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Pension plans: Risk and governance

Inaugural dissertation submitted by Oliver Dichter in fulfillment of the requirements for the degree of Doctor rerum oeconomicarum at the Faculty of Business, Economics and Social Sciences of the University of Bern.

Submitted by
Oliver Dichter
from Wangen bei Olten, Solothurn

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I. Pension Risk and Corporate Investment

Oliver Dichter*

October, 2015

This paper studies the relation of systematic pension risk (pension beta) and corporate investment in a large sample of U.S. firms. We present evidence of a negative impact of pension risk on investment, which is consistent with the view that firms forego valuable investment opportunities because they fail to notice that systematic pension risk causes an upward bias in the discount rates they use in capital budgeting decisions. The pension risk bias in investment is economically relevant and not limited to financially constrained firms. The study can be generalized to all firms that base their investment decisions on a firm-wide discount rate without noticing the different sources of systematic risk.

Keywords: Defined benefit pension plan; Corporate investment; Capital budgeting; Cost of capital

JEL codes: G23, G31

*Oliver Dichter (corresponding author, oliver.dichter@ifm.unibe.ch), the Institut für Finanzmanagement, University of Bern. This paper is part of my dissertation with Claudio Loderer, the Institut für Finanzmanagement, University of Bern and the Swiss Finance Institute. I wish to thank Claudio Loderer, Lukas Roth, Aleksandar Andonov, and Demian Berchtold for valuable comments. All errors are mine.

1. Introduction

The bad news is that standard cost of capital calculations used in corporate finance, which do not distinguish between the operating asset risk and pension plan risk, can greatly overestimate the discount rate for net present value analysis of operating projects. […] In such cases, capital projects with positive net present value could be rejected by management.

– Jin, Merton, and Bodie, Journal of Financial Economics, 2006 –

This paper presents empirical evidence for the distortion of corporate investment induced by inappropriate factoring in of the risk of defined benefit pension plans. According to the standard textbook formula, the value of an operating project depends on both its expected cash flows and its risk, which is reflected in the project's discount rate (Krüger, Landier, and Thesmar (2015)). Most firms do not estimate this rate for each project separately but use the weighted average cost of capital (WACC) as a single, firm-wide discount rate instead (Bierman (1993) and Graham and Harvey (2001)). According to Jin, Merton, and Bodie (2006), the WACC of firms that sponsor defined benefit (DB) plans is a biased estimate of the discount rate for net present value (NPV) analyses of operating projects. By discounting operating cash flows at the WACC, firms effectively assign their total risk to their business operations, although part of that risk actually comes from the pension assets and liabilities (Merton (2006)). Moreover, the standard calculation of the WACC understates the firm's leverage ratio because it does not take into account the pension liabilities, which are off-balance sheet items (Merton (2006)). The cumulative effect of these distortions is that the WACC generally overestimates the discount rate for operating projects, which could induce firms to forgo valuable investment opportunities (Jin, Merton, and Bodie (2006) and Merton (2006)).

DB pension plans are the largest off-balance sheet risk of corporate America (Shivdasani and Stefanescu (2010)). Despite the recent trend to freeze DB plans, the aggregate value of corporate (DB) pension liabilities reached an all-time high of USD 3.1 trillion in 2012 (Investment Company Institute (2014)). The distortion of corporate investment by pension risk could hence be substantial.

For the years 2003 to 2012 our analysis covers 1,562 U.S. COMPUSTAT firms that sponsor a DB pension plan. Our empirical strategy relies on a regression analysis with firm and year fixed effects. We measure corporate investment by capital expenditures (net of depreciation) and estimate the pension risk bias in the discount rate as suggested by Jin, Merton, and Bodie (2006). We predict that investment is decreasing in systematic pension risk because the discount rate (WACC) increases with pension beta, which reduces the number of positive NPV projects that the firm effectively invests in. Our empirical evidence supports this prediction. The distortion of corporate investment because of pension plan risk is economically large, robust to different empirical specifications, and emerges independently of plan freezes and financial distress of the plan sponsoring firm. On average, pension firms forgo valuable investment opportunities.

Our findings are consistent with Krüger, Landier, and Thesmar (2015), who show that firms underinvest in low risk divisions and overinvest in high risk divisions because they use a single discount rate that overestimates the systematic risk of low risk segments, while it underestimates the systematic risk of high risk segments.

Our results could reflect the financing constraints of the sponsoring firm. Rauh (2006) shows that investment of financially constrained firms declines with mandatory pension contributions, which are payments to the pension plan that cannot be altered or postponed. Campbell, Dhaliwal, and Schwartz (2012) explain Rauh’s finding with the effect of mandatory contributions on the WACC of financially constrained firms. Since our data do not allow us to control for mandatory pension contributions, the negative relation we observe between pension risk and investment activities could also be caused by the correlation of pension risk with these contributions. We address this concern by separately estimating the sensitivity of investment to pension risk for financially constrained as well as financially unconstrained firms. If our results were explained by unobserved mandatory contributions, pension risk should only affect the investment of financially constrained firms. Contrary to that, however, we find that pension risk affects corporate investment regardless of financing constraints. This indicates that the distortion of investment by pension risk is not driven by the presence of mandatory contributions.

A further endogeneity concern is the potential correlation between pension risk and unobserved investment opportunities. Firms with larger pension plans and higher systematic pension risk are typically older than firms with smaller pension obligations. Older firms might have fewer investment opportunities (Loderer, Stulz, and Waelchli (2015)). The correlation of an explanatory variable of investment with unobserved investment opportunities is well-known in the literature (e.g. Kaplan and Zingales (1997), Kaplan and Zingales (2000), Erickson and Whited (2000), and Rauh (2006)). To address this concern we match each firm with a DB plan in our sample to a firm without such a plan. The systematic pension risk is naturally zero for firms that do not sponsor DB pension plans (Jin, Merton, and Bodie (2006)) and therefore uncorrelated with investment opportunities. The results of this matched-sample analysis are consistent with our previous findings, which is inconsistent with the claim that the pension risk sensitivity of investment we observed is the consequence of a correlation of pension risk with unobserved investment opportunities.

The last section of the paper examines whether nonpension firms seize the investment opportunities that firms with DB pension plans forgo. Rauh (2006) finds that forgone investment by financially constrained firms is undertaken by firms that are not financially constrained. We test this prediction by regressing capital expenditures (net of depreciation) of nonpension firms on the aggregate pension risk of pension firms in the same Fama French 48 industry. We find that the investment of nonpension firms is indeed positively related to industry pension risk.

We contribute to the investment literature by showing that corporate investment is distorted by the risk of an important nonoperating activity of listed firms in the U.S. We show that this distortion is consistent with the effect of pension risk on the standard estimate of project discount rates (WACC), as described by Jin, Merton, and Bodie (2006). Moreover, we show that pension risk distorts corporate investment on top of the distortion from mandatory contributions, as identified by Rauh (2006). Krüger, Landier, and Thesmar (2015) find that using the WACC as a single discount rate distorts the within firm allocation of resources. We extend their work by showing that discounting with the WACC distorts the resource allocation between firms as well. Hence, we believe that the relevance of this paper goes beyond pension economics. Our findings apply to all firms that base their investment decisions on a single discount rate without taking into account the different sources of systematic risk.

The rest of the paper is organized as follows. Section 2 discusses the theoretical background of the paper. Section 3 describes the data and the sample selection. Section 4 presents our empirical method and the main variables. Section 5 shows the empirical results and their discussion. Last, section 6 concludes.

2. Theoretical background

There is a substantial strand of literature that studies whether the values of DB pension assets and liabilities are reflected in the market value of the sponsoring firms. Representative studies concerning the firm’s equity value include Oldfield (1977), Feldstein and Seligman (1981), Feldstein and Morck (1983), and Bulow, Morck, and Summers (1987). Carroll and Niehaus (1998) present similar evidence with respect to the debt market. Jin, Merton, and

Bodie (2006) extend this literature by showing that capital markets also account for the systematic risk of pension assets and liabilities. They show formally as well as empirically that a firm's systematic capital risk (β_{D+E}) is related to the firm's systematic pension risk (PR).[1]

$$\beta_{D+E} = \beta_{OA}\frac{OA}{D+E} + PR\,, \tag{1}$$

where β_{OA} is the systematic risk of operating assets (OA), E is the market value of equity, and D is the market value of debt. The systematic pension risk is the value weighted difference between the systematic risk of pension assets (β_{PA}) and the systematic risk of pension liabilities (β_{PL}).

$$PR = \beta_{PA}\frac{PA}{D+E} - \beta_{PL}\frac{PL}{D+E} \tag{2}$$

Fundamentally, the WACC is affected by pension risk because firms estimate their cost of capital based on past return data that reflect the systematic risk of their DB pension plans (Jin, Merton, and Bodie (2006)).

According to Bierman (1993), Graham and Harvey (2001), Brealey, Meyers, and Allen (2005), Ross, Westerfield, Jaffe, and Jordan (2010), and Krüger, Landier, and Thesmar (2015), standard capital budgeting techniques rely on the WACC as a single, firm-wide operating discount rate. Equation (1) states that the WACC is positively related to systematic pension risk. Since pension risk is unrelated to the risk of a firm's operating activity, the WACC is a biased estimate of the discount rate for the NPV estimation of operating projects

[1] Jin, Merton, and Bodie (2006) state the relation in risk terms and as a pre-tax cost of capital. They argue that this approach has the advantage that it removes the impact of financing which makes it more comparable across firms that potentially have different capital structures.

(Jin, Merton, and Bodie (2006)). Following this argument we predict that corporate investment is negatively related to pension risk as an increase in systematic pension risk (PR) increases the hurdle rate required for project acceptance, leading the firm to reject valuable low-risk projects. What follows documents the specification and the results of an empirical test of this prediction.

3. Data

Firms in the U.S. can choose between two types of retirement saving instruments – defined contribution (DC) and defined benefit (DB) plans. In a DB plan, the firm guarantees its employees specific benefits upon retirement. This commitment represents a debt-like liability of the firm (Jin, Merton, and Bodie (2006)). Since 1974, firms are obligated by the Employment Retirement Income Security Act (ERISA) to guarantee their pension liability with assets on a legally segregated account. The difference between pension assets and pension liabilities determines the funding status of a plan. A funding status of less than zero represents an underfunding. Whenever a plan is underfunded, the firm must cover for the deficit by deficit reduction contributions (Rauh (2006)). Until 2006, ERISA required that firms amortize the underfunding of a DB plan within thirty years (Rauh (2006)). The Pension Protection Act (PPA) of 2006 reduced this amortization period to seven years (Campbell, Dhaliwal, and Schwartz (2012)). In addition to the deficit reduction contributions to underfunded plans, firms are required to cover the plans' normal cost, which is the present value of pension benefits accrued during the year (Rauh (2006)). The sum of deficit reduction contributions and normal cost determines the firm's mandatory pension contributions.[2] When a firm fails to meet its mandatory contributions, the Pension Benefit Guarantee Corporation (PBGC) is entitled to recover the outstanding amount by filing a claim against the firm. In a

[2] Rauh (2006) and Campbell, Dhaliwal, and Schwartz (2012) provide a detailed discussion on mandatory pension contributions.

bankruptcy case, the PBGC claim has the most senior status (Shivdasani and Stefanescu (2010)).

The firm's responsibility in the case of DC plans is fundamentally different. The firm is simply committed to pay regular and fixed contributions to the employees' retirement accounts. Upon retirement, the employees receive whatever amount of money (contributions plus interest) has accumulated on their behalf. The uncertainty about the level of retirement benefits is borne entirely by the employees. Besides the regular contributions, the firm faces no further obligations (Shivdasani and Stefanescu (2010)). Consistent with previous research, including Rauh (2006), Jin, Merton, and Bodie (2006), and Campbell, Dhaliwal, and Schwartz (2012), we exclude DC plans from our analysis. Throughout this paper, we consequently use the terms *pension plan* and *defined benefit pension plan* interchangeably.

Our sample of DB pension sponsoring firms builds on data from the COMPUSTAT North America Pension database and the CRSP/COMPUSTAT Merged file. The COMPUSTAT North America Pension file contains firm level accounting data on DB pension plans. Corporate pension accounting in the U.S. is regulated by the Financial Accounting Standards Board (FASB). Pension assets and liabilities are recorded off-balance sheet in the footnotes of the 10-K annual statements (Shivdasani and Stefanescu (2010)). The FASB requires that pension assets be measured by their market value. Moreover, since the release of FAS 132(R) in 2003, firms are required to disclose pension assets along the categories *equities*, *bonds*, *real estate*, and *other investments*. Pension liabilities on the other hand have to be estimated as the actuarial present value of the promised benefits. However, the rate at which firms discount the pension benefits has to reflect current interest rate levels (Jin, Merton, and Bodie (2006) and Carmichael and Graham (2012)). This makes sure that the actuarial value of pension liabilities is close to their fair, economic value. There are two common measures of a firm's pension liability – the Projected Benefit Obligation (PBO) and the Accumulated Benefit Obligation (ABO). While the ABO only consists of the present value of the benefits that have already been earned by the employees, the PBO additionally includes the present value of the projected benefits that are attributable to future salary increases. The general obligation to disclose the ABO ended in 1998, when FAS 132 required disclosure only in case of severely underfunded plans. Even though Bodie (1990) argues that the ABO is the most accurate measure of the economic value of the pension liability, the lack of data after 1998 makes it an unfeasible measure during our observation period. We therefore quantify the pension liability by the PBO, which is in line with recent studies on corporate pension plans, including Franzoni and Marín (2006), Campbell, Dhaliwal, and Schwartz (2010), Campbell, Dhaliwal, and Schwartz (2012), and An, Huang, and Zhang (2013).

We limit our analysis to pension sponsoring firms and firm-years where complete accounting data are reported (book assets, market value of equity, book value of debt, capital expenditures, net income, depreciation, pension assets, pension asset classes, PBO, and pension contributions). Moreover, we exclude foreign firms with American Depository Receipts (ADRs). Our initial sample consists of 10,100 observations and covers the years from 2003 to 2012. The sample period starts in 2003 because the information on pension asset allocations is not available for previous years. This information is essential in our estimation of the systematic risk of pension assets. We do not exclude financial firms because Krüger, Landier, and Thesmar (2015) argue that these firms most likely base their investment decisions on discounted value evaluation techniques as well. Since we normalize our main variables by beginning-of-year assets, we require information about assets in at least two consecutive sample years, which reduces our sample size by 330 firm-years to 9,770 observations concerning 1,562 firms.

4. Empirical method and variable construction

In section 2, we argue that corporate investment is distorted by pension risk because of a pension risk bias in the operating discount rate. An empirical test of the relation between pension risk and investment hence requires that the investment measure reflects decisions that are presumably made on the basis of criteria such as NPV or internal rate of return (IRR). According to Bierman (1993), Graham and Harvey (2001), and Krüger, Landier, and Thesmar (2015), NPV and IRR are predominantly used in decisions about capital expenditures. Research and development expense (R&D), however, is often set as a fixed fraction of sales (Anthony and Govindarajan (2007)). Since NPV and IRR are hence less relevant in these investment decisions, R&D expense should be less sensitive to a bias in the operating discount rate. Therefore, we primarily measure investment by capital expenditures. In our main specification, we deduct depreciation expense to focus on decisions that do not reflect routine replacement activities for which NPV and IRR considerations might also be less relevant.

4.1. Measures of pension risk

Our empirical specification of the systematic pension risk follows Jin, Merton, and Bodie (2006) and is based on the functional relation reported in equation (2). First, we estimate the systematic pension asset risk (β_{PA}) of firm j in year t as the weighted average (CAPM) beta of the pension asset classes.

$$\beta_{PA_{j,t}} = \frac{\beta_{\text{Equities}} \times \text{Equities}_{j,t}}{PA_{j,t}} + \frac{\beta_{\text{Bonds}} \times \text{Bonds}_{j,t}}{PA_{j,t}} + \frac{\beta_{\text{Real Estate}} \times \text{Real estate}_{j,t}}{PA_{j,t}} + \frac{\beta_{\text{Alternatives}} \times \text{Alternatives}_{j,t}}{PA_{j,t}} \qquad (3)$$

The weight of each asset class is obtained from the COMPUSTAT North America Pension database. The betas we assume for equities, fixed income, and real estate are from Jin, Merton, and Bodie (2006).[3] The values are $\beta_{\text{Equities}} = 1$, $\beta_{\text{Bonds}} = 0.175$, and $\beta_{\text{Real estate}} = 0.15$. The asset class *alternatives* equals the COMPUSTAT category *other*, which comprises all assets that are not equity, bond, or real estate investments (COMPUSTAT (2004)). Since Jin, Merton, and Bodie (2006) do not consider alternative assets in their study, we rely on the beta of *alternatives* in Mohan and Zhang (2014). They argue that investments in alternative assets of DB pension plans predominantly consist of private equity, venture capital, and commodity investments, which have a beta of 1.2 on average.[4]

For systematic pension liability risk we again rely on Jin, Merton, and Bodie (2006). Based on the systematic risk of 30-year treasury bonds, they suggest two alternative point estimates; $\beta_{\text{PL1}} = 0.18$ and $\beta_{\text{PL2}} = 0.46$. In addition to these numbers, we consider a third estimate where we assume that the pension liability beta equals zero. Although this potentially underestimates the systematic risk of pension liabilities, it has the advantage that the systematic pension risk from equation (2) is reduced to the systematic risk of pension assets, which is independent from the pension liability definition (PBO vs. ABO) and the actuarial assumption on the discount rate of pension benefits. Our three estimates of systematic pension risk hence are

$$\text{PR1}_{\text{j,t}} = \frac{\beta_{\text{PA}_{\text{j,t}}} \times \text{PA}_{\text{j,t}} - 0.18 \times \text{PL}_{\text{j,t}}}{\text{A}_{\text{j,t}}}, \tag{4}$$

$$\text{PR2}_{\text{j,t}} = \frac{\beta_{\text{PA}_{\text{j,t}}} \times \text{PA}_{\text{j,t}} - 0.46 \times \text{PL}_{\text{j,t}}}{\text{A}_{\text{j,t}}}, \text{ and} \tag{5}$$

$$\text{PR3}_{\text{j,t}} = \frac{\beta_{\text{PA}_{\text{j,t}}} \times \text{PA}_{\text{j,t}}}{\text{A}_{\text{j,t}}}. \tag{6}$$

[3] Jin, Merton, and Bodie (2006), Table 5, p. 9.
[4] Mohan and Zhang (2014), Table 5, p. 407.

The variation in these estimates stems from changes in the asset allocation, changes in the pension funding status, and changes in the size of the pension plan relative to the size of the sponsoring firm. The assumptions about asset class and liability betas are constant over time and do not vary between firms. We normalize by book value of firm assets (A) because Rauh (2006) argues that all variables in the investment regression, which we discuss in the next section, should be scaled by the same quantity. Since book assets are also the denominator of Tobin's Q, Rauh (2006) suggest to scale both pension and firm variables by the book value of assets.

4.2. Regression model and discussion of the control variables

Following a large body of investment literature, including Fazzari, Hubbard, and Petersen (1988), Kaplan and Zingales (1997), Kaplan and Zingales (2000), Baker, Stein, and Wurgler (2003), and Rauh (2006), we examine the pension risk sensitivity of investment in a linear regression model. The investment of firm j in year t, scaled by beginning-of-year book value of firm assets, is hence given by

$$\mathrm{Inv}_{\mathrm{j,t}} = \alpha_{\mathrm{j}} + \alpha_{\mathrm{t}} + \beta \mathrm{PR}_{\mathrm{j,t-1}} + \boldsymbol{\Gamma}'\mathbf{X}_{\mathbf{j,t}} + \epsilon_{\mathrm{j,t}} , \quad (7)$$

where, α_{j} identifies firm fixed effects, α_{t} represents year fixed effects, $\mathbf{X}$ is a set of control variables, $\boldsymbol{\Gamma}'$ is a vector of coefficients, and ϵ is a stochastic error term. β identifies the pension risk sensitivity of investment. We expect that the value of β is negative. We consider the beginning-of-year systematic pension risk $\left(\mathrm{PR}_{\mathrm{j,t-1}}\right)$ because the end-of-year pension risk cannot yet be reflected in the stock returns that the firm uses to estimate the current year discount rate.

Including year fixed effects allows controlling for macroeconomic effects. Firm fixed effects control for firm specific differences in investment levels that remain constant over time. Our continuous control variables include the market-to-book ratio of the beginning-of-year asset values (average Tobin's Q), the current year nonpension cash flow (NPC), and the funding status of the firm's beginning-of-year pension liability. Q is a standard control for the firm's investment opportunities. NPC and funding status are controls for investment opportunities that are unobserved by Q (Rauh (2006)).

Consistent with Rauh (2006), we define nonpension cash flow and funding status by

$$\mathrm{NPC}_{j,t} = \frac{\text{Net income}_{j,t} + \text{Depreciation \& Amortization}_{j,t} + \text{Pension expense}_{j,t}}{A_{j,t-1}} \text{ and} \tag{8}$$

$$\text{Funding status}_{j,t-1} = \frac{PA_{j,t-1} - PBO_{j,t-1}}{A_{j,t-1}}, \tag{9}$$

where PBO is the Projected Benefit Obligation.

Furthermore, we control for financial leverage. Lang, Ofek, and Stulz (1996) find that investment is negatively related to financial leverage. Shivdasani and Stefanescu (2010) show that leverage is in turn negatively related to the ratio of pension liabilities to total assets, which is an important determinant of systematic pension risk. Failing to control for financial leverage could cause the coefficient of systematic pension risk to suffer from omitted variable bias.

We also consider the possibility that mandatory pension contributions could crowd out the investment of firms that face external financing constraints (Rauh (2006)). That could explain our findings. Yet we cannot control for these contributions because COMPUSTAT does only provide information on total pension contributions, which are the sum of mandatory and voluntary payments. We can, however, distinguish between predicted and unexpected *total* contributions. According to Rauh (2006), financially constrained firms could take measures to secure additional finance for predictable mandatory contributions. If so, the sensitivity of investment to mandatory contributions should primarily stem from unexpected mandatory contributions. In analogy to Rauh (2006), we therefore expect that total pension contributions affect the firms' willingness to invest mainly via their unexpected rather than via their expected component. We measure unexpected contributions by the difference between the pension contributions that the firm effectively pays during a given year and the amount of contributions it has planned to pay as of the beginning of the fiscal year. Consistent with the normalization of the other variables in our regression model, we scale unexpected pension contributions by the beginning-of-year book value of firm assets. We inquire into the potential distortion of the pension risk sensitivity of investment by unobserved mandatory contributions in the section (5.3) on endogeneity.

Finally, we control for the natural logarithm of the beginning-of-year book value of assets (Ln firm size) and the natural logarithm of firm age (Ln firm age) because firms with large pension plans, and hence large pension risk, are typically larger and older than firms with small pension plans (Rauh (2006)). Older firms might have fewer investment opportunities (Loderer, Stulz, and Waelchli (2015)). Detailed definitions of all the variables are displayed in Table 8 of the appendix.

5. Empirical results

5.1. Descriptive sample statistics

Table 1 shows summary statistics for our sample of pension sponsoring firms from 2003 to 2012. All continuous variables are winsorized at the 1st and the 99th percent level of their pooled distribution to eliminate outliers. On average, capital expenditures (Gross investment) correspond to 4.5 percent of firm assets and capital expenditures minus depreciation (Net investment) amount to 0.8 percent of assets. The positive mean of net investment indicates that the average firm in our sample is growing (Lang, Ofek, and Stulz (1996)). The mean aggregate ratio of pension liabilities to firm assets (Pension Liability) is 0.15 which, compared to an average financial leverage ratio (Leverage) of 0.25, illustrates that pension plans are an important corporate liability. The average funding status of -0.03 indicates that the average firm's pension liability is underfunded by 3 percent of the firm's assets, which corresponds to an average underfunding of pension liabilities by 22 percent. The mean and the median systematic pension risk are positive for all specifications, which supports the claim of Jin, Merton, and Bodie (2006) that the systematic pension risk causes the WACC of the average pension sponsoring firm to exceed the value that would be appropriate for the firm's operating business. The mean value of PR1 (0.060) amounts to 9 percent of the average asset beta of U.S. firms, which, according to Damodaran (2015), is 0.67. The 90th percentile of PR1 (0.158) corresponds to 24 percent of the average asset beta, which shows that for some firms, the distortion of the discount rate by pension risk could be substantial.

A large part of the overall variation in our main variables stems from within-firm variation over time. This is important because, in our main regression model (7), cross-sectional variation is eliminated by firm fixed effects (Baltagi (2013)). The within-firm standard deviation of PR1 (0.026) equals one third of the overall standard deviation of PR1 (0.077). In case of PR2 and PR3, the within-firm variation amounts to 58 and 29 percent of the overall variation, respectively. The within-firm standard deviation of our main investment variable (Net investment) corresponds to 61 percent of its overall standard deviation. For R&D, however, the within-firm variation only amounts to 25 percent of the total variation, which supports the view that R&D expense is stickier than capital expenditures, possibly because it is set as a fixed fraction of sales.

[insert Table 1 here]

Table 2 reports the pairwise Pearson correlation statistics between selected variables. The correlation between systematic pension risk and net investment is significantly negative for all pension risk measures, which represents univariate evidence of the distortion of corporate investment by DB pension plans. The correlation between net investment and the ratio of pension liabilities to firm assets is significantly negative as well. This is consistent with Jin, Merton, and Bodie (2006) and Merton (2006) who argue that pension firms overstate the discount rate (WACC) for operating projects because the standard calculation of WACC understates the leverage of these firms. The strong positive correlation between systematic pension risk and the ratio of pension liabilities to firm assets reflects that systematic pension risk is related to the size of the pension plan relative to the size of the firm. On average, firms with larger pension plans have higher pension risk.

R&D expense is negatively correlated with systematic pension risk well. Compared to net investment, the correlation is however weaker and only significant when we measure pension risk by PR2.

All our control variables are significantly correlated with net investment. With the sole exception of leverage, the control variables are also significantly correlated with systematic pension risk. This supports our approach to study the pension risk sensitivity of corporate investment in a multivariate regression analysis.

[insert Table 2 here]

5.2. The pension risk sensitivity of investment

Table 3 studies the relation between corporate investment and systematic pension risk. The statistical significance of the coefficients is determined based on a two-tailed test with standard errors clustered at the firm level. Column (1) displays the results of a regression of gross investment (capital expenditures) on systematic pension risk (PR1) and controls, including year fixed effects. The coefficient on pension risk takes a negative value of -0.052 and is statistically significant at the 1 percent level. This is consistent with the hypothesis that investment decisions are distorted because firms are discounting expected project cash flows at a rate that is sensitive to systematic pension risk. Moreover, this result is in line with Krüger, Landier, and Thesmar (2015), who find that segment investment of conglomerate firms is distorted because firms fail to adjust the discount rates for the difference in the systematic risk between their various business segments.

