $|D_{jt}^{K,2}| \geq |D_{jt}^{K,1}|$ *for all* j *with a strict inequality for at least one* j. *Additionally, assume that* $\mathcal{M}_{ijt}^{1}$ *is constant across insurers. Then the total issuance of product* i *satisfies*

$$Q_{it}^{2} < Q_{it}^{1} \qquad \text{if } D_{it} > D_{jt}^{A} \text{ for all } j$$

$$Q_{it}^{2} > Q_{it}^{1} \qquad \text{if } D_{it} < D_{jt}^{A} \text{ for all } j$$

Proof: See Appendix A.6.

Therefore, according to Proposition 4, a decline in rates that renders all insurers' duration gap more negative leads to a reduction in market issuance for long-duration policies but increases market issuance for short-duration policies. With these results in hand, we now turn to our empirical setting: life insurance markets during the post-GFC, low-interest-rate period.

4 The State of Life Insurance After the Financial Crisis

Equipped with the model predictions, we now turn to our empirical analysis. We begin by discussing our data sources and our method for identifying exposed insurers. We then present results on liability rebalancing and issuance dynamics for exposed and non-exposed insurance groups. We end with an exploration of aggregate issuance dynamics and the evolution of life insurance markets over the last two decades.

4.1 Data Construction

Statutory Filings Much of our data are sourced from life insurers' statutory filings, which we access through S&P Global. Every insurer in the United States must prepare these filings annually for the National Association of Insurance Commissioners (NAIC), who then provides these data to institutions for research purposes.

We pull from a variety of exhibits in the statutory filings. Our primary data is from the Exhibit of Life Insurance, which provides detailed information on policies and coverage issued and in force (gross and net). The exhibit separately identifies ordinary life (term and whole life policies) and group life lines of business. For the latter, there are two policy categories: group contracts and group certificates. Contracts reflect insurer-firm relationships, while certificates are a measure of the number of insured individuals. When using policy-level data, we use certificates.

We complement these data with reserve positions, premiums, and commissions for each product category. Reserves are taken from the Aggregate Reserves for Life Contracts. The filings record the reserve positions (gross and net) for each product category at the end of the fiscal year. Premiums and commissions come from Exhibit 1.

Data on variable annuity issuance and holdings come from the General Interrogatories. These filings record the total related account value for each annuity product sold as well as the reserves held in the general account by the insurer. Note that the account values and reserves only reflect minimum return guarantees since insurers hold the principal of the annuities in their separate accounts.

Finally, we use information on insurers' assets and liabilities, which further allows us to produce leverage ratios. For summary statistics, we use data from the Interest Sensitive Life Insurance Products Report. We also use stock prices of publicly traded insurers[16] from CRSP. Data on Treasury yields and annual GDP are taken from FRED.

Insurance Prices Our data on insurance prices are taken from Compulife, a software system used by insurance agents that generates quotes for various product categories and insurance companies. We collect monthly quotes for 10-, 15-, 20-, and 30-year term life products between January 2008 and December 2022. Quotes are for 40-year-old, non-smoking men in regular health. The data only contain quotes for a select sample of insurers. We discuss the representativeness later in this section, and Wenning (2024) provides a detailed breakdown of a broader subset of annual data.

Sample Construction Our unit of analysis is an insurance group. We choose to use insurance groups over individual companies for two reasons. First, many insurance groups organize their subsidiaries according to their product specialization. For example, among the

[16]We adopt the same list of insurers as in Koijen and Yogo (2022).

subsidiaries of the insurance group MetLife Inc., Brighthouse Financial was a large issuer of variable annuities and variable life insurance. Separating Brighthouse Financial from other subsidiaries, such as the flagship company Metropolitan Life Insurance Company, would paint an incomplete picture of MetLife as a whole. Second, insurers are publicly traded at the insurance group level. Since most public insurers also issued variable annuities, it is consistent with existing evidence on duration gaps and stock returns to use insurance groups (e.g., Hartley et al., 2016; Koijen and Yogo, 2022; Li, 2024).

Our theory predicts that insurers whose liabilities are more convex are more exposed to interest rate risk. Variable annuities are a particularly convex liability due to their minimum return guarantees as discussed in Section 2.2. We therefore split insurance groups by their variable annuities exposure, measured as the total related account value of their variable annuities divided by their total liabilities. We label an insurer as "exposed" if its variable annuity share of liabilities is in the top decile of insurers between 2005 and 2007. Note that only about 25% of insurance groups in our sample issue variable annuities during this time period, so our cutoff corresponds to approximately the top 40% of variable annuity issuers.

Note that we exclude insurers that were not in an insurance group between 2005-2007 for most of the analysis. This is done to provide a clean comparison between exposed and non-exposed insurers prior to the crisis. We bring these insurers back into the sample when we explore aggregate product market trends for completeness.

We also exclude captive reinsurers from our insurance group definitions. This is of little consequence when studying trends in product issuance since reinsurers typically do not issue new policies. However, as we will see later in this section, adding them back into the sample when studying market-level trends does not change the results in the time series. This exclusion also prevents large jumps in the exposed insurers' insurance in force due to the split between MetLife and RGA.

Summary Statistics We provide summary statistics for our primary sample in Table 1. We split the table on two dimensions. First, we report summary statistics for exposed and non-exposed insurance groups separately. Second, we report the statistics for 2005-2008 and 2009-2023 separately. We refer to the first time period as the pre-crisis period and the second time period as the post-crisis period. There are 26 (25) exposed insurers and 239 (198) non-exposed insurers in the pre-crisis (post-crisis) period. Exposed insurers are relatively more represented in the Compulife data, although there is still sufficient variation across the two groups in both time periods.

Exposed insurers are systematically larger than non-exposed insurers. In particular, the

Table 1: Summary Statistics

	Exposed Insurers		Non-Exposed Insurers	
	2005-2008	2009-2023	2005-2008	2009-2023
Number of Groups				
Full Sample	26	25	239	198
Compulife Sample	12	15	39	43
Assets	94.68	100.30	8.31	14.57
Surplus	5.09	5.39	0.67	1.25
Leverage Ratio	19.62	19.17	6.56	8.97
Leverage Ratio (Weighted)	20.13	21.15	17.94	16.26
VA Liability Share	0.57	0.50	0.01	0.01
IS Reserve Share	0.67	0.65	0.24	0.25
Issuance Market Share				
Ordinary	0.43	0.29	0.54	0.61
Group	0.45	0.42	0.54	0.51
In Force Market Share				
Ordinary	0.38	0.29	0.37	0.39
Group	0.48	0.44	0.49	0.47

Note: This table reports summary statistics for our primary sample. Assets and surplus are reported in billions of dollars. All variables except market shares, the weighted leverage ratio, and the number of groups are unweighted averages across insurers. The weighted leverage ratio is weighted by insurer assets within each period. Market shares are calculated across all years within each period.

average exposed insurer is 11.4 times as large as the average non-exposed insurer in the pre-crisis period and 6.88 times as large in the post-crisis period. This is consistent with variable annuity issuance being dominated by large insurers: since variable annuities are among the most complex products issued by life insurers, it is likely that only large insurance groups have adequate resources to manage them. Exposed insurers also have more capital (surplus), though only by an order of 7.6 and 4.3 in the pre- and post-crisis periods, respectively. This difference suggests that exposed insurers are more levered: the average leverage ratio, calculated as liabilities divided by surplus, is 3 and 2.1 times the average leverage of non-exposed insurers in the pre- and post-crisis periods, respectively. When weighting insurers' leverage by assets, exposed insurers still have nearly 12.2% higher leverage in the pre-crisis period and 30.1% higher leverage in the post-crisis period.

Consistent with our definition of variable annuity exposure, exposed insurers have substantially higher variable annuity liabilities as a share of total liabilities.[17] This is not surprising, as the majority of non-exposed insurers do not issue variable annuities at all. That being said, certain life insurance products are also recorded as interest-sensitive and may be exposed to the low-rate environment in the post-crisis period. The table suggests that insurers exposed to variable annuities also have a substantially higher exposure to interest-sensitive life insurance policies.

Table 1 also preempts our findings across product markets. In the pre-crisis period, exposed insurers, despite being small in number, accounted for 43% of total ordinary life insurance issuance. The remaining 90% of insurance groups accounted for 54% of the issuance, with the remainder being issued by small non-group companies. The numbers for group life issuance are similar. However, in the post-crisis period, exposed insurers only issued 29% of new ordinary life insurance coverage, with non-exposed insurers increasing their share to 61%. On the other hand, group life issuance shares remained relatively stable.

The decline in ordinary life issuance is echoed when considering life insurance in force. Exposed insurers decreased their market share of life insurance coverage in force from 38% to 29% between the two periods, while non-exposed insurers' market share increased from 37% to 39%. Group life insurance in force again remained relatively stable. Note that the numbers for ordinary life only add up to 75%; the majority of the remaining insurance was held by reinsurers, and within reinsurers was largely held by RGA, a prior subsidiary of MetLife until their split in 2008.

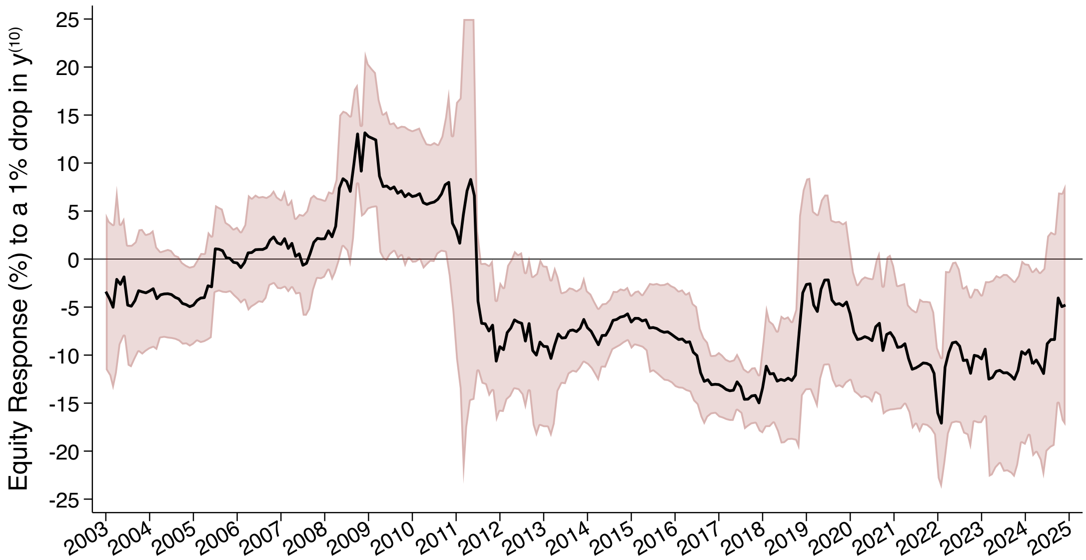
4.2 The Evolution of Duration Gaps

We begin our analysis by documenting the widening of life insurers' duration gaps following the financial crisis. Many studies (Berends et al., 2013; Ozdagli and Wang, 2019; Huber, 2022; Li, 2024) provide evidence of larger duration gaps using rolling estimates of interest rate betas for a portfolio of public insurers' returns. We replicate this finding in Figure 3, where we estimate

$$\text{Insurance Portfolio Returns}_t = \alpha + \beta \times \text{Market Returns}_t - \gamma \times \Delta\text{10-Year Yield}_t + \varepsilon_t. \quad (10)$$

[17] Note that insurers' liabilities are calculated differently than variable annuity liabilities. The share reported in the table and used for our classification is merely meant to separate those with high exposure from those with low exposure relative to their size.

Figure 3: Rolling Estimates of Life Insurers' Duration Gaps



Note: This figure reports rolling regression estimates of equation (10). The black line plots estimates of the interest rate sensitivity, γ, for two-year (24-month) rolling windows. Red bands report 95% confidence intervals using heteroscedasticity-robust standard errors.

We use monthly returns and consider rolling two-year intervals to calculate β and γ. We plot the results, along with 95% confidence bands, in Figure 3. As in the aforementioned studies, we find that after approximately 2011, the life insurance portfolio exhibits a consistently negative duration gap for all time periods except 2019.

However, our theory predicts that certain insurance companies were more exposed to interest rate risk than others. In particular, as we highlighted in the previous section and as shown by Sen (2023), insurers that are heavily exposed to variable annuities should have faced even more intense interest rate risk exposure. It is difficult to show this using stock returns since most public companies issued variable annuities prior to the crisis. Therefore, we instead rely on a more direct estimation strategy following Huber (2022). Recall that the duration gap of insurer j at time t can be written as

$$D_{jt}^{K} \equiv D_{jt}^{A} + \frac{L_{jt}}{K_{jt}} \times (D_{jt}^{A} - D_{jt}^{L}). \tag{11}$$

Therefore, in order to estimate insurers' duration gaps directly, we need three objects. First, we use liabilities, L_{jt}, and surplus capital, K_{jt}, from insurers' statutory filings. Second, we use liability duration estimates, D^L_{jt}, from Huber (2022), which run from 2005-2020.[18] Finally, we approximate D^A_{jt} using the duration of insurers' corporate bond portfolios, the main asset class held by insurers, which we calculate directly using information on their corporate bond holdings (Schedule D in the statutory filings).

We plot the cumulative change in insurers' (asset-weighted) average duration gaps for both exposed and non-exposed insurance groups in panel (a) of Figure 4.[19] The average duration gap across the full sample of insurers exhibits similar patterns as in Figure 3: a brief uptick in the duration gap for the first two years after the financial crisis, followed by a persistent decrease in the duration gap that is only mitigated around 2018 and 2019. The consistency across methods is encouraging evidence that duration gaps at the industry level became increasingly negative after the financial crisis.

When we split the sample into exposed and non-exposed insurers, we see that the majority of the decline after the financial crisis was driven by exposed insurers. After the crisis, exposed insurers experienced a large and persistent decline in their duration gaps, while non-exposed insurers saw a brief decline in 2011 and virtually no changes until 2020. This is consistent with our discussion in Section 2.3 and the previous section that highlighted the interest rate sensitivity of variable annuity issuers' liabilities.

While these results are suggestive, we would like to highlight two caveats with this measure of duration gaps. First, the liability duration measure from Huber (2022) actually excludes variable annuity minimum return guarantee reserves. The differences are therefore driven primarily by changes in leverage and other liabilities. Including reserves for minimum return guarantees would likely exacerbate differences across exposed and non-exposed groups, further validating our findings.

