The Review of Financial Studies

Outraged by Compensation: Implications for Public Pension Performance

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Public pension boards fear inciting stakeholder outrage if they compensate internal investment managers with market-level salaries. We derive theoretical implications in an agency-portfolio-choice model motivated by inequality aversion. In a global sample, relaxing the effect of outrage on contracting leads to an average annual incremental value-added of $49 million generated through 11 bps in higher excess returns from risky assets, at the cost of $302,429 in additional compensation. Governance reforms that address outrage by reducing political appointees or requiring independent skills-based boards can increase the annual value-added. These findings are orthogonal to costly political distortions from underfunding and pay-to-play schemes. (*JEL* G11, G23, G3)

Received December 10, 2019; editorial decision June 3, 2021 by Editor Holger Mueller. Authors have furnished an Internet Appendix, which is available on the Oxford University Press Web site next to the link to the final published paper online.

Public pensions and sovereign funds hold $21.5 trillion in assets (Official Monetary and Financial Institutions Forum 2018). When these public funds fail to manage these assets effectively, public sector workers and taxpayers bear the costs to support promised pension payouts to retirees. This paper follows Romano (1993) and a long-standing literature in examining the potential for

We thank Lukasz Pomorski for significant contributions to the ideas developed in this paper. We thank Tetyana Balyuk and Michele Dathan for excellent research assistance. We also thank Katharina Lewellen, Josh Rauh, Enrichetta Ravina, and Scott Weisbenner and participants at the AFA meetings, the NBER Law and Economics meeting, Financial Intermediation Research Society conference, the Berkeley-Haas seminar, the Red Rock Finance Conference, the Western Economics Association Meetings, Ackerman Corporate Governance Conference, and the ASU-Sonoran Finance Conference for helpful comments and suggestions. We thank CEM for providing the data used in this analysis. The views expressed in this paper are solely those of the authors and do not represent the position or policies of the Department of the Treasury. Supplementary data can be found on *The Review of Financial Studies* web site. Send correspondence to Alexander Dyck, adyck@rotman.utoronto.ca.

The Review of Financial Studies 35 (2022) 2928–2980

doi: 10.1093/rfs/hhab109 Advance Access publication October 19, 2021

agency costs to erode performance of public pension funds. We complement the recent literature, which has focused on distortions arising from politicians' extraction of private benefits from pension funds' asset management,[1] by theoretically and empirically focusing on a novel human capital channel.

Specifically, we hypothesize that pension trustees fear the triggering of public outrage if they compensate their investment managers at a market rate level. This tension arises because the market rate pay of investment managers is large relative to pension beneficiaries and local taxpayers, triggering an income inequality aversion. Fears of this outrage cause pension trustees (who have career concerns sensitive to public perception) to hire lower-skilled internal managers and offer suboptimal incentive contracts. This talent hiring friction reduces pension performance, thereby causing future costs to workers and pension retirees. Notably, outrage costs are not distributed equally, but fall harder on *Main Street* communities which already exhibit lower local wages and greater income inequality.

To illustrate the outrage constraint in action, consider the dilemma of the Oregon State Treasurer in his service as the chair of the state pension fund. *The Oregonian* newspaper reports:

> Unspoken, but also politically inconvenient is the compensation to attract talent from the private sector. The state's existing investment officers are some of the best paid public employees, making an average of $200,000 a year. But Treasury officials quietly complain that staff is underpaid by industry standards, and bristle about having to explain and get approval from the Legislature to release performance-based pay each year.... As Treasurer Read pleads: "If we have the talent, we will be able to make the decisions better."

Attempts by Treasurer Reed to hire better-paid investment professionals were rebuffed, with concerns about compensation exceeding members' wages and public pay scales, that is, outrage (Sickinger 2013). Table A.1 in the appendix provides a sampling of similar anecdotes. What is remarkable about the anecdotes is how similar tensions arise across many different types of pension systems and many different geographies of pension funds.

We begin with an agency model of portfolio choice and investment manager selection. Public pension trustees must hire and compensate an investment manager who constructs the portfolio over three assets – a mean-variance efficient risky asset, a political risky asset that is nonfrontier in returns, and fixed income. Boards choose the skill level (ability to capture the risk premium) of

[1] Hochberg and Rauh (2013) and Bradley, Pantzalis, and Yuan (2016) present evidence of pension fund overinvestment in local assets. They find that overinvestment leads to lower returns. Andonov, Hochberg, and Rauh (2018) document that the presence of politicians on pension fund boards leads to weaker performance in private equities. A theoretical and empirical understanding of the importance of underfunding and the resultant risking-up pressures for public pensions is found in Rauh (2009), Novy-Marx and Rauh, (2011), Ang, Chen, and Sundaresan (2012), Addoum, van Binsbergen, and Brandt (2015), and Adonov, Bauer, and Cremers (2017).

the investment manager. Boards then set the manager's compensation contract to attract the desired skill level and incentivize the optimal risk-taking in the portfolio. However, an outrage pay constraint (i.e., a threat of outrage repercussions implying an ex ante constraint) on skill binds for some public pension funds. Because trustees fear private costs from outrage if they pay market-level salaries, they instead choose to hire managers below a skill threshold to avoid compensation breaching the outrage pay constraint. In the model, we also incorporate the previously documented effects of political private benefit taking and unfunded liabilities to demonstrate how all the agency frictions might be independently important in asset management.[2]

The model produces comparative statics relating board agency to intermediate outcomes (investment manager skill and the riskiness of asset allocations) and then to the ultimate outcome of portfolio performance. Of particular interest are the predictions arising from tightening the outrage pay constraint. If the outrage constraint binds, the public pension fund hires lower skill managers, which then implies an inability to capture the full risk premium.

To test the theoretical predictions, we use a global sample of 176 public pension funds that account for \$5.4 trillion in assets at the end of our sample period and that cover the U.S., Canada, Oceania, and Europe for 1995–2014. The average (median) fund has \$45 (\$14) billion in assets under management (AUM). We hand-collect data on investment manager compensation.

Outrage is a latent concept; hence, we employ a standard technique in latent estimation from labor econometrics. We extract a latent variable *outrage* as the common component (the shared variation) among a set of measurement variables that in turn has maximal power to explain manager-adjusted compensation.[3] This is an application of confirmatory factor analysis (CFA), a special form of factor analysis used in Heckman, Stixrud, and Urzua (2006) and many other works. We then use the extracted outrage factor to study performance. This technique is different from the instrumental variables (IV) technique that could be applied to our setting if we used the measurement variables for outrage as excluded instruments for compensation. The IV approach would regress compensation on instruments, extracting the union of correlations between the instruments and compensation. In contrast, our CFA approach extracts only the intersection of correlations among the measurement variables that further relates to compensation, implying different exogeneity conditions than the IV exclusion restriction. We estimate all equations of our

[2] Private benefit incentives emerge from political motives (local economy building and direct vote chasing) to tilt investments locally, as documented by Bradley, Pantzalis, and Yuan (2016), Bernstein, Lerner, and Schoar (2013), Hochberg and Rauh (2013), Brown, Pollet, and Weisbenner (2015), and Dyck and Morse (2011). In addition, private-benefit-taking can emerge from pay-to-play schemes generating campaign contributions or direct side payments (Andonov, Hochberg, and Rauh 2018). Underfunding affects the risk preferences of boards (Andonov, Bauer, and Cremers 2017), as modeled in swinging-for-the-fences or gambling-for-resurrection models of Ang, Chen, and Sundaresan (2013) and Binsbergen and Brandt (2015).

[3] We normalize compensation by the median finance job salary in each location, to level the pay scale.

empirical model simultaneously within a structural equation modeling (SEM) model using generalized methods of moments (GMM).

