Financial Intermediaries and Demand for Duration

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Abstract

Stocks with long-term cash flows earn lower expected returns because they hedge fluctuations in investment opportunities. We study the role of financial institutions in shaping this duration premium using equity holdings of primary dealers, pension funds, banks, and insurance companies. We find that intermediaries' demand for equity duration varies systematically with their risk-bearing capacity. In the time series, institutions reduce their demand for long-duration claims and increase their exposure to reinvestment risk when aggregate capital ratios are low. Such a result extends cross-sectionally: better-capitalized and better-performing institutions tilt their portfolios more strongly toward long-duration stocks than their constrained peers. These patterns align with an ICAPM framework in which hedging demand declines with risk aversion. Counterfactual exercises show that shifts in intermediaries' preferences generate monotonic changes in expected returns across duration deciles, with especially large effects when demand shocks operate at the holding-company level.

Keywords: Institutional demand, Equity duration, Long-term investors, Capital constraints

JEL Classification: G10, G11, G20

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1 Introduction

The timing at which stocks pay out their cash flows has important implications for their pricing. Long-duration equities are valuable to long-term investors because they serve as natural hedges against changes in the investment opportunity set, due to their high sensitivity to discount rate variation (Gonçalves, 2021a). This hedging demand creates a "duration premium" – i.e., a differential in expected returns between U.S. short- and long-duration stocks – that varies with economic conditions and subsumes other dimensions of risk (Gonçalves, 2021b; Gormsen, 2021; Gormsen and Lazarus, 2023).

Due to the structure of their balance sheets, financial intermediaries tend to be associated with a preference for claims that pay off in the distant future. For example, institutions such as pension funds are commonly regarded as "long-term" investors given the maturity of their typical liabilities. The growing share of intermediated capital suggests that these institutions may play a key role in determining the equilibrium prices of long-term claims (Haddad and Muir, 2025a,b). However, their demand for long-term assets does not necessarily imply that they maintain net long-duration exposures in the equity market. Moreover, financial institutions differ substantially in their capital-absorbing capacity and regulatory constraints – both across and within types – which implies that their portfolio decisions may not align with a representative-agent framework. Therefore, it is not obvious whether, or to what extent, this class of investors collectively contributes to the level and dynamics of the duration premium.

In this paper, we take a granular look at the demand for equity duration from banks, primary dealers, insurance companies, and pension funds by studying their equity holdings. We show that intermediaries seek long-duration stocks during periods of high capital availability, while they concentrate their holdings in short-duration equities during periods of stress. These patterns are consistent with predictions from an ICAPM framework with reinvestment risk, where the hedging demand is decreasing in investors' risk aversion. Using counterfactual analysis, we show that shocks to intermediaries' risk-bearing capacity explain a significant portion of the duration premium. This allows us to quantify how changes in institutional asset allocation in response to shocks to their risk-bearing capacity impact market prices and shape the cross-section of equities.

We begin by considering the intertemporal CAPM model of Gonçalves (2021a), which expresses the optimal equity allocation of a long-term investor as the sum of two components: a myopic component and a component that hedges changes in the investment opportunity set, which we model as a linear function of a stock's equity duration. We then move beyond this representative-agent setting by allowing for heterogeneity in intermediaries' demand though the demand-system framework of Koijen and Yogo (2019). Consistent with the ICAPM, we incorporate an asset's beta and its duration as key characteristics explaining institutional demand, alongside a stock's market capitalization to pin down demand elasticities. We measure stock duration as the weighted average timing of cash payouts to investors, following the approach of Gonçalves (2021b), and estimate the demand system using quarterly 13F equity holdings of banks, applying the GMM instrumental variable approach of Koijen et al. (2024). Our sample includes 929 institutions from the Thomson Reuters database, collectively holding about one-third of total stock market capitalization over the 1980–2023 period. Among the banks, we separately identify primary dealers based on the New York Fed's list of trading counterparties in the implementation of monetary policy, prompted by research that highlights their role as marginal investors across multiple asset classes (He et al., 2017).

The demand-system estimates reveal that banks display the lowest duration demand unconditionally. Conversely, insurance companies exhibit the strongest average tilt in their portfolios toward high-duration stocks. The average demand coefficients of primary dealers and pension funds fall in between the other types. These average figures, however, mask a great deal of variation throughout the sample period as well as across institutions. Motivated by recent evidence in the intermediary asset pricing literature (Adrian et al., 2014; He et al., 2017; He and Krishnamurthy, 2018; Baron and Muir, 2021), we investigate whether such fluctuations in demand for duration are captured by institutional risk tolerance, as predicted by the ICAPM, which we measure by their type-specific equity capital ratio. In the time series, the capital ratios co-move strongly with aggregate proxies of risk aversion and risk appetite, such as the GZ spread of Gilchrist and Zakrajšek (2012) and Shiller's CAPE ratio. However, unlike these aggregate measures, the capital ratio is directly rooted in institutional balance sheet conditions and can be constructed at the type and institution level, thereby enabling cross-sectional tests.

We find that during times of low equity capital, institutions curtail their demand for long-term claims and, as a consequence, become more exposed to short-lived shocks. Importantly, intermediaries' capital constraints affect the demand for long-term cash flows even after controlling for proxies of reinvestment risk, which reassures us that we are not merely capturing portfolio rebalancing due to a predicted deterioration in economic conditions. As we elaborate below, the negative association between risk-bearing capacity and demand for long-duration stocks is also confirmed by cross-sectional analyses that exploit measures of capital constraints and regulatory frameworks specific to each institutional type.

With regard to preferences for betas, in times of scarce capital availability, institutions tilt toward higher market beta stocks. This pattern shares similarities with the mechanism in Frazzini and Pedersen (2014), where constrained institutional investors seek higher (implied) leverage by investing in high-beta stocks. Thus, our results imply that duration and betas capture two separate dimensions of portfolio choice for investors facing leverage constraints. Despite long-duration stocks having higher market betas, we document that when the balance sheet constraint is binding (i.e., equity capital is low), investors prefer stocks with high beta *and* low duration.

Our demand system framework allows us to quantify the economic impact of shifts in financial intermediaries' preferences on the duration premium. To this end, we compute counterfactual equilibrium prices across the cross section of stocks under hypothetical changes in demand for long-duration claims by institutional investors. Specifically, we simulate market outcomes had primary dealers, banks, insurance companies, and pension funds altered their preferences for stock duration (and possibly market beta) in response to shocks to their risk-bearing capacity.

To compute the resulting equilibrium prices, we begin by estimating the demand functions of all other investor types – including mutual funds, investment advisers, and households – using holdings data from Thomson Reuters. We then impose market clearing in each quarter to solve for the prices consistent with hypothetical shifts in demand from our intermediaries. We consider both a positive shift in duration demand, where we set institutional demand coefficients during periods of capital stress equal to their groups' historical averages, and a negative shift, which mirrors the observed sensitivity of preferences to equity capital, potentially combined with an increase in demand for market beta.

These counterfactual exercises yield several insights. The hypothetical shifts in intermediary demand have large and systematic effects on stock prices and expected returns across duration deciles. An increase in demand for long-duration equities leads to a steepening of the cross-sectional return curve, lowering prices and increasing expected returns for short-duration stocks relative to long-duration ones. The differential in expected returns rises by as much as 5% when periods of capital scarcity are identified with quarters in which the capital ratio of He et al. (2017) is below the bottom tercile of its historical distribution. Conversely, when demand for long-duration stocks experiences a negative shift coupled with a positive one for market beta, the overall effect over the full sample period is a decline in the duration premium of about 2.5%. This is about one-third of the 7.5% annualized spread between low- and high-duration stocks in our sample, an estimate in line with Gonçalves (2021b).

In the Thomson Reuters database, stock holdings are aggregated at the parent company level. The magnitudes reported above can thus be regarded as an upper bound on the effect of intermediaries' shifts in demand on the duration premium, under the assumption that shocks to risk-bearing capacity affect the entire holding company – either directly or indirectly through the impairment of internal capital markets following severe distress at a subsidiary, as discussed in He et al. (2017).

To obtain a more conservative benchmark of the effect, we leverage the granularity of the FactSet database, which retains information on the specific reporting unit in the 13F filings. This alternative reporting level implies that the same groups of institutions i) account for a smaller 9% of overall stock market capitalization, and ii) have potentially different demand elasticities and marginal role on price formation in the stock market. The effect of these differences on equilibrium prices is, therefore, ultimately an empirical question.

We repeat the counterfactual experiments using FactSet data and find that the resulting effect on the difference between high- and low-duration portfolios is about 0.5%, roughly one-fifth of the effect observed in Thomson Reuters. Given that assets under management for the same groups of institutions are about one-third of those in the Thomson Reuters database, this comparison suggests that shocks to the demand for duration at the holding-company level generate more than proportional redistribution effects across the equity market.

We provide several analyses that corroborate and substantiate our empirical approach. First, for the demand system, we verify that the latent demand (i.e., the portion of weights driven by non-observable characteristics) resulting from our proposed setup with duration is lower than that obtained from including an array of commonly used characteristics – book-to-market, profitability, investment, and payout. This finding, which resonates with evidence in Gormsen and Lazarus (2023) from standard cross-sectional asset pricing tests, supports the view that duration is a relevant and encompassing dimension of institutional risk and capital allocation. Moreover, we provide an out-of-sample validation of our capacity to capture the demand for long-term cash flows in the equity market by examining target-date funds (Parker et al., 2023). We find that target-date funds with longer time to maturity have substantially higher demand for duration than funds with a closer expiration date, and that this result is not driven by the gradual shifting from stocks to bonds as the target date approaches.

Second, to examine how intermediaries' demand relates to their risk-bearing capacity, we exploit the fact that much of the variation in portfolio choices – especially for equity duration – is institution-specific. We therefore analyze the cross-sectional determinants of equity duration demand separately for each intermediary type. Using the FactSet holdings database, which provides institution-level portfolio detail and offers a comprehensive coverage and data quality, we construct proxies for risk tolerance based on asset/liability structures, income sources, and regulatory constraints, enabling cleaner identification at the firm level.

For banks and primary dealers, tighter constraints – such as a lower equity-to-assets ratio or higher VaR-to-assets – reduce demand for duration while increasing demand for high-beta stocks, underscoring how capital availability shapes portfolio allocation. To strengthen causal inference, we exploit the introduction of the Supplementary Leverage Ratio (SLR) in 2014Q3, which required banks to hold Tier 1 capital against all assets, including off-balance-sheet exposures. We find that capital-constrained banks with below-median equity ratios responded by shifting toward short-duration, high-beta stocks – confirming that capital constraints induce opposite adjustments in duration and beta demand.

Among insurance companies, declining equity capital is again linked to reduced demand for duration. To isolate exogenous shocks, we focus on firms that undertook major re-capitalizations during the 2008–09 crisis (e.g., through public equity issuance or dividend suspension), signaling binding constraints. These insurers subsequently raised their demand for equity duration, suggesting that improved risk-bearing capacity drives a greater tilt toward long-term claims. This effect remains even after controlling for capital and loss ratios, in line with ICAPM predictions that hedging demand rises with risk tolerance. Unlike banks, however, insurers' beta demand is not strongly tied to capital ratios. Instead, capital-constrained insurers facing operating losses – our proxy for adverse financial shocks – shift toward lower-beta, safer portfolios, reflecting heightened aversion to risk.

Finally, corporate defined benefit (DB) pension plans respond to capital shocks – such as low funding ratios or large losses – by favoring low-duration, high-expected-return stocks and increasing beta exposure. Public DB funds, unlike private ones, can base discount rates on expected asset returns under GASB rules. Following Andonov et al. (2017), we use the share of retired plan members to capture funding pressure: a higher retired share increases the need to justify high discount rates. Consistent with this, we find that as the share of retirees rises, public DB funds shift toward riskier, lower-duration equities. This behavior reflects an incentive to maintain favorable funding status through elevated expected returns. Notably, this shift is more pronounced in duration than beta demand, suggesting the adjustment mainly targets the hedging dimension.

Related Literature: Our paper relates to three strands of the finance literature. First, we add to the literature on equity duration.[1,2] Dechow et al. (2004) use market equity values and expected future cash flows to infer a weighted average of discounted future payouts paid by the stock—i.e., a cash-flow-implied equity duration. Weber (2018) finds a negative relation between this duration measure and expected stock returns, suggesting a behavioral mispricing explanation. In contrast, Gonçalves (2021a,b) propose that long-duration stocks provide a hedge against reinvestment risk, prompting long-term investors to optimally hold them, thereby lowering their expected returns in equilibrium. This literature, however, is silent on the identity of these long-term investors in the equity market. For instance, while pension funds are natural candidates, their balance sheets may be only *net* long-term, holding very high-duration bonds while maintaining short-term equity exposures. We specifically examine the demand for long-duration equity by financial intermediaries and document a novel link between time-varying exposure to reinvestment risk and risk-bearing capacity. Li and Xu (2024) examine how the term structure of equity yields, measured from dividend futures, relates to an intermediary-based SDF in the time series. By contrast, we focus on

[1] The concept of equity duration is equivalent to the traditional Macaulay duration measure for bonds: $D_t = \sum_{h=1}^{N} \frac{CF_{t+h}}{(1+r)^h} h$. The key difference is that the deterministic future cash flow CF_{t+h} is stochastic in the equity market. Thus, we must rely on an econometric model for expected future cash flows to infer equity duration.

[2] In this paper, we focus on the equity term structure in the cross-section of individual stocks. A complementary literature—initiated by Binsbergen et al., 2012, 2013—studies the aggregate market equity term structure using derivatives on dividend claims (see van Binsbergen and Koijen, 2017 for a review). More recently, Giglio et al. (2024) construct aggregate and portfolio dividend claims from a well-specified asset pricing model.

the cross section of individual stocks and complement their evidence by showing that the price of reinvestment risk varies systematically with intermediary capital.

Second, we contribute to the literature on long-term investing and the role of intermediaries in asset markets. Cochrane (2022) discusses portfolio theory for long-term investors and its empirical challenges. To advance understanding of the principles driving portfolio practice, he advocates focusing on the stream of payoffs an investment can generate, rather than on one-period returns, and considering portfolios from a general-equilibrium perspective. Our study takes a step in this direction by examining the demand for duration – which accounts for the entire stream of expected payoffs – using actual institutional investor portfolios through the lens of a demand-system framework grounded in intertemporal optimization. With respect to the growing literature on intermediary asset pricing (He and Krishnamurthy, 2013; Brunnermeier and Sannikov, 2014), our findings are consistent with studies linking institutional balance sheets to the cross-section of bonds, equities, and highly intermediated asset classes (Adrian et al., 2014; He et al., 2017; Haddad and Muir, 2021). Our paper also contributes to the market macrostructure agenda by identifying the key intermediaries trading long-duration stocks, assessing how their behavior shapes the cross-section of equities in equilibrium, and quantifying how shocks to their financial health impact equity demand (Haddad and Muir, 2025b).

Finally, we relate to studies employing a demand-system asset pricing framework (Koijen and Yogo, 2019). This literature seeks to match asset prices with investors' holdings.[3] Recently, Koijen et al. (2024) emphasize the heterogeneity in investors' demand functions. We confirm that such heterogeneity extends to institutional demand for duration, while also identifying capital availability as a common driver of time variation in this demand. Our estimation further shows that incorporating equity duration into the system streamlines the framework and discussion without sacrificing relevant information compared to major characteristics.

[3] Tobin (1958, 1969); Brainard and Tobin (1968) are early contributions to this line of research.

The rest of the paper is organized as follows. Section 2 describes our duration-based demand system, discusses the estimates and validates them with several tests. Section 3 investigates the role of capital availability in driving variation in intermediaries' preference for duration. Section 4 studies the effect of a shift in intermediaries' preferences for duration on asset prices. Section 5 presents the cross-sectional analysis linking intermediaries' balance sheets and their demand for duration. Section 6 concludes.

2 Duration and Demand-based Asset Pricing

In Section 2.1, we outline the theoretical framework linking portfolio allocation decisions and equity duration. We use this framework to explain the equity holdings of U.S. banks, primary dealers, insurance companies, and pensions funds from 13(f) mandatory filings, which we describe in Section 2.2. Section 2.3 describes our estimation strategy, while Section 2.4 presents the resulting demand-based coefficients. Finally, Section 2.5 provides a battery of tests that validate our approach.

2.1 A Demand System with Duration

Our starting point is the Intertemporal Capital Asset Pricing Model (ICAPM) of Gonçalves (2021a), who extends the model in Campbell (1993) to account for reinvestment risk. In Gonçalves (2021a), a long-term (i.e., infinitely lived) investor with Epstein-Zin recursive preferences maximizes her long-term wealth facing trade-offs between current wealth and future expected wealth growth. The investor demands long-duration stocks in her portfolio to hedge for "reinvestment risk" – i.e., unfavorable shifts in future investment opportunities, such as an increase in discount rates – which, in turn, lowers current-period risk premia to hold long-term claims.[4] We review the main implications of Gonçalves's (2021a) model for

[4]In a similar spirit, although less formally, Dechow et al. (2004) argue that investors may prefer to hold long duration equities because they have long investment horizons and wish to immunize themselves from reinvestment risk arising from expected return shocks.

the optimal allocation as they serve the basis for our empirical investigation of intermediaries' equity demand.

Formally, the investor chooses consumption C_t and portfolio allocation w_t to maximize her lifetime utility, subject to the budget constraint $W_{t+1} = (W_t - C_t) \cdot R_{w,t+1}$, where $R_{w,t}$ represents the investor's wealth portfolio. Following most of the previous literature, we assume that the marginal investor is fully invested in equities, so that the wealth portfolio is equal to the equity market portfolio. Furthermore, we allow the investor's risk-aversion coefficient γ_t to potentially vary over time, which is key to capture the observed time-variation in the equity term structure. The log stochastic discount factor m_{t+1} implied by the ICAPM optimality conditions can then be written as:[5]

$$m_{t+1} = a_t - \gamma_t \cdot r_{w,t+1} - (\gamma_t - 1) \cdot vw_{t+1}$$

where $r_{w,t} = \log(R_{w,t})$ are log returns on the investor's wealth portfolio, $vw_t = \log(V_t/W_t)$ is the log value-wealth ratio (i.e., the ratio of the value function to current wealth), and a_t is a function of the preference parameters.

Under the assumption that m_t and log returns to a given asset n are conditionally jointly normal, the ICAPM Euler condition implies that expected returns in excess of the risk-free asset (denoted by r_f) satisfies the following relationship:

$$E_t\left[r^e_{n,t+1}\right] + 0.5\left(\mathrm{Var}_t\left(r_{n,t+1}\right) - \mathrm{Var}_t\left(r_{f,t+1}\right)\right) = \gamma_t \mathrm{Cov}_t\left(r_{w,t+1}, r^e_{n,t+1}\right) + (\gamma_t - 1)\,\mathrm{Cov}_t\left(vw_{t+1}, r^e_{n,t+1}\right).$$

Staking the $n = 1, 2, \ldots, N$ asset-by-asset equations and applying the Campbell et al. (2003) portfolio approximation, we obtain an expression for the vector of expected excess returns

[5] See Gonçalves (Internet Appendix, Section I, 2021a). In particular, the derivation uses the log-linear approximation to the investor's consumption-wealth ratio in Campbell (1993).

(up to the Jensen's term):[6]

$$E_t\left[r^e_{t+1}\right] + \frac{1}{2}\mathrm{diag}\left(\Sigma_{r^e,t}\right) = \gamma_t \Sigma_{r^e,t}\omega^*_t + \left(\gamma_t - 1\right)\mathrm{Cov}_t\left(vw_{t+1}, r^e_{t+1}\right) \ .$$

This equation reveals that equilibrium expected excess returns have a two-factor structure. Long-term investors require compensation for holding stocks whose returns are positively correlated with the contemporaneous returns to the wealth portfolio – i.e., stocks with high exposure to market risk. All else equal, expected excess returns are also higher for stocks with a positive covariance with the log value-wealth ratios – i.e., stocks that bear higher reinvestment risk. Long-term claims have typically higher market betas, but their higher sensitivity to variation in discount rates makes them a good hedge for a deterioration in investment opportunities. For claims that pay off in the far distant future, the reinvestment risk channel dominates and expected excess returns to high duration equities are lower than those of short-term claims.

We can solve for the resulting investor's optimal equity allocation ω^*_t:

$$\omega^*_t = \underbrace{\frac{1}{\gamma_t}\Sigma^{-1}_{r^e,t}\left(E_t\left[r^e_{t+1}\right] + 0.5\cdot\mathrm{diag}\left(\Sigma_{r^e,t}\right)\right)}_{\omega^{\text{myopic}}_t} + \underbrace{\frac{1}{\gamma_t}\left(1-\gamma_t\right)\Sigma^{-1}_{r^e,t}\mathrm{Cov}_t\left(vw_{t+1}, r^e_{t+1}\right)}_{\omega^{\text{hedging}}_t}, \qquad (1)$$

Analogous to risk premia, the optimal demand is driven by two components. The first component, ω^{myopic}_t, corresponds to the standard static CAPM and depends on the covariance between a stock's return and the wealth portfolio. The second component, $\omega^{\text{hedging}}_t$, depends on the coefficient from a projection of vw_{t+1} on r^e_{t+1}, which we denote $\phi_t\left(vw_{t+1}, r^e_{t+1}\right) \equiv \Sigma^{-1}_{r^e,t}\mathrm{Cov}_t\left(vw_{t+1}, r^e_{t+1}\right)$. This component captures the portion of long-term investors' equity demand driven by hedging motives.

To make further progress, we follow Gonçalves (Section B.1, 2021a) and model the log

[6]Recall that $R_{w,t+1} = \omega_t^\intercal R_{t+1} + \left(1 - \sum_n \omega_{n,t}\right) R_{f,t} = \omega_t^\intercal R^e_{t+1} + R_{f,t}$. For log returns, portfolio aggregation is not additive, as $r_{w,t} = \log\left(R_{w,t}\right) = \log\left(\omega_t^\intercal R_{t+1}\right) \neq \omega_t^\intercal r_{t+1}$. Campbell et al. (2003) approximate excess returns to the wealth portfolio as $r_{w,t+1} - r_{f,t} \approx \omega_t^\intercal r^e_{t+1} + \frac{1}{2}\omega_t^\intercal \mathrm{diag}\left(\Sigma_{r^e,t}\right) - 0.5\omega_t^\intercal \Sigma_{r^e,t}\omega_t$, which implies that $r_{w,t+1} - r_{f,t} \approx \omega_t^\intercal r^e_{t+1}$ plus a time-t term.

value-wealth ratio as a function of a set of state variables, s_t and risk aversion:

$$vw_{t+1} = a_{v,t} + \frac{1}{\gamma_t - 1} \cdot s_{t+1}. \tag{2}$$

For the optimal weight on asset n, this assumption allows us to write

$$\omega_t^*(n) = \omega_t^{\text{myopic}}(n) - \frac{1}{\gamma_t}\phi_t\left(s_{t+1}, r_{n,t+1}^e\right), \tag{3}$$

which reveals that the hedging component of demand for stock n depends on its covariance with the variables that track changes in the investment opportunity set.[7] Stocks that pay off more in the far future provide hedging benefits as an increase in today's expected returns is associated with a drop in today's valuation and a contemporaneous increase in the continuation value through higher future returns. This argument suggests to treat the projection term in the hedging demand as a linear function of a stock's equity duration, or

$$\phi_t\left(s_{t+1}, r_{t+1}^e\right) = -b\, dur_t(n)$$

with $b > 0$. If we plug this expression in Eq. (2.1), we finally obtain:

$$\omega_t^*(n) = \omega_t^{\text{myopic}}(n) + \frac{b}{\gamma_t} dur_t(n). \tag{4}$$

In equilibrium, long-term investors demand more strongly stocks with higher duration as they negatively correlate with reinvestment risk, which in turn implies that these stocks trade at relatively low risk premia conditional on their market risk.

The ICAPM framework has been developed and tested by Gonçalves (2021a) in a representative agent setting to explain the downward-sloping unconditional relation between

[7] As emphasized by Gonçalves (2021a), the formulation also has appealing implications for the countercyclicality of the equity term structure as it implies that only the price for market risk varies over time. Thus, when risk aversion is high, the market risk channel dominates and the term structure flattens or becomes upward sloping.

duration and expected returns (Gonçalves, 2021b), as well as its steepening during crisis times (Gormsen, 2021). Given their balance sheet structure, it is natural to regard financial intermediaries – such as banks, insurance companies, and pension funds – as long-term investors that are concerned about the effects of changes in the investment opportunity set and actively manage duration exposures. However, financial institutions differ significantly in their asset and liability structure and regulatory framework, which suggests that their demand for duration might be heterogeneous both across and, potentially, within different institutional types.

To capture differences in beliefs and portfolio decisions, we rely on the demand system approach of Koijen and Yogo (2019) and its extension by Koijen et al. (2024). In their framework, under certain parameter restrictions, an institution's optimal portfolio weight on a given stock n at time t relative to outside assets, $\frac{w_{i,t}(n)}{w_{i,t}(0)}$, takes the form a characteristics-based demand. This is the product of an exponentially affine function of the stock's characteristics times an institution's latent demand $\epsilon_{i,t}(n)$, which captures the investor demand for unobserved (by the econometrician) characteristics of the asset and can be regarded to as a measure of sentiment.

The optimal allocation from our Intertemporal CAPM in Eq. (2.1) suggests the following parsimonious setup:

$$\frac{w_{i,t}(n)}{w_{i,t}(0)} = \exp\left\{\theta_{mkt,i,t}\ \beta_{mkt,t}(n) + \theta_{dur,i,t}\ dur_t(n) + \theta_{me,i,t}\,\mathrm{me}_t(n) + \beta_{K,i,t}\right\}\epsilon_{i,t}(n), \qquad (5)$$

where me is the stock's log market equity, market beta β_{mkt} measures exposure to contemporaneous shocks to the aggregate portfolio, and dur is log duration. The investor- and time-specific coefficients θs capture the portfolio tilt toward a given characteristic, holding all others fixed. The first characteristic in the system captures demand for market beta (β_{mkt}), which accounts for the myopic component. The second term, our primary focus, captures demand for equity duration or long-term dividend claims. When long-term investors have a positive positive $\theta_{dur,i,t}$, it indicates that they tilt their portfolio toward long-duration stocks to capture their reinvestment risk hedging properties. Finally, we include in the system the stock's market capitalization to pin down price elasticity, and time-investor fixed effects $\beta_{K,i,t}$.

We use the demand-system framework to study heterogeneity in equity demand for market and reinvestment risk through institutional holdings data, which we describe next.

2.2 Holdings Data and Summary Statistics

We characterize the portfolio decisions of institutional investors using data from 13(f) mandatory filings (hereafter 13F). Our analysis centers on four institutional types: primary dealers, banks, insurance companies, and pension funds. We treat primary dealers separately from banks because recent research emphasizes their role as marginal investors in asset markets (Adrian et al., 2014; He et al., 2017). For banks, insurers, and pension funds, the long-term nature of their liabilities motivate substantial equity exposure and active duration management of their assets to meet their financial obligations. For example, defined-benefit pension plans allocate roughly half of their assets to corporate equities as of 2015 (Klingler and Sundaresan, 2019). We exclude delegated investors – such as hedge funds, mutual funds, and other investment advisers – from our main analysis because their liability structure and the nature of the constraints they face are quite different (He et al., 2017). We estimate their demand functions in Section 4 to compute counterfactual equilibrium prices.

We use two data sources for quarterly filings. The first source is Thomson Reuters, whose Institutional Holdings Database (s34) has been widely utilized in previous studies, including Koijen and Yogo (2019). The sample period spans from 1980Q1 to 2023Q4. For the shorter period starting in 1999Q1, we also rely on holdings data compiled by FactSet. Appendix A.1 provides details on how we group institutions from the two databases into institutional types.

Given its longer time span, we treat the Thomson Reuters database as the benchmark sample for our time-series and counterfactual analyses. However, we use FactSet data for our cross-sectional tests linking institutional demand to financial constraints, as FactSet offers several advantages. First, FactSet retains the identity of the actual reporting institution, unlike Thomson Reuters, which aggregates positions at the holding company level.[8] Since measures of financial constraints, such as the underfunding ratio for pension funds, tend to be institution-specific, the granularity of FactSet data enables us to trace an institution's stock holdings and its financial and income status with greater precision.

Second, research by Ben-David et al. (2021) has documented reporting gaps in aggregate holdings within the Thomson Reuters data, which are absent in FactSet. Third, FactSet provides a unified identifier across its various sections, maximizing sample availability for measuring the financial constraints of reporting institutions and their holding companies. Finally, since FactSet data are organized at the security level, they allow us to include information on exchange-traded fund (ETF) holdings, which are reported in 13F filings but cannot be mapped to a CUSIP. This information helps us rule out concerns that intermediaries achieve a target duration through ETF positions rather than direct stock holdings. Following Ben-David et al. (2018), we recognize ETFs in CRSP by ticker and compute ETF stock weights using Thomson Reuters s12 quarterly holdings. We identify 436 unique ETFs held by our sample of institutional investors.

We merge the institutional filings with equity data and fundamentals from the CRSP-Compustat database. Our investment universe consists of ordinary common shares (share codes 10 and 11) trading on the NYSE, Amex, and Nasdaq, with the addition of ETFs for the FactSet sample. The outside asset w_0 consists of securities other than common stocks (share codes 12 and 18) and stocks for which we cannot compute duration due to missing data in their characteristics.

Table 1 provides relevant summary statistics of our working sample, which results from

[8]For example, in 2006:4, Thomson Reuters contains a single entry for "J.P. Morgan Chase & Co." classified as a Bank. In contrast, FactSet reports separate holdings for entities such as "J.P. Morgan Investment Management" (classified as Mutual Fund), "J.P. Morgan Securities" (Bank), "J.P. Morgan Trust Bank" (Bank), "J.P. Morgan Asset Management" (Investment Adviser), and "J.P. Morgan Chase Bank, North America" (Investment Adviser).

merging the holdings and equity data. Panel A refers to the Thomson Reuters database over the full 1980Q1–2023Q4 period. Overall, we analyze 929 distinct institutions that hold, on average, about 36% of the public equity market. Banks account for more than half of total observations, followed by insurance companies and pension funds, which together represent one-third. Primary dealers are the least represented group with 63 institutions. However, they account for a sizeable 10% of the total market capitalization on their own, which represents nearly 25% of our sample in terms of assets under management.[9] Furthermore, their portfolio decisions appear rather peculiar, as the median number of stocks they hold, at 1,500, is nearly an order of magnitude larger than that of all other institutions, which is comparable at 227 to that in Koijen and Yogo (2019).

Panel B of the table reports statistics for Thomson Reuters over the 1999Q1–2023Q4 period to facilitate comparison with FactSet. The number of reporting institutions decreases to 509, collectively accounting for nearly 30% of overall market capitalization.

Panel C summarizes statistics for the FactSet database. The number of distinct institutions and the median AUM are higher than for Thomson Reuters, with the largest increases originating from pension funds and banks. However, in FactSet, these four types of institutions account for less than 9% of overall market capitalization.[10] This is because FactSet, as mentioned above, does not aggregate holdings at the holding company level, ensuring that the AUM reflects only the assets managed by specific reporting entities rather than those of the broader parent organization. The number of stocks per institutional type is relatively comparable across the two sources, with primary dealers once again exhibiting a disproportionately broader set of holdings compared to other institutional types.

Panel A of Table 2 summarizes the portfolio weights from the Thomson Reuters database. The average stock weight in institutional portfolios is highest at 0.19% for pension funds and

[9] In Koijen and Yogo (2019), primary dealers are classified as banks and, less frequently, as investment advisors or mutual funds.

[10] For FactSet, the AUM incorporates indirect stock holdings through ETFs. The average ETF portfolio weight, conditional on holding at least one ETF, is about 2%. Without ETFs, these institutions hold on average about 8.50% of the market.

Panel A: Thomson Reuters, 1980-2023						
			AUM ($MM)		N. stocks	
	N. institutions	% market	median	p90	median	p90
Primary Dealers	63	10.13	14,169	133,220	1,503	2,722
Banks	594	15.50	412	9,176	198	1,051
Insurances	158	5.13	816	16,511	157	1,175
Pension Funds	114	5.51	3,023	24,281	438	1,468
Total	929	36.26	711	21,990	227	1,541

Panel B: Thomson Reuters, 1999-2023						
			AUM ($MM)		N. stocks	
	N. institutions	% market	median	p90	median	p90
Primary Dealers	34	9.44	22,377	183,901	2,167	3,748
Banks	292	12.53	422	13,094	206	1,393
Insurances	97	3.64	1,016	26,257	189	1,660
Pension Funds	86	3.95	4,612	34,765	605	1,974
Total	509	29.56	855	34,034	258	2,157

Panel C: FactSet, 1999-2023						
			AUM ($MM)		N. stocks	
	N. institutions	% market	median	p90	median	p90
Primary Dealers	34	1.44	10,994	60,605	1,951	3,464
Banks	359	4.00	704	20,307	336	2,008
Insurances	95	0.69	1,453	9,680	160	1,071
Pension Funds	116	2.80	3,368	39,814	547	1,919
Total	604	8.92	1,351	25,493	359	2,226

TABLE 1: **Summary statistics of portfolio holdings data**: The table presents summary statistics of quarterly portfolio holdings data from Thomson Reuters for the full 1980Q1 to 2023Q4 sample period (Panel A), from Thomson Reuters for the 1999Q1-2023Q4 subsample (Panel B), and from FactSet, which is available only from 1999Q1 to 2023Q4 (Panel C). For each institutional type, we report: the number of distinct reporting institutions, the average share of total public equity market held, and the median and 90th percentile of assets under management (AUM, in millions of dollars) and the number of stocks held by a given institution.

lowest at 0.05% for primary dealers. The average weight of the outside asset ranges between 27% and 30%. This implies that the resulting demand estimates are directly comparable across groups, as they are scaled by about the same amount. FactSet portfolio weights display similar properties; see Appendix Table A.1.

	$w_i(n)$					w_0	
	mean	std	p10	p50	p90	mean	std
Primary Dealers	0.05	0.34	0.00	0.00	0.08	30.54	8.79
Banks	0.18	1.02	0.00	0.00	0.39	26.96	10.11
Insurances	0.19	1.21	0.00	0.01	0.39	27.51	10.16
Pension Funds	0.12	0.76	0.00	0.01	0.23	27.28	10.06

TABLE 2: **Summary statistics of portfolio weights**: The table presents descriptive statistics of portfolio weights, expressed in percentage, by institutional type for the Thomson Reuters database from 1980Q1 to 2023Q4. The statistics are computed for weights of all stocks in the investment universe $w_i(n)$ and of the outside asset w_0 in each quarter, and then averaged across all quarters in the sample period.

2.3 Estimation Strategy

The demand system in Eq. (5) includes as characteristics a stock's market capitalization, its market beta, and its duration. We estimate market beta from a regression of monthly returns in excess of the 1-month Tbill rate onto excess market returns using a 60-month moving window, requiring at least 24 months of non-missing observations.

Our benchmark measure of equity duration is from Gonçalves (2021b), which adapts the traditional Macaulay duration used for fixed-income securities to equities. The duration of stock n in quarter t is computed as the weighted average of time multiplied by the fraction of future (expected) payoffs over current market equity. The main difference with respect to bond duration is that, for equity, future payoffs are not deterministic. We follow Gonçalves (2021b) and model expected future payouts with a vector autoregressive model; see Appendix A.2 for details. We verify the validity of this measure for the purpose of capturing investors' demand for cash flows at different horizons. Stocks classified as short-duration do indeed pay a larger fraction of their market price in a much shorter time frame compared to their long-duration counterparts; see Appendix Figure A.1. Our main findings extend to other duration measures, as long as they share this feature. For FactSet data, we treat ETFs as separate securities with their own demand functions and compute an ETF's duration as the weighted average duration of its constituent stocks. Over 1980-2023, the average value-weighted equity duration is 44 years, in line with the 39 years estimate of Gonçalves (2021b) for 1973-2017. We observe an upward trend in stock duration from 30 years in 1980 to 50 years in 2000, with fluctuations in the 45-55 years range afterwards; see Appendix Figure A.2. In the demand system, we use log duration, that averages 3.66 with a standard deviation of 0.42 years; see Appendix Table A.2.

Equipped with institutional holdings and stock characteristics, we estimate the demand system in Eq. (5) by GMM using the procedure in Koijen et al. (2024). In particular, we pool holdings for a given manager across four quarters (from $t-3$ to t) and use quarter fixed effects to pin down the quarter-specific intercept. In a given quarter, we perform the estimation separately for each manager with at least 2,000 holdings in the pooled four-quarter window. The 2,000 threshold empirically ensures convergence of the GMM algorithm. For managers below the threshold, we employ a two-step procedure. First, we sort institutions within the same type by their AUM, obtain bins with at least 2,000 holdings, and estimate the demand system for each type-AUM bin. Next, we use a shrinkage estimator to obtain manager-specific demand coefficients. The shrinkage targets are the demand coefficients from the first step, with penalty parameters that are selected by cross-validation. We verify that the investment universe of our set of institutional investors is persistent over time, as the percentage of stocks held in the current quarter that were ever held in the previous one to eleven quarters is 96% across all AUM deciles; see Appendix Table A.3. This evidence is consistent with the presence of exogenous investment mandates, which forms the basis for the instrumental variables approach of Koijen and Yogo (2019) on (log) market equity. We winsorize the estimated demand coefficients at the top and bottom 1% to avoid outliers to exert undue influence. Appendix B provides further details on the estimation.

It is worth emphasizing that while the two-step procedure facilitates the convergence of GMM, it retains considerable heterogeneity in demand coefficients both over time for the same institution and across institutions within a given quarter, as we highlight in the next section. When using FactSet data, we further refine the shrinkage estimation by forming the

AUM-based bins within subgroups of institutional types (e.g., property and casualty versus life insurance companies) to better capture cross-sectional differences in demand.

2.4 Estimates of Demand System

In Table 3, we summarize the estimated demand functions by intermediary type. Panel A reports the coefficient on equity duration, θ_{dur}. Among all quarter-institution observations, insurance companies exhibit the strongest tilt toward high-duration stocks, with an average estimate of 0.14, followed by primary dealers with an average of 0.05. Pension fund portfolios show a slight preference for lower-duration stocks, with an almost symmetric average coefficient of -0.04, while banks display the most negative estimate at -0.17. Panel B presents the demand coefficients on market beta, which are negative across institutions. Banks again show the most negative tilt, with an average of -0.25, while the other intermediary types have coefficients slightly below zero. In terms of economic magnitude, these estimates imply that, considering insurance companies as an example, a 10% increase in a stock's beta is accompanied by a 0.6% decrease in weight (relative to the outside asset), while a 10% increase in its duration raises the weight by 1.4%.[11]

A notable feature of institutional demand for duration is its considerable heterogeneity, as evidenced by its large standard deviation and wide percentiles, which far exceed those of demand for beta. This heterogeneity arises from both aggregate variation and cross-sectional differences in demand. To capture the former, the top plot of Figure 1 displays the time series of the cross-sectional AUM-weighted average of θ_{dur} for each institutional type. We observe a general increase in institutions' tilt toward high-duration stocks in the first part of the sample, a decline in demand leading up to the Global Financial Crisis and around the pandemic (particularly pronounced for insurance companies and primary dealers), and a subsequent recovery in the later period. The first principal component of the series explains

[11] Taking logs of Eq. (5) yields a log-level model for market beta and a log-log model for duration. However, since the average market beta is approximately one, the economic magnitude of the estimates is ultimately directly comparable.

Panel A: Demand for Duration (θ_{dur})				
	Primary Dealers	Banks	Insurances	Pension Funds
mean	0.05	-0.17	0.14	-0.04
std	0.69	1.14	0.98	0.65
p10	-0.75	-1.51	-0.83	-0.70
p50	0.03	-0.14	0.05	-0.05
p90	0.82	1.14	1.30	0.61
N	2,620	21,097	6,784	5,896
Panel B: Demand for β_{mkt} (θ_{mkt})				
	Primary Dealers	Banks	Insurances	Pension Funds
mean	-0.05	-0.25	-0.06	-0.04
std	0.38	0.56	0.48	0.35
p10	-0.49	-0.92	-0.57	-0.38
p50	-0.02	-0.23	-0.06	-0.05
p90	0.36	0.39	0.46	0.33
N	2,620	21,097	6,784	5,896

TABLE 3: **Summary statistics of estimated demand coefficients**: The table reports summary statistics of the demand coefficients on log equity duration (θ_{dur}, Panel A) and market beta (θ_{mkt}, Panel B) from the system in Eq. (5). The coefficients are estimated over 1980Q1–2023Q4 by GMM for each institution-quarter using the procedure described in Section 2.3.

about 50% of the overall variation, confirming the presence of common driving factors in institutional portfolio decisions. Recessions and economic downturns are characterized by an overall decline in the quest for equity duration. Despite these common patterns, several periods exhibit type-specific movements in demand.[12]

To quantify the extent of cross-sectional heterogeneity in demand for duration, we project the coefficients on type-quarter fixed effects and display, in the bottom plot of Figure 1, the histogram of the corresponding residuals. We show separate bars for institutions with more than 2'000 holdings in a given quarter and for other institutions, whose coefficients are estimated using the two-stage shrinkage procedure. The figure reveals substantial variation in institution-specific demand beyond aggregate factors; indeed, the type-quarter fixed effects

[12] Appendix Figure A.3 displays analogous series for investors' demand for market beta. The patterns are broadly similar to those documented by Koijen and Yogo (2019).

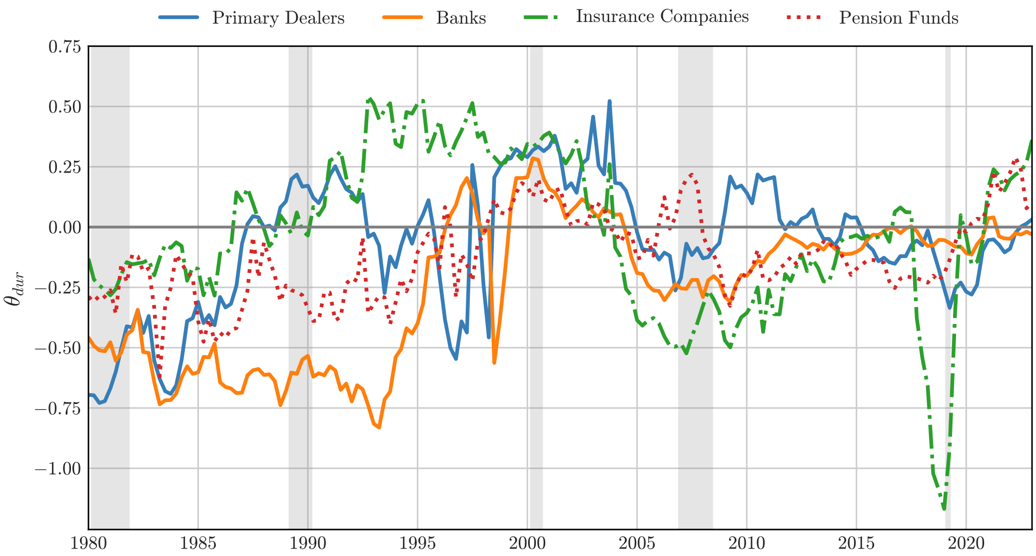
explain only about 10% of the overall variance. We also note that the two histograms largely overlap, which reassures us that the extent of heterogeneity is a genuine feature of institutional portfolios and not an artifact of the estimation approach. In Section 3 below, we link this wealth of variation in institutions' demand for duration to financial constraints and capital availability.

2.5 Validation Tests

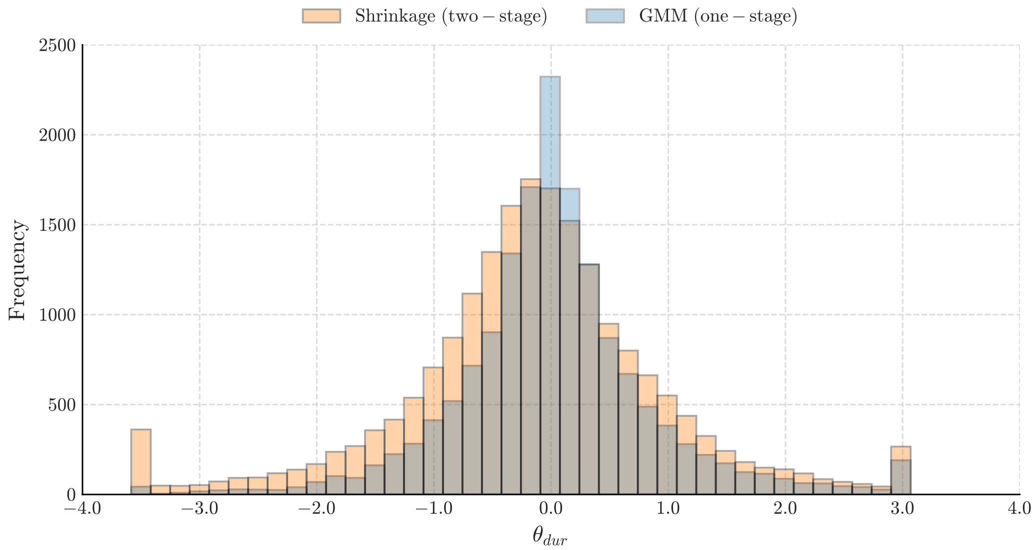
Before proceeding, we provide three analyses that corroborate the validity of our approach. First, regarding the specifics of the demand system framework, we ask whether adding stock duration to market beta is sufficient to capture the heterogeneity in institutional investors' portfolio choices. To this end, we compare our specification with the multi-dimensional system proposed by Koijen and Yogo (2019), where, in addition to market beta, four other characteristics (book equity, profitability, investment, and dividends-to-book equity) also drive institutional demand. We estimate their model on our sample and examine the standard deviation of the unexplained latent demand $\epsilon_{i,t}$ in Eq. (5) across all holdings in a given quarter for a given institutional type. This standard deviation captures the extent of variability in the investor-specific component of demand left unexplained by the model.

Figure 2 plots the resulting cross-sectional standard deviation over our sample period for both our two-variable specification ("mkt+dur," solid line) and their system (five characteristics, dotted line). The two series track each other closely, with the standard deviation from our parsimonious specification being smaller for most quarter-type observations. This result demonstrates that duration is an encompassing dimension of institutions' equity allocation, extending the evidence in Gormsen and Lazarus (2023) from standard cross-sectional asset pricing tests.

Next, we provide out-of-sample validation for our approach by analyzing target-date funds. These funds explicitly state the investment horizon of their asset allocation strategy, making them a natural vehicle for managing retirement savings. We expect target-date funds

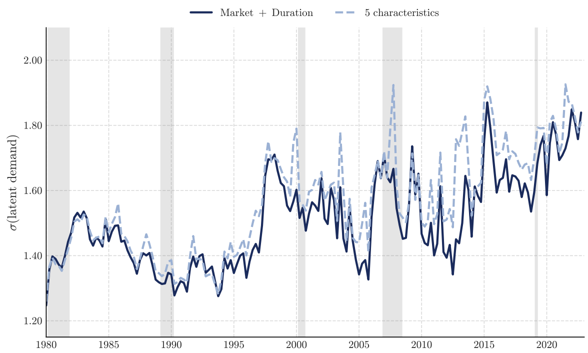


(A) Aggregate demand for duration.

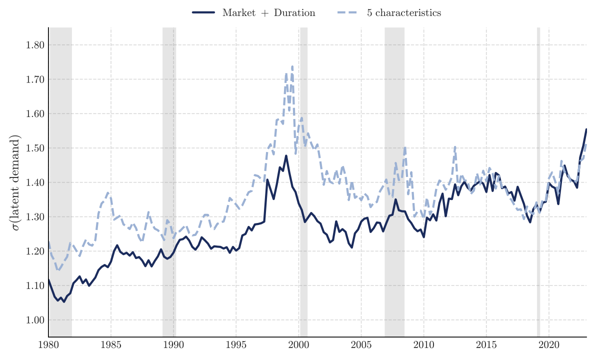


(B) Institution-specific demand for duration.

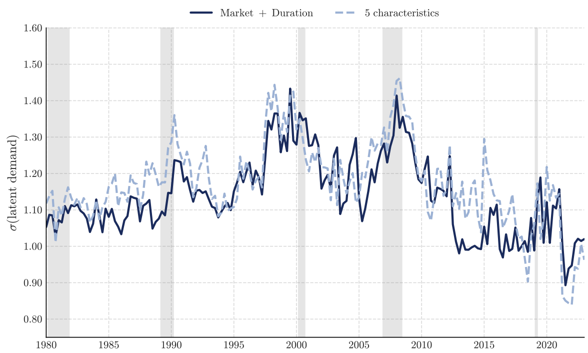
FIGURE 1: **Variation in demand for duration**. Panel A shows the time series of AUM-weighted demand coefficients on duration (θ_{dur}) by institutional type over 1980Q1–2023Q4. Shaded regions indicate NBER recessions. Panel B presents the residuals from regressing institution-quarter demand coefficients on quarter-type fixed effects, with separate histograms for institutions whose demand coefficients are estimated individually (one-stage) or using the two-stage shrinkage procedure (two-stage).



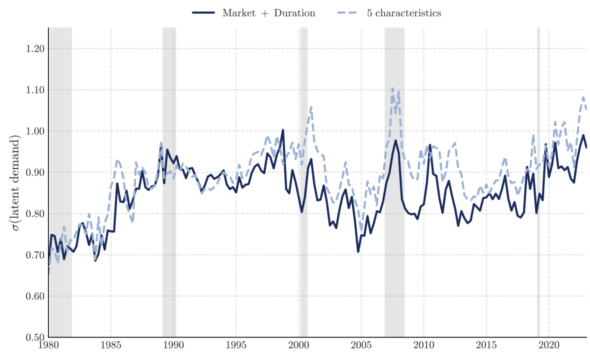
(A) Primary dealers.



(B) Banks.



(C) Insurance companies.



(D) Pension funds.

FIGURE 2: **Latent demand of characteristic-based models.** The figure displays the time series of the cross-sectional standard deviation of log latent demand by institutional type over 1980Q1–2023Q4, resulting from estimating a demand system with market beta and duration ("Market+Duration," solid line) and with the five characteristics in Koijen and Yogo (2019) ("5 characteristics," dashed line). Shaded regions indicate NBER recessions.

with longer horizon mandates to exhibit stronger hedging demand and a greater preference for long-term cash-flow exposure. We identify 1,271 target-date funds from the CRSP Mutual Funds Survivorship-Bias-Free database beginning in 2010Q1, with details of the identification procedure provided in Appendix A.3. Since target-date funds are known to hold a sizable fraction of mutual funds in their portfolios (Parker et al., 2023), we treat mutual fund holdings as separate securities and compute their duration as the weighted average duration of their constituent stocks.[13] To ensure that duration exposure arises within the equity market, rather than from shifts between equity and fixed income, we restrict the analysis to target-date funds with three or more constituent equity funds (defined as funds with at least 80% equity holdings).

We define a target-date fund's time to maturity (TTM) as the difference between its target date and the current quarter. In each quarter, we classify short-horizon funds as those with TTM less than 10 years, mid-horizon funds as those with TTM between 10 and 19 years, and long-horizon funds as those with TTM greater than 19 years. We then estimate our demand system in Eq. (5) separately for these three groups.

The left panel of Figure 3 shows the average of θ_{dur} for the three groups. The demand for duration increases monotonically from low- to high-TTM funds, with the difference between the two extreme groups statistically significant at the 10% level. This direct test supports the interpretation of θ_{dur} as capturing demand for long-term cash flows in the equity market. In the right panel, we repeat the test but include non-equity positions (such as bonds and mortgage-backed securities) in the outside good of the demand system, as available in the CRSP Mutual Funds Survivorship-Bias-Free database. Accounting for these positions preserves the observed difference in θ_{dur} across groups.

Finally, we run additional tests to verify that the results are not mechanically driven by adjustments in equity shares. The positive relation between demand for equity duration and target-date funds' investment horizon holds: (i) when restricting the sample to target-

[13]Information on the stocks held by each mutual fund is taken from the CRSP Mutual Fund Holdings database.

date funds with at least 75% equity, (ii) when estimating fund-level demand coefficients and explicitly controlling for equity share (see Appendix Table A.4), and (iii) when using only direct stock holdings and scaling by the AUM directly invested in equity.

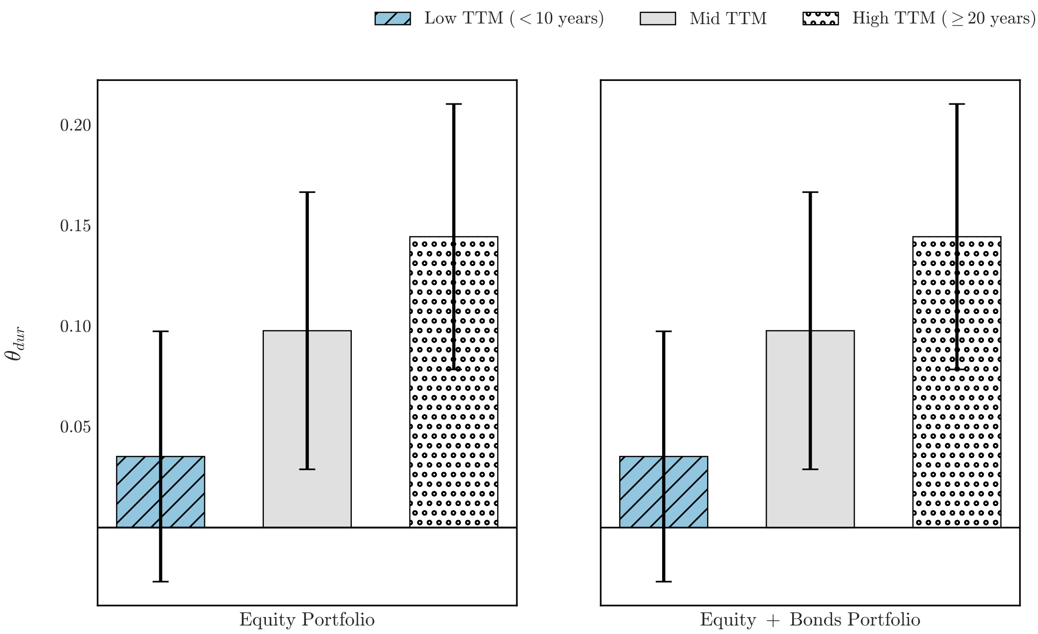


FIGURE 3: **Demand for duration of target-date funds.** The figure displays the average duration demand coefficient (θ_{dur}) for target-date funds with time-to-maturity (TTM) lower than 10 years (blue hatched bar), target-date funds with TTM between 10 and 19 years (gray solid bar) and funds with TTM equal or greater than 20 years (white bar with circles). The demand system in Eq. (5) is estimated pooling target-date funds by their TTM. The sample period is from 2010Q1 to 2023Q4. The leftmost plot reports results considering equity-only holdings, while the rightmost plot also includes holdings of other assets, such as fixed income positions.

Finally, in Appendix Figure A.4, we report the rank correlation between an institution's preference for duration, $\theta_{dur,i,t}$, and its aggregated portfolio duration, computed as the dollar holding-weighted average stock duration. We observe a clear positive relation, indicating that higher θ_{dur} coefficients are broadly associated with greater holdings of long-duration stocks. Nevertheless, the imperfect correlation suggests that the θ_{dur} coefficient captures demand for long-term cash flows beyond a simple weighted average portfolio duration.

3 Demand for Duration and Institutional Risk Aversion

In this section, we investigate the key drivers of institutional portfolio choices. In the ICAPM framework, the hedging component of the optimal equity allocation in Eq. (1) decreases with the investor's relative risk aversion, γ_t. This result suggests that variations in intermediaries' demand for equity duration ought to be driven by their risk-bearing capacity. We test this prediction by regressing the aggregate (average) demand for duration of each institutional type on type-specific measures of risk aversion in a time-series setting. In Section 5.1 below, we refine our identification further by conducting this analysis at a more granular institutional level using a panel regression framework.

We proxy the risk-bearing capacity of financial intermediaries by their capital ratio, i.e., the ratio of equity capital to total assets. Our choice is motivated by evidence from the intermediary asset pricing literature (He et al., 2017) that the capital ratio of primary dealers is a priced risk factor in the cross-section of expected returns across a wide array of asset classes. Adrian and Shin (2010, 2014) use bank's leverage ratio (i.e., one minus the capital ratio) as proxy for financial constraints and show that banks' leverage is a good measure to capture their Value-at-Risk, as reported on banks' filings. For the cross section of insurance companies, Koijen and Yogo (2015) show that leverage ratio emerges as the most important characteristic (among a group that includes risk-based capital relative to guideline, net equity inflow, and asset growth) in driving their shadow cost of capital. These studies suggest that better-capitalized financial institutions are more tolerant to take on risks from intermediating in financial markets.[14]

We employ distinct capital ratios for the institutional types we study. For primary dealers, we directly use the series from He et al. (2017). For the other groups, we compute aggregate series from Compustat based on SIC classification. Specifically, we first construct

[14]The relevance of intermediary leverage may also stem from its correlation with the stochastic discount factor of average and wealthy households (Lettau et al., 2019; Santos and Veronesi, 2022).

an institution's capital ratio, following He et al. (2017), as the ratio of its market equity to the market value of its assets. The aggregate capital ratio for a given type is then computed as the value-weighted average of individual ratios, with weights equal to each company's market-valued asset share. Since the number of pension funds identified from SIC codes is rather limited, we employ a single ratio for the combined group of insurance companies and pension funds. We provide further details in Appendix A.4.

Figure 4 plots the three capital ratios over the 1980-2023 sample period. The series clearly share common patterns, and tend to drop in correspondence to economic downturns such as the Global Financial Crisis and the pandemic outbreak. This evidence implies that our measure of intermediary leverage are countercyclical, consistent with predictions from several theoretical models (Bernanke and Gertler, 1989; He and Krishnamurthy, 2012, 2013; Brunnermeier and Sannikov, 2014). The capital ratios of banks and primary dealers track each other very closely (correlation of 0.95), while that of insurances and pension funds shows more type-specific dynamics and is more disconnected from the others (correlation of about 0.60).

We regard the capital ratios as inverse proxies for intermediaries' risk aversion. To validate this interpretation, we examine their relation with several stock market predictors and risk aversion measures: Shiller's cyclically adjusted price-to-earnings ratio (CAPE); the University of Michigan consumer sentiment index (Sentiment); the market dividend-price ratio (DP Ratio); the inverse relative wealth, computed as in Ilmanen (1995) (INVRELW); the GZ spread of Gilchrist and Zakrajšek (2012), also used in Haddad and Muir (2021) to proxy for intermediary risk aversion; and the risk aversion index from Bekaert et al. (2022) (ra^{BEX}). High CAPE and Sentiment are associated with periods of lower risk premia (i.e., high risk appetite), while the DP Ratio, INVRELW, GZ spread, and ra^{BEX} reflect higher levels of risk aversion. Table 4 reports univariate and joint regressions of the capital ratios of Primary Dealers, Banks, and the common series for Insurance companies and Pension funds on these variables. As expected, we find positive coefficients for CAPE and Sentiment, and

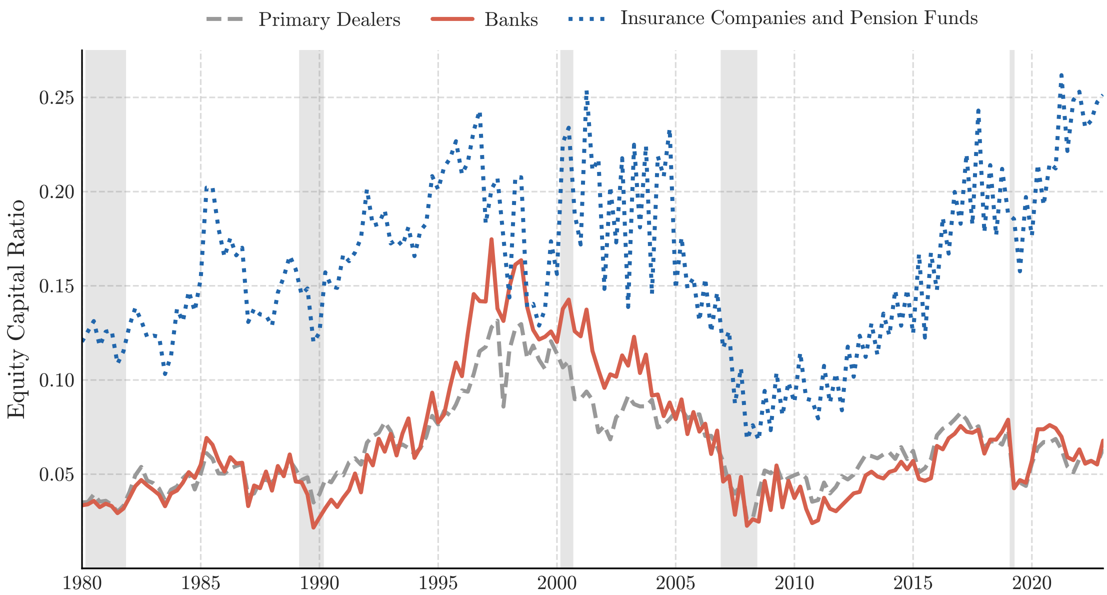


FIGURE 4: **Intermediaries' Capital Ratios**. This figure plots the aggregate capital ratios (i.e., the ratio of market equity to the market value of assets) for primary dealers (gray dashed line), banks (red solid line), and the combined group of insurance companies and pension funds (blue dotted line) over 1980Q1–2023Q4. The series are computed as AUM-weighted averages of individual companies' capital ratios, as detailed in Appendix A.4.

broadly negative loadings for DP Ratio, INVRELW, GZ spread, and ra^{BEX}, with several statistically significant coefficients and some differences across types. Collectively, these variables capture a large fraction, ranging from 47% to 77%, of the variance of capital ratios. Importantly, capital ratios capture institutional balance sheet conditions and can be computed at both the type level, used in the following time-series analysis, and the institution level, which allows us to perform cross-sectional tests in Section 5.1.

We next study the relationship between demand for equity duration and intermediaries' risk aversion in a time-series regression setting. Specifically, we project the average AUM-weighted demand for duration (θ_{dur}) of a given institutional type in a given quarter from the Thomson Reuters database onto its corresponding capital ratio series, η. To mitigate concerns of spurious correlation, we lag the capital ratios by one quarter and standardize them to have a mean of zero and a unit standard deviation, facilitating the interpretation of the coefficients.

Variable	(1) Univariate	(2) Joint	(3) Univariate	(4) Joint	(5) Univariate	(6) Joint
	Primary Dealers		Banks		Insurances and Pension funds	
CAPE	1.822***	0.945***	2.416***	0.679	2.234***	1.704
	(0.215)	(0.339)	(0.332)	(0.588)	(0.475)	(1.425)
Sentiment	1.492***	1.247***	2.041***	2.121***	1.306*	0.745
	(0.274)	(0.159)	(0.408)	(0.263)	(0.790)	(0.535)
DP Ratio	-1.544***	-1.163***	-2.024***	-2.707***	-1.877***	-1.493
	(0.230)	(0.401)	(0.363)	(0.764)	(0.337)	(1.715)
INVRELW	-0.706**	0.474**	-0.629	1.065***	-0.725	2.976***
	(0.344)	(0.213)	(0.559)	(0.407)	(0.659)	(0.767)
GZ	-0.073	-0.264	0.073	-0.757**	-1.207**	-3.617***
	(0.373)	(0.169)	(0.548)	(0.314)	(0.546)	(0.722)
ra^{BEX}	-0.562***	-0.138	-0.370	0.321	-1.169**	-0.116
	(0.182)	(0.218)	(0.352)	(0.424)	(0.534)	(0.541)
R^2 (Univ./Joint)	[0.0926–0.616]	0.772	[0.034–0.502]	0.713	[0.0257–0.244]	0.471

Table 4: **Capital Ratios and Risk Premia**: The table reports OLS coefficients, with standard errors in parentheses, from regressions of intermediaries' type-specific capital ratios (η) on the following variables: Shiller's cyclically adjusted price-to-earnings ratio (CAPE); the University of Michigan consumer sentiment index (Sentiment); the market dividend-price ratio (DP Ratio); the inverse relative wealth, computed as in Ilmanen (1995) (INVRELW); the GZ spread of Gilchrist and Zakrajšek (2012) (GZ); and the risk aversion index from Bekaert et al. (2022) (ra^{BEX}). Columns 1, 3, and 5 report estimates from univariate regressions and, in the last row, the range in univariate R^2. Columns 2, 4, 6 report estimates from multivariate regressions and, in the last row, the corresponding R^2. All explanatory variables are standardized and expressed in percentages. The sample period is 1980Q1–2023Q4, except for ra^{BEX}, whose series begins in 1986Q2. One, two, and three asterisks indicate significance at the 10%, 5%, and 1% level, respectively.

These univariate regressions are reported in Panel A of Table 5, in every first column of each institutional type (i.e., columns 1, 3, 5, and 7). Across all types, we find consistently positive and highly statistically significant coefficients on η. This result aligns with the ICAPM prediction, implying that institutions increase their demand for long-term dividend claims in periods of high capital ratios, when intermediaries' risk-bearing capacity is stronger. Conversely, when equity capital is low – such as, during the Global Financial Crisis – risk aversion surges, and intermediaries reduce their demand for long-term equity, thereby becoming more exposed to short-lived shocks. The effect of changes in risk-bearing capacity on portfolio tilt is strongest for banks, where a one-standard-deviation increase in η is associated with a 0.150 increase in demand for duration, and weakest for pension funds, where the same shock leads to a more modest 0.051 shift.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
				Panel A: Demand for equity duration				
	Primary dealers		Banks		Insurances		Pension funds	
η_{t-1}	0.091**	0.091**	0.150***	0.145***	0.134***	0.134***	0.051**	0.052***
	(0.039)	(0.039)	(0.033)	(0.033)	(0.033)	(0.033)	(0.020)	(0.017)
$FinUnc_{t-1}$		0.017		0.058*		-0.015		0.048**
		(0.020)		(0.031)		(0.043)		(0.020)
Obs.	172	172	172	172	172	172	172	172
R^2	0.122	0.126	0.273	0.314	0.168	0.170	0.085	0.160
				Panel B: Demand for market beta				
	Primary dealers		Banks		Insurances		Pension funds	
η_{t-1}	-0.035***	-0.035***	-0.026	-0.025	-0.043**	-0.044**	-0.036***	-0.036***
	(0.009)	(0.008)	(0.017)	(0.017)	(0.021)	(0.019)	(0.013)	(0.012)
$FinUnc_{t-1}$		-0.005		-0.016		-0.038**		-0.021**
		(0.011)		(0.011)		(0.015)		(0.010)
Obs.	172	172	172	172	172	172	172	172
R^2	0.130	0.133	0.051	0.070	0.079	0.143	0.129	0.177

TABLE 5: **Institutional Demand and Capital Ratios**: The table reports OLS coefficients, with standard errors in parentheses, from regressions of the average AUM-weighted intermediaries' demand for equity duration (θ_{dur}, Panel A) and market beta (θ_{mkt}, Panel B) on the lagged type-specific capital ratio η. In columns 2, 4, 6, and 8, we additionally include the financial uncertainty index of Jurado et al. (2015). All explanatory variables are lagged by one quarter and standardized. The sample period is 1980Q1–2023Q4, and the demand coefficients for each institution-quarter observation in the Thomson Reuters database are obtained following the procedure described in Section 2.3. One, two, and three asterisks indicate significance at the 10%, 5%, and 1% level, respectively.

We further address the concern that the capital ratio might capture shifts in future investment opportunities rather than changes in preferences. If periods of low intermediaries' risk aversion (i.e., high capital ratios) are associated with a perception of heightened reinvestment risk, our findings might actually reflect a preference for long-duration stocks from hedging motives. To rule out this alternative explanation, in every second column of Panel A, we report estimates from a multivariate regression that includes the financial uncertainty index of Jurado et al. (2015) as a proxy for variations in expected investment conditions, which we also lag by one quarter and standardize. We find that our main results hold in this augmented specification. As expected, the uncertainty index is associated with stronger demand for long-duration stocks for three out of four institutional types, with significant coefficients for banks and pension funds (the exception being insurance companies, for which the coefficient is negative but insignificant). The inclusion of the uncertainty index does not diminish the importance of capital ratios, whose coefficients remain close to those in the univariate specification. Overall, the finding that institutions increase demand for long-term dividend claims when capital ratios are high supports the ICAPM implication that lower risk aversion, proxied by the inverse of the capital ratio, boosts demand for long-term equity.

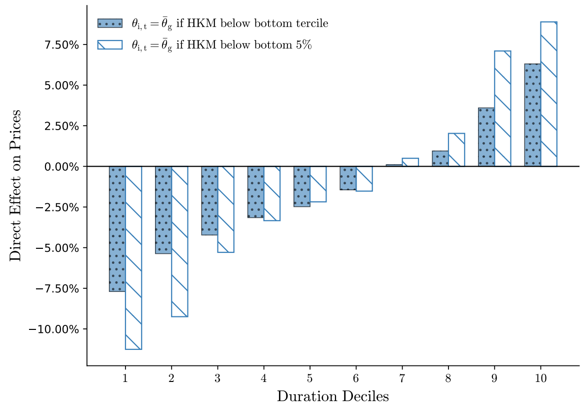
Panel B of Table 5 reports estimates using the same predictors to explain variation in intermediaries' demand for market beta. We find that capital ratios have a consistently negative impact. The coefficients are statistically significant for all types except banks, but are smaller in magnitude than those for equity duration, ranging between -0.045 and -0.025. The negative sign indicates that a decrease in intermediaries' risk-bearing capacity is accompanied by a stronger tilt toward high-beta stocks. This result does not align with the ICAPM optimal portfolio policy in Eq. (1), where the myopic component of an investor's demand decreases with her risk aversion coefficient. Rather, our finding shares similarities with the mechanism in Frazzini and Pedersen (2014), where constrained institutional investors achieve higher (implied) leverage by overweighting high-beta stocks, leading to a flattening of the security market line. Our evidence that intermediaries increase their demand for market beta in times of scarce capital suggests a more complex structure of the portfolio's myopic component, one where changes in intermediaries' risk aversion interact with leverage and margin constraints.

4 Intermediaries' Demand and Duration Premium

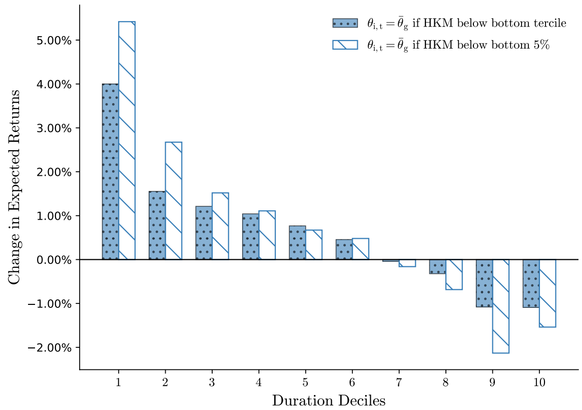
We quantify the effect of financial intermediaries' preferences for long-term equity claims on the duration premium through a counterfactual analysis. Specifically, using the equilibrium conditions from the demand system, we compute market-clearing prices in a given quarter under alternative assumptions about intermediaries' demand for duration. To this end, we estimate the demand function of all other investors in the Thomson Reuters dataset over the 1980–2023 period—namely mutual funds, investment advisers, other 13F institutions, and the residual household group—following Koijen and Yogo (2019) in the identification of institutional types. We then compute counterfactual equilibrium prices by applying the algorithm in Appendix C of Koijen and Yogo (2019), iterating until convergence is achieved.

We conduct two experiments corresponding to either a positive or a negative shift in their demand for duration. In the first experiment, we bring the demand coefficients for (log) duration, $\theta_{dur,i,t}$, during periods of reduced capital-bearing capacity to their within-group unconditional average for all institutions in our sample – primary dealers, banks, pension funds, and insurance companies. We identify the affected periods as quarters in which the aggregate intermediaries capital ratio of He et al. (2017) falls in the bottom of its distribution. Preferences for market betas and demand elasticities are left unchanged at the their estimated values throughout. As documented in the previous section, periods of reduced capital-bearing capacity are associated with below-average demand for duration. Therefore, this counterfactual scenario corresponds to a positive shift in duration preference relative to actual demand in quarters when capital ratios are low, resulting in stronger demand pressure for long-term claims by intermediaries.

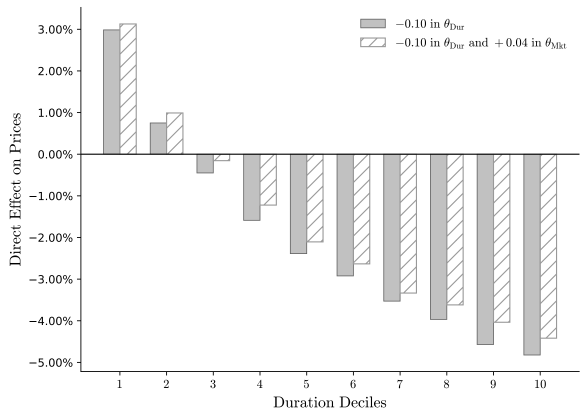
Panel A of Figure 5 displays the annualized difference between counterfactual and actual prices, expressed as a percentage of the latter, across duration deciles. These differences are averaged across quarters in which the capital ratio of He et al. (2017) falls in either the bottom tercile or the bottom 5% of its unconditional distribution. Several findings are worth noting. First, the shift leads to a decline in the prices of short-duration stocks and an increase in the prices of long-duration stocks, consistent with the mechanism described above and our regression analysis. Second, the re-pricing is monotonic across duration deciles. This pattern is genuine and not mechanically driven by the structure of the demand system, as the equilibrium impact of the preference shift depends on the joint distribution of holdings, demand coefficients, and price elasticities across all investors holding a given stock. Third,



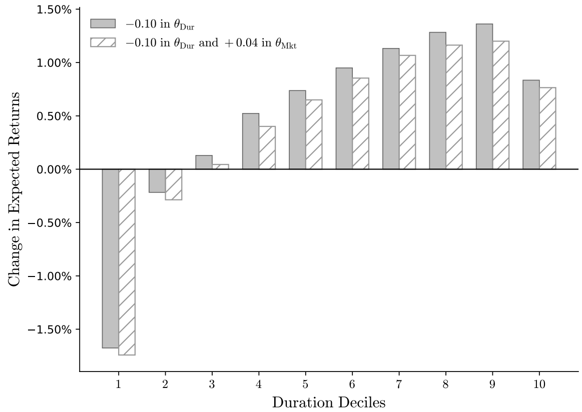
(A) Price impact.



(B) Change in Expected returns.



(C) Price impact.



(D) Change in Expected returns.

FIGURE 5: **Price Impact and Change in Expected Return of a Shift in Preference for Duration.** The figure displays the annualized impact on stock prices (left panels) and the corresponding change in annualized expected returns (right panels) across duration deciles, following shifts in duration demand by primary dealers, banks, insurance companies, and pension funds. In the top panels, each institution's duration demand coefficient is set equal to its average value over the sample period, while holding constant preferences for market beta and demand elasticity. Price and expected return differences are averaged across quarters in which the capital ratio of He et al. (2017) falls within either the bottom tercile or the bottom 5% of its unconditional distribution. In the bottom panels, the duration demand coefficient $\theta_{dur,i,t}$ is decreased by 0.10 (solid bars), with an alternative scenario that simultaneously increases the market beta demand coefficient by 0.04 (white bars). The counterfactual prices are computed for each quarter during 1980Q1–2023Q4, using holdings data from the Thomson Reuters database.

the magnitude of the price effects grows as the positive shifts in duration preferences relative to actual demand become larger – for example, when considering quarters in which the capital ratio, as defined by He et al. (2017), falls in the bottom 5% rather than the bottom tercile of its unconditional distribution. This effect is most pronounced in the extreme low- and high-duration portfolios.

In the second experiment, we decrease the duration demand coefficient $\theta_{dur,i,t}$ of our intermediaries by 0.10, a value that corresponds to the expected change following a one-standard-deviation decline in equity capital, as reported in Table 5. This counterfactual is meant to capture a shift in preferences toward equity claims with nearer-term payoffs, reflecting reduced hedging demand for long-duration assets.

Moreover, Table 5 also shows that the myopic component of equity demand, proxied by sensitivity to market beta, responds to shocks in risk-bearing capacity in the opposite direction to duration demand. The pricing implications of these shifts depend on the correlation between duration and market beta exposures in the cross section of stocks, potentially amplifying or dampening the overall effect. To account for this interaction, we consider a joint adjustment: we reduce $\theta_{dur,i,t}$ by 0.10 while simultaneously increasing $\theta_{mkt,i,t}$ by 0.04, consistent with the estimates in Table 5.

Panel C of Figure 5 reports the percentage change in counterfactual prices relative to actual prices, averaged over the sample period and across duration deciles. The solid bars correspond to the scenario in which only $\theta_{dur,i,t}$ is shifted by -0.10. As shown, this adjustment leads to a marked increase in the prices of short-duration stocks, by approximately +3% on an annualized basis, which declines monotonically across deciles, turning negative by the third decile and reaching around −5% for the highest-duration stocks. These effects are only partially offset when we account for the simultaneous increase in market beta demand, shown in the white bars. That is, incorporating the change in myopic demand has only a limited impact on the overall pricing pattern.

The hypothetical shifts in demand by our intermediaries translate into counterfactual prices through market clearing, with the differences from actual prices reflecting the market pressure following changes in holdings. These price effects can be interpreted as changes in expected returns, assuming they follow a first-order autoregressive process with root ϕ. Based on the literature on stock return predictability (Campbell and Shiller, 1988; Cochrane, 2008), the implied change in expected returns can be obtained by multiplying the price effects by a factor of $(\rho\phi - 1)$, where $\rho = 1/(1 + e^{\overline{dp}})$ and dp is the log dividend-to-price ratio.[15] We estimate ρ and ϕ for each duration decile using their dividend-to-price ratio, and apply this transformation to convert the counterfactual price differences into expected return units. The right plots in Figure 5 show the resulting annualized expected returns across duration deciles implied by each scenario. Since $(\rho\phi - 1)$ is approximately -0.20 across deciles, the price shocks translate into revisions in expected returns that are smaller in magnitude and of the opposite sign. Furthermore, because ρ and ϕ are decile-specific, the resulting changes in expected returns are no longer strictly monotonic, though they remain nearly so empirically.

As shown in Panel B, the hypothetical increase in duration demand during bad times in the first scenario implies an almost 4% increase in expected returns for low-duration stocks, and a 1% decrease for those in the top two deciles. In contrast, Panel D illustrates that a decline in duration demand accompanied by an increase in demand for market beta generates a negative shock of −1.50% in expected returns for stocks in the bottom duration decile, and a 1% positive shock for those in the top duration decile. To put this full-sample figure in perspective, note that the annualized difference in average returns between the top and bottom duration deciles over our sample period is approximately −7.50%, consistent with the findings in Gonçalves (2021b). Therefore, our second experiment that reduces intermediaries' demand for high-duration stocks throughout the sample dampens the unconditional duration premium by about one third, i.e., 2.5% out of 7.5%. Overall, our analysis reveals that hypothetical shifts in institutional preferences – whether toward or away from duration – can have a sizable impact on the duration premium.

An important consideration in interpreting these counterfactual exercises is that, in the Thomson Reuters database, stock holdings are aggregated at the parent-company level. As such, the magnitudes we report should be viewed as an upper bound on the effect of intermediaries' demand shifts on the duration premium, assuming that shocks to risk-bearing

[15] We acknowledge the ongoing debate on estimating demand elasticities and the view that investors may respond more strongly to price changes at longer horizons (van der Beck, 2022; Binsbergen et al., 2025). Inferring expected returns from counterfactual prices via the present-value identity allows us to capture how persistent shifts in discount rates map into asset prices.

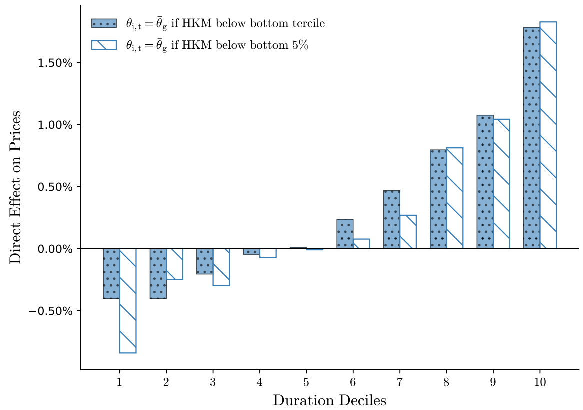
capacity spill over to the entire holding company – directly or through impaired internal capital markets triggered by severe subsidiary distress, as noted by He et al. (2017).

To evaluate how shifts in the preferences of the individual entities reporting in 13F filings influence asset prices, we exploit the finer granularity offered by the FactSet database. As noted earlier, using this alternative reporting unit reduces the share of the stock market attributable to these institutions to 9% of total market capitalization. Moreover, these entities may respond to price changes with different elasticities and exert a distinct marginal influence on price formation. How such differences translate into equilibrium prices is, therefore, an empirical matter.

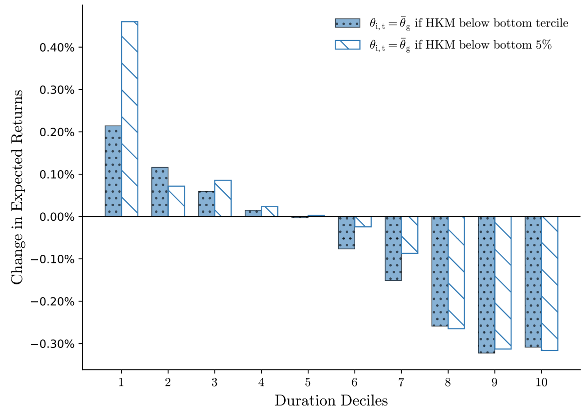
We repeat the two counterfactual experiments using the FactSet dataset and present the results in Figure 6, maintaining the same layout as in Figure 5. The figure shows once again that preference shifts exert a monotonic impact on prices and a broadly monotonic impact on expected returns across equity duration deciles. In both scenarios, the resulting effect on the return spread between high- and low-duration portfolios is about 0.5%, i.e., roughly one-fifteenth of the duration premium and one-fifth of the effect obtained using Thomson Reuters data. Given that assets under management for the same institutional groups are about three times larger in the Thomson Reuters database, these findings suggest that shocks to duration demand at the holding-company level lead to more than proportional redistribution effects across the equity market.

5 Cross-sectional analysis and alternative measures of duration

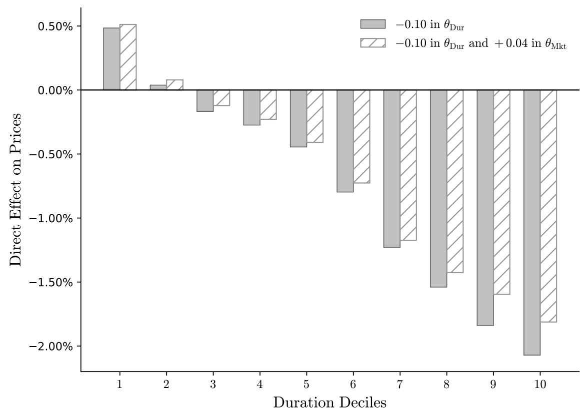
Our demand-system estimates indicate that a significant fraction of the variation in intermediaries' portfolio choices, especially with respect to equity duration, is specific to individual institutions. In Section 5.1, we relate each institution's demand for duration to



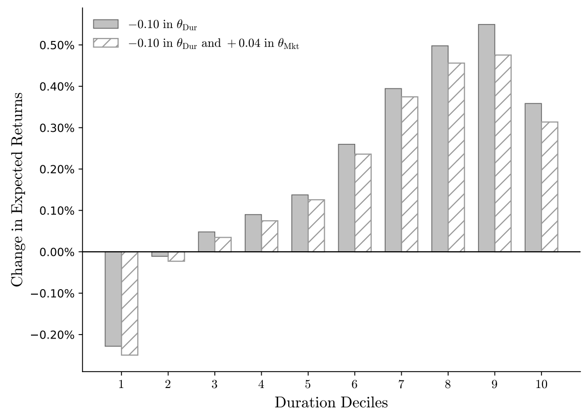
(A) Price impact.



(B) Change in Expected returns.



(C) Price impact.



(D) Change in Expected returns.

FIGURE 6: **Price Impact and Change in Expected Return of a Shift in Preference for Duration, FactSet data**. The figure displays the annualized impact on stock prices (left panels) and the corresponding change in annualized expected returns (right panels) across duration deciles, following shifts in duration demand by primary dealers, banks, insurance companies, and pension funds. In the top panels, each institution's duration demand coefficient is set equal to its average value over the sample period, while holding constant preferences for market beta and demand elasticity. Price and expected return differences are averaged across quarters in which the capital ratio of He et al. (2017) falls within either the bottom tercile or the bottom 5% of its unconditional distribution. In the bottom panels, the duration demand coefficient $\theta_{dur,i,t}$ is decreased by 0.10 (solid bars), with an alternative scenario that simultaneously increases the market beta demand coefficient by 0.04 (white bars). The counterfactual prices are computed for each quarter during 1999Q1–2023Q4, using holdings data from the FactSet database.

its capital availability using the FactSet holdings database, which provides institution-level portfolio detail and, as discussed above, offers the most comprehensive coverage and data quality across multiple dimensions. This finer approach strengthens the case for a causal interpretation of our results. Section 5.2 examines alternative duration measures and presents an additional, instrumented specification of our model.

5.1 Cross-sectional Analysis

To begin, we verify that our time-series evidence extends to the 1999Q1–2023Q4 period covered by FactSet. Appendix Table A.5 reports specifications analogous to those in Table 5, but with the dependent variable being the average demand from FactSet holdings data. Over this reduced time span, we continue to find that equity capital broadly predicts demand for equity duration positively and demand for market beta negatively.

Next, we investigate the determinants of intermediaries' demand for equity duration by separately analyzing banks and primary dealers, insurance companies, and finally pension funds. Our analysis draws from prior studies that examine their asset/liability structure, income sources, and regulatory framework to identify key drivers of portfolio decisions. Accordingly, we construct institution-level explanatory variables that capture intermediaries' risk tolerance and use panel regressions that pool all available institution-quarter observations. In all models, we lag the regressors by one quarter and include time fixed effects to absorb systematic patterns and isolate cross-sectional variation.

5.1.1 Banks and Primary Dealers

We begin by analyzing the combined group of banks and primary dealers.[16] In Panel A of Table 6, we relate their demand for equity duration to risk-absorption capability from the balance sheet (of the bank's holding company). In column 1, we capture risk tolerance with the institution-level equity capital ratio, η. We obtain a positive and significant coefficient, indicating that the time-series result of Table 5 extends to the cross-sectional level. We find a similar, albeit statistically weaker, result for the Tier 1 equity-to-assets ratio in column 2. Finally, in column 3, we consider Value-at-Risk to assets, a widely used alternative to balance sheet metrics (Adrian and Shin, 2010). The negative and significant coefficient suggests that tighter capital constraints prompt banks to tilt their portfolios toward short-

[16]We pool the two groups since primary dealers are too few to draw inference in a panel setting. For simplicity, in this section, we refer to all these institutions simply as 'banks'.

	(1)	(2)	(3)
Panel A: Demand for equity duration			
Capital Ratio	0.454***		
	(0.037)		
Tier 1 Capital Ratio		0.351*	
		(0.183)	
Value-at-Risk			-7.583**
			(3.713)
Obs.	7,158	8,034	965
R^2	0.067	0.046	0.045
Time FE	Yes	Yes	Yes
Panel B: Demand for market beta			
Capital Ratio	-0.715***		
	(-0.037)		
Tier 1 Capital Ratio		-2.847***	
		(-0.195)	
Value-at-Risk			12.454***
			(-3.177)
Obs.	7158	8034	965
R^2	0.118	0.062	0.078
Time FE	Yes	Yes	Yes

Table 6: **Determinants of Banks and Primary Dealers' Demand**: The table reports OLS coefficients, with time-clustered standard errors in parentheses, from panel regressions of banks and primary dealers' demand for equity duration (Panel A) and market beta (Panel B) on the following institution-specific explanatory variables: the capital ratio; the tier 1 capital ratio; and Value-at-Risk. The regressors are lagged by one quarter. The sample period is 1999Q1–2023Q4, and the demand coefficients for each institution-quarter observation in the FactSet database are obtained following the procedure described in Section 2.3. One, two, and three asterisks indicate significance at the 10%, 5%, and 1% level, respectively.

duration equities.

Panel B of the table shows that the relationship between financial constraints and asset demand differs between beta and duration. While tighter constraints, such as a lower equity-to-capital ratio or higher VaR-to-assets, lead banks to reduce their duration exposure (as shown in Panel A), they also lead to increased demand for higher-beta stocks. This inverse relationship underscores the distinct ways capital availability influence banks' portfolio allocation decisions.

We further exploit changes in banking regulation to reinforce a causal interpretation of our findings. Specifically, we examine the portfolio shifts triggered by the Supplementary

	(1)	(2)
Panel A: Demand for equity duration		
SLR × Post	-0.044*	0.084**
	(0.025)	(0.036)
SLR × Post × LowCapital		-0.298***
		(0.058)
Post × LowCapital		0.187***
		(0.043)
Obs.	11,602	8,400
R^2	0.383	0.358
Time FE	Yes	Yes
Manager FE	Yes	Yes
Panel B: Demand for market beta		
SLR × Post	-0.126***	-0.250***
	(0.019)	(0.027)
SLR × Post × LowCapital		0.293***
		(0.045)
Post × LowCapital		-0.092***
		(0.033)
Obs.	11,602	8,400
R^2	0.461	0.431
Time FE	Yes	Yes
Manager FE	Yes	Yes

TABLE 7: **Determinants of Banks and Primary Dealers' Demand and SLR Policy**: The table reports OLS coefficients, with time-clustered standard errors in parentheses, from panel regressions of banks and primary dealers' demand for equity duration (Panel A) and market beta (Panel B) on institution-specific explanatory variables based on interactions between: a dummy variable for institutions subject to SLR policy, SLR; a post-2014Q3 dummy variable, Post; a dummy variable for institutions with below-median equity capital ratios, Low Capital. The sample period is 1999Q1–2023Q4, and the demand coefficients for each institution-quarter observation in the FactSet database are obtained following the procedure described in Section 2.3. One, two, and three asterisks indicate significance at the 10%, 5%, and 1% level, respectively.

Leverage Ratio (SLR), a regulatory policy introduced in 2014Q3 that requires Tier 1 capital to cover on-balance-sheet assets and certain off-balance-sheet exposures. Walz (2024) shows that these stricter capital requirements caused banks to shift their portfolios toward long-term government bonds. We study how the policy impacted their equity portfolio demand. We construct a dummy variable, SLR, which takes the value of one for treated institutions – those with assets exceeding \$250 billion or foreign exposures above \$10 billion. A post-2014Q3 dummy, Post, captures the SLR implementation period.

In Panel A of Table 7, we relate equity demand to the interaction between (lagged) SLR and Post, further including institution-level fixed effects to isolate the effect of regulation. The coefficient is negative and marginally significant, implying that the overall effect of the policy was to somewhat reduce the demand of treated SLR banks for high-duration stocks. In column 2, we investigate the differential effect of the regulatory shock for banks with below-median equity capital ratios (Low Capital). The coefficient on the interaction term (Post × LowCapital) is positive at 0.187, indicating that non-SLR, constrained banks increased their demand for duration. Conversely, the coefficient on the triple interaction term (SLR × Post × LowCapital) is larger in magnitude at −0.298 and highly significant. This suggests that regulated, capital-constrained banks shifted toward stocks providing short-term cash flows.

Panel B of the table confirms, within the SLR framework, that shocks to risk-absorbing capacity affect demand for beta and duration in opposite directions. The shift toward higher-beta assets is concentrated among SLR banks with low capital, while the coefficient for risk-based capital is insignificant.

5.1.2 Insurances

In Panel A of Table 8, we study the drivers of equity demand by Property&Casualty (P&C), Life, and Multiline insurers. Column 1 shows that a decline in an institution's equity capital is again associated with a decrease in its demand for duration, which implies that as institutions become more capital constrained they reduce their hedging demand. Consistently, in column 2 we find that an increase in the Loss Ratio (the ratio of operating losses to total assets) is accompanied with lower tilt toward long-term dividend claims. In column 3, we incorporate the approach proposed by Ge and Weisbach (2021), where insurers' operating losses serve as proxies for negative shocks to their financial conditions. We attempt to capture this effect with an interaction term between the Loss Ratio and a dummy variable that is one for companies with below-median capital ratio, LowCapital. The corresponding coefficient is negative as expected, but statistically insignificant.

In Panel B of the table we estimate the models on insurers' demand for beta. The portfolio effects are particularly pronounced when using operating losses as financial shock, as in Ge and Weisbach (2021), especially for Property& Casualty insurers. This evidence aligns with the nature of P&C businesses, where weather-related disasters can trigger significant claims in regions where insurers have a substantial market presence. The positive and statistically significant coefficient on Loss indicates that, following operating losses, P&C insurers tend to increase their holdings of riskier, high-beta stocks. However, there is considerable heterogeneity in the effect of losses on portfolio decision. Specifically, the significant interaction term reveals that capital constrained insurers (with below-median equity capital) respond to losses by shifting toward safer, lower-beta portfolios. For these insurers, losses have a greater impact, leading to a reduction in risk-taking behavior.

As we did with banks, we focus on episodes that allow for a cleaner identification of shocks to capital-absorption capacity. Koijen and Yogo (2015) study insurance companies that undertook significant recapitalization efforts between September 2008 and July 2009, such as applying for government assistance, issuing public equity, or suspending dividends. Taking such drastic measures signals that these firms were likely financially constrained during the financial crisis. Ge and Weisbach (2021) also use the financial crisis as a laboratory, when financing frictions were more severe, making the effect of operating losses on insurers' allocations more pronounced. We construct a dummy variable, Recapitalize, that equals one for all insurance companies (and their subsidiaries) that successfully raised new capital in the post-crisis period, which we denote with the Post dummy.

Column 1 of Table 9 shows that the demand for equity duration by recapitalized insurance companies significantly increased in the post-crisis period. This result indicates that insurers with revamped risk-bearing capacity tilted their portfolios more heavily toward long-duration equities. In column 2, we find that for these select companies, the effect of recapitalization remains intact after controlling for the capital ratio and loss ratio. This finding lends further support to the theoretical prediction that the hedging component of demand of long-term

	(1)	(2)	(3)	(4)
		Panel A: Demand for equity duration		
Companies:	All	All	All	P&C and Multiline
Capital Ratio	0.503***		0.695***	0.522***
	(0.086)		(0.159)	(0.186)
Loss Ratio		-8.943**	-3.680	3.169
		(3.530)	(3.352)	(3.216)
LowCapital × Loss Ratio			-5.784	-3.657
			(7.371)	(6.785)
Obs.	1,110	1,189	1,023	493
R^2	0.185	0.166	0.187	0.494
Time FE	Yes	Yes	Yes	Yes
		Panel B: Demand for market beta		
Companies:	All	All	All	P&C and Multiline
Capital Ratio	-0.002		0.704***	1.323***
	(0.103)		(0.208)	(0.323)
Loss Ratio		0.557	3.818	11.970***
		(2.626)	(3.087)	(3.647)
LowCapital × Loss Ratio			-8.136	-20.147***
			(5.066)	(6.633)
Obs.	1,110	1,189	1,023	493
R^2	0.185	0.131	0.135	0.282
Time FE	Yes	Yes	Yes	Yes

TABLE 8: **Determinants of Insurance Companies' Demand**: The table reports OLS coefficients, with time-clustered standard errors in parentheses, from panel regressions of insurance companies' demand for equity duration (Panel A) and market beta (Panel B) on the following institution-specific explanatory variables: the capital ratio; the ratio of losses to total assets, Loss Ratio; an interaction term between the Loss Ratio and a dummy variable for institutions with below-median equity capital ratios, Low Capital. The regressors are lagged by one quarter. The sample period is 1999Q1–2023Q4, and the demand coefficients for each institution-quarter observation in the FactSet database are obtained following the procedure described in Section 2.3. One, two, and three asterisks indicate significance at the 10%, 5%, and 1% level, respectively.

investors (insurance companies, in this case) is positively related to their risk tolerance.

5.1.3 Pension Funds

In columns 1 and 2 of Table 10 we study the determinants of equity portfolio demand from U.S. corporate defined benefit (DB) pension plans. Our analysis builds on Franzoni and Marin (2006), who study the effect of pension plan underfunding on their sponsoring company valuations in a portfolio-sorting setting. In the first column, we use the funding

	(1)	(2)
Recapitalize × Post	0.453***	0.219**
	(0.081)	(0.110)
Capital Ratio		-0.215
		(0.261)
Loss Ratio		4.420
		(3.018)
Obs.	2,382	1,022
R^2	0.489	0.488
Time FE	Yes	Yes
Manager FE	Yes	Yes

TABLE 9: **Insurance Companies' Demand for Duration and Capital Shock**: The table reports OLS coefficients, with time-clustered standard errors in parentheses, from panel regressions of banks and primary dealers' demand for equity duration on explanatory variables. In column 1, the regressor is the interaction term between a dummy variable for companies that successfully recapitalized in 2009 (Recapitalize) and post-2009 dummy (Post). In column 2, we add the companies' capital ratio and loss ratio, lagged by one quarter. The sample period is 1999Q1–2023Q4, and the demand coefficients for each institution-quarter observation in the FactSet database are obtained following the procedure described in Section 2.3. One, two, and three asterisks indicate significance at the 10%, 5%, and 1% level, respectively.

level of pension funds as a proxy for their risk-bearing capacity. Following Franzoni and Marin (2006), we compute the funding ratio as the difference between a pension plan's assets and liabilities, scaled by the firm's market capitalization. This measure is constructed using accounting data on the fair value of plan assets and the projected benefit obligation and is motivated by the Employee Retirement Income Security Act of 1974, which mandates that underfunded plans be restored within 3–5 years through required contributions. Since funds that accumulate losses on their assets are more likely to enter underfunded status, we include a fund's period loss (computed in FactSet as actual return on assets over beginning of period plan assets) as an alternative explanatory variable in the second column.

Our panel regression estimates reveal that funds experiencing negative shocks to their capital availability – whether due to high losses or a low funding ratio – adjust their portfolios by shifting toward low-duration, high-expected-return stocks (Panel A), while simultaneously increasing their exposure to stocks with higher beta (Panel B).

In column 3, we focus on U.S. public defined benefit (DB) pension funds, which have greater discretion in setting liability discount rates compared to private pension funds. Under

	(1)	(2)		(3)
Panel A: Demand for equity duration				
Corporate DB pension plans			Public DB pension plans	
Funding Ratio	1.357***		%Retired	-0.146***
	(0.225)			(0.042)
Period Loss		-3.895***		
		(0.536)		
Obs.	332	209	Obs.	1,540
R^2	0.300	0.481	R^2	0.118
Time FE	Yes	Yes	Time FE	Yes
Panel B: Demand for market beta				
Corporate DB pension plans			Public DB pension plans	
Funding Ratio	-0.540**		%Retired	0.010
	(0.229)			(0.021)
Period Loss		2.214**		
		(0.907)		
Obs.	332	209	Obs.	1,540
R^2	0.324	0.413	R^2	0.072
Time FE	Yes	Yes	Time FE	Yes

TABLE 10: **Determinants of Pension Funds' Demand**: The table reports OLS coefficients, with time-clustered standard errors in parentheses, from panel regressions of pension funds' demand for equity duration (Panel A) and market beta (Panel B) on explanatory variables. For corporate defined benefit pension plans, in columns 1 and 2, the regressors are respectively a fund's funding ratio and period loss. For public defined benefit pension plans, in column 3, the regressor is the percentage of retired pension plan members (%Retired). The sample period is 1999Q1–2023Q4, and the demand coefficients for each institution-quarter observation in the FactSet database are obtained following the procedure described in Section 2.3. One, two, and three asterisks indicate significance at the 10%, 5%, and 1% level, respectively.

Government Accounting Standards Board guidelines, public pension funds can base their liability discount rates on the expected return of their assets, unlike private pension funds, which must use a combination of upper-medium and high-grade long-term corporate bond yields (Rauh, 2006; Brown and Wilcox, 2009).

Following Andonov et al. (2017), we use the percentage of retired pension plan members – %Retired, which they refer to as "fund maturity" – as a determinant of portfolio decisions.[17]

[17] We thank Aleks Andonov for providing us with a crosswalk to collapse the PPD data (https://publicplansdata.org/) from the pension plan level to the pension fund (retirement system) level.

Funds with a higher proportion of retired members face greater risks of future underfunding, which, under the regulatory incentives hypothesis, should increase their incentives to maintain a high liability discount rate.

Our panel regression analysis supports this prediction: as the percentage of retired members increases, public DB funds shift toward lower-duration, riskier stocks. This increased risk-taking allows constrained public pension funds to justify higher discount rates (and report a better funding status). The coefficient on equity duration is negative at −0.146 and highly statistically significant, implying that a 10 percentage points increase in the percentage of retired members in U.S. public pension funds is associated with a reduction in the demand for equity duration in the next quarter by 0.0146. Interestingly, the effect is muted for demand for beta, suggesting that the adjustment takes place only in the hedging component of demand.

5.2 Endogeneity concerns and alternative measures of duration

Our baseline results rely on the measure of duration proposed by Gonçalves (2021b), computed as:

$$Dur_t(n) = \sum_{h=1}^{\infty} \frac{E_t\left[PO_{t+h}(n)\right] e^{-h \cdot dr_t(n)}}{ME_t(n)} \times h \tag{6}$$

where dr is the discount rate, PO is the cash flows (proxied by firms' payout) to equity investors and ME is the firm's market equity.

The definition in Eq. (6) might raise the concern that correlated demand shocks (e.g., high investors' sentiment) that enter latent demand may impact duration through their effect on market equity, thereby making duration an endogenous regressor. We address this concern by instrumenting the equity duration with a measure that employs the counterfactual market equity capitalization from variation in investment mandates in Eq. (6). We denote with $\widehat{dur}_t(n)$ this duration measure purged by correlated demand shocks. As a consequence, our identifying assumption becomes now: $\mathbb{E}[\epsilon_{i,t}(n) | \widehat{me}_{i,t}(n), \widehat{dur}_t(n), \beta_{mkt,t}(n)] = 1$.

In Table 8, we show that our conclusions are robust to endogeneity concerns over equity duration for intermediaries with 2,000 holdings within four quarters.[18] The positive coefficients on η_{t-1} confirm the role of intermediaries' constraints on their demand for long-term cash flows. For comparison, in Table 8 we report estimates on the same sample of intermediaries when assuming *dur* as exogenous characteristic (i.e., without instrumenting it). As we can see, the two sets of coefficients are very similar in magnitude and statistical significance. These findings reassure us that our estimates are not unduly biased.

We also assess the sensitivity of our results to two key modifications of Gonçalves (2021b)'s approach. First, we change the discounting method by applying a constant 12% discount rate across firms, in order to eliminate discount rate variation and isolate the timing of cash flows. Second, we demean all state variables at the firm level before estimating the VAR, akin to including firm fixed effects. This procedure removes cross-sectional differences and ensures that the resulting dynamics reflect only the time-series behavior of cash flows, rather than persistent cross-sectional variation in their levels. Further details are provided in Appendix C.

Panel A of Appendix Table A.6 shows a consistently positive – though somewhat weaker – relationship between lagged equity capital ratios and demand for the timing component of cash flows across eight specifications. The effect is strongest for banks and pension funds (columns 3, 4, 7, and 8). While the link with timing is evident, the explanatory power is lower (e.g., the R^2 for banks is about 12%) compared to that for duration demand (approximately 27%). This suggests that risk-bearing capacity is more closely tied to duration demand than to timing demand, consistent with a reinvestment risk interpretation. Turning to demand for market beta, Panel B of the table reveals a negative relationship with lagged equity capital ratios, which is significant for all investor types except banks. The explanatory power remains strong for insurance firms and pension funds but drops for primary dealers

[18] The second-stage demand-system estimation for intermediaries with fewer than 2,000 holdings only shrinks coefficients on exogenous characteristics. Since in this section we treat equity duration as an endogenous characteristic, we can estimate the demand system only for intermediaries with at least 2,000 holdings within a four-quarter window.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Panel A: Demand for equity duration, instrumented							
	Primary dealers		Banks		Insurances		Pension funds	
η_{t-1}	0.101**	0.101**	0.168***	0.165***	0.135***	0.135***	0.026	0.027
	(0.040)	(0.039)	(0.028)	(0.029)	(0.039)	(0.039)	(0.024)	(0.024)
$FinUncert_{t-1}$		0.012		0.027		-0.021		0.052**
		(0.025)		(0.023)		(0.048)		(0.021)
Obs.	172	172	172	172	172	172	172	172
R^2	0.118	0.120	0.369	0.378	0.122	0.125	0.018	0.088
	Panel B: Demand for equity duration, non-instrumented							
η_{t-1}	0.091**	0.092**	0.158***	0.153***	0.125***	0.124***	0.014	0.015
	(0.039)	(0.039)	(0.033)	(0.034)	(0.038)	(0.037)	(0.024)	(0.023)
$FinUncert_{t-1}$		0.017		0.058*		-0.022		0.048**
		(0.020)		(0.030)		(0.046)		(0.021)
Obs.	172	172	172	172	172	172	172	172
R^2	0.124	0.128	0.295	0.334	0.118	0.122	0.005	0.063

TABLE 11: **Institutional Demand and Capital Ratios, Instrumented Duration**: The table reports OLS coefficients, with standard errors in parentheses, from regressions of the average AUM-weighted intermediaries' demand for equity duration (instrumented in Panel A, and non instrumented in Panel B) on the lagged type-specific capital ratio η. In columns 2, 4, 6, and 8, we additionally include the financial uncertainty index of Jurado et al. (2015). All explanatory variables are lagged by one quarter and standardized. The sample period is 1980Q1–2023Q4 for intermediaries with more than 2'000 holdings, and the demand coefficients for each institution-quarter observation in the Thomson Reuters database are obtained following the procedure described in Section 2.3. One, two, and three asterisks indicate significance at the 10%, 5%, and 1% level, respectively.

compared to Table 5.

Finally, we also consider alternative measures of duration, in particular the one proposed by Weber (2018). Weber's duration measure holds the discount rate constant across firms, forecasts cash flows over a finite horizon of fifteen years, and uses observed prices to infer post-horizon cash flows. While the timing of cash flow distributions shows the expected relationship with the measure in univariate sorts, we find it less effective once we double-sort by firm size; see Appendix C for details. Given that our demand system controls for market equity, we consider the measure by Gonçalves (2021b) to be more appropriate for our purposes.

6 Concluding Remarks

Stocks with distant payoffs offer a natural hedge against fluctuations in investment opportunities, giving rise to a differential in average returns between low- and high-duration stocks that is both sizeable and captures other dimensions of risk. Yet, the questions of *which* investors demand such hedging properties, and how their behavior shapes the time-series and cross-sectional dynamics of the duration premium, remain largely unanswered. Our study takes a step in this direction by examining the drivers and asset pricing implications of shifts in demand for equity duration by financial institutions through their equity holdings.

We show that institutions with higher capital ratios systematically tilt toward long-duration equities, while constrained intermediaries reduce their exposure to long-term claims and become more sensitive to short-term shocks. These patterns align with an ICAPM framework, where hedging demand declines with risk aversion and risk-bearing capacity. Using counterfactual experiments, we quantify the impact of shifts in intermediaries' risk-bearing capacity on equilibrium prices and expected returns. Positive shocks to duration demand steepen the cross-sectional return curve, lowering prices and increasing expected returns for short-duration stocks relative to long-duration ones. Negative shocks have the opposite effect, reducing the duration premium. Notably, shocks at the holding-company level produce disproportionately larger effects than those at the subsidiary level, reflecting the nonlinear transmission of balance sheet constraints to market prices.

Overall, our study highlights that heterogeneity across and within intermediary types has significant consequences for asset pricing. Variations in demand for long-term claims – whether driven by capital shocks, regulatory changes, or institutional constraints – directly affect stock prices, expected returns, and the cross-sectional distribution of the duration premium. These findings underscore the central role of intermediaries in shaping the risk-return tradeoff across equities of different durations.

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Appendix

A Data and Variable Construction

A.1 Classification of Financial Intermediaries

We focus our analysis on the following four institutional types: primary dealers, banks, insurance companies, and pension funds.

Primary dealers are identified as in He et al. (2017) using the historical list available on the New York Fed's website, which outlines the institutions serving as counterparties in its implementation of monetary policy. For the Thomson Reuters database, we hand-match dealers to portfolio holdings from their holding companies' manager number (mgrno). We match 88% (99/112) of primary dealers appearing at least once in the New York Fed's list of primary dealers between 1980 and 2023. These 99 matches correspond to 63 unique managers in our Thomson Reuters data. This is because some primary dealers are reported multiple times in the New York Fed's list due to, for instance, mergers or acquisitions.[19] We manually verified that the 13 primary dealers we cannot match to 13F managers are likely Treasury dealers only, such as Aubrey G. Lanston & Co Inc. For FactSet, the hand-matching is performed via an institution's `factset_entity_id`. We identify 94% (34/36) of primary dealers appearing at least once in the primary dealers list during 1999-2023.

For Thomson Reuters, we assign the other institutional types based on their type code. We then apply manual corrections as explained in Appendix D of Koijen and Yogo (2019). We also follow their approach in classifying investment advisors, mutual funds, and the residual household sector, whose demand functions are needed to compute the equilibrium counterfactual prices in Section 4.

For FactSet, we use the `entity_sub_type` variable (hereafter `est`) to categorize institutional investors. Since this variable provides more granular classifications than the institutional types in Thomson Reuters, we regroup them as follows. Companies classified as investment banks, brokers, bank investment divisions, or subsidiary branches are grouped under Banks. Insurance companies are identified by an `est` value of "IN" (insurance). These are further divided into Property & Casualty (P&C), Life Insurance (LI), or multi-line insurers based on the types of products they offer. We create a residual group for a select few insurance companies that are headquartered outside the U.S., and for which the classification is ambiguous. Pension funds are identified by an `est` value of "PF" (pension funds), and are further categorized into private defined benefits (e.g., ExxonMobil's defined benefit pension plan), public defined benefits (e.g., California State Teachers' Retirement System), or others (e.g., defined contribution plans) based on the company entity structure reported in FactSet. Additionally, we cross-check institutions against the list of the top 300 pension funds worldwide and reclassify as pension funds 17 companies that were originally labeled as "Corporation," "Foundation," or "Others" in FactSet.

[19] For example, Warburg Dillon Read & Co. was acquired by Swiss Bank Corporation in 1997. In the NY Fed's list, Warburg Dillon Read & Co. appears as primary dealer from June 3, 1996 to September 2, 1997, and SBC Warburg Dillon Read & Co. appears as primary dealer from September 3, 1997 to June 28, 1998. These two entries are both matched to a single 13F manager (mgrno=82550).

A.2 Equity Duration

Gonçalves (2021b) proposes a measure of equity duration based on a VAR(1) model with twelve firm characteristics useful to forecast payouts to shareholders. The characteristics are obtained from CRSP/Compustat and summarized below:

– book to market: $\log(BE_t/ME_t)$

– payout yield: $\log(1 + PO_t/ME_t)$

– sales yield: $\log(REVT_t/ME_t)$

– book equity growth: $\log(BE_t/BE_{t-1})$

– asset growth: $\log(AT_t/AT_{t-1})$

– sales growth: $\log(REVT_t/REVT_{t-1})$

– clean surplus profitability: $\log(1 + CSE_t/BEt - 1)$

– return on equity: $\log(1 + \frac{E_t}{0.5BE_t + 0.5BE_{t-1}})$

– gross profitability: $\log(1 + \frac{GP_t}{0.5AT_t + 0.5AT_{t-1}})$

– market leverage: $B_t/(ME_t + B_t)$

– book leverage: B_t/AT_t

– cash holdings: C_t/AT_t.

Here, PO_t are total payouts; $REVT_t$ are total revenues; AT_t are total assets; CSE_t are clean surplus earnings $(PO_t + \Delta BE_t)$; E_t is income before extraordinary items (ib); GP_t is gross profits ($revt - cogs$); B_t is total book debt ($dltt + dlc$); and C_t denotes cash plus short-term investment (che).

In the estimation of the VAR parameters, we only include companies with at least two years of data available to minimize backfiling concerns (see Fama and French, 2015) and exclude microcaps, i.e., firms with market equity below the 20% NYSE quantile based on NYSE breakpoint in a given quarter.[20] We refer to Gonçalves (2021b, Section 2) for further details on data cleaning.

The main assumptions underlying the duration measure are: (i) total stocks' payouts to investors are given by dividends + equity repurchases - issuances, (ii) firms' log profitability and growth evolve linearly (as in Vuolteenaho, 2002 and Campbell et al., 2009). We estimate the VAR parameters equation by equation from Fama-MacBeth regressions weighting each cross-section by the number of firms in that cross-section, and using an expanding window approach employing data from 1965:1 to the quarter t for which we estimate equity duration. Crucially, as in Gonçalves (2021b), the VAR parameters for the equity duration measure are the same for all firms in each cross-section. Hence, all cross-sectional variation in equity duration in quarter t only comes from heterogeneity in the twelve firm characteristics at that time. This is relevant for our purpose, as estimating different VAR parameters for each firm could introduce heterogeneity in measurement errors and generate a misleading ranking of equity duration. Using the same VAR parameters for all firms prevents us from this risk.

[20]The exclusion of microcaps is solely for computing the VAR parameters. Following Gonçalves (2021b), we also calculate their equity duration.

We obtain the discount rate using a root-finding algorithm that equates current market capitalization to the present discounted sum of future expected payouts from the VAR model, as in Gonçalves (2021b). Despite the differences in data frequency and sample period, the resulting distribution of firm-quarter discount rates matches quite closely that in Figure IA.1 of its Internet Appendix.

Figure A.1 plots the cumulative percentage of market equity that is paid to investors (dividends plus repurchases less issuances) from t to $t + 10$ for the top and bottom duration deciles. Panel A sorts firms into deciles based on duration. As expected, stocks in the lowest decile pay off about one-third of their market equity within ten years, compared to the nearly zero fraction paid by their high-duration counterparts. In Panel B, we perform a conditional (within-quarter) double sort: we first split firms above and below the median market capitalization using NYSE breakpoints, and then construct duration deciles within each group. While the difference in payouts across duration deciles becomes less pronounced, it remains substantial. This evidence reassures us that the measure properly captures the timing of equity cash flows.

A.3 Target-Date Funds

Target-date funds typically disclose their target date in their name. For instance, the fun "Schwab Capital Trust Target 2045" has a pre-specified target date in 2045. We exploit this pattern to identify target-date funds in the CRSP Mutual Funds survivorship-bias-free database. In particular, from January 2010 to December 2023, we identify a target date for each fund isolating the "20XX" term in their name. Then, we validate the target date keeping only mutual funds expiring between 2011 and 2099. Following this procedure, we obtain 21,693 fund-quarter observation between 2010 and 2023 and 1,271 unique funds. The farthest target date is in 2070 (4 funds). For each fund-quarter observation, we obtain time to maturity as the distance between the target and the current quarter's year.

A.4 Aggregate Measures of Financial Constraints

In the time-series analysis, we use the aggregate capital ratio, η, as main proxy for an institutional type's risk-bearing capacity. An institution's capital ratio is constructed following He et al. (2017) as the ratio of its market equity to the market value of its assets. The market value of assets is computed as the sum of market equity and book debt, where book debt is defined as total assets (*at*) minus common equity (*ceq*) from Compustat. The aggregate capital ratio is the value-weighted average of individual ratios, with weights equal to a company's market-valued asset share.

For primary dealers, we use directly the series from He et al. (2017). The alternative of computing the measure based on our sample of primary dealers yields quantitatively similar results. For the other types, we compute the aggregate capital ratio across all companies in the corresponding Fama and French 49 industry classification of financials. Specifically, for banks, we select companies with SIC code in ranges 6020-6036, 6040-6062, 6080-6082, 6090-6099 and 6120-6129. For insurance companies, the relevant SIC code ranges are 6310-6331, 6350-6351, 6360-6361, 6390-6399 and 6300, while pension funds are identified by SIC codes between 6370 and 6379. Using the pension funds filter on Compustat during 1980-2023 restricts the aggregate pension fund sector to only few companies per quarter, which renders the resulting series rather noisy. For this reason, we construct a single capital ratio for the combined group of insurance companies and pension funds.

B Estimation of Demand System

In Equation (5), log-market equity is endogenous. For instance, high sentiment around a particular stock would induce high latent demand ϵ and trigger an increase in market valuation (me), thereby violating the OLS orthogonality condition. To solve this endogeneity issue, Koijen and Yogo (2019) propose an instrumental variable (IV) based on investors' mandates. The idea behind the IV is that holdings that reflect investment mandates are exogenous to changes in stock market equity. Assuming investors allocate an equal weight to each stock in their investment mandate, the overall hypothetical wealth invested in a given stock would serve as valid instrument for its market equity. Concretely, the instrument for stock j and investor i is constructed as:

$$\widehat{ME}_i(j) = \sum_{\substack{l=1 \\ l \neq i, l \neq HH}}^{I} AUM_l \cdot \frac{\mathbb{I}_l(j)}{1 + \sum_{n=1}^{N} \mathbb{I}_l(n)}, \tag{B.1}$$

where $\mathbb{I}_l(j)$ is an indicator function equal to one if investor l has stock j in its investment universe. Following Koijen et al. (2024), we exclude households direct holdings in computing the IV. This choice reduces potential endogenity coming from the households' investment universe, and significantly increases the first-stage t-statistics.

Defying the set of stocks in an investor's investment universe is an empirical challenge. Koijen and Yogo (2019) observe that the set of stocks held by institutional is highly persistent. For instance, on average, among all stocks held by an institutional at time t, 95% of them were also in the investor's portfolio 11 quarters earlier. This suggests as investors' set of choices for their investment decision (i.e., investment universe) is somewhat limited and persistent over time. Exploiting these facts allows us to define an investors' universe at time t as the set of stocks ever held between t and $t - Q$ quarters. We follow Koijen et al. (2024) using $Q = 11$ quarters to define the instrument for market equity. Increasing Q reduces endogeneity concerns around the instrument, as higher lags allow better identification of an investor's universe. However, adding more lags requires a longer sample of data to construct the first observation of the instrument. In Table A.3, we verify that the persistence in holdings is present also in our sample of institutions and time period.

The demand system in Eq. (5) requires defining an outside asset. Since 13(f) filings only contain positions in certain securities (primarily stocks and a few derivatives), we do not observe cash or other non-equity assets. Therefore, we classify as outside assets all stocks for which we cannot construct the measure of equity duration. This simplifying assumption may be a concern if investors exhibit different demand coefficients for the securities we include in the outside asset. For instance, an investor could hold only high-duration stocks within the subset we label as outside assets. While we cannot directly rule out this concern – since, by definition, we do not observe the duration of stocks in the outside asset – we amke the following two observations. First, in our definition, the outside asset primarily consists of micro-cap stocks and those missing standard Compustat data items, such as total assets (at). Given this, we find it reasonable to assume that institutional investors in our sample do not specifically seek duration exposure through such micro-cap or relatively obscure stocks. Second, we examine whether certain institutions in our sample exhibit particularly high exposure to stocks in the outside asset. However, we find no significant differences in investor weights associated with stocks for which we do not observe duration. The weight in the outside asset remains stable both over time and across institutions.

We estimate the demand system by GMM, pooling portfolio holdings across four quarters; i.e., the demand coefficients for quarter t use holdings from $t-3$ to t. For institutions with more than 2,000 holdings over this four-quarter period, we estimate manager-specific demand coefficients and quarter-specific fixed effects. For other institutions, we rely on the two-step GMM procedure of Koijen et al. (2024). In the first step, we pool institutions of the same type by AUM and form bins with at least 2,000 combined holdings. We then estimate the demand system and obtain bin-specific coefficients. In the second step, we use a shrinkage approach to obtain manager-specific demand coefficients. The moment condition for the shrinkage estimator is:

$$\mathbb{E}\left[\left(\widehat{\delta}_{i,t}(n)\exp\left(-\boldsymbol{\alpha}'_i\boldsymbol{e}_t - \theta_{mkt,i,t}\beta_{mkt,t}(n) - \theta_{dur,i,t}dur_t(n)\right) - 1\right)\begin{pmatrix}\boldsymbol{e}_t\\ \beta_{mkt,t}(n)\\ dur_t(n)\end{pmatrix}\right] - \frac{\lambda}{|\mathcal{N}_i|^{\xi}}\begin{pmatrix}\mathbf{0}\\ \boldsymbol{\theta}_i - \widehat{\boldsymbol{\theta}}_1\end{pmatrix} = \mathbf{0}.$$

where $\widehat{\delta}_{i,t}(n) = \frac{w_{i,t}(n)}{w_{i,t}(0)}\exp(-\theta_{me,i,t}me_t(n))$, $\mathcal{N}_i$ is the number of stocks in investor i's investment universe, and $\boldsymbol{e}_t$ are quarter fixed effects. The shrinkage target, $\widehat{\boldsymbol{\theta}}_1$, is the bin-specific estimate from the first step. We select the shrinkage hyperparameters by cross-validation. Specifically, we randomly split the sample in half in each quarter. We then estimate the demand coefficients using only one half of the sample and select the parameters (λ, ξ) that minimize the median squared error in the second half (test sample). The resulting penalty parameters are $\lambda = 120$ and $\xi = 1.07$, which are similar to those reported by Koijen et al. (2024). In the estimation, we also impose the restriction $\theta_{me,i,t} < 1$ for all investors. Koijen and Yogo (2019) show that this condition is sufficient to ensure that the system of market-clearing equations has a unique solution in prices.

C Alternative measures of duration

We use an alternative measure of duration, building on Gonçalves but modifying two key components: the discounting method and the VAR estimation approach. Specifically, we isolate cash-flow timing using a uniform discount rate and apply firm-level demeaned VARs to isolate time-series dynamics. Below, we explain the motivations for these adjustments and how they address potential limitations in the original approach.

First, in Gonçalves (2021b), the discount rate for each stock is calibrated so that the present value of forecasted cash flows equals the observed market price. This implies that firm-specific discount rates enter directly into the duration computation. To remove this source of variation, we replace the stock-specific discount rates with a constant discount rate of 12% across all firms. Correspondingly, instead of using observed market prices we calculate the implied price using the forecasted cash flows discounted at this uniform rate.

Second, Gonçalves (2021b) uses pooled VAR estimations that do not control for unconditional cross-sectional differences in cash-flow levels across firms. This can confound time-series persistence with persistent cross-sectional differences. For instance, growth stocks, which are characterized by high market-to-book ratios, typically exhibit higher earnings-to-book equity ratios (profitability) than value stocks. This cross-sectional relation persists over time. Duration, by construction, aims to capture the dynamics (i.e., early vs. late arrival of cash flows) rather than levels (i.e., high vs. low cash flows). If one mistakenly attributes persistent cross-sectional levels to dynamic cash-flow timing, this will mechanically overstate future cash flows for firms with high values of variables correlated with cash-flow levels, such as the market-to-book ratio. If these same variables are also correlated with discount rates, this mis-attribution can induce a spurious negative relationship between returns and estimated duration. To address this issue, we follow Chen et al. (2013) and demean all state variables at the firm level before estimating the VAR (akin to firm fixed effects). This approach strips out cross-sectional differences and ensures that the estimated dynamics reflect only time-series variation. Our resulting duration measure thus avoids the pitfalls of pooled VAR estimation and more cleanly isolates the timing component of expected cash flows.

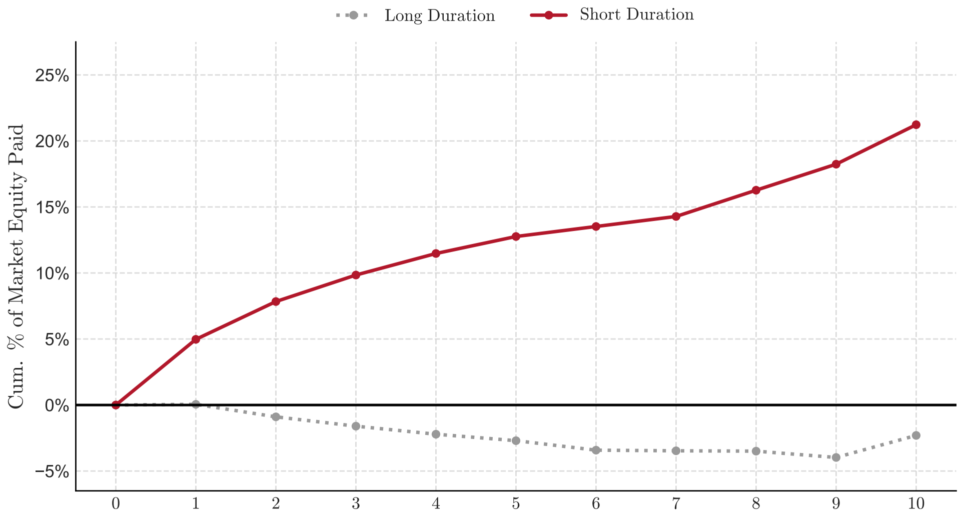
Alternative measures of duration have been proposed in the literature. Notably, Dechow et al. (2004) first extended the concept of duration to equities, and Weber (2018) later adapted this framework to study the cross-section of duration and stock returns. These measures differ from the approach in Gonçalves (2021b) along several dimensions.

Both Dechow et al. (2004) and Weber (2018) forecast firm-level cash flows over finite horizons—10 years in Dechow et al. and 15 years in Weber. After this forecasting window, the present value of projected cash flows is subtracted from the market price (assumed to equal the present value of all future cash flows), and the residual is assumed to be paid out as a level perpetuity. This approach effectively uses observed prices to infer post-horizon cash flows and attributes all variation in price to variation in expected cash flows rather than in discount rates. Although the discount rate is held constant across firms, this approach relies on observed prices to impute residual cash flows.

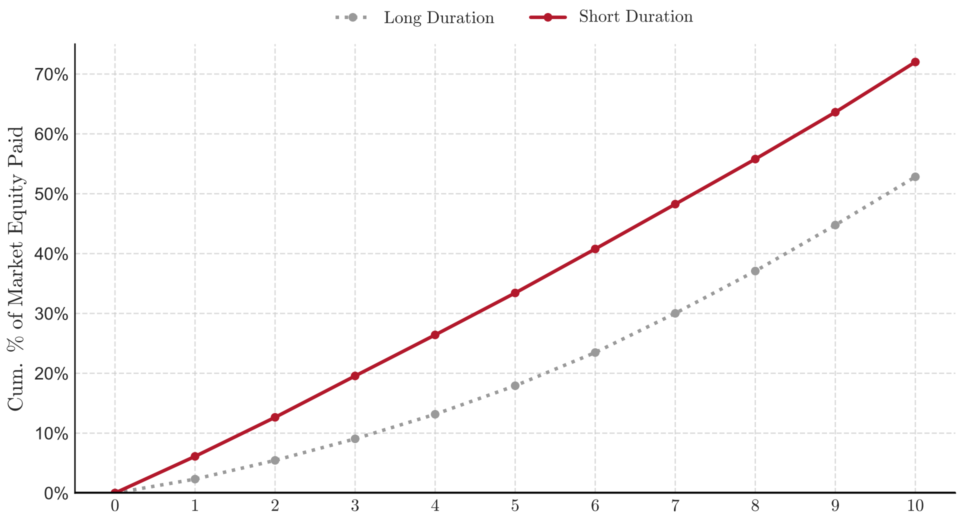
In contrast, Gonçalves (2021b) builds on the logic of Dechow et al. (2004) but extends the forecast horizon to 1,000 years using a VAR-based forecasting model. He also calibrates firm-specific discount rates such that the present value of forecasted cash flows matches the observed market price. This design allows duration to reflect both cash flow timing and discount rate variation in an internally consistent manner.

To compare these measures empirically, Figure A.5 mirrors the format of Panel A of Figure A.1, but uses the Weber duration instead. This figure plots the cumulative payouts of portfolios sorted by duration over time, expressed as a percentage of market equity at the time of portfolio formation. Intuitively, firms with short-duration equity – i.e., early-timing portfolios – should deliver more payouts in the near term, while long-duration equity delivers more in the distant future. This pattern holds for both the Weber and Gonçalves measures, though the spread is noticeably larger for Gonçalves (2021b).

It is important to note that our demand system controls for market equity, effectively making our empirical strategy closer to a double sort on size and duration. Gonçalves (2021b) shows that the short-duration premium is robust to size controls (see his Table 3). To evaluate this further, in Panel B of Figure A.5 we replicate the corresponding panel of Figure A.1 using the Weber measure with a double sort on duration and market capitalization. While the qualitative pattern remains—long-duration firms pay out later—the spread in payouts is narrower, and the two curves intersect around year 8. This contrasts with the Gonçalves (2021b) measure, where the separation in payout timing persists beyond year 10.



(A) Single-sorting by duration



(B) Double-sorting by duration and market capitalization.

FIGURE A.1: **Cumulative share of market equity paid within 10 years, Gonçalves (2021b) measure**. The figure displays the cumulative fraction of total realized payouts (dividends plus equity repurchases less issuances) over ME_t for stocks in the top duration decile (solid line) and bottom duration decile (dotted line) using the measure from Gonçalves (2021b) as of December of each year t. Panel A sorts firms into deciles based on duration. In Panel B, we perform a conditional (within-quarter) double sort: we first split firms above and below the median market capitalization using NYSE breakpoints, and then construct duration deciles within each group. For all stocks in the top and bottom deciles we compute the cumulative ratio of PO_{t+h}/ME_t for $h = 1, \ldots, 10$ and plot the average over 1980Q1–2023Q4 for the top and bottom duration deciles.

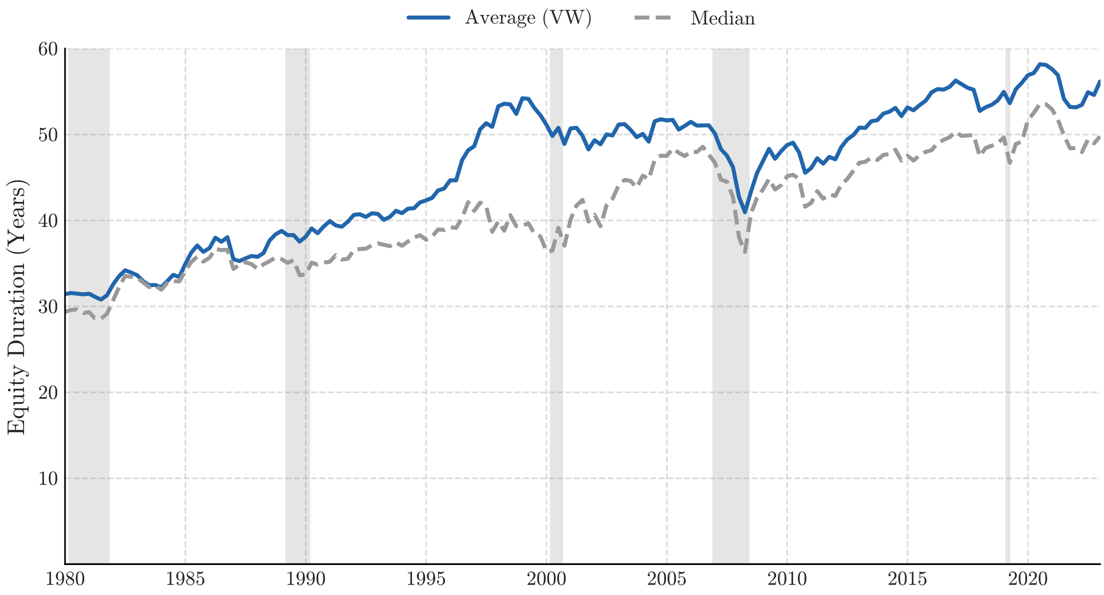


FIGURE A.2: **Time series of equity duration**. The figure displays the quarterly time series of value-weighted average (solid line) and median (dotted line) equity duration over 1980Q1–2023Q4. Shaded regions are recessions as defined by the National Bureau of Economic Research.

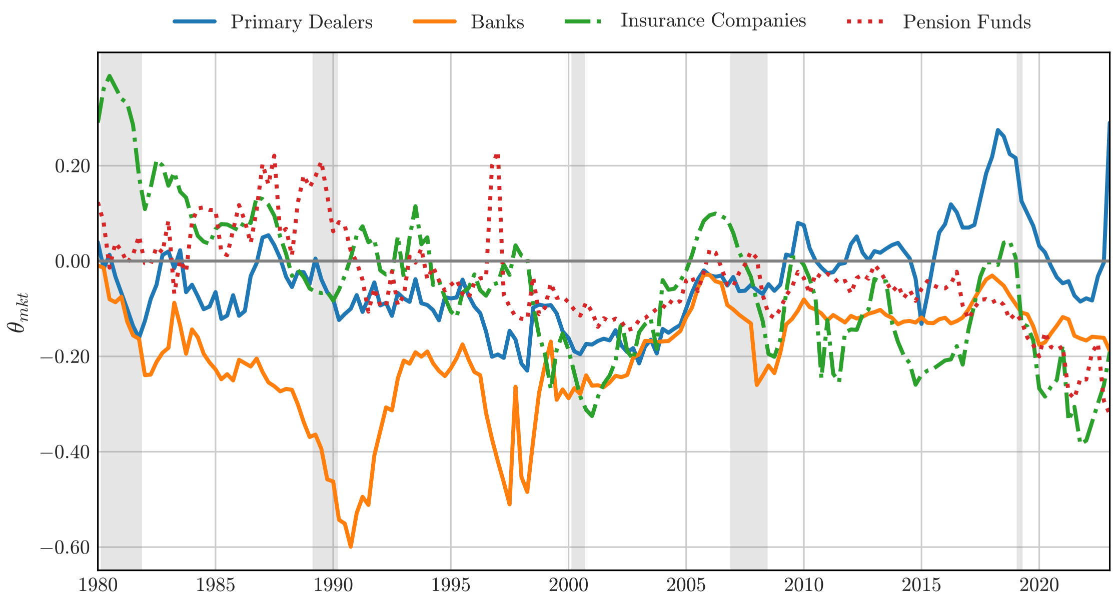


FIGURE A.3: **Investors' demand for market beta**. The figure displays the time series of AUM-weighted demand coefficients on equity beta (θ_{mkt}) by institutional type over 1999Q1–2023Q4. Shaded regions indicate NBER recessions.

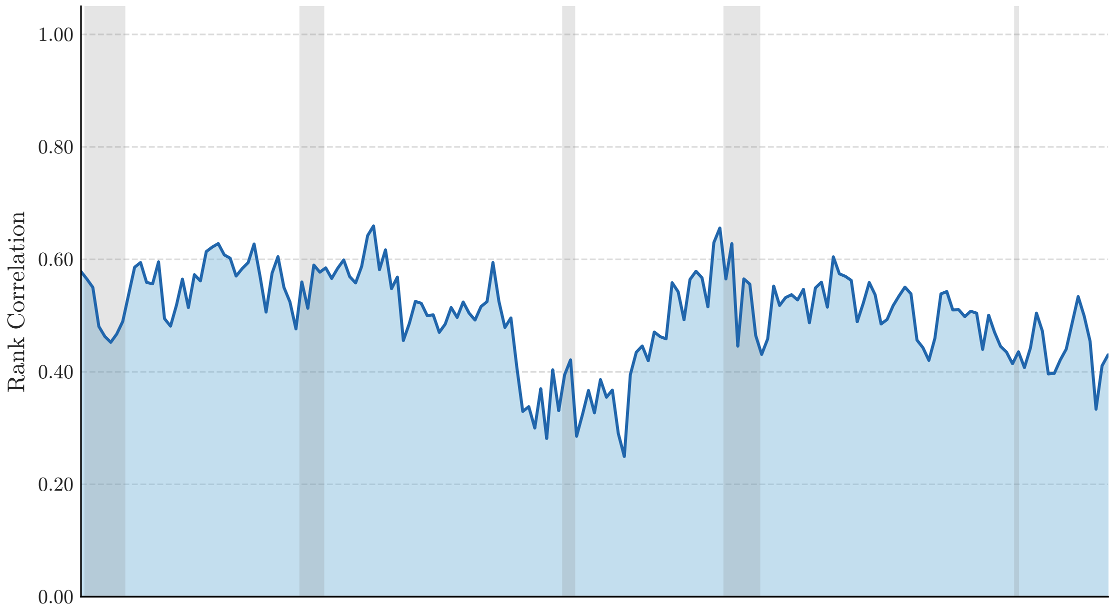
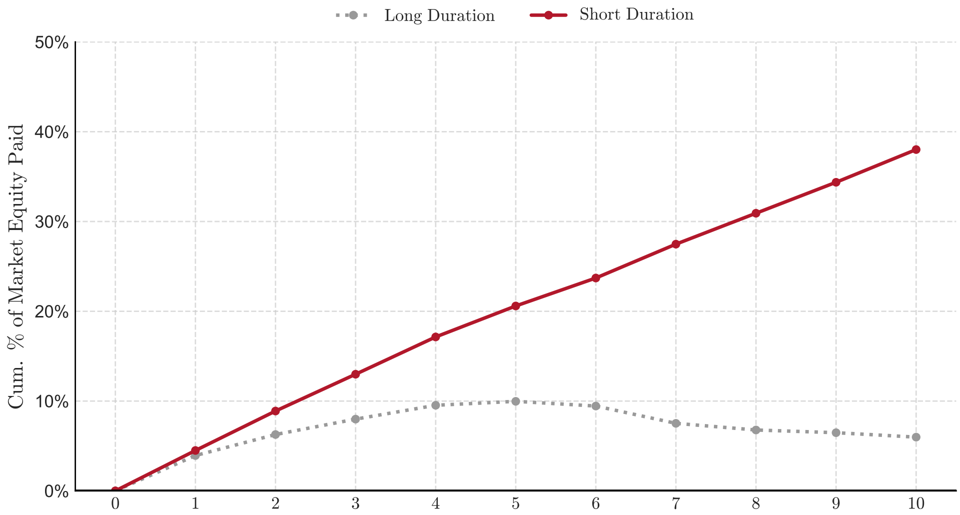
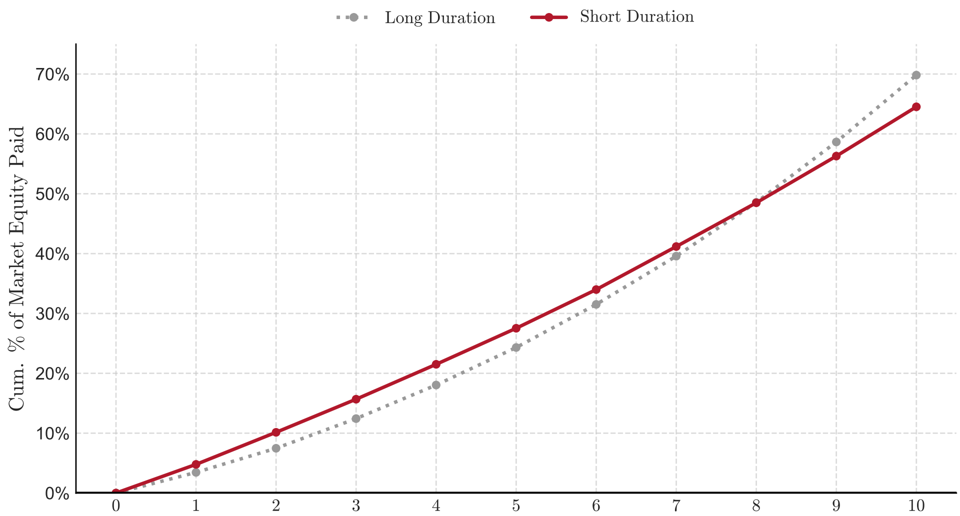


FIGURE A.4: **Rank correlation between demand for duration and weighted average portfolio duration.** The figure displays the time series of the cross-sectional rank correlation between an institution's demand for duration, $\theta_{dur,i,t}$, estimated from the demand system in Eq. (5), and its weighted average portfolio equity duration, $\sum_{n=1}^{N} w_{i,t}(n) \cdot Dur_t(n)$, over 1980Q1–2023Q4. Shaded regions indicate NBER recessions.



(A) Single-sorting by duration



(B) Double-sorting by duration and market capitalization.

FIGURE A.5: **Cumulative share of market equity paid within 10 years, Weber (2018) measure.** The figure displays the cumulative fraction of total realized payouts (dividends plus equity repurchases less issuances) over ME_t for stocks in the top duration decile (solid line) and bottom duration decile (dotted line) using the measure from Gonçalves (2021b) as of December of each year t. Panel A sorts firms into deciles based on duration. In Panel B, we perform a conditional (within-quarter) double sort: we first split firms above and below the median market capitalization using NYSE breakpoints, and then construct duration deciles within each group. For all stocks in the top and bottom deciles we compute the cumulative ratio of PO_{t+h}/ME_t for $h = 1, \dots, 10$ and plot the average over 1980Q1–2023Q4 for the top and bottom duration deciles.

	$w_i(n)$					w_0	
	mean	std	p10	p50	p90	mean	std
Primary Dealers	0.04	0.34	0.00	0.00	0.05	29.84	8.53
Banks	0.10	0.72	0.00	0.00	0.16	29.08	11.37
Insurances	0.18	1.15	0.00	0.01	0.39	27.30	9.71
Pension Funds	0.10	0.92	0.00	0.01	0.17	27.72	9.83

Table A.1: **Portfolio weights, FactSet data**: The table presents descriptive statistics of portfolio weights, expressed in percentage, by institutional type for the FactSet database over 1999Q1–2023Q4. The statistics are computed for weights of all stocks in the investment universe $w_i(n)$ and of the outside asset w_0 in each quarter, and then averaged across all quarters in the period.

	Pooled (EW)			VW	
	mean	sd	median	mean	sd
Market beta	1.12	0.75	1.05	1.01	0.49
Duration	42.99	24.04	38.22	51.33	14.42
log (Duration)	3.66	0.42	3.64	3.90	0.27

Table A.2: **Summary statistics of stock characteristics**: The table reports summary statistics of market beta (β_{mkt}) and equity duration (in levels and logs) over 1980Q1–2023Q4. We compute mean, standard deviation, and median across the pooled panel of stock-quarter observations, and mean and standard-deviation when value-weighting stocks by market capitalization.

AUM decile	Quarters Lag 1	2	3	4	5	6	7	8	9	10	11
1	89	91	92	93	94	94	95	95	96	96	96
2	91	92	93	94	94	95	95	96	96	96	97
3	91	92	93	94	95	95	96	96	96	96	97
4	90	92	92	93	93	94	95	95	95	96	96
5	91	92	93	94	94	94	94	95	95	96	96
6	91	92	93	94	94	95	95	95	95	96	96
7	91	93	93	94	94	95	95	95	95	96	96
8	91	92	93	94	94	95	95	95	96	96	96
9	92	93	93	94	94	95	95	95	96	96	96
10	91	92	93	94	94	95	95	95	95	96	96

Table A.3: **Persistence of institutional holdings**: The table reports average portfolio holdings' persistence for institutional investors grouped by AUM deciles. For each quarter over 1980Q1–2023Q4, we compute the percentage of stocks held in the current quarter that were ever held in the previous one to 11 quarters, and average this figure over the sample. For instance, on average, 96% of stocks held in the current quarter by institutionals in the fourth decile of AUM were held at least once in the past 11 quarters.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
TTM	0.036***		0.036***					
	(0.007)		(0.006)					
Lag TTM		0.037***		0.037***				
		(0.004)		(0.004)				
Equity weight			-0.587				-0.602	
			(0.630)				(0.624)	
Lag Equity weight				0.231				0.189
				(0.216)				(0.195)
Mid TTM					0.121		0.265**	
					(0.094)		(0.121)	
High TTM					0.062		0.414***	
					(0.297)		(0.127)	
Lag Mid TTM						0.178***		0.168***
						(0.039)		(0.041)
Lag High TTM						0.356***		0.354***
						(0.055)		(0.059)
Obs.	460	401	460	401	618	489	460	401
R^2	0.763	0.885	0.765	0.886	0.622	0.847	0.748	0.870
Fund FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

TABLE A.4: **Demand for duration and target date funds' investment horizon**: The table reports OLS coefficients, with standard errors in parentheses, from regressions of target date funds' demand for equity duration on their investment horizon. In columns 1–4, the horizon is measured by a fund's time to maturity. In columns 5–8, we use a dummy specification for funds with medium and high time to maturity, defined as in Figure 3. Columns 3, 4, 7, and 8 additionally control for the fund's equity portfolio weight. The sample period is 2010Q1-2023Q4, and demand coefficients for each fund–quarter observation are obtained following the procedure in Section 2.3.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
				Panel A: Demand for equity duration				
	Primary dealers		Banks		Insurances		Pension funds	
η_{t-1}	-0.008	-0.000	0.088***	0.087***	0.031	0.029	0.066***	0.065***
	(0.028)	(0.018)	(0.019)	(0.015)	(0.032)	(0.032)	(0.021)	(0.020)
$FinUnc_{t-1}$		0.068***		0.034**		0.035		0.024
		(0.019)		(0.014)		(0.035)		(0.018)
Obs.	96	96	96	96	96	96	96	96
R^2	0.003	0.205	0.476	0.547	0.019	0.044	0.233	0.265
				Panel B: Demand for market beta				
	Primary dealers		Banks		Insurances		Pension funds	
η_{t-1}	-0.106***	-0.111***	-0.060***	-0.059***	-0.025	-0.025	-0.049***	-0.047***
	(0.017)	(0.017)	(0.010)	(0.009)	(0.020)	(0.020)	(0.012)	(0.008)
$FinUnc_{t-1}$		-0.037		-0.021**		0.001		-0.037***
		(0.029)		(0.008)		(0.020)		(0.006)
Obs.	96	96	96	96	96	96	96	96
R^2	0.309	0.345	0.495	0.558	0.031	0.031	0.436	0.683

TABLE A.5: **Institutional Demand and Capital Ratios, FactSet data**: The table reports OLS coefficients, with standard errors in parentheses, from regressions of the average AUM-weighted intermediaries' demand for equity duration (θ_{dur}, Panel A) and market beta (θ_{mkt}, Panel B) on the lagged type-specific capital ratio η. In columns 2, 4, 6, and 8, we additionally include the financial uncertainty index of Jurado et al. (2015). All explanatory variables are lagged by one quarter and standardized. The sample period is 1980Q1–2023Q4, and the demand coefficients for each institution-quarter observation in the FactSet database are obtained following the procedure described in Section 2.3.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
				Panel A: Demand for equity duration				
	Primary dealers		Banks		Insurances		Pension funds	
η_{t-1}	0.001	0.001	0.025***	0.026***	0.014	0.013	0.022**	0.022***
	(0.008)	(0.008)	(0.008)	(0.008)	(0.016)	(0.015)	(0.009)	(0.007)
$FinUnc_{t-1}$		-0.000		-0.003		-0.012		0.018*
		(0.009)		(0.010)		(0.017)		(0.009)
Obs.	172	172	172	172	172	172	172	172
R^2	0.001	0.001	0.122	0.123	0.011	0.020	0.078	0.129
				Panel B: Demand for market beta				
	Primary dealers		Banks		Insurances		Pension funds	
η_{t-1}	-0.016**	-0.016**	-0.017	-0.016	-0.049***	-0.050***	-0.040***	-0.040***
	(0.008)	(0.008)	(0.013)	(0.013)	(0.018)	(0.017)	(0.012)	(0.011)
$FinUnc_{t-1}$		0.005		-0.005		-0.031**		-0.017
		(0.011)		(0.011)		(0.015)		(0.011)
Obs.	172	172	172	172	172	172	172	172
R^2	0.031	0.034	0.024	0.026	0.111	0.156	0.150	0.177

TABLE A.6: **Institutional Demand and Capital Ratios**: This table replicates the analysis in Table 5, but replaces the preference for duration with a measure of preference for the timing of cash flows (a modified version of the Goncalves measure). The table reports OLS coefficients, with standard errors in parentheses, from regressions of the average AUM-weighted intermediaries' demand for equity duration (θ_{dur}, Panel A) and market beta (θ_{mkt}, Panel B) on the lagged type-specific capital ratio η. In columns 2, 4, 6, and 8, we additionally include the financial uncertainty index of Jurado et al. (2015). All explanatory variables are lagged by one quarter and standardized. The sample period is 1980Q1–2023Q4, and the demand coefficients for each institution-quarter observation in the Thomson Reuters database are obtained following the procedure described in Section 2.3.