Column (2) shows the results of our regression model (7) that controls also for firm fixed effects. In this analysis, the coefficient on systematic pension risk is only significant at the 5 percent level and takes a value of -0.036, which is below the estimate from the preceding regression. In Columns (3) and (4), we estimate the pension risk sensitivity of net investment (capital expenditures minus depreciation). Without firm fixed effects, the pension risk coefficient takes a value of -0.059. Including firm fixed effects slightly changes the estimate to -0.052. In both regressions, the relation between net investment and pension risk is statistically significant at the 1 percent level. The comparably stronger pension risk sensitivity of net investment is in line with our expectation that pension risk primarily affects decisions on new investment because NPV and IRR evaluation techniques are less frequently used in pure replacement activities.

The coefficients of the control variables in the regressions that include firm fixed effects are in line with previous studies regardless of whether investment is measured gross or net of depreciation. In accordance with Rauh (2006), we find that investment increases with the pension funding status, the nonpension cash flow, and Tobin's Q, while it decreases with unexpected pension contributions.[5] Moreover, consistent with the findings in Lang, Ofek, and Stulz (1996), we find that investment is negatively related to leverage. With the exception of firm size and firm age, all coefficients are significantly different from zero. Firm size is significantly negatively related to gross investment but unrelated to net investment. The coefficient on firm age is insignificant and close to zero in magnitude.

Columns (5) and (6) test for the robustness of our results to the assumption concerning the systematic risk of pension liabilities. We find that net investment is negatively related to both alternative measures of systematic pension risk – PR2 and PR3. We obtain a similar result in not tabulated regressions of gross investment.

The distortion of corporate investment by systematic pension risk is also of economic significance. Based on the coefficients from Column (4), a one standard deviation increase of pension risk (0.08) decreases net investment by 0.11 standard deviations (1 SD = 0.036). This represents a decrease in the ratio of capital expenditures to assets by 19 percent. Given the total asset value of our sample firms of USD 2.3 trillion, this corresponds to an annual USD amount of 90 billion.

The remainder of Table 3 concerns the causality of the relation between pension risk and investment. A detailed discussion on endogeneity and further considerations on causality follow in the remainder of this paper. Columns (7) and (8) test for the pension risk sensitivity of R&D expense. Unlike capital expenditures, R&D is often set as a fixed fraction of sales (Anthony and Govindarajan (2007)). Therefore, it should be fairly insensitive to distortions in the discount rate. Indeed, regardless of whether the regression includes firm fixed effects,

[5] In not tabulated regressions, we replace unexpected contributions by the mandatory pension contribution estimate of Campbell, Dhaliwal and Schwartz (2012). This results in a positive contribution coefficient, which is inconsistent with Rauh (2006). We also receive positive coefficients for the Campbell et al. measure when we exactly replicate the regressions in Campbell, Dhaliwal and Schwartz (2012), p. 891, Table 2, Colum (3) and Column (4) for our sample period. We explain this result in the way that the measure of Campbell, Dhaliwal and Schwartz (2012) is based on the number of years a firm is allowed to take to eliminate an underfunding. In 2008, this period changed from 30 to 7 years. The Campbell et al. measure thus potentially suffers from a structural break within our sample period (2003 – 2012). Replacing unexpected pension contributions by a simple measure of overall pension contributions has virtually no impact on the coefficient on systematic pension risk.

R&D expense is not significantly related to systematic pension risk (PR1). In not tabulated regressions, we receive a similar result for PR2 and PR3.

According to Jin, Merton, and Bodie (2006), the WACC is affected by systematic pension risk because firms estimate their cost of capital based on past return data that reflect the risk of their DB pension plans. Since firms usually determine the WACC based on more than one year of past stock return information (Brotherson, Eades, Harris, and Higgins (2013)), investment decisions should also be sensitive to systematic pension risk measures lagged by two years. The evidence in Column (9) supports this prediction. We find that net investment is significantly negative related to both the one-year and the two-year lag of systematic pension risk. We obtain a similar result when we measure systematic pension risk by PR2 and PR3, respectively, or when we use gross investment as the dependent variable (not tabulated). The regression in Column (10) additionally includes the current year systematic pension risk. In section 4.2, we argue that investment should not be affected by the current year systematic pension risk because this information is not yet reflected in the stock returns that the firm uses to estimate the discount rate. Consistent with this reasoning the coefficient on end-of-year systematic pension risk is not statistically different from zero. In a not tabulated regression, we receive a similar result for the one year lead systematic pension risk.

[insert Table 3 here]

5.3. Endogeneity

This section elaborates further on the causality of the relation between pension risk and investment and discusses potential endogeneity concerns.

5.3.1. Pension freezes and financial distress

In recent years, many firms have frozen their DB pension plans and replaced new DB promises by contributions to DC plans, where the uncertainty about future retirement benefits lies entirely with the employees (Rauh, Stefanescu, and Zeldes (2013)). When a DB plan is frozen, future accruals are discontinued. The firm's existing (DB) pension obligations, however, remain (Shivdasani and Stefanescu (2010)). Since the WACC reflects the risk from existing pension assets and liabilities (Jin, Merton, and Bodie (2006)), the relation between pension risk and investment should be robust to pension freezes. We test for this robustness by separately estimating the pension risk sensitivity of investment of firms with frozen DB plans and firms with open DB plans. The first two columns of Table 4 present the results.

We identify firms with frozen plans by the reported rate of compensation increase, which is the firms' estimate of the increase in the employees' salaries that will affect future pension plan payments.[6] According to FAS 87, a firm is only required to disclose this item if it sponsors pay-related plans, which are plans where the benefits increase with the salary of the employees. Since wage related benefit adjustments are explicitly discontinued in (hard) frozen plans, we conclude that firms that do not report the rate of compensation increase have frozen their DB plans.[7] Based on this identification criterion, we find that the number of firms with frozen DB plans has steadily increased from 90 in 2003 to 320 in 2012. In an average year, 25 firms freeze their DB pension plans. However, there is also a small number of 7 firms per year that unfreeze their DB plans.

We find that the coefficient on PR1 is negative and statistically significant in both subsamples. In not tabulated regressions, we receive a similar result for gross investment and our alternative pension risk measures. This shows that the pension risk sensitivity of

[6] The reported rate of compensation increase is represented by the COMPUSTAT item *pprci*.

[7] In a soft freeze, it is only future accruals from additional years of service that are discontinued. The pension obligation is still adjusted for increases in the employees' salaries (Rauh, Stefanescu, and Zeldes (2013)).

investment is robust to plan freezes. Moreover, it implies that a further increase in the number of firms with frozen DB plans is unlikely to diminish the economic importance of our finding.

In the second part of Table 4, we test whether our results could be driven by financially distressed firms. Bodie, Light, Morck, and Taggart (1985), Bodie (1996), and Jin, Merton, and Bodie (2006) argue that financially distressed pension sponsors face a put option on their pension liability by the PBGC. In case the plan sponsor goes bankrupt, the PBGC takes over the pension liabilities. This pension put potentially incentivizes firms in financial distress to invest the pension assets in securities with higher systematic risk (Jin, Merton, and Bodie (2006)). According to Hoshi, Kashyap, and Scharfstein (1990), financially distressed firms have also a tendency to cut investment. The negative relation between pension risk and investment could hence be driven by financially distressed firms that simultaneously increase their pension risk and reduce their investment. To address this concern, we repeat our analysis for nondistressed firms only. Following Jin, Merton, and Bodie (2006), we measure financial distress based on (1) book to market ratio, (2) return on investment, and (3) financial leverage.[8] Based on each of these criteria, we construct a subsample of firms where investment decisions are unlikely affected by financial distress. In each year, we exclude the 90th percentile of firms that appeared the most severe financially distressed the year before. In Columns (3) to (5) of Table 4, we show the estimates for these subsamples of nondistressed firms. In all regressions, we estimate a negative and statistically significant coefficient on systematic pension risk that is close to -0.05 in magnitude, which is comparable to the pension risk sensitivity of net investment in in the entire sample. We receive a similar result for gross investment and PR2 and PR3 in not tabulated regressions. This indicates that our results are not driven by firms in financial distress.

[insert Table 4 here]

[8] The definitions are given in Table 8 of the appendix.

5.3.2. Unobserved mandatory contributions

Rauh (2006) finds that investment of financially constrained firms declines with mandatory pension contributions because firms that lack access to external financing resources face a trade-off between spending a dollar on investment and spending this dollar on mandatory pension contributions. Campbell, Dhaliwal, and Schwartz (2012) show that Rauh's result is explained by the positive impact of mandatory contributions on the WACC of financially constrained firms. So far, we control for this effect by including unexpected pension contributions as a control variable. This might however be insufficient to control for the correlation between mandatory contributions and pension risk. Our finding could hence still be a consequence of an omitted variable bias. If this was the case, the pension risk sensitivity of investment could actually reflect the crowding out of investment by mandatory contributions, as argued by Rauh (2006). It would however also imply that the pension risk sensitivity of investment is limited to financially constrained firms. Unconstrained firms can easily replace cash outflows by additional external financing. Their investment activity should thus be unaffected by mandatory contributions. On the other hand, if the sensitivity of investment to pension risk is caused by a discount rate bias in the capital budgeting process, it should affect pension sponsoring firms at large regardless of financing constraints. To distinguish between these two explanations, we follow Rauh (2006) and repeat our analysis for different subsamples where we sort the firms along different possible proxies for financing constraints. Table 5 presents the results. In each panel, the most severe financially constrained firms constitute the first subsample, partially constrained firms form the second subsample, and comparably unconstrained firms make up the third subsample. We display the results of regressions of net investment on PR1 and controls, including year and firm fixed effects. Our findings remain qualitatively unaffected when we consider regressions of gross investment or alternatively measure systematic pension risk by PR2 and PR3.

Overall, we find that pension risk is negatively related to investment independently of financing constraints, which is consistent with the existence of a discount rate bias. Unexpected pension contributions, however, primarily affect the investment of financially constrained firms, which is in line with the results of Rauh (2006) for mandatory pension contributions.

The first panel, where we sort firms along their median age, shows a significant negative impact of pension risk on investment for middle aged and old firms, which are comparably unconstrained. On the other hand, the coefficient on unexpected pension contributions is only statistically significant for young firms, where its magnitude (-0.182) is more twenty times its value in the sample of the oldest firms (-0.009).

Hadlock and Pierce (2010) argue that firm age, together with firm size, is a particularly good indicator of financing constraints. Hence, they combine the two measures in their size-age (SA) index, which we use to sort the firms in the second panel.[9] We find that investment decreases with systematic pension risk both for firms with high SA index values (higher financing constraints) and firms with low SA index values. The coefficient on unexpected pension contributions is however only statistically significant for firms with high SA index values. The investment of firms with low SA index values is essentially unaffected by unexpected pension contributions.

The third panel classifies firms according to their S&P credit rating into firms with no credit rating, firms with a credit rating that is below investment grade (BBB-), and investment grade firms. While we find a negative and statistically significant pension risk sensitivity of investment in all subsamples, the coefficient on unexpected pension contributions is only statistically significant for firms with no rating.

The fourth panel divides the sample along the median dividend-to-assets ratio. In analogy to Rauh (2006), the investment of firms with low dividend ratios displays the

[9] The detailed SA index calculation is provided in Table 8 of the appendix.

strongest sensitivity to unexpected pension contributions, while we observe no significant relation for high dividend paying firms. On the other hand, investment of high dividend as well as low dividend firms is significantly negatively affected by systematic pension risk.

Altogether, the results of the analyses in this section represent evidence against an explanation of the pension risk sensitivity of investment by unobserved mandatory contributions. On the contrary, they indicate that pension risk affects corporate investment on top of the distortion from mandatory contributions, as identified by Rauh (2006).

[insert Table 5 here]

5.3.3. Unobserved investment opportunities

A further endogeneity concern is that our results could be caused by the correlation of systematic pension risk with unobserved investment opportunities of the sponsoring firm. The correlation of an explanatory variable of investment with unobserved investment opportunities is a well-known issue in the investment literature. Historically, the problem primarily concerned the analysis of the cash flow sensitivity of investment (e.g. Kaplan and Zingales (1997), Kaplan and Zingales (2000), Erickson and Whited (2000), and Rauh (2006)). The same endogeneity concern applies to our study because firms with higher systematic pension risk are typically older than firms with small pension plans and these firms might have fewer investment opportunities (Loderer, Stulz, and Waelchli (2015)). To mitigate the concern that systematic pension risk is correlated with unobserved investment opportunities, we match each pension-sponsoring firm to a comparable nonpension firm, i.e., a firm that does not sponsor a DB pension plan. Our sample of nonpension firms includes all COMPUSTAT firms that are not listed on the COMPUSTAT North America Pension database. Systematic pension risk is zero for these firms (Jin, Merton, and Bodie (2006)) and therefore by definition uncorrelated with unobserved investment opportunities. Thus, a negative and significant coefficient on pension risk in this matched sample is unlikely the consequence of a correlation of pension risk with unobserved investment opportunities.

We perform a one-to-one matching of pension to nonpension firms based on propensity scores for each fiscal year, without replacement, using a caliper of 1 percent. Our matching variables are Tobin's Q, cash flow, book leverage, book value of assets, firm age, and Fama French 48 industry membership. We use these criteria because they represent important determinants of corporate investment and pension firms typically differ from nonpension firms along these dimensions. Firms that sponsor a DB plan tend to be older and larger than firms without DB plan (Rauh (2006)). Therefore, they likely have fewer growth opportunities (Loderer, Stulz, and Waelchli (2015)). Moreover, pension firms are more leveraged and have higher operating cash flows than nonpension firms (Shivdasani and Stefanescu (2010)). Finally, due to the historical evolution of pension plans and the emergence of DC plans, firms that sponsor a DB plan usually belong to more traditional industries. We calculate the propensity scores based on the method suggested by Abadie and Imbens (2006) and originally developed by Rosenbaum and Rubin (1983).[10] The procedure provides us with a sample of 9,548 observations (4,774 pension firm-years and 4,774 nonpension firm-years).

Table 6 reports the results of the analysis. In Column (1), we regress net investment on an identifier variable for pension firms (b DBP sponsor), the continuous nonpension variables from equation (7), and year fixed effects. The coefficient on b DBP sponsor takes a value of -0.007. This suggests that net investment (relative to assets) of pension firms is on average approximately 0.7 percentage points lower than net investment of nonpension firms. This corresponds to 14 percent of the average capital expenditures to assets ratio (0.049) of the firms in this matched sample.

[10] This is implemented in the STATA module *psmatch2* of Leuven and Sianesi (2014)

In Column (2), we extend the regression by the systematic pension risk (PR1). In this regression, the coefficient on b DBP sponsor decreases to -0.004. The coefficient on pension risk is negative and statistically significant. This indicates that a sizable part of the relatively lower investment of pension firms is explained by systematic pension risk, which is consistent with the interpretation that pension firms underinvest, on average. In Columns (3) to (5), we test our original two-way fixed effects specification from equation (7) and estimate the pension risk sensitivity based on our three pension risk measures (PR1, PR2, and PR3). Similar to our main analysis in section 5.2, including firm fixed effects allows controlling for omitted variables that remain constant over time. The pension risk sensitivity of investment is negative and statistically significant in all estimations. Column (6) shows that pension risk is also negatively related to gross investment.[11] The consistency of these results with our main findings in section 5.2 alleviates the concern that the pension risk sensitivity of investment is induced by a negative correlation between pension risk and unobserved investment opportunities.

[insert Table 6 here]

5.4. Reaction of nonpension firms

As pointed out in the introduction, we finally examine the reaction of firms that do not sponsor a DB plan (nonpension firms) to the distortion of investment at pension firms. Our explanation of the pension risk sensitivity of investment implies that pension firms forgo economically valuable investment opportunities. Rauh (2006) presents evidence that nonpension firms capture the forgone investment opportunities of financially constrained

[11] In not tabulated regressions, we receive a similar result for PR2 and PR3.

pension firms in their industry. With a similar argument, we therefore expect that nonpension firms seize the investment opportunities that pension firms forgo because of the pension risk bias in their discount rate. We test this prediction by inquiring into whether the investment activities of nonpension firms are positively related to the industry-wide pension risk.

For each Fama-French 48 industry (h), we first sum the not normalized pension risk of all pension firms in that industry. Since the magnitude of aggregate industry pension risk is important to nonpension firms only if it is large relative to the aggregate size of their own balance sheet assets, we then normalize this amount by the beginning-of-year aggregate book value of assets of all nonpension firms in that industry. Equation (10) shows the calculation.[12]

$$\text{Industry PR}_{h,t} = \frac{\sum_{j \in h, DB} \beta_{PA_{j,t}} \times PA_{j,t} - \beta_{PL} \times PL_{j,t}}{\sum_{i \in h, i \notin DB} A_{i,t-1}}, \tag{10}$$

where DB is the set of firms that sponsor a defined benefit pension plan, j is an identifier of pension firms, and i is an identifier of nonpension firms. Consistent with our previous analyses, we measure PA by the market value of pension assets, PL by the PBO, and A by the book value of firm assets. We determine the industry pension risk for an assumed pension liability beta (β_{PL}) of 0.18 (Industry PR1), 0.46 (Industry PR2), and 0 (Industry PR3), respectively. Similarly, we also estimate the industry unexpected pension contributions. Detailed definitions of all industry pension measures are provided in Table 8 of the appendix.

Table 7 reports the results of our regressions of investment of nonpension firms on industry pension risk, industry unexpected pension contributions, Tobin's Q, cash flow, leverage, and firm size. In analogy to Rauh (2006), we cluster the standard errors at the industry level. In the first three columns of Table (7), we examine the sensitivity of net investment to our three specifications of industry pension risk. We find a significant positive

[12] We calculate industry pension risk analogous to the calculation of industry mandatory contributions in Rauh (2006).

industry pension risk sensitivity of net investment of nonpension firms in all regressions. Column (4) shows that gross investment of nonpension firms is significantly positively related to industry pension risk as well.[13] Although it is not statistically significant, the coefficient on industry unexpected pension contributions is positive. This is consistent with Rauh (2006), who shows that nonpension firms capture investment that is crowded out by mandatory pension contributions. With the exception of the natural logarithm of age, which is significantly negatively related to investment of nonpension firms, the signs and the statistical significance of the coefficients on the remaining control variables are comparable to the estimates in or preceding analyses.

We interpret the results from Table (7) as supporting evidence for our explanation of the pension risk sensitivity of investment with a discount rate bias. It shows that the forgone valuable investment by firms that use distorted discount rates is undertaken by firms whose capital budgeting process is not analogously biased.

[insert Table 7 here]

[13] In not tabulated regressions, where we estimate the industry pension risk based on PR2 and PR3, respectively, we receive qualitatively similar results.

6. Conclusion

By analyzing a large sample of U.S. firms, we find that corporate investment is negatively affected by systematic pension risk and that pension firms underinvest, on average. The effect is economically large and not limited to firms that have not yet frozen their defined benefit pension plans. Our results are consistent with the interpretation that using the WACC as a firm-wide discount rate distorts capital budgeting decisions because the standard calculation of WACC fails to ignore the size and the systematic risk of pension assets and liabilities, which are both unrelated to a firm's operating business. Our findings cannot be explained by financing constraints of the sponsoring firms or the correlation of systematic pension risk with unobserved investment opportunities. The distortion of investment by pension risk occurs on top of the effect from mandatory contributions, as identified by Rauh (2006). The forgone investment by pension firms is partially seized by firms that do not suffer from a pension risk bias in their capital budgeting process.

We contribute to the investment literature by showing that corporate investment is distorted by an important nonoperating systematic risk in the computation of discount rates. Existing research finds that using a single discount rate in the NPV analysis of investment projects distorts the resource allocation within the firm. We find that it can distort the resource allocation between firms as well. Hence, we believe that the relevance of this paper goes beyond pension economics. Our findings apply to all firms that base their investment decisions on a single discount rate without noticing the different sources of systematic risk.

Appendix: Tables

Table 1: Descriptive sample statistics

This table shows the descriptive sample statistics of our main variables. Variable definitions are in Table 8. The data refer to 2003 to 2012. The sample is restricted to firms that sponsor a defined benefit pension plan. All variables are winsorized at the 1st and the 99th percentile of their pooled sample distribution.

	Mean	Median	Min	10th %ile	90th %ile	Max	Std. (overall)	Std. (within)	N
Gross investment	0.045	0.032	0.000	0.002	0.099	0.259	0.046	0.023	9,770
Net investment	0.008	-0.000	-0.053	-0.022	0.048	0.180	0.036	0.022	9,770
R&D	0.029	0.016	0.000	0.000	0.078	0.196	0.038	0.010	4,563
PR1	0.060	0.031	-0.003	0.002	0.158	0.395	0.077	0.026	9,770
PR2	0.017	0.004	-0.052	-0.008	0.059	0.202	0.038	0.022	9,770
PR3	0.087	0.048	0.001	0.005	0.222	0.542	0.107	0.031	9,770
Pension liability	0.154	0.093	0.002	0.010	0.374	0.967	0.180	0.046	9,770
Funding status	-0.031	-0.016	-0.226	-0.084	0.000	0.061	0.045	0.023	9,770
Unexpected contributions	0.001	0.000	-0.020	-0.002	0.007	0.031	0.006	0.006	9,719
NPC	0.090	0.085	-0.171	0.007	0.189	0.377	0.085	0.053	9,770
Q	1.499	1.261	0.730	0.949	2.343	5.004	0.723	0.328	9,770
Leverage	0.245	0.224	0.000	0.031	0.483	0.835	0.177	0.069	9,770
Firm assets ($m)	15,677	2,544	36	275	29,016	427,452	51,647	8,277	9,770
Firm age	34	32	5	9	62	86	22	2	9,770

Table 2: Pearson correlation matrix

This table shows the pairwise Pearson correlations between selected variables. Variable definitions are in Table 8. The data refer to 2003 to 2012. The sample is restricted to firms that sponsor a defined benefit pension plan. Subscripts indicate the number of lagged periods. The asterisk denotes statistical significance at the 5% level using a two-tailed test. All variables are winsorized at the 1st and the 99th percentile of their pooled sample distribution.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)
(1) Net investment	1												
(2) R&D	-0.1238*	1											
(3) PR1 $_{-1}$	-0.0849*	-0.0283	1										
(4) PR2$_{-1}$	-0.0325*	-0.0386*	0.8478*	1									
(5) PR3$_{-1}$	-0.0932*	-0.0244	0.9918*	0.7774*	1								
(6) Pension liability $_{-1}$	-0.1064*	-0.0091	0.9067*	0.5612*	0.9516*	1							
(7) Funding status$_{-1}$	0.1079*	-0.019	-0.3791*	0.0583*	-0.4668*	-0.6585*	1						
(8) Unexpected contributions	-0.0258*	-0.0143	0.1672*	0.0576*	0.1872*	0.2236*	-0.2456*	1					
(9) NPC	0.2077*	0.0559*	0.1603*	0.1141*	0.1648*	0.1647*	-0.1271*	0.1170*	1				
(10) Q$_{-1}$	0.1101*	0.3114*	0.1327*	0.1319*	0.1266*	0.1073*	-0.0297*	0.0648*	0.5977*	1			
(11) Leverage$_{-1}$	-0.0292*	-0.2144*	-0.0083	-0.0041	-0.0080	-0.0065	-0.0180	-0.0122	-0.0757*	-0.0959*	1		
(12) Ln firm size $_{-1}$	0.0569*	-0.0694*	-0.0773*	-0.0033	-0.0906*	-0.1183*	0.1631*	0.0234*	-0.1028*	-0.1061*	0.1272*	1	
(13) Ln firm age	0.0773*	-0.0620*	0.2598*	0.2075*	0.2605*	0.2450*	-0.1257*	0.0839*	0.0840*	0.0106	-0.0291*	0.1958*	1

Table 3: The pension risk sensitivity of investment

This table shows the results of our main regressions of corporate investment on the distortion in the WACC by systematic pension risk according to Jin, Merton, and Bodie (2006), and controls. The data refer to 2003 to 2012. Variable definitions are in Table 8. The sample is restricted to firms that sponsor a defined benefit pension plan. Subscripts indicate the number of lagged periods. All variables are winsorized at the 1st and the 99th percentile of their pooled distribution. Asterisks denote statistical significance at the 1% (***), 5% (**), and 10% (*) level using a two-tailed test. Standard errors (in parentheses) are clustered at the firm level.

Dependent variable	Gross investment		Net investment				R&D		Net investment	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
$PR1_{-1}$	-0.052***	-0.036**	-0.059***	-0.052***			-0.009	0.004	-0.044***	-0.031*
	(0.011)	(0.018)	(0.009)	(0.016)			(0.012)	(0.009)	(0.016)	(0.017)
$PR2_{-1}$					-0.062***					
					(0.019)					
$PR3_{-1}$						-0.043***				
						(0.013)				
$PR1_{-2}$									-0.022***	-0.022***
									(0.008)	(0.008)
PR1										-0.019
										(0.013)
Unexpected contributions	-0.237***	-0.056	-0.123*	-0.082*	-0.079*	-0.084*	-0.077	-0.005	-0.115*	-0.110*
	(0.079)	(0.051)	(0.066)	(0.045)	(0.045)	(0.044)	(0.081)	(0.024)	(0.059)	(0.059)
Funding $status_{-1}$	0.021	0.052**	0.059***	0.065***	0.081***	0.056***	-0.045*	0.001	0.061***	0.056**
	(0.017)	(0.022)	(0.014)	(0.020)	(0.022)	(0.019)	(0.024)	(0.012)	(0.023)	(0.024)
NPC	0.257***	0.060***	0.113***	0.038***	0.038***	0.038***	-0.090***	-0.010	0.038***	0.039***
	(0.019)	(0.009)	(0.014)	(0.009)	(0.009)	(0.009)	(0.021)	(0.008)	(0.010)	(0.010)
Q_{-1}	-0.007***	0.012***	-0.002	0.012***	0.012***	0.012***	0.020***	0.004***	0.011***	0.011***
	(0.002)	(0.002)	(0.001)	(0.002)	(0.002)	(0.002)	(0.003)	(0.001)	(0.002)	(0.002)
$Leverage_{-1}$	0.033***	-0.041***	-0.003	-0.040***	-0.040***	-0.041***	-0.045***	-0.006	-0.042***	-0.042***
	(0.005)	(0.007)	(0.004)	(0.006)	(0.006)	(0.006)	(0.008)	(0.004)	(0.008)	(0.008)
Ln firm $size_{-1}$	-0.002***	-0.007***	0.001**	0.001	0.002	0.000	-0.000	-0.012***	-0.000	-0.001
	(0.001)	(0.002)	(0.000)	(0.002)	(0.002)	(0.002)	(0.001)	(0.002)	(0.002)	(0.002)
Ln firm age	0.006***	0.003	0.005***	-0.001	-0.001	-0.001	-0.003*	0.000	0.005	0.006
	(0.001)	(0.006)	(0.001)	(0.005)	(0.005)	(0.005)	(0.002)	(0.004)	(0.008)	(0.008)
Firm fixed effects	No	Yes	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	8,076	8,076	8,076	8,076	8,076	8,076	3,798	3,798	6,585	6,585
R2 adjusted	0.197	0.765	0.098	0.626	0.626	0.626	0.163	0.938	0.640	0.640

Table 4: Plan freezes and financial distress

This table shows the results of our tests on whether the pension risk sensitivity of investment is affected by plan freezes and financial distress of the sponsoring firms. The data refer to 2003 to 2012. Variable definitions are in Table 8. Subscripts indicate the number of lagged periods. All variables are winsorized at the 1st and the 99th percentile of their pooled distribution. Asterisks denote statistical significance at the 1% (***), 5% (**), and 10% (*) level using a two-tailed test. Standard errors (in parentheses) are clustered at the firm level.