Under income inequality aversion (Fehr and Schmidt 1999), a disutility depends on the extent to which an observed wage for someone else exceeds an individuals' reference wage. The concept of inequality aversion has strong micro-foundations in the theoretical and empirical literature that finds it to be a persistent and important behavioral factor in decision-making.[4] Our five measurement variables for outrage are the average wage level of working beneficiaries, the median local household income, and three variables capturing the composition of occupations among the trustees: the percentage of teachers, municipal workers, and budget civil servants (i.e., trustees who have responsibility for budget setting in local government). These trustee composition variables capture the notion that trustees particulary sensitive to relative wages and income inequality will be more sensitive to the threat of outrage.

Two exogeneity conditions must be satisfied to make a claim that outrage affects performance through compensation. First, we need the standard latent variable assumption that components of the measurement variables that are orthogonal to the CFA outrage factor – the measurement errors – are uncorrelated with compensation. We discuss possible violations to this assumption and show robustness drawing only from subsets of the measurement variables. We also estimate an IV form of the model to show robustness to specific CFA identification concerns. Second, our identification requires that outrage affects performance only through the compensation channel. Because outrage is the common component of the measurement variables, the second CFA exogeneity condition is less restrictive than the exclusion restriction in IV analysis. Indeed, the IV approach would require the exogeneity of each measurement variable, while our approach only requires the exogeneity of the common factor of the measurement variables that can explain compensation.

We find that a one-standard-deviation lower *Outrage* (associated with having—all at once—higher *Local income*, higher *Relative wages*, and a lower proportion of trustees who are *Municipal workers, Teachers,* and *Budget civil servants*) translates into higher compensation around the mean of $302,429. Further, the exogenous variation of compensation due to outrage has a significant relationship on performance. A one-standard-deviation lower *Outrage*, passing through *Adjusted log(compensation)*, produces a 11-basis-point (bps) higher portfolio excess returns over benchmark. In terms of economic magnitudes, relaxing the outrage constraint by one-standard-deviation at a cost of $302,429, the average fund would realize $49 million per year in value-added using the method of Berk and van Binsbergen (2015), and $15.5 million in value-added at the median.

[4] For example, see the theoretical advancements of Bolton and Ockenfels (2000) and the experimental evidence of Loewenstein, Thompson, and Bazerman (1989). Fehr and Schmidt (2003) summarize the empirical literature.

These results are robust to alternative ways to define the compensation variable, alternative ways to define the measurement variables for outrage, and the use of IV as opposed to SEM analysis.[5] In particular we note that our results are distinct from Andonov, Hochberg, and Rauh (2018), who study the impact of trustee composition on pension fund performance in private equity, characterizing trustees into nine categories based on both who appointed them and their profession. The clearest indication that our results derive from an alternative channel is our test where we exclude all trustee composition variables from the group of measurement variables for outrage, finding results that are consistent with our baseline estimates. Further, our trustee composition variables are defined differently as they are exclusively based on six professional categories of the trustee member. When using trustee categories as measurement variables for outrage, our results are robust to alernative trustee categorizations, and to dropping the trustee category closest to the political category used in Andonov, Hochberg, and Rauh (2018).

In further tests, we find that the positive performance associated with reductions in outrage at the portfolio level is driven by a strong link between outrage, log compensation, and within-asset-class excess returns in two risky asset classes: alternatives and public equities. Indeed, a one-standard-deviation lower *Outrage* implies 33-bps higher net returns on alternatives and 12-bps higher net returns on public equities. Further, we document that our results are not driven by funds that are insulated from the outrage threat having greater realized risk. We calculate a fund-specific tracking error and use it to define the *Information ratio* that we use as a measure of risk-adjusted performance. Consistent with our prior results, we find that an increase in *log(Compensation)* caused by a reduction in outrage has a positive effect on the *Information ratio*. More precisely, a standard deviation lower *Outrage* implies one-fifth larger *Information ratio* relative to the average.

Finally, we note that these results arise even while allowing, consistent with the prior literature, for distortions arising from politicians' payoffs to local investment and for distortions arising from underfunding to affect asset allocation and returns. Consistent with Andonov, Hochberg, and Rauh (2018) and Hochberg and Rauh (2013), we find that politicization has a direct negative effect on excess returns in alternatives asset classes: a one-standard-deviation increase in our *Political* variable reduces excess returns in alternatives by 41 bps, although this result is statistically insignificant in our sample. Overall, we interpret our results as complementing these papers, showing an important and neglected human capital channel through which political costs from compensation setting can also undermine returns.

In exploring the impact of politicized governance for public fund outcomes, our paper contributes to a large literature. Romano (1993), for example,

[5] In the IV analysis, we use the measurement variables as excluded instruments for compensation.

hypothesizes and finds that politicization affects fund performance. The author's study focuses on a sample of 50 public funds in the 1980s. Andonov, Hochberg, and Rauh (2018) find that political trustees are associated with lower fund performance in private equity. While our results are broadly consistent, our focus on the human capital channel leads us to consider limiting exposure to outrage, and thus different policy conclusions from these papers.

1. Institutional Background

This section provides background context regarding pension fund management contracting and the role of management and the board in investment decisions.

1.1 Manager contracting

Funds use several approaches to management contracting, oversight, and governance. In perhaps the most common structure, the board appoints a CEO who is responsible for both investment management and pension administration. In another common structure, the board appoints an investment manager (designated a CEO or CIO) responsible only for investment, separating the investment organization structure from that of pension administration. In a third structure, the board appoints a CEO responsible for the administrative activities and appoints a CIO responsible for investment.[6] One thing these structures have in common is the responsibility of the board of trustees to hire and supervise the activities of a manager (CEO or CIO), who will be totally or partially responsible for the selection of securities in the pension fund portfolio.

1.2 The process for making investment decisions in pension plans

Investment decisions in pension management involves the setting of asset class allocation targets and benchmarks and the selection of assets. The nominal power of the board in these investment decisions is laid out in statutes and in board investment policy statements. In most plans, board approval is required for the asset allocation targets and benchmarking. Boards also provide direct input on the delegation of mandates to outside managers, both in the extent to which and in the to whom decisions. In addition, boards indirectly assert delegated authority in the form of expenditure authorization on fees and/or setting dollar limits on asset class mandates. At CalPERS, for example, the 2019 board investment policy (a) specifies policy allocation targets of 50% public equity, 8% private equity, 28% fixed income, and 13% real assets, (b) delineates benchmarks therein, and (c) limits private equity and real asset external delegation to \$200 million and \$100 million, respectively.

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[6] Examples of the first structure include Australia Super, Canada Pension Plan Investment Board, Teacher Retirement System of Texas, New York State Teachers' Retirement System, School Employees Retirement System of Ohio, and Teachers' Retirement System of Oklahoma. Examples of the second structure include State of Wisconsin Investment Board and Massachusetts Pension Reserves Investment Management Board. Examples of the third structure include CalPERS and CalSTRS.

In practice, management retains significant influence over allocation decisions and benchmarking. The board's setting of targets involves an explicit provision for incorporating management advice, and, in practice, boards often fully defer to management. Likewise, most external mandate proposals to the board come from management, implying that this decision too is often implicitly deferred to management. When boards are more involved in the allocation targets, the board policy generally specifies a range around the targets and for tracking error from benchmarks, providing for management discretion. (At CalPERs, the board sets a range of $\pm 4\%$ for private equity and $\pm 5\%$ for real assets.) Thus, management has the authority to build portfolios with asset selection and to enact asset class decisions with some flexibility over time.

One takeaway from this section is that the pension plan investment process is complex and while it allows for board influence, it provides significant scope for management skill to affect outcomes. The next section introduces a model for portfolio choice that captures these characteristics. A second takeaway is that management control over decision-making may vary across plans. The empirical section explores whether our results are robust to such variation.