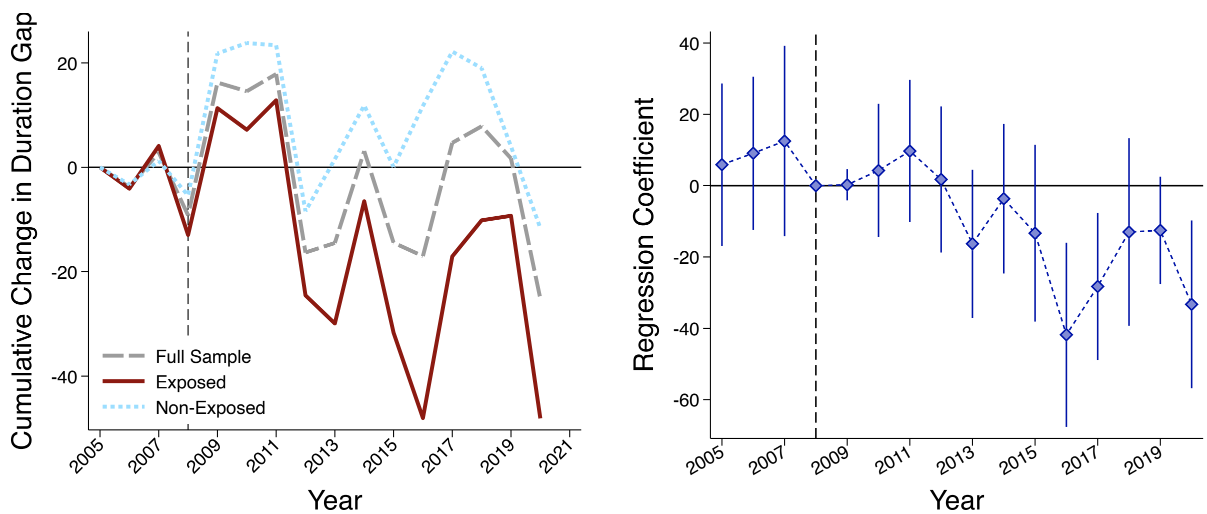
Second, the magnitudes of the changes in duration gaps using this measure are large in some years. This is primarily due to our use of surplus capital (as opposed to economic capital) in measuring leverage ratios. As we highlighted in Table 1, exposed insurers experienced an increase in their (asset-weighted) surplus-based leverage ratio from roughly 20.1 to 21.1 ($\approx 5\%$), while non-exposed insurers experienced a decline in their leverage ratio from 17.9 to 16.3 ($\approx -10\%$). Therefore, even small differences in asset and liability durations are exacerbated by differences in leverage across insurers and across time.[20] We therefore

[18]We would like to thank Max Huber for making these data publicly available to researchers.

[19]We trim insurer-level duration gaps at the 1% and 99% level across the full sample to account for substantial outliers. However, as shown in Appendix Figure C.1, the patterns in the figure remain unchanged even when these observations are not removed.

[20]We explore a decomposition of duration gap changes in Appendix B.1 and show in Figure C.2 that leverage played an important role in exacerbating the difference in duration gaps across groups.

Figure 4: Changes in Duration Gaps by Exposure



(a) Duration Gap Changes, Raw Data

(b) Gap Differences, Regression Output

Note: This figure documents changes in the insurers' duration gaps [panel (a)] and estimates $\{\beta_\tau\}_{\tau=2005}^{2020}$ from regression (12) [panel (b)] over time. Duration gaps are constructed as in equation (11). Red lines reflect exposed insurers, blue dotted lines reflect non-exposed insurers, and gray dashed lines reflect the full sample. The figures report within-sample, asset-weighted averages. In panel (b), blue spikes represent 95% confidence intervals using standard errors clustered at the yearly level.

interpret our results as suggestive evidence of insurers' differential exposure to interest rate risk rather than interpreting the magnitudes as precise.

To further address concerns regarding mismeasurement, we also consider a more rigorous test of the changes in duration gaps across exposed and non-exposed insurers. We estimate the regression

$$D_{jt}^K = \sum_{\tau=2005}^{2020} \beta_\tau \mathbf{1}\{t = \tau\} \times \text{Exposed}_j + \delta_j + \delta_t + \varepsilon_{jt}, \tag{12}$$

which allows us to control for mismeasurement in the average level of duration gaps due to our inclusion of time fixed effects and mismeasurement at the insurer-level due to our inclusion of insurer fixed effects. The estimates $\{\beta_\tau\}_{\tau=2005}^{2020}$ are therefore a more robust measure of how duration gaps changed across our two groups over time. As with the raw data, we weight observations by insurers' assets. We set 2008 to be our omitted year, so the estimates are relative to the year 2008. We present our results graphically in panel (b) of Figure 4. There is a clear negative trend that begins after 2011, which is consistent with the raw differences in panel (a) and the results in Figure 3. The estimates are large and significant at the 5% level for many of the years after 2011, although some years have more noise than others.

Overall, the evidence in this section points to an increase in interest rate risk exposure by insurance companies that were exposed to variable annuities. Having validated this, we now turn to an analysis of insurers' product market behavior.

4.3 The Effect of Interest Rate Risk on Insurance Pricing

Our theory predicts that the changes in relative duration gaps across insurers should manifest in their pricing decisions. This section tests this formally using the model as a guide. Consider two products, ℓ and s, in which $D_{\ell t} > D_{st}$. Lemma 2 suggests that to first order, we have that[21]

$$\overbrace{\underbrace{\mathbb{E}_{\text{Ex}}\left[\log\frac{P_{\ell jt}/V_{\ell t}}{P_{sjt}/V_{st}}\right]}_{\substack{\text{long-short markup}\\ \text{for exposed insurers}}} - \underbrace{\mathbb{E}_{\text{NonEx}}\left[\log\frac{P_{\ell jt}/V_{\ell t}}{P_{sjt}/V_{st}}\right]}_{\substack{\text{long-short markup}\\ \text{for non-exposed insurers}}}}^{\text{long-short markup spread}}$$

$$\approx \sigma^2_{t+1} \times \left(\mathbb{E}_{\text{Ex}}\left[\bar{\lambda}'_{jt} D^K_{jt}\right] - \mathbb{E}_{\text{NonEx}}\left[\bar{\lambda}'_{jt} D^K_{jt}\right]\right) \times (D_{\ell t} - D_{st}) \quad (13)$$

We refer to the left-hand-side of equation (13) as the (ℓ, s) markup spread. It measures how long-duration products are priced relative to short-duration products, comparing exposed insurers to non-exposed insurers. Our theory makes two predictions. First, it predicts that, due to their larger interest rate risk exposures, exposed insurers will have a higher long-short markup spread. This implies that (13) should be positive. Second, our theory predicts that this markup spread should increase when interest rates are low, as exposed insurers have more convex capital. Therefore, (13) should co-move negatively with long-term yields.

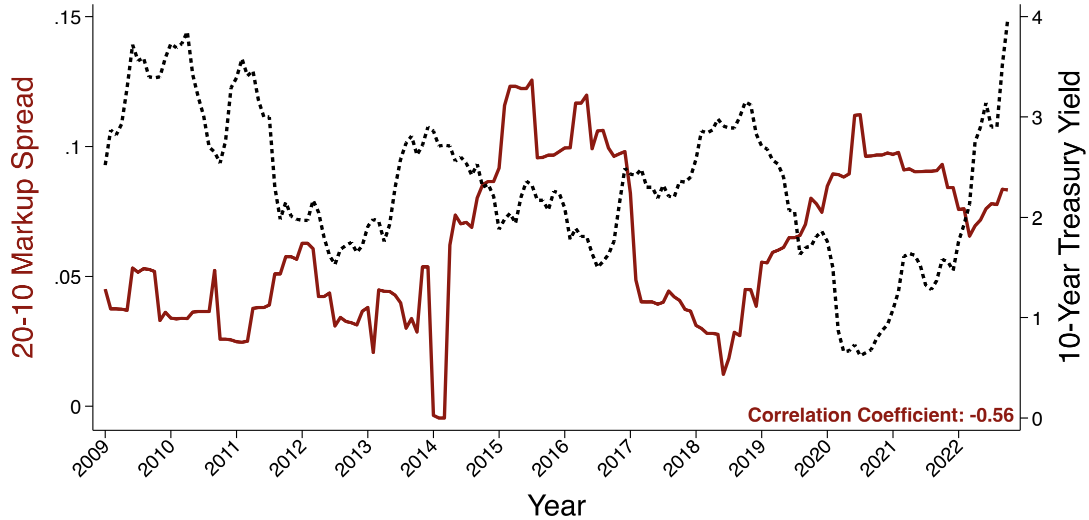
We plot the (20,10) markup spread for each month between January 2009 and December 2022 in Figure 5.[22] When calculating averages as in equation (13), we weight insurers by their total assets to capture the most important insurers within each group.[23] We also overlay monthly 10-year Treasury yields. Both conditions highlighted above hold: the markup spread is positive each month, is generally higher during periods of low interest rates, and co-moves

[21] See Section B.2 for a complete derivation.

[22] As Koijen and Yogo (2015) show, the late months of 2008 displayed some extraordinary pricing behavior due to regulatory accounting practices. In particular, more distressed insurers reduced prices of long-term products relative to short-term products as a result of the differences between actuarially fair and statutory reserve values. We therefore start our sample in 2009 to avoid this episode and focus on time periods where interest rate risk was more pronounced.

[23] Note that using assets as weights may put too much emphasis on large insurers that are not very active in ordinary life insurance markets. We show in Appendix Figure C.3 that our results hold (and in fact, become slightly stronger) when weighting by ordinary life insurance in force as well.

Figure 5: The Interaction between Product Pricing and Interest Rate Risk



Note: This figure plots the (20,10) markup spread (red, left axis) and the 10-year Treasury yield (black dotted, right axis) for each month between January 2009 and December 2022. When calculating the markup spread, averages are weighted by assets.

negatively with 10-year yields.[24] Over the full sample, the correlation coefficient between the relative markup spread and yields is -0.56.

There could be other explanations for why the relative markup spread co-moves negatively with rates. For example, perhaps exposed insurers are simply larger and more active in long-duration term insurance markets, giving them more pricing power over time. It is also possible that, since exposed insurers are more likely to issue complex products, there may be product-specific unobservables driving the results. As such, we test our results formally in a regression framework, which allows us to control for such differences. We consider a triple difference specification in which we compare two product categories at a time,

$$\log \text{Price}_{ijt} = \beta \times y_t^{(10)} \times \text{Exposed}_j \times \text{Long}_i + \delta_{jt} + \delta_{it} + \delta_{ij} + \varepsilon_{ijt}, \tag{14}$$

where $y_t^{(10)}$ is the monthly 10-year Treasury yield, Exposed_j is an indicator for whether insurer j is in the exposed group and Long_i is an indicator for whether product i has the longer duration of the two product categories. We interpret $\beta < 0$ as evidence that exposed insurers (relative to non-exposed insurers) set higher prices on longer-maturity products (relative to shorter-maturity products) when interest rates are low. We consider every long-

[24]We verify in Appendix Figure C.4 that this result is not driven by changes in interest rate uncertainty, as equation (13) suggests could be the case.

short combination of 10-, 15-, and 20-year term policies.[25] To avoid potential issues from the differences in samples across policies, our baseline results consider insurer-month observations in which the insurer sells all policy categories, though we show in Table C.1 that our results are robust to including the remaining observations. Finally, as in Figure 5, we weight the regression by insurers' assets; however, the results are robust to using ordinary life insurance in force as weights, as presented in Appendix Table C.2.

We include several granular fixed effects to address the myriad concerns above. First, we include insurer × month fixed effects, which remove insurer-specific time variation (e.g., differences in size). Second, we include insurer × product fixed effects, which removes unobservable, time-invariant differences in product characteristics (e.g., convertibility clauses or renewal benefits). Third, we include month × product fixed effects, which absorb differences in demand over time for different product categories.

To further validate the prediction of Proposition 1 (i.e., interest rate uncertainty increases insurers' interest rate risk exposure and relative maturity markups), we estimate a similar triple difference specification where the 10-year Treasury yield is replaced with the monetary policy uncertainty (MPU) index by Husted et al. (2020),

$$\log \text{Price}_{ijt} = \beta \times \text{MPU}_t \times \text{Exposed}_j \times \text{Long}_i + \delta_{jt} + \delta_{it} + \delta_{ij} + \varepsilon_{ijt}. \quad (15)$$

We use the uncertainty in monetary policy decisions as a proxy for the uncertainty in interest rates. Proposition 1 predicts $\beta > 0$, meaning that exposed insurers (relative to non-exposed insurers) set higher prices on longer-maturity products (relative to shorter-maturity products) when interest rate uncertainty is high.

We report our results in Table 2. For all long-short pairs, we find that exposed insurers set higher markups for longer-duration products relative to non-exposed insurers when 10-year Treasury yields are low. Additionally, for each long-term category, we find a monotonically declining relationship between relative prices and duration differences. For example, the (20,15) category has a coefficient of -0.018, while the (20,10) category has a coefficient of -0.023. Furthermore, exposed insurers increase markups for longer-duration products when monetary policy uncertainty is higher. The effects are again more pronounced for product

[25] We exclude 30-year term policies in our main analysis for two reasons. First, about 20% of companies in the Compulife data do not sell or report any 30-year term life prices. Given the small initial number of firms, this may bias the results. Second, due to the high cost of 30-year policies, households may be less willing to let these policies lapse, as repurchasing them at an older age would be significantly more costly relative to a shorter-term policy. This may lower the effective duration of 30-year policies, making the cross-maturity comparison less insightful. Nevertheless, we present results for 30-year policies in Appendix Table C.4, where we find similar monotonicity patterns as our main results.

Table 2: Interest Rates and Prices — Regression Results

	Dependent Variable: $\log \text{Price}_{ijt}$					
(Long, Short) Category:	(15,10)	(20,15)	(20,10)	(15,10)	(20,15)	(20,10)
$y_t^{(10)} \times \text{Exposed}_j \times \text{Long}_i$	−0.006***	−0.018***	−0.023***			
	(0.002)	(0.002)	(0.003)			
$\text{MPU}_t \times \text{Exposed}_j \times \text{Long}_i$				0.007***	0.013***	0.020***
				(0.002)	(0.002)	(0.003)
Insurer × Month FE	✓	✓	✓	✓	✓	✓
Insurer × Product FE	✓	✓	✓	✓	✓	✓
Month × Product FE	✓	✓	✓	✓	✓	✓
Observations	8956	8956	8956	8956	8956	8956
Within-R^2	0.001	0.023	0.020	0.002	0.014	0.017

Note: This table reports regression results for equations (14) and (15). The dependent variable is the log of the premium quote for product i sold by insurer j in month t. $y_t^{(10)}$ is the monthly 10-year Treasury yield, MPU_t is the Husted et al. (2020) monetary policy uncertainty index normalized to a mean of 1, Exposed_j is an indicator equal to 1 if insurer j is in the exposed group, and Long_i is an indicator equal to 1 if product i has the longer maturity of the two product categories in the regression. Standard errors clustered at the product-time level are reported in parentheses. Observations are weighted by insurer j's assets in the year corresponding to month t. * $p < 0.1$ ** $p < 0.05$ *** $p < 0.01$.

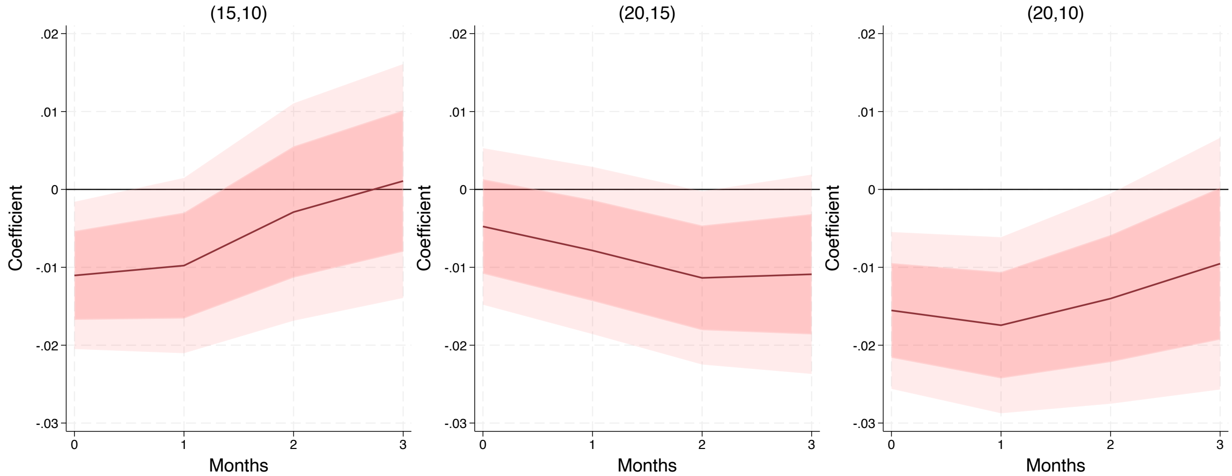
pairs with a larger duration difference. The coefficients estimated for the (20,15) and (20,10) categories are 0.013 and 0.020, respectively. Appendix Table C.2 shows that the results are as striking when we weight observations by ordinary life insurance in force instead of assets. Appendix Table C.3 further confirms that the results are robust to controlling for insurers' total assets in a similar triple-interacted specification.

To further sharpen the identification, we examine the responses of insurance prices to movements in $y_t^{(10)}$ around FOMC meetings. Formally, we estimate the local projection

$$\Delta_3^k \log \text{Price}_{ijt} = \beta \times \left(\Delta_3 y_t^{(10)} \mid \text{FOMC}\right) \times \text{Exposed}_j \times \text{Long}_i + \delta_{jt} + \delta_{it} + \delta_{ij} + \varepsilon_{ijt}, \quad (16)$$

where the dependent variable $\Delta_3^k \log \text{Price}_{ijt}$ is the change in the log price of product i issued by insurer j from month $t-3$ to $t+k$. $\left(\Delta_3 y_t^{(10)} \mid \text{FOMC}\right)$ is the total change in the 10-year Treasury yield over two-day windows around FOMC meetings as in Li (2024), aggregated

Figure 6: Price Responses to Long Rates



Note: This figure reports the β coeffcients estimated from (16), for horizons $k \in \{0, 1, 2, 3\}$. The three panels correspond to estimates for the (15,10), (20,15), and (20,10) categories, respectively. Standard errors are clustered at the product-time level, and the shaded areas indicate 68% and 90% confidence intervals.

over the three-month period between month $t-3$ and month t.[26] Figure 6 plots the impulse responses of prices to FOMC innovations in the long rate. The estimates for $k = 0$ capture the contemporaneous responses of quarterly log price changes to quarterly long rate shocks, while those for $k > 1$ capture longer-term price reactions. Consistent with Table 2, insurers mark up long-term products following negative innovations in the long rate. The responses persist for at least 3 months and are larger and more significant for product pairs with larger maturity differences.

We have therefore shown that insurance companies set prices at least in part based on their interest rate risk exposure. Longer-term products are marked up more when interest rate risk increases, such as in low-interest-rate and high-uncertainty periods. While this mechanism is telling, we need to verify that the pricing behavior we observe is consistent with their product issuance. Unfortunately, we do not observe term life issuance at the insurer-maturity level. However, instead of looking *within* product categories, we can instead focus our analysis on *across* product categories to capture differences in product duration. We therefore turn to a comparison of ordinary and group life issuance.

[26] Hillenbrand (2025) shows that such movements in the long rate around FOMC meetings are not predictable by macroeconomic news.

4.4 Liability Rebalancing

We have now shown that insurers exposed to variable annuities before the crisis experienced larger declines in their duration gaps and, in the context of term life insurance, passed through the additional interest rate risk to their prices. But how did they adjust their entire portfolio of liabilities? Our theory suggests that exposed insurers may have an incentive to shift their product issuance toward low-duration policies in response to an increase in interest rate risk. Group life policies, which are typically renewable yearly, have low reserve valuations, and carry very little interest rate risk, provide a natural alternative to long-dated term or whole life insurance policies.[27]

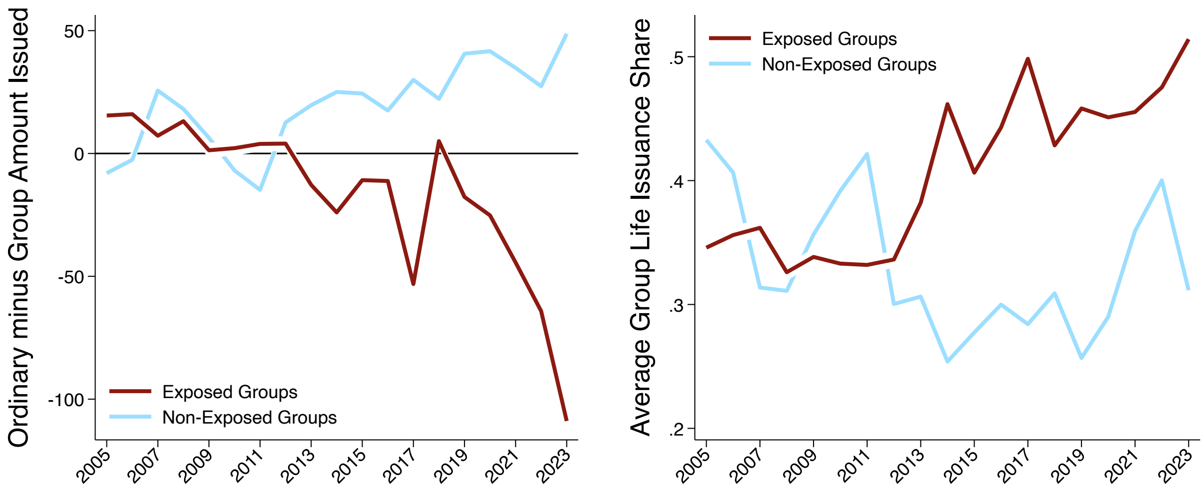
We therefore begin by exploring how the issuance of ordinary and group life products changed throughout the post-crisis period in Figure 7. Panel (a) reports the asset-weighted average difference between ordinary and group life issuance for each group of insurers, conditional on issuing both types of products. Units are in billions of nominal dollars. Panel (b) reports the asset-weighted average share of group life issuance across insurers within each group.

The figure strongly confirms the predictions of the theory. On average, exposed insurance groups reduced their relative issuance of ordinary life insurance coverage over the sample period. Notably, the decline begins after the financial crisis and accelerates after the drop in yields and the widening of duration gaps in 2011. At the same time, we see that non-exposed groups began to *increase* their relative issuance after 2011. This is consistent with non-exposed groups capturing demand that was previously allocated to exposed groups.

The results on relative product issuance could be influenced by a few large firms. For example, MetLife was a large factor in the strong growth in group life issuance. Panel (b) shows that this is not a driving factor: average group life issuance *shares* also increased substantially for exposed groups while remaining stable for non-exposed groups. Since shares remove potentially large size differences across insurers, we interpret this as suggestive evidence of liability rebalancing. We further check for robustness in Appendix Figure C.6 by reproducing Figure 7 using unweighted averages within exposed and non-exposed groups, as well as in Appendix Figure C.7, where we reproduce the figure after excluding MetLife from the sample. The trends are generally unchanged: relative issuance of ordinary life insurance declines for exposed groups, while increasing or remaining unchanged for non-exposed groups. Similarly, group life issuance shares increased from 35% to over 50% for exposed groups while remaining virtually unchanged for non-exposed groups.

[27] See Appendix B.3 for a discussion on the differences in reserve values across these two product categories.

Figure 7: Product Issuance Across Insurance Groups Over Time



(a) Relative Issuance

(b) Issuance Shares

Note: This figure reports average ordinary life insurance issuance relative to average group life issuance [panel (a)] and average group life issuance shares [panel (b)] for exposed (red) and non-exposed (blue) insurers over time. Averages are weighted by assets within each class of insurers. For panel (a), units are in billions of US dollars.

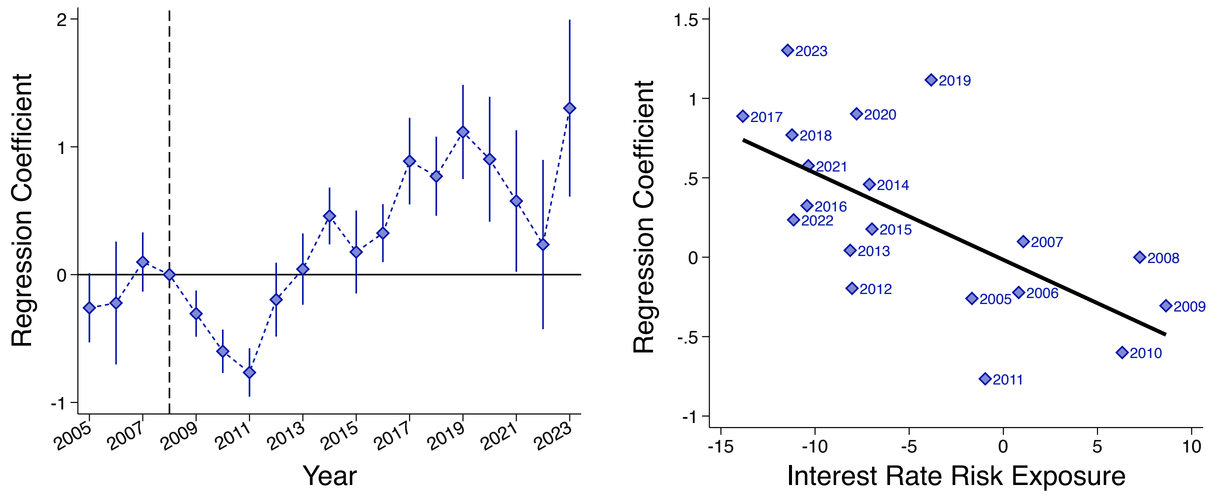
We now turn to a more careful analysis of liability rebalancing. We estimate the Poisson regression

$$\log \mathbb{E}[\text{Issuance}_{ijt}] = \sum_{\tau=2005}^{2023} \beta_\tau \mathbf{1}\{\tau = t\} \times \text{Exposed}_j \times \text{Group}_i + \delta_{ij} + \delta_{jt} + \delta_{it} + \varepsilon_{ijt} \qquad (17)$$

where Exposed_j is an indicator equal to 1 if insurer j is exposed to variable annuities in the pre-crisis period and Group_i is an indicator for group life policies. For our main specification, we use insurance coverage as our dependent variable. We choose a Poisson specification to flexibly accommodate zeros in the dependent variable, as some insurers do not underwrite one of the two categories of insurance.

We include a variety of fixed effects that alleviate a number of concerns with the descriptive analysis above. First, insurer-product fixed effects, δ_{ij}, control for the possibility that some insurers have a particular connection to certain product types. For example, some insurers are more likely to sell simple term products, while others are more likely to sell complicated universal life policies equipped with guarantees. Second, insurer-year fixed effects, δ_{jt}, help control for time-varying insurer characteristics that may influence the types of products they issue, such as age or the size of their distribution system. Third, product-time

Figure 8: Liability Rebalancing — Regression Results



(a) Regression results over time

(b) Regression results vs Risk Exposures

Note: This figure reports regression results for equation (17). The regression estimates a set of coefficients $\{\beta_\tau\}_{\tau=2005}^{2023}$, each of which represents the relative insurance coverage issued (i) between exposed and non-exposed insurers and (ii) between group and ordinary life insurance products. Panel (a) plots the estimated coefficients across time, while panel (b) reports the estimated coefficients against the yearly average interest rate risk exposure taken from the estimates in Figure 3. In panel (a), blue spikes represent 95% confidence intervals using standard errors clustered at the product-time level.

fixed effects, δ_{it}, control for time-varying demand conditions for each product market that could drive issuance.

Our estimates of interest are the parameters $\{\beta_\tau\}_{\tau=2005}^{2023}$, which represent the difference in issuance across group and non-group life insurance policies, as well as across exposed and non-exposed groups, for each year. A positive coefficient for a given year indicates that, relative to non-exposed groups, exposed groups issue relatively more group life insurance than ordinary life insurance in that year relative to the base year, which we take to be 2008. Given the granular set of fixed effects and the evidence of interest rate risk differences across exposed and non-exposed insurers documented in the previous two sections, we interpret a positive β_τ as evidence of liability rebalancing as a risk management strategy.

We present our estimates in Figure 8. Panel (a) displays the estimates $\{\beta_\tau\}_{\tau=2005}^{2023}$ for each year along with 95% confidence intervals. There are two clear patterns to note. First, during and immediately after the financial crisis, the coefficient is negative, which implies exposed insurers shifted toward ordinary life issuance and away from group life issuance relative to non-exposed insurers. However, this relationship flipped after about 2013, implying a shift toward group life issuance for exposed insurers relative to non-exposed insurers.

This finding is consistent with the results on changes in duration gaps in Section 4.2. Recall that duration gaps actually became positive for a brief period of time after the financial crisis before becoming negative after around 2011 (see Figure 3 and panel (b) of Figure 4). We show this directly by plotting the estimated coefficients $\{\beta_\tau\}_{\tau=2005}^{2023}$ against the average yearly interest rate risk exposure taken from Figure 3. We find a strong negative correlation: when interest rate risk exposures (duration gaps) are negative, exposed insurers are more biased toward group life issuance relative to non-exposed insurers. The opposite is true when duration gaps are positive, as our theory would predict.

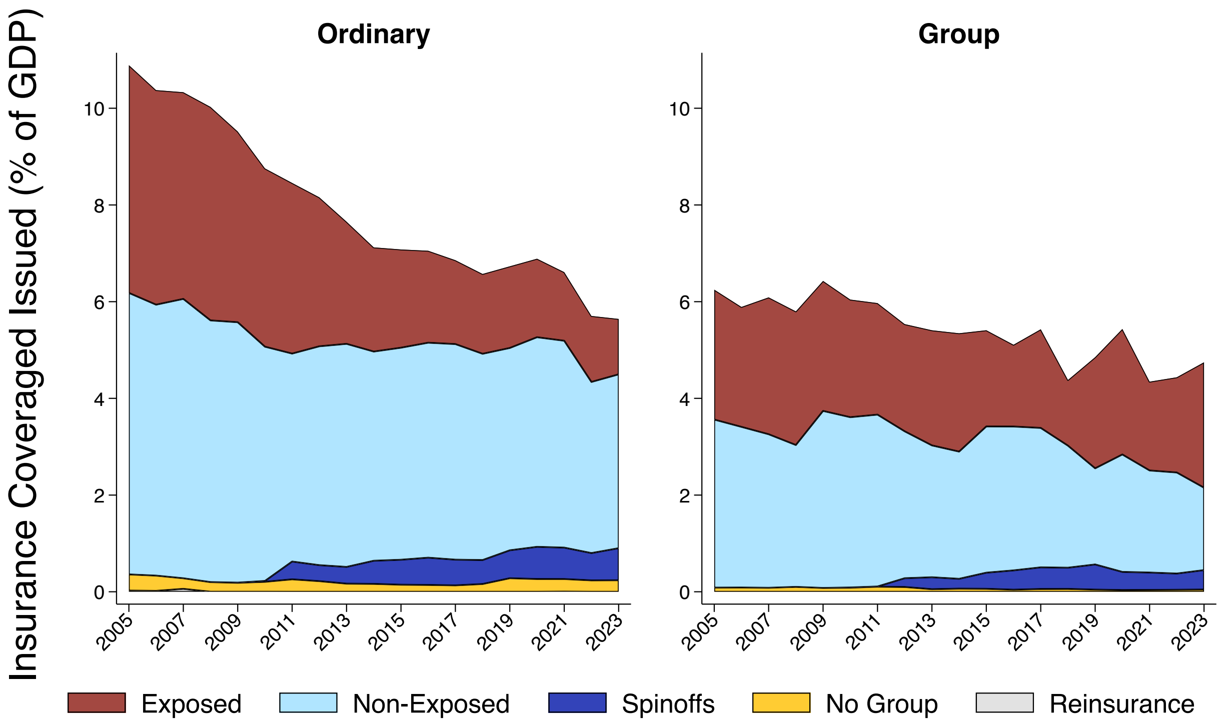
We conduct a variety of robustness exercises to validate our claim. Appendix Figure C.8 reports the results using number of policies instead of the dollar amount of insurance coverage, which gives similar results. Appendix Figure C.9 excludes MetLife from the analysis to ensure that MetLife's dramatic shift toward group life issuance does not drive the results. We find that the results hold for both insurance coverage and the number of policies, although the estimates are slightly noisier. Appendix Figure C.10 confirms that the patterns are robust to further controlling for insurer size interacted with time and Group_i. Finally, Appendix Figure C.11 reports the results of an OLS regression using log issuance, and considers both weighted and unweighted samples. Again, the results hold broadly, indicating strong support for liability rebalancing as an interest rate risk management channel.

4.5 Aggregate Product Market Dynamics

We now turn to the market-level effects of liability rebalancing. In addition to exposed and non-exposed groups, we also consider the issuance of three other categories of insurance companies for completeness. First, we include spin-offs, e.g., companies that were part of either an exposed or non-exposed insurance group in the pre-crisis period but later left to form their own group (e.g., Brighthouse Financial departing with MetLife in 2017). Second, we include insurers that are not a part of an insurance group. Third, we include reinsurers.

Although ordinary life issuance declined for exposed insurers, this does not necessarily imply that aggregate issuance declined. In particular, if non-exposed insurers more than picked up the slack, it could be that issuance was stable over time at the market level. The first panel of Figure 9 suggests this is not the case: measured as a percentage of GDP, aggregate ordinary life issuance fell by 48.2% (10.9% to 5.6% of GDP) between 2005 and 2023. This was predominantly driven by exposed insurers (4.7% to 1.15% of GDP), while issuance also fell for non-exposed insurers to a much lesser extent (5.8% to 3.6% of GDP). Spinoffs and no-group insurers added only a small amount relative to the issuance of the

Figure 9: Aggregate Issuance By Product



Note: This figure reports real aggregate life insurance issuance as a percentage of real GDP from 2005 to 2023. The first panel reflects ordinary life issuance, and the second panel reflects group life issuance. Red areas represent exposed insurance groups, light blue areas represent non-exposed insurance groups, dark blue areas reflect insurance companies that belonged to either the exposed or non-exposed insurance groups in the pre-crisis period but have since spun off, yellow areas represent insurers not in a life insurance group, and gray areas reflect reinsurance companies.

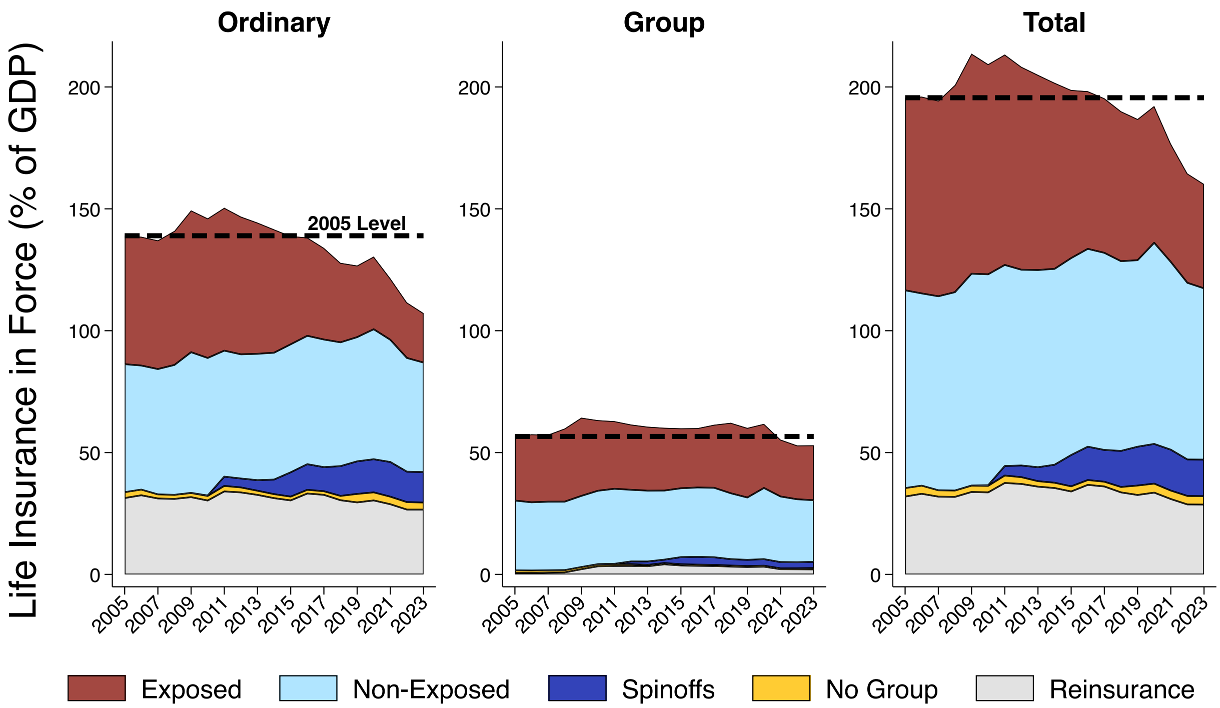
primary groups in the sample.

Group life issuance, seen in the second panel of Figure 9, also fell as a percentage of GDP, but by a smaller amount (6.2% to 4.7% of GDP). Importantly, exposed insurers, aside from a few years, only slightly decreased their issuance relative to 2005 (2.7% to 2.6% of GDP). Non-exposed insurers decreased their issuance more (3.5% to 1.7% of GDP), likely due to increased competition in this market by exposed insurers.

In Appendix Figure C.13, we report changes in nominal issuance for both ordinary and group life. Nominal ordinary life issuance increased by 31% for non-exposed insurers and declined by 48% for exposed insurers, while nominal group life issuance increased by 4% for non-exposed insurers and increased by a sizable 105% for exposed insurers. The discrepancy between the nominal trends and Figure 9 is the relative growth rate of GDP, which grew faster than the issuance of both products.

While issuance declined as a percentage of GDP for both ordinary and group life insur-

Figure 10: Aggregate Life Insurance Market Dynamics



Note: This figure reports real aggregate gross life insurance in force as a percentage of real GDP from 2005 to 2023. The first panel reflects ordinary life issuance, the second panel reflects group life issuance, and the third panel reflects the sum of ordinary and group life insurance. Red areas represent exposed insurance groups, light blue areas represent non-exposed insurance groups, dark blue areas reflect insurance companies that belonged to either the exposed or non-exposed insurance groups in the pre-crisis period but have since spun off, yellow areas represent insurers not in a life insurance group, and gray areas reflect reinsurance companies. Dashed black lines represent aggregate insurance in force within each panel as of 2005.

ance, this does not necessarily translate into a decline in market-level insurance coverage. Issuance may have declined due to insurers already reaching a large fraction of households; if households are not lapsing on their policies, there will be fewer households to reach in a given year, and therefore, less issuance would be expected. We therefore examine the dynamics of insurance coverage in force over time.

As shown in Figure 10, this was not the case. Although ordinary life insurance in force increased in the early part of the post-crisis period, peaking around 150.4% of GDP, it ultimately fell to 107% of GDP by 2023. While both exposed and non-exposed groups were responsible, the vast majority of the decline can be explained by exposed insurers: their life insurance in force fell from 52.7% of GDP in 2005 to 20.3% of GDP in 2023, accounting for three quarters of the decline.

Unlike issuance, group life insurance in force remained stable throughout most of the post-crisis period, only moderately declining relative to initial levels after the COVID-19 crisis. Consistent with our hypothesis, exposed insurers were the key difference between ordinary and group life market dynamics. Putting the two together, life insurance in force at the industry level fell from 213.2% to 160% of GDP.

5 Conclusion

Interest rate risk is of first order to many financial institutions. During the low interest rate period that accompanied the recovery from the financial crisis, exposure to interest rate risk grew for many of these institutions. In particular, due to the long-term nature of their liabilities and issues of market incompleteness and regulatory frictions, many life insurance companies had their equity squeezed by low rates.

We provide theory and evidence that insurers with especially convex liabilities, such as variable annuities, may retreat from long-duration product markets to reduce their exposure to interest rate risk. While they substitute toward short-duration products to an extent, the industry as a whole may not remain stable if there are substantive differences in product market characteristics. This appears to be the case for life insurers today: group life insurance markets did not grow enough to offset the decline in ordinary life insurance markets, resulting in a shrunken system.

Our analysis, while telling, abstracts from simultaneous fluctuations and trends in insurance demand. This is a relevant omission since the low interest rate environment coincided with a sharp economic recession and a sluggish recovery, both of which would put downward pressure on already declining tastes for standard insurance products. That being said, our model is amenable to estimation that would allow us to separate demand-based changes from supply-side contractions due to duration mismatch and interest rate risk. We intend to carry out such an analysis in the future.

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A Model Proofs

A.1 Proof of Lemma 1

The proposition follows from the first order condition for product i. In particular, note that we can combine terms and rewrite i's contribution to the objective function as

$$\left[\Big(1+\mathbb{E}_t[\lambda_{jt+1}(R^A_{jt+1}-\overline{R}^A_{jt+1})]\Big)P_{ijt}-\Big(1+\mathbb{E}_t[\lambda_{jt+1}(R_{it+1}-\overline{R}_{it+1})]\Big)V_{it}\right]Q_{ijt}$$
$$-\Big(1+\mathbb{E}_t[\lambda_{jt+1}(R^A_{jt+1}-\overline{R}^A_{jt+1})]\Big)F_{ijt}.$$

The first order condition with respect to P_{ijt} is then

$$\Big(1+\mathbb{E}_t[\lambda_{jt+1}(R^A_{jt+1}-\overline{R}^A_{jt+1})]\Big)Q_{ijt}$$
$$+\left[\Big(1+\mathbb{E}_t[\lambda_{jt+1}(R^A_{jt+1}-\overline{R}^A_{jt+1})]\Big)P_{ijt}-\Big(1+\mathbb{E}_t[\lambda_{jt+1}(R_{it+1}-\overline{R}_{it+1})]\Big)V_{it}\right](1-\varepsilon_{it})\frac{Q_{ijt}}{P_{ijt}}=0.$$

Rearranging, we have

$$P_{ijt}=\left(\frac{\varepsilon_{it}}{\varepsilon_{it}-1}\right)\mathcal{M}_{ijt}V_{it},\qquad \mathcal{M}_{ijt}\equiv\frac{1+\mathbb{E}_t[\lambda_{jt+1}(R_{it+1}-\overline{R}_{it+1})]}{1+\mathbb{E}_t[\lambda_{jt+1}(R^A_{jt+1}-\overline{R}^A_{jt+1})]}.$$

Next, we'll solve for the optimal number of agents hired, T_{ijt}. Substituting our expression for P_{ijt} into the objective function, note that the component corresponding to product i becomes

$$\frac{1+\mathbb{E}_t[\lambda_{jt+1}(R_{it+1}-\overline{R}_{it+1})]}{\varepsilon_{it}-1}\overline{Q}_{ijt}\left(\frac{\varepsilon_{it}}{\varepsilon_{it}-1}\right)^{-\varepsilon_{it}}\mathcal{M}_{ijt}^{-\varepsilon_{it}}\kappa(T_{ijt})-\Big(1+\mathbb{E}_t[\lambda_{jt+1}(R^A_{jt+1}-\overline{R}^A_{jt+1})]\Big)\eta_{it}T_{ijt}.$$

Define $\mathcal{E}_{it}\equiv\varepsilon_{it}^{-\varepsilon_{it}}(\varepsilon_{it}-1)^{\varepsilon_{it}-1}$. The first order condition with respect to T_{ijt} can therefore be written, conditional on $T_{ijt}>0$,

$$\Big(1+\mathbb{E}_t[\lambda_{jt+1}(R_{it+1}-\overline{R}_{it+1})]\Big)\mathcal{E}_{it}\overline{Q}_{ijt}\mathcal{M}_{ijt}^{-\varepsilon_{it}}\kappa'(T_{ijt})=\Big(1+\mathbb{E}_t[\lambda_{jt+1}(R^A_{jt+1}-\overline{R}^A_{jt+1})]\Big)\eta_{it}.$$

Rearranging to isolate T_{ijt} on the left-hand side, we have

$$\kappa'(T_{ijt})=\frac{\eta_{it}}{\mathcal{E}_{it}\overline{Q}_{ijt}\mathcal{M}_{ijt}^{1-\varepsilon_{it}}}.$$

The solution follows from inverting $\kappa'(\cdot)$ to solve for T_{ijt}. Of course, $T_{ijt} \geq 0$, so the solution must be bounded below by 0. □

A.2 Proof of Lemma 2

From the approximation in (7), note that we can write the numerator of $\mathcal{M}_{ijt}$ as

$$\begin{aligned} & 1 + \mathbb{E}_t[\bar{\lambda}_{jt+1}(R_{it+1} - \overline{R}_{it+1})] - \mathbb{E}_t[\bar{\lambda}'_{jt+1} D^K_{jt} \Delta R_{t+1}(R_{it+1} - \overline{R}_{it+1})] \\ = \; & 1 + \mathbb{E}_t[\bar{\lambda}_{jt+1}(-D_{it}\Delta R_{t+1})] - \mathbb{E}_t[\bar{\lambda}'_{jt+1} D^K_{jt} \Delta R_{t+1}(-D_{it}\Delta R_{t+1})] \\ = \; & 1 - (D_{it}\bar{\lambda}_{jt+1})\mathbb{E}_t[\Delta R_{t+1}] + \bar{\lambda}'_{jt+1} D^K_{jt} \mathbb{E}_t[(\Delta R_{t+1})^2] D_{it} \end{aligned}$$

Since R_{t+1} follows a martingale process, $\mathbb{E}_t[\Delta R_{t+1}] = 0$. Substituting this into the above equation gives

$$1 + (\bar{\lambda}'_{jt+1} D^K_{jt} \sigma^2_{t+1}) D_{it}$$

as claimed. A similar series of calculations for the denominator term delivers the results. □

A.3 Proof of Proposition 1

Note that under the approximation in (7), we can write the derivative of the risk management markup with respect to σ^2_{t+1} as

$$\begin{aligned} \frac{\partial \mathcal{M}_{ijt}}{\partial \sigma^2_{t+1}} &\propto \bar{\lambda}'_{jt+1} D^K_{jt} D_{it}(1 + \bar{\lambda}'_{jt+1} D^K_{jt} D^A_{jt} \sigma^2_{t+1}) - (1 + \bar{\lambda}'_{jt+1} D^K_{jt} D_{it} \sigma^2_{t+1}) \bar{\lambda}'_{jt+1} D^K_{jt} D^A_{jt} \\ &= \bar{\lambda}'_{jt+1} D^K_{jt} \left[D_{it} - D^A_{jt} \right] \end{aligned}$$

If $D_{it} > D^A_{jt}$, then the above is expression is positive. Hence, $\mathcal{M}_{ijt}$ is increasing in σ^2_{t+1} when $D_{it} > D^A_{jt}$, so P_{ijt} is increasing in σ^2_{t+1}. From the expression for T_{ijt}, it also follows that T_{ijt} is declining in $\mathcal{M}_{ijt}$, so κ_{ijt} is declining in σ^2_{t+1}. It therefore follows that Q_{ijt} is declining in σ^2_{t+1}. On the other hand, if $D_{it} < D^A_{jt}$, then the expression above is declining. Therefore, prices decline and market penetration increases, resulting in a higher Q_{ijt}. This completes the proof. □

A.4 Proof of Proposition 2

This proof is identical to the proof of Proposition 1 except for replacing the derivative with respect to D^K_{jt}. By assumption, the two insurers are identical except for their convexity, so prior to the shift in rates, $D^{K,0}_{jt} = D^{K,0}_{j't}$. After the shift, their durations are $D^K_{jt} \approx$

$D_{jt}^{K,0} - |\gamma_{jt}^K|(R_t^2 - R_t^1) < D_{j't}^{K,0} - |\gamma_{j't}^K|(R_t^2 - R_t^1) \approx D_{j't}^K$ and therefore, $|D_{jt}^K| > |D_{j't}^K|$. Since $\bar{\lambda}'_{jt+1} < 0$ for both j, it further follows that we can write $\bar{\lambda}'_{jt+1} D_{jt}^K = |\bar{\lambda}'_{jt+1}| \times |D_{jt}^K| > 0$. Hence, it suffices to show the relationships for a change in $|D_{jt}^K|$, which will have the same form as in the proof of Proposition 1. □

A.5 Proof of Proposition 3

We will separate this proof into a few parts. First, we derive a closed form expression for the price level, $\mathcal{P}_{it}$. We then show that it's unique conditional on the change in markups. Finally, given the implied change in the price level, we show the existence and uniqueness of the cutoffs.

A.5.1 Part 1: Deriving a Closed Form Expression for the Price Level

Let $\kappa(T) = 1 - e^{-T}$. First, note that

$$\alpha_j \frac{Y_{it}}{P_{ijt}} \left(\frac{P_{ijt}/V_{it}}{\mathcal{P}_{it}} \right)^{1-\varepsilon_{it}} (P_{ijt} - V_{it}) = \alpha_j Y_{it} \mathcal{P}^{\varepsilon_{it}-1} \left(\frac{P_{ijt}}{V_{it}} \right)^{-\varepsilon_{it}} \left(\frac{P_{ijt}}{V_{it}} - 1 \right)$$
$$= \alpha_j Y_{it} \mathcal{P}^{\varepsilon_{it}-1} \mathcal{E}_{it} \mathcal{M}_{ijt}^{1-\varepsilon_{it}}$$

where $\mathcal{E}_{it} \equiv \varepsilon_{it}^{-\varepsilon_{it}} (\varepsilon_{it} - 1)^{\varepsilon_{it}-1}$. It follows then from the FOC for T_{ijt} that

$$\alpha_j Y_{it} \mathcal{P}_{it}^{\varepsilon_{it}-1} \mathcal{E}_{it} \mathcal{M}_{ijt}^{1-\varepsilon_{it}} (1 - \kappa_{ijt}) \leq \eta_{it}.$$

Therefore,

$$\kappa_{ijt} = \max \left\{ 1 - \frac{\eta_{it} \mathcal{P}_{it}^{1-\varepsilon_{it}}}{\mathcal{E}_{it} Y_{it} \alpha_j \mathcal{M}_{ijt}^{1-\varepsilon_{it}}},\ 0 \right\}.$$

We can use this expression to explicitly solve for $\mathcal{P}_{it}^{1-\varepsilon_{it}}$. Let $\mathcal{J}_{it} \subset \mathcal{J}$ denote the set of insurers that are active in product market i at time t (e.g., $j \in \mathcal{J}_{it}$ if $\kappa_{ijt} > 0$). Then we have

$$
\begin{aligned}
\mathcal{P}_{it}^{1-\varepsilon_{it}} &= \alpha_{it}^{0} + \sum_{j\in\mathcal{J}} \alpha_j \kappa_{ijt} \left(\frac{P_{ijt}}{V_{it}}\right)^{1-\varepsilon_{it}} \\
&= \alpha_{it}^{0} + \sum_{j\in\mathcal{J}_{it}} \alpha_j \left[1 - \frac{\eta_{it}\mathcal{P}_{it}^{1-\varepsilon_{it}}}{\mathcal{E}_{it} Y_{it} \alpha_j \mathcal{M}_{ijt}^{1-\varepsilon_{it}}}\right] \varepsilon_{it}\mathcal{E}_{it}\mathcal{M}_{ijt}^{1-\varepsilon_{it}} \\
&= \alpha_{it}^{0} + \varepsilon_{it}\mathcal{E}_{it} \sum_{j\in\mathcal{J}_{it}} \alpha_j \mathcal{M}_{ijt}^{1-\varepsilon_{it}} - \frac{\varepsilon_{it}\eta_{it}}{Y_{it}} |\mathcal{J}_{it}| \mathcal{P}_{it}^{1-\varepsilon_{it}}.
\end{aligned}
$$

Solving for $\mathcal{P}_{it}^{1-\varepsilon_{it}}$, it follows that

$$
\mathcal{P}_{it}^{1-\varepsilon_{it}} = \frac{\alpha_{it}^{0} + \varepsilon_{it} \sum_{j\in\mathcal{J}_{it}} \alpha_j \mathcal{E}_{it} \mathcal{M}_{ijt}^{1-\varepsilon_{it}}}{1 + \frac{\varepsilon_{it}\eta_{it}}{Y_{it}} |\mathcal{J}_{it}|}.
$$

A.5.2 Part 2: Uniqueness of the set $\mathcal{J}_{it}$

Order the set of insurers as follows: $j > j'$ if and only if $\alpha_j \mathcal{M}_{ijt}^{1-\varepsilon_{it}} > \alpha_{j'} \mathcal{M}_{ij't}^{1-\varepsilon_{it}}$. We claim that there exists a cutoff j_{it} such that $j \in \mathcal{J}_{it}$ if and only if $j \geq j_{it}$.

To show this, suppose first that there are no firms currently active in the market, $\mathcal{J}_{it} = \varnothing$. Then $\mathcal{P}_{it}^{1-\varepsilon_{it}} = \alpha_{it}^{0}$. This is an equilibrium if and only if no insurer j would find it optimal to enter, i.e.

$$
\alpha_j \mathcal{E}_{it} \mathcal{M}_{ijt}^{1-\varepsilon_{it}} < \frac{\eta_{it}\alpha_{it}^{0}}{Y_{it}} \equiv \Gamma_{it}.
$$

Note that Γ_{it} therefore defines a lower bound on j. Suppose that this condition does not hold for a positive subset of $\mathcal{J}$. For all such insurers, let $1 + \mu_{ijt} = \alpha_j \mathcal{E}_{it} \mathcal{M}_{ijt}^{1-\varepsilon_{it}} / \Gamma_{it} > 1$. Then for a given set $\mathcal{J}_{it} \subseteq \mathcal{J}$, we can express $\mathcal{P}_{it}^{1-\varepsilon_{it}}$ as

$$
\mathcal{P}_{it}^{1-\varepsilon_{it}} = \alpha_{it}^{0} \left[\omega_{it} + (1-\omega_{it}) \frac{1}{|\mathcal{J}_{it}|} \sum_{j\in\mathcal{J}_{it}} (1+\mu_{ijt})\right], \quad \omega_{it} \equiv \left(1 + \frac{\varepsilon_{it}\eta_{it}}{Y_{it}} |\mathcal{J}_{it}|\right)^{-1}.
$$

If $j \notin \mathcal{J}_{it}$, then this price index is an equilibrium price index if and only if

$$
(1+\mu_{ijt}) \frac{\alpha_{it}^{0}\eta_{it}}{Y_{it}} < \frac{\alpha_{it}^{0}\eta_{it}}{Y_{it}} \left[\omega_{it} + (1-\omega_{it}) \frac{1}{|\mathcal{J}_{i}|} \sum_{j\in\mathcal{J}_{it}} (1+\mu_{ijt})\right]. \tag{18}
$$

However, if $j' \in \mathcal{J}_{it}$ and $j > j'$, then (18) cannot hold. Hence, there exists a cutoff j_{it} that determines the equilibrium set of market participants: $\mathcal{J}_{it} = \left\{ j \in \mathcal{J} \;\middle|\; j \geq j_{it} \right\}$. □

A.5.3 Part 3: Proof of the Proposition

Let $\psi_{ijt} \equiv \mathcal{M}^2_{ijt}/\mathcal{M}^1_{ijt}$. Note that we can write

$$
\begin{aligned}
P^2_{ijt}Q^2_{ijt} &= \alpha_j Y_{it}\kappa^2_{ijt}\left(\frac{P^2_{ijt}/V_{it}}{\mathcal{P}_{it}}\right)^{1-\varepsilon_{it}} \\
&= \frac{\eta_{it}\varepsilon_{it}}{Y_{it}}\left[\frac{\mathcal{E}_{it}(Y_{it}+\eta_{it}\varepsilon_{it}|\mathcal{J}_i|)\alpha_j\psi_{ijt}^{1-\varepsilon_{it}}\mathcal{M}_{ijt}^{1-\varepsilon_{it}}}{\eta\alpha^0_{it}+\eta_{it}\varepsilon_{it}\mathcal{E}_{it}|\mathcal{J}_i|\mathbb{E}_{it}[\alpha_{j'}\psi_{ij't}^{1-\varepsilon_{it}}\mathcal{M}_{ij't}^{1-\varepsilon_{it}}]} - 1\right] \\
&= \frac{\eta_{it}\varepsilon_{it}}{Y_{it}}\left[\frac{\mathcal{E}_{it}(Y_{it}+\eta_{it}\varepsilon_{it}|\mathcal{J}_i|)\alpha_j\mathcal{M}_{ijt}^{1-\varepsilon_{it}}}{\eta\alpha^0_{it}\psi_{ijt}^{\varepsilon_{it}-1}+\eta_{it}\varepsilon_{it}\mathcal{E}_{it}|\mathcal{J}_i|\mathbb{E}_{it}\left[\alpha_{j'}\left(\frac{\psi_{ijt}}{\psi_{ij't}}\right)^{\varepsilon_{it}-1}\mathcal{M}_{ij't}^{1-\varepsilon_{it}}\right]} - 1\right]
\end{aligned}
$$

under the assumption that $\mathcal{J}_{it}$ does not change. Further, note that if $\psi_{ijt} = \max_{j'}\{\psi_{ij't}\}$, then the denominator in the above expression is strictly larger than the denominator when $\psi_{ijt} = 1$ for all j (environment 1). Therefore, $P^2_{ijt}Q^2_{ijt} < P^1_{ijt}Q^1_{ijt}$, and since $P^2_{ijt} > P^1_{ijt}$, it follows that $Q^2_{ijt} < Q^1_{ijt}$. Therefore, the insurer whose prices most respond to interest rate risk reduces their issuance with certainty.

On the other hand, note that if $\psi_{ijt} = 0$ in environment 2, then the denominator is strictly less than in environment 1, implying that $P^2_{ijt}Q^2_{ijt} > P^1_{ijt}Q^1_{ijt}$. Since $P^2_{ijt} = P^1_{ijt}$, it follows that $Q^2_{ijt} > Q^1_{ijt}$. Therefore, since the denominator is strictly increasing in ψ_{ijt}, there must exist a cutoff $\overline{\psi}_{it}$ such that if $\psi_{ijt} < \overline{\psi}_{it}$, issuance increases; otherwise, issuance declines. This completes the proof. □

A.6 Proof of Proposition 4

From the expression for the outside option share of expenditures, we have

$$
\begin{aligned}
\sigma^0_{it} = \frac{Q^0_{it}}{Y_{it}} = \alpha^0_{it}\mathcal{P}_{it}^{\varepsilon_{it}-1} &= \frac{\alpha^0_{it}}{\alpha^0_{it}+\sum_{j\in\mathcal{J}}\alpha_j\kappa_{ijt}(P_{ijt}/V_{it})^{1-\varepsilon_{it}}} \\
&= \left(\frac{\alpha^0_{it}}{\alpha^0_{it}+\varepsilon_{it}\sum_{j\in\mathcal{J}_{it}}\alpha_j\mathcal{E}_{it}\mathcal{M}_{ijt}^{1-\varepsilon_{it}}}\right)\left(1+\frac{\varepsilon_{it}\eta_{it}}{Y_{it}}|\mathcal{J}_{it}|\right).
\end{aligned}
$$

Clearly, if $\mathcal{M}_{ijt}$ increases for all insurers (which is the case when $D_{it} > D_{jt}^A$), then Q_{it}^0/Y_{it} increases as well. Since total market expenditures constant, note that $Y_{it} = \sigma_{it}^{0,1} Y_{it} + (1 - \sigma_{it}^{0,1}) Y_{it} = \sigma_{it}^{0,2} Y_{it} + (1 - \sigma_{it}^{0,2}) Y_{it}$. Since $\sigma_{it}^{0,2} > \sigma_{it}^{0,1}$, it follows that

$$\sum_{j \in \mathcal{J}_i} P_{ijt}^1 Q_{ijt}^1 = (1 - \sigma_{it}^{0,1}) Y_{it} > (1 - \sigma_{it}^{0,2}) Y_{it} = \sum_{j \in \mathcal{J}_i} P_{ijt}^2 Q_{ijt}^2.$$

Therefore, expenditures on product i decline. Note further that since $\mathcal{M}_{ijt}^1 = \mathcal{M}_{it}^1$ for all j, we necessarily have $P_{ijt}^1 = P_{it}^1$ for all j. It follows from the above inequality that

$$Q_{it}^2 = \sum_{j \in \mathcal{J}_{it}} Q_{ijt}^2 < \frac{1}{P_{it}^1} \sum_{j \in \mathcal{J}_{it}} P_{ijt}^2 Q_{ijt}^2 < \frac{1}{P_{it}^1} \sum_{j \in \mathcal{J}_{it}} P_{ijt}^1 Q_{ijt}^1 = \sum_{j \in \mathcal{J}_{it}} Q_{ijt}^1 = Q_{it}^1.$$

The result for $D_{it} < D_{jt}^A$ follows an analogous argument. □

B Additional Figures and Analyses

B.1 The Drivers of Duration Gaps

This section decomposes changes in duration gaps across the exposed and non-exposed groups into several components. To do so, define $\text{Lev}_{jt} = L_{jt}/K_{jt}$ to be insurer j's leverage ratio, and let $G_{jt} = D^A_{jt} - D^L_{jt}$ be the difference between their asset duration and liability duration (but not their duration gap, which is the duration of their capital). Note that the change in the duration gap can be written

$$\begin{aligned} \Delta D^K_{jt} &= \Delta D^A_{jt} + \Delta\Big[\text{Lev}_{jt} \times G_{jt}\Big] \\ &= \underbrace{\Delta D^A_{jt}}_{\substack{\text{Asset Duration} \\ \text{Component}}} + \underbrace{\Delta \text{Lev}_{jt} \times G_{jt}}_{\substack{\text{Leverage} \\ \text{Component}}} + \underbrace{\text{Lev}_{jt} \times \Delta G_{jt}}_{\substack{\text{Duration Mismatch} \\ \text{Component}}} + \underbrace{\Delta \text{Lev}_{jt} \times \Delta G_{jt}}_{\text{Residual}} \end{aligned} \tag{19}$$

Figure C.2 plots the four components cumulatively for exposed [panel (a)] and non-exposed [panel (b)] insurers, using 2005 as the base year. A striking pattern that emerges is that leverage is the primary driver of the decline in duration gaps for exposed insurers. This is consistent with variable annuities receiving higher capital requirements in the aftermath of the financial crisis (Koijen and Yogo, 2022), thereby exacerbating these insurers' leverage. Therefore, any pre-existing duration mismatch would be amplified. This amplification would be even more dramatic if the liability duration measure of Huber (2022) included variable annuities and other interest-sensitive liabilities.

B.2 Derivation for the Relative Markup Spread

This section derives equation (13) using the result in Lemma 2. To first order, we can write the log markup over fair value as

$$\log \frac{P_{ijt}}{V_{it}} \approx \log\left(\frac{\varepsilon_{it}}{\varepsilon_{it}-1}\right) + \bar{\lambda}'_{jt+1} D^K_{jt} \sigma^2_{t+1}\Big(D_{it} - D^A_{jt}\Big). \tag{20}$$

Consider two products ℓ and s in which $D_{\ell t} > D_{st}$. Differencing across products implies that

$$\log \frac{P_{\ell jt}/V_{\ell t}}{P_{sjt}/V_{st}} \approx \log \frac{\varepsilon_{\ell t}(\varepsilon_{\ell t}-1)^{-1}}{\varepsilon_{st}(\varepsilon_{st}-1)^{-1}} + \bar{\lambda}'_{jt+1} D^K_{jt} \sigma^2_{t+1}\Big(D_{\ell t} - D_{st}\Big). \tag{21}$$

We can then take averages across exposed and non-exposed insurers separately and take the difference between the two. This final step gives the expression in the text.

B.3 Reserve Valuation Across Products

We begin by exploring how the product-level reserve values of exposed insurers changed after the financial crisis. As we showed in Table 1, exposed groups had substantially more exposure to interest-sensitive life insurance policies in addition to their variable annuities, so we should expect their ordinary life insurance reserves to be sensitive to interest rate changes. Group life insurance, on the other hand, is yearly renewable, so its valuation should not systematically change with interest rates.

Figure C.5 confirms this finding. Panel (a) plots the average reserve value of ordinary and group life policies separately for each year in our sample.[28] Three patterns emerge. First, group life policies require substantially fewer reserves than ordinary life policies. This is due to their shorter maturities. Second, average ordinary life reserve values for exposed insurers increased by 34% (0.031 to 0.043) between 2010 and 2023, consistent with the decline in yields and the sensitivity of their reserves to interest rates. Ordinary reserve values also increased over the same time period for non-exposed insurers, but only by 11% (0.047 to 0.052). Third, despite the increase in ordinary life reserve values over the post-crisis period, exposed insurers' total reserve value remained stable. This is suggestive of liability rebalancing: as reserve values increase for ordinary life insurance, the threat of future rate changes incentivizes exposed insurers to shift their issuance away from long-duration policies and toward short-duration policies. We explore liability rebalancing in detail in the following section.

[28] Note that these averages are weighted by the total amount of insurance in force for each insurer. Insurers who have small positions in a particular category tend to have high reserve values due to a lack of diversification. Additionally, reserve values are inflated when life insurance in force is close to 0, which creates outliers.

C Additional Tables and Figures

Table C.1: Interest Rates and Prices Regression — Unbalanced

	Dependent Variable: $\log \text{Price}_{ijt}$					
(Long, Short) Category:	(15,10)	(20,15)	(20,10)	(15,10)	(20,15)	(20,10)
$y_t^{(10)} \times \text{Exposed}_j \times \text{Long}_i$	−0.006***	−0.016***	−0.019***			
	(0.002)	(0.002)	(0.003)			
$\text{MPU}_t \times \text{Exposed}_j \times \text{Long}_i$				0.005***	0.012***	0.011***
				(0.002)	(0.002)	(0.002)
Insurer × Month FE	✓	✓	✓	✓	✓	✓
Insurer × Product FE	✓	✓	✓	✓	✓	✓
Month × Product FE	✓	✓	✓	✓	✓	✓
Observations	10614	10798	11830	10614	10798	11830
Within-R^2	0.002	0.021	0.011	0.002	0.013	0.005

Note: This table reports regression results for equations (14) and (15) using the full sample of insurer-product observations. The dependent variable is the log of the premium quote for product i sold by insurer j in month t. $y_t^{(10)}$ is the monthly 10-year Treasury yield, MPU_t is the Husted et al. (2020) monetary policy uncertainty index normalized to a mean of 1, Exposed_j is an indicator equal to 1 if insurer j is in the exposed group, and Long_i is an indicator equal to 1 if product i has the longer maturity of the two product categories in the regression. Standard errors clustered at the product-time level are reported in parentheses. Observations are weighted by insurer j's assets in the year corresponding to month t. * $p < 0.1$ ** $p < 0.05$ *** $p < 0.01$.

Table C.2: Interest Rates and Prices Regression — In Force Weighted

	Dependent Variable: $\log \text{Price}_{ijt}$					
(Long, Short) Category:	(15,10)	(20,15)	(20,10)	(15,10)	(20,15)	(20,10)
$y_t^{(10)} \times \text{Exposed}_j \times \text{Long}_i$	-0.012^{***}	-0.018^{***}	-0.028^{***}			
	(0.002)	(0.002)	(0.003)			
$\text{MPU}_t \times \text{Exposed}_j \times \text{Long}_i$				0.009^{***}	0.013^{***}	0.017^{***}
				(0.002)	(0.002)	(0.002)
Insurer × Month FE	✓	✓	✓	✓	✓	✓
Insurer × Product FE	✓	✓	✓	✓	✓	✓
Month × Product FE	✓	✓	✓	✓	✓	✓
Observations	10614	10798	11830	10614	10798	11830
Within-R^2	0.006	0.026	0.022	0.004	0.014	0.010

Note: This table reports regression results for equations (14) and (15). The dependent variable is the log of the premium quote for product i sold by insurer j in month t. $y_t^{(10)}$ is the monthly 10-year Treasury yield, MPU_t is the Husted et al. (2020) monetary policy uncertainty index normalized to a mean of 1, Exposed_j is an indicator equal to 1 if insurer j is in the exposed group, and Long_i is an indicator equal to 1 if product i has the longer maturity of the two product categories in the regression. Standard errors clustered at the product-time level are reported in parentheses. Observations are weighted by insurer j's ordinary life insurance in force in the year corresponding to month t. * $p < 0.1$ ** $p < 0.05$ *** $p < 0.01$.

Table C.3: Interest Rates and Prices Regression — Size Control

	Dependent Variable: $\log \text{Price}_{ijt}$					
(Long, Short) Category:	(15,10)	(20,15)	(20,10)	(15,10)	(20,15)	(20,10)
$y_t^{(10)} \times \text{Exposed}_j \times \text{Long}_i$	-0.004*	-0.015***	-0.019***			
	(0.002)	(0.002)	(0.003)			
$y_t^{(10)} \times \text{Assets}_{jt} \times \text{Long}_i$	-0.005***	-0.006***	-0.011***			
	(0.001)	(0.001)	(0.002)			
$\text{MPU}_t \times \text{Exposed}_j \times \text{Long}_i$				0.008***	0.012***	0.020***
				(0.002)	(0.002)	(0.003)
$\text{MPU}_t \times \text{Assets}_{jt} \times \text{Long}_i$				0.004***	-0.003***	0.001
				(0.001)	(0.001)	(0.001)
Insurer × Month FE	✓	✓	✓	✓	✓	✓
Insurer × Product FE	✓	✓	✓	✓	✓	✓
Month × Product FE	✓	✓	✓	✓	✓	✓
Observations	8956	8956	8956	8956	8956	8956
Within-R^2	0.009	0.044	0.052	0.011	0.023	0.017

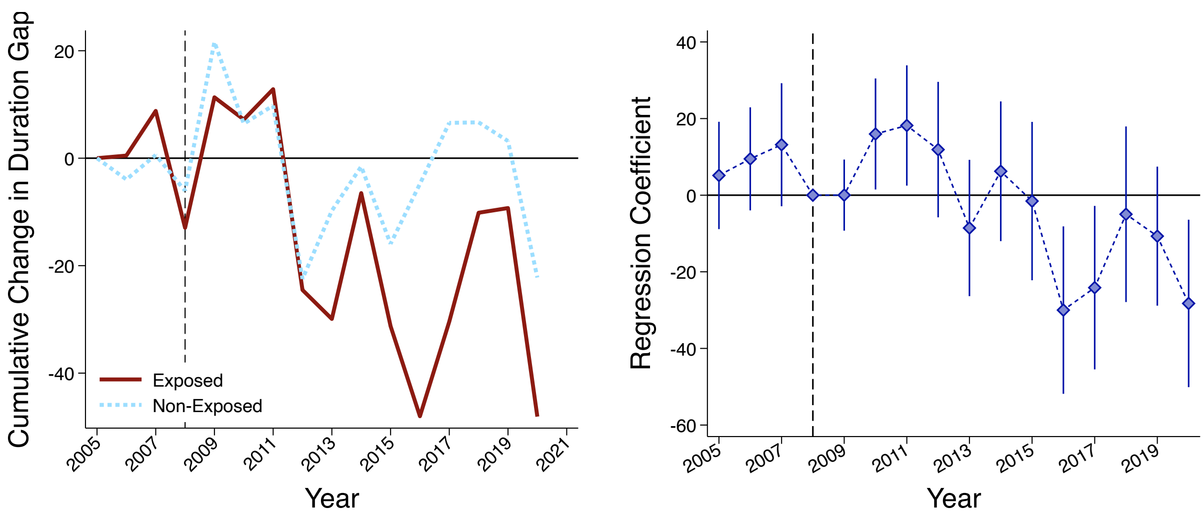
Note: This table reports regression results for equations (14) and (15). The dependent variable is the log of the premium quote for product i sold by insurer j in month t. $y_t^{(10)}$ is the monthly 10-year Treasury yield, MPU_t is the Husted et al. (2020) monetary policy uncertainty index normalized to a mean of 1, Exposed_j is an indicator equal to 1 if insurer j is in the exposed group, Asset_{jt} is the standardized total assets of insurer j in period t, and Long_i is an indicator equal to 1 if product i has the longer maturity of the two product categories in the regression. Standard errors clustered at the product-time level are reported in parentheses. Observations are weighted by insurer j's assets in the year corresponding to month t. * $p < 0.1$ ** $p < 0.05$ *** $p < 0.01$.

Table C.4: Interest Rates and Prices Regression — 30-year Policies

	Dependent Variable: $\log \text{Price}_{ijt}$					
(Long, Short) Category:	(30,20)	(30,15)	(30,10)	(30,20)	(30,15)	(30,10)
$y_t^{(10)} \times \text{Exposed}_j \times \text{Long}_i$	0.014***	−0.003	−0.009*			
	(0.004)	(0.005)	(0.005)			
$\text{MPU}_t \times \text{Exposed}_j \times \text{Long}_i$				−0.001	0.012**	0.019***
				(0.004)	(0.005)	(0.005)
Insurer × Month FE	✓	✓	✓	✓	✓	✓
Insurer × Product FE	✓	✓	✓	✓	✓	✓
Month × Product FE	✓	✓	✓	✓	✓	✓
Observations	8956	8956	8956	8956	8956	8956
Within-R^2	0.011	0.000	0.002	0.000	0.007	0.008

Note: This table reports regression results for equations (14) and (15) for 30-year term life policies. The dependent variable is the log of the premium quote for product i sold by insurer j in month t. $y_t^{(10)}$ is the monthly 10-year Treasury yield, MPU_t is the Husted et al. (2020) monetary policy uncertainty index normalized to a mean of 1, Exposed_j is an indicator equal to 1 if insurer j is in the exposed group, and Long_i is an indicator equal to 1 if product i has the longer maturity of the two product categories in the regression. Standard errors clustered at the product-time level are reported in parentheses. Observations are weighted by insurer j's ordinary life insurance in force in the year corresponding to month t. * $p < 0.1$ ** $p < 0.05$ *** $p < 0.01$.

Figure C.1: Changes in Duration Gaps by Exposure — Includes Outliers

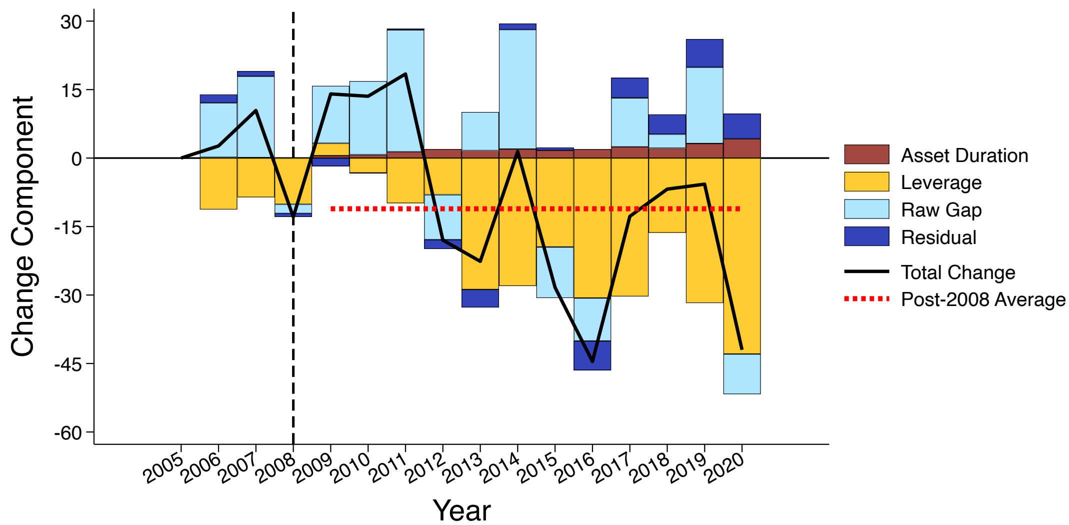


(a) Duration Gap Changes, Raw Data

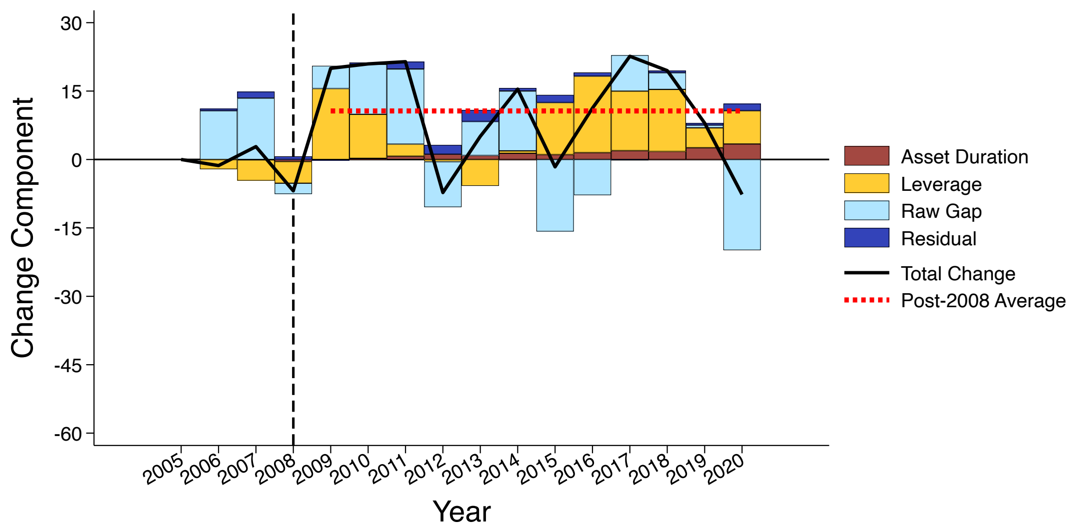
(b) Gap Differences, Regression Output

Note: This figure documents changes in the insurers' duration gaps [panel (a)] and estimates $\{\beta_\tau\}_{\tau=2005}^{2020}$ from regression (12) [panel (b)] over time. Duration gaps are constructed as in equation (11). Red lines reflect exposed insurers, blue dotted lines reflect non-exposed insurers, and gray dashed lines reflect the full sample. The figures report within-sample, asset-weighted averages. In panel (b), blue spikes represent 95% confidence intervals using standard errors clustered at the yearly level. We do not exclude outliers in these plots as we do in the main text.

Figure C.2: Change in Duration Gap Components



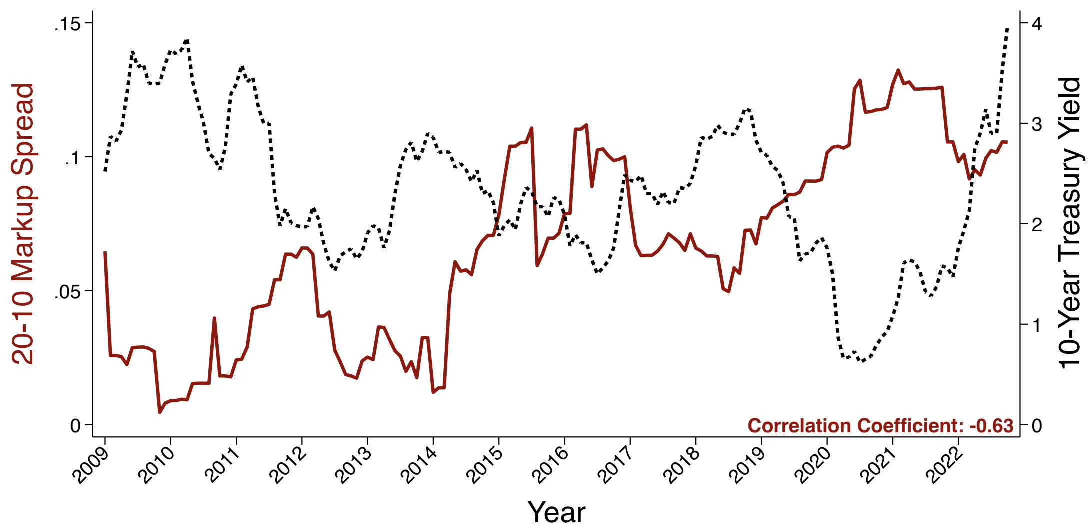
(a) Exposed Insurers



(b) Non-Exposed Insurers

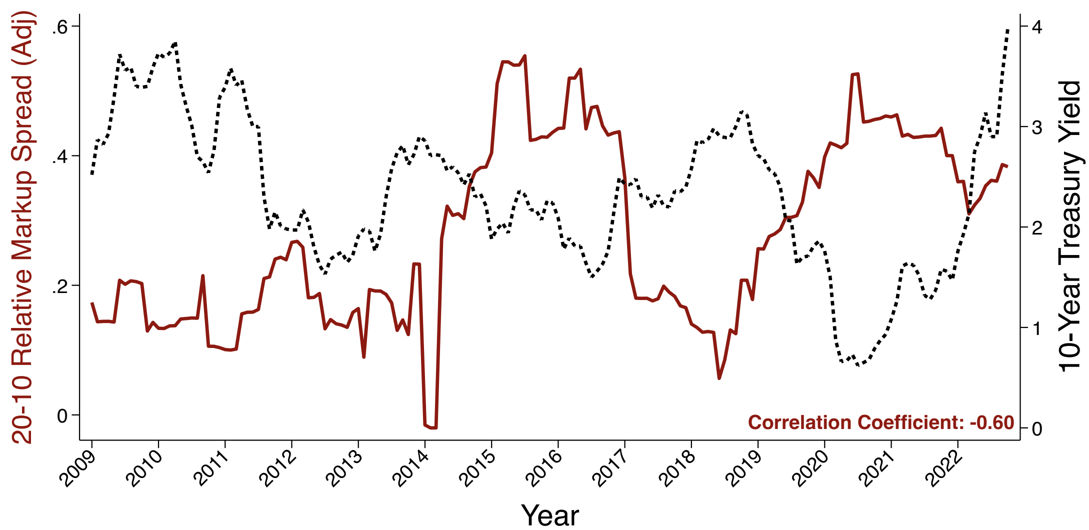
Note: This figure reports the decomposition of average duration gap changes as in equation (19) for exposed [panel (a)] and non-exposed [panel (b)] insurers. Red bars reflect changes in asset duration, light blue bars reflect changes in the duration mismatch component, yellow bars reflect changes in the leverage component, and dark blue bars reflect the residual component. Black lines represent the average change in duration gaps, while dotted gray lines reflect the average duration gap change relative to 2005 between 2008 and 2023. Averages are weighted by insurers' assets within exposed and non-exposed groups.

Figure C.3: The Interaction between Product Pricing and Interest Rate Risk



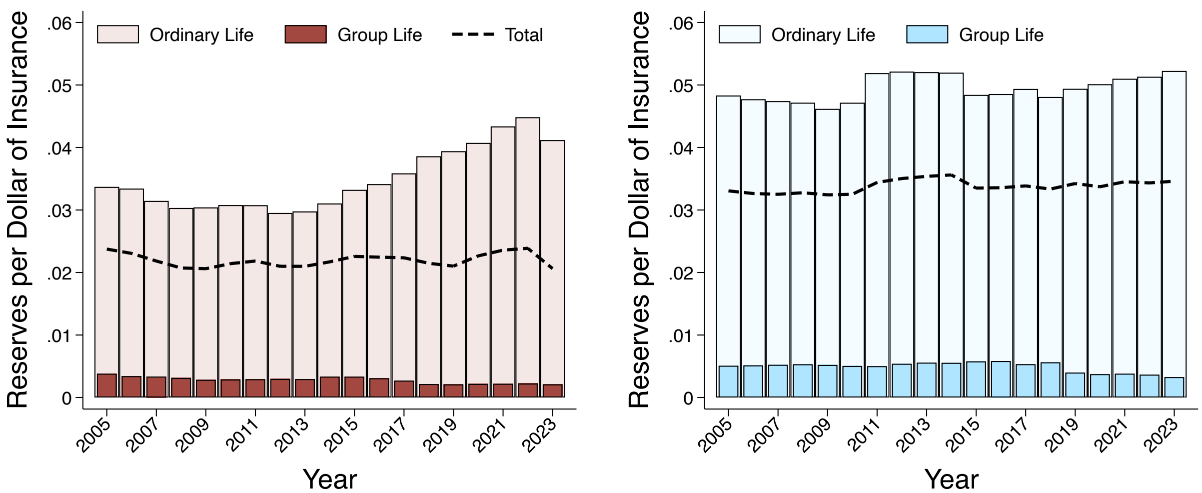
Note: This figure plots the (20,10) markup spread (red, left axis) and the 10-year Treasury yield (black dotted, right axis) for each month between January 2008 and December 2022. When calculating the markup spread, averages are weighted by ordinary life insurance in force.

Figure C.4: The Interaction between Product Pricing and Interest Rate Risk



Note: This figure plots the (20,10) markup spread (red, left axis) and the 10-year Treasury yield (black dotted, right axis) for each month between January 2008 and December 2022. When calculating the markup spread, averages are weighted by assets. We report the markup spread dividing by interest rate volatility, which we estimate each year using a 24-month rolling window.

Figure C.5: Reserve Value Across Products Over Time



(a) Exposed (b) Non-Exposed

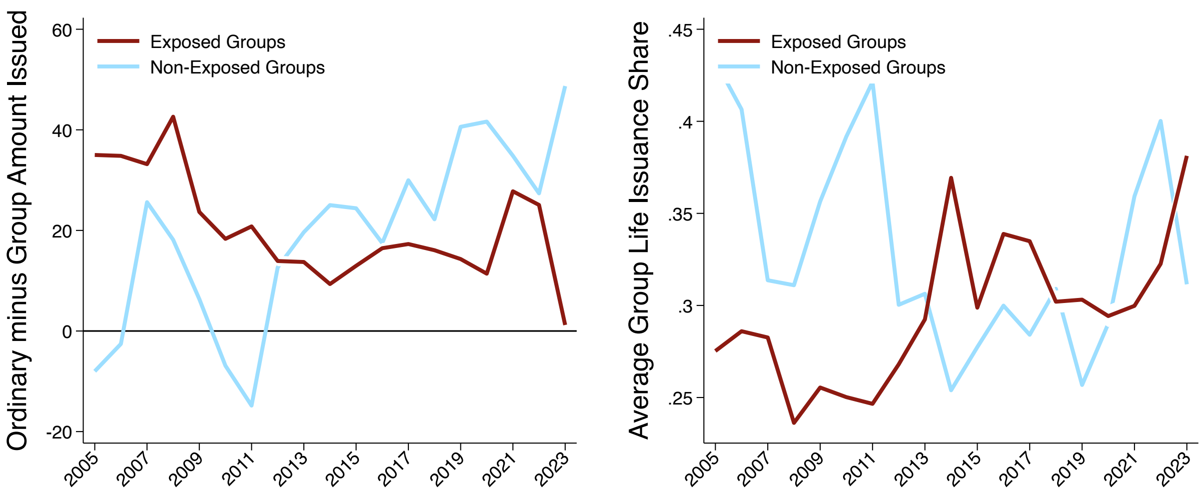
Note: This figure reports average reserve values for ordinary, group, and combined life insurance among exposed insurance groups. Reserve value is calculated as gross reserves divided by life insurance coverage in force. Panel (a) reports reserve values for exposed insurance groups, while panel (b) reports reserve values for non-exposed groups. Dark bars represent average group life reserve values, light bars represent average ordinary life reserve values, and the dashed black line represents the average of the total. Reserve values are weighted by life insurance in force within each category of insurance groups to avoid outliers.

FIGURE C.6: RELATIVE PRODUCT ISSUANCE — UNWEIGHTED

(a) Ordinary Life (b) Group Life

Note: This figure reports average ordinary life insurance issuance relative to average group life issuance [panel (a)] and average group life issuance shares [panel (b)] for exposed (red) and non-exposed (blue) insurers over time. Averages are unweighted. For panel (a), units are in billions of US dollars.

Figure C.7: Relative Product Issuance — Excludes Metlife

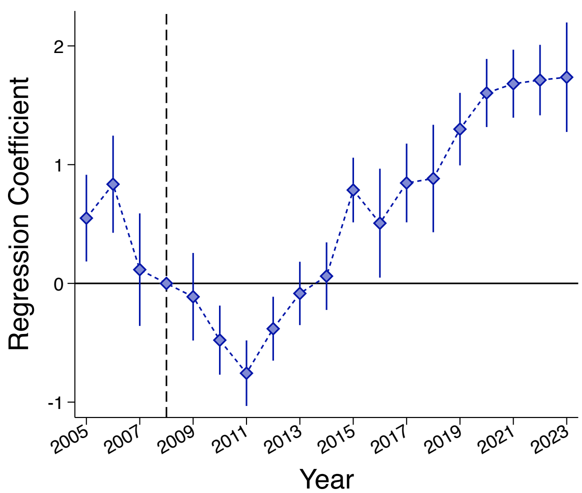


(a) Ordinary Life

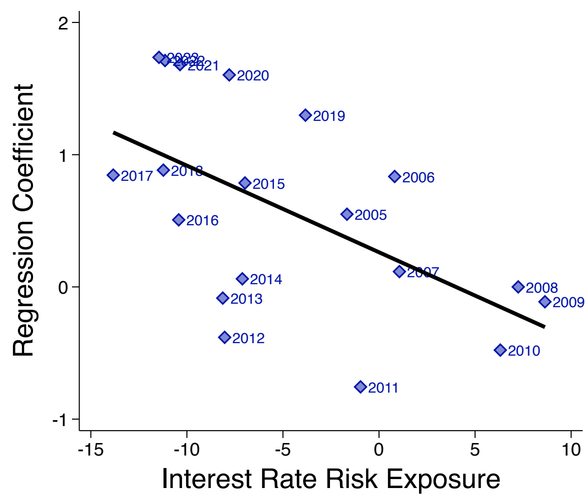
(b) Group Life

Note: This figure reports average ordinary life insurance issuance relative to average group life issuance [panel (a)] and average group life issuance shares [panel (b)] for exposed (red) and non-exposed (blue) insurers over time. Averages are weighted by assets within each class of insurers. For panel (a), units are in billions of US dollars. Metlife is excluded from calculations.

Figure C.8: Liability Rebalancing — Number of Policies



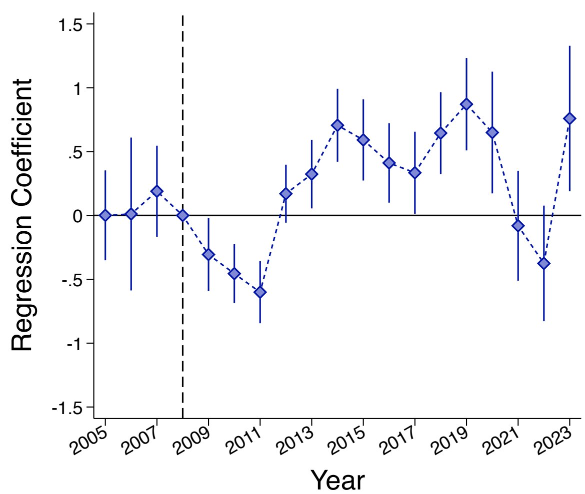
(a) Regression results over time



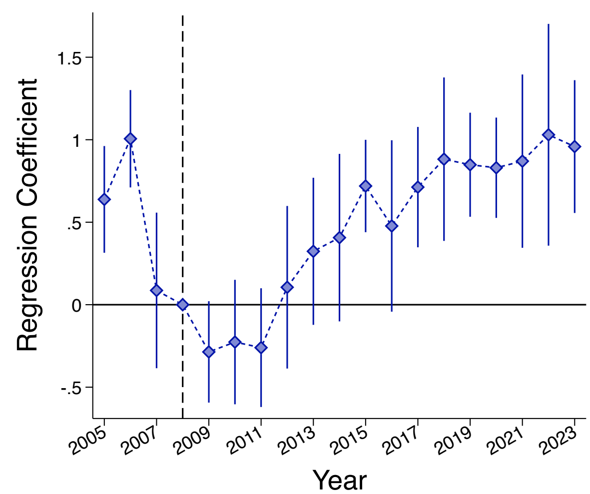
(b) Regression results vs Risk Exposures

Note: This figure reports regression results for equation (17). The regression estimates a set of coefficients $\{\beta_\tau\}_{\tau=2005}^{2023}$, each of which represents the relative number of insurance policies issued (i) between exposed and non-exposed insurers and (ii) between group and ordinary life insurance products. Panel (a) plots the estimated coefficients across time, while panel (b) reports the estimated coefficients against the yearly average interest rate risk exposure taken from the estimates in Figure 3. In panel (a), blue spikes represent 95% confidence intervals using standard errors clustered at the product-time level.

Figure C.9: Liability Rebalancing — Excludes Metlife



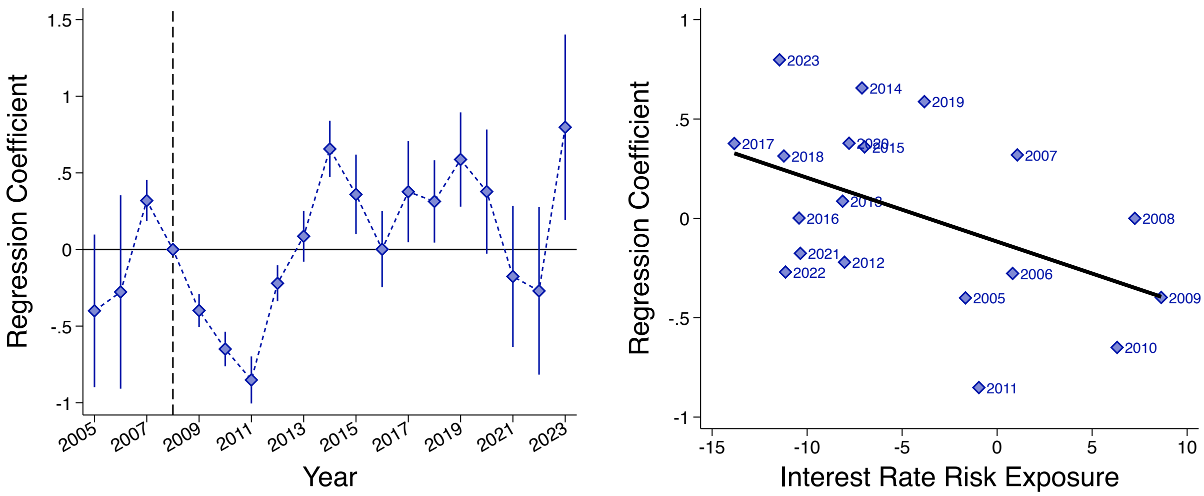
(a) Insurance Coverage



(b) Number of Policies

Note: This figure reports regression results for equation (17). The regression estimates a set of coefficients $\{\beta_\tau\}_{\tau=2005}^{2023}$, each of which represents the relative insurance coverage [panel (a)] or number of insurance policies [panel (b)] issued (i) between exposed and non-exposed insurers and (ii) between group and ordinary life insurance products. blue spikes represent 95% confidence intervals using standard errors clustered at the product-time level. These figures exclude Metlife from the regressions.

FIGURE C.10: LIABILITY REBALANCING — SIZE CONTROL

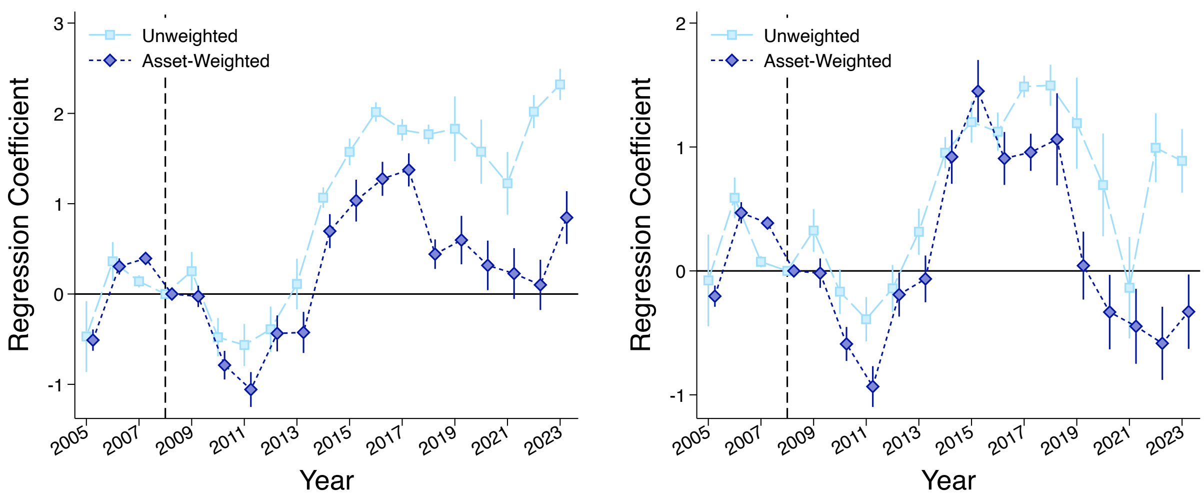


(a) Regression results over time

(b) Regression results vs Risk Exposures

Note: This figure reports regression results for equation (17), while controlling for the additional variables $\mathbf{1}\{\tau = t\} \times \text{Assets}_{jt} \times \text{Group}_i, t \in \{2005, 2006, \dots, 2023\}$, where Assets_{jt} is the (standardized) total assets of insurer j in period t. The regression estimates a set of coefficients $\{\beta_\tau\}_{\tau=2005}^{2023}$, each of which represents the relative insurance coverage issued (i) between exposed and non-exposed insurers and (ii) between group and ordinary life insurance products. Panel (a) plots the estimated coefficients across time, while panel (b) reports the estimated coefficients against the yearly average interest rate risk exposure taken from the estimates in Figure 3. In panel (a), blue spikes represent 95% confidence intervals using standard errors clustered at the product-time level.

FIGURE C.11: LIABILITY REBALANCING — OLS AND LOGS

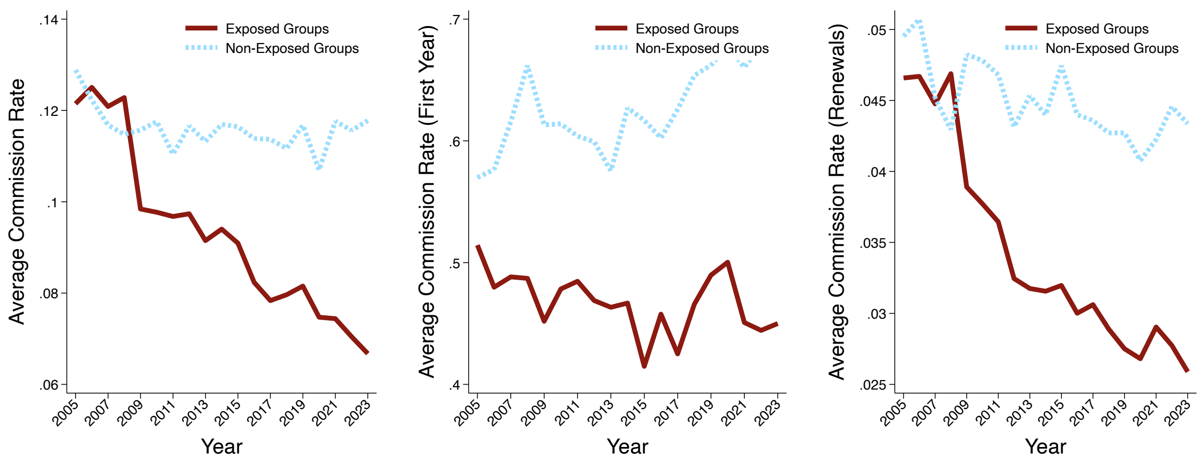


(a) Insurance Coverage

(b) Number of Policies

Note: This figure reports OLS regression results for equation (17) with $\log \text{Issuance}_{ijt}$ as the dependent variable. The regression estimates a set of coefficients $\{\beta_\tau\}_{\tau=2005}^{2023}$, each of which represents the relative insurance coverage [panel (a)] or number of insurance policies [panel (b)] issued (i) between exposed and non-exposed insurers and (ii) between group and ordinary life insurance products. Spikes represent 95% confidence intervals using standard errors clustered at the product-time level. Dark blue diamonds represent regression results that are weighted by insurer j's assets in year t, while light blue squares represent unweighted regression results.

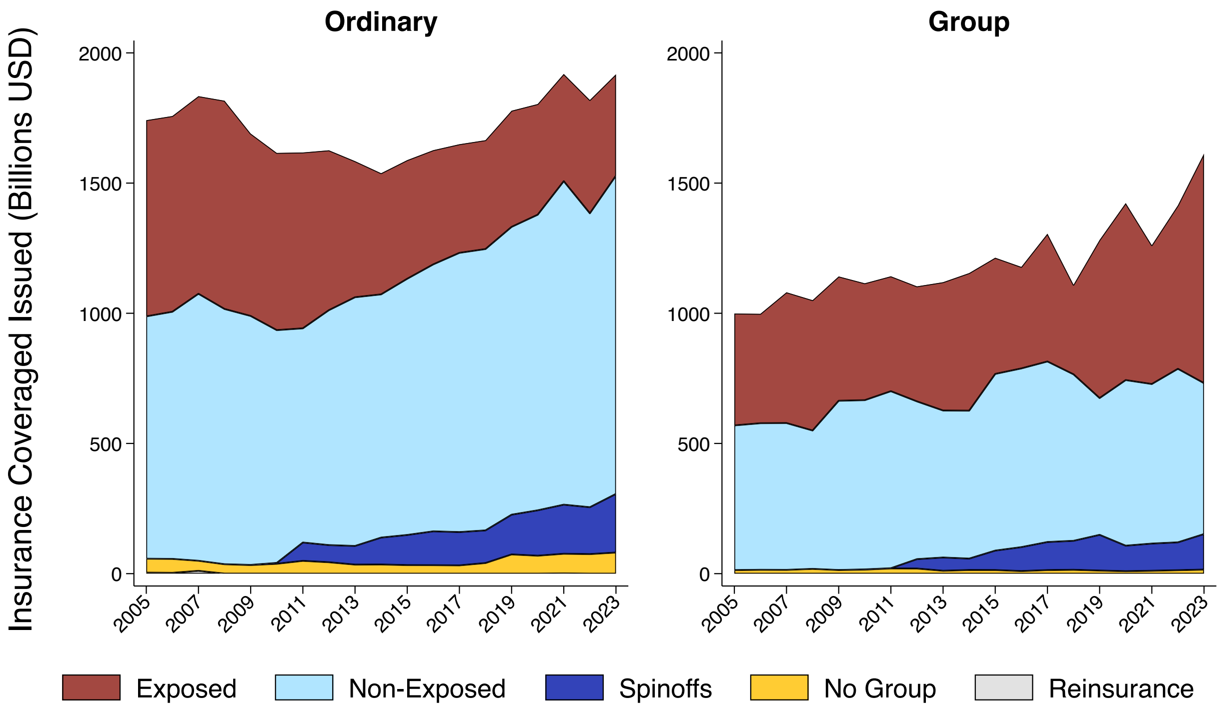
Figure C.12: Average Commission Rates



(a) Total (b) New Policies (c) Renewals

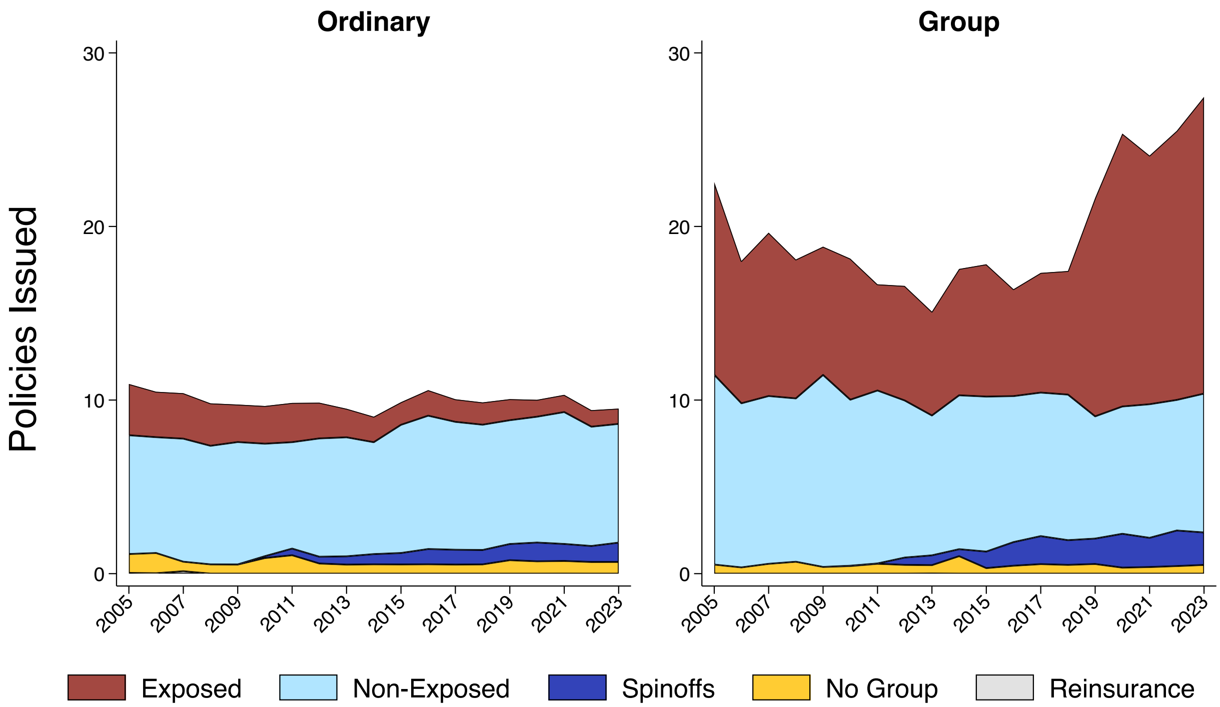
Note: This figure reports average commission rates for exposed (red) and non-exposed (blue) insurance groups from 2005 to 2023. Panel (a) reports total commission rates, panel (b) reports commissions on policies issued in the current year, and panel (c) reports commissions on policy renewals. Commission rates are calculated as direct commissions paid to agents divided by direct premium revenues. The data are winsorized at the 1% and 99% level to avoid outliers.

Figure C.13: Aggregate Nominal Issuance by Product Category



Note: This figure reports nominal aggregate life insurance issuance from 2005 to 2023. The first panel reflects ordinary life issuance, and the second panel reflects group life issuance. Red areas represent exposed insurance groups, light blue areas represent non-exposed insurance groups, dark blue areas reflect insurance companies that belonged to either the exposed or non-exposed insurance groups in the pre-crisis period but have since spun off, yellow areas represent insurers not in a life insurance group, and gray areas reflect reinsurance companies. The figures are reported in billions of US dollars.

Figure C.14: Aggregate Policy Issuance by Product Category



Note: This figure reports aggregate life insurance policy issuance from 2005 to 2023. The first panel reflects ordinary life issuance, and the second panel reflects group life issuance. Red areas represent exposed insurance groups, light blue areas represent non-exposed insurance groups, dark blue areas reflect insurance companies that belonged to either the exposed or non-exposed insurance groups in the pre-crisis period but have since spun off, yellow areas represent insurers not in a life insurance group, and gray areas reflect reinsurance companies. The figures are reported in millions of policies.