Dependent variable	Net investment				
	(1)	(2)	(3)	(4)	(5)
Sample restricted to	Firms that sponsor hard frozen plans	Firms that sponsor not hard frozen plans	Nondistressed firms by book-market ratio	Nondistressed firms by ROI	Nondistressed firms by financial leverage
$PR1_{-1}$	-0.069*	-0.046***	-0.056***	-0.056***	-0.045***
	(0.041)	(0.017)	(0.016)	(0.017)	(0.017)
Unexpected contributions	-0.126	-0.054	-0.077*	-0.076	-0.087*
	(0.177)	(0.050)	(0.045)	(0.047)	(0.048)
Funding $status_{-1}$	-0.010	0.072***	0.062***	0.076***	0.064***
	(0.038)	(0.022)	(0.021)	(0.021)	(0.021)
NPC	0.038**	0.044***	0.038***	0.043***	0.040***
	(0.016)	(0.011)	(0.009)	(0.010)	(0.010)
Q_{-1}	0.015***	0.011***	0.012***	0.013***	0.012***
	(0.004)	(0.002)	(0.002)	(0.002)	(0.002)
$Leverage_{-1}$	-0.043***	-0.037***	-0.042***	-0.036***	-0.039***
	(0.016)	(0.007)	(0.007)	(0.006)	(0.008)
Ln firm $size_{-1}$	0.004	0.001	0.000	0.000	0.001
	(0.005)	(0.002)	(0.002)	(0.002)	(0.002)
Ln firm age	0.020	0.001	-0.001	0.002	-0.001
	(0.018)	(0.006)	(0.006)	(0.005)	(0.006)
Firm fixed effects	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes
Observations	1,834	6,242	7,294	7,334	7,311
R2 adjusted	0.614	0.644	0.622	0.634	0.622

Table 5: The pension risk sensitivity of investment in subsamples defined by alternative indicators of financing constraints

This table shows the results of regressions of net investment on systematic pension risk and controls with the sample divided by hypothesized a priori indicators of financing constraints. Variable definitions are in Table 8. The data refer to 2003 to 2012. The sample is restricted to firms that sponsor a defined benefit pension plan. All variables are winsorized at the 1^{st} and the 99^{th} percentile of their pooled distribution. Subscripts indicate the number of lagged periods. All regressions control for year and firm fixed effects. Standard errors are clustered at the firm level.

				Explanatory variables															
Dependent variable				PR$_{-1}$		Unexpected contributions		Funding status$_{-1}$		NPC		Q$_{-1}$		Leverage$_{-1}$		Ln firm size$_{-1}$		Ln firm age	
Net investment	Count	Min	Max	Coeff	(t-Stat)	Coeff	(t-Stat)	Coeff	(t-Stat)	Coeff	(t-Stat)	Coeff	(t-Stat)	Coeff	(t-Stat)	Coeff	(t-Stat)	Coeff	(t-Stat)
Panel 1: Sorting by median firm age																			
Age (youngest)	2,721	5	19	-0.029	(-0.89)	-0.182	(-1.66)	0.009	(0.33)	0.034	(2.93)	0.015	(3.63)	-0.052	(-5.86)	0.006	(2.06)	-0.000	(-0.03)
Age (middle)	2,728	20	45	-0.095	(-2.62)	-0.120	(-1.20)	0.088	(1.86)	0.035	(2.50)	0.011	(3.17)	-0.025	(-2.18)	-0.004	(-1.04)	0.045	(1.05)
Age (oldest)	2,627	46	86	-0.040	(-1.97)	-0.009	(-0.19)	0.078	(3.21)	0.048	(2.12)	0.012	(3.97)	-0.052	(-4.96)	-0.001	(-0.18)	-0.146	(-2.29)
Panel 2: Sorting by median SA index																			
High SA index	2,695	-3.798	-2.907	-0.068	(-1.88)	-0.199	(-1.72)	0.046	(1.20)	0.037	(3.13)	0.011	(2.69)	-0.051	(-5.61)	0.003	(1.02)	-0.008	(-0.69)
Middle SA index	2,686	-4.808	-3.798	-0.052	(-1.47)	-0.098	(-1.13)	0.061	(1.20)	0.034	(2.15)	0.014	(4.66)	-0.022	(-1.84)	-0.001	(-0.22)	0.043	(1.79)
Low SA index	2,695	-6.496	-4.808	-0.044	(-2.20)	-0.005	(-0.10)	0.067	(2.83)	0.048	(2.18)	0.011	(3.79)	-0.051	(-5.09)	-0.000	(-0.14)	-0.092	(-1.52)
Panel 3: Sorting by median S&P credit rating																			
No S&P credit rating	3,119	-	-	-0.070	(-2.18)	-0.182	(-2.18)	0.046	(1.23)	0.027	(1.62)	0.012	(3.19)	-0.046	(-3.94)	-0.001	(-0.17)	-0.004	(-0.42)
S&P credit rating (low)	1,778	D	BB+	-0.046	(-1.63)	-0.054	(-0.55)	0.080	(2.01)	0.033	(2.67)	0.017	(4.85)	-0.040	(-4.49)	0.003	(1.19)	0.006	(0.58)
S&P credit rating (high)	3,179	BBB-	AAA	-0.033	(-2.13)	-0.005	(-0.09)	0.077	(3.28)	0.070	(4.89)	0.009	(4.22)	-0.030	(-3.18)	0.000	(0.04)	0.010	(1.55)
Panel 4: Sorting by median dividend-to-asset ratio																			
Low dividend	2,696	0.000	0.002	-0.090	(-3.27)	-0.211	(-2.07)	0.039	(1.15)	0.022	(1.46)	0.014	(2.98)	-0.050	(-4.87)	0.002	(0.57)	-0.002	(-0.15)
Middle dividend	2,684	0.002	0.014	-0.018	(-0.56)	-0.072	(-0.75)	0.058	(1.63)	0.065	(4.02)	0.016	(4.65)	-0.036	(-3.69)	0.002	(0.56)	0.007	(0.92)
High dividend	2,696	0.014	0.868	-0.041	(-1.88)	-0.037	(-0.66)	0.088	(2.70)	0.037	(2.58)	0.009	(3.34)	-0.030	(-3.16)	-0.001	(-0.19)	0.003	(0.27)

Table 6: Matched sample regressions

This table shows the results of regressions of corporate investment on the distortion in the WACC by systematic pension risk in a matched sample of firms that sponsor a defined benefit pension plan and firms that do not sponsor a defined benefit pension plan. Variable definitions are in Table 8. The data refer to 2003 to 2012. The sample is constructed by a one-to-one matching of pension to nonpension firms based on propensity scores for each fiscal year, without replacement, and a caliper of 1 percent. The matching variables are Tobin's Q, cash flow, financial leverage, book value of assets, firm age, and Fama French 48 industry membership. The calculation of the propensity scores is based on the methodology suggested by Abadie and Imbens (2006). All variables are winsorized at the 1st and the 99th percentile of their pooled distribution. Subscripts indicate the number of lagged periods. Asterisks denote statistical significance at the 1% (***), 5% (**), and 10% (*) level using a two-tailed test. Standard errors (in parentheses) are clustered at the firm level.

Dependent variable	Net investment					Gross Investment
	(1)	(2)	(3)	(4)	(5)	(6)
b DBP sponsor	-0.007***	-0.004**				
	(0.002)	(0.002)				
$PR1_{-1}$		-0.061***	-0.093**			-0.091**
		(0.013)	(0.041)			(0.044)
$PR2_{-1}$				-0.114**		
				(0.054)		
$PR3_{-1}$					-0.076**	
					(0.034)	
Unexpected contributions			-0.186*	-0.177	-0.189*	-0.183
			(0.109)	(0.109)	(0.108)	(0.123)
Funding status$_{-1}$			0.074	0.108*	0.056	0.052
			(0.051)	(0.057)	(0.049)	(0.052)
NPC	0.077***	0.078***	0.046***	0.046***	0.046***	0.061***
	(0.012)	(0.012)	(0.013)	(0.013)	(0.013)	(0.013)
Q_{-1}	0.003***	0.003***	0.011***	0.011***	0.011***	0.011***
	(0.001)	(0.001)	(0.002)	(0.002)	(0.002)	(0.003)
Leverage$_{-1}$	-0.002	-0.002	-0.051***	-0.051***	-0.051***	-0.047***
	(0.004)	(0.004)	(0.010)	(0.010)	(0.010)	(0.010)
Ln firm size$_{-1}$	0.002***	0.001***	-0.004	-0.004	-0.004	-0.012***
	(0.000)	(0.000)	(0.004)	(0.004)	(0.004)	(0.004)
Ln firm age	-0.002	-0.001	-0.005	-0.006	-0.005	0.000
	(0.001)	(0.001)	(0.010)	(0.010)	(0.010)	(0.010)
Firm fixed effects	No	No	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	8,617	8,578	8,578	8,578	8,578	8,578
R2 adjusted	0.059	0.062	0.566	0.566	0.566	0.717

Table 7: Investment response of nonpension firms

This table reports the results of regressions of corporate investment of nonpension firms on aggregate industry pension risk and controls. Variable definitions are in Table 8. The data refer to 2003 to 2012. Columns (1) to (3) are regressions on net investment; regression (4) is a regression on gross investment. All variables are winsorized at the 1st and the 99th percentile of their pooled distribution. Subscripts indicate the number of lagged periods. Asterisks denote statistical significance at the 1% (***), 5% (**), and 10% (*) level using a two-tailed test. Standard errors (in parentheses) are clustered at the industry level.

Dependent variable	Net investment			Gross investment
	(1)	(2)	(3)	(4)
Industry $PR1_{-1}$	0.004***			0.004**
	(0.001)			(0.001)
Industry $PR2_{-1}$		0.007*		
		(0.004)		
Industry $PR3_{-1}$			0.002***	
			(0.001)	
Industry unexpected contributions	0.007	0.019	0.004	0.006
	(0.055)	(0.049)	(0.057)	(0.057)
Cash flow	0.020**	0.020**	0.020**	0.018
	(0.010)	(0.010)	(0.010)	(0.012)
Q_{-1}	0.005***	0.005***	0.005***	0.006***
	(0.001)	(0.001)	(0.001)	(0.001)
$Leverage_{-1}$	-0.055***	-0.055***	-0.055***	-0.053***
	(0.016)	(0.016)	(0.016)	(0.018)
Ln firm $size_{-1}$	-0.008	-0.008	-0.008	-0.019***
	(0.006)	(0.006)	(0.006)	(0.005)
Ln firm age	-0.023**	-0.023**	-0.023**	-0.013
	(0.010)	(0.010)	(0.010)	(0.010)
Firm fixed effects	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
Observations	21,692	21,692	21,692	21,692
R2 adjusted	0.554	0.554	0.554	0.667

Table 8: Variable definitions

This table summarizes the variable definitions. The data are from the CRSP/COMPUSTAT Merged database and the COMPUSTAT North America Pension database. Subscripts indicate the number of lagged periods.

Variable	Definition (incl. COMPUSTAT mnemonics)
b DBP sponsor	Binary variable that takes a value of one when the firm is listed on the COMPUSTAT North America Pension database in the current year, and zero otherwise.
Net investment	Difference of capital expenditures (capx) and depreciation (dp) normalized by beginning-of-year assets (at_{-1}).
Gross investment	Capital expenditures (capx) normalized by beginning-of-year assets (at_{-1}).
R&D	Research and development expense (xrd) normalized by beginning-of-year assets (at_{-1}).
Pension asset beta	Weighted average (CAPM) beta of the firm's pension assets (pplao + pplau). Asset class weights in percent are equities (pnate), fixed income securities (pnatd), real estate (pnatr), and other assets (pnato). The category *other assets* contains all residual assets, which are mainly alternative assets. The assumed asset betas are as follows: Beta equities = 1 (Jin, Merton, and Bodie (2006)) Beta fixed income securities =0.175 (Jin, Merton, and Bodie (2006)) Beta real estate = 0.15 (Jin, Merton, and Bodie (2006)) Beta other = 1.2 (Mohan and Zhang (2014)) $\beta_{PA} = \frac{(1 \times \text{pnate} + 0.175 \times \text{pnatd} + 0.15 \times \text{pnatr} + 1.2 \times \text{pnato})/100}{\text{pplao} + \text{pplau}}$
Pension liability	Pension liabilities (pbpro + pbpru) to assets (at).
PR1	Systematic pension risk 1: Pension asset beta (β_{PA}) multiplied by the market value of pension assets (pplao + pplau) minus 0.18 times the PBO (pbpro + pbpru) normalized by assets (at).
PR2	Systematic pension risk 2: Pension asset beta (β_{PA}) multiplied by the market value of pension assets (pplao + pplau) minus 0.46 times the PBO (pbpro + pbpru) normalized by assets (at).
PR3	Systematic pension risk 3: Pension asset beta (β_{PA}) multiplied by the market value of pension assets (pplao + pplau) normalized by assets (at).
Unexpected contributions	Difference between effective employer contributions to defined benefit pension plans (pbec) and the beginning-of-year expectation of pension contributions ($pbece_{-1}$) normalized by beginning-of-year assets (at_{-1}). If the beginning-of-year expectation of pension contributions is missing, it is replaced by the previous year effective contribution ($pbec_{-1}$).
Funding status	Difference between the market value of pension assets (pplao + pplau) and the PBO (pbpro + pbpru) normalized by assets (at).
NPC	Nonpension cash flow according to Rauh (2006): Sum of net income (ni), depreciation and amortization (dp), and pension expense (xpr) normalized by beginning-of-year assets (at_{-1}).
Q	Average Tobin's Q: Market value of equity (csho × prcc_f) plus assets (at) minus the book value of common equity including deferred taxes (ceq + txdb) normalized by assets (at).
Leverage	Financial leverage: The ratio of book value of debt (dltt + dlc) to assets (at).
Ln firm size	Natural logarithm of assets in million USD (at).
Ln firm age	Natural logarithm of the difference between the current fiscal year and the year of birth of the firm. The year of birth is calculated as the first year the firm appears on the CRSP tapes or on the COMPUSTAT files or a link is indicated on the CRSP/COMPUSTAT Merged database.
Book market ratio	Book value of equity (at- dltt – dlc) divided by market value of equity (csho × prcc_f).
ROI	Return on investment: Net investment (ni) divided by book value of assets (at).
SA Index	-0.737 times the natural logarithm of assets (at) plus 0.043 times the squared natural logarithm of assets (at^2) minus 0.04 times the firm's age in years.
Dividend-to-asset ratio	Dividends paid (dvc) divided by book value of assets (at).
Cash-to-asset ratio	Cash and equivalents (che) divided by book value of assets (at).
Industry PR	The sum of the not normalized systematic pension risk of all pension firms in a Fama-French 48 industry divided by the beginning-of-year aggregated assets of nonpension firms in the same industry. $\text{Industry PR}_{h,t} = \frac{\sum_{j \in h,DB} \beta_{PA_{j,t}} \times (\text{pplao}_{j,t} + \text{pplau}_{j,t}) - \beta_{PL} \times (\text{pbpro}_{j,t} + \text{pbpru}_{j,t})}{\sum_{i \in h, i \notin DB} \text{at}_{i,t-1}}$ Industry PR1 assumes a pension liability beta (β_{PL}) of 0.18, Industry PR2 assumes a pension liability beta of 0.46, and Industry PR3 assumes a pension liability beta of 0.
Industry unexpected contributions	The sum of the unexpected pension contributions of all pension firms in a Fama-French 48 industry divided by the aggregated assets of nonpension firms in the same industry. $\text{Industry unexpected contributiions}_{h,t} = \frac{\sum_{j \in h,DB} (\text{pbec}_{j,t} - \text{pbece}_{j,t-1})}{\sum_{i \in h, i \notin DB} \text{at}_{i,t}}$

II. The Duration Gap Matters: How Pension Duration Affects Equity Returns

Oliver Dichter*

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This paper empirically studies whether equity returns of U.S. nonfinancial firms reflect the systematic interest rate risk of the sponsored defined benefit pension plans. It is not obvious that they should. Pension accounting rules are complex and pension assets and liabilities are held separately from the firm's operating assets. We find that the gap between the duration of pension assets and pension liabilities affects the interest rate exposure of the sponsoring firm without bias. This is consistent with the hypothesis of informationally efficient capital markets. Our results are robust to a wide range of assumptions regarding the duration of pension liabilities and pension asset classes and are not driven by firms with negligibly small pension plans or firms in financial distress. Besides, our results are neither caused by the recent financial crisis nor explained by the subsequent years of historically low interest rates.

Keywords: Defined benefit pension plan; Interest rates; Exposure; Market efficiency

JEL codes: G12, G14, G23, G32

*Oliver Dichter (corresponding author, oliver.dichter@ifm.unibe.ch) is from the Institut für Finanzmanagement, University of Bern. This paper is part of my dissertation with Claudio Loderer, the Institut für Finanzmanagement, University of Bern and the Swiss Finance Institute. I wish to thank Claudio Loderer for valuable comments. All errors are mine.

1. Introduction

For many [pension] schemes unhedged interest rate and inflation sensitivity are the biggest risks they're running within their investment portfolio.

– John Belgrove, senior partner of Aon Hewitt, Professional Pensions, 20 March 2014 –

The sharp fall in interest rates during the period 2008 to 2012 has resulted in a USD 460 billion increase in the liabilities of corporate defined benefit (DB) pension plans in the United States (Investment Company Institute (2014)). According to J.P. Morgan (2015), the interest rate risk of DB pension plans ranks among the top ten "striking facts" that firms should consider by setting their corporate finance strategies for 2015. This paper empirically examines whether capital markets are aware of the interest rate sensitivity of DB plans by studying whether the interest rate exposure of the sponsoring firm reflects the duration gap between the firm's pension assets and liabilities. It is not obvious that markets process pension information without bias. Pension assets and liabilities are reported off-balance sheet and held separately from the firm's operating assets, which could prevent that capital markets fully recognize the risk of DB pension plans.

The efficiency of capital markets with respect to pension plan risk is controversially discussed in the academic literature. While Jin, Merton, and Bodie (2006) find that capital markets recognize the systematic risk (beta) of pension plans without bias, Franzoni and Marín (2006) show that equity markets underestimate the financial risk from pension shortfalls.

Besides its relevance for the hypothesis of efficient capital markets, our research also contributes to a better understanding of the interest rate exposure of nonfinancial firms. With the notable exceptions of Sweeney and Warga (1986) and Bartram (2002), most of the existing research on interest rate exposures concerns financial firms (e.g., Flannery and James

(1984), Bae (1990), Madura and Zarruk (1995), and Memmel (2011)). We believe that the main reason for this limitation lies in the great difficulty to assess the interest rate sensitivity of the operating activity of nonfinancial firms. Changes in interest rates simultaneously affect operating cash flows, cost of capital, investment decisions, and the firm's competitive position relative to other firms (Bartram (2002)). The joint effect of these channels of influence on firm value predominantly depends on the specific business model characteristics of the firm and its competitors (Bartram (2002)). Since these factors are mostly unobservable, an empirical prediction of the interest rate exposure of nonfinancial firms is quite difficult. Our paper investigates the interest rate sensitivity of corporate DB pension plans. That sensitivity is one component of the firm's overall interest rate exposure that can be measured with publicly available information.

Based on the considerations about the relation between systematic firm risk and systematic pension risk in Jin, Merton, and Bodie (2006), we show formally that there should be a one-to-one relation between the systematic interest rate risk of the pension plan and the interest rate exposure of the sponsoring firm. We determine the systematic interest rate risk of pension plans by the difference between the duration of pension assets and the duration of pension liabilities. Our estimation approach is comparable to the calculation of systematic pension risk by Jin, Merton, and Bodie (2006). Our measure of the firm's interest rate exposure is the sensitivity of stock returns to shifts of the yield curve (i.e., the firm's equity duration).

We test our prediction of a one-to-one relation between the firm's interest rate exposure and the systematic interest rate risk of its pension plan with a panel of 224 nonfinancial U.S. COMPUSTAT firms that sponsor DB pension plans. We cover the years from 2003 to 2012 (1,195 firm-years). Our sample excludes firms whose pension plans might hedge their interest rate risk with derivatives. Unobserved hedging would substantially distort our estimate of the duration gap between pension assets and liabilities. Such distortions would make it difficult to tell whether deviations from the theoretical relation between the interest rate exposure of the firm and the systematic interest rate risk of the pension plan stem from market inefficiencies or simply from measurement error.

In agreement with the hypothesis that capital markets are informationally efficient, we present evidence that the firm’s interest rate exposure reflects the pension duration gap without bias. Moreover, we show that this is no different for firms with heavily underfunded pension plans. Hence, our results are consistent with efficient capital markets but in contradiction with Franzoni and Marín (2006), who argue that capital markets overvalue the equity of firms with deficits in their pension funding.

We test the robustness of our results in different ways. We cannot find that the relation between interest rate exposure and pension duration gap crucially depends on our assumptions about the duration of pension assets and liabilities. Similarly, we cannot find that our results are distorted by firms with negligibly small pension plans. Jin, Merton, and Bodie (2006) expect that pension risk has no first-order impact on equity returns of firms with small pension responsibilities. Furthermore, the relation between interest rate exposure and pension duration gap is not driven by firms in financial distress. Possibly, distressed firms simultaneously reduce their operating interest rate risk and the interest rate risk of their pension plans. Rauh (2009) shows that distressed firms reduce their pension risk to lower the expected cost of bankruptcy. Nance, Smith, and Smithson (1993) find that distressed firms increase risk hedging in general. At last, our results are robust to the financial crisis and the subsequent period of low interest rates. Bartram (2002) shows that both the direction and the magnitude of corporate interest rate exposures are sensitive to different periods in time.

Our paper adds to a long tradition of research about the impact of DB pension plans on the value of the sponsoring firm. Most notably, this includes Oldfield (1977), Feldstein and

Seligman (1981), Feldstein and Morck (1983), and Bulow, Morck, and Summers (1987). Furthermore, we contribute to the existing work of Jin, Merton, and Bodie (2006). They find that stock returns reflect the systematic risk of DB pension plans. We show that capital markets also recognize the systematic *interest rate* risk of DB plans, which represents a specific risk factor that has not yet been fully understood in the context of nonfinancial firms. Thus, we also contribute to the comparably scarce literature on interest rate exposures of nonfinancial firms (e.g., Sweeney and Warga (1986) and Bartram (2002)).

The rest of the paper is organized as follows. Section 2 describes the institutional background of U.S. corporate pension plans. Section 3 discusses the theoretical relation between the interest rate exposure of the firm and the systematic interest rate risk of the pension plan. Section 4 presents our empirical strategy. Section 5 describes the data. Section 6 presents the results and their discussion. Last, section 7 concludes.

2. Institutional background

Firms in the United States can choose between two types of retirement saving instruments – defined benefit (DB) plans and defined contribution (DC) plans. In a DB plan, the firm guarantees the employees specific and unconditional benefits upon retirement. This commitment represents a debt-like liability for the firm (Jin, Merton, and Bodie (2006)). Since 1974 firms are obligated by the Employment Retirement Income Security Act (ERISA) to guarantee their pension liabilities with assets in a segregated account. Whenever these assets are insufficient to cover the liabilities, the pension plan is underfunded, the firm must cover the deficit with deficit-reducing contributions (Rauh (2006)). Additionally, the firm must cover the discounted value of the pension benefits that have accrued during the current fiscal year (Rauh (2006)).[14] When a firm fails to meet its mandatory contributions, the Pension Benefit Guarantee Corporation (PBGC) is entitled to recover the outstanding amount by filing a claim against the firm. In a bankruptcy case, the PBGC claim has the most senior status (Shivdasani and Stefanescu (2010)).

The firm's responsibility in the case of DC plans is fundamentally different. When sponsoring a DC plan, the firm is simply committed to pay regular and fixed contributions to the employees' retirement accounts. Upon retirement, the employees receive whatever amount of money (contributions plus interest) has accumulated on their behalf. The uncertainty about the level of retirement benefits is borne entirely by the employees. The firm faces no further obligation besides that of the regular contributions (Shivdasani and Stefanescu (2010)). Consistent with previous research, including Rauh (2006), Jin, Merton, and Bodie (2006), Franzoni and Marín (2006), and Campbell, Dhaliwal, and Schwartz (2012), we exclude DC plans from our analysis. Throughout this paper we consequently use the terms *pension plan* and *defined benefit pension plan* interchangeably.

3. Theoretical considerations

Pension assets and liabilities are recorded off-balance sheet in the footnotes of 10-K annual statements (Shivdasani and Stefanescu (2010)). Nevertheless, firms are fully economically responsible for the risk of their pension plans. Firm and pension plan form a consolidated entity (Jin, Merton, and Bodie (2006)). Accordingly, financial analysts and rating agencies (e.g., Credit Suisse (2011) and Smyth (2013)) adjust their estimates of firm value by the values of pension assets and liabilities. There is also a large body of literature showing that the market value of pension sponsoring firms reflects the values of their pension assets and liabilities. Representative studies include Oldfield (1977), Feldstein and Seligman

[14] Rauh (2006) and Campbell, Dhaliwal, and Schwartz (2012) provide a detailed discussion on mandatory pension contributions.

(1981), Feldstein and Morck (1983), Bulow, Morck, and Summers (1987), and Carroll and Niehaus (1998).

Jin, Merton, and Bodie (2006) extend this literature by showing that stock returns not only reflect the values of pension assets and liabilities but also their systematic risk. They derive the following relation between the systematic risk of equity (β_E), the systematic risk of operating assets (β_{OA}), the systematic risk of debt (β_D), the systematic risk of pension assets (β_{PA}), and the systematic risk of pension liabilities (β_{PL}):

$$\beta_{E_j} = \frac{OA_j}{E_j}\beta_{OA_j} - \frac{D_j}{E_j}\beta_{D_j} + \left[\frac{PA_j}{E_j}\beta_{PA_j} - \frac{PL_j}{E_j}\beta_{PL_j}\right], \tag{1}$$

where OA_j is the value of operating assets, E_j is the value of equity, D_j is the value of debt, PA_j is the value of pension assets, and PL_j is the value of pension liabilities of firm j. The systematic risk of equity (β_{E_j}) is the equity beta from the Sharpe (1964) capital asset pricing model (CAPM), which implies that a firm's stock return $\left(R_{E_j}\right)$ in excess to the risk free rate (R_F) is given by

$$R_{E_j} - R_F = \alpha_j + \beta_{E_j}(R_{EM} - R_F) + \epsilon_j, \tag{2}$$

where R_{EM} is the return on an equity market index (a proxy for the return on the market portfolio), α_j is an intercept, and ϵ_j represents an error term. All returns are continuously compounded. Equation (1) shows that a firm's CAPM beta is positively related to the firm's systematic pension risk, which is represented by the expression in brackets.

A substantial part of the systematic risk of corporate DB pension plans consists of systematic *interest rate* risk that stems from the duration mismatch between pension assets and pension liabilities (Cornett and Saunders (2008) and Adams and Smith (2009)). To analyze the reflection of the pension duration gap in the stock returns of the sponsoring firm, we decompose the CAPM beta from equation (2) into a pure equity market risk component and a pure interest rate risk component. We do so following Stone (1974). In that APT model, a firm's stock return $\left(\mathrm{R}_{\mathrm{E_j}}\right)$ is modeled as

$$\mathrm{R}_{\mathrm{E_j}} = \alpha_\mathrm{j} + \beta'_{\mathrm{E_j}}\mathrm{R}_{\mathrm{EM}} + \gamma_{\mathrm{E_j}}\mathrm{R}_{\mathrm{BM}} + \epsilon_\mathrm{j} \, , \tag{3}$$

where R_{EM} is the continuously compounded return on an equity market index and R_{BM} is the continuously compounded return on a debt market index. $\beta'_{\mathrm{E_j}}$ and $\gamma_{\mathrm{E_j}}$ measure the responsiveness of stock returns to movements of the equity and debt market, respectively. Stone (1974) shows that the CAPM beta from equation (2) is a combination of the equity and interest rate risk component from equation (3), namely

$$\beta_{\mathrm{E_j}} = \beta'_{\mathrm{E_j}} + \gamma_{\mathrm{E_j}} \frac{\mathrm{Cov}(\mathrm{R}_{\mathrm{BM}}, \mathrm{R}_{\mathrm{EM}})}{\mathrm{Var}(\mathrm{R}_{\mathrm{EM}})} = \beta'_{\mathrm{E_j}} + \gamma_{\mathrm{E_j}}\beta_\mathrm{I} \, . \tag{4}$$

If we analogously decompose the betas in equation (1) and solve the expression for the systematic interest rate risk of equity $(\gamma_{\mathrm{E_j}})$, we obtain

$$\gamma_{\mathrm{E_j}} = \frac{\mathrm{OA_j}}{\mathrm{E_j}}\gamma_{\mathrm{OA_j}} - \frac{\mathrm{D_j}}{\mathrm{E_j}}\gamma_{\mathrm{D_j}} + \left[\frac{\mathrm{PA_j}}{\mathrm{E_j}}\gamma_{\mathrm{PA_j}} - \frac{\mathrm{PL_j}}{\mathrm{E_j}}\gamma_{\mathrm{PL_j}}\right] , \tag{5}$$

where γ_{OA} is the systematic interest rate risk of operating assets, γ_D is the systematic interest rate risk of the firm's debt, γ_{PA} is the systematic interest rate risk of pension assets, and γ_{PL} is the systematic interest rate risk of pension liabilities. We henceforth refer to γ_E as the interest rate exposure of the firm, which is consistent with the terminology in Bartram (2002). The derivation of equation (5) is shown in Appendix A.

Provided that capital markets process the relevant information on systematic interest rate risk of pension plans without bias, equation (5) implies a one-to-one relation between the firm's interest rate exposure (γ_E) and the systematic interest rate risk of its pension plan ($\gamma_{Pension}$), namely

$$\gamma_{E_j} = \gamma_{Pension_j} + \frac{OA_j}{E_j}\gamma_{OA_j} - \frac{D_j}{E_j}\gamma_{D_j}\ , \tag{6}$$

with

$$\gamma_{Pension_j} = \frac{PA_j}{E_j}\gamma_{PA_j} - \frac{PL_j}{E_j}\gamma_{PL_j}\ . \tag{7}$$

In the next section, we discuss our empirical strategy to test whether equation (6) holds in practice.

4. Regression specification and variable construction

4.1. Regression model

The linearity of equation (6) enables us to test the relation between the firm's interest rate exposure $\left(\gamma_{E_j}\right)$ and the systematic interest rate risk of its pension plan $\left(\gamma_{Pension_j}\right)$ with the linear regression

$$\gamma_{\mathrm{E}_{\mathrm{j,T}}} = a_{\mathrm{j}} + a_{\mathrm{T}} + b\gamma_{\mathrm{Pension}_{\mathrm{j,T}}} + \boldsymbol{\Gamma}'\mathbf{X}_{\mathbf{j,T}} + \epsilon_{\mathrm{j,T}} \,, \tag{8}$$

where b represents the sensitivity of the firm's interest rate exposure to the systematic interest rate risk of the pension plan, a_{j} and a_{T} identify firm and year fixed effects, respectively, $\boldsymbol{\Gamma}'\mathbf{X}_{\mathbf{j,T}}$ measures the firm's time variant nonpension interest rate risk, and $\epsilon_{\mathrm{j,T}}$ is a stochastic error term. If capital markets process the information about the interest rate risk of pension plans without bias, b has to equal one in magnitude.

4.2. Interest rate exposure of the firm

Consistent with previous research, including Flannery and James (1984), Sweeney and Warga (1986), Bae (1990), and Bartram (2002), we specify the firm's interest rate exposure as an economic exposure. However, while these studies estimate interest rate exposure as the sensitivity of stock returns to changes in a specific interest rate, we estimate it as the sensitivity of stock returns to a change in the entire yield curve. The reason is that pension liabilities and pension asset classes have different maturity structures and are thus unequally sensitive to changes in interest rates of different maturities. A measure that only reflects one specific interest rate would not capture the entire interest rate risk of pension plans.

We estimate the sensitivity of stock returns to changes in the 1-year, 5-year, 10-year, and 30-year default-free yield to maturity, while controlling for the return on a broad equity market index. In analogy to Jin, Merton, and Bodie (2006), we run the following regression for each firm j and year T based on weekly CRSP data (up to 52 observations):

$$\mathrm{R}_{\mathrm{E}_{\mathrm{j,t}}} = \alpha_{\mathrm{j}} + \beta'_{\mathrm{E}_{\mathrm{j}}}\mathrm{R}_{\mathrm{EM}_{\mathrm{t}}} + \gamma_{\mathrm{E1}_{\mathrm{j}}}\Delta\mathrm{I}_{1,\mathrm{t}} + \gamma_{\mathrm{E5}_{\mathrm{j}}}\Delta\mathrm{I}_{5,\mathrm{t}} + \gamma_{\mathrm{E10}_{\mathrm{j}}}\Delta\mathrm{I}_{10,\mathrm{t}} + \gamma_{\mathrm{E30}_{\mathrm{j}}}\Delta\mathrm{I}_{30,\mathrm{t}} + \epsilon_{\mathrm{j,t}} \,, \tag{9}$$

where $R_{E_{j,t}}$ is the continuously compounded total stock return of firm j in week t, α_j is an intercept, R_{EM_t} is the continuously compounded total return on the CRSP value-weighted U.S. stock market index in week t, and $\Delta I_{1,t}, \Delta I_{5,t}$, $\Delta I_{10,t}$, and $\Delta I_{30,t}$ measure the weekly change in the continuously compounded yield to maturity of the 1-, 5-, 10-, and 30-year CRSP fixed term index in week t, respectively. β'_{E_j} measures the responsiveness of stock returns to movements of the equity market. The coefficients γ_{E1_j}, γ_{E5_j}, γ_{E10_j}, and γ_{E30_j} denote the sensitivity of equity returns to a ceteris paribus change in the 1-, 5-, 10-, and 30-year yield to maturity, respectively. The sum of these partial interest rate sensitivities $\left(\hat{\gamma}_{E_j} = \gamma_{E1_j} + \gamma_{E5_j} + \gamma_{E10_j} + \gamma_{E30_j}\right)$ represents the interest rate exposure of the firm. Since this measure determines the sensitivity of stock returns to shifts in the yield curve, it meets the standard textbook definition of modified duration.

4.3. Systematic interest rate risk of the pension plan – the pension duration gap

Consistent with the estimation of the firm's interest rate exposure in the previous section, we empirically specify the systematic interest rate risk of pension assets and liabilities by their duration statistics. Our estimate of the systematic interest rate risk of pension plans ($\hat{\gamma}_{Pension}$) in equation (7) thus reflects the difference between the duration of pension assets and the duration of pension liabilities, weighted by the value of pension assets and liabilities, respectively, and normalized by firm equity. Henceforth, we refer to $\hat{\gamma}_{Pension}$ as the duration gap between pension assets and liabilities.

The empirical specification of the pension duration gap requires that we make assumptions about the duration of pension liabilities and the duration of pension asset classes. In accordance with Jin, Merton, and Bodie (2006), we assume that the pension liability duration is 13, which reflects the interest rate sensitivity of a 30-year government bond portfolio. This assumption is in line with Novy-Marx and Rauh (2011), who estimate the average pension liability duration at 13 as well, but base their estimation on detailed information on the actual maturity structure of the benefits of a sample of public pension plans.

Based on the pension asset classes in our data set, which we discuss in section 5, we estimate a firm's pension asset duration as the value-weighted average duration of the assets invested in bonds, equities, and real estate. According to Adams and Smith (2009), the duration of the bond portfolio of a representative pension plan is 6, while the duration of the equity portfolio is 0. The assumption about bond duration corresponds to the modified duration of a broad bond market index, e.g., the Barclays Global Aggregate Bond Index (Barclays (2014)). The equity duration assumption is consistent with our estimate of the average equity duration of COMPUSTAT firms in section 5.3. There is little agreement in the literature on the duration of real estate portfolios of pension plans. Estimates range from 0 to over a 100 (Hartzell, Shuhnan, Langetieg, and Leibowitz (1988), Chen and Chan (1989), and Chaney and Hoesli (2010)). In our main specification, we assume a real estate duration of 6, which equals our assumption about the duration of bond portfolios. In section 6.2 on the robustness of our results to the duration assumptions, we show that our results are however robust to real estate duration assumptions from 0 to 100.

Based on equation (7), aforementioned duration assumptions, and in analogy to the estimation of systematic pension risk by Jin, Merton, and Bodie (2006), we define the pension duration gap of firm j in year T by

$$\hat{\gamma}_{\mathrm{Pension}_{j,T}} = \mathrm{PA}_{j,T}\left(\frac{0 \times \mathrm{equities}_{j,T} - 6 \times \mathrm{bonds}_{j,T} - 6 \times \mathrm{real\ estate}_{j,T}}{\mathrm{E}_{j,T}}\right) - \mathrm{PL}_{j,T}\left(\frac{-13}{\mathrm{E}_{j,T}}\right), \qquad (10)$$

where $\text{equities}_{j,T}$, $\text{bonds}_{j,T}$, and real $\text{estate}_{j,T}$ represent the fraction of pension assets of firm j that are invested in these asset classes at year T. $E_{j,T}$ is the current year market value of the firm's equity, $PA_{j,T}$ is the market value of pension assets, and $PL_{j,T}$ is the actuarial value of pension liabilities. The duration statistics (-6 and -13) carry a negative sign because they represent negative interest rate sensitivities.

Following the terminology in Bartram, Brown, and Conrad (2011), our pension duration gap measure represents a gross (pre-hedging) exposure of the net pension plan value to changes in the interest rate. This measure differs across firms and over time because of differences in asset allocation, in pension funding status, and in size of the pension plan relative to the equity of the sponsoring firm. The assumptions we made about asset class and liability duration are constant over time and do not vary between firms.

4.4. Control variables

The control variables in the regression equation (8) should represent the term $\frac{OA}{E}\gamma_{OA} - \frac{D}{E}\gamma_D$ in equation (6), i.e., the fraction of the variation in corporate interest rate exposures that is caused by the interest rate sensitivity of operating assets and debt. Following the empirical specification of the test of the relation between systematic firm risk and systematic pension risk in Jin, Merton, and Bodie (2006), we use a series of proxies.

We proxy the term $-\frac{D}{E}\gamma_D$ by the ratio of short-term book value of debt to market value of equity and the ratio of long-term book value of debt to market value of equity. We expect that the firm's interest rate exposure increases with these ratios. This is because the interest rate sensitivity of debt (γ_D) is negative and multiplied by minus one in the expression $-\frac{D}{E}\gamma_D$. Moreover, since long-term debt has a higher duration than short-term debt, we expect that the coefficient on long-term debt to equity exceeds the coefficient on short-term debt to equity.

As for the interest rate risk of the operating business $\left(\frac{\mathrm{OA}}{\mathrm{E}}\gamma_{\mathrm{OA}}\right)$, Bartram (2002) argues that changes in interest rates simultaneously affect operating cash flows, cost of capital, investment decisions, and the competitive position of the firm relative to other firms. We proxy for these effects by controlling for the cash flow to assets ratio and the natural logarithm of the book value of assets ($\mathrm{Firm\ size}$). Our reasoning is that interest rate exposures are significantly related to cash flows (Bartram (2002)) and that firm size could serve as a very general proxy for a firm's business model and competitive position.

Finally, our regression model (8) includes firm fixed effects, year fixed effects, and the one-year lag of interest rate exposure. Firm fixed effects control for firm specific differences in interest rate exposures that remain constant over time, e.g., industry effects. Year fixed effects allow controlling for macroeconomic effects that affect all firms in a similar way. The one-year lag of interest rate exposure controls for firm specific trends and shifts in interest rate exposures. All variable definitions are in Table 9 of Appendix B.

5. Data

5.1. Data source

Our sample builds on data from the COMPUSTAT North American Pension database, the COMPUSTAT/CRSP Merged database, the CRSP daily fixed term index files, and the CRSP daily stock files. The COMPUSTAT database contains firm level reporting data based on 10-K annual statements.

The FASB requires that pension assets be measured by their market value, while pension liabilities have to be estimated as the actuarial present value of the promised benefits. The rate at which firms discount their pension liabilities has to reflect current interest rate levels (Jin, Merton, and Bodie (2006) and Carmichael and Graham (2012)). There are two common definitions of a firm's pension liability – the Projected Benefit Obligation (PBO) and the Accumulated Benefit Obligation (ABO). While the ABO is defined as the present value of the benefits on the assumption that the pension plan is to be terminated immediately, the PBO additionally reflects the estimated remaining service life of employees, their projected salary increases, and their mortality rates. Since the issuance of FAS 87 in 1985, pension assets and pension liabilities are disclosed in the footnotes of annual financial statements. The general obligation to disclose the ABO ended in 1998. Even though the ABO is the most accurate measure of the economic value of pension liabilities (Bodie (1990)), it is potentially affected by a selection bias. We therefore quantify pension liabilities by the PBO measure in most of our analysis. This approach is in line with recent studies on corporate pension plans, including Franzoni and Marín (2006), Campbell, Dhaliwal, and Schwartz (2010), Campbell, Dhaliwal, and Schwartz (2012), and An, Huang, and Zhang (2013). Nevertheless, we show that our results remain virtually unchanged if we measure the pension liabilities by the ABO measure instead.

5.2. Sample selection

We limit our sample to nonfinancial firms, remove observations from foreign firms with American Depository Receipts (ADRs), and exclude firm-years with incomplete information on the relevant firm and pension accounting data (firm assets, firm debt, pension assets, pension asset allocation and PBO). We also exclude observations where the market value of equity is missing or the firm's stock has not been traded in more than 43 weeks of the year. The estimated betas of firms that are infrequently traded are not meaningful. Furthermore, we exclude observations where the previous year interest rate exposure is missing.

The sample period starts in 2003 because the information on pension asset allocations is not available for previous years. FAS 132 (R) requires that firms disclose pension assets along the categories *equities*, *bonds*, *real estate*, and *other*. The residual category *other* includes all assets that are not equity, bond, or real estate investments (COMPUSTAT (2004)). Consequently, this includes derivative positions held for interest rate hedging purposes. We exclude firms that report a partial allocation of their pension assets to *other* assets. Unobserved derivative hedging would make it impossible to determine whether deviations from the theoretical relation between the pension duration gap and the firm's interest rate exposure stem from market inefficiencies or simply from an inability to measure the duration gap correctly. This leaves us with a sample of 224 firms and 1,195 firm-years.

5.3. Descriptive sample statistics

The first three panels of Table 1 show the summary statistics for our sample of pension sponsoring firms from 2003 to 2012. All variables are winsorized at the 1^{st} and the 99^{th} percent level of their pooled distribution to eliminate outliers. Panel A displays our main variables. Panel B shows additional pension plan characteristics. Panel C shows the characteristics of the firms in our sample. In Panel D, we additionally display these firm characteristics for a broad sample of pension and nonpension COMPUSTAT firms. This sample consists of both financial and nonfinancial firms that satisfy the nonpension selection criteria from the previous section.

The median interest rate exposure in our sample is 0.97, which suggests that the equity value of the average sample firm increases by 1 percent in reaction to an upward shift of the yield curve by one percentage point. Stated differently, the median equity duration of our sample firms is approximately -1. The median interest rate exposure of our broad sample of COMPUSTAT firms (Panel D) is 0.22, which is close to our assumption that the average duration of a broadly diversified equity portfolio is 0. In section 6.2, we show that our results are virtually unaffected if we assume that the duration of equity portfolios is -0.22.

Despite the fact that interest rate exposures almost neutralize on average, they differ substantially between firms. In our pension firm sample, they range from -58.6 to 63.2. The average pension duration gap (normalized by firm equity) amounts to 3.19 if we measure pension liabilities by the PBO and 3.22 if we measure pension liabilities by the ABO instead. The 90th percentile is 7.3 (PBO) and 7.5 (ABO), respectively. The smaller number of observations in the case of the ABO duration gap (954 vs. 1,195 in the case of the PBO measure) reflects the fact that, as mentioned above, firms are not generally required to report the ABO. The positive minimum values of 0.025 (PBO) and 0.018 (ABO) illustrate that the duration gap is strictly positive, which implies that the duration of pension liabilities is always higher than the duration of pension assets.

The average ratio of pension liabilities (PBO) to firm assets is 0.16. This compares to an average financial leverage of 0.27, which documents the relative importance of pension plans as a corporate liability. On average, only 76 percent of the PBO is backed by pension assets, which indicates that the average pension plan is substantially underfunded. The average pension asset allocation is dominated by equity investments (60 percent). Fixed income securities only account for 39 percent and real estate investments for 1 percent of the average pension asset allocation. The predominant allocation of pension assets to equities represents the main reason for the distinct positive duration gap of corporate DB pension plans in the U.S. While the duration of pension liabilities is 13 (Jin, Merton, and Bodie (2006)), the duration of diversified equity portfolios is 0 (Adams and Smith (2009)).

The descriptive statistics of the firm characteristics show that our sample consists of an average set of COMPUSTAT firms. Based on the comparison of the median values, we find that the average firm in our pension sample has a higher cash flow to asset ratio and is slightly larger and more leveraged than the average COMPUSTAT firm. However, such a bias is common in studies on corporate DB pension plans and in line with previous studies (e.g., Rauh (2006) and Shivdasani and Stefanescu (2010)).

[insert Table 1 here]

Table 2 reports the pairwise Pearson correlations between selected variables. The correlation between the pension duration gap and the interest rate exposure is positive, statistically significant and almost identical for both pension duration gap measures. This represents first, univariate evidence of the recognition of the pension duration gap by capital markets. Similarly, the ratios of short-term and long-term debt to equity are significantly positively correlated with the firm's interest rate exposure. The control variables for the firm's operating interest rate risk are not significantly related to interest rate exposure. They are, however, significantly correlated with both the PBO and the ABO based duration gap measure. The high correlation (0.997) between these two measures of the interest rate sensitivity of pension plans indicates that the pension liability definition might be of little consequence to our analysis.

[insert Table 2 here]

6. Empirical analysis

In this section, we empirically test whether the relation between interest rate exposure and pension duration gap, which we have identified under the assumption of efficient capital markets, also holds in practice.

6.1. Main results

In this section, we empirically study the relation between the firm's interest rate exposure and the pension duration gap. We run regressions based on different specifications of equation (8). All regressions are controlled for firm and year fixed effects. The statistical significance of the coefficients is determined based on a two-tailed test with standard errors clustered at the firm level. Table 3 displays the results. The p-value at the bottom of the table refers to a two-tailed Wald test of whether the coefficient on the pension duration gap equals one.

In the first two columns, we run regression of interest rate exposure on pension duration gap and our proxies for the interest rate sensitivity of firm debt (short-term and long-term debt to equity). In these analyses, the variation in a firm's operating interest rate risk is reflected in the error term. In the second two columns, we display the results of regressions that include our entire set of control variables from section 4.4. We measure pension liabilities by the PBO in Columns (1) and (3) and by the ABO in Columns (2) and (4). For both duration gap measures, and regardless of whether we include our proxies for the firm's operating interest rate risk, the coefficient on pension duration gap is significantly larger than zero and not statistically different from one. These results support our prediction of a one-to-one relation between interest rate exposure and pension duration gap and are therefore consistent with the hypothesis that capital markets are informationally efficient. Considering our control variables, we find that interest rate exposure is significantly related to cash flows, firm size, and the one-year lag interest rate exposure. In line with Bartram (2002), who shows that the statistical relation between interest rate exposure and financial leverage is weak, we can not find that the firm's interest rate exposure is significantly affected by the short-term and long-term debt to equity ratio, respectively.[15]

In Columns (5) and (6), we test for the robustness of our results to the proxies that are used by Jin, Merton, and Bodie (2006) to control for systematic risk (CAPM beta) of operating assets. Possibly, these factors might explain the systematic *interest rate* risk of operating assets as well. These additional controls include the percentage of industry total sales that is earned by the firm (Market share), the capital intensiveness of the firm's operating business, cash holdings divided by total assets (Cash position), the growth rate of asset (Growth rate), the ratio of current assets to current liabilities (Liquidity), the ratio of research and development expense to assets, and the ratio of advertising expense to assets. We provide detailed definitions of these variables in Table 9 of Appendix B. For both the PBO and the ABO based duration gap measure, the coefficient is positive and statistically not different from one. None of the coefficients on the proxies for systematic risk is significantly different from zero. In our further analyses, we thus rely on our original regression specification in section 4.

The impact of the pension duration gap on the interest rate exposure of the sponsoring firm is also of economic significance. Based on the coefficients estimated in Column (3), a one standard deviation increase in the duration gap (6.0) leads to a rise in the interest rate exposure (equity duration) of the firm by 7.6 (0.2 standard deviations).

[insert Table 3 here]

In the next step of our analysis, we address the concern of Franzoni and Marín (2006) that capital markets overvalue the equity of firms with large deficits in their pension funding.

[15] If we control for financial leverage instead of short-term and long-term debt to equity, we receive a similar result (not tabulated).

We do so by sorting our sample along the pension funding ratio and repeating our analysis for different groups of firms whose pension liabilities are the most severely underfunded. The funding ratio is the ratio of pension assets to PBO. Table 4 reports the results for the PBO based duration gap measure. Our findings do not change if we alternatively measure pension liabilities by the ABO (not tabulated). Columns (1) to (4) show the estimates for four alternative subsamples where we keep 80, 60, 40, and 20 percent of the firms with the most underfunded pension plans, respectively. In all regressions, the duration gap coefficient is significantly positive but statistically not different from one. These results indicate that capital markets are informationally efficient with respect to the systematic interest rate risk of pension plans regardless of their funding status.

[insert Table 4 here]

6.2. Robustness to the duration assumptions

Figure 1 considers the robustness of our findings to our assumptions about the duration of pension assets and liabilities. Graph A shows the duration gap coefficient as a function of the pension liability duration assumption. Graphs B, C, and D display the estimate as a function of the assumption about the duration of pension assets invested in bonds, equities, and real estate, respectively. We measure pension liabilities by the PBO. However, the shapes of the graphs are unaffected if we alternatively measure pension liabilities by the ABO.

For pension liabilities, the literature mentions both longer and shorter duration assumptions than the 13 we use. Ryan and Fabozzi (2002) and Pennacchi and Rastad (2011) assume that the duration of state and local pension liabilities is 15. Mercer (2014) estimates the average duration of corporate pension liability at 17, Jared Gross, a former chief financial economist at the Pension Benefit Guarantee Corporation (PBGC), arrives at an estimate of 12 (Jin, Merton, and Bodie (2006)). Graph A plots the coefficient on the pension duration gap for an assumed pension liability duration from 10 to 20. A liability duration of 12 is associated with a duration gap coefficient of 1.41; a duration of 15 yields an estimate of 1.09; and a liability duration of 17 results in a coefficient of 0.95. None of these values is significantly different from one.

The impacts of bond, equity, and real estate duration assumptions are comparably small as well. If we set the bond duration at 18, which implies that pension plans uniquely invest in bonds with maturities over 20 years (Barclays (2014)), the duration gap coefficient takes a value of 1.71. If we assume that the entire fixed income portfolio is invested in cash equivalents (duration of 0), we obtain a point estimate of 1.12. Both estimates are statistically not different from one.

For equity securities, we test for both positive and negative duration statistics. Leibowitz (1986) argues that equity durations could also be negative. Graph C plots the pension duration gap coefficient as a function of equity durations between -20 and 6. We set the upper level of equity duration at 6 because we believe it is highly unlikely that equity portfolios have a higher duration than broadly diversified bond portfolios. The lower level (-18) reflects a reduction of our main equity portfolio duration assumption (0) by one standard deviation of the equity duration of COMPUSTAT firms in Table 1. The resulting point estimates range from 0.77 (equity duration of -18) to 1.58 (equity duration of 6). They are never statistically different from 1. If we assume that the equity duration equals the median equity duration of our sample of COMPUSTAT firms in Table 1 (-0.22), the coefficient on the pension duration gap takes a value of 1.27, which is virtually identical to the estimate of 1.28 in Column (1) of Table 3.

In section 4.3, we show that the literature mentions duration assumptions for real estate portfolios between 0 and 100. Though this reflects a considerable uncertainty about the duration of real estate investments, it does not compromise our analysis as the duration gap coefficient is virtually insensitive to changes in the assumed real estate duration. The duration gap coefficient is 1.28 if we assume that real estate has a duration of 0. In comparison, it takes a value of 1.27 if we set real estate duration to 100. The reason why the duration gap coefficient is almost insensitive to the assumption about the duration of real estate investments is that real estate only accounts for a small fraction of the average pension asset allocation.

Altogether, the results in this section indicate that our findings are robust to the assumptions about the duration of pension liabilities and pension asset classes.

[insert Figure 1 here]

6.3. Robustness to negligibly small pension responsibilities

For some firms in our sample, the size of the pension plan is small compared to the size of the sponsoring firm. The pension duration gap might therefore not have a first-order impact on equity returns of these firms. According to Jin, Merton, and Bodie (2006), these observations are not likely to add information to the analysis, which could dilute the fit of our regressions to estimate the true relation between pension duration gap and interest rate exposure. We address this concern by repeating our analysis for different subsamples of firms with comparably large pension plans. We measure the size of pension plans by both the ratio of pension liabilities (PBO) to book value of firm assets and the pension duration gap.

Table 5 displays the results. In Columns (1) and (2), we exclude firm-years, where the ratio of PBO to firm assets is smaller than the 10^{th} and 20^{th} percentile of the pooled distribution, respectively. Similarly, Columns (3) and (4) report the results for a sample of firm-years where the pension duration gap is larger than in 10 and 20 percent of the observations, respectively. In all regressions, the coefficient on the pension duration gap (PBO) is positive and not statistically different from one. We receive a similar result if we measure pension liabilities by the ABO (not tabulated). We thus conclude that our results are not distorted by firms with negligibly small pension plans.

[insert Table 5 here]

6.4. Robustness to financial distress

This section considers whether our results could be driven by financial distress of the pension sponsoring firms. According to Rauh (2009), financially distressed firms try to reduce the expected cost of bankruptcy by allocating a larger portion of pension assets to bonds, which implies a reduction of the pension duration gap. Nance, Smith, and Smithson (1993) find that distressed firms increase their hedging activities in general. The positive relation between the pension duration gap and the interest rate exposure of the sponsoring firm might therefore be caused by financially distressed firms that simultaneously reduce the interest rate risk of their business and their pension plan. We test for this potential bias by examining the relation between pension duration gap and corporate interest rate exposure for different subsamples of nondistressed firms only. We use the same indicators of financial distress as Jin, Merton, and Bodie (2006), namely book to market value, return on investment, and financial leverage.[16] In each sample year, we exclude either the decile or the quintile of firms that appear to be most severely financially distressed the year before. These are the firms with the highest book to market ratio, the highest financial leverage, and the lowest return on investment, respectively.[17] In total, we consider six different subsamples of non-distressed firms.

Table 6 presents the results. Columns (1) and (2) sort the firms by their book to market value, Columns (3) and (4) by return on investment, and Columns (5) and (6) by financial leverage. In all regressions, the pension duration gap (PBO) coefficient is positive and not statistically different from one. In not tabulated regressions, we receive a similar result for the ABO based pension duration gap measure. These results indicate that the relation between interest rate exposure and duration gap is not caused by financial distress of the sponsoring firms.

[insert Table 6 here]

6.5. Robustness to the sample period

Our sample period (2003 to 2012) covers a fairly dynamic interest rate environment characterized by large macroeconomic disturbances, central bank interventions, and a massive decrease in interest rates. Bartram (2002) finds that the interest rate exposure of nonfinancial firms differs considerably between different periods in time. This raises the concern that our results could heavily depend on our observation period. Therefore, we estimate the relation between duration gap and interest rate exposure for different subperiods of our sample. First,

[16] The definitions are given in Table 9 of Appendix B.

[17] The value of the 10th percentile of return on investment is negative in all sample years. The value of the 20th percentile is negative in the majority of the years.

we test whether our results are driven by the recent financial crisis. Second, we compare the pension duration gap coefficient in high interest rate periods with the estimate in a low interest rate period. Table 7 displays the estimates of regressions on the PBO based duration gap measure. However, we receive similar results if we measure pension liabilities by the ABO (not tabulated).

In Column (1), we exclude the year 2008, when Lehman brothers filed for bankruptcy and global stock markets collapsed. In Column (2), we further exclude the year 2007, when the first manifestations of a liquidity crisis took place, and 2009, when the stock market bottomed out and started to recover. In Columns (3) and (4), we split the sample into years with comparably high interest rate levels (2004, 2005, 2006, 2007, and 2008) and years with comparably low interest rate levels (2003, 2009, 2010, 2011, and 2012). In all regressions, we find a positive pension duration gap coefficient that is not statistically different from one. This indicates that our findings are persistent over time and do not depend on a specific macroeconomic environment.

[insert Table 7 here]

6.6. Out of sample analyses

In this last section of our paper, we investigate the relation between interest rate exposure and pension duration gap in two alternative samples of pension sponsoring firms. Table 8 displays the results. Column (1) considers nonfinancial firms with pension plans that potentially hedge their interest rate risk with derivatives. As we discuss in section 5.2, these represent the firms that report a partial allocation of their pension assets to *other* assets. The duration gap coefficient in this sample is significantly negative, which conflicts with our prediction that, under the assumption of market efficiency, the estimate should be equal to one. However, as we argued before, we cannot rule out that this result is due to unobserved derivative hedging. We therefore believe that this evidence is not sufficient to reject the hypothesis of efficient capital markets.

Column (2) displays the estimates for financial firms with pension plans that do not hedge their interest rate risk with derivatives. Although the coefficient on pension duration gap is negative, it is not significantly different from one. This is because the standard error is approximately six times as large as in our main sample of nonfinancial pension sponsoring firms. Thus, we cannot reject the market efficiency hypothesis based on this sample either.

[insert Table 8 here]

7. Conclusion

This paper examines whether the interest rate exposure of nonfinancial firms is related to the systematic interest rate risk of their pension plans, and whether the relation is one-to-one. The interest rate sensitivity of DB pension plans is among the biggest nonoperating risks of corporate America. Based on publicly available information and controlling for other factors that potentially affect interest rate exposures, we find that equity returns reflect the duration gap between pension assets and liabilities without bias – even in cases where previous research has questioned market efficiency. Our finding is economically important and robust to the assumptions about the duration of pension assets and liabilities. Moreover, our analysis is not distorted by negligibly small pension plans, financial distress of the sponsoring firms, or changes in the macroeconomic environment.

This paper contributes to the ongoing discussion about the efficiency of capital markets with respect to information about DB pension plans. We extend this literature to systematic interest rate risk, which represents a risk factor that has not yet been fully understood in its impact on nonfinancial firms. Therefore, we also contribute to the comparatively limited literature on the interest rate exposure of nonfinancial firms.

Since our data do not provide us with information on derivative hedging in corporate DB pension plans, we limit our main analysis to firms whose pension plans do not invest in derivative securities at all. Still, we cannot exclude that these firms hedge the interest rate risk of their pension plans on their own accounts. However, even if this was the case, it could not explain our findings. On the contrary, it would rather weaken the relation between pension duration gap and interest rate exposure of the firm. Future research might have access to more comprehensive data on interest rate hedging and might therefore be able to test the relation between pension duration gap and interest rate exposure of the sponsoring firm in a larger sample of firms.

Appendix A: Derivation of the equations in section 3

If we decompose β_{OA}, β_{D}, β_{PA} and β_{PL} in analogy to equation (4) into their equity and their interest rate risk components, we can rewrite equation (1) by

$$\begin{aligned}\beta'_{E_j} + \gamma_{E_j} \cdot \beta_I = \frac{OA_j}{E_j}\left(\beta'_{OA_j} + \gamma_{OA_j} \cdot \beta_I\right) - \frac{D_j}{E_j}\left(\beta'_{D_j} + \gamma_{D_j} \cdot \beta_I\right) \\ + \left[\frac{PA_j}{E_j}\left(\beta'_{PA_j} + \gamma_{PA_j} \cdot \beta_I\right) - \frac{PL_j}{E_j}\left(\beta'_{PL_j} + \gamma_{PL_j} \cdot \beta_I\right)\right]. \qquad \text{(A1)}\end{aligned}$$

Rearranging the expression yields

$$\beta'_{E_j} + \gamma_{E_j} \cdot \beta_I = \frac{OA_j}{E_j}\beta'_{OA_j} - \frac{D_j}{E_j}\beta'_{D_j} + \left[\frac{PA_j}{E_j}\beta'_{PA_j} - \frac{PL_j}{E_j}\beta'_{PL_j}\right]$$
$$+ \frac{OA_j}{E_j}\gamma_{OA_j} \cdot \beta_I - \frac{D_j}{E_j}\gamma_{D_j} \cdot \beta_I + \left[\frac{PA_j}{E_j}\gamma_{PA_j} \cdot \beta_I - \frac{PL_j}{E_j}\gamma_{PL_j} \cdot \beta_I\right], \tag{A2}$$

which we decompose into an interest rate risk part

$$\gamma_{E_j} \cdot \beta_I = \frac{OA_j}{E_j}\gamma_{OA_j} \cdot \beta_I - \frac{D_j}{E_j}\gamma_{D_j} \cdot \beta_I + \left[\frac{PA_j}{E_j}\gamma_{PA_j} \cdot \beta_I - \frac{PL_j}{E_j}\gamma_{PL_j} \cdot \beta_I\right] \tag{A3}$$

and an equity risk part

$$\beta'_{E_j} = \frac{OA_j}{E_j}\beta'_{OA_j} - \frac{D_j}{E_j}\beta'_{D_j} + \left[\frac{PA_j}{E_j}\beta'_{PA_j} - \frac{PL_j}{E_j}\beta'_{PL_j}\right]. \tag{A4}$$

If $\beta_I \neq 0$, we can divide equation (A3) by β_I, which yields an interest rate risk expression analogous to the expression for systematic risk in equation (1):

$$\gamma_{E_j} = \frac{OA_j}{E_j}\gamma_{OA_j} - \frac{D_j}{E_j}\gamma_{D_j} + \left[\frac{PA_j}{E_j}\gamma_{PA_j} - \frac{PL_j}{E_j}\gamma_{PL_j}\right] \tag{5}$$

Appendix B: Tables and figures

Table 1: Descriptive sample statistics

This table shows the descriptive sample statistics of our main variables. Variable definitions are in Table 9. The data refer to 2003 to 2012. In Panels A, B, and C, the sample is restricted to firms that sponsor defined benefit pension plans that do not hedge their interest rate risk with derivatives. The sample in Panel D consists of both pension and nonpension COMPUSTAT firms. All variables are winsorized at the 1st and the 99th percentile of their pooled sample distribution.

	Mean	Median	Min	10th %ile	90th %ile	Max	Std.	N
Panel A: Main Variables								
Interest rate exposure	1.903	0.974	-58.597	-16.469	22.523	63.194	18.051	1,195
Pension duration gap (PBO)	3.187	1.307	0.025	0.207	7.313	42.674	5.912	1,195
Pension duration gap (ABO)	3.220	1.299	0.018	0.221	7.459	42.504	6.023	954
Panel B: Pension plan characteristics								
Pension liabilities (PBO) to firm assets	0.157	0.106	0.005	0.017	0.351	0.896	0.167	1,195
Funding ratio	0.758	0.757	0.184	0.527	0.978	1.492	0.199	1,195
Asset allocation to equities (in %)	59.855	62.000	0.000	42.000	74.000	99.040	15.455	1,195
Asset allocation to bonds (in %)	39.116	37.000	0.960	24.300	56.600	100.000	15.534	1,195
Asset allocation to real estate (in %)	1.006	0.000	0.000	0.000	5.000	12.016	2.633	1,195
Panel C: Firm characteristics								
Financial leverage	0.271	0.249	0.000	0.014	0.515	0.941	0.199	1,195
Short-term debt to equity	0.069	0.010	0.000	0.000	0.137	1.830	0.221	1,195
Long-term debt to equity	0.579	0.242	0.000	0.000	1.342	7.855	1.096	1,195
Cash flow to assets	0.078	0.082	-0.418	-0.008	0.179	0.329	0.098	1,195
Book value of firm assets (in $m)	4,570	1,501	27	164	13,494	55,746	8,778	1,195
Panel D: Characteristics of COMPUSTAT firms								
Interest rate exposure	0.667	0.217	-59.993	-18.477	20.728	61.925	18.325	21,510
Financial leverage	0.230	0.194	0.000	0.000	0.500	0.907	0.198	21,510
Short-term debt to equity	0.202	0.018	0.000	0.000	0.440	5.163	0.657	21,510
Long-term debt to equity	0.608	0.230	0.000	0.000	1.359	8.939	1.246	21,510
Cash flow to assets	0.057	0.063	-0.510	-0.021	0.160	0.377	0.113	21,510
Book value of firm assets (in $m)	9,381	1,387	10	96	19,125	227,097	29,171	21,510

Table 2: Pearson correlation matrix

This table shows the pairwise Pearson correlations between selected variables. Variable definitions are in Table 9. The data refer to 2003 to 2012. The sample is restricted to firms that sponsor defined benefit pension plans that do not hedge their interest rate risk with derivatives. The asterisk denotes statistical significance at the 10% level using a two-tailed test. All variables are winsorized at the 1st and the 99th percentile of their pooled sample distribution.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
(1) Interest rate exposure	1						
(2) Pension duration gap (PBO)	0.1531*	1					
(3) Pension duration gap (ABO)	0.1566*	0.9974*	1				
(4) Short-term debt to equity	0.0483*	0.4493*	0.4746*	1			
(5) Long-term debt to equity	0.1151*	0.4877*	0.5006*	0.3442*	1		
(6) Cash flow to assets	-0.0139	-0.3548*	-0.3711*	-0.3375*	-0.3487*	1	
(7) Firm size	-0.0341	-0.0833*	-0.0902*	-0.0760*	0.0547*	0.1804*	1

Table 3: The relation between interest rate exposure and the pension duration gap

This table reports the results of our main regressions of interest rate exposure on pension duration gap and controls. Variable definitions are in Table 9. The data refer to 2003 to 2012. The sample is restricted to firms that sponsor defined benefit pension plans that do not hedge their interest rate risk with derivatives. All variables are winsorized at the 1st and the 99th percentile of their pooled distribution. Asterisks denote statistical significance at the 1% (***), 5% (**), and 10% (*) level using a two-tailed test. Standard errors (in parentheses) are clustered at the firm level. The p-value at the bottom of the table refers to a two-tailed Wald test of whether the coefficient on the pension duration gap equals one.

Dependent variable	Interest rate exposure					
	(1)	(2)	(3)	(4)	(5)	(6)
Pension duration gap (PBO)	0.775**		1.283***		1.417***	
	(0.354)		(0.308)		(0.284)	
Pension duration gap (ABO)		0.864**		1.319***		1.295***
		(0.370)		(0.314)		(0.313)
Short-term debt to equity	-10.774	-12.641	-8.523	-9.284	-10.282	-11.133
	(9.132)	(12.017)	(9.156)	(11.785)	(9.950)	(12.046)
Long-term debt to equity	0.839	0.225	0.866	0.116	-0.365	0.024
	(1.898)	(1.803)	(2.030)	(1.947)	(1.732)	(1.815)
Cash flow to assets			40.235***	39.416**	40.491***	41.571**
			(13.080)	(15.881)	(14.537)	(16.966)
Firm size			7.461**	7.047**	10.132***	8.840**
			(3.050)	(3.375)	(3.388)	(3.891)
Lag interest rate exposure			-0.146**	-0.144**	-0.157***	-0.148**
			(0.058)	(0.072)	(0.059)	(0.073)
Market share (in %)					-23.353	-28.237
					(18.976)	(20.070)
Capital intensiveness					0.667	1.700
					(17.379)	(20.085)
Cash position					16.199	11.252
					(16.383)	(19.593)
Growth rate					-0.193	-1.371
					(5.283)	(6.129)
Liquidity					-0.873	-1.635
					(1.132)	(1.474)
Advertisement to assets					11.036	-80.366
					(55.474)	(118.091)
R&D to assets					48.034	78.839
					(79.347)	(76.313)
Firm fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1'195	954	1,195	954	1,168	938
R2 adjusted	0.074	0.074	0.116	0.114	0.118	0.112
p-value (b = 1)	0.526	0.713	0.360	0.310	0.143	0.347

Table 4: Impact of the pension funding ratio

This table reports the results of regressions of interest rate exposure on pension duration gap and controls for firms with low pension funding. Variable definitions are in Table 9. The data refer to 2003 to 2012. The sample is restricted to firms that sponsor defined benefit pension plans that do not hedge their interest rate risk with derivatives. All variables are winsorized at the 1st and the 99th percentile of their pooled distribution. Asterisks denote statistical significance at the 1% (***), 5% (**), and 10% (*) level using a two-tailed test. Standard errors (in parentheses) are clustered at the firm level. The p-value at the bottom of the table refers to a two-tailed Wald test of whether the coefficient on the pension duration gap equals one.

Dependent variable	Interest rate exposure			
Include the P^{th} percentile of observations with the lowest funding ratio	P = 80 (1)	P = 60 (2)	P = 40 (3)	P = 20 (4)
Pension duration gap (PBO)	1.231***	1.104**	1.392***	1.561***
	(0.312)	(0.501)	(0.516)	(0.485)
Short-term debt to equity	-12.170	-9.642	2.153	18.984
	(10.258)	(10.998)	(22.074)	(11.996)
Long-term debt to equity	1.257	2.247	2.627	2.217
	(2.267)	(2.635)	(2.834)	(3.762)
Cash flow to assets	33.140**	21.202	15.772	21.269
	(16.372)	(20.057)	(25.139)	(32.541)
Firm size	5.710	8.593*	6.849	6.949
	(3.668)	(4.435)	(5.307)	(8.397)
Lag interest rate exposure	-0.172**	-0.204**	-0.181	-0.342**
	(0.068)	(0.085)	(0.125)	(0.157)
Firm fixed effects	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
Observations	956	717	478	239
R2 adjusted	0.114	0.140	0.062	0.288
p-value (b = 1)	0.459	0.836	0.449	0.250

Table 5: Robustness to negligibly small pension responsibilities

This table reports the results of regressions of interest rate exposure on pension duration gap and controls excluding firms with negligibly small pension responsibilities. Variable definitions are in Table 9. The data refer to 2003 to 2012. The sample is restricted to firms that sponsor defined benefit pension plans that do not hedge their interest rate risk with derivatives. All variables are winsorized at the 1st and the 99th percentile of their pooled distribution. Asterisks denote statistical significance at the 1% (***), 5% (**), and 10% (*) level using a two-tailed test. Standard errors (in parentheses) are clustered at the firm level. The p-value at the bottom of the table refers to a two-tailed Wald test of whether the coefficient on the pension duration gap equals one.

Dependent variable	Interest rate exposure			
Include observations where	PBO to firm assets >		Pension duration gap (PBO) >	
	10th percentile (1.7%)	20th percentile (3.1%)	10th percentile (0.21)	20th percentile (0.38)
	(1)	(2)	(3)	(4)
Pension duration gap (PBO)	1.132***	1.202***	1.231***	1.247***
	(0.334)	(0.355)	(0.325)	(0.308)
Short-term debt to equity	-7.020	-6.197	-8.187	-9.272
	(9.924)	(11.179)	(9.260)	(9.567)
Long-term debt to equity	1.576	0.468	0.830	0.141
	(2.392)	(2.518)	(2.083)	(1.763)
Cash flow to assets	38.931***	38.077***	39.241***	32.746**
	(14.203)	(14.416)	(13.729)	(14.411)
Firm size	4.862	4.745	4.584	4.350
	(3.243)	(3.495)	(3.167)	(3.630)
Lag interest rate exposure	-0.137**	-0.142**	-0.144**	-0.150**
	(0.063)	(0.067)	(0.065)	(0.070)
Firm fixed effects	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
Observations	1,076	956	1,076	956
R2 adjusted	0.107	0.133	0.110	0.097
p-value (b = 1)	0.693	0.569	0.478	0.424

Table 6: Robustness to financial distress

This table reports the results of regressions of interest rate exposure on pension duration gap and controls excluding firms that are potentially financially distressed. Variable definitions are in Table 9. The data refer to 2003 to 2012. The sample is restricted to firms that sponsor defined benefit pension plans that do not hedge their interest rate risk with derivatives. All variables are winsorized at the 1st and the 99th percentile of their pooled distribution. Asterisks denote statistical significance at the 1% (***), 5% (**), and 10% (*) level using a two-tailed test. Standard errors (in parentheses) are clustered at the firm level. The p-value at the bottom of the table refers to a two-tailed Wald test of whether the coefficient on the pension duration gap equals one.

Dependent variable	Interest rate exposure					
Measure of financial distress	Book to market ratio		Return on investment		Financial leverage	
In each year, exclude the Pth percentile of the most financially distressed firms the year before	10	20	10	20	10	20
	(1)	(2)	(3)	(4)	(5)	(6)
Pension duration gap (PBO)	1.001**	1.024*	1.087***	1.049*	1.294***	1.344***
	(0.467)	(0.578)	(0.404)	(0.584)	(0.296)	(0.288)
Short-term debt to equity	-10.435	-8.731	-14.834*	-10.526	-18.620**	-13.112
	(12.566)	(14.855)	(8.228)	(9.420)	(8.431)	(8.806)
Long-term debt to equity	2.059	0.635	0.435	-0.320	1.081	1.212
	(2.073)	(2.206)	(2.213)	(2.290)	(1.993)	(3.111)
Cash flow to assets	38.844***	44.222**	50.766***	44.206***	39.128***	38.060**
	(14.471)	(18.245)	(13.669)	(14.954)	(13.828)	(14.955)
Firm size	5.881*	7.259**	7.206**	6.975**	6.833**	5.836*
	(3.059)	(3.261)	(3.251)	(3.266)	(3.321)	(3.290)
Lag interest rate exposure	-0.152**	-0.187***	-0.140**	-0.144**	-0.136**	-0.162***
	(0.062)	(0.071)	(0.056)	(0.065)	(0.062)	(0.060)
Firm fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1,078	959	1,079	960	1,078	959
R2 adjusted	0.108	0.110	0.123	0.138	0.105	0.130
p-value (b = 1)	0.998	0.968	0.829	0.933	0.321	0.234

Table 7: Robustness to financial crisis

This table reports the results of regressions of interest rate exposure on pension duration gap and controls for different macroeconomic environments. Variable definitions are in Table 9. The data refer to 2003 to 2012. The sample is restricted to firms that sponsor defined benefit pension plans that do not hedge their interest rate risk with derivatives. All variables are winsorized at the 1st and the 99th percentile of their pooled distribution. Asterisks denote statistical significance at the 1% (***), 5% (**), and 10% (*) level using a two-tailed test. Standard errors (in parentheses) are clustered at the firm level. The p-value at the bottom of the table refers to a two-tailed Wald test of whether the coefficient on the pension duration gap equals one.

Dependent variable	Interest rate exposure			
Included years	All ex 2008	All ex. 2007 to 2009	2004 to 2008	2003 and 2009 to 2012
	(1)	(2)	(3)	(4)
Pension duration gap (PBO)	1.152***	0.925**	0.683**	1.269**
	(0.307)	(0.428)	(0.300)	(0.563)
Short-term debt to equity	2.730	-5.157	-6.946	1.278
	(5.874)	(11.459)	(8.807)	(17.533)
Long-term debt to equity	3.166***	4.792*	0.940	1.761
	(0.805)	(2.613)	(1.712)	(3.552)
Cash flow to assets	43.418***	39.255**	22.594**	52.495
	(9.588)	(19.956)	(11.116)	(35.020)
Firm size	7.646***	6.618*	1.917	10.934*
	(1.837)	(3.648)	(3.840)	(6.046)
Lag interest rate exposure	-0.166***	-0.188***	-0.284***	-0.204*
	(0.038)	(0.063)	(0.054)	(0.105)
Firm fixed effects	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
Observations	1'077	964	605	590
R2 adjusted	0.121	0.188	0.252	0.061
p-value (b = 1)	0.633	0.861	0.293	0.633

Table 8: Out of sample analyses

This table reports the results of regressions of interest rate exposure on pension duration gap and controls for financial firms and firms with pension plans that likely hedge their interest rate risk with derivatives, respectively. Variable definitions are in Table 9. The data refer to 2003 to 2012. All variables are winsorized at the 1st and the 99th percentile of their pooled distribution. Asterisks denote statistical significance at the 1% (***), 5% (**), and 10% (*) level using a two-tailed test. Standard errors (in parentheses) are clustered at the firm level. The p-value at the bottom of the table refers to a two-tailed Wald test of whether the coefficient on the pension duration gap equals one.

Dependent variable	Interest rate exposure	
Sample	Nonfinancial firms with plans that potentially hedge with derivatives	Financial firms with plans that do not hedge with derivatives
	(1)	(2)
Pension duration gap (PBO)	-0.171**	-1.993
	(0.078)	(1.888)
Short-term debt to equity	3.285	-3.485
	(3.248)	(2.197)
Long-term debt to equity	0.071	-0.194
	(0.624)	(1.826)
Cash flow to assets	15.024***	-14.297
	(5.176)	(20.295)
Firm size	-0.999	-4.744
	(1.202)	(4.247)
Lag interest rate exposure	-0.146***	-0.185
	(0.021)	(0.118)
Firm fixed effects	Yes	Yes
Year fixed effects	Yes	Yes
Observations	7,296	426
R2 adjusted	0.082	0.048
p-value (b = 1)	0.000	0.117

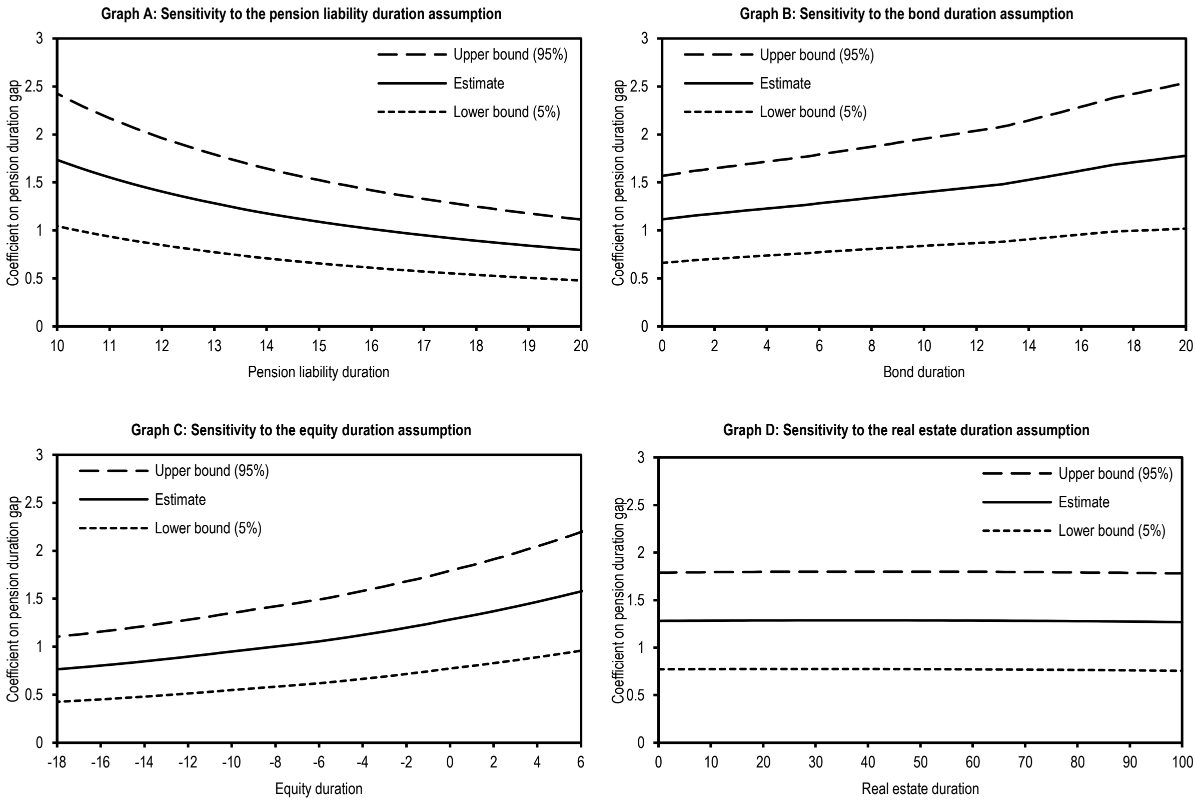
Table 9: Variable definitions

This table summarizes the variable definitions. The data are from the from the COMPUSTAT North American Pension database, the COMPUSTAT/CRSP Merged database, the CRSP daily fixed term index files, and the CRSP daily stock files.

Variable	Calculation (incl. COMPUSTAT mnemonics)
	Panel A: Main variables
Interest rate exposure ($\hat{\gamma}_E$)	The interest rate exposure of the firm's equity is the sum of the partial sensitivities of stock returns to absolute changes in the yield to maturity of the 1-, 5-, 10- and 30-year CRSP fixed term index, while controlling for the return on the CRSP value-weighted total stock market index. The regression run is: $R_{E_{j,t}} = \alpha_j + \beta'_{E_j} \cdot R_{EM_t} + \gamma_{E1_j}\Delta I_{1,t} + \gamma_{E5_j}\Delta I_{5,t} + \gamma_{E10_j}\Delta I_{10,t} + \gamma_{E30_j}\Delta I_{30,t} + \epsilon_{j,t}$ The estimation is made for each firm and year based on weekly CRSP data (up to 52 observations). All returns are continuously compounded. We exclude stocks that have not been traded for more than 43 weeks during a year. $\hat{\gamma}_{E_j} = \gamma_{E1_j} + \gamma_{E5_j} + \gamma_{E10_j} + \gamma_{E30_j}$
Pension duration gap ($\hat{\gamma}_{Pension}$)	The pension duration gap is the difference between the duration of pension assets (normalized by the ratio of pension assets to firm equity) and the duration of pension liabilities (normalized by the ratio of pension liabilities to firm equity). The duration of pension assets is the weighted average duration of pension assets invested in equities (pnate), bonds (pnatd), and real estate (pnatr). Pension assets are measured by their market value (pplao+pplau). Pension liabilities are primarily measured by the PBO (pbpro + pbpru). In an alternative specification, pension liabilities are measured by the ABO (pbaco + pbacu). The duration assumptions are: Duration Pension liabilities = 13 Duration equities = 0 Duration bonds = 6 Duration real estate = 6 Pension duration gap (PBO) $= \left(\frac{(0 \cdot \text{pnate} - 6 \cdot \text{pnatd} - 6 \cdot \text{pnatr})(\text{pplao+pplau}) - (-13(\text{pbpro+pbpru}))}{100(\text{csho} \cdot \text{prcc\_f})}\right)$ Pension duration gap (ABO) $= \left(\frac{(0 \cdot \text{pnate} - 6 \cdot \text{pnatd} - 6 \cdot \text{pnatr})(\text{pplao+pplau}) - (-13(\text{pbaco +pbacu}))}{100(\text{csho} \cdot \text{prcc\_f})}\right)$
	Panel B: Control variables
Short-term debt to equity	Book value of short-term debt (dlc) divided by market value of equity (csho x prcc_f).
Long-term debt to equity	Book value of long-term debt (dltt) divided by market value of equity (csho x prcc_f).
Cash flow to assets	Sum of net income and depreciation and amortization (ni + dp) divided by total assets (at).
Firm size	Natural logarithm of total assets (at).
Market share (in %)	Sales (sale) divided by total sales of firms in the same Fama-French 48 industry multiplied by hundred.
Capital intensiveness	Current assets (act) divided by total assets (at).
Cash position	Cash and short-term investments (che) divided by total assets (at).
Growth rate	Natural logarithm of total assets divided by lagged total assets (ln(at / at_{t-1})).
Liquidity	Current assets (act) divided by current liabilities (lct).
Advertisement to assets	Advertising expense (xad) divided by total assets (at). If advertising expense is missing it is set to zero.
R&D to assets	Research and development expense (xrd) divided by total assets (at). If research and development expense is missing it is set to zero.
	Panel C: Further variables
PBO to firm assets	The PBO (pbpro + pbpru) divided by total assets (at).
Funding ratio	Pension assets (pplao + pplau) divided by the PBO (pbpro + pbpru).
Book-market ratio	Book value of equity (at-dlc-dltt) divided by market value of equity (csho x prcc_f).
Return on investment	Net income (ni) divided by total assets (at).
Financial leverage	Book value of debt (dltt + dlc) divided by total assets (at).

Figure 1: Coefficient on pension duration gap as a function of the duration assumptions

This figure plots the coefficient on the pension duration gap as a function of the assumption about the duration of pension assets and liabilities. We run regressions of interest rate exposure on pension duration gap and controls, including year and firm fixed effects. We measure pension liabilities by the PBO. The set of control variables includes short-term debt to equity, long-term debt to equity, cash flow to assets, firm size, and the one-year lag of interest rate exposure. Variable definitions are in Table 9. The data refer to 2003 to 2012. The sample is restricted to firms that sponsor defined benefit pension plans that do not hedge their interest rate risk with derivatives. Graph A displays the duration gap coefficient as a function of the pension liability duration assumption, Graphs B, C, and D plot the estimate as a function of the assumption about the duration of pension assets invested in bonds, equities, and real estate, respectively. All variables are winsorized at the 1st and the 99th percentile of their pooled distribution. Standard errors are clustered at the firm level.



III. How the Chairman's Personal Preferences Affect Public Pension Risk

Oliver Dichter*

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Based on the analysis of 343 changes of chairpersons in 110 U.S. state and local government pension boards of trustees, this paper shows that the risk from the mismatch between pension assets and liabilities reflects the personal risk preferences of the chairman of the board (COB). We find that pension risk is negatively affected by an increase in COB age, and that it is lower if the COB is a woman. We also find that pension risk is higher if the COB is an annuitant of the plan, consistent with an incentive of retirees to gamble for higher benefits. Finally, we observe that the risk of public pension plans is higher if the COB is an ex officio trustee, possibly because reporting rules enable politicians to avoid tax increases or spending cuts by boosting the risk of the fund's assets. The current underfunding problems faced by public pension plans are hence partially a consequence of past decisions of pension COBs. Our results are robust to different definitions of pension risk, economically relevant, and particularly strong for more poorly governed pension plans. We find no evidence of an endogenous selection of COBs to pension plans that match their risk preferences.

Keywords: Public pension plans; Asset-liability management; Chairman of the board; Risk preferences; Governance

JEL codes: G23, G34, J14, J16

*Oliver Dichter (corresponding author, oliver.dichter@ifm.unibe.ch), the Institut für Finanzmanagement, University of Bern. This paper is part of my dissertation with Claudio Loderer, the Institut für Finanzmanagement, University of Bern and the Swiss Finance Institute. I wish to thank Claudio Loderer for valuable comments and suggestions. All errors are mine.

1. Introduction

Failure to take account of the mismatch between the assets in defined benefit pension plans (primarily equities) and the liabilities (deferred fixed annuities) has long been a major unrecognized source of financial instability. The underfunding problems now facing state and local government pension plans [...] are a direct consequence of this conceptual and practical failure.

– Zivi Bodie (2012), International Journal of Central Banking –

The vast majority of public defined benefit (DB) pension plans in the U.S. are substantially underfunded. In 2013, the asset value of the 126 largest public DB pension plans equaled USD 2.86 trillion, while the reported pension liabilities amounted to USD 3.81 trillion (Public Fund Survey (2015)). The economic value of pension obligations was even 15 to 40 percent higher (Novy-Marx and Rauh (2011)). In most states of the U.S., the deficit in pension funding is larger than the general obligation debt and equals several years of tax revenues (Novy-Marx and Rauh (2011)). The shortfall in public pension funding has also been a major contributor to some of the most recent bankruptcies of U.S. cities, including Detroit (IL), Harrisburg (PA), Mammoth Lakes (CA), Stockton (CA), and Central Falls (RI) (Spangler (2013)). Unfunded public pension liabilities are thus a huge burden on tax payers. Moreover, these shortfalls imply that future generations of taxpayers must pay for today's pension liabilities, which violates the fundamental principle of public finance that each generation should pay for the services it consumes (Bader (2015)).

The main cause of the underfunding problems of public pension plans is the historic mismatch between pension assets and liabilities (Bodie (2012)). According to Pennacchi and Rastad (2011), this asset-liability mismatch determines the total financial risk of a DB pension plan. Despite the far-reaching consequences of this risk, we know little about its determinants

(Bodie (2012)). Existing research, including Cronqvist, Makhija, and Yonker (2012), Baxamusa and Jalal (2015), and Cain and McKeon (2015) shows that the risk of listed firms is explained by the private preferences of corporate executives. This paper investigates if pension risk is partially explained by the personal preferences of pension executives as well. We focus on the chairman of the board (COB) because the chairperson represents the pension executive with the strongest influence on asset-liability decisions. While the chief investment officer (CIO) is involved in the everyday management of pension assets, the COB is the president of the board that effectively determines investment allocations, actuarial valuations, system operations, and often plan benefits (Mitchell (2001)). The pension COB is comparable to a corporate chairman with his superior access to new information and his substantial influence on meeting agendas and committee decisions (Parker (1990)).[18]

To the best of our knowledge, we are the first to inquire into whether the individual preferences of the COB explain pension risk. Existing empirical studies typically rely on pension plan or board characteristics to explain risk taking at pension plans but largely ignore the possible role of individual executives. Pennacchi and Rastad (2011) show that pension risk increases with the fraction of beneficiaries on the pension board of trustees, possibly because beneficiaries have an incentive to gamble for higher benefits. Park (2009), Weller and Wenger (2009), and Mohan and Zhang (2014) present evidence that pension managers tend to follow trends and peer group norms in their risk taking decisions. There is also a wide literature on moral hazard of pension managers, which argues that public pension accounting rules allow managers to manipulate the value of the pension liabilities and the amount of required pension contributions (e.g., Lucas and Zeldes (2009), Pennacchi and Rastad (2011), Novy-Marx and Rauh (2011), and Mohan and Zhang (2014)).

Our estimation of the risk from mismatched pension assets and liabilities closely follows the method suggested by Pennacchi and Rastad (2011). Thus, we define pension risk

[18] If not stated differently, masculine pronouns refer to women and men alike.

by the volatility of the return difference between pension assets and liabilities. We find that pension assets are generally poorly matched to pension liabilities. On average, pension risk therefore even exceeds the volatility of equity market returns. We test whether the personal risk preferences of COBs affect the risk of the plans they are responsible for based on four empirical predictions derived from the literature. (1) Pension risk is negatively affected by an increase in COB age, (2) pension risk is lower if the COB is a woman, (3) pension risk is higher if the COB is an annuitant (retiree of the plan), and (4) pension risk is higher if the COB is a politician (ex officio trustee).

According to a substantial body of literature in finance and economics, including Campbell (2006), Serfling (2014), and Berger, Kick, and Schaeck (2014), risk taking is decreasing with an individual's age. We therefore predict that pension risk is negatively affected by an increase in the COB's age.

There is also a large literature on gender specific differences in risk attitudes, which mostly finds that women are intrinsically less risk loving than men (Apicella, Dreber, Campbell, Gray, Hoffman, and Little (2008), Sapienza, Zingales, and Maestripieri (2009), and Croson and Gneezy (2009)). Two notable exceptions are Adams and Funk (2012) and Berger, Kick, and Schaeck (2014), who show that female board members of large, listed firms seek higher levels of risk than their male counterparts. Adams and Funk (2012) explain this finding by the personal costs of women who choose a career that eventually leads to a board seat. Women who decide to bear these costs are less tradition and security oriented than the average woman in the population (Adams and Funk (2012)). Another study deviating from the notion that women are universally less risk seeking than men is the survey of Jörg (2005). It shows that women are only more risk averse than men in investment type decisions but appear to be more risk loving than men in insurance situations. For public pension plans, however, we expect that female COBs are less risk loving than their male counterparts. This is because we cannot observe that the women in our sample face a trade-off between family and career as they are often rank and file employees of the pension sponsor. Furthermore, existing empirical evidence on risk taking in personal retirement accounts indicates that women prefer lower levels of pension risk than male account holders (Sundén and Surette (1998) and Agnew, Balduzzi, and Sundén (2003)).

Our prediction that pension risk is higher if the COB is an annuitant of the pension plan is based on the fact that payments to retirees are upward flexible, while at the same time they are downside protected (Monahan (2010)). Such optionality creates an incentive to speculate for higher benefits by increasing the risk of the pension plan (Pennacchi and Rastad (2011)).

Finally, we expect that pension risk is higher if the COB is an ex officio trustee. Politicians are not concerned about long-term funding issues because they operate under a relatively short time horizon (Giertz and Papke (2007)). Knowing that they will likely be gone from office when a potential pension underfunding problem becomes critical, they will rather reduce public pension contributions than raise taxes or cut programs, (Giertz and Papke (2007)). The preference of politicians for low pension contributions entails a preference for high pension risk because pension accounting rules allow public pension sponsors to reduce their contributions by choosing a riskier asset allocation (Lucas and Zeldes (2009), Pennacchi and Rastad (2011), Novy-Marx and Rauh (2011), and Mohan and Zhang (2014)).

We test our empirical predictions in a difference-in-difference (DID) analysis. The DID estimator compares the pension risk in a treatment group to the pension risk in a control group both before and after treatment. It allows controlling for omitted variables that affect both groups in a similar way or remain constant over time (Berger, Kick, and Schaeck (2014)). In this study, the treatment group consists of observations where the COB risk preferences change in the current fiscal year. Board-years with no contemporaneous change in COB preferences form the control group. We address the concern that pension trustees might simultaneously decide on the risk of the pension plan and appoint a COB who matches their risk preferences by considering a reduced sample of pension plans where COB changes are exogenous.

We test our empirical predictions with data from the Public Plans Database (PPD) of the Center for Retirement Research at Boston College (2015), which covers 90 percent of the pension members and assets of state and local DB plans in the U.S. Our initial sample covers the years from 2001 to 2013 and consists of 1,310 observations about 110 pension boards of trustees. We then hand collected information on COBs from over 2,000 public information sources and personal contacts with the pension plans. We obtained complete COB information for 1,159 observations of our initial sample (88 percent). This sample covers 343 changes of chairpersons.

Our empirical results are consistent with the hypothesis that pension risk reflects the personal risk preferences of the COB. Pension risk is negatively affected by an increase in COB age, lower if the COB is a woman, higher if the COB is an annuitant, and higher if the COB is an ex officio trustee. We do not find that COBs are appointed by the board of trustees based on their personal attitude towards pension risk. However, we find evidence that pension risk is predominantly affected by COB risk preferences if pension governance is weaker. We consider a plan as more weakly governed if decisions on pension risk are not made by a separate and professionalized investment board. This finding is consistent with Cronqvist, Makhija, and Yonker (2012), who shows that CEOs primarily align the risk of the firm with their personal risk preferences when corporate governance is poor.

Our main contribution is that we are the first to present evidence that the personal risk preferences of pension COBs are a predictor of public pension plan risk. Since pension risk is the main driver of today's substantial funding shortfalls, taxpayers should be wary of who is in charge of the pension plans in their community.

The rest of the paper is organized as follows. In Section 2, we provide an overview of the public pension system in the U.S. In Section 3, we derive our empirical predictions and discuss the relevant literature. In Section 4, we describe the sample selection and how we measure pension risk. In Section 5, we discuss our empirical strategy. In Section 6, we present the results and their discussion. Last, we conclude in Section 7.

2. Public pension plans in the U.S.

While many firms in the U.S. have moved away from DB plans and opened defined contribution (DC) retirement schemes such as 401(k) plans instead, the public sector has seen very limited movement in this direction (Novy-Marx and Rauh (2011)). In 2014, total public DC assets (USD 533 billion) only amounted to 15 percent of state and local DB assets (USD 3.6 trillion) (Hoops, Stefanescu, and Vidangos (2015)). In a DB plan, the employer guarantees its employees specific benefits upon their retirement. If pension assets prove to be insufficient to cover these benefits, the sponsor has to make additional contributions. Therefore, DB plans represent a financial risk to their sponsors. This risk increases with the mismatch between pension assets and liabilities (Pennacchi and Rastad (2011)). The sponsor's responsibility under DC plans is fundamentally different. It only consists of the sponsor's commitment to pay regular and fixed contributions to the employees' retirement accounts. At retirement, the employees receive whatever amount of money (contributions plus interest) has accumulated on their behalf. The uncertainty about the level of retirement benefits lies entirely with the employees (Shivdasani and Stefanescu (2010)). We therefore exclude these retirement schemes from our analysis, which is consistent with previous research on public pension risk (e.g., Pennacchi and Rastad (2011) and Mohan and Zhang (2014)). Throughout this paper, we consequently use the terms *pension plan* and *defined benefit pension plan* interchangeably.

3. Literature review and empirical predictions

Recent empirical evidence shows that corporate executives imprint their private risk preferences on the firms they manage. Cronqvist, Makhija, and Yonker (2012) find that CEOs align the corporate capital structure with their preference for leverage. Cain and McKeon (2015) show that firms have a higher equity return volatility if the CEO possesses a private pilot's license, which they consider a proxy for personal risk-taking. Baxamusa and Jalal (2015) find that firms issue more debt and are geographically and operationally less diversified if the CEO plays a risky sport. Based on this evidence, we expect that pension COBs align the risk of the plan with their personal attitudes towards risk as well. In our empirical analysis of the relation between COB risk preferences and pension risk, we test four predictions that have been suggested by the literature. The following sections discuss these predictions in detail.

3.1. COB age

A large body of literature identifies a negative relation between individual age and risk taking. Campbell (2006) finds that older households invest a lower fraction of their total wealth into equity securities than younger households. Agnew, Balduzzi, and Sundén (2003) observe a similar pattern in over 7,000 401(k) accounts. In a survey among more than 500 business executives, MacCrimmon and Wehrung (1990) document a negative relation between executive age and corporate risk taking as well. This is supported by recent empirical evidence. Serfling (2014) shows that stock return volatility is negatively affected by CEO age and Berger, Kick, and Schaeck (2014) find that the risk of bank portfolios decreases with the fraction of older directors on the board.

There are also studies that contradict the view of higher risk aversion among older executives. Chevalier and Ellison (1999) find that younger mutual fund managers are less risk tolerant than their older colleges because they face a higher probability of being dismissed for poor performance. Hong, Kubik, and Solomon (2000) show that inexperienced security analysts are conservative in their forecasts because they are more likely terminated for inaccurate forecasts that deviate from the consensus estimates than their more experienced counterparts. We do however not expect career concerns to have a first order impact on the personal risk preference of pension COBs. First of all, pension COBs are typically much older than managers and analysts of mutual funds, which makes it less likely that a onetime negative event destroys their reputation. While the average fund manager in Chevalier and Ellison (1999) is 44 of age, the mean age of our pension COBs is 57 years. Secondly, we cannot observe that past performance is of any consequence to pension COB replacements. Table 1 shows that the probability the COB is replaced after a year of relatively bad investment performance does not differ from the probability that he is replaced after a year of good performance. Past performance is also irrelevant for COB replacements if we only consider the 50 percent youngest or the 25 percent youngest COBs in our sample.[19] We therefore predict that pension risk decreases with COB age.

[insert Table 1 here]

3.2. COB gender

Studies on gender differences in risk attitudes mostly find that women prefer lower levels of risk than men (Croson and Gneezy (2009)). In their recent analysis of risk taking in banks, Palvia, Vähämaa, and Vähämaa (2014) document that female CEOs choose less risky capital structures (higher Tier 1 capital and higher equity capital) than men. Apicella, Dreber,

[19] We receive a similar result if we consider three years of past performance instead of one year.

Campbell, Gray, Hoffman, and Little (2008) and Sapienza, Zingales, and Maestripieri (2009)) find that risk taking of individuals is increasing in their testosterone level. Dwyer, Gilkeson, and List (2002) and Jörg (2005) present evidence that the higher risk-aversion of women is explained by knowledge disparities.

In contradiction to the notion of greater risk aversion among women, Adams and Funk (2012) and Berger, Kick, and Schaeck (2014) find that female directors of large, listed firms are more risk loving than male directors because the trade-off between having a family and choosing a career path that eventually leads to a board seat is more costly for women than for men. Women who choose career over family are less tradition and security oriented, and therefore less risk averse than their male counterparts. More risk loving women thus self-select into the pool of eligible director candidates (Adams and Funk (2012)). In our sample of public pension COBs, career constraints are however not as evident as they are in case of directors of large corporations. The typical pension COB is selected from among the members of the board of trustees, which primarily consists of beneficiaries of the pension plan, which are often rank and file employees of the pension sponsor (Pennacchi and Rastad (2011)). In our sample, 72 of 99 female COBs (73 percent) are pension beneficiaries. The trade-off between family and career is likely small for these women. We thus expect that the selection bias discussed in Adams and Funk (2012) is not of first order relevance in our analysis.

The second reservation to the notion that women are generally more risk averse than men is formulated by Jörg (2005). In a large survey of over 2,316 households, she observes that women are only more risk averse than men in investment type decisions but appear relatively more risk loving in insurance situations. Decisions on pension risk entail both an investment and an insurance component. From an investment perspective, higher pension risk is associated with a higher expected return on pension assets. From an insurance perspective, reducing pension risk without simultaneously reducing pension benefits requires that the lower expected return on pension assets is compensated by higher pension contributions (Brown and Wilcox (2009)). These contributions represent an insurance premium. Which perspective on pension risk dominates is ultimately an empirical question. Empirical evidence on asset allocation decisions in personal retirement accounts indicates that women prefer lower levels of pension risk than men (Sundén and Surette (1998) and Agnew, Balduzzi, and Sundén (2003)). Based on these considerations, we predict that pension risk is lower if the COB is female.

3.3. Annuitants

Bulow and Scholes (1983), Bodie (1990), and Carroll and Niehaus (1998) show that beneficiaries of overfunded plans tend to bargain for higher benefit levels. Since the probability of future funding surpluses increases with the mismatch between pension assets and liabilities, Pennacchi and Rastad (2011) argue that beneficiary trustees have a preference for higher pension risk. However, this is only strictly the case for beneficiaries that are already retired (annuitants). Beneficiaries that are still working for the pension sponsor (actives) face the downside of higher pension risk as well. The reason is that an increase in pension risk also increases the probability of future funding shortfalls. Highly underfunded plans are allowed to reduce the level of future benefits (Monahan (2010)). Moreover, the sponsors of these plans, which are the employers of active beneficiaries, might have to cut wages and discharge employees because the deficit reduction contributions drain their financial resources.

In contrast, the pension risk preferences of annuitants are not ambiguous. Benefits that are already earned (vested) are downside protected in most U.S. states (Monahan (2010)).[20] Furthermore, annuitants do not need to worry about their salaries or their jobs as they already retired. Since speculating for higher benefits comes at no personal costs for retirees, we

[20] An exception are Texas and Indiana, where past pension benefit levels are not legally protected (Monahan, 201 0).

expect annuitants to have a preference for high pension risk. We thus predict that pension risk is higher if the COB is an annuitant.

3.4. Politicians

According to the Governmental Accounting Standards Board Statement No. 25 (GASB 25), public pension plans are allowed to discount pension benefits at the expected rate of return on pension assets. Since this rate reflects the risk of pension assets, higher pension risk reduces the actuarial value of pension liabilities and improves the pension funding status (Mohan and Zhang (2014)). Moreover, since the annual required contribution (ARC) of sponsors of underfunded plans depends on the extent of the underfunding (GASB 25), higher pension risk also enables a reduction in the ARC. Brown and Wilcox (2009) point out that the current funding deficits of public pension plans are partially explained by the lack of contributions due to artificially high discount rates.

Politicians prefer to avoid unpopular decisions in the short-term, such as raising taxes or cutting social benefits (Giertz and Papke (2007)). Since riskier pension assets help justify a higher discount rate, which allows lower pension contributions, politicians have a personal incentive to increase the share of risky assets in the pension asset allocation, which increases the overall risk of the pension plan. When low contributions and high risk eventually materialize in underfunded pension liabilities, the politician who is responsible for the decision will likely be gone from office (Giertz and Papke (2007)). We therefore predict that pension risk is higher if the COB is an ex officio trustee.

4. Data and variable construction

4.1. Sample selection

Our main data source is the Public Plans Database (PPD) of the Center for Retirement Research at Boston College (2015).[21] The PPD contains data from Comprehensive Annual Financial Reports (CAFRs) of 150 state and local defined benefit pension plans. It covers 90 percent of the pension members and assets of U.S. public DB plans for the years from 2001 to 2013. We aggregate all plans where asset-liability risk decisions are made by the same board of trustees. This is to avoid double counting certain COB changes that affect several pension plans at the same time. With the sole exception of North Dakota, all plans that share the same (investment) board of trustees have also the same asset allocation.[22] We then match the aggregated PPD data with hand collected information on the COB from more than 2,000 public information sources, including CAFRs, CVs, company webpages, newspaper articles, obituaries, public records databases (*intelius.com* and *dobsearch.com*), and social networking services (*linkedin.com* and *facebook.com*). Furthermore, we personally contacted 53 pension boards where we could not gather all the information from public sources. In total, our data cover 1,503 observations among 116 pension boards of trustees.

We limit our analysis to observations where complete accounting data are reported (asset value, asset allocation, liability value, number of beneficiaries, and number of annuitants). This reduces the number of observations by 112. It has however no effect on the number of boards in our sample. We further exclude pension plans where the board of trustees does not have full discretion over asset allocation decisions but receives risk targets by a separate pension administration board or delegates asset allocation decisions to a separate

[21] The data are publicly available at: http://publicplansdata.org/public-plans-database/download-full-data-set/

[22] In North Dakota, we only use the data for the largest system of the state, the North Dakota Public Employees Retirement System.

investment administration council.[23] For these plans, it is unclear which authority is effectively responsible for asset-liability matching decisions. We do however not exclude plans where the board of trustees is complemented by a separate pension administration board if the administration board does not set explicit targets for the risk of plan assets.[24] This leaves us with an initial sample of 1,310 observations among 110 pension boards (142 pension plans). We have complete COB information in 88 percent of these observations (1,159 board-years).

4.2. Measuring pension risk

Our estimation of public pension risk closely follows the method suggested by Pennacchi and Rastad (2011), which defines the risk of plan i in year t as the volatility of the return difference between pension assets and pension liabilities (tracking error).

$$\text{Pension risk}_{i,t} = 100 \times \sqrt{\sigma^2_{A_{i,t}} + \sigma^2_{L_{i,t}} - 2\rho_{AL_{i,t}}\sigma_{A_{i,t}}\sigma_{L_{i,t}}}\,, \tag{1}$$

where, $\sigma^2_{A_{i,t}}$ is the variance of the pension asset returns of plan i in year t, $\sigma^2_{L_{i,t}}$ is the variance of the relative change in the value of pension liabilities, and $\rho_{AL_{i,t}}$ is the annual correlation between pension asset and pension liability returns. We multiply by 100 to obtain an expression in percentage points.

[23] This includes the Alaska Public Employees’ Retirement System and the Alaska Teachers’ Retirement System for the years 2001 to 2005, the North Dakota Public Employees Retirement System, the Oregon Public Employees Retirement System, the Employees' Retirement System of Rhode Island, the Rhode Island Municipal Employees’ Retirement System, the South Carolina Public Employee Benefit Authority, the South Dakota Retirement System, and the Employees Retirement System of Texas.

[24] This includes the State Employees’ Retirement System of Illinois, the Iowa Public Employees Retirement System, the Massachusetts State Employee Retirement System, the Massachusetts Teachers' Retirement System, the Minnesota State Retirement System, the Montana Public Employee Retirement Administration, the Montana Teachers' Retirement System, Nebraska Public Employees Retirement Systems, the North Dakota Public Employees Retirement System, the West Virginia Public Employees Retirement System, the West Virginia teachers’ Retirement System, the Wisconsin Retirement System, and the Los Angeles County Employees Retirement Association.

The variance of pension assets is defined by

$$\sigma^2_{A_{i,t}} = \boldsymbol{\omega}'_{\mathbf{i,t}}\boldsymbol{\Omega}\boldsymbol{\omega}_{\mathbf{i,t}} , \quad (2)$$

where $\boldsymbol{\Omega}$ is the variance-covariance matrix of the asset class returns and $\boldsymbol{\omega}_{\mathbf{i,t}}$ is the asset allocation vector of the pension plan i in year t. The PPD data provide us with asset allocation information along the categories equities, bonds, real estate, short-term securities, and alternative assets. Following Pennacchi and Rastad (2011), we estimate the variance-covariance matrix based on monthly time series of asset class returns over our entire observation period (2001 to 2013). Equity returns are total returns of a composite of U.S. and international equities. Two thirds of this composite consist of the *Vanguard Total Stock Market Index Fund* (Institutional share class). The other third consists of the *Vanguard Total International Stock Index Fund* (Investor share class). This weighting corresponds to the average international diversification of equity securities of pension plans that separately report their allocations to U.S. and international equities. These plans are approximately half of the plans in our sample. Our bond composite consists of 90 percent the *Merrill Lynch U.S. Corporate & Government Master Index* and 10 percent the *Merrill Lynch Global Broad Market ex US Dollar Index*. As in the estimation of equity returns, this weighting reflects the average international diversification of fixed income investments of the plans in our sample. Real estate returns are the returns of the *Vanguard REIT Index Fund* (Investor share class), and short term security returns are the returns of the *Vanguard Short-Term Investment-Grade Fund* (Institutional share class). Finally, the return series for alternative investments reflects the equally weighted average returns of the *Thomson Reuters Core Commodity CRB Total Return Index*, the *Thomson Reuters Private Equity Buyout Research Index*, and the *Thomson Reuters Venture Capital Research Index*. Mohan and Zhang (2014) show that the risk of alternative investments of public pension plans is best approximated by the equally weighted average risk of commodities, private equity, and venture capital. Some plans report a residual fraction of assets that is not assigned to any of the above asset classes. Following Pennacchi and Rastad (2011), we ignore these investments and proportionally increase the weights of the other asset classes. Since unassigned assets only account for one percent of the average asset allocation in our sample, the measurement error from this approximation is small. Our investment return estimates are indeed highly correlated with the actual investment returns reported by the pension plans. The correlation statistic is 0.92 for the entire sample (1,159 board-years), 0.94 for pension plans that allocate more than 10 percent of their assets to alternative assets (395 board-years), and 0.88 for plans that report a positive allocation to other assets (247 board-years).

The economic value of pension liabilities is the present value of expected future benefit payments. This value is sensitive to both interest rate changes and changes in the growth rate of government wages (Pennacchi and Rastad (2011)). Interest rates affect the rate at which pension benefits should be discounted. Changes in the wage growth affect future benefit payments because pension benefits are defined as a percentage of an employee's last salary before retirement. Pennacchi and Rastad (2011) define the economic risk (variance) of pension liabilities by

$$\sigma^2_{L_{i,t}} = \sigma^2_{LW_{i,t}} + \sigma^2_{LB_{i,t}} + 2\rho_{WB}\sigma_{LW_{i,t}}\sigma_{LB_{i,t}}\ , \tag{3}$$

where $\sigma_{LW_{i,t}}$ measures the volatility of wage increases at plan i in year t, $\sigma_{LB_{i,t}}$ measures bond return volatility (interest rate risk), and ρ_{WB} is the correlation between wage increases and bond holding period returns over the entire observation period (2001-2013). Pennacchi and Rastad (2010) show that both wage and interest rate risk of pension liabilities depend on the ratio of active (working) beneficiaries to total pension beneficiaries. Obligations to annuitants do not bear any wage risk because retirees cannot receive wage increases. Moreover, the interest rate risk (duration) of annuitant liabilities is lower than the interest rate risk of obligations to actives because annuitants have a shorter life expectancy than actives. Pension liability risk is therefore increasing in the ratio of active to total beneficiaries. In line with Pennacchi and Rastad (2010), we define the wage risk component of pension liability risk by

$$\sigma_{LW_{i,t}} = \sigma_W \left[\frac{N_{E_{i,t}}}{N_{E_{i,t}} + N_{A_{i,t}}}\right] \tag{4}$$

and the interest rate risk component by

$$\sigma_{LB_{i,t}} = \sigma_B \left[0.4 + 0.558\left(\frac{N_{E_{i,t}}}{N_{E_{i,t}} + N_{A_{i,t}}}\right) + 0.0425\left(\frac{N_{E_{i,t}}}{N_{E_{i,t}} + N_{A_{i,t}}}\right)^2\right], \tag{5}$$

where $N_{E_{i,t}}$ is the number of active employees of plan i in year t, $N_{A_{i,t}}$ is the number of annuitants, σ_W is the annualized volatility of quarterly changes in the Bureau of Labor Statistics (2015) seasonally adjusted Employment Cost Index for State and Local Government Workers,[25] and σ_B is the annualized volatility of the monthly holding period return of a 15-year zero government bond. Both σ_W and σ_B are scalars that we estimate based on return information over our entire observation period. A detailed derivation of the nominal factors in equation (5) is provided in Pennacchi and Rastad (2010).

According to Pennacchi and Rastad (2011), liabilities of public pension plans reflect nominal interest rate risk when the pension plan does not provide Cost of Living Adjustments

[25] The data are available at http://www.bls.gov/data/#wages

(COLAs) but reflect real interest rate risk when the plan provides COLAs. Therefore, we estimate two different versions of the interest rate risk of pension liabilities. The first is based on the nominal holding period return of a 15-year U.S. government bond from Thomson Reuters Eikon (USGOV15YZ). The second is based on the holding period return of a 15-year zero TIPS from Gürkaynak, Sack, and Wright (2008).[26] We thus also define two different estimates of public pension risk. The first (Nominal PR) reflects nominal interest rate risk in case of all pension plans. Our second pension risk measure (COLA adj PR) measures real interest rate risk if the pension plan reports that it adjusts pension benefits for changes in the Consumer Price Index (CPI), and nominal interest rate risk otherwise.

Table 2 displays the correlations and standard deviations of the return series we use to estimate public pension risk. The reported correlation statistics are consistent with the values from the return series used by Pennacchi and Rastad (2011). On the one hand, we find that equities, real estate, and alternative investments are highly correlated. On the other hand, the correlation of these asset classes with bond returns and wage growth is weak. Most intuitively, the correlation between bond portfolio returns and 15-year zero bond returns is high. The only asset class that is substantially correlated with wage increases is short term securities. During the years 2001 to 2013, 15-year zero bonds had a volatility of 0.151, while the standard deviation of equity securities was 0.109. This highlights the relative importance of interest rate risk during our observation period.

[insert Table 2 here]

[26] The data are available at http://www.federalreserve.gov/econresdata/feds/2008/index.htm

5. Empirical strategy

5.1. Regression model

We assess the impact of COB risk preferences on pension risk in a difference-in-difference analysis. The DID estimator compares the change in the risk of pension plans in the current and subsequent fiscal years in which the COB risk preferences have changed (treatment group) with the contemporaneous change in the risk of pension plans for which the risk preferences of the COB have not changed (control group). This enables to control for omitted variables that similarly affect the risk taking of the treated and the untreated group (Berger, Kick, and Schaeck (2014)). Moreover, it enables to control for unobserved pension plan characteristics that remain constant over time. We define our main regression equation in first difference (FD) form by

$$\Delta\,\text{Pension risk}_{\text{i,t}} = \boldsymbol{\alpha}_{\text{t}} + \boldsymbol{\beta}' \cdot \Delta\,\textbf{COB RP}_{\text{i,t}} + \boldsymbol{\Gamma}' \cdot \Delta\,\mathbf{X}_{\mathbf{COB}_{\text{i,t}}} + \boldsymbol{\Upsilon}' \cdot \Delta\,\mathbf{X}_{\mathbf{Pension}_{\text{i,t}}} + \epsilon_{\text{i,t}}\,, \quad (6)$$

where $\Delta\,\text{Pension risk}_{\text{i,t}}$ is the change in the risk of the plans managed by board i from year $\text{t} - 1$ to year t. $\Delta\,\textbf{COB RP}_{\text{i,t}}$ represents a matrix of year-to-year changes in our set of COB risk preference measures, $\Delta\,\mathbf{X}_{\mathbf{COB}}$ is a set of variables that controls for changes in other COB characteristics, and $\Delta\,\mathbf{X}_{\mathbf{Pension}}$ consists of controls for changes in pension plan characteristics. $\boldsymbol{\beta}'$, $\boldsymbol{\Gamma}'$, and $\boldsymbol{\Upsilon}'$ are vectors of coefficients, $\boldsymbol{\alpha}_{\text{t}}$ represents year fixed effects, and $\epsilon_{\text{i,t}}$ is a stochastic error term.

Following our considerations in Section 3, the COB risk preference measures include changes in COB age ($\Delta\,\text{COB age}$) and gender ($\Delta\,\text{COB female}$), changes in the COB's status as an annuitant of the plan ($\Delta\,\text{COB annuitant}$), and changes in whether he is an ex officio trustee ($\Delta\,\text{COB ex officio}$) or not. We define $\Delta\,\text{COB age}$ as the difference between the age of the new COB and the age of his predecessor in years. $\Delta\,\text{COB female}$, $\Delta\,\text{COB annuitant}$, and $\Delta\,\text{COB ex officio}$ represent changes in the respective COB characteristics. All of these variables take a value of one if the new COB exhibits the characteristic while the predecessor did not, a value of minus one if the new COB does not exhibit the characteristic while the predecessor did, and a value of zero otherwise. All COB variables take a value of zero if the COB does not change. Throughout our analysis, we allocate COB changes to the first fiscal year when the new COB presides the pension board for at least half of the fiscal year (six months).

Based on our empirical predictions from Section 3, we expect a negative coefficient on $\Delta\,\text{COB age}$ and $\Delta\,\text{COB female}$ and a positive coefficient on $\Delta\,\text{COB annuitant}$ and $\Delta\,\text{COB ex officio}$, respectively.

5.2. Control variables

The first variable in our set of COB controls ($\Delta\,\mathbf{X_{COB}}$) is a binary variable (New COB) that identifies board-years where the COB has changed in the current fiscal year. We control for the COB change per se to distinguish between changes in pension risk that stem from the COB change alone and effects that reflect changes in COB risk preferences (treatment effect). Furthermore, we control for the change in a binary variable that indicates whether the COB is a beneficiary of the plan ($\Delta\,\text{COB beneficiary}$). This is necessary to distinguish between a general preference of pension beneficiaries for higher pension risk, as suggested by Pennacchi and Rastad (2011), and our prediction that only annuitants prefer higher levels of pension risk. Finally, our set of pension plan controls includes the change in a binary variable that indicates whether the COB is financially literate ($\Delta\,\text{Financial literacy}$). Existing research indicates that financial literacy matters in portfolio decisions. For instance, poorly educated households invest less in equity securities and avoid financial strategies for which they feel unqualified

(Campbell (2006)). We classify a COB as financially literate when he holds a Master's degree or a PhD in economics, finance, or business administration, and/or is a Chartered Financial Analyst (CFA), a Certified Financial Planner (CFP), a Certified Public Accountant (CPA), or a Certified Public Financial Advisor (CPFA).[27]

The first two variables in our set of pension plan controls ($\Delta\,\mathbf{X_{Pension}}$) are the change in the previous year funding ratio ($\Delta\,\text{Funding ratio}_{-1}$) and the change in the previous year reported investment return ($\Delta\,\text{Return}_{-1}$). Rauh (2009) and Mohan and Zhang (2014) present evidence that pension risk is correlated with the previous period funding ratio and the previous period investment return because of either risk management or risk transfer incentives of pension managers. A risk management view implies that pension risk is positively affected by both previous year funding ratio and investment return because an increase in these variables increases the risk carrying capacity of the pension plan (Rauh (2009)). A risk transfer view on the other hand implies that pension risk is negatively affected by funding status and past return, respectively. This is because pension managers try to improve the funding ratio in the short run by raising the expected return on pension assets and shifting the risk to future tax payers (Mohan and Zhang (2014)). We estimate the funding ratio as the market value of assets divided by the value of pension liabilities. In line with Pennacchi and Rastad (2011), we measure pension liabilities by their actuarial value under GASB standards because pension plans do not report the fair, economic value of their liabilities.

We also control for changes in the natural logarithm of pension assets ($\Delta\,\text{Ln size}$). According to Mohan and Zhang (2014), larger pension plans enjoy economies of scale for transaction fees, which increases their incentive to invest in equity and alternative assets, which both lead to higher pension risk. Finally, we include year fixed effects to control for

[27] For 21 COBs in our main sample, we could not find any information about their education. We assume that those COBs do not fulfill our criteria of financial literacy. When we alternatively exclude these observations, our results remain virtually identical.

macro-economic effects that affect all pension plans in a similar way. Table 9 of the appendix provides detailed definitions of all the variables we use in this paper.

6. Empirical results

6.1. Descriptive statistics

Table 3 shows the summary statistics for our sample of public pension plans from 2001 to 2013. All continuous variables are winsorized at the 1st and the 99th percent level of their pooled distribution to eliminate outliers. Panel A displays descriptive statistics of selected pension plan and COB characteristics, Panel B does so for the first differences of our continuous pension variables, Panel C shows descriptive statistics of the changes in COB characteristics, and Panel D reports the pairwise Pearson correlation coefficients between changes in COB characteristics.

The average nominal pension risk (Nominal PR) of the plans in our sample is 14.3 percent; the average COLA adjusted pension risk (COLA adj PR) is 13.4. Pension risk therefore exceeds the volatility of a broadly diversified equity portfolio (10.9 percent) and corresponds to the volatility of alternative assets (14.4 percent) in Table 2. Table 3 also shows that, while the majority of pension assets are invested in equities (56 percent) and alternative assets (8 percent), only 28 percent of pension assets are allocated to bonds. This predominant allocation of pension assets to equities and alternatives explains the distinct mismatch between pension assets and liabilities. As we show in Table 2, both equity returns and returns on alternative assets are only weakly correlated with wage changes and long-term bond returns.

Consistent with previous studies of public pension plans, including Pennacchi and Rastad (2011) and Mohan and Zhang (2014), we observe a pronounced underfunding of pension obligations. On average, only 77 percent of the actuarial pension liabilities are funded by assets. According to Novy-Marx and Rauh (2011), the funding ratio would be even lower if pension liabilities were measured by their fair value.

The typical (median) COB age is 56. The probability that the chairperson is a woman, an annuitant, an ex officio trustee, or a beneficiary, is 22, 12, 22, and 71 percent, respectively. The probability that he is financially literate is 21 percent. The summary statistics of the 343 COB changes (Panel C) indicate that the average new COB is 2 years younger than his predecessor. In 166 cases, the new COB is younger than the incumbent chairman. In 143 cases, it is the other way around. The distributions of the first differences in the remaining measures of COB risk preferences indicate that these changes are fairly symmetrical. 52 out of 100 changes in COB gender are from man to woman, 38 out of 81 changes of the COB's annuitant status are from non-annuitant to annuitant, and 12 out of 28 changes in the chairman's ex officio status are from non-ex officio to ex officio. The distributions of the COB control variables are symmetrical as well. 47 out of 98 changes in the beneficiary status are from a non-beneficiary to a beneficiary COB. In 56 out of 103 changes in financial literacy, the new COB is financially educated while the replaced chairman was not.

Panel D shows that changes in COB characteristics, except for changes in COB age, are not significantly correlated with the decision to replace the COB ($\mathrm{New\,COB}$). However, changes in COB characteristics are correlated with each other. Older COBs are more likely annuitants and beneficiaries but less likely financially educated. Female COBs are less likely financially literate than men, which is consistent with the findings in Dwyer, Gilkeson, and List (2002) and Jörg (2005). Female chairpersons are however more likely annuitants, who are in turn less likely ex officio trustees, more likely beneficiaries, and less likely financially educated. Finally, ex officio trustees are more likely financially literate.

[insert Table 3 here]

6.2. *Main results*

Table 4 studies the relation between COB risk preferences and the risk of public pension plans. We run FD regressions based on equation (6). Since we make directional predictions on the relation between pension risk and our measures of COB risk preferences, we determine the statistical significance of the respective coefficients based on one-tailed tests. In contrast, the statistical significance of the coefficients on the control variables is determined with a two-tailed test. T-statistics are provided in parentheses. We use robust standard errors clustered at the pension board level.

Column (1) displays the results of an FD regression of nominal pension risk on COB risk preferences and controls. Column (2) shows the estimates of a similar regression of COLA adjusted pension risk. In both regressions, the coefficients on $\Delta\,\mathrm{COB}$ age and $\Delta\,\mathrm{COB}$ female are negative, while the coefficients on $\Delta\,\mathrm{COB}\,\mathrm{annuitant}$ and $\Delta\,\mathrm{COB}$ ex officio are positive. This is in line with our predictions that pension risk is decreasing in COB age, lower if the COB is a woman, higher if the COB is an annuitant, and higher if the COB is an ex officio trustee. With the sole exception of the coefficient on $\Delta\,\mathrm{COB}$ ex officio in Column (2), all coefficients are statistically significant.

The change of the COB per se $(\mathrm{New\,COB})$ has no impact on public pension risk. The COB's beneficiary status has no significant effect on pension risk as well, which supports our expectation that the pension risk preferences of active beneficiaries are ambiguous. The financial literacy of the COB has no significant impact on pension risk either.[28] We explain this result by the fact that pension COBs have easy access to advice from investment professionals. Most public pension plans employ a professional investment consultant (Goyal and Wahal (2008)). Gaudecker (2015) shows that financial literacy only matters in investment decisions by individuals who do not seek outside advice.

[28] In a not tabulated regression, where we extend our definition of financial literacy to bachelor degrees in finance, economics, and business administration, we receive a similar result.

Furthermore, we find that the change in pension risk is significantly positively related to both the lagged change in pension funding ratio and the lagged change in investment return. This supports the risk-management view of Rauh (2009) that pension risk is positively related to the risk carrying capacity of the plans. It also supports Weller and Wenger (2009) who show that managers of underfunded pension plans do not systematically chase returns. Moreover, public pension risk is significantly positively related to pension plan size. This is in line with Mohan and Zhang (2014), who argue that larger pension plans have higher pension risk as they benefit from economies of scale for transaction fees, which increases their incentive to invest in equities and alternative assets.

The impact of COB risk preferences on pension risk is also of economic importance. Based on the coefficients in Column (1), we find that a one standard deviation increase in COB age (13 years) results in a reduction of pension risk by 1.42 percent, which equals 0.12 standard deviations of pension risk. A change in COB gender results in a 0.08 standard deviation change of pension risk, a change in the COB's annuitant status explains 0.15 standard deviations of pension risk, and a change in the COB's ex officio status leads to a change in pension risk by 0.16 standard deviations. In comparison, a one standard deviation change in funding status explains a 0.20 standard deviation change in pension risk. The replacement of an old, female, non-ex officio COB by a young, male, ex officio trustee causes an increase in pension risk by 4.31 percent (0.36 standard deviations of pension risk). To put this in perspective, in case of California, this would result in an increase of the one year 97.5 percent value at risk of public pension plans by USD 8.5 billion or 7.2 percent of the annual Californian tax revenue.[29]

[insert Table 4 here]

[29] In 2009, the State of California faced a fair value of pension liabilities of USD 700 billion and achieved annual tax revenues of USD 117 billion (Novy-Marx and Rauh, 2011).

6.3. *Representativeness*

The median COB turnover in our sample is 3 years. However, there are considerable differences between boards. While in some plans, the chairman changes every year, he only changes once during our observation period in other plans. Plans with high COB turnover hence appear more often in the treatment group than plans where the COB turnover is low. If the treatment group is repeatedly composed of the same, small number of pension plans, our findings could thus be non-representative for U.S. public pension plans in general.

We address this concern in Table 5 by repeating the regressions in the previous section for a restricted sample of pension board-years where the replaced COB served a minimum term of two years (Columns (1) and (2)). This insures that our analysis is not dominated by plans with extremely high COB turnover. Consistent with the results in Table 4, we find that the coefficients on the COB risk preference measures support our predictions. The estimates on COB age and COB gender remain statistically significant. The coefficient on the COB's ex officio status, which is not statistically significant in the COLA adjusted pension risk regressions in Table 4, is now significantly larger than zero for both pension risk measures. The coefficient on changes in the COB's annuitant status is however not statistically significant anymore. In Columns (3) and (4), we further restrict our sample to board-years where the replaced COB served a minimum term of three years. With the sole exception of the coefficient on the COB's ex officio status, which loses its statistical significance in the nominal pension risk regression in Column (3), the magnitude and statistical significance of the estimates remain virtually unchanged.

Based on these results, we conclude that our findings in the previous section are representative for public pension plans at large.

[insert Table 5 here]

6.4. *Causality*

The main concern regarding our interpretation that personal COB risk preferences affect pension risk is that the relation between pension risk and COB risk preferences could reflect the risk preferences of the board of trustees, which usually appoints the COB and formally decides on the risk of the pension plan. Potentially, the board simultaneously changes the risk of the plan and appoints a new COB whose preferences match the new risk policy. Table 6 addresses this concern.

In Columns (1) and (2), we extend our set of controls by changes in the discount rate of pension liabilities ($\Delta\,\text{Discount rate}$), changes in the amortization period for an underfunding of pension liabilities ($\Delta\,\text{Amortization period}$), and changes in the smoothing period for the recognition of an investment loss ($\Delta\,\text{Smoothing period}$). Mohan and Zhang (2014) show that pension risk is increasing in those policy variables. Since changes in risk policies require board approval, these policy variables reflect the risk preferences of the board of trustees. If these preferences were driving the results, their inclusion should weaken the relation between pension risk and our measures of COB risk preferences. Contrary to that, we find that the coefficients are virtually identical to the estimates in Table 4. None of the coefficients on the policy variables is statistically significant. However, both the estimate on changes in the discount rate and the coefficient on changes in the amortization period are positive, which supports the findings in Mohan and Zhang (2014).

In Columns (3) and (4), we also control for changes in the composition of the board of trustees. We conduct this analysis in a subsample of state-wide pension plans for which we have information about changes in the fraction of board seats held by women ($\Delta\,\text{Female seats}$), annuitants ($\Delta\,\text{Annuitant seats}$), ex officio trustees ($\Delta\,\text{Ex officio seats}$), and beneficiaries ($\Delta\,\text{Beneficiary seats}$), respectively.[30] We cannot control for changes in

[30] We like to thank Caroline Ruprecht from the University of Bern for her help in collecting these data.

average board member age because we often lack this information. Consistent with our argumentation about the preferences of individual COBs in Section 3, we expect that boards with a higher percentage of female board members are more risk averse, while boards with a higher percentage of annuitant and/or ex officio trustees prefer higher levels of pension risk. We control for the fraction of beneficiary board seats because Pennacchi and Rastad (2011) find that pension risk increases with the ratio of beneficiary trustees to total board members. In this subsample of state-wide plans, we can also control for changes in the S&P state credit rating (Δ S&P rating), which is a measure of fiscal constraints of the sponsoring state government (Mohan and Zhang (2014)). According to Mohan and Zhang (2014), government sponsors that experience tight fiscal constraints have an incentive to increase the risk of pension assets. Following GASB standards, increasing the risk of pension assets allows sponsors to justify a higher liability discount rate. Higher discounting reduces the actuarial value of pension liabilities, improves the funding status, and reduces the annual required contribution. Higher pension risk hence allows government sponsors to substitute general obligation debt for pension debt. We measure Δ S&P rating by the number of notches the current year rating is higher (positive values) or lower (negative values) than the previous year rating.

We find that sign, magnitude, and statistical significance of the coefficients on our measures of COB risk preferences are almost identical to the estimates in Columns (1) and (2). None of the board composition measures is significantly related to nominal pension risk. In the regression of COLA adjusted pension risk, we find that the fraction of annuitant trustees is significantly negatively related to pension risk. We interpret this result as the reflection of a higher risk aversion of older pension boards. Consistent with the explanation that fiscally constrained pension sponsors substitute general obligation debt for pension debt, we find a negative relation between S&P rating and pension risk. However, the coefficient on rating changes is only statistically significant in Column (4).

[insert Table 6 here]

To further mitigate the concern that the relation between pension risk and COB preferences reflects the risk preferences of the board of trustees, we consider a subsample of pension plans with ex officio COB. Ex officio chairmen hold their position because they are elected to a specific public office (e.g., governor, mayor, or superintendent of the school district). The pension board has no say in their appointment. Therefore, changes in ex officio COB risk preference cannot be a reflection of changes in the risk preference of the board. Since ex officio chairpersons are never annuitants but always ex officio trustees, our set of risk preference measures is reduced to COB age and gender. Similarly, we do not control for the COB's beneficiary status because in this subsample there is no within-board variation in this variable. Columns (1) and (2) of Table 7 report the estimates from regressions of nominal pension risk and COLA adjusted pension risk, respectively. In line with our previous results, we find that pension risk decreases with COB age and is lower if the COB is a woman. Both effects are statistically significant above the 5 percent level. The magnitude of the coefficients suggests that the reflection of COB preferences in pension risk is more pronounced for plans where the COB holds his position ex officio than for the average plan in our sample. A one standard deviation increase in COB age (8.5 years) causes a reduction in pension risk by 3.5 percent (0.30 standard deviations) and the replacement of a male COB by a female chairperson reduces pension risk by 6.8 percent (0.58 standard deviations). These values are more than twice the magnitude of the economic effects of COB age and gender in our main analysis (Table 4). We explain these differences by the fact that in one third of the observations in the ex officio sample, the COB serves as a sole trustee. The personal influence of sole trustees is likely higher than the influence of chairpersons of large pension boards because sole trustees do not need their fellow board members to support their decisions. In Columns (3) and (4), we test this explanation by repeating our regressions for the subsample of plans where the COB is the sole trustee. In this sample, we find an even stronger economic effect of COB risk preferences, which supports our contention that the magnitude of the relation between pension risk and COB risk preferences depends on the personal influence of the COB. A one standard deviation increase in COB age (8.3 years) reduces pension risk by 0.66 standard deviations and a switch from male to female COB reduces the pension risk by 0.87 standard deviations.

[insert Table 7 here]

The results from Tables 7 and 8 are inconsistent with the hypothesis that COBs are selected to match the board of trustees' risk preferences. Moreover, our findings suggest that individual COB risk preferences have a more pronounced effect on public pension risk when the COB has a higher influence on asset liability matching decisions. The next section elaborates further on how our results are affected by the governance structure of the pension plan.

6.5. Impact of pension governance

Cronqvist, Makhija, and Yonker (2012) show that CEOs primarily imprint their personal preferences on the managed firm when corporate governance is weak. In this section, we inquire into whether good pension governance keeps COBs from imprinting their personal preferences on the plans they are responsible for as well. Table 8 compares the impact of

COB risk preferences on pension risk between plans with a single board of trustees and plans with an (investment) board of trustees that is complemented by a separate pension administration board. In the first governance model, all decision power lies with the same board. In the second governance model, asset-liability matching decisions lie with the investment board, which, however, has to coordinate its decisions with the pension administration board (Miller and Funston (2014)). The administration board hence serves as an implicit supervisory unit for pension risk decisions of the investment board. Moreover, the administration board is responsible for decisions on pension benefit levels. Plans with two separate boards are also larger and more professionalized (Miller and Funston (2014)). Thus, we consider plans with a separate investment board as better governed. Consistent with Cronqvist, Makhija, and Yonker (2012), we expect the risk of these plans to be less affected by personal preferences of the chairman (of the investment board).

Columns (1) and (2) of Table 8 show the estimates for a sample of pension plans with a single board of trustees. In both regressions, all coefficients on the COB risk preference measures are statistically significant and consistent with our empirical predictions. Columns (3) and (4) show the results for plans with a separate investment board. In the regression of nominal pension risk, none of the COB risk preference measures is significantly related to pension risk. In the regression of COLA adjusted pension risk, the coefficient on the COB's annuitant status is significantly positive and the estimate on the COB's ex officio status is significantly negative. All other coefficients are not significantly different from zero.

Overall, we interpret these results as supporting evidence for the view that pension risk is predominantly affected by COB preferences when pension governance is weaker.

[insert Table 8 here]

7. Conclusion

This paper studies whether the overall risk of public pension plans is affected by the personal risk preferences of the chairman of the board of trustees. In line with Pennacchi and Rastad (2011), we define pension risk by the volatility of the difference between pension asset and pension liability returns. Our empirical analyses of the relation between pension risk and COB risk preferences are based on four empirical predictions that have been suggested by the literature. Pension risk is (1) negatively affected by an increase in COB age, (2) lower if the COB is a woman, (3) higher if the COB is an annuitant, and (4) higher if the COB is a politician (ex officio trustee).

The evidence we present in this paper supports these predictions. Moreover, we cannot find that the relation between pension risk and COB risk preferences is explained by an endogenous appointment of COBs to plans that match their personal preferences. Similarly, our results are not explained by the risk preferences of the board of trustees. The impact of COB preferences on pension risk is economically important. Each of our risk preference measures explains a considerable amount of the variation in pension risk. Consistent with previous results on the impact of individual manager preferences on corporate risk taking, we find that the reflection of COB preferences in pension risk is particularly evident among more weakly governed plans, which are plans not governed by a separate and professionalized investment board.

Our main contribution is the identification of personal COB risk preferences as a predictor of the risk from mismatched pension assets and liabilities. Since current funding problems of state and local government pension plans are mainly the consequence of asset and liability mismatching, it is in the interest of taxpayers to be wary of who is in charge of the asset-liability risk decisions in public pension plans.

Appendix: Tables

Table 1: Probability of COB replacement

This table compares the probability that the COB is replaced after a year of high investment performance with the probability of a COB replacement after a year of low investment performance. An observation is classified as *high performance* when the previous year performance is above the median investment performance of public pension plans in that year, and as *low performance* otherwise. Panel A includes observations from pension plans with old as well as young COBs. Panel B only includes observations where the COB is younger than the median COB (56 years). Panel C only includes observations where the COB is younger than the 25th percentile of the COBs (50 years). The numbers in parentheses indicate t-statistics of a parametric test of mean difference.

	Probability of COB replacement			
	After a year of high performance	After a year of low performance	Difference	
Panel A: All observations	0.327	0.316	0.012	(0.42)
Panel B: COB age < 56 years	0.313	0.308	0.005	(0.13)
Panel C: COB age < 50 years	0.314	0.307	0.008	(0.14)

Table 2: Correlations and standard deviations of asset returns and wage growth

This table shows the correlations statistics and the standard deviations of the return series used to estimate the risk of public pension plans. The data refer to 2001 to 2013. (1) Equity returns are composed of two thirds the returns of the *Vanguard Total Stock Market Index Fund* (Institutional share class) and one third the returns of the *Vanguard Total International Stock Index Fund* (Investor share class). (2) Bond returns are composed of nine tenth the returns of the *Merrill Lynch U.S. Corporate & Government Master Index* and one tenth the returns of the *Merrill Lynch Global Broad Market ex US Dollar Index*. (3) Real estate returns are the returns of the *Vanguard REIT Index Fund* (Investor share class). (4) Short term returns are the returns of the *Vanguard Short-Term Investment-Grade Fund* (Institutional share class). (5) The returns of alternative investments are the equally weighted average returns of the *Thomson Reuters Core Commodity CRB Total Return Index*, the *Thomson Reuters Private Equity Buyout Research Index*, and the *Thomson Reuters Venture Capital Research Index*. (6) The Nominal 15 year zero bond returns are monthly holding period returns extracted from the yields to maturity of a 15 year zero government bond in Thomson Reuters Eikon (USGOV15YZ). (7) The real 15 year zero bond returns are monthly holding period returns extracted from the yield to maturity of the 15 year zero coupon TIPS reported by Gürkaynak, Sack, and Wright (2008). (8) The wage growth is estimated by the Bureau of Labor Statistics quarterly seasonal adjusted Employment Cost Index for State and Local Government Workers.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
(1) Equities	1.000							
(2) Bonds	-0.082	1.000						
(3) Real Estate	0.700	0.111	1.000					
(4) Short term	-0.147	0.055	-0.092	1.000				
(5) Alternatives	0.946	-0.405	0.640	-0.222	1.000			
(6) Nominal 15 year zero bond	-0.266	0.883	-0.084	0.019	0.128	1.000		
(7) Real 15 year zero bond	0.008	0.804	0.138	0.010	-0.172	0.731	1.000	
(8) Wages	-0.292	0.110	-0.171	0.809	-0.272	-0.083	0.023	1.000
Standard deviation	0.109	0.041	0.233	0.005	0.144	0.151	0.107	0.060

Table 3: Descriptive sample statistics

This table shows the descriptive sample statistics of our main variables and further pension plan characteristics. Variable definitions are in Table 9. The data refer to 2001 to 2013. The sample is restricted to pension plans where the board of trustees is fully responsible for asset allocation decisions. Panel A displays the summary statistic of selected pension plan and COB characteristics, Panel B provides the summary statistics of the continuous pension variables, Panel C depicts descriptive statistics of the changes in COB characteristics for board-years where the COB changes, and Panel D shows the pairwise Pearson correlations between changes in COB characteristics. The asterisk denotes statistical significance at the 10% level using a two-tailed test. All continuous variables are winsorized at the 1st and the 99th percentile of their pooled sample distribution.

Panel A: Summary statistics of plan and COB characteristics

Year	Mean	Median	Min	Max	Std.	N
Nominal PR (in %)	14.31	14.30	10.62	18.83	1.69	1,159
COLA adj PR (in %)	13.38	13.28	8.79	18.34	2.08	1,159
Alloc. to equities (in %)	56.02	57.27	24.32	73.27	9.58	1,159
Alloc. to bonds (in %)	28.32	27.73	13.22	53.58	7.51	1,159
Alloc. to real estate (in %)	5.37	5.28	0.00	17.87	4.33	1,159
Alloc. to short term securities (in %)	2.10	1.23	0.00	13.69	2.51	1,159
Alloc. to alternatives (in %)	7.98	5.49	0.00	38.43	8.51	1,159
Funding ratio	0.768	0.768	0.376	1.275	0.184	1,159
Market value of assets (in $m)	22,346	10,401	798	163,438	30,105	1,159
COB age (in years)	57	56	37	82	10	1,159
COB female	0.223	0.000	0.000	1.000	0.416	1,159
COB annuitant	0.119	0.000	0.000	1.000	0.324	1,159
COB ex officio	0.223	0.000	0.000	1.000	0.417	1,159
COB beneficiary	0.714	1.000	0.000	1.000	0.452	1,159
COB financial literacy	0.210	0.000	0.000	1.000	0.407	1,159

Panel B: Summary statistics of continuous pension variables

	Mean	Median	Min	Max	Std.	N
Δ Nominal PR	-0.048	0.208	-4.238	3.513	1.876	1,052
Δ COLA adj PR	-0.074	0.115	-4.206	3.503	1.826	1,052
Δ Funding ratio	-0.018	0.009	-0.324	0.140	0.097	1,052
Δ Return	0.015	0.017	-0.360	0.486	0.170	1,052
Δ Ln size	0.035	0.073	-0.356	0.236	0.126	1,052

Panel C: Summary statistics of changes in COB characteristics provided a change of the COB

	Mean	Std.	N	N (nonzero)	N (positive)	N (negative)
New COB	1.000	0.000	343	343	343	0
Δ COB age (in years)	-1.818	13.077	331	309	143	166
Δ COB female	0.012	0.541	343	100	52	48
Δ COB annuitant	-0.015	0.486	343	81	38	43
Δ COB ex officio	-0.012	0.286	343	28	12	16
Δ COB beneficiary	-0.012	0.535	343	98	47	51
Δ COB financial literacy	0.026	0.548	343	103	56	47

Panel D: Correlations between changes in COB characteristics

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
(1) New COB	1.000						
(2) Δ COB age (in years)	-0.114*	1.000					
(3) Δ COB female	0.018	0.014	1.000				
(4) Δ COB annuitant	-0.025	0.511*	0.067*	1.000			
(5) Δ COB ex officio	-0.034	-0.028	-0.019	-0.169*	1.000		
(6) Δ COB beneficiary	-0.018	0.144*	-0.010	0.292*	0.000	1.000	
(7) Δ COB financial literacy	0.039	-0.107*	-0.089*	-0.120*	0.150*	-0.119*	1.000

Table 4: COB characteristics and pension risk

This table shows the results of our main first difference regressions of pension risk on the risk preferences of the COB and controls. Variable definitions are in Table 9. The data refer to 2001 to 2013. The sample is restricted to pension plans where the board of trustees is fully responsible for asset allocation decisions. Subscripts indicate the number of lagged periods. Asterisks denote statistical significance at the 1% (***), 5% (**), and 10% (*) level using one-tailed tests with respect to the COB risk preference measures and two-tailed tests in case of all other variables. T-statistics are provided in parentheses. Standard errors are clustered at the pension board level. All continuous variables are winsorized at the 1st and the 99th percentile of their pooled sample distribution.

Dependent variable	Δ Nominal PR (in %)	Δ COLA adj PR (in %)
	(1)	(2)
New COB	-0.057	-0.046
	(-0.856)	(-0.614)
Δ COB age	-0.015***	-0.015***
	(-2.915)	(-2.597)
Δ COB female	-0.137*	-0.192**
	(-1.434)	(-1.781)
Δ COB annuitant	0.251**	0.347***
	(1.827)	(2.505)
Δ COB ex officio	0.277**	0.271
	(1.752)	(0.956)
Δ COB beneficiary	-0.008	-0.056
	(-0.078)	(-0.477)
Δ COB financial literacy	-0.157	-0.138
	(-1.395)	(-1.170)
Δ Funding ratio$_{-1}$	3.362***	2.006**
	(3.981)	(2.528)
Δ Return$_{-1}$	1.313***	0.889**
	(3.718)	(2.511)
Δ Ln size	4.206***	4.868***
	(9.992)	(10.662)
Year fixed effects	Yes	Yes
Observations	935	935
COB changes	302	302
R2 adjusted	0.761	0.693

Table 5: Representativeness

This table shows the results of first difference regressions of pension risk on the risk preferences of the COB and controls for a subsample of board-years where the COB does not change or the replaced COB served a minimum term of two and three years, respectively. Variable definitions are in Table 9. The data refer to 2001 to 2013. The sample is restricted to pension plans where the board of trustees is fully responsible for asset allocation decisions. Subscripts indicate the number of lagged periods. Asterisks denote statistical significance at the 1% (***), 5% (**), and 10% (*) level using one-tailed tests with respect to the COB risk preference measures and two-tailed tests in case of all other variables. T-statistics are provided in parentheses. Standard errors are clustered at the pension board level. All continuous variables are winsorized at the 1st and the 99th percentile of their pooled sample distribution.

Sample	In case the COB changes, the predecessor served a minimum term of 2 years		In case the COB changes, the predecessor served a minimum term of 3 years	
Dependent variable	Δ Nominal PR (in %)	Δ COLA adj PR (in %)	Δ Nominal PR (in %)	Δ COLA adj PR (in %)
	(1)	(2)	(3)	(4)
New COB	-0.036	-0.027	-0.093	-0.066
	(-0.418)	(-0.282)	(-0.807)	(-0.530)
Δ COB age	-0.013**	-0.014**	-0.018**	-0.014*
	(-1.979)	(-1.820)	(-2.170)	(-1.382)
Δ COB female	-0.236**	-0.262**	-0.387***	-0.301**
	(-2.203)	(-2.007)	(-3.485)	(-2.293)
Δ COB annuitant	0.106	0.151	0.171	0.080
	(0.601)	(0.827)	(0.667)	(0.313)
Δ COB ex officio	0.320**	0.311**	0.208	0.256*
	(2.103)	(1.691)	(1.262)	(1.490)
Δ COB beneficiary	0.037	-0.038	0.003	-0.077
	(0.226)	(-0.208)	(0.014)	(-0.338)
Δ COB financial literacy	-0.258	-0.254	-0.166	-0.246
	(-1.479)	(-1.492)	(-0.936)	(-1.333)
Δ Funding ratio$_{-1}$	3.515***	2.190**	3.540***	2.323**
	(3.716)	(2.475)	(3.480)	(2.440)
Δ Return$_{-1}$	1.285***	0.819**	1.344***	0.954**
	(3.275)	(2.096)	(3.117)	(2.224)
Δ Ln size	4.318***	5.003***	4.067***	4.783***
	(8.764)	(8.993)	(7.472)	(7.658)
Year fixed effects	Yes	Yes	Yes	Yes
Observations	820	820	748	748
COB changes	187	187	115	115
R2 adjusted	0.753	0.683	0.757	0.685

Table 6: Simultaneity to changes in pension policies and changes in the board composition

This table shows the results of our tests for the endogeneity of our findings to changes in pension policies or changes in the composition of the board of trustees. Variable definitions are in Table 9. The data refer to 2001 to 2013. The initial sample is restricted to pension plans where the board of trustees is fully responsible for asset allocation decisions. Subscripts indicate the number of lagged periods. Asterisks denote statistical significance at the 1% (***), 5% (**), and 10% (*) level using one-tailed tests with respect to the COB risk preference measures and two-tailed tests in case of all other variables. T-statistics are provided in parentheses. Standard errors are clustered at the pension board level. All continuous variables are winsorized at the 1st and the 99th percentile of their pooled sample distribution.

Sample	Entire sample		State pension plans only	
Dependent variable	Δ Nominal PR (in %)	Δ COLA adj PR (in %)	Δ Nominal PR (in %)	Δ COLA adj PR (in %)
	(1)	(2)	(3)	(4)
New COB	-0.061	-0.048	-0.089	-0.050
	(-0.906)	(-0.640)	(-0.990)	(-0.492)
Δ COB age	-0.015***	-0.015***	-0.019**	-0.016**
	(-2.923)	(-2.605)	(-2.208)	(-1.692)
Δ COB female	-0.136*	-0.192**	-0.163*	-0.223**
	(-1.418)	(-1.773)	(-1.337)	(-1.797)
Δ COB annuitant	0.254**	0.351***	0.260*	0.319**
	(1.845)	(2.536)	(1.359)	(1.700)
Δ COB ex officio	0.295**	0.290	0.300*	0.161
	(1.766)	(1.002)	(1.395)	(0.498)
Δ COB beneficiary	-0.005	-0.054	-0.047	-0.078
	(-0.052)	(-0.463)	(-0.365)	(-0.521)
Δ COB financial literacy	-0.160	-0.141	-0.079	-0.075
	(-1.403)	(-1.180)	(-0.579)	(-0.534)
Δ Female seats			-0.117	-0.177
			(-0.236)	(-0.326)
Δ Annuitant seats			-1.052	-2.150*
			(-0.972)	(-1.924)
Δ Beneficiary seats			-2.160	-1.767
			(-1.210)	(-0.967)
Δ Ex officio seats			-1.684	-2.730
			(-0.564)	(-0.850)
Δ Discount rate	25.570	27.443	13.374	13.560
	(0.985)	(1.075)	(0.396)	(0.413)
Δ Amortization period	0.003	0.003	0.008	0.009
	(0.403)	(0.334)	(1.025)	(0.993)
Δ Smoothing period	-0.007	0.021	0.028	0.023
	(-0.169)	(0.491)	(0.639)	(0.523)
Δ Funding $ratio_{-1}$	3.355***	2.020**	3.167***	1.591*
	(3.925)	(2.502)	(3.016)	(1.714)
Δ $Return_{-1}$	1.313***	0.889**	1.422***	0.971**
	(3.674)	(2.473)	(3.231)	(2.209)
Δ Ln size	4.219***	4.898***	4.317***	5.124***
	(9.947)	(10.591)	(7.563)	(8.223)
Δ S&P Rating			-0.017	-0.026**
			(-1.438)	(-2.336)
Year fixed effects	Yes	Yes	Yes	Yes
Observations	935	935	629	629
COB changes	302	302	190	190
R2 adjusted	0.761	0.692	0.773	0.699

Table 7: Ex officio appointment of the COB

This table shows the results of our test for a deliberate selection of COBs by the board of trustees. Variable definitions are in Table 9. The data refer to 2001 to 2013. The initial sample is restricted to pension plans where the board of trustees is fully responsible for asset allocation decisions and the COB is appointed ex officio. Subscripts indicate the number of lagged periods. Asterisks denote statistical significance at the 1% (***), 5% (**), and 10% (*) level using one-tailed tests with respect to the COB risk preference measures and two-tailed tests in case of all other variables. T-statistics are provided in parentheses. Standard errors are clustered at the pension board level. All continuous variables are winsorized at the 1st and the 99th percentile of their pooled sample distribution.

Sample	Plans with ex officio COB		Plans where the COB is the sole trustee	
Dependent variable	Δ Nominal PR (in %)	Δ COLA adj PR (in %)	Δ Nominal PR (in %)	Δ COLA adj PR (in %)
	(1)	(2)	(3)	(4)
New COB	-0.306**	-0.256	-0.290	-0.261
	(-2.048)	(-1.472)	(-0.651)	(-0.486)
Δ COB age	-0.059**	-0.058**	-0.129***	-0.135***
	(-1.858)	(-1.736)	(-4.239)	(-4.188)
Δ COB female	-0.965***	-0.957**	-1.415**	-1.054*
	(-2.451)	(-2.002)	(-2.069)	(-1.619)
Δ COB financial literacy	0.162	0.220	0.434*	0.492
	(0.604)	(0.698)	(1.654)	(1.320)
Δ Funding ratio$_{-1}$	4.067**	2.768**	5.532	2.075
	(2.523)	(2.048)	(1.380)	(0.576)
Δ Return$_{-1}$	1.204**	0.187	1.459	0.501
	(1.994)	(0.274)	(1.046)	(0.388)
Δ Ln size	3.579**	4.949**	6.088***	8.863***
	(2.431)	(2.226)	(3.467)	(4.386)
Year fixed effects	Yes	Yes	Yes	Yes
Observations	176	176	62	62
COB changes	29	29	11	11
R2 adjusted	0.833	0.736	0.737	0.631

Table 8: Impact of pension governance

This table displays the results of first difference regressions of pension risk on measures of COB risk preferences and controls separately for pension plans that are governed by a single board of trustees and plans that are governed by two separate boards – one for investment and one for pension administration decisions. The initial sample is restricted to pension plans where the (investment) board of trustees is fully responsible for asset allocation decisions. The COB variables refer to the chairperson of the board that decides on asset-liability matching. The data refer to 2001 to 2013. Subscripts indicate the number of lagged periods. Asterisks denote statistical significance at the 1% (***), 5% (**), and 10% (*) level using one-tailed tests with respect to the COB risk preference measures and two-tailed tests in case of all other variables. T-statistics are provided in parentheses. Standard errors are clustered at the pension board level. All continuous variables are winsorized at the 1st and the 99th percentile of their pooled sample distribution.

Sample	Single board of trustees		Separate pension administration board	
Dependent variable	Δ Nominal PR (in %)	Δ COLA adj PR (in %)	Δ Nominal PR (in %)	Δ COLA adj PR (in %)
	(1)	(2)	(3)	(4)
New COB	-0.071	-0.049	0.180	0.218
	(-0.981)	(-0.605)	(1.104)	(1.085)
Δ COB age	-0.016***	-0.017***	-0.001	0.005
	(-2.834)	(-2.653)	(-0.097)	(0.405)
Δ COB female	-0.152*	-0.237***	0.136	0.190
	(-1.505)	(-2.344)	(0.499)	(0.533)
Δ COB annuitant	0.274**	0.338**	0.068	0.427*
	(1.839)	(2.278)	(0.251)	(1.473)
Δ COB ex officio	0.343**	0.482**	-0.222	-0.827**
	(1.928)	(1.776)	(-1.125)	(-2.216)
Δ COB beneficiary	-0.014	-0.052	-0.291***	-0.556***
	(-0.131)	(-0.402)	(-2.816)	(-3.991)
Δ COB financial literacy	-0.133	-0.106	-0.315*	-0.527**
	(-1.011)	(-0.807)	(-1.767)	(-2.490)
Δ Funding ratio-1	3.055***	1.675*	3.978**	4.125**
	(3.120)	(1.845)	(2.202)	(2.295)
Δ Return-1	1.497***	1.020**	-0.292	0.331
	(3.627)	(2.492)	(-0.986)	(1.316)
Δ Ln size	4.173***	4.978***	3.959***	3.437***
	(8.969)	(9.935)	(9.181)	(4.466)
Year fixed effects	Yes	Yes	Yes	Yes
Observations	839	839	96	96
COB changes	272	272	30	30
R2 adjusted	0.755	0.692	0.816	0.723

Table 9: Variable definitions

This table summarizes the variable definitions. The first column provides the variable names, the second column displays the definitions, and the third column shows the source of the data.

Variable	Definition	Data source
Δ Nominal PR	Year-to-year change in the annualized volatility of the monthly difference between pension asset returns and pension liability returns (tracking error). The estimation is made according to Pennacchi and Rastad (2011). The expression is in percentage points. The interest rate risk of pension liabilities reflects nominal interest rate risk. A detailed description is provided in Section 4.2.	Center for Retirement Research at Boston College (2015)[31], Thomson Reuters Eikon, Bureau of Labor Statistics[32]
Δ COLA adj PR	Equivalently defined to Δ Nominal PR with the exception that the interest rate risk of pension liabilities reflects real interest rate risk for plans that provide CPI related Cost of Living Adjustments (COLAs) of their pension benefits. A detailed description is provided in Section 4.2.	Center for Retirement Research at Boston College (2015)[31], Thomson Reuters Eikon, Gürkaynak, Sack, and Wright (2008)[33], Bureau of Labor Statistics[32]
New COB	Binary variable that takes 1 in the first fiscal year when the new COB presides the pension board for at least half of the fiscal year (six month), and 0 otherwise.	Hand collected
Δ COB age	Difference in age between the new COB and the predecessor in years. COB age is defined as the difference between the current fiscal year and the year of birth of the COB. The variable is set to 0 if New COB is 0.	Hand collected
Δ COB female	Difference in gender between the new COB and the predecessor. The variable takes a value of 1 if the COB changes from a man to a woman, -1 if the COB changes from a woman to a man, and zero otherwise.	Hand collected
Δ COB annuitant	Difference in the annuitant status between the new COB and the predecessor. The variable takes a value of 1 if the COB changes from a non-annuitant trustee to an annuitant trustee, -1 if the COB changes from an annuitant trustee to a non-annuitant trustee, and 0 otherwise.	Hand collected
Δ COB ex officio	Difference in the ex officio status between the new COB and the predecessor. The variable takes a value of 1 if the COB changes from a non-ex officio trustee to an ex officio trustee, -1 if the COB changes from an ex office trustee to a non-ex officio trustee, and 0 otherwise.	Hand collected
Δ COB beneficiary	Difference in the beneficiary status between the new COB and the predecessor. The variable takes a value of 1 if the COB changes from a non-beneficiary to a beneficiary of the plan, -1 if the COB changes from a beneficiary to a non-beneficiary, and 0 otherwise.	Hand collected
Δ COB financial literacy	Difference in the financial literacy between the new COB and the predecessor. The variable takes a value of 1 if the COB changes from a not financially literate person to a financially literate person, -1 if the COB changes from a financially literate person to a financially not literate person, and 0 otherwise. A COB is considered financially literate when he holds a Master's and/or a PhD degree in economics, finance, or business administration, or when he is a Chartered Financial Analyst (CFA), a Certified Financial Planner (CFP), a Certified Public Accountant, or a Certified Public Financial Advisor (CPFA). For 21 COBs in our sample, we could not find an indication about their education. We assume that those COBs do not fulfill our criteria of financial literacy.	Hand collected
Δ Female seats	Year-to-year change in the ratio of female board members to total board members.	Hand collected
Δ Annuitant seats	Year-to-year change in the ratio of annuitant board members to total board members.	Hand collected
Δ Beneficiary seats	Year-to-year change in the ratio of beneficiary board members to total board members.	Hand collected

(continued on next page)

31 The data are available at: http://publicplansdata.org/public-plans-database/download-full-data-set/
32 The data are available at: http://www.bls.gov/data/#wages
33 The data are available at: http://www.federalreserve.gov/econresdata/feds/2008/index.htm

Table 9 *(continued)*

Variable	Definition	Data source
Δ Ex officio seats	Year-to-year change in the ratio of ex officio board members to total board members.	Hand collected
Δ Funding ratio	Year-to-year change in the funding ratio. Funding ratio is defined as the market value of pension assets divided by the actuarial value of pension liabilities following GASB.	Center for Retirement Research at Boston College (2015)[34]
Δ Return	Year-to-year change in the one year investment return.	Center for Retirement Research at Boston College (2015)[34]
Δ Ln size	Year-to-year change in the natural logarithm of the market value of assets.	Center for Retirement Research at Boston College (2015)[34]
Δ S&P rating	Year-to-year change in the S&P credit rating of the state the pension plan is incorporated in. The change is measured by the number of notches the current year rating is higher (positive values) or lower (negative values) than the previous year rating.	The Pew Charitable Trusts[35] and Office of the Chief Financial Officer of the District of Columbia[36]
Δ Discount rate	Year-to-year change in the expected rate of return on pension plan assets, which is equivalent to the discount rate on pension benefits (GASB 25). If the discount rate is reported missing, we assume no change in the discount rate (24 observations).	Center for Retirement Research at Boston College (2015)[34]
Δ Amortization period	Year-to-year change in the number of years the plan is allowed to take for the amortization of an underfunding of pension liabilities. If the amortization period is missing, we assume it does not change (138 observations).	Center for Retirement Research at Boston College (2015)[34]
Δ Smoothing period	Year-to-year change in the smoothing period (in years) for recognizing pension investment losses. If the smoothing period is missing, we assume it does not change (47 observations).	Center for Retirement Research at Boston College (2015)[34]

[34] The data are available at: http://publicplansdata.org/public-plans-database/download-full-data-set/

[35] The data are available at: http://www.pewtrusts.org/en/research-and-analysis/blogs/stateline/2014/06/09/sp-ratings-2014

[36] The data are available at: http://cfo.dc.gov/sites/default/files/dc/sites/ocfo/publication/attachments/Current%20Historical%20GO-IT%20Credit%20Ratings_073113.pdf

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Oliver Dichter

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