2. Model of Portfolio Choice with Political Agency Costs

Consider a setup in which a pension fund optimally invests in a mean-variance efficient portfolio over a risky asset and fixed income. The board of trustees for this pension fund achieves this objective by making manager-contracting choices to maximize beneficiaries' utility subject to manager participation and incentives. In our setting, because the pensions are *public pension funds*, being in the political domain can affect trustees' incentives and decisions. Although trustees have a fiduciary duty to act in the best interests of their beneficiaries, political private costs and benefits from their funds' choices create incentives to deviate from a strict interpretation of this duty. We call the resultant distortions *political agency costs*.

Our model and empirical analysis consider three political agency costs. The first emerges from outrage, the inability of politicized boards to pay optimally for investment manager skill because of political costs emerging from workers, retirees, and voters in the community. The second political agency cost emerges from politicized boards' preference for investing in political assets. Political assets are defined as investments that generate private benefits for a political board member, either in the form of local-tilted assets (which generate positive media attention, reputation, and ultimately votes and legacies) or in the form of pay-to-play allocations (which produce kickbacks from asset managers to politicians or political campaigns in return for asset allocations). The third political agency cost emerges from the pressure of liabilities that can induce public pension fund boards to risk-up portfolios to meet funding needs (e.g., to pay pensioners) rather than to have to face disclosure of shortfalls.

The focus of our model is on how these political agency costs affect asset allocations and pension plan performance, working through the mechanism of hiring and compensating an investment manager.

2.1 Assets and investment manager heterogeneity

A public pension fund board hires and sets a linear compensation contract for an investment manager to allocate the pension's capital among assets. Managers are risk averse and are assumed to have the same risk aversion as the beneficiaries of the pension fund, λ. Managers are heterogeneous in one dimension, their skill in the selection of assets within each asset class, represented by the parameter s. Skill levels are transparent, and their supply is perfectly competitive. A manager of type s has an outside option $O(s)$, where $O(\cdot)$ is an increasing function such that skilled managers have higher outside options.

The manager chooses portfolio weights among three assets: fixed income, a mean-variance efficient risky security (MV security) and a political asset. Fixed income pays a riskless return r_f:

$$\textit{Fixed income}: E\left[R_f\right]=r_f.$$

The MV security has variance σ_{MV}^2 and risk premium φ_{MV}:

$$\textit{MV security}: E[R_{MV}]=r_f+s\varphi_{MV}.$$

The political asset is also risky but has variance σ_P^2 and risk premium φ_P.

$$\textit{Political asset}: E[R_P]=r_f+s\varphi_P.$$

We assume that $\varphi_P/\sigma_P<\varphi_{MV}/\sigma_{MV}$ so that the MV security dominates the political asset in Sharpe ratio terms.

In both risky securities, managers earn a fraction s of the potential risk premium, in proportion to their skill. Only managers with maximal skill (i.e., $s=1$) can capture the full risk premium with their asset selections. This assumption is empirically motivated; while some investment managers in public pension funds have significant financial experience from working previously in a finance position in a public pension fund or the private sector, others' prior experience is limited to a managerial or civil servant role with no asset management responsibilities.

Managers form portfolios by selecting the weights on MV-efficient securities, political assets, and fixed income as w_{MV}, w_P, and $(1-w_{MV}-w_P)$, respectively.[7] For tractability, and consistent with Hochberg and Rauh (2013), we assume that the MV security and political assets have a joint normal distribution with correlation ρ, which is large enough to prevent hedging between asset classes.[8]

[7] A pension fund not affected by agency problems would invest in a combination of the MV security and fixed income.

[8] See the appendix for the explicit restriction that prevents the portfolio manager from taking short positions in any asset class.

2.2 Utility and political agency costs

Under the assumption of mean-variance preferences, the utility of the board equals that of beneficiaries if no political agency costs are at work:

$$U_{board}^{no\ agency} = U_{beneficiaries} = E\left[R - manager\ pay\right] - \frac{1}{2}\lambda Var\left[R - manager\ pay\right], \tag{1}$$

where R is the total return of the portfolio; manager pay is the compensation paid to an investment manager; and λ is the risk aversion of beneficiaries.[9] We introduce three political agency costs that cause the board's utility to deviate from that of the beneficiaries.

2.2.1 Outrage pay constraints. First, trustees in public pension funds are in a political domain, and this leads them to consider potential political costs arising from their choices. In the typical pension plan, trustees are either beneficiaries or politicians who employ and pay the beneficiaries. Costs arise for trustees if beneficiaries or others in the community who elect politicians become outraged by the compensation of the top executives of the public pension. The prospect of negative media attention and the resultant negative reputation consequences ensure trustees consider potential outrage in setting compensation.[10]

A basis for outrage of beneficiaries and those in the community is inequality aversion. Fehr and Schmidt (2003) cite voluminous experimental evidence consistent with inequality aversion. This evidence includes examples where subjects make choices to avoid inequality even when they know it will hurt them. See Loewenstein, Thompson, and Bazerman (1989), for example, or Engelmann and Strobel (2004), who find that most people value equality more than efficiency.

If the board were to set compensation sufficiently high such that outrage occurred, it would have to bear some utility cost:

$$U_{board} = E\left[R - manager\ pay\right] - \frac{1}{2}\lambda Var\left[R - manager\ pay\right] - outrage\ cost. \tag{2}$$

If trustees' utility consequences of outrage are sufficiently large, they would want to preclude the possibility of outrage altogether. The easiest way for trustees to ensure that compensation, which is stochastic, does not go over the

[9] The level of monetary compensation is not necessarily an indication of skill in pension plans and could reflect other forms of utility enjoyed by pension plan managers (e.g., they may put a higher value on the connections they make than nonpension managers would).

[10] The model abstracts from the fact that fees to outside managers may also generate outrage. It is difficult for beneficiaries and taxpayers to discover the compensation level of such outside managers, not least because their compensation usually depends on fees from multiple asset owners and is not subject to public disclosure requirements. Nevertheless, all that really matters is that outrage be more intensely focused against the compensation of internal managers' as opposed to that of external managers.

outrage threshold is to hire lower quality managers. To model this intuition, we assume that each fund has a threshold on skill, $s^{outrage}$. Thus, the board's utility reverts to Equation (1), but with a constraint:

$$U_{board} = E\left[R - manager\ pay\right] - \frac{1}{2}\lambda Var\left[R - manager\ pay\right],$$

subject to

$$(outrage\ \ constraint): s \leq s^{outrage}. \tag{3}$$

For some funds, the threshold is large and never binding. This is more likely if the reference wage level of beneficiaries or others in the community is sufficiently high.

2.2.2 Private benefits from politicized investing. Second, allocation choices can create private benefits for political trustees. These private benefits include votes garnered from investing locally and creating employment opportunities for local citizens, or side payments (e.g., in the form of campaign contributions or direct payouts) from pay-to-play arrangements.[11] We incorporate the political agency cost from private benefits from politicized investing in our model by assuming that the board receives a riskless, private benefit worth κ dollars for each dollar invested in political assets:

$$U_{Board} = E\left[R - manager\ pay\right] - \frac{1}{2}\lambda Var\left[R - manager\ pay\right] + \kappa w_P. \tag{4}$$

2.2.3 Liability-driven preference for risk. Finally, effective board risk aversion, λ_{board}, can be affected by liability obligations of the pension fund. Ang, Chen, and Sundaresan (2013) model the tensions pensions face due to the constant need to fund payments to retirees. Their main inference is that when funding is low, pension boards have a lower effective risk aversion, that is, a desire to "swing for the fences." The friction often at work is that boards face a personal reputational cost if they have to go back to legislatures to request funds to cover a down year of returns. The resultant risk-taking behavior is similar to the gambling for resurrection ideas of van Binsbergen and Brandt (2015). Such increased risk taking in the presence of underfunded liabilities has been found in U.S. public pension funds, for example, by Andonov, Bauer, and Cremers (2017).

We assume that underfunded status results in a higher risk appetite:

$$\lambda_{board} = \frac{\lambda}{\theta}, \tag{5}$$

[11] Andonov, Bauer, and Cremers (2017) find that U.S. pension funds with political boards tend to invest in local and less profitable private equity funds, and Dyck and Pomorski (2011) and Bernstein, Lerner, and Schoar (2013) show a similar pattern in the investments of sovereign wealth funds. Bradley, Pantzalis, and Yuan (2016) show not only a local bias but also a bias to invest in politically connected firms.

where θ is an exogenous politically determined variable that captures the risking-up pressure. The final utility formulation for the board, incorporating all political agency issues, is thus given by

$$U_{board} = E\left[R - manager\ pay\right] - \tfrac{1}{2}\lambda_{board} Var\left[R - pay\right] + \kappa w_P,$$

subject to

$$(outrage\ \ constraint): s \leq s^{outrage} \text{ if reference wages are low.} \tag{6}$$

2.3 Solving for the optimal contract and manger skill

We solve the model by considering the post-hiring portfolio choice, assuming that a manager with skill s already is hired. The board asserts its preferences for risk and for political investments by offering a compensation contract to the investment manager to induce the preferred portfolio choice. We derive this optimal contract for any skill level s. Next, we calculate the optimal manager skill s chosen by the board, from which we can determine the resultant asset allocation.

We restrict our model to linear contracts. The manager receives a cash salary c, independent of her performance. In addition, the board gives a share $1-a$ of the realized financial return to the manager to induce risk-taking. The board also asserts its political preferences by giving the manager an additional transfer of b dollars for each dollar invested in political assets.[12] Linear compensation is given by

$$manager\ pay(R, w_P | c, a, b) = c + (1-a)R + b w_P. \tag{7}$$

Like the beneficiaries, we assume that the investment manager has CARA utility with risk aversion λ. Thus, the manager chooses risk and political asset weight (w_{MV}, w_P) solving the following program:

$$\max_{w_{MV}, w_P} U_{manager} = \max_{w_{MV}, w_P} \left\{ E\left[manager\ pay\right] - \frac{1}{2}\lambda Var\left[manager\ pay\right] \right\}. \tag{8}$$

The board maximizes the expected monetary payoff penalized by the variance, with penalizing factor $\lambda_{board} = \lambda/\theta$, which depends on the risking-up pressure θ. The optimization problem is restricted by (a) the manager's incentive constraint and (b) the manager's participation constraint, which obligates the board to offer a contract that generates an expected utility for the manager not smaller than her outside option $O(s)$.

12 One interpretation of a political investment is a socially responsible investment, where trustees place a greater weight on the noninvestment impact than do management and where the investment has lower expected returns than does a nonpolitical risky investment.

The participation constraint is the channel connecting political asset investing to manager contracting. Because political assets are dominated in performance by the MV security, more politicized boards realize smaller utility increments from the skill of managers. Thus, the higher the political benefits κ are, the less willing is the board to pay compensation for skill.

The underlying program, which defines the optimal contract and the indirect utility $V_{board}(s)$ of the board when hiring the manager with skill s, is given by

$$V_{board}(s) \equiv \max_{c,a,T} U_{board} = E\left[R - manager\ pay\right] - \frac{1}{2}\lambda_{board} Var\left[R - manager\ pay\right]$$

$$+\kappa w_P = (\kappa - b) w_P + aE[R] - c - \frac{1}{2}\lambda_B a^2 Var[R], \qquad (9)$$

subject to

$$\textit{(participation constraint)}\ c + (1-a)E[R] + bw_P - \frac{1}{2}\lambda_M (1-a)^2 Var[R] \geq O(s)$$

$$\textit{(incentive constraint)}\ \{w_{MV}, w_P\} = \underset{w_{MV}, w_P}{\operatorname{argmax}} \left\{U_{manager} | c, a, b\right\}.$$

In the appendix, we show that the optimal contract is given by

$$a^* = \frac{\lambda}{\lambda + \lambda_{board}}, \qquad (10)$$

$$b^* = \left(1 - a^*\right)\kappa.$$

The optimal payment factor a^* reflects the standard sharing rule in which the less risk averse agent receives a larger component of the risky outcome. In the optimal contract, the manager receives the same fraction $1-a^*$ of the financial return R and of the political return κ. The resultant base salary c^* is the number that makes the participation constraint binding.

Finally, the board will choose the manager skill that satisfies the outrage constraint (if local reference wages are low) and maximizes their ex ante utility:

$$\max_s V_{board}(s), \text{s.t.} s \leq s^{outrage}. \qquad (11)$$

If the outrage constraint is not binding, then marginal disturbances around the optimal s^* are such that the marginal increase on the squared Share ratio is equal to the marginal cost of hiring a slightly better manager.[13] This first order condition allows for the calculation of the optimal value of s^* in the unconstrained case, which we call s^{free}. If outrage is binding, the public pension fund will hire the best manager they can within the constraints imposed by public outrage. Therefore, the general solution for the manager quality is given by

$$s^* = \min\{s^{free}, s^{outrage}\}. \tag{12}$$

2.4 Comparative statics

The solution (12) illustrates how funds differ in their cost-performance trade-off when choosing manager skill. For instance, both boards that face high private benefits κ from political investing, as well as boards that face an outrage constraint on compensation, prefer to hire managers with lower skill compared to the optimal manager for the beneficiaries. On the other hand, boards facing a personal cost from not having enough returns to cover pension liabilities might optimally choose a higher-skilled manager to benefit from risking-up the portfolio. Table 1 reports these comparative statics, focusing not just on how the agency issues affect manager contracting of skill, but to how ultimately these frictions translate into portfolio choice effects – allocations and performance.

Panel A isolates the effect of a binding outrage constraint on performance and allocations. The mechanical consequence of a binding outrage constraint is that the board of an outrage-prone pension fund hires a less-skilled manager ($\Delta s < 0$). The lower-skilled manager realizes lower risky asset returns ($\Delta R_{MV} < 0, \Delta R_P < 0$); thus, the board optimally sets a contract to induce more portfolio weight on fixed income ($1 - \Delta w_{MV} - \Delta w_P > 0$). There is no point in paying compensation for extra risk not rewarded with the capture of extra risk premium. The combination of more investment in fixed income and weaker managerial skill adds (on both counts) to a portfolio with poorer overall expected performance ($\Delta R < 0$).

Panel B looks at the partial derivatives with respect to changes in the other political agency issues. Boards with greater benefits from investments in political assets ($\partial\kappa$) hire less-skilled managers, since the expected return payoff from skill is lower in the portfolio tilted toward the political asset. Lower skill leads to smaller within-asset-class expected returns ($\Delta R_{MV} < 0, \Delta R_P < 0$) and less investment in the MV security ($\Delta w_{MV} < 0$). In addition, these boards design contracts to incentivize greater investment in the political asset ($\Delta w_P > 0$), which further reduces overall performance ($\Delta R < 0$).

By contrast, boards with higher liability-driven risk-up pressure (larger θ) hire more-skilled managers to take more advantage of the risky asset classes

[13] In the appendix, we show that this leads to the following first-order condition on the marginal payment to managers: $O'(s^*) = \frac{\left(\sigma_P^2 \varphi_{MV}^2 - 2\rho\sigma_P\sigma_{MV}\varphi_{MV}\varphi_P + \sigma_{MV}^2\varphi_P^2\right)s^* + \left(\sigma_{MV}^2\varphi_P - \rho\sigma_P\sigma_{MV}\varphi_{MV}\right)\kappa}{\lambda\sigma_P^2\sigma_{MV}^2\left(1-\rho^2\right)}$.

Table 1
Comparative statics: Political agency variables role

A. Effect of intensifying the budget outrage constraint

Variable	Model notation	Predicted change to row variable with: ∂ outrage	Test of prediction
Manager skill	∂s	≥ 0 (>0 if constraint is binding)	Tables 7, 11
Allocations			
Weight on MV security	∂(wMV)	≥ 0 (>0 if constraint is binding)	Table 12
Weight on political asset	∂(wP)	≥ 0 (>0 if constraint is binding)	Table 12
Weight on fixed income	∂(1-wP+wMV)	≤ 0 (<0 if constraint is binding)	Table 12
Weight on all risky	∂(wP+wMV)	≥ 0 (>0 if constraint is binding)	Table 12
Performance			
E[return on MV security]	∂(RMV)	≥ 0 (>0 if constraint is binding)	Tables 8, 10
E[return on political asset]	∂(RP)	≥ 0 (>0 if constraint is binding)	Tables 8, 10
E[portfolio return]	∂(R)	≥ 0 (>0 if constraint is binding)	Tables 8, 10

B. Effect of other political agency costs

		Partial derivative of row variable with respect to:		
Variable	Model notation	∂κ (κ: private benefits of political asset)	∂Θ (Θ: liability-induced preference for risk)	Test of prediction
Manager skill	∂s	<0	>0	Tables 7, 11
Allocations				
Weight on MV security	∂(wMV)	<0	>0	Table 12
Weight on political asset	∂(wP)	>0	?	Table 12
Weight on fixed income	∂(1-wP+wMV)	?	<0	Table 12
Weight on all risky	∂(wP+wMV)	?	>0	Table 12
Performance				
E[return on MV security]	∂(RMV)	<0	>0	Tables 8, 10
E[return on political asset]	∂(RP)	<0	>0	Tables 8, 10
E[portfolio return]	∂(R)	<0	>0	Tables 8, 10

This table lays out model predictions to show the comparative statics of how manager skill, portfolio choice, and returns change in the model with changes in political agency variables. The political agency issue of outrage is considered in panel A. Because outrage is a binding-or-not constraint, the comparative statics reflect a discrete change from not binding to binding. In panel B, the political agency issues of private benefits of political assets and underfunding are considered. In panel B, the comparative statics show the partial derivatives of a change in manager skill, allocations, or performance with respect to a change in agency, that is, the private benefits of political asset investing (κ) and the board preference for risk, driven by pension liabilities (Θ). The right column relates the prediction to the table of reference for empirical results.

($\Delta s > 0$, $\Delta w_{MV} + \Delta w_P > 0$), hence increasing within-asset-class and overall performance ($\Delta R_{MV} > 0, \Delta R > 0$). The extra risk that these boards induce may be rewarded with realization of additional returns, but it is above the level of risk desired by the beneficiaries. As stakeholders and taxpayers, beneficiaries may find themselves bailing out pension liabilities from taxes when bad returns realizations occur.

Although we do not explicitly include the cross partials in Table 1, one final piece of intuition is worth highlighting. When public pension funds have high liability pressures, the effect of an outrage constraint is very damaging: in this situation, public boards incentivize a poorly skilled investment manager to take on more risk. In doing so, the manager ends up with a riskier portfolio that underperforms in the risky asset classes.

3. Empirical Methodology

Our goal is to estimate how agency affects public pension fund outcomes. While we are interested in the impact of all of the political agency issues on fund performance, we set up our empirical methodology to focus on the novelty of our paper vis-à-vis the prior literature, namely, the compensation contract mechanism and the constraints arising from outrage. Because outrage is a latent concept, we employ standard techniques in latent variable estimation from labor econometrics to measure outrage. We then use a linear SEM of performance that builds off this latent extraction of outrage.

3.1 Confirmatory factor analysis

The stereotypical example of latent variable estimation in labor econometrics is the identification of latent ability in wage specifications. Because latent variable techniques are not employed as commonly in financial econometrics, we describe the identification foundations using a simplified version of Heckman, Stixrud, and Urzua (2006). In this work, labor earnings Y_i are determined by a vector of cognitive and noncognitive latent abilities, denoted by factor vector f_i, and a set of controls:

$$Y_i = f_i\theta + controls + \varepsilon_i. \tag{13}$$

The labor econometrician has a number of measurement variables for latent ability, such as test scores, grades, promotions, and other performance indicators. Her goal is to use the measurement variables to extract the shared variance that can be attributed to the ability factor f_i. Principal component analysis (PCA) can be used to solve the econometric problem by extracting the common factor of the measurement variables, that is, the factor that maximizes the common variance explained. However, PCA faces the critique that the factor is not easily interpretable in economic terms.

The preferred labor approach is confirmatory factor analysis (CFA), a special form of factor analysis used in Heckman, Stixrud, and Urzua (2006) and many other works.[14] The objective of CFA, like other factor techniques, is to capture the shared variance of measurement variables of a latent concept. However, differing from PCA, CFA restricts that shared variance to be that which in turn correlates with a dependent variable according to an economic model. The objective in Heckman, Stixrud, and Urzua (2006) is to extract the latent factor that can be interpreted in causal terms, as long as an exogeneity condition is satisfied. For example, in Heckman, Stixrud, and Urzua (2006), performance scores are the measurement variables for ability, and the CFA extracts the common component of performance scores that also explains wages. The

[14] See Carneiro, Hansen, and Heckman (2003), Hansen, Heckman, and Mullen (2004), Heckman, Stixrud, and Urzua (2006), Hanushek and Woessmann (2008), Cunha, Heckman, and Schennach (2010), and Heckman, Pinto, and Savelyev (2013) for other examples.

exogeneity condition that allows for causal inference from ability to wages is that the components of the measurement variables that are orthogonal to the CFA ability factor—the measurement errors—should be uncorrelated with wages.

3.2 Measurements variables of outrage

We use CFA to identify the impact of the latent variable outrage. A first step in implemening CFA is to introduce measurement variables for the latent variable outrage. We hypothesize outrage when an investment manager's wage exceeds reference wages. This is based on inequality aversion, where disutility depends on the extent to which an observed wage for someone else exceeds an individuals' reference wage (see Fehr and Schmidt 1999). The observed wage that we focus on is that given to the investment manager of the pension plan.

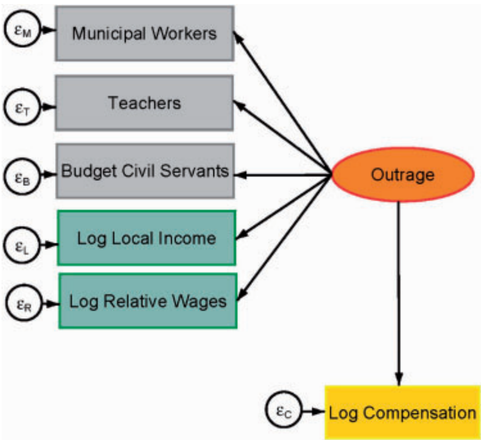
We use two measures of local wages as measurement variables for outrage. The first local wage measure is the average wage level of working beneficiaries. Trustees have fiduciary obligations to act in the best interests of beneficiaries who are likely to use their own wages as reference wages. The second local wage measure is the median local household income. This is a common reference wage for anyone living in the community.

The next three measurement variables for outrage capture the idea that a board of trustees will be more sensitive to outrage if it has a higher percentage of trustees of an occupation that is particularly sensitive to outrage. As discussed at greater length in the data section (Section 4), we categorize trustees by occupation from biographical information. We posit that boards that have a greater percentage of trustees with one of the following three occupations will be particularly sensitive to the threat of outrage: budget civil servants, teachers, municipal employees. We define as budget civil servants, trustees with titles such as controller, auditor, treasurer, revenue commissioner, and finance directors, as all such individuals are involved in the setting of compensation across multiple government agencies, and thus are acutely aware of pay differentials.[15] Teachers and municipal workers are well-known to be low-paid professions. Trustees from these occupations rise to the trustee level from within the organization of workers, rather than through an appointment like trustees who are finance professionals, lawyers, respected local corporate executives, or professors. In addition, the wages of teachers and municipal workers are tightly distributed, making the average wage salient in determining outrage and making the idea of extreme right-skewed wages of an investment manager more exceptional to their own and thus incomparable.

15 For example, when considering compensation levels for CALPERs investment officials, the budget civil servant trustee strongly argued for lower wages (option 1, in the following quote) using a reference wage argument: "So I ask this Committee to keep that in mind. These are State employees. We value the work they do. They do a fantastic job, and they're certainly worth the money that they make, but out of equity for the broader State workforce, who are also California civil service employees that do a great job for the State of California every day, I would respectfully ask the Board to consider option 1, and I would move option 1."

3.3 SEM and identification

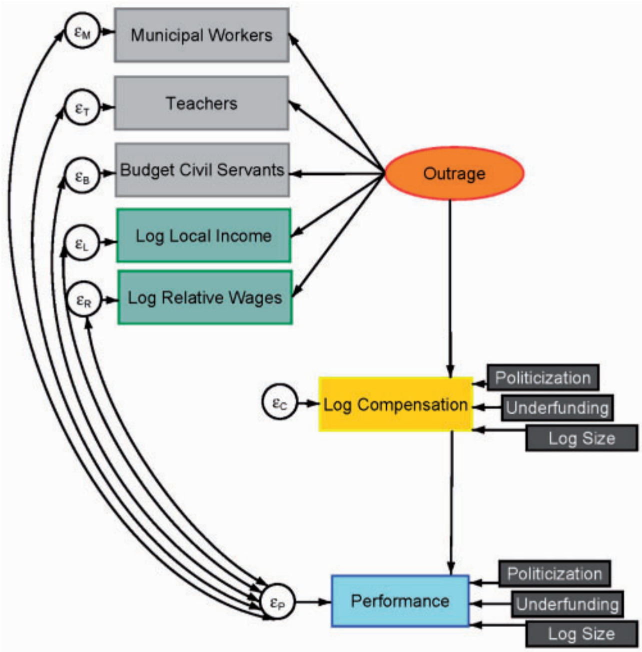
We apply CFA to estimate latent outrage from measurement variables for outrage and then use the outrage factor to study performance and other pension fund outcomes. We estimate both of these steps simultaneously within a SEM model using GMM. Graphical SEM representation 1 solely focuses on the CFA extraction of outrage, and graphical SEM representation 2 provides the full model, including the performance equation, and covariates for the compensation and performance regressions. In the graphical SEM, the object of pointing arrows is that which is explained, namely, a dependent variable. Residuals from all estimated relationships are represented by an ε term. Lines with double arrows that connect error terms represent allowed correlations, which are estimated in the system.



(1)

In graphical Equation (1), the latent variable *Outrage* is represented by an oval. The variable is extracted as the measurement variables' common component that has maximal power to explain manager-adjusted compensation. As noted, the five measurement variables for outrage are log local income, log beneficiary wages, municipal workers as a percentage of the board, teachers as a percentage of the board, and budget civil servants as a percentage of the board. As described in more detail in Section 4, we measure compensation as the log of the ratio of investment manager compensation to the median pay level of finance professionals in the same region. This adjusted compensation measure is intended to remove the influence of local labor markets and cost-of-living location effects.

Graphical Equation (2), the complete model, visually shows how the latent outrage factor affects compensation and then performance. The compensation equation includes the agency variables of politicization and underfunding, as well as fund size. The performance equation regresses performance on the latent outrage factor and includes the same covariates.



(2)

We can describe outrage as having a causal impact on performance so long as two exogeneity conditions are satisfied. First, the error term in the estimated compensation equation (ε_C) must be uncorrelated with the components of the measurement variables that are orthogonal to outrage ($\varepsilon_M, \varepsilon_T, \varepsilon_B, \varepsilon_L, \varepsilon_R$). This is the standard latent variable assumption.

One concern about the validity of this assumption might arise as a result of fund size. A larger fund may pay higher manager wages, independent of the pay of locals in the financial sector. Another example arises if the local economy has performed well, and the fund is home-biased, then both local reference wages and fund size may have grown in tandem, independently of our outrage mechanism. Including fund size directly in the compensation equation removes that variation, forcing the extracting of the shared variance to not load on that relationship. We also show robustness of results to normalizing compensation by fund size, rather than local financial professionals' median income to address concerns about nonlinearities and to appeal to the idea that, as an agent, a manager's compensation is a proportion of valuation.

Another concern might be that politicization could drive both the composition of trustees and compensation. For example, more politician trustees mechanically reduce all the other percentages. Again, by controlling for politicization in the compensation estimation, we force the extraction of the outrage factor not to load on that relationship. Yet, this topic is particularly important given the findings in Andonov, Hochberg, and Rauh (2018) that politicization affects performance in private equity. Thus, we return to other ways that politicization could enter the results throughout, since one measure of politicization may not capture the subtleties of relationships.

The second exogeneity condition is that the latent outrage must be uncorrelated with ε_P.In other words, it is assumed that outrage affects performance only through the compensation channel. On the other hand, the components of the measurement variables that are orthogonal to outrage can affect performance directly, as the latent variable-CFA formulation allows for and estimates the correlation between the measurement variables residuals $(\varepsilon_M, \varepsilon_T, \varepsilon_B, \varepsilon_L, \varepsilon_R)$ and ε_P.

3.4 Comparing SEM and instrumental variable estimation

The second CFA exogeneity condition is less restrictive than the exclusion restriction in instrumental variable (IV) analysis in which all measurement variables are used as instruments. In IV analysis, the total variation of each measurement variable would be used to estimate the effect of compensation on performance, thus requiring the exogeneity of all measurement variables for performance. The CFA formulation only uses the variation of latent outrage to estimate the effect of compensation on performance, which allows for potential correlations between the measurement variables and performance.

An example illustrates the less restrictive nature of this exogeneity condition in SEM. Suppose one of the measurement variables for outrage also plausibly had a direct connection to performance (e.g., the measurement variable is also correlated with a greater use of inefficient pay-to-play schemes or value-destroying asset allocation interference). If this economic connection is not shared by all the other measurement variables, we can estimate this set of correlations in the system, and therefore we do not need to worry about bias in our relationship between outrage-induced compensation and performance.[16] The second exogeneity condition instead just imposes that the common component of the measurement variables for outrage that maximally explains compensation be unrelated to performance except through compensation. This is a much easier argument, given that by construction the CFA factor is that which is maximally correlated with investment manager compensation and a common variation across all measures.[17]

This latter point—that our result is independent of the relationship between politicization and performance—is particularly important given the work of Andonov, Hochberg, and Rauh (2018). They show that state officials' representation on pension fund boards is negatively related to the performance of private equity investments. Our identification of outrage as a latent factor is exogenous to this effect, and a complement to their work.

16 Such a bias could arise with a relationship between fund size and fund performance, as explored in Gârleanu and Pedersen (2007), Dyck and Morse (2011), Pástor, Stambaugh and Taylor (2015), and Zhu (2018), among others.

17 In the data, the realized correlation between the latent outrage CFA factor and compensation is 0.60.

4. Data

4.1 Public pension funds sample

Our sample of pension funds emerges from the union of two data sets – the set of U.S. public pension funds covered by the Center for Retirement Research (CRR) of Boston College and the set of global public pension funds with at least $10 billion in assets identified in Pensions & Investments in 2011. Because of the need to search manually for the personal characteristics and compensation of trustees and managers, we limit the sample to funds in North America, Oceania, and Europe. Table A.2 in the appendix defines all variables and lists their sources. We convert all monetary data to 2010 U.S. dollars.

Table 2 reports statistics about our sample of public pension funds. Panel A, our estimation sample consisting of funds with compensation and trustee data, covers 112 public pension funds and 461 fund-year observations. As a comparison, panel B reports the full 176 funds and 2,644 fund-year observations. The funds in our estimation sample have a mean and median of $92 and $30 billion in assets, respectively, and are larger compared to the funds in panel B, where the mean and median size of funds are $45 billion and $14 billion in assets, respectively. The mean gross portfolio returns are similar between the estimation sample (panel A), where the gross return is 5.5%, and the full data set with a 5.8% gross return (panel B). As both panels show, 59% of the funds are from the United States and 41% are dispersed across Canada, Europe, and Oceania.

4.2 Allocations data

We collect each fund's asset allocation and performance over 1995–2011 in three asset classes: (a) alternatives (hedge funds, private equity and real estate); (b) public equities; and (c) fixed income. We order these asset classes in decreasing risk. The data come from a combination of sources: annual reports, funds' current and cached websites, direct requests to the funds, the Boston College CRR data set and CEM Benchmarking. In the estimation sample, reported in panel A of Table 3, on average funds have 23% of assets allocated to alternatives, 59% to public equities, and 33% to fixed income.

We also collect information on the fraction of assets managed via delegation, for the subset of funds that have this information. On average, the fractions of assets managed via delegation are 0.495 for fixed income, 0.733 for equities, and 0.756 for alternatives (excluding hedge funds, which are all outsourced).

4.3 Performance

Panel B of Table 3 reports the key performance statistics we use of gross returns, benchmark returns, and net returns. Benchmark returns are expressed for the asset class and are a weighted average of multiple sub-asset-class benchmarks within the indicated asset class, with the weights calculated using the beginning period weights. Trustees of a fund select the benchmarks, not the asset manager,

Table 2
Pension fund profile statistics

A. Sample with compensation and trustee data

	Assets under management ($billion)						Gross portfolio returns				
	Number of funds	Fund-year observations	Mean	25th percentile	Median	75th percentile	Fund-year observations	Mean	25th percentile	Median	75th percentile
Canada	10	97	49.68	13.14	33.78	81.30	97	0.0618	0.0074	0.0900	0.1296
Europe	17	108	255.58	15.36	62.89	253.20	108	0.0558	0.0090	0.0287	0.0870
Oceania	11	45	19.99	10.11	15.48	18.21	45	0.0570	0.0024	0.0856	0.1107
United States	74	211	44.00	11.76	26.13	59.64	211	0.0506	−0.0256	0.0604	0.1378
Total	112	461	92.42	11.89	29.51	71.78	461	0.0548	−0.0005	0.0559	0.1280

A. Full sample

	Assets under management ($billion)						Gross portfolio returns				
	Number of funds	Fund-year observations	Mean	25th percentile	Median	75th percentile	Fund-year observations	Mean	25th percentile	Median	75th percentile
Canada	16	290	47.34	11.82	19.67	73.94	217	0.0595	0.0040	0.0767	0.1219
Europe	39	431	106.53	9.14	18.46	59.93	304	0.0449	0.0049	0.0349	0.0825
Oceania	17	213	31.67	8.53	15.70	28.17	160	0.0578	0.0011	0.0832	0.1318
United States	104	1710	29.99	5.93	12.34	32.21	1188	0.0606	0.0012	0.0678	0.1383
Total	176	2644	44.51	7.14	14.10	38.90	1869	0.0577	0.0019	0.0565	0.1282

This table reports the assets under management and portfolio returns by region of the pension fund. Panel A presents these statistics for the pension funds for which we have manager compensation and trustee profile data. Panel B reports statistics for the full set of funds in our sample.

Table 3
Performance and allocation statistics

	Portfolio			Alternatives			Equities			Fixed income		
	Mean	SD	N	Mean	SD	N	Mean	SD	N	Mean	SD	N
A. Portfolio allocation												
Weights				0.230	(0.174)	261	0.589	(0.176)	305	0.326	(0.125)	262
Delegation fraction	0.336	(0.577)	158	0.756	(0.327)	216	0.733	(0.360)	192	0.495	(0.468)	182
B. Performance												
Gross returns	0.053	(0.111)	355	0.063	(0.130)	260	0.060	(0.211)	305	0.063	(0.046)	262
Benchmark returns	0.051	(0.106)	355	0.064	(0.119)	260	0.056	(0.207)	305	0.057	(0.029)	262
Net returns	0.002	(0.037)	355	−0.001	(0.081)	260	0.004	(0.020)	305	0.006	(0.031)	262
C. Risk-adjusted performance												
Tracking error	0.054	(0.212)	128	0.081	(0.066)	97	0.026	(0.044)	112	0.024	(0.019)	101
Information ratio	0.220	(1.036)	339	0.044	(1.078)	244	0.199	(1.049)	284	0.353	(1.229)	240

This table reports summary statistics of the portfolio weights and performance at the portfolio level and by asset classes. Asset classes are (a) alternatives, defined as hedge funds, real estate, and private equity; (b) public equities; and (c) fixed income. In panel A, we present the weights and the fractions of each asset class delegated to outside management. Panel B reports performance across three metrics: gross returns, benchmark returns, and net returns over the benchmark. Panel C report the cross-sectional tracking error and the time-varying information ratio.

mitigating some concerns about biases in benchmark selection. At the portfolio level, mean gross and net returns are 5.3%, benchmark returns are 5.1%, and net returns are 0.2%. Net returns are −0.1% in alternatives, 0.4% in equities and 0.6% in fixed income.

Panel C reports the performance measure adjusted by realized risk, the in-sample tracking error (*TE*). Tracking error is defined as the time-series sample standard deviation of the net returns of each fund. In mathematical terms, the tracking error of each fund i is defined as

$$TE_i = \sqrt{\frac{1}{T}\sum_{t=1}^{T}\left(R_{i,t} - R_{benchmark,t}\right)^2 - \left[\frac{1}{T}\sum_{t=1}^{T}\left(R_{i,t} - R_{benchmark,t}\right)\right]^2}. \quad (14)$$

The static tracking error calculation uses all the data available for each fund. According to panel C of Table 3, the average tracking error in our sample is 5.4%, similar to the tracking error of equity mutual funds, which are concentrated in the interval between 4% and 6% according to Petajisto and Cremers (2009).

We apply the tracking error to define the *Information ratio*, a measure of risk-adjusted performance. To construct the fund *Information ratio* for each fund and each year, we divide the fund's net return by the fund's static tracking error. This way, a fund that invests in riskier securities on average will be penalized with a larger static denominator. In mathematical terms, the *Information ratio* for the fund i and the year t by the equation:

$$IR_{it} = \frac{R_{i,t} - R_{benchmark,t}}{TE_i}. \quad (15)$$

Panel C reports that the average portfolio information ratio is 0.220, with a standard deviation of 1.036. This indicates that the average fund overperforms even when its performance is adjusted to its risk exposure, but many funds underperform. The risk-adjusted performance is heterogeneous across asset classes, with higher-risk asset classes (e.g., alternatives) having lower information ratios.

4.4 Investment manager compensation and skill data

The model assumes a single investment manager. In practice, the management team may include a CEO and a Chief Investment Officer (CIO). In some plans, the CEO focuses on administrative decisions and the CIO dictates investment decisions, whereas in other plans the CEO acts without a CIO to make investment decisions. Rather than *ex ante* specifying the title for the investment manager with responsibility for investment decisions, we collect the compensation information for holders of the CIO and CEO positions and use the compensation of the highest paid executive for our tests, assuming the highest paid executive will be most involved in investment decisions. For example, in

Table 4
Compensation, trustee occupation, reference wage, and other agency statistics

	Count	Mean	SD	25th percentile	Median	75th percentile
Manager compensation						
Manager compensation ($)	461	684,831	749,247	276,650	490,962	769,944
Median local CIO compensation ($)	461	172,561	34,634	147,439	174,374	194,820
log(Adjusted manager compensation)	461	1.06	0.879	0.40	1.04	1.54
Outrage: Reference wages						
Local income	461	54,530	17,889	40,277	47,154	66,001
log(Local income)	461	10.86	0.313	10.60	10.76	11.10
Relative wages	461	48,184	15,036	39,106	45,345	55,181
log(Relative wages)	461	10.74	0.294	10.57	10.72	10.92
Outrage: Trustee occupations						
Municipal workers (% trustees)	461	0.050	0.084	0.000	0.000	0.100
Teachers (% trustees)	461	0.112	0.167	0.000	0.077	0.143
Budget civil servants (% trustees)	461	0.11	0.146	0.00	0.09	0.15
Other agency variables						
Political	461	0.204	0.283	0.000	0.000	0.375
Underfunded index	461	0.182	1.124	−0.195	0.000	0.231

This table reports the summary statistics of the main variables characterizing the governance of pension funds in our sample. *Manager compensation* is defined based on the highest paid executive (CEO or CIO) for the public fund. The two outrage income measures—*Worker wages* and *Local income*—are, respectively the average wages of workers and the median income in the municipality where the fund is located. *Municipal workers* is the percentage of the board whose career is in the municipal labor force, defined as police, fire department, hospitals, libraries, and other noncivil servant positions. *Budget civil servant* is the percentage of the board whose background is in public sector financial positions (e.g., city controllers, auditors). *Teachers* is the percentage of the pension board who are teachers. *Political* is the average of two variables: a dummy taking a value of one if the chair is appointed by the government and the fraction of board members appointed by the government. *Underfunded index* is an index constructed by taking the mean across the standardized value of one minus the funded ratio and age, following Andonov, Bauer, and Cremers (2017).

CALPERS case the CIO is the most highly paid executive, and that is the salary we use.

We hand-collect compensation data for investment managers. For funds with mandated disclosure, we access compensation from annual reports and public filings. For funds that do not provide this compensation information, we conduct an exhaustive search for each named manager and public pension fund, using the name of the fund, the name of the executive, and a variety of keywords related to compensation. Newspapers are an important source, sometimes being able to access compensation information based on freedom of information requests. The resultant sample covers 112 public pension funds with a total panel of 461 observations, including all geographies spanned by our sample. Table 4 reports summary statistics on investment manager compensation in our data set. The median total compensation of the investment managers is $490,962 USD, with a mean of $684,831. For a quarter of the funds, managers make salary of $276,650 or less. This level of compensation is less than that for CIOs of nonprofit endowments, who earn on average $826,000 in the sample of Binfare and Harris (2020).

To normalize the scale of pension plan investment managers' compensation relative to the local market for finance professionals, we gather compensation of chief investment officers in the city where the fund is headquartered from PayScale. When city-level comparison data are unavailable, we use

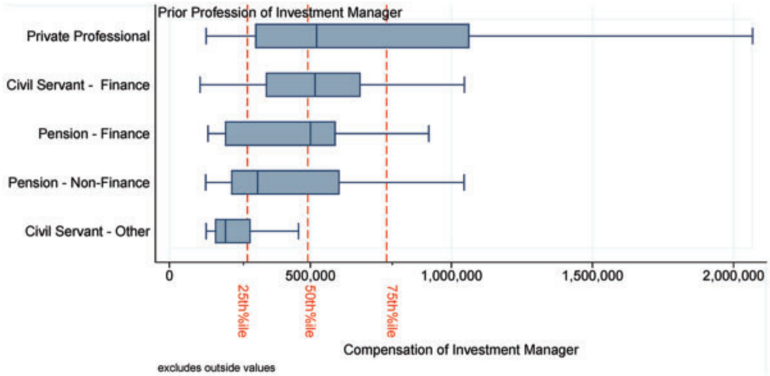


Figure 1
Investment manager compensation by prior profession
This figure graphs the distribution of investment manager compensation for each category of managers' prior professions. The box plot displays the mean (box center line) as well as the first (box edges) and second (stem edges) standard deviations. The dashed line indicates the overall sample's 25th, 50th, and 75th percentiles. The distribution of the sample is as follows (also reported in Table 6, along with the more detailed titles of the professions under the categories): Private Professionals, Finance Civil Servants, Pension (Finance), Pension (nonfinance), and Civil Servant (nonfinance).

country-level data from other sources (SalaryExpert for Denmark and Norway, SalaryExplorer for Sweden, Hudon for New Zealand, and PayScale for Netherlands). As shown in Table 4, the median local CIO compensation is $172,561. To remove the local geography effect, the variable we use in estimation is the natural log of the ratio between *Manager compensation* and *Median local CIO compensation*.

Our model assumes managers with greater skill have a higher opportunity cost and in equilibrium receive higher compensation. We hand gather data on the professions of all investment managers immediately prior to their current role as an investment manager, as one way to validate links between skill and compensation. As reported in Table 5, panel A, for almost two-thirds of the fund managers their immediate prior experience was in finance in some capacity. About one in three (31%) were private professionals with a senior role in investing in the private sector, about one in three (30%) worked in the civil service as a bureaucrat with budgetary responsibility, and 5% worked as a senior investment manager at another pension fund. Notably for the other one-third of investment managers, their prior experience was either as a civil servant with no financial expertise (16.4%) or as a nonfinancial executive in a pension fund (18%).

Figure 1 depicts box plots of the distribution of compensation by prior profession categories. The red dashed vertical lines represent the quartile cutoffs from the full sample. As expected, the group with the highest median pay, is the group of pension managers that were private professionals. Those without prior professional finance experience clearly earn lower compensation. The median

Table 5
Professions of investment managers and trustees

Occupation	Description	Professions represented	%
A. Investment managers' professions			
Prior pension executives			
Pension: Investment executive	Investment manager from another pension fund	Director of Investment, CEO, CIO	4.9
Pension: Other executive	Other executive position in another pension fund	Assistant General Counsel, Assistant Executive Director, Deputy Executive Director, Chief of Staff, COO	18.0
Prior private firm finance professionals or executives			
Private firm profes-sional	Financial position from privately firm	CEO, CIO, Director, Managing Partner, Accountant, Actuary, Auditor, Consultant, CRO	31.1
Civil servants			
Civil servant (budget)	Civil servant with financial experience	Treasurer, Auditor, Accountant, Controller, Budget Officer, Finance Director, Public Institution Professor	29.5
Civil servant (nonbud-get)	Civil servant without financial experience	City Council CEO, City Manager, Executive Director, Department of Correction Administrator, Deputy Chief of Staff, Director, Executive Commissioner, Natural Resource Advisor, Teacher, Senator	16.4
B. Trustees' professions			
Civil servants			
Politician	Includes any representative or elected official of municipal, state, or federal government	Senator, House Representative, Mayor, Governor, Lieutenant Governor, Secretary of State, Attorney General, Assembly Speaker, State Representative, Secretary, Minister, Borough President, City Manager, Assistant Deputy Minister, Deputy Governor, Premier Deputy Chief of Staff, Deputy Minister, City Council, County Commissioner, Deputy City Manager, Deputy General Counsel,	6.4
Budget civil servant	Civil servant with financial experience	Treasurer, Auditor, Accountant, Controller, Budget Officer, State Finance Director	34.4
Other civil servant	Civil servant without financial experience	Judge, Prosecutor, Clerk, Commissioner, Assistant Commissioner, Professor, Dean	13.7
Noncivil servants			
Teacher	Teachers	Teachers	14.7
Municipal worker	Workers providing services to city residents, union labor	Police Officer, Fire Officer, Jail Worker, Railway, Steel, Construction, Electrician, Mail Employee, Librarian, Miner, Bus Driver, Chimney Sