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The Review of Financial Studies

The Return Expectations of Public Pension Funds

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The return expectations of public pension funds are positively related to cross-sectional differences in past performance. This positive relation operates through the expected risk premium, rather than the expected risk-free rate or inflation rate. Pension funds act on their beliefs and adjust their portfolio composition accordingly. Persistent investment skills, risk taking, efforts to reduce costly rebalancing, and fiscal incentives from unfunded liabilities cannot fully explain the reliance of expectations on past performance. The results are consistent with extrapolative expectations, since the dependence on past returns is greater when executives have personally experienced longer performance histories with the fund. (*JEL* G02, G11, G23, G28, H75, D83, D84)

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What do institutional investors believe about the expected returns of the asset classes in which they invest, and how do they set these expectations? Investor beliefs about expected returns are key parameters of portfolio choice models (Black and Litterman 1992; Pastor 2000; Ang, Ayala, and Goetzmann 2018). A growing literature on individual investors reflects the importance of past experience in determining forward-looking expectations and asset allocations (e.g., Benartzi 2001; Vissing-Jorgensen 2003;

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Malmendier and Nagel 2011; Greenwood and Shleifer 2014; Kuchler and Zafar 2019; Giglio et al. 2021). Studies of expectations usually focus only on a single asset class or macroeconomic variable, such as listed equities, housing, or inflation (e.g., Amromin and Sharpe 2014; Malmendier and Nagel 2016; Bailey et al. 2019). However, there has been little direct, large-sample evidence about the beliefs of institutional investors across a range of asset classes, and limited analysis of how those beliefs affect the allocation decisions of institutional investors in multiple asset classes.

In this paper, we provide the first evidence that institutional investors rely on past performance in setting future return expectations and that these expectations affect their target asset allocation policy. The U.S. Governmental Accounting Standards Board Statement 67 (GASB 67) requires U.S. public pension plans, a group of institutional investors that manage approximately $4 trillion of assets, to report long-term expected returns by asset class beginning in the 2014 fiscal year. This required disclosure separately reveals the pension fund's target portfolio composition and return expectations in individual asset classes such as public equity, fixed income, private equity, real assets, and hedge funds.

Our objective is to examine cross-sectional differences in return expectations among pension funds as function of cross-sectional differences in past returns. To remove aggregate time-series predictability on an asset class level (Brennan, Schwartz, and Lagnado 1997; Campbell and Viceira 1999; Barberis 2000), we control for time (year-reporting-month) fixed effects. We find that asset allocation weights explain only 27% of cross-sectional variation in portfolio expected returns, leaving a large role for variation across pension funds in beliefs about expected returns within the different asset classes. Cross-sectional variation in the average returns experienced in the past 10 years adds substantial explanatory power, both to expectations for the whole portfolio and to expectations for a given asset class. Specifically, each additional percentage point of past return relative to other pension funds raises the long-term portfolio expected return by 30 basis points, even after controlling for the percentage allocated to each risky asset class, the volatility of past returns, and fiscal incentives from unfunded liabilities.

The disclosures enable us to decompose the portfolio expected return into two components: the expected inflation rate and the expected real return, or alternatively the expected risk-free rate and the expected risk premium. We find that the positive effect of past returns on the real return assumption drives the positive relation with the overall portfolio expected return, as opposed to an effect of past returns on expected inflation or the expected risk-free rate. Pension funds with higher past performance expect a higher risk premium when they invest in risky assets, and they expect a higher risk premium in a given asset class when they have achieved stronger past performance in that asset class relative to other pension funds.

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We structure our analysis around four possible rational explanations for these findings. First, the positive relation between past returns and future expected returns might reflect persistent cross-sectional differences in investment skill among pension funds. Pension funds might rationally extrapolate past performance in asset classes where past cross-sectional outperformance predicts future outperformance, due to better skill or access to higher-quality external managers. We test this hypothesis by examining the persistence in the cross-section of realized returns, again controlling for year-reporting-month fixed effects to remove aggregate time-series predictability on an asset class level (e.g., predictability based on the dividend yield). Although there is evidence of one-year persistence in the cross-section of performance, this persistence disappears completely and even reverses when looking at horizons beyond one year. Carhart (1997) and Busse, Goyal, and Wahal (2010) also find that even when there is limited persistence on a single-year horizon, that persistence disappears when the analysis is extended to a two-to-three-year horizon.

Our investigations of individual asset classes provide further evidence against the persistent skills hypothesis as an explanation for our main results. For example, pension plans rely on past returns in public equity, which makes up around half of their portfolios, when formulating expected returns in this asset class. However, we find no persistence in the cross-section of pension fund performance in public equity. The extrapolation in public equity expectations also contradicts prior evidence that shows no long-run persistence in performance delivered by public equity asset managers and no ability by public pension plans to time the hiring and firing of asset managers (Goyal and Wahal 2008; Busse, Goyal, and Wahal 2010).

The positive short-term (one-year) persistence we find on an overall fund level is due to persistence within alternative assets, such as real assets and private equity. Genuine variation in skill in alternative asset investing (Cavagnaro et al. 2019) may explain this short-term persistence in reported alternative asset returns, as might less frequent fair valuation of alternative investments and smoothing of the reported returns (e.g., Geltner 1991; Getmansky, Lo, and Makarov 2004; Ang et al. 2018). Nevertheless, even within alternative assets there is no evidence of long-term persistence. These results indicate that the persistent skills hypothesis cannot explain our findings of extrapolation in the cross-section.

Second, the extrapolation of past returns when forming future expectations could capture higher risk taking in the past that will continue in the future. We include asset class weights as one set of risk controls, but these variables do not capture the possibility of differential risk taking within asset classes. We address this alternative hypothesis by adding to our baseline specification controls for the standard deviation of past performance, exposure to market risk, and exposure to Fama-French factors. These risk controls do not explain the positive relation between expected returns and past performance.

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Third, pension funds might adjust their expectations based on past performance to justify asset allocation decisions that would reduce costly rebalancing. Under this hypothesis, past performance mechanically affects actual asset class allocations. If pension funds want to reduce rebalancing costs, they might adjust their target asset allocations in the same direction that realized returns move the actual asset allocation, in which case the target asset allocation might also reflect the past returns. Pension plan return expectations would then not necessarily reflect their true expectations, and they would be linked to past returns simply to justify the adjustment of the target asset allocation policy toward the actual asset allocation policy. A key prediction of this hypothesis is that target asset allocation weights should converge toward actual asset allocation weights. When we estimate regression specifications to study how lagged deviations between target and actual asset allocations correlate with changes in actual and target asset allocations, respectively, we find some evidence for this effort to reduce costly rebalancing. However, we find significantly stronger evidence for a convergence of actual asset allocation to target allocation, which suggests that target weights reflect pension funds' true beliefs and are not merely justifications of mechanical changes in asset allocation.

These findings would still leave open the possibility that pension funds adjust target allocation weights ex ante in anticipation of future mechanical changes in actual allocation weights, rather than to reflect their true beliefs. To address this issue, we decompose the changes in the actual asset allocation into mechanical and residual changes. The reduce-costly-rebalancing hypotheses would predict that the mechanical changes in actual asset allocation explain how the actual asset allocation weights converge to the target weights. We find, however, that actual allocations converge toward target asset allocations through active residual changes in asset allocation, rather than through mechanical changes. This again suggests that target weights reflect pension funds' true beliefs, as pension funds implement active decisions to reach their targets.

Fourth, pension plans might extrapolate past returns when forming future expectations to reduce the amount of reported unfunded pension liabilities and fiscal pressure.[1] Incentives by fiscally stressed governments to maintain higher pension discount rates may induce them to state higher return expectations as justification (Brown and Wilcox 2009; Novy-Marx and Rauh 2011; Naughton, Petacchi, and Weber 2015; Andonov, Bauer, and Cremers 2017). We indeed

[1] GASB instructs pension plans to use an expected return as a discount rate to measure the present value of promised pension benefits, despite the conceptual mismatch of using an expected return on assets as a discount rate for a contractually prespecified, market-invariant stream of liability cash flows (Novy-Marx and Rauh 2011). Even though this regulation creates a link between expected returns and discount rates, we find that the expected return is significantly more variable than the pension discount rate. Moreover, the processes for setting the two assumptions also differ. The expected return informs the setting of the discount rate, but the pension plan board decides on the discount rate independently in consultation with its actuaries. In contrast, the expected return is prepared by pension plan investment staff with the support of investment consultants.

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find that unfunded liabilities are positively related to portfolio expected returns. Specifically, an unfunded liability equal to an additional year of total government revenue raises the expected return by 16 basis points. However, the effect of fiscal stress operates primarily through an effect on inflation assumptions as opposed to real returns, and it does not mitigate the extrapolative effect of past returns. While fiscal incentives are important, we find no evidence that fiscal incentives drive the extrapolative effects of past returns.

Having found that none of these hypotheses can explain the full effect of relative past returns on relative return expectations, we turn to the alternative hypothesis that pension funds may place excessive weight on their own past returns, relative to what can be justified by differences in risk taking, persistence in performance, or an attempt to reduce costly rebalancing. The arguments behind this excessive extrapolation hypothesis build on the literature on individual investors (e.g., Vissing-Jorgensen 2003; Kaustia and Knüpfer 2008; Choi et al. 2009; Malmendier and Nagel 2011; Greenwood and Shleifer 2014; Kuchler and Zafar 2019). We find evidence consistent with excessive extrapolation, as pension funds whose investment executives have personally experienced the funds' past returns display an economically stronger relation between past performance and future expectations.[2] An interaction term of executive tenure and past returns explains around one third of the positive relation between past returns and future expectations on a pension fund level. These results are in line with the literature that links the experiences and decisions of individual investors, corporate executives (Malmendier, Tate, and Yan 2011; Gennaioli, Ma, and Shleifer 2016), mutual fund managers (Greenwood and Nagel 2009), and securitization agents (Cheng, Raina, and Xiong 2014). Our contribution is to document that personal experience and beliefs of decision makers play a role in setting return expectations also in an institutional investor setting.

Next, we show that public pension funds act on their extrapolated return assumptions. Specifically, higher past returns affect target long-term allocations to risky assets and induce more risk taking among institutional investors. Each percentage point higher expected risk premium across all risky assets correlates with target allocations that are higher by approximately 1 percentage point. Each percentage point of higher past return increases the target allocation to risky assets overall by 2 percentage points. Putting these results together, under an identifying assumption that past returns affect target asset allocations only through their effects on beliefs, we calculate that for each percentage point of additional risk premium that is driven by extrapolated past returns, overall risk

[2] We control for the investment consultants employed by the pension plans, since consultants provide recommendations and contribute to the process of formulating return expectations (Jenkinson, Jones, and Martinez 2016).

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allocations are higher by 2 to 3 percentage points. The sensitivity of portfolio composition to the return expectations of institutional investors is similar to the sensitivity for individual retail investors (Amromin and Sharpe 2014; Giglio et al. 2021).

We conclude with a discussion of the potential costs to pension funds and their stakeholders that may arise from formulating expectations as a function of past returns. Cross-sectional differences in past performance are correlated with substantial differences in target asset allocation policy. For example, comparing the target allocation to public equity of pension funds that have achieved top quintile public equity returns with those in the bottom quintile reveals a difference of 18 percentage points. Explicitly measuring the welfare cost of these differences in asset allocation is challenging, because it requires specifying the frictions that lead deviations from optimal allocation to impact utility. Investing in any financial asset at a fair price is a zero net present value (NPV) transaction, so for asset allocation to have important quantitative effects there must be reasons why different distributions of outcomes have different welfare consequences. Although it is beyond the scope of this paper to solve this welfare calculation in its entirety, we argue that substantial deviations from the optimal allocation driven by past-return-induced variation in return expectations can be costly for four reasons.

First, in the Campbell and Viceira (2002) myopic model, deviating from an optimal asset allocation has relatively small costs. Being off of the asset allocation optimum by 10 percentage points reduces the certainty equivalent by only 1–3 basis points under standard parameters. However, since defined benefit pension funds have liabilities and funding requirements, there are additional economic costs to shortfalls. Van Binsbergen and Brandt (2016) find that when incorporating funding constrains in the analysis, the certainty equivalent costs of deviating from the optimal allocation could be as large as 150 basis points for differences of 10 percentage points in the allocation to risky assets. Second, since pension plans tend to achieve their new past-performance-driven target weights not only through mechanical changes but also through active rebalancing and asset management transitions, they pay transition costs when they change their asset allocation, which we estimate may be on the order of 33 basis points per mandate transition. Third, recent research suggests that the price elasticity of the aggregate stock market is small and the aggregate multiplier from flows is large (Koijen and Yogo 2019; Ben-David et al. 2021; Gabaix and Koijen 2021). If the aggregate multiplier is high, pension fund flows in and out of risky assets induced by extrapolative expectations could have quantitatively large effects on aggregate asset prices. Fourth, in alternative asset classes, public pension funds that increase their allocation may take marginal investments in which they underperform investors with superior access to the best funds (e.g., Lerner, Schoar, and Wongsunwai 2007; Sensoy, Wang, and Weisbach 2014).

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1. Portfolio and Asset-Class Expected Returns

GASB Statement No. 67 (GASB 2012) describes the required disclosure and provides an example, which we show in Internet Appendix Figure A.1.[3] We collect the asset-class specific disclosures for 228 U.S. state and local government pension plans over the period 2014 to 2017. Pension plans present this information in their Comprehensive Annual Financial Reports (CAFRs) or in separate GASB 67 statements. Under the requirements of GASB 67, pension sponsors can choose to designate the disclosed asset-class expected returns as either arithmetic or geometric. If returns are lognormally distributed with mean μ and variance σ^2, then the difference between these two expressions converges as T gets large to approximately $\sigma^2/2$ under the standard statistical properties of the normal distribution. Although some pension plans report asset-class-based expected returns on a nominal basis and others report them on a real basis, all plans disclose the underlying inflation assumption. We first harmonize all disclosures to a nominal basis and then examine the inflation rate assumption and real rate of return assumptions separately.

The disclosure allows us to calculate the portfolio expected return (Portfolio ER) of a pension fund as the inner product of the two vectors: (i) a vector of the fund's expected returns by asset class; and (ii) a vector of the fund's target weights on each asset class. The stated purpose of the GASB 67 disclosure is to provide some justification for the discount rate that pension plans use to calculate the present value of their pension liabilities. From a process standpoint, however, the Portfolio ER informs the setting of the pension discount rate (Pension DR) only in part. The pension plan board decides on the Pension DR based in part on the Portfolio ER and in consultation with its actuaries. Pension fund investment staff typically develop the Portfolio ER, often with the support of investment consultants.[4] As a result, we find significantly more variation in the Portfolio ER than in the Pension DR. The Portfolio ER generally does not match the Pension DR, exceeding it in 67% of observations and falling short of it in 25% of cases.[5] While some of these

[3] Specifically, GASB Statement No. 67 requires that "the following information should be disclosed. … (c) The long-term expected rate of return on pension plan investments and a description of how it was determined, including significant methods and assumptions used for that purpose. … (f) The assumed asset allocation of the pension plan's portfolio, the long-term expected real rate of return for each major asset class, and whether the expected rates of return are presented as arithmetic or geometric means, if not otherwise disclosed."

[4] In the text of the reports, we document differences in the process of formulating Portfolio ER and Pension DR. For instance, in the Florida Retirement System (FRS) for 2017, the Portfolio ER is 7.10% and the Pension DR is 7.50%. The Comprehensive Annual Financial Report of FRS states: "The 7.10 percent reported investment return assumption differs from the 7.50 percent investment return assumption chosen by the 2017 FRS Actuarial Assumption Conference for funding policy purposes, as allowable under governmental accounting and reporting standards" (Florida Retirement System 2017). In Florida, the Actuarial Assumption Conference refers to a meeting of the state pension fund board (known as the State Board of Administration), the state's actuary, and the state's investment consultant.

[5] Internet Appendix Table A.1. shows a mismatch between the Portfolio ER and the Pension DR for 93% of pension plans reporting the Portfolio ER on an arithmetic basis and for 91% of pension plans reporting on a geometric basis.

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differences may be due to reporting of geometric versus arithmetic expectations, most of the differences appear to be due to real differences between expected returns and the discount rate.

Figure 1 plots the Pension DR against the Portfolio ER, with panel A showing the 571 pension plans that report the Portfolio ER on an arithmetic basis and panel B showing the 319 pension plans that report the Portfolio ER on a geometric basis. The beta of the Portfolio ER with respect to the Pension DR is 0.267 and 0.869, respectively, and the R-squared statistics are 0.015 and 0.207. Thus, the Portfolio ER and Pension DR are positively related, but there is also considerable variation in the Portfolio ER that is not explained by a plan's choice of Pension DR. For example, there are 49 pension plans in our sample that report the same Pension DR of 7.50% in 2014, but their arithmetic Portfolio ER range from 7.19% to 11.32%. Overall, the fact that there is considerably more variation in the Portfolio ER provides an opportunity to analyze the drivers of heterogeneity in the formation of return assumptions.

Table 1 and Figure 2 show summary statistics separately for the sample of pension plans choosing an arithmetic Portfolio ER versus a geometric Portfolio ER. In both cases, the expected return equals the sum of the assumed inflation rate and expected real return. In order to achieve a homogeneous set of asset classes, we aggregate all disclosures into seven categories: fixed income, cash, public equity, real assets, hedge funds, private equity, and other risky assets. Real assets include real estate, infrastructure, and natural resources. Hedge funds include hedge funds with different styles as well as tactical asset allocation mandates. Private equity includes buyout and venture capital investments. Other risky assets include undifferentiated portfolios of alternative assets, commodities, and derivatives.

The arithmetic and geometric subsamples have similar portfolio composition. Fixed income represents on average 24.6% of the portfolio for systems reporting the ER on an arithmetic basis and 25.0% for systems reporting the ER on a geometric basis. Public equity averages 46.6% and 46.9%, respectively. Among the alternative asset classes, the arithmetic systems are slightly more likely to invest in real assets and private equity, whereas the geometric are slightly more likely to invest in in hedge funds and other risky assets.

The Portfolio ERs are 8.18% on average in the arithmetic sample and 7.53% in the geometric sample. Geometric returns are therefore on average 0.65% lower than arithmetic, which under the standard approximation would imply volatility of 11.40% (if $\sigma^2/2=0.0065$, then $\sigma=0.1140$). In comparison, the average of the time-series standard deviation of returns of funds in our sample is slightly higher at 12.09%. The fact that expected return differences are most pronounced in the risky asset classes and minimal in fixed income and cash supports the hypothesis that the differences between the arithmetic and geometric samples are a result of the difference between arithmetic and geometric disclosures, as opposed to different underlying assumptions about return moments in these two samples.

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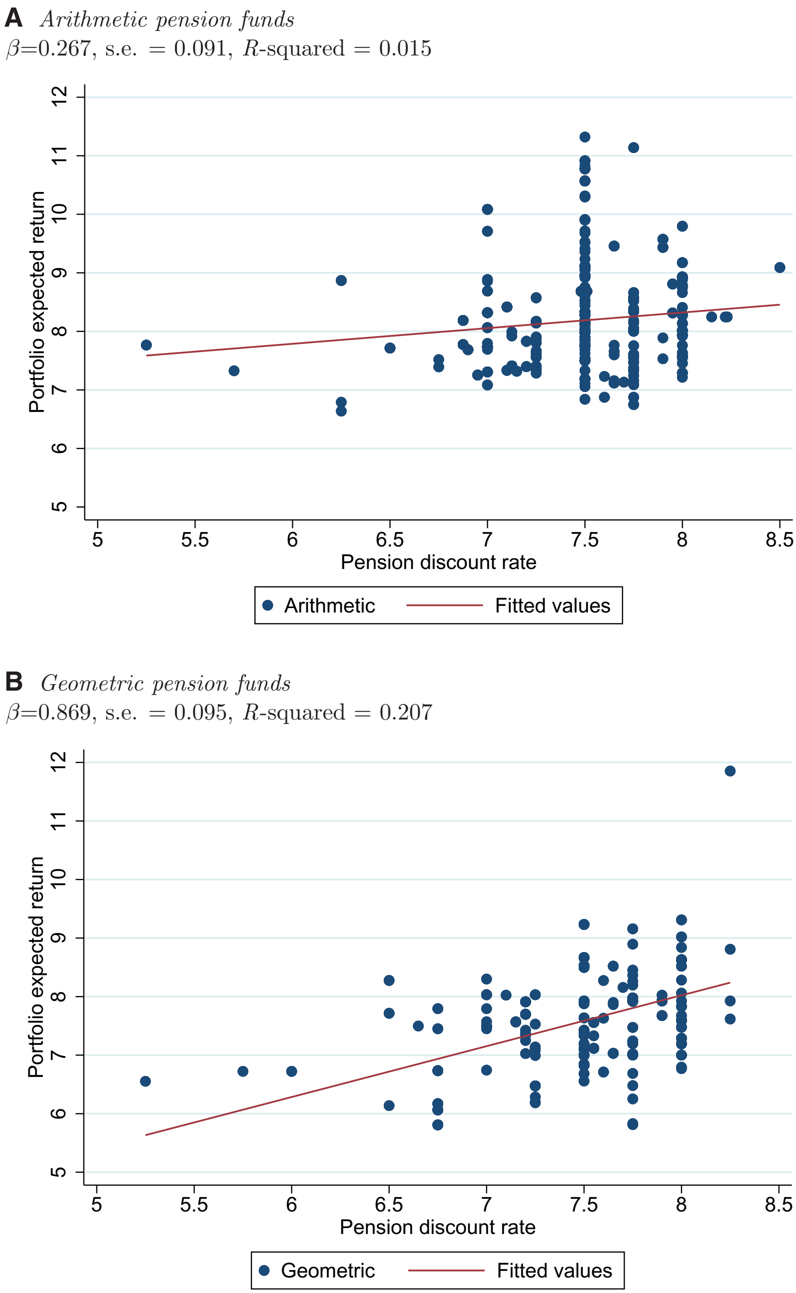

Figure 1
Portfolio expected return and pension discount rate

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Table 1
Summary statistics

	Arithmetic				Geometric			
	PFs	Mean	Median	SD	PFs	Mean	Median	SD
A. Portfolio expected return								
Portfolio ER	571	8.183	8.048	0.840	319	7.533	7.497	0.862
Inflation rate	571	2.793	2.750	0.358	319	2.718	2.750	0.417
Real return	571	5.390	5.298	0.885	319	4.816	4.699	0.964
%Fixed income	571	0.246	0.240	0.071	319	0.250	0.250	0.105
%Cash	571	0.011	0.000	0.014	319	0.002	0.000	0.068
%Equity	571	0.466	0.440	0.096	319	0.469	0.500	0.148
%Real assets	571	0.108	0.100	0.063	319	0.075	0.075	0.058
%Hedge funds	571	0.066	0.040	0.073	319	0.080	0.040	0.134
%Private equity	571	0.082	0.090	0.064	319	0.060	0.070	0.049
%Other risky assets	571	0.021	0.000	0.055	319	0.065	0.000	0.088
ER Fixed income	571	4.793	4.750	1.186	313	4.705	4.600	0.829
ER Cash	279	3.355	3.150	1.292	155	2.758	2.950	0.804
ER Equity	571	9.430	9.337	1.112	311	8.527	8.625	0.817
ER Real assets	502	7.966	7.775	0.992	228	7.623	7.500	1.265
ER Hedge funds	301	7.336	7.150	1.374	163	6.669	6.617	0.869
ER Private equity	432	11.984	11.800	1.599	207	10.104	9.950	1.272
ER Other risky assets	136	9.723	8.650	2.537	150	7.936	8.400	2.504
B. Pension plan and state variables								
Past return	571	6.524	6.430	1.010	319	6.710	6.710	1.305
Past standard deviation	571	12.071	12.160	1.533	319	12.107	12.180	1.294
PF size ($ bil.)	571	11.596	1.705	23.423	319	16.408	2.578	44.043
Past return equity	506	7.733	7.749	1.289	280	8.027	7.739	1.415
Past return RA	368	6.803	6.930	2.528	163	6.982	6.956	3.019
Past return PE	306	11.352	11.445	3.317	157	12.839	13.543	3.465
Unfunded liability / Revenue	571	1.649	1.529	0.787	319	1.753	1.497	0.878
Unfunded liability / GSP	571	0.214	0.193	0.102	319	0.219	0.206	0.112
GSP per capita ($ thousands)	571	49.034	45.493	12.742	319	50.964	52.863	15.080

Panel A presents summary statistics separately for pension plans reporting on an arithmetic and geometric basis. The two main components of the portfolio expected return (Portfolio ER) are the assumed inflation rate and the expected real return. The Portfolio ER is calculated using the target weights and expected returns by asset class. We organize asset allocation into seven asset classes: fixed income, cash, equity, real assets, hedge funds, private equity, and other risky assets. For every asset class, we present the target allocation and the expected nominal return. Column *PFs* reports the number of observations, and it decreases when we present the expected returns by asset class because some pension plans do not invest in every asset class. In Panel B, we report summary statistics for the other variables used in our analysis. *Past return* measures the average annual arithmetic return in the previous 10-year period. *Past standard deviation* measures the standard deviation of the returns in the previous 10-year period. *Past return equity*, *Past return RA*, and *Past return PE* measure the average arithmetic return in equity, real assets, and private equity. *PF size* ($ bil.) presents the assets under management. *Unfunded liability / Revenue* and *Unfunded liability / GSP* are ratios of unfunded liabilities of public pension funds relative to state revenue or Gross State Product.

Table 1 also shows summary statistics for the *Past return*, which is the average of the 10-year arithmetic return, and *Past standard deviation*, which is the standard deviation of the annual returns in the previous 10-year period. We measure the annual returns as total net investment income divided by

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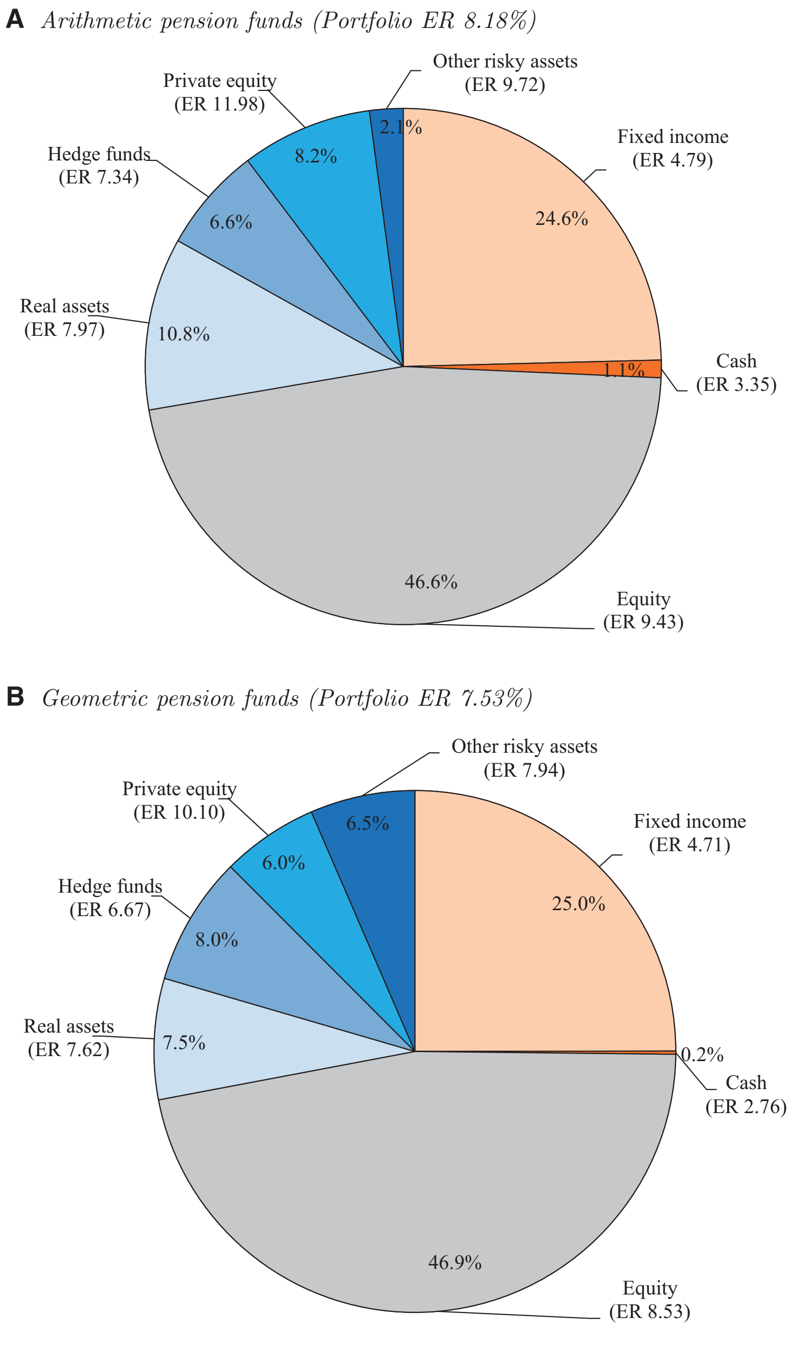

Figure 2
Pension plan asset allocation and portfolio expected return (ER)

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beginning-of-year assets.[6] Pension funds disclose this information in the Financial Section of their CAFRs. The average *Past return* is 6.53% in the arithmetic and 6.71% in the geometric sample, well below the Portfolio ERs of 8.18% in the arithmetic and 7.53% in the geometric sample. For arithmetic (geometric) systems, 90% (72%) of the systems have a Portfolio ER that exceeds the 10-year average past return.[7] Overall, pension plans appear to be optimistic in their beliefs about future returns relative to the returns of the past 10 years.

Past return equity, Past return RA, and *Past return PE* are collected from the Investments Section of the CAFRs. These variables capture the average arithmetic return in public equity, real assets, and private equity. In equity, almost all pension plans report a time series of annual returns, and we estimate the average past return in equity using the returns in the previous 10-year period. In real assets and private equity, the reported history of returns may be shorter than 10 years if a pension plan did not invest for the entire 10 years in the asset class or did not report returns for the entire period. We require a time series of at least five returns in order to estimate the average return in these asset classes.

Finally, Table 1 shows measures of unfunded pension liabilities scaled either by state revenue from taxes and fees or by Gross State Product (GSP). The unfunded liability measure revalues each state and local government's accrued liabilities using the point on the Treasury yield curve that matches the plan-specific duration (Rauh 2017). These liability measures therefore depend not on the discount rate chosen by the pension fund but rather on market bond yields. The average value of unfunded liability is 1.65 years of state and local own-generated revenue for arithmetic plans and 1.75 years for geometric plans, or 21.4% of annual GSP for arithmetic plans and 21.9% for geometric plans.

2. Explaining the Portfolio Expected Return

2.1 Portfolio expected return

In this section, we analyze the determinants of the cross-sectional variation in the portfolio expected return. Our starting null hypothesis is that the only determinants of expected returns are the arithmetic or geometric disclosure basis and the chosen target asset allocation. We test this null hypothesis against the alternative hypothesis that past returns experienced by the pension plan affect the expected return. Our data allow us to disentangle the effects of asset allocation and past performance on return expectations because we observe target weights that are not subject to mechanical market movements (actual allocation weights will be mechanically affected by performance). A further

[6] In addition to real differences in within-asset-class performance, these return measurements could be affected by differences in timing of contributions and pension benefit payments during the year, as well as how systems choose to mark the value of their unrealized stakes in private equity funds and other funds involving illiquid assets.

[7] Furthermore, when we calculate the past return as a geometric past return, for 287 (or 90%) of the 319 systems the assumed geometric return exceeds the past geometric return.

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null hypothesis that we test is that within a given asset class, the past return affects neither the expected return nor the expected risk premium.

We begin the empirical investigation by examining the relation between the portfolio expected return (*ER*) and *Geometric* as an indicator variable for geometric reporting, as well as $\boldsymbol{\omega}$, a 5-vector of target allocations to public equity, real assets, private equity, hedge funds, and other risky assets. The omitted asset categories are fixed income and cash. We estimate the following equation:

$$ER_{it} = \alpha_t + \beta Geometric_{it} + \boldsymbol{\theta}' \boldsymbol{\omega}_{it} + \gamma Controls_{it} + \varepsilon_{it}. \tag{1}$$

We control for *PF size*, which is the natural logarithm of pension fund assets under management. We measure the asset under management on the level of the institution that invests the assets. If the assets of multiple pension plans are pooled and invested by one institution, we use the sum of the assets when calculating pension fund size.[8] We use year-by-reporting-month fixed effects because pension plans have different fiscal-year ending dates and cluster the standard errors by pension plan.

Column 1 of Table 2 shows that pension plans reporting on a geometric basis have a 75 basis points lower portfolio expected return than pension plans reporting on an arithmetic basis. In addition, pension plans that invest more in risky assets expect higher returns overall. Relative to a 100% portfolio of fixed income and cash, each percentage point increase in the allocation to public equity, real assets, hedge funds, and other risky assets raises the expected return by 3.4, 4.1, 3.0, and 4.7 basis points, respectively. The asset allocation variables combined with the geometric indicator and control for pension fund size explain 27.2% of the variation in the portfolio expected return.

In column 2, we replace the asset allocation variables with *Past return*, defined as the 10-year arithmetic average of prior annual returns. This equation has an adjusted R-squared that is only slightly lower than Equation (1), at 25.3%. When we include both asset allocation weights and the term for *Past return* in column 3, we can explain 30.2% of the variation in the expected return and the *Past return* coefficient is statistically significant. Each additional percentage point of past return increases the Portfolio ER by 26 basis points. Column 4 adds a control for the past standard deviation, to examine whether past risk taking might be the driving factor instead of past return. If some funds generally take more risk within asset classes than others, they might also expect higher future returns, but one would also expect to see that linked to higher past

[8] For example, four pension plans in Connecticut (JRS, MERS, SERS, and TRS) have different target asset allocation policies and sometimes different expected returns by asset class. However, the execution of their investments is done together by the State of Connecticut Retirement Plans and Trust Funds. In our analysis, we consider the size of the assets managed by the joint entity, State of Connecticut Retirement Plans and Trust Funds, because it reflects better potential investment experience, negotiating power, and (dis)economies of scale.

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Table 2
Portfolio expected return

	Portfolio expected return							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Geometric	−0.745***	−0.653***	−0.750***	−0.739***	−0.752***	−0.745***	−0.757***	−0.752***
	[0.122]	[0.097]	[0.117]	[0.117]	[0.116]	[0.114]	[0.115]	[0.113]
Past return		0.300***	0.258***	0.265***	0.292***	0.305***		
		[0.055]	[0.068]	[0.066]	[0.064]	[0.065]		
Benchmark past return							0.174***	0.177***
							[0.030]	[0.030]
Abnormal past return							0.281***	0.294***
							[0.067]	[0.067]
Past standard deviation				−0.034	−0.052	−0.061	−0.061*	−0.069**
				[0.045]	[0.043]	[0.043]	[0.031]	[0.032]
Unfunded liability / Revenue					0.167***		0.144***	
					[0.054]		[0.050]	
Unfunded liability / GSP						1.391***		1.210***
						[0.443]		[0.413]
GSP per capita					0.006*	0.006*	0.008**	0.008***
					[0.003]	[0.003]	[0.003]	[0.003]
PF size	−0.073	−0.166***	−0.099*	−0.109**	−0.133***	−0.133***	−0.112**	−0.113**
	[0.052]	[0.055]	[0.054]	[0.053]	[0.051]	[0.051]	[0.050]	[0.050]
%Equity	3.443***		2.094***	2.252**	2.036**	1.953**	2.642***	2.567***
	[0.817]		[0.805]	[0.872]	[0.909]	[0.894]	[0.799]	[0.789]
%Real assets	4.088***		3.164***	3.516***	3.408***	3.325***	3.795***	3.736***
	[1.222]		[1.100]	[1.134]	[1.093]	[1.086]	[1.016]	[1.014]
%Private equity	0.655		−0.610	−0.369	−0.126	−0.089	−0.056	−0.051
	[1.040]		[1.040]	[1.128]	[1.071]	[1.061]	[1.017]	[1.015]
%Hedge funds	3.008***		2.772***	2.888***	2.880***	2.826***	3.463***	3.427***
	[0.692]		[0.661]	[0.717]	[0.745]	[0.724]	[0.621]	[0.609]
%Other risky assets	4.682***		2.921**	3.142**	2.841**	2.848**	2.927**	2.924**
	[1.275]		[1.304]	[1.275]	[1.221]	[1.209]	[1.232]	[1.224]
Year × Reporting month FE	Yes	Yes	Yes	Yes	Yes	Yes	No	No
Observations	890	890	890	890	890	890	890	890
Adjusted *R*-squared	0.272	0.253	0.302	0.303	0.325	0.327	0.301	0.302

This table presents results of regressions in which the dependent variable is the portfolio expected return of pension plans during the 2014–2017 period. *Geometric* is an indicator variable for pension plans reporting geometric portfolio expected returns (the omitted category is plans reporting arithmetic expected returns). *Past return* and *Past standard deviation* measure the average arithmetic return and the standard deviation of the annual returns in the previous 10-year period. *PF size* is the natural logarithm of pension fund assets under management. *Unfunded liability / Revenue* and *Unfunded liability / GSP* are ratios of unfunded liabilities of state and local pension funds relative to state revenues or Gross State Product. *GSP per capita* is Gross State Product per capita in \$ thousands. *%Equity*, *%Real assets*, *%Private equity*, *%Hedge funds*, and *%Other risky assets* measure the percentage allocated to different risky asset classes (the omitted categories are fixed income and cash). In columns 7 and 8, we decompose the *Past return* into two parts. *Benchmark past return* captures the average past return of all pension plans reporting performance in the same year and in the same month. *Abnormal past return* equals the difference between the *Past return* of pension fund i and the *Benchmark past return* in that year and reporting month. Columns 1–6 include year-reporting-month fixed effects, as pension plans have different fiscal-year ending dates. Standard errors clustered by pension plan are reported in brackets. $^{*}p<.1$; $^{**}p<.05$; $^{***}p<.01$.

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volatility. We find, however, that the coefficient on past standard deviation is insignificant, while the coefficient on *Past return* remains significant.[9]

Columns 5 and 6 of Table 2 augment the regression equation further by including variables for the unfunded pension liability of the sponsoring state or local government, scaled by revenue or GSP, respectively. A one standard deviation increase in the unfunded liabilities scaled by revenues, which corresponds to 0.82 of a year of revenues, raises the Portfolio ER by 14 basis points. Scaling the unfunded liability by GSP yields similar conclusions. Each additional 10 percentage points of *Unfunded liability / GSP* raises the Portfolio ER also by 14 basis points, consistent with the hypothesis that fiscally stressed governments face pressure to maintain higher expected rates of return. The inclusion of these controls does not attenuate the *Past return* coefficient, and in fact the coefficient increases to 30 basis points. Our analysis shows that if pension plan A had 1 percentage point higher past return than pension plan B, pension plan A will expect around 30 basis points higher return in the future even after controlling for asset allocation.

Our object of inquiry is cross-sectional differences in return expectations as a function of cross-sectional differences in past returns, so we examine the relation between return expectations, past returns, and target weights while controlling for year-reporting-month fixed effects in order to isolate the asset-class-level movements. If there is return predictability on an asset class level (e.g., predictability based on the dividend price ratio), investors would optimally incorporate this predictability into their strategic asset allocation models, which would generate different expected returns at different investment horizons (Brennan, Schwartz, and Lagnado 1997; Campbell and Viceira 1999; Barberis 2000). In our analysis, the year-reporting-month fixed effects absorb any time-series variation in asset class returns and market valuation measures since the aggregate asset-class-level performance will be the same for all pension funds reporting at that moment of time. Using these fixed effects enables us to study cross-sectional differences in return expectations as a function of cross-sectional differences in past returns.[10]

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9 In at least one report we find explicit evidence of the use of past returns in setting expected returns. According to the Tennessee Consolidated Retirement System (TCRS): “The long-term expected rate of return on pension plan investments was established by the TCRS Board of Trustees in conjunction with the June 30, 2012 actuarial experience study by considering the following three techniques: (1) the 25-year historical return of the TCRS at June 30, 2012, (2) the historical market returns of asset classes from 1926 to 2012 using the TCRS investment policy asset allocation, and (3) capital market projections that were utilized as a building-block method in which best-estimate ranges of expected future real rate of return (expected returns, net of pension plan investment expense and inflation) are developed for each major asset class” (Tennessee Consolidated Retirement System 2014).

10 One disadvantage of the year-reporting-month fixed effects is that we cannot control for marketwide fundamental valuation as these fixed effects absorb the time-series variation in any market valuation measure. In Internet Appendix Table A.6, we examine the role of fundamental valuation and add the cyclically adjusted price-to-earnings (CAPE) ratio as a control variable. In this analysis, we only use separate fixed effects for year and reporting month. The data on the CAPE ratio come from the Robert Shiller’s data set. Our results on the cross-sectional differences in expected returns are not affected by the inclusion of the CAPE ratio. We still find that

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To illustrate the rationale behind this, in columns 7 and 8 in Table 2, we decompose the past return into two parts but remove the year-reporting-month fixed effects. The first component, *Benchmark past returns*, equals the average past return of all pension plans reporting performance in the same year and in the same month. This component captures common movements in returns over time. The second component, *Abnormal past returns*, equals the difference between the past return of pension fund i and the *Benchmark past return* in that year and reporting month.

We find that the coefficient on *Abnormal past return* in columns 7 and 8 is almost the same as the coefficient on *Past return* in columns 5 and 6 of Table 2. This implies that the year-reporting-month fixed effects used in columns 5 and 6 capture the common component in past performance and enable us to examine the cross-sectional differences in pension fund asset allocation and return expectations. The extrapolation of past returns when formulating return expectations is not due to asset-class-level return predictability (e.g., predictability based on the dividend price ratio). Our empirical setting with year-reporting-month fixed effects allows us to derive conclusions that pension funds' relative differences in past performance affect their relative return expectations.[11]

The specifications in Table 2 include an indicator for pension plans reporting geometric expectations, effectively imposing a linearity assumption and also assuming that any selection mechanism between plans that report arithmetic versus geometric forecasts does not affect the estimates of the past return effect. In Internet Appendix Table A.2, we estimate a logit model in which the dependent variable equals one if a pension plan reports geometric expected returns. We find that our main variable of interest, *Past return*, is not related to the probability of reporting geometric expectations. Among the other variables, larger pension plans are significantly more likely to report geometric expectations. In Internet Appendix Table A.3, we estimate a robustness test of our results separately for arithmetic and geometric plans. We find that the relation between the expected return and past return is significant in both subsamples.

In our analysis, we estimate average past returns using equal weights on the realized returns in each of the previous 10 years. Other studies that analyze the relation between experience and expectations, such as Malmendier and Nagel (2011), estimate weighting parameters over different past time horizons relative

pension plans with higher past performance expect higher returns than pension funds with lower past performance. We also find that the CAPE ratio is negatively related to the portfolio expected return, so the relation between aggregate expectations of pension funds and market level seems to differ from the results of Greenwood and Shleifer (2014) on individual investors. These results should be taken with caution because we have a short time series of expectations.

[11] The coefficient on *Benchmark past return* is also positive and significant. The sample has substantial cross-sectional variation, but the time series is short. Hence it is not appropriate to assign a strong economic interpretation to this result.

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to the time at which the expectation is set. In Internet Appendix Table A.4, we calculate average past performance using alternative weights (e.g., more weight on recent observations instead of equal weights). We find that the different weighting schemes do not affect our results materially; in fact, the equal weights provide the greatest explanatory power.[12]

In the robustness test in Internet Appendix Table A.5, we limit attention to the subsample of observations in the year 2014 to address two issues. First, observations in 2014 are the first reported expectations based on the new disclosure requirement. Any strategic incentives to adjust expectations in response to feedback received on the initial disclosure are therefore less likely to affect these disclosures. Second, our specification focuses on the *Past return* variable, a moving average of the returns in the previous 10-year period that is therefore quite stable over our sample period. If some pension plans maintain the same portfolio expected return over time, one potential worry is that using panel data increases the sample size and that clustering of standard errors by pension plan might not sufficiently address this issue. In Internet Appendix Table A.5, we obtain estimates similar to those in our main models from Table 2 in terms of both economic and statistical significance.

2.2 Components of portfolio expected return

Next, we study the mechanisms behind the positive relation between the portfolio expected return and past returns by analyzing the different components of the portfolio expected return. The expected return on any asset can be decomposed into an expected inflation rate plus an expected real return, or into an expected risk-free return plus an expected risk premium.

The first possible mechanism is that pension plans with higher past returns might assume higher expected returns in all asset classes. If pension plans with higher past returns assume a higher inflation rate or a higher risk-free rate, they can raise the expected returns on all assets in the portfolio. The second possible mechanism is that pension plans with higher past returns can assume a higher expected real rate of return or a higher risk premium for investing in risky assets. This mechanism implies that pension plans with higher past returns will expect higher returns for investing in risky assets, but they will expect similar inflation rate and returns for investing in nonrisky assets.

We distinguish between these two mechanisms by decomposing the expected return in two ways. First, we consider separately variation in real return assumptions versus variation in expected inflation: $ER_t = E\pi_t + Er_t$, where π_t represents the inflation rate and r_t represents real returns. We examine whether

[12] The exact formula for the weighting function follows Malmendier and Nagel (2011), Equation (1). We note that the coefficients when L is not equal to zero do not have a direct economic interpretation, hence we focus on comparing the adjusted R-squared across the different models to assess fit. Compared to Malmendier and Nagel (2011), the time period in our analysis is considerably shorter, covering only the returns in the previous 10-year period, as opposed to entire human lifetimes. This short time period may explain why different weighting functions yield similar results.

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Table 3
Expected inflation rate and return on safe assets (components of Portfolio ER)

	Expected inflation rate			Expected risk-free rate		
	(1)	(2)	(3)	(4)	(5)	(6)
Geometric	−0.076*	−0.074	−0.095**	0.082	0.103	0.081
	[0.045]	[0.045]	[0.040]	[0.127]	[0.128]	[0.126]
Past return		−0.020	0.010		−0.223***	−0.217***
		[0.026]	[0.020]		[0.084]	[0.078]
Past standard deviation			−0.045***			0.005
			[0.014]			[0.054]
Past state inflation			−0.319***			
			[0.097]			
Unfunded liability / GSP			0.849***			1.315**
			[0.176]			[0.530]
GSP per capita			0.009***			0.004
			[0.002]			[0.003]
PF size	−0.045**	−0.043**	−0.036*	−0.224***	−0.204***	−0.209***
	[0.021]	[0.021]	[0.020]	[0.065]	[0.066]	[0.066]
Year × Reporting month FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	890	890	890	884	884	884
Adjusted R-squared	0.199	0.200	0.336	0.072	0.095	0.107

This table presents results of regressions in which the dependent variables are components of the portfolio expected return. In columns 1–3, the dependent variable is the expected inflation rate. In columns 4–6, the dependent variable is the expected return in fixed income and cash assets of pension plans. *Geometric* is an indicator variable for pension plans reporting geometric portfolio expected return (the omitted category is plans reporting arithmetic expected return). *Past return* and *Past standard deviation* measure the average arithmetic return and the standard deviation of annual returns in the previous 10-year period. *Past state inflation* is the average annual inflation rate in the state in the previous 10-year period. *PF size* is the natural logarithm of pension fund assets under management. *Unfunded liability / GSP* is the ratio of unfunded liabilities of state and local pension funds relative to state revenues or Gross State Product. *GSP per capita* is Gross State Product per capita in $ thousands. The specifications include year-reporting-month fixed effects. Standard errors clustered by pension plan are reported in brackets. *$p<.1$; **$p<.05$; ***$p<.01$.

higher past returns are increasing the real component or inflation component of future expected returns:

$$E\pi_{it}=\alpha_t+\beta_1\overline{R_{i,t-j}}+\mathbf{\Gamma_1'}\mathbf{X}+\epsilon_{it}, \tag{2}$$

$$Er_{it}=\alpha_t+\beta_1\overline{R_{i,t-j}}+\mathbf{\Gamma_1'}\mathbf{X}+\epsilon_{it}, \tag{3}$$

where $R_{i,t-j}$ is the average realized return of the fund over the previous j years (10 years in our setting). Second, we consider separately the risk-free rate and the expected risk premium: $ER_t = Er_{ft} + E(R_t - r_{ft})$, where r_{ft} represents the risk-free rate:

$$Er_{f,it}=\alpha_t+\beta_1\overline{R_{i,t-j}}+\mathbf{\Gamma_1'}\mathbf{X}+\epsilon_{it}, \tag{4}$$

$$E(R_{it}-r_{f,it})=\alpha_t+\beta_1\overline{R_{i,t-j}}+\mathbf{\Gamma_1'}\mathbf{X}+\epsilon_{it}. \tag{5}$$

In all these equations, note that we observe the expectations of return components directly in the data.

Table 3 examines the first mechanism and the dependent variable measures either the expected inflation rate or the expected risk-free rate of return. In columns 1–3, we document that the coefficient on *Past return* is economically and statistically insignificant, so the relation between past returns on future

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return expectations does not operate through inflation assumptions. Pension plans with higher past return are not simply adding higher inflation rates to their real return assumptions.

The specification in column 3 shows that the ratio of unfunded liabilities is significantly related to expected inflation. The economic magnitude is around 60% of the unfunded liability effects found in Table 2. A one standard deviation increase in this fiscal pressure variable is 10.6% of annual GSP, and so would be consistent with an inflation rate assumption that is higher by 9 basis points. This suggests that pension funds in states with large unfunded liabilities relative to their resources tend to justify higher return assumptions in part by using higher inflation assumptions.[13]

In columns 4–6 of Table 3, we focus on the risk-free rate as a potential mechanism to extrapolate past performance and raise the expected returns on all assets in the portfolio. We proxy for the expected risk-free rate of return using the expected return on cash and fixed income assets. We estimate Equation (4) and arrive at an even starker conclusion that past returns have a negative effect on the expected returns on fixed income and cash. That is, pension plans with higher past return expect lower future return on safe assets. The results for unfunded liabilities confirm that pension plans located in states with higher fiscal pressure have incentives to formulate higher expectations. The ratio of unfunded liabilities to state GSP is positively related to the expected return on fixed income and cash assets. Our interpretation of these results is that fiscally constrained pension plans have incentives to assume a higher inflation rate or higher risk-free rate of return in order to raise the expected returns on all assets in the portfolio.

Table 4 examines the second mechanism and the dependent variable measures either the expected real rate of return or the expected risk premium. We estimate the expected real returns and risk premium for all risky assets together as well as separately for equity, real assets, and private equity. The expected real return or risk premium on all risky assets is a weighted average of the expected risk premium in equity, real assets, private equity, hedge funds, and other risky assets.

In column 1 we estimate Equation (4) by considering the expected real rate of return and find contrasting effects to those seen in the inflation analysis. Pension funds with higher past returns expect higher real returns in risky assets. The coefficient on past returns in this analysis is essentially the same as it is in

[13] The third column also adds a control for average past inflation in the state of the pension plan in the previous 10-year period. This variable is based on the Consumer Price Index for All Urban Consumers (CPI-U) reported by the Bureau of Labor Statistics on a combined statistical area level. Past local inflation might affect the inflation beliefs of pension funds if fund officials have lived most of their lives in the local area or if pension funds tend to overweight local investments in their portfolios (Hochberg and Rauh 2013; Brown, Pollet, and Weisbenner 2015; Malmendier and Nagel 2016). That said, there is no evidence that pension fund performance depends on the cross-sectional differences in inflation rates across U.S. states, let alone evidence that inflation is persistent within U.S. regions. We find no evidence that pension plans make higher inflation assumptions on the basis of higher past regional inflation.

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Table 4
Expected real return and risk premium (components of Portfolio ER)

	Expected real return				Expected risk premium			
	All (1)	Equity (2)	RA (3)	PE (4)	All (5)	Equity (6)	RA (7)	PE (8)
Geometric	−0.648***	−0.644***	−0.172	−1.910***	−1.036***	−0.985***	−0.666***	−2.469***
	[0.129]	[0.132]	[0.193]	[0.255]	[0.125]	[0.126]	[0.169]	[0.232]
Past return	0.316***	0.167*	0.313***	0.339*	0.272***	0.183**	0.425***	0.364***
	[0.069]	[0.088]	[0.102]	[0.187]	[0.103]	[0.080]	[0.113]	[0.112]
Past return equity		0.066				0.525***		
		[0.138]				[0.109]		
Past return RA			0.153***				0.036	
			[0.030]				[0.037]	
Past return PE				0.192***				0.113***
				[0.041]				[0.040]
Past standard deviation	−0.051	−0.111**	−0.150**	−0.011	−0.152***	−0.161**	−0.046	−0.065
	[0.046]	[0.053]	[0.065]	[0.207]	[0.056]	[0.066]	[0.083]	[0.206]
Unfunded liability / GSP	0.735	0.557	1.297**	0.719	−0.583	−0.061	0.756	0.314
	[0.465]	[0.520]	[0.615]	[1.340]	[0.575]	[0.510]	[0.639]	[1.410]
GSP per capita	−0.001	−0.001	0.018***	0.007	0.006	0.004	0.019***	0.010*
	[0.004]	[0.005]	[0.003]	[0.006]	[0.006]	[0.004]	[0.005]	[0.006]
PF size	−0.105*	−0.140	0.092	−0.088	0.100	0.014	0.076	0.060
	[0.055]	[0.088]	[0.109]	[0.176]	[0.070]	[0.071]	[0.107]	[0.192]
Asset allocation controls	Yes	No	No	No	Yes	No	No	No
Year × Reporting month FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	890	786	531	463	884	786	531	463
Adjusted R-squared	0.264	0.163	0.305	0.333	0.366	0.258	0.292	0.447

This table presents results of regressions in which the dependent variables are components of the portfolio expected return. In columns 1–4, the dependent variable is the expected real rate of return. In columns 5–8, the dependent variable is the risk premium expected by pension plans, measured as the difference between the expected return on a risky asset minus the expected return on fixed income and cash. We estimate the expected real returns and risk premium for all risky assets together as well as separately for equity, real assets (RA), and private equity (PE). Risky assets include equity, real assets, private equity, hedge funds, and other risky assets. *Geometric* is an indicator variable for pension plans reporting geometric portfolio expected return (the omitted category is plans reporting arithmetic expected return). *Past return* and *Past standard deviation* measure the average arithmetic return and the standard deviation of the annual returns in the previous 10-year period. *Past return equity* measures the average arithmetic return in public equity in the previous 10-year period. *Past return RA* and *Past return PE* capture the average arithmetic return in real assets and private equity. *PF size* is the natural logarithm of pension fund assets under management. *Unfunded liability / GSP* is the ratio of unfunded liabilities of state and local pension funds relative to the state revenues or Gross State Product. *GSP per capita* is Gross State Product per capita in $ thousands. The specifications also include controls for the percentage allocated to different risky asset classes and year-reporting-month fixed effects. Standard errors clustered by pension plan are reported in brackets. $^{*}p<.1$; $^{**}p<.05$; $^{***}p<.01$.

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the nominal portfolio expected return analysis of Table 2, indicating that the extrapolation of past returns operates completely through increasing the pension fund's real expected return assumptions. In column 5 of Table 4, we extend the analysis and focus on the expected risk premium for investing in risky asset classes. The risk premium equals the difference between the expected return on a risky asset minus the expected return on fixed income and cash. Similar to the results for expected real return, we find a positive relation between past returns and expected risk premium of around 27 basis points.

Columns 2–4 of Table 4 focus separately on public equity, real assets, and private equity, which are the largest (and more homogeneous) categories within risky assets. Column 2 shows that each percentage point higher whole-portfolio past return leads to an assumption going forward that equity returns will be 17 basis points higher. Column 3 shows stronger effects within real assets: an additional percentage point of overall past return increases the assumed return on real assets by 31 basis points. The past return on real assets has an additional extrapolative effect of 15 basis points. Finally, in column 4, we find that each percentage point higher whole-portfolio past return and past return in private equity lead to 34 and 19 basis points higher expected return on private equity, respectively. Columns 6–8 of Table 4 estimate Equation (5) focusing only on single asset classes. Similar to the results on expected real returns, we find that the expected risk premium on an asset class level is positively related to the overall past return on the whole portfolio as well as to the past return in the analyzed asset class.

Contrary to the results in Table 3, we find that the relative value of unfunded liabilities is generally not significantly related to the expected real return or expected risk premium. This confirms that pension funds in states with large unfunded liabilities relative to their resources tend to justify higher return expectations primarily using a higher inflation assumption or a higher expected risk-free rate of return. This mechanism enables fiscally constrained pension plans to assume higher expected returns in all asset classes. We conclude that fiscal incentives and past return extrapolation appear to be distinct phenomena, since past return and fiscal pressure variables are related to different components of the portfolio expected return.

In sum, the positive relation between portfolio expected return and past performance operates through the expected real return or risk premium for investing in risky assets. We also find evidence of extrapolation of past returns both in public equity and in alternative assets traded in private markets.

3. Hypotheses and Explanations of the Extrapolation of Past Performance

We posit four rational hypotheses that could explain why the return expectations are a function of prior performance. The first hypothesis is that the positive relation between past returns and future expected returns reflects persistent cross-sectional differences in investment skill among pension funds. The second

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hypothesis is that the extrapolation of past returns captures higher risk taking in the past that will continue in the future. The third hypothesis is that pension plans make asset allocation decisions in order to reduce costly rebalancing, and the incentive to do so drives the correlation between past performance and future expectations, which are set to justify the avoidance of rebalancing. The fourth hypothesis is that pension plans extrapolate past returns when forming future expectations to reduce the amount of reported unfunded pension liabilities and fiscal pressure. In this section, we examine the empirical support for these hypotheses.

3.1 Cross-sectional performance persistence

The first hypothesis is that the positive relation between past returns and future expectations could be justified if there is long-term cross-sectional persistence in the investment skill among pension funds. If a pension fund that outperforms other pension funds in a given asset class in a given year also tends to do so in future years, this pension fund could formulate future expected returns in this asset class as a function of relative cross-sectional past performance. However, aggregate time-series predictability on an asset class level based on the dividend yield or other indicator (Brennan, Schwartz, and Lagnado 1997; Campbell and Viceira 1999; Barberis 2000) should be the same for all pension funds and should not lead to cross-sectional differences in return expectations.

One challenge for our analysis of persistence in pension plan performance is that the expected returns refer to a long, 20–30 year horizon and we do not observe realized returns over a sufficiently long horizon to test for this kind of long-horizon persistence. Our analysis relies on persistence tests on a shorter horizon of one to 10 years, as well as the findings of prior literature. We use specifications with year-reporting-month fixed effects to isolate aggregate time-series predictability on an asset class level and to focus only on cross-sectional performance persistence.

In columns 1–4 of Table 5, we estimate the persistence in overall pension plan performance on a different horizons using the lagged plan return from $t-1$, from $t-3$ to $t-1$, from $t-5$ to $t-1$, and from $t-10$ to $t-1$. We find only modest positive persistence on a short, one-year horizon, which is driven by pension fund performance in alternative asset classes. The results for the medium- and long-term persistence do not support the extrapolation of past performance based on persistence in overall pension plan performance; if anything, the negative coefficient on lagged returns from $t-5$ to $t-1$ suggests that pension plans with higher past performance underperform later. Pension fund performance is positively related to risk taking, as proxied by the strategic asset allocation to risky assets. We conclude that after controlling for risk taking, prior cross-sectional performance has no explanatory power for future cross-sectional performance on a medium- or long-term horizon, and hence does not justify the extrapolation of past returns when formulating return expectations.

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Table 5
Persistence in pension fund performance

Return in year t:	Overall pension plan level				Equity		Real assets		Private equity	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Lags in average past return:	t-1	t-3	t-5	t-10	t-1	t-3	t-1	t-3	t-1	t-3
Past return	0.105**	−0.017	−0.204*	0.095	−0.034	−0.315*	0.202***	0.220	0.128	0.053
	[0.041]	[0.055]	[0.107]	[0.095]	[0.078]	[0.187]	[0.054]	[0.140]	[0.103]	[0.134]
PF size	0.199**	0.226**	0.238***	0.234***	−0.108	−0.214*	0.735**	0.399	0.752**	0.280
	[0.092]	[0.090]	[0.089]	[0.086]	[0.071]	[0.112]	[0.303]	[0.290]	[0.368]	[0.316]
%Equity	0.066***	0.071***	0.082***	0.102***						
	[0.015]	[0.015]	[0.016]	[0.012]						
%Real assets	0.001	−0.005	−0.018	0.042***						
	[0.014]	[0.015]	[0.015]	[0.014]						
%Private equity	0.093***	0.100***	0.108***	0.101***						
	[0.013]	[0.014]	[0.015]	[0.016]						
%Hedge funds	0.010	0.006	0.004	0.006						
	[0.018]	[0.019]	[0.019]	[0.013]						
%Other risky assets	0.055***	0.055***	0.058***	0.063***						
	[0.015]	[0.017]	[0.019]	[0.014]						
Year × Reporting month FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	2,730	2,542	2,206	1,161	2,409	2,005	1,611	1,300	1,441	1,154
Adjusted R-squared	0.967	0.968	0.969	0.961	0.970	0.971	0.790	0.790	0.701	0.851

This table presents results of regressions in which the dependent variable is pension fund performance in year t. In columns 1–4, we examine persistence in performance on an overall pension fund level. In columns 5–10, we examine persistence in performance on an asset class level, and we focus on equity, real assets, and private equity. The *Lags in average past performance* line reports the number of lagged observations used to estimate the average past performance. When analyzing persistence on an overall pension plan level, we control for the performance in the previous year (t-1), average performance in the previous three years (t-3 to t-1), average performance in the previous five years (t-5 to t-1), and average performance in the previous ten years (t-10 to t-1). When analyzing persistence on an asset class level, we control for the performance in the previous year (t-1) and average performance in the previous three years (t-3 to t-1). *PF size* is the natural logarithm of pension fund assets under management. When analyzing overall persistence, we control for the asset allocation in year t. *%Equity*, *%Real assets*, *%Private equity*, *%Hedge funds*, and *%Other risky assets* measure the percentage allocated to different risky asset classes (the omitted categories are fixed income and cash). The specifications include year-reporting-month fixed effects. Standard errors clustered by pension plan are reported in brackets. $^*p<.1$; $^{**}p<.05$; $^{***}p<.01$.

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In columns 5–10 of Table 5, we estimate persistence in performance on an asset class level. We focus on public equity, real assets, and private equity, and report the results only using the lagged plan returns from $t-1$ and from $t-3$ to $t-1$. For the estimation of cross-sectional performance persistence by asset class, the sample size decreases significantly if we extend the analysis to a longer horizon. There is no evidence of persistence in pension plan performance in public equity, which suggests that the extrapolation of past returns in this asset class is not justified. This result is in line with prior literature: pension funds have no skills in their selection and termination decisions of public equity asset managers (Goyal and Wahal 2008; Jenkinson, Jones, and Martinez 2016), and there is no performance persistence in institutional public equity mandates (Busse, Goyal, and Wahal 2010). The fact that extrapolation occurs strongly within public equities shows that the persistent investment skills hypothesis cannot fully explain our findings. There is no evidence to support the idea that it would be justified for a pension plan A to expect 30 basis points higher return in public equity than pension plan B in the future because pension plan A had 1 percentage point higher past return in public equity.

We find that the positive one-year persistence on an overall plan level is due to persistence within alternative assets, such as real assets and private equity. The persistence in pension plan performance in alternative assets may be due either to genuine variation in skills in alternative assets (Cavagnaro et al. 2019) or to less frequent fair valuation of alternative private investments and smoothing of the reported returns. Smoothing bias due to the appraisal process and partial adjustments of valuations to market prices leads to serial correlation in reported returns on private investments. Smoothing bias is documented in the returns of institutional investors in real estate, hedge funds, and private equity (e.g., Geltner 1991; Getmansky, Lo, and Makarov 2004; Ang et al. 2018).

In Internet Appendix Table A.7, we explore further whether the persistent skills hypothesis can explain our findings in private equity. For every pension plan, we calculate the average performance separately for investments in private equity funds made three to eight years ago, nine to 13 years ago, and more than 13 years ago, respectively. We find that the performance of the oldest funds has predictive power for the expected risk premium even though the decision to invest in these funds was made in the distant past. We would expect the performance of the middle group of funds to be most predictive in a skills framework, since there is sufficient time to incorporate cash distributions into the reported returns but the investment decision was made somewhat more recently. However, pension funds do not appear to incorporate the performance of the medium-term private equity funds into their expected risk premiums. Furthermore, we observe that pension plans with less experience in private equity expect higher future returns, contrary to the established positive link between limited partners' experience in this asset class and returns

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(Lerner, Schoar, and Wongsunwai 2007; Sensoy, Wang, and Weisbach 2014; Dyck and Pomorski 2016). These results indicate that the persistent skills hypothesis cannot fully explain our findings.

3.2 Higher risk taking

The second hypothesis is that the extrapolation of past returns when forming future expectations captures higher risk taking in the past that will continue also in the future. In the analysis throughout, we always control for percentage allocated to risky assets. These variables capture differences in risk taking based on differences in asset allocation policy, but they do not capture differences in risk taking within an asset class. If some pension funds generally take more risk within asset classes than others, they might also expect higher future returns. To capture risk taking within asset classes, we control for the standard deviation of past performance.

In Internet Appendix Table A.8, we examine further whether the positive relation between past performance and portfolio expected return is due to higher risk taking. We replicate Tables 2 and 4, but instead of controlling for the standard deviation of past returns, we include beta coefficients as measures of risk taking. We estimate these beta coefficients separately for every pension plan using the annual returns in the previous 10-year period. The robustness analysis includes either only a market beta coefficient estimated using the capital asset pricing model (CAPM), or market, small minus big (SMB), and high minus low (HML) betas estimated using the Fama-French model. Our results show that the past standard deviation is a good measure of risk taking, since the correlation between past standard deviation and CAPM market beta is 0.84. The relation between expected returns and market beta coefficient is negative and insignificant. Only the coefficients on HML betas are positive and significant, but even when controlling for them and the SMB betas, the coefficient on *Past return* remains statistically significant and around 30 basis points. This robustness test confirms that the positive relation between portfolio expected return and past performance is not due to higher risk taking.

One potential additional concern is that if higher past performance is due to higher risk taking, we should not necessarily expect that the risk variables will drive out the past performance variable, especially given the relatively short time series of annual returns. In Internet Appendix Table A.9, we control for past standard deviation or market betas but do not include the past return variable. We find that the coefficients on past standard deviation as well as market betas remain negative and insignificant in all models and that they do not explain the positive relation between past performance and future expectations. Overall, cross-sectional differences in risk taking cannot explain why the return expectations are a function of prior performance.

3.3 Costly rebalancing

The third hypothesis is that the positive relation between past returns and future expectations reflects pension fund asset allocation decisions implemented

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to reduce costly rebalancing. This hypothesis argues that past performance mechanically affects the actual allocation weights, so pension plans link their return expectations to past returns just to avoid rebalancing the portfolio. Our analysis throughout relies on target allocation weights, which are not directly subject to mechanical changes from market movements.[14] However, a pension fund can potentially reduce costly rebalancing by adjusting the target asset allocation weights in the direction of the actual asset allocation weights. This adjustment implicitly creates a positive relation between the target asset allocation and past returns. Thus, the reduce-costly-rebalancing hypothesis implies that pension plan return expectations do not necessarily reflect pension plan true expectations.

We test the reduce-costly-rebalancing hypothesis against an alternative hypothesis that the disclosed pension plan expected returns reflect their true beliefs and the target asset allocation weights reflect their true future strategic asset allocation policy. Distinguishing between the two hypotheses is challenging since both predict that there will be a positive relation between the stated asset class return expectations and past performance. We distinguish between them with tests of the direction of the changes in actual asset allocation weights versus changes in target asset allocation weights.

Under the reduce-costly-rebalancing hypothesis, pension plans will adjust ex post their target asset allocation weights in the direction of their actual weights, which enables them to justify their current asset allocation and reduce the costs for portfolio rebalancing. For example, if the target asset allocation weight in year $t-1$ is higher than the actual asset allocation weight in year $t-1$, a pension plan will reduce the target asset allocation in year t in order to shrink the distance between the target and actual allocation without rebalancing the portfolio. In contrast, the true-beliefs hypothesis predicts that pension plans will adjust the actual asset allocation weights toward the target weights since the target weight reflects their true long-term asset allocation strategy and pension plans make investment decisions in order to implement this strategy. For example, if the target asset allocation weight in year $t-1$ is higher than the actual asset allocation weight in year $t-1$, a pension plan will increase the actual asset allocation in year t as it implements the new policy (Internet Appendix Figure A.2 provides a stylized example of the empirical predictions).

We test the predictions of both hypotheses using the following empirical setting. The explanatory variable $\omega t_{i,t-1} - \omega a_{i,t-1}$ is the same for both hypotheses. It captures the distance between the target weight in asset class i in year $t-1$ ($\omega t_{i,t-1}$) and the actual weights in asset class i in year $t-1$

[14] Analyzing only the connection between actual (realized) asset allocation and past performance is insufficient because the dependence of asset allocation on past returns could be explained if pension funds act similarly to individuals and do not fully rebalance their portfolios (Calvet, Campbell, and Sodini 2009). For example, Rauh (2009) documents that after the technology crash in 2000, corporate pension funds allowed the share allocated to equities to drift downward. Such a finding would be consistent with status quo bias in asset allocation (Samuelson and Zeckhauser 1988), inertial investing (Choi et al. 2002), or costs to portfolio rebalancing.

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($\omega a_{i,t-1}$). For the reduce-costly-rebalancing hypothesis, the dependent variable $\omega t_{i,t} - \omega t_{i,t-1}$ captures the change in the *target* weight in asset class i between year $t-1$ and year t. This version of the reduce-costly-rebalancing hypothesis predicts that b_1 will have a negative coefficient in this estimation so the target weight converges to the actual weight:

$$\omega t_{i,t} - \omega t_{i,t-1} = \alpha_t + b_1 \left(\omega t_{i,t-1} - \omega a_{i,t-1}\right) + \mathbf{\Gamma}'\mathbf{X} + \varepsilon_{it} \tag{6}$$

For the true-beliefs hypothesis, the dependent variable $\omega a_{i,t} - \omega a_{i,t-1}$ captures the change in the actual weight in asset class i between year $t-1$ and year t. The true-beliefs hypothesis predicts that b_1 will have a positive coefficient in the following estimation, with the result being that the actual weight converges to the target weight:

$$\omega a_{i,t} - \omega a_{i,t-1} = \alpha_t + b_1 \left(\omega t_{i,t-1} - \omega a_{i,t-1}\right) + \mathbf{\Gamma}'\mathbf{X} + \varepsilon_{it} \tag{7}$$

This analysis requires data on actual pension fund allocation, in addition to the strategic target allocation. From the pension fund annual reports, we obtain information on both actual and target allocations over at least a 10-year horizon for 209 out of the 228 pension plans that disclose return expectations. We estimate the specification for overall allocation to risky assets, as well as separately for the allocation to equity, real assets, and private equity. Recognizing that the entire convergence does not need to occur in one period, we also estimate a robustness test by including a lagged term capturing differences between the target and actual asset allocation in year $t-2$.

Table 6 presents the results for both hypotheses. In panel A, we find some evidence in line with the reduce-costly-rebalancing hypothesis since the coefficient for overall risky assets is −0.10. This suggest that target weights to some extent move in the direction of previous year actual weights, but the economic magnitude is small, and incentives to reduce rebalancing costs do not seem to be the main driver of target asset allocation decisions. In panel B, we find significantly stronger evidence in line with the true-beliefs hypothesis since the changes in actual asset allocation are positively related to the difference between the target weight in the previous year and the actual weight in the previous year. The coefficient for overall risky assets is 0.41. Comparison of the coefficients indicates that the reliance on past returns in setting expected returns and target asset allocations cannot be fully explained by efforts to reduce costly rebalancing and rationalize past mechanical changes in actual allocation weights.

However, this test still leaves open the possibility that pension funds may adjust target allocation weights ex ante in anticipation of future mechanical changes in actual allocation weights. This version of the reduce-costly-rebalancing hypothesis also predicts that the actual asset allocation weights move in the direction of the target asset allocation weights. We distinguish between the second version of the reduce-costly-rebalancing hypothesis and the true-beliefs hypothesis by decomposing the changes in the actual asset

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Table 6
Changes in actual and target asset allocation

	All risky		Equity		Real assets		Private equity	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
A. Changes in target asset allocation								
$\omega t_{i,t-1} - \omega a_{i,t-1}$	−0.099***		−0.057***		−0.110***		−0.019	
	[0.015]		[0.012]		[0.022]		[0.013]	
$\omega t_{i,t-2} - \omega a_{i,t-2}$		0.001		0.005		−0.070***		0.009
		[0.021]		[0.023]		[0.019]		[0.015]
PF size	−0.027	−0.064	0.112*	0.083	−0.075**	−0.081**	0.014	0.025
	[0.048]	[0.049]	[0.059]	[0.059]	[0.036]	[0.036]	[0.028]	[0.029]
Year × Reporting month FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	2,187	1,987	2,187	1,987	2,187	1,987	2,187	1,987
Adjusted *R*-squared	0.093	0.087	0.091	0.079	0.077	0.065	0.097	0.098
B. Changes in actual asset allocation								
$\omega t_{i,t-1} - \omega a_{i,t-1}$	0.407***		0.289***		0.227***		0.156***	
	[0.030]		[0.023]		[0.021]		[0.012]	
$\omega t_{i,t-2} - \omega a_{i,t-2}$		0.083***		0.153***		0.090***		0.180***
		[0.017]		[0.018]		[0.016]		[0.015]
PF size	−0.224***	−0.105**	−0.123**	−0.050	0.008	−0.018	0.051*	0.057**
	[0.062]	[0.044]	[0.058]	[0.060]	[0.036]	[0.029]	[0.027]	[0.028]
Year × Reporting month FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	2,187	1,987	2,187	1,987	2,187	1,987	2,187	1,987
Adjusted *R*-squared	0.331	0.128	0.477	0.403	0.282	0.201	0.149	0.161
C. Mechanical changes in actual asset allocation								
$\omega t_{i,t-1} - \omega a_{i,t-1}$	0.009*		−0.004		0.013		−0.006	
	[0.005]		[0.004]		[0.008]		[0.007]	
$\omega t_{i,t-2} - \omega a_{i,t-2}$		0.010*		0.003		0.002		0.019*
		[0.006]		[0.004]		[0.009]		[0.010]
PF size	0.026	0.024	−0.116***	−0.085***	0.038**	0.025	0.050**	0.036
	[0.021]	[0.023]	[0.026]	[0.027]	[0.015]	[0.017]	[0.023]	[0.023]
Year × Reporting month FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1,750	1,601	1,750	1,601	1,750	1,601	1,750	1,601
Adjusted *R*-squared	0.912	0.911	0.916	0.920	0.531	0.538	0.374	0.399
D. Residual changes in actual asset allocation								
$\omega t_{i,t-1} - \omega a_{i,t-1}$	0.364***		0.320***		0.243***		0.176***	
	[0.034]		[0.028]		[0.033]		[0.018]	
$\omega t_{i,t-2} - \omega a_{i,t-2}$		0.110***		0.140***		0.010		0.172***
		[0.021]		[0.016]		[0.020]		[0.017]
PF size	−0.233***	−0.107*	0.007	0.058	−0.067	−0.051	−0.029	−0.007
	[0.069]	[0.059]	[0.069]	[0.065]	[0.051]	[0.045]	[0.032]	[0.033]
Year × Reporting month FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1,750	1,601	1,750	1,601	1,750	1,601	1,750	1,601
Adjusted *R*-squared	0.442	0.301	0.317	0.165	0.178	0.084	0.189	0.188

This table analyzes the changes in allocation to overall risky assets as well as separately the changes in allocation to public equity, real assets, and private equity. In panel A, the dependent variable is the difference between the target weight in asset class *i* in year *t* and the target weight in asset class *i* in year *t-1*. In panel B, the dependent variable is the difference between the actual weight in asset class *i* in year *t* and the actual weight in asset class *i* in year *t-1*. In panels C and D, we decompose the change in the actual weights into mechanical and residual changes. The explanatory variable $\omega t_{i,t-1} - \omega a_{i,t-1}$ is the same in all panels and captures the difference between the target weight in asset class *i* in year *t-1* and the actual weights in asset class *i* in year *t-1*. We also add lagged values of this variable for *t-2* period. *PF size* is the natural logarithm of pension fund assets under management. The specifications include year-reporting-month fixed effects. Standard errors clustered by pension plan are reported in brackets. $*p<.1$; $**p<.05$; $***p<.01$.

allocation into mechanical and residual changes. The second version of the reduce-costly-rebalancing hypothesis predicts that the mechanical changes in actual asset allocation explain how the actual asset allocation weights converge to the target weights. For example, a pension fund may expect a higher return on risky assets than on safe assets and may define target weights in advance

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that account for the anticipated mechanical increase in allocation to risky assets over time. In this case, if the target weight on risky assets in year $t-1$ is higher than the actual weight on risky assets in year $t-1$, this difference will predict a mechanical increase in the actual weight in year t. The true-beliefs hypothesis argues that pension funds implement active decisions to reach their strategic asset allocation policy, so nonmechanical residual changes explain how the actual weights converge to the target weights. For example, a pension fund may actively reduce its allocation to safe assets and increase its allocation to risky assets in order to implement a riskier target asset allocation policy. In this case, if the target weight on risky assets in year $t-1$ is higher than the actual weight on risky assets in year $t-1$, a pension plan will implement active (nonmechanical) investment decisions to increase the actual weight in year t.

We test the predictions of both hypotheses using the following empirical setting. The explanatory variable $\omega t_{i,t-1} - \omega a_{i,t-1}$ remains again the same for both hypotheses, but we decompose the change in the actual weight in asset class i between year $t-1$ and year t, $\omega a_{i,t} - \omega a_{i,t-1}$ into mechanical and residual changes. For the mechanical change in actual weight in asset class i between year $t-1$ and year t, we assume that the dollar amount invested in an asset class in year t is determined mechanically by the dollar amount invested in the asset class in year $t-1$ and the pension plan return in this asset class from $t-1$ to t, $(\omega a\_mechanical_{i,t} - \omega a_{i,t-1})$. This alternative version of the reduce-costly-rebalancing hypothesis predicts that b_1 will have a positive coefficient in this estimation, with the result being that the actual weight mechanically converges to the target weight:

$$\omega a\_mechanical_{i,t} - \omega a_{i,t-1} = \alpha_t + b_1\left(\omega t_{i,t-1} - \omega a_{i,t-1}\right) + \mathbf{\Gamma}'\mathbf{X} + \varepsilon_{it} \qquad (8)$$

The residual change in actual weight in asset class i between year $t-1$ and year t equals the difference between the actual weight in asset class i in year t and the mechanical weight in asset class i in year t, $(\omega a_{i,t} - \omega a\_mechanical_{i,t})$. The residual change captures changes in portfolio weights due to active rebalancing investment decisions, shifts from investing new contribution inflows, and shifts from selling decisions to pay pension benefits. The true-beliefs hypothesis predicts that b_1 will have a positive coefficient in the following estimation, with the result being that the actual weight converges to the target weight through active residual investment decisions:

$$\omega a_{i,t} - \omega a\_mechanical_{i,t} = \alpha_t + b_1\left(\omega t_{i,t-1} - \omega a_{i,t-1}\right) + \mathbf{\Gamma}'\mathbf{X} + \varepsilon_{it} \qquad (9)$$

These tests require data for the allocations and returns in all asset classes in year t and $t-1$. From the 2,187 observations of actual changes in asset allocation policy in panel A of Table 6, we are able to decompose 1,750 observations into mechanical and residual changes. Our results in panels C and D of Table 6 show that the actual allocation converges toward the target asset allocation policy through the residual changes in the asset allocation, and not through the mechanical changes. The difference between the target

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weight in asset class i in year $t-1$ and the actual weights in asset class i in year $t-1$ predicts the residual changes in asset allocation (panel D), while the relation with mechanical changes is economically insignificant (panel C). Overall, target weights seem to reflect pension plans' true long-term portfolio strategy. Pension plans make active investment decisions in order to implement this strategy, while the anticipation of future mechanical movements in the asset allocation plays a limited role.

In Internet Appendix Table A.10, we estimate a robustness test that illustrates the limited explanatory power of the reduce-costly-rebalancing hypothesis. The intuition behind this estimation is that the mechanical changes in asset allocation are a quadratic function of the weight on risky assets (see also Parker, Schoar, and Sun 2020 on rebalancing strategies of target-date-funds). A pension fund holding a portfolio of 50% risky assets and 50% safe assets is most sensitive to returns on the stock market. There are no mechanical changes if a portfolio consists of either 100% risky assets or 100% safe assets. Based on this intuition, we argue that if the use of past performance in forming return expectations is simply a mechanism to reduce costly rebalancing, pension funds closer to the 50/50 asset allocation should be more likely to rely on past returns, as they are more sensitive to the mechanical changes in percentage terms than pension funds closer to either extreme.

We test this hypothesis in Internet Appendix Table A.10 by splitting pension funds into terciles based on their actual allocation to risky assets. The bottom group has a mean allocation to risky assets of 68%, the middle group 75%, and the high group 81%. We replicate the estimates from Tables 2 and 4 separately for each tercile. The positive relation between return expectations and past performance is economically and statistically strongest in the tercile with the highest allocation to risky assets, where the allocation is farthest away from 50/50. It is weakest in the tercile with the lowest allocation to risky assets, where the allocation is closest to 50/50. This result is the opposite of what the reduce-costly-rebalancing hypothesis would predict, since pension funds closer to the 50/50 asset allocation are the ones whose mechanical changes in allocation are most sensitive to market movements.

We also point out that the reduce-costly-rebalancing hypothesis predicts that only recent returns should matter, and pension funds should not extrapolate return realizations from the more distant past. Internet Appendix Table A.4 explores different weighting scheme of returns over the past 10 years. Column 5 shows a positive relation between return expectations and past returns even when we put more weight on distant past returns.

In conclusion, we find some support for the reduce-costly-rebalancing hypothesis, which argues that pension funds might extrapolate past performance in order to adjust target allocation weights ex post and rationalize past mechanical changes. However, the reliance on past returns in setting expected returns and target asset allocations cannot be fully reconciled with the

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reduce-costly-rebalancing hypothesis, since pension plans make economically significant active investment decisions.

3.4 Fiscal incentives

We find that unfunded liabilities are positively related to the portfolio expected return, consistent with the hypothesis that fiscally stressed governments have incentives to maintain higher expected rates of return to justify a higher pension discount rate. The effect of fiscal incentives from unfunded liabilities operates primarily through an effect on inflation assumptions or expected risk-free rate, and it does not explain the positive relation between past returns and future real return expectations. Our explanation is that this mechanism enables fiscally constrained pension plans to assume higher expected returns in all asset classes. In this section, we discuss three robustness tests that confirm that fiscal incentives and past return extrapolation are distinct effects, and that fiscal pressure is not the main driver of return extrapolation.

First, in Internet Appendix Table A.11, we include an interaction term between the measures of past return and unfunded liabilities. We find no evidence that past returns positively affect the portfolio expected return when the unfunded liabilities are large. Second, in Internet Appendix Table A.12, we focus on the subsample of pension plans with below-median unfunded liabilities instead of including an interaction term. If fiscal pressure is the main reason to extrapolate past performance, we should observe a weaker (or insignificant) relation between the past return and future expectations in this subsample of better funded pension plans. However, even within this subsample, the coefficient on past return remains around 25 basis points and significant, and the positive relation exists only with expected real rate of return and not with the expected inflation rate. Third, in Internet Appendix Table A.13 we directly control for the pension discount rate, whose value affects required contribution payments and the state budget. We find that the results on the relation between past return and expected return remain qualitatively unchanged. That is, even for a given pension discount rate, there is just as much evidence of past return extrapolation in the portfolio expected return.

In addition to fiscal incentives, pension plans may also have incentives to display favoritism toward certain asset managers, perhaps for political reasons or impact investing objectives (Andonov, Hochberg, and Rauh 2018; Barber, Morse, and Yasuda 2021). Although displays of favoritism might be more likely in transferring assets from one manager to another within an asset class without changing the strategic asset allocation, it is possible that such favoritism might affect the allocation of assets across asset classes. However, we believe that favoritism toward certain asset managers cannot explain our extrapolation results because it would bias our results downward. Pension plans displaying favoritism in the selection of asset managers based on political and other nonfinancial objectives might well obtain lower returns, but these pension plans would not reduce their future target allocation weights after

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observing lower past returns. Thus, investing based on favoritism toward certain underperforming managers would if anything reduce the cross-sectional differences in target allocation weights between pension plans with high performance and pension plans with low performance.

We conclude that fiscal incentives from unfunded liabilities are important in explaining the differences in expected inflation rate of risk-free rate of return. However, past return and fiscal pressure variables are related to different components of the portfolio expected return, so the fiscal incentives and reliance on past returns in setting expected returns appear to be distinct phenomena.

4. The Role of Executive Director Experience

In the previous section, we document that the positive relation between past returns and future expected returns cannot be fully explained by performance persistence, differences in risk taking, efforts to reduce costly rebalancing, and fiscal incentives. This suggests that the extrapolation of past returns when forming future expectations might reflect at least in part excessive expectations among pension funds. The excessive extrapolation hypothesis argues that pension funds put an excessive (too high) weight on their own past returns, and this leads to extrapolation of cross-sectional differences in pension fund past performance. Based on the prior literature on extrapolation among individual investors (e.g., Vissing-Jorgensen 2003; Malmendier and Nagel 2011; Kuchler and Zafar 2019), we hypothesize that pension fund executives with a longer tenure are more prone to extrapolate past returns because they have personally experienced these returns. Our analysis focuses on pension fund investment executives because they lead the process of designing target asset allocation and formulating return expectations.

To test the role of personal experience, we collect data on the highest-ranking staff member of the pension fund responsible for making investment decisions—namely, the executive director or CEO. If the investment decisions for one or multiple pension plans are made by a separate entity (investment board), we collect data about the executive director of this investment entity. For example, the State of Wisconsin Investment Board (SWIB) manages the assets of the Wisconsin Retirement System (WRS) and "is responsible for setting long-term investment policies, asset allocation, benchmarks, and fund level risk, and monitoring investment performance" (State of Wisconsin Investment Board 2016).

We collect data on pension fund investment executives from the pension fund CAFRs and websites, as well as newswire articles on the Pensions & Investments website. Using these sources, we document exact starting and ending dates for each CEO and estimate the tenure of the executive for each pension fund-year observation. The variable *CEO tenure* measures the tenure of the executive director in years at the fiscal-year ending date when the pension fund expectations are reported. In nine investment institutions (covering 35

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Table 7
Portfolio expected return and executive directors

	Portfolio expected return					
	(1)	(2)	(3)	(4)	(5)	(6)
Geometric	−0.740***	−0.889***	−0.882***	−0.997***	−0.889***	−0.893***
	[0.115]	[0.127]	[0.127]	[0.128]	[0.120]	[0.122]
Past return	0.304***	0.286***	0.196***		0.192**	0.112
	[0.064]	[0.064]	[0.075]		[0.081]	[0.083]
CEO tenure		0.034***	−0.033		−0.045**	−0.029
		[0.006]	[0.023]		[0.022]	[0.022]
CEO tenure × Past return			0.010***		0.010***	0.008**
			[0.003]		[0.003]	[0.004]
CEO past return adjusted				0.343***		
				[0.078]		
CEO treasury		0.321**	0.324**	0.292*	0.365*	0.333*
		[0.154]	[0.152]	[0.154]	[0.188]	[0.178]
CEO interim		0.219**	0.223**	0.269***	0.120	0.138
		[0.106]	[0.106]	[0.103]	[0.104]	[0.107]
Past standard deviation	−0.059	−0.001	0.008	0.011	−0.009	0.003
	[0.042]	[0.041]	[0.041]	[0.040]	[0.044]	[0.044]
Unfunded liability / GSP	1.393***	1.303***	1.322***	1.272***	1.696***	1.330***
	[0.442]	[0.416]	[0.414]	[0.439]	[0.439]	[0.435]
GSP per capita	0.006*	0.010***	0.010***	0.017***	0.007**	0.007**
	[0.003]	[0.003]	[0.003]	[0.004]	[0.003]	[0.003]
PF size	−0.118**	−0.080*	−0.069	−0.071	−0.034	−0.037
	[0.050]	[0.047]	[0.047]	[0.046]	[0.047]	[0.049]
CEO tenure FE	No	No	No	Yes	No	No
Consultant FE	No	No	No	No	Yes	Yes
Consultant FE × Past return	No	No	No	No	No	Yes
Asset allocation controls	Yes	Yes	Yes	Yes	Yes	Yes
Year × Reporting month FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	883	883	883	883	883	883
Adjusted R-squared	0.324	0.383	0.389	0.468	0.431	0.456

This table presents results of regressions in which the dependent variable is the portfolio expected return. *Geometric* is an indicator variable for pension plans reporting geometric expected return. *Past return* measures the average arithmetic return in the previous 10-year period. *CEO tenure* measures the tenure of the executive director in years at the fiscal-year ending date when the pension fund expectations are reported. *CEO past return adjusted* is an alternative measure to calculate the average arithmetic return experienced by the CEOs. This measure takes into account the returns personally experienced by the CEO in the pension plan, but if the CEO has a tenure of less than 10 years, it replaces the earlier missing returns before the appointment of the CEO with the average return of all pension plans reporting in the same month. In this specification, we include CEO tenure fixed effects. *CEO treasury* is an indicator variable for pension funds that do not have investment staff members and are managed by the state treasurer. *CEO interim* is an indicator variable for executive directors who were appointed initially as interim directors. Column 5 includes general investment consultant fixed effects. Column 6 controls also for interaction terms between the general investment consultants and past returns. All specifications include year-reporting-month fixed effects. Standard errors clustered by pension plan are reported in brackets. $^{*}p<.1$; $^{**}p<.05$; $^{***}p<.01$.

pension plans), the state treasury or related office manages investment policy, as opposed to a separate executive director. We define *CEO treasury* as an indicator variable for pension funds that do not have investment staff members and are instead managed by the state treasurer and the associated office. *CEO interim* is an indicator variable for executive directors who were appointed initially as interim directors.

Column 1 of Table 7 replicates the baseline analysis of the relation between portfolio expected return and past return for the sample of pension plans with

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data on the executive director.[15] We still find that each additional percentage point of average past return increases the overall expected return by 30 basis points. Our objective is to examine whether the coefficient on past returns can be absorbed by an interaction term with CEO tenure. In all models, we control for the percentage allocation to risky assets as well as fiscal incentives.

Column 2 shows the relations between the CEO variables and portfolio expected return without including an interaction term. For each additional year of CEO tenure, the expected return is 3.4 basis points higher. When the executive director is the treasurer, expected returns are higher, perhaps due to the state government's desire to maintain high discount rates. When the executive director starts with an interim appointment, expected returns are also higher.

Column 3 adds an interaction term of CEO tenure with past return. Each additional year of tenure increases the sensitivity of portfolio expected return to past return by 1 basis point. This interaction term reduces the economic significance of the baseline relation between past returns and future expectations. For instance, the average executive tenure is around seven years and the standard deviation is around eight years. Based on the estimated coefficients, each additional percentage point of past return increases the expected return by 20 basis points for a new CEO, by 27 basis points for a CEO with mean tenure, and by 35 basis points for a CEO with tenure that is one standard deviation above the mean. A pension fund of a CEO with 20 years' tenure, a level which is exceeded in only four cases, would increase the expected return by 40 basis points for each additional percentage point of past return.[16]

Since the past return variable is the 10-year past return of the fund regardless of the CEO's tenure, one concern is that the interaction term between CEO tenure and past return might be an insufficient proxy for past performance personally experienced by the executive director. To address this concern, we follow the insights of Malmendier and Nagel (2011) and construct additional measures of past returns personally experienced by the CEO. *CEO past return adjusted* calculates the average of past returns personally experienced by the CEO in the pension plan, but if the CEO has a tenure of less than 10 years, it replaces the earlier missing returns before the appointment of the CEO with the average return of all pension plans reporting in the same month. For example, if a CEO has a tenure of five years as an executive director of a pension plan, *CEO past return adjusted* will be an arithmetic average of the five annual returns personally experienced by the CEO in this pension plan and the average return of all pension plans reporting in the same month for the five annual returns

15 Internet Appendix Table A.14 presents a robustness test of Table 7 where the dependent variable is the pension plan expected real rate of return instead of overall expected return. The results and conclusions remain the same.

16 In Internet Appendix Table A.17 we show that our results are robust to excluding the four executive directors with tenure longer than 20 years.

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that the CEO did not personally experience. This measure ensures that the past return of all CEOs is estimated using a time series of 10 annual observations.[17]

CEO past return adjusted has a somewhat lower standard deviation even than *Past return* at the fund level, since for years before the current CEO's tenure of each fund, the same returns are being averaged in for every fund, and so this is particularly true for situations of short CEO tenure. In Internet Appendix Table A.15, we compare the summary statistics for these measures of past performance across different subsamples of CEO tenure.

In column 4 of Table 7, we replace *CEO tenure* and the interaction term simply with *CEO past return adjusted*. For each percentage point difference in the CEO's experienced past return as reflected in the *CEO past return adjusted* variable, the portfolio expected return is higher by 34 basis points.[18] This estimate is close to the original estimates using the fund *Past return*, suggesting that past returns experienced by the CEOs are indeed driving the estimated effects of past returns on future expectations.

Another potential concern is that the interaction term of executive directors' tenure and past return actually proxies for decisions made by other agents involved in the process of forming return expectations—namely, investment consultants, from whom the pension fund's professional staff receive advice and reports (Jenkinson, Jones, and Martinez 2016). For example, the Louisiana State Employee Retirement System (LASERS) "works closely with its investment consultant to conduct a thorough asset allocation and liability review on an annual basis" (Louisiana State Employees Retirement System 2017). To address this concern, we collect data from the pension fund annual reports on the general investment consultants hired by each pension plan over the sample period.[19] Columns 5 and 6 of Table 7 present two robustness tests of the main specification in column 3.

In the first robustness specification, we augment our main specification with general investment consultant fixed effects. In column 5, we find that some consultant fixed effects are significant, but these relations do not affect our conclusion on the role of executive director personal experience, since the interaction term between CEO tenure and past return remains positive and

[17] An alternative measure, *CEO past return*, calculates average past returns using the annual returns personally experienced by the CEO regardless of the tenure length. However, especially for CEOs with short tenure, this variable will be heavily influenced by the second moment (volatility) of the return distribution. Indeed, the standard deviation of *CEO past return* is 5.97 percentage points for CEOs with less than two years of tenure and 2.14 percentage points for CEOs with two to 10 years of tenure. In Internet Appendix Table A.16 we present regressions using *CEO past return*, despite its imprecision. Given the addition of noise, it is unsurprising that the estimates are smaller and that they approach the estimates in Table 7 when we limit the sample to longer-tenured CEOs.

[18] In this estimation, we include CEO tenure fixed effects that correspond to the age fixed effects included by Malmendier and Nagel (2011) when analyzing the expectations of individual investors.

[19] The general investment consultant is involved in the determination of asset allocation policy and formulation of macroeconomic outlook. Pension funds usually have only one general investment consultant but may hire a number of other consultants who specialize in managerial selection within asset classes. Of the 228 pension plans in our sample, 221 use a general investment consultant.

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significant. Column 6 extends the analysis by adding a set of interaction effects between the investment consultant fixed effects and the past return variable. The baseline effect of past return declines to 11 basis points (about one third of its original value in column 1) and is insignificant. However, we still find that the interaction term between CEO tenure and past return remains positive and significant. Controlling for investment consultants seems to matter, but it does not explain why the relation between return expectations and past performance is stronger for pension funds led by investment executives with longer tenure.

Overall, CEO tenure is a distinct channel through which pension plans extrapolate past performance when formulating return expectations. Our results provide evidence consistent with the excessive extrapolation hypothesis by documenting that the relation between return expectations and performance is stronger in plans whose executives have personally experienced the pension plan's past returns.

5. Implications

5.1 The impact of return assumptions on target asset allocation

In this section, we study the extent to which return expectations based on past performance affect target asset allocation. Pension fund investment staff propose changes in target asset allocation policy to the fund's board and these recommendations reflect expected returns by asset class. Our setting allows us to observe the target allocation weights, which are not mechanically affected by past performance (on the contrary, actual allocation weights mechanically depend on past returns). We examine whether past returns matter for future target asset allocation by estimating the following specification for fund i at time t:

$$\Omega_{it} = \alpha_t + \mu * \overline{R_{i,t-j}} + \mathbf{\Gamma}'\mathbf{X} + \varepsilon_{it} \tag{10}$$

where Ω_{it} is the target allocation to all risky assets, which is an aggregation of equity, real assets, private equity, hedge funds, and other risky assets into one category of risky assets. We interpret a finding that $\mu > 0$ in the presence of year-by-reporting-month fixed effects (α_t) as evidence that cross-sectional variation in pension fund past returns affects the cross-section of target asset allocation.

We then consider the narrower asset classes of equity, real assets, and private equity separately, using a similar specification:

$$\omega_{it} = \alpha_t + \mu * \overline{R_{\omega,i,t-j}} + \nu * \overline{R_{i,t-j}} + \mathbf{\Gamma}'\mathbf{X} + \varepsilon_{it} \tag{11}$$

where ω_{it} represents the allocation to asset class ω by pension plan i at time t, and $R_{\omega,i,t-j}$ is the past return that pension plan i experienced in asset class ω over the previous j years. We interpret a finding that $\mu > 0$ as evidence that a pension fund's past returns in a given asset class affect the fund's allocation to that asset class. If $\nu > 0$ in the presence of time fixed effects (α_t), then there

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are spillovers to the allocation to asset class ω from the relative performance of pension fund i in other asset classes.

Next, we ask how the target allocation to risky assets is correlated with the expected risk premium of pension fund i at time t. Again, we perform this analysis either at the level of aggregated risky assets and the whole fund's expected risk premium

$$\Omega_{it} = \alpha_t + \lambda * E(R_{it} - r_{f,it}) + \mathbf{\Gamma}'\mathbf{X} + \varepsilon_{it}, \tag{12}$$

or at the level of the individual asset classes,

$$\omega_{it} = \alpha_t + \lambda * E(R_{\omega,it} - r_{f,it}) + \mathbf{\Gamma}'\mathbf{X} + \varepsilon_{it} \tag{13}$$

This analysis asks whether systems with higher expected returns on risky assets do in fact invest more in risky assets. Theory gives some guidance as to what we should expect from the parameter λ. In the standard myopic portfolio choice model reviewed by Campbell and Viceira (2002), the optimal allocation is:

$$\omega_t = \frac{\mathrm{E}r_{t+1} - r_{f,t-1} + \sigma_t^2/2}{\gamma\sigma_t^2}, \tag{14}$$

where $r = \ln(1+R), \gamma$ is the constant of relative risk aversion and σ is the volatility of the risky asset. If risky assets have $\sigma = 0.2$, then an increase in the risk premium by 1 percentage point increases allocation to the risky asset by $25/\gamma$ percentage points. For example, for a log investor with $\gamma = 1$, the expected change in allocation to risky assets when the risk premium increases by 1 percentage point is a full 25 percentage points. For a risk-average investor with $\gamma = 10$, the effect would be only 2.5 percentage points, and to explain an effect of only 1 percentage point would require $\gamma = 25$.

Combining the approaches in Equations (10) and (12), a two-stage least squares (2SLS) estimation reveals how variation in expected risk premium driven only by the past return affects the target allocation to asset class ω at time t by fund i. The second stage is simply Equation (12), while the first stage is Equation (5), which presents the relation between the expected risk premium and past return. The measure of λ using the two-stage estimation is the measure of the effect that return extrapolation has on target allocation through its effect on the expected risk premium.

Table 8 shows estimates of the reduced form relationship in Equation (10). Column 1 shows that each percentage point of past return increases the target allocation to risky assets by 2.0 percentage points. Columns 2–4 examine this relation for narrower asset classes as in Equation (11): public equity, real assets, and private equity. In column 2, we find very strong and statistically significant effects of both the past equity return and the past overall return on allocation to equities. A 1 percentage point higher past equity return increases the allocation to equity by 5.5 percentage points, even keeping the overall past portfolio return held constant. At the same time, a 1 percentage point higher overall past

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Table 8
Target allocation to risky assets and past return

	Target allocation to risky assets			
	All (1)	Equity (2)	RA (3)	PE (4)
Geometric	0.002	−0.015	−0.021***	−0.023***
	[0.010]	[0.013]	[0.007]	[0.005]
Past return	0.020***	0.049***	0.009	0.008***
	[0.004]	[0.008]	[0.005]	[0.003]
Past return equity		0.055***		
		[0.011]		
Past return RA			−0.000	
			[0.002]	
Past return PE				0.003***
				[0.001]
Past standard deviation	0.019***	−0.013**	0.013***	0.013***
	[0.005]	[0.005]	[0.003]	[0.003]
Unfunded liability / GSP	0.043	0.230***	−0.075**	−0.118***
	[0.041]	[0.047]	[0.035]	[0.023]
GSP per capita	−0.001**	−0.000	0.000	−0.000**
	[0.000]	[0.000]	[0.000]	[0.000]
PF size	−0.004	−0.023***	0.001	0.016***
	[0.005]	[0.007]	[0.005]	[0.004]
Year × Reporting month FE	Yes	Yes	Yes	Yes
Observations	890	786	531	463
Adjusted R-squared	0.233	0.420	0.196	0.541

This table presents results of regressions in which the dependent variable is the target allocation to risky assets of pension plans. Risky assets in column 1 include equity, real assets, private equity, hedge funds, and other risky assets. Columns 2–4 separately examine the target allocation to equity, real assets (RA), and private equity (PE). *Geometric* is an indicator variable for pension plans reporting geometric portfolio expected return (the omitted category is plans reporting arithmetic expected return). *Past return* and *Past standard deviation* measure the average arithmetic return and the standard deviation of annual returns in the previous 10-year period. *Past return equity* measures the average arithmetic return in public equity in the previous 10-year period. *Past return RA* and *Past return PE* capture the average arithmetic return in real assets and private equity. *PF size* is the natural logarithm of pension fund assets under management. *Unfunded liability / GSP* is the ratio of unfunded liabilities of pension funds relative to Gross State Product. *GSP per capita* is Gross State Product per capita in $ thousands. The specifications include year-reporting-month fixed effects. Standard errors clustered by pension plan are reported in brackets. *$p<.1$; **$p<.05$; ***$p<.01$.

portfolio return increases the allocation to equity by 4.9 percentage points, even keeping the past equity return held constant. Thus, high past returns on other asset classes seem to spill over into inducing higher allocations to equity. In column 4, we observe for private equity a smaller but still statistically significant effect of the past return on the target allocation. Each percentage point of higher private equity return is correlated with only a 0.3 percentage point higher allocation to private equity. The fact that overall allocations to private equity are much smaller than allocations to public equity can explain part of the difference in the economic effect between columns 2 and 4.

To explore further the deviation in asset allocation due to formulating expectations as a function of past performance, we split pension plans into quintiles based on their past performance in public equity and to look at differences in target allocation to equity. We focus on public equity because there is neither performance persistence nor smoothing of reported returns in this asset class. Internet Appendix Table A.18 shows a gradually increasing pattern in the allocation to equity as we move from quintiles with lower past

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Table 9
Target allocation to risky assets and risk premium

	Target allocation to risky assets					
	All (1)	Equity (2)	RA (3)	PE (4)	All (5)	All (6)
Geometric	0.008	0.027*	−0.019***	−0.016***	0.040***	0.024**
	[0.009]	[0.016]	[0.007]	[0.006]	[0.013]	[0.012]
Risk premium all risky assets	0.010***				0.041***	0.027***
	[0.003]				[0.009]	[0.007]
Risk premium equity		0.031***				
		[0.004]				
Risk premium RA			0.006*			
			[0.003]			
Risk premium PE				0.002		
				[0.002]		
Past standard deviation	0.022***	0.000	0.015***	0.012***		0.023***
	[0.005]	[0.005]	[0.003]	[0.003]		[0.005]
Unfunded liability / GSP	0.039	0.204***	−0.079**	−0.130***		0.044
	[0.040]	[0.061]	[0.034]	[0.025]		[0.037]
GSP per capita	−0.001**	−0.001**	0.000	−0.000***		−0.001**
	[0.000]	[0.000]	[0.000]	[0.000]		[0.000]
PF size	−0.001	−0.024***	0.001	0.019***	−0.007	−0.002
	[0.005]	[0.007]	[0.005]	[0.005]	[0.007]	[0.005]
Year × Reporting month FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	884	786	531	463	884	884
Adjusted *R*-squared	0.238	0.210	0.202	0.475		

This table presents results of regressions in which the dependent variable is the target allocation to risky assets of pension plans. Risky assets in columns 1, 5, and 6 include equity, real assets, private equity, hedge funds, and other risky assets. Columns 2–4 separately examine the target allocation to equity, real assets (RA), and private equity (PE). In columns 5 and 6, we estimate a 2SLS analysis using *Past return* as an instrument for the expected risk premium. In the first stage, we regress the risk premium for all risky assets on past return (and all the other controls). *Geometric* is an indicator variable for pension plans reporting geometric portfolio expected return (the omitted category is plans reporting arithmetic expected return). *Risk premium* variables measure the expected risk premium for all risky assets together as well as separately for equity, real assets, and private equity. *Past standard deviation* measures the standard deviation of the annual returns in the previous 10-year period. *PF size* is the natural logarithm of pension fund assets under management. *Unfunded liability / GSP* is the ratio of unfunded liabilities of pension funds relative to the Gross State Product. *GSP per capita* is Gross State Product per capita in $ thousands. The specifications include year-reporting-month fixed effects. Standard errors clustered by pension plan are reported in brackets. *$p<.1$; **$p<.05$; ***$p<.01$.

performance in equity to quintiles with higher past performance in equity. For example, pension plans in the top quintile have an 18 percentage points higher future target allocation to public equity than pension plans in the bottom quintile. This robustness test confirms that cross-sectional differences in past performance are correlated with substantial differences in the target asset allocation policy.

Table 9 shows estimates of coefficients in Equations (8) and (9), which examine the relation between the expected risk premium and target allocation to risky assets. In column 1, we estimate that the target allocation to risky assets is 1.0 percentage point higher for each percentage point higher overall risk premium that pension plans assume on their assets ($\lambda \approx 1$), which would represent very high risk aversion ($\gamma = 25$).[20] Column 2 shows that in public

[20] This is rather consistent with Mehra (2003), whose model requires $\gamma = 10$ to explain a risk premium of 1.4%. In our setting, the average expected risk premium on risky assets is 25.

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equity alone, this effect is stronger with an estimated λ coefficient of 3.1 percentage points, consistent with a constant relative risk aversion coefficient of 8.1. In real assets we find similar magnitudes to the baseline, while in private equity we do not find an effect specifically of the risk premium in private equity on the target allocation.

Columns 5 and 6 of Table 9 estimate Equation (10) by 2SLS. It shows the relation between the expected risk premium and target allocation to risky assets, considering only variation in the expected risk premium that is driven by the past return.[21] Here the λ coefficients are somewhat larger, ranging from 2.7 to 4.1 percentage points of reaction in target asset allocation to a 1 percentage point higher expected risk premium when that increase in the risk premium is driven by an increase in the past return. These coefficients therefore represent an estimate of how an additional percentage point of extrapolative expectations affects target asset allocations.

5.2 The costs of formulating expectations from past returns

Asset allocation decisions based on unjustified extrapolated return expectations would push pension funds away from their optimal asset allocation policy. In this section, we discuss several costly implications of deviations from optimal allocation driven by past-return-induced, cross-sectional variation in return expectations. The general challenge with analyzing welfare consequences is that for deviations from optimal allocation to have major utility consequences, there must be substantial frictions. That is, investing in any financial security at a fair price would be a zero NPV transaction, so for asset allocation to have important quantitative effects there must be reasons why different distributions of outcomes have different welfare consequences. Although it is beyond the scope of this paper to estimate all welfare consequences, we consider four reasons why deviations from optimal asset allocation may be costly.

First, using a myopic asset allocation model such as Campbell and Viceira (2002) as a baseline, it is well known that deviating from the optimal allocation has relatively small costs. For example, referring to Equations (2.24) and (2.25) of Campbell and Viceira (2002), under baseline parameters of risk aversion of 2.00, an expected return on the risky asset of 0.07, and a risk-free rate of 0.02, the optimal allocation to risky assets is 65%. Exceed that by 10 percentage points reduces the agent's certainty equivalent by 1.4 basis points; falling short of it by 10 percentage points reduces the agent's certainty equivalent by 2.7 basis points.

Such baseline costs appear modest. However, this type of model ignores the economic costs of mismatches between assets and liabilities faced by underfunded defined-benefit pension funds. For example, such costs may arise

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[21] The regression in column 5 of Table 4 constitutes the first stage for this 2SLS procedure.

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when regulations require pension systems to remediate unfunded liabilities at times when the cost of doing so is high. Van Binsbergen and Brandt (2016) incorporate pension liabilities in the asset allocation problem and analyze the differences between myopic asset allocation and dynamic allocation that takes these funding constraints into account. They find that after incorporating liabilities and funding constraints into the analysis, the certainty equivalent costs of deviating from the optimal allocation could be as large as 150 basis points for differences of 10 percentage points in the risky asset allocation.

We do not aim to identify the extent to which actual asset allocations of pension plans deviate from the optimal allocation. Prior research documents that U.S. pension plans do not sufficiently consider asset and liability management (ALM) and likely overinvest in risky assets relative to these models (Novy-Marx and Rauh 2011; Pennacchi and Rastad 2011; Andonov, Bauer, and Cremers 2017). Furthermore, other models find large utility losses when managers and other stakeholders have different levels of risk tolerance (Basak, Pavlova, and Shapiro 2007). If pension plans that are already above the optimal allocation to risky assets further increase risk taking due to excessive extrapolation from past returns, the certainty equivalent costs could be large.

Second, to the extent that pension funds implement active changes in allocation in response to past returns, they incur rebalancing costs in doing so. We show in Table 8 and Internet Appendix Table A.18 that pension plans adjust their future target asset allocation based on the past performance and in Table 6 that pension plans reach their target weights primarily through the residual changes in actual asset allocation. Thus, our results suggest that pension plans will achieve their new past-performance-driven target weights not merely through mechanical changes or managing inflows/outflows but in large part through active rebalancing and asset management transitions. According to data from MJ Hudson, a major transition manager, the average transition management fees are 33 basis points per mandate.[22] Public pension plans in the United States have around \$4.4 trillion assets under management. If we assume that they transition 1% of their assets between asset classes based on signals from past performance, this corresponds to transitioning \$44 billion in assets and paying around \$150 million in transition costs per year. These costs include fees paid to transition managers, direct costs (commission, taxes, etc.), indirect execution costs (market impact and bid-ask spread), and exposure to potential changes in asset value due to market movements (MJ Hudson 2021). They come on top of fees paid to legacy portfolio managers and new portfolio managers.

[22] MJ Hudson's transition manager survey data provides summary statistics on around 2,200 individual transition events, executed between 2014 and 2019, worth collectively nearly \$900 billion (MJ Hudson 2021). The sample includes transitions in equity and fixed income mandates and the majority of clients is pension funds.

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Third, our findings on the positive relation between future target asset allocation and past performance raises the possibility that pension fund rebalancing affects asset prices and returns. When the price elasticity of demand of the aggregate stock market is small, pension fund inflows and outflows in equities can have large price impacts (Koijen and Yogo 2019; Ben-David et al. 2021; Gabaix and Koijen 2021). For instance, Gabaix and Koijen (2021) estimate that the aggregate price multiplier is around 5, which implies that extrapolative pension fund return expectations can have potentially long-lasting impact on asset prices. Pension plans with lower past performance in equities may exit their underperforming mandates, which will have a negative impact on the prices of the underlying securities, while pension plans with higher past performance in equities may allocate more capital to their mandates, which will have a positive impact on the prices of the underlying securities.

Finally, in alternative asset classes, public pension funds increasing their allocation may take marginal investments in which they underperform the investors with superior access to the best funds (e.g., Lerner, Schoar, and Wongsunwai 2007; Sensoy, Wang, and Weisbach 2014). In our setting, we cannot directly examine whether the increases in alternative asset classes in response to performance are connected to worse outcomes in the asset class, as it takes a long time for the commitments made to private equity funds to materialize in cash flows and performance. The quality of the alternative investments made at the margin may be a concern.

Overall, we find substantial variation in target portfolio allocation among pension plans due to extrapolation of cross-sectional differences in past performance. These deviations in the asset allocation policy could be costly for pension plan participants and taxpayers (implicitly) supporting pension funds for multiple reasons, but we leave the quantitative measurement of the welfare implications of pension funds' deviating from their optimal allocation policy to future research.

6. Conclusion

Expected returns are one of the fundamental inputs to many canonical asset pricing models. Recent literature has established that past stock market performance affects the forward-looking expectations of individual investors (Vissing-Jorgensen 2003; Malmendier and Nagel 2011; Greenwood and Shleifer 2014). Although these papers and others clearly demonstrate a relation between beliefs and past experience for retail investors, our study is the first that we are aware of to make this determination for institutional investors.

The empirical specifications in this analysis identify this dependence in the cross-section: a pension fund that has achieved higher past returns than another in the same time period reports higher expected returns going forward, both at the pension fund level and at the level of individual asset classes. We find that the positive relation between past performance and future expectations operates

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through the expected real return or risk premium for investing in risky assets, rather than the inflation assumption or expected risk-free rate.

The extrapolation of past performance does not seem justified by differences in risk or the minimal amount of cross-sectional performance persistence that we find, which disappears completely and even reverses when looking at horizons beyond one year. Target asset allocation policy formed based on extrapolated expectations seems primarily to reflect the institution's true investment beliefs as actual asset allocation weights converge to target weights over time, and pension funds implement this convergence to target weights through active rebalancing. Our results provide evidence consistent with the excessive extrapolation hypothesis as the relation between return expectations and past performance is stronger in pension funds whose executives have personally experienced the pension fund's past returns.

Pension funds act on their past-return-induced expectations as funds with higher expected risk premiums have a higher target allocation to risky assets. Consistent with the literature on the equity premium puzzle, relatively high degrees of risk aversion would be required to explain the magnitude of the slope of this correlation. However, pension funds appear to respond more strongly in reallocating assets when the updating of their beliefs about the equity risk premium comes through their extrapolation of past returns—even though most of this updating is unlikely to be justifiable by the extent to which cross-sectional variation in past returns predict future returns.

These findings raise many questions for future research. One is the effect of marketwide fundamentals on pension fund expectations. In robustness analysis, we examine the relation between pension fund asset allocation and fundamental ratios such as the cyclically adjusted price-to-earnings (CAPE) ratio, but our time series is too short for a complete analysis. We leave the analysis of how pension plan expectations as a whole and in individual asset classes evolve over time for future research.

Furthermore, if public pension fund reallocations actually reduce their wealth on average in a predictable fashion, then such return-chasing flows would bear resemblance to what is found in the mutual fund literature (e.g., Frazzini and Lamont 2008). If return-chasing behavior aggregates to the level of the entire asset management industry and institutional investors are neglecting the equilibrium effects of such behavior, it could potentially lead to amplified swings in valuations at the asset class level (Shiller 2000; Barberis et al. 2018). However, if other institutional investors do not engage in excessive extrapolation and recognize the tendency of public pension funds to do so, they could potentially engage in stabilizing arbitrage and erase any amplification effect. Understanding the likely aggregate effects and welfare implications would require broader research on how institutional investors beyond public pension funds consider past returns in formulating the return expectations and adjusting their asset allocation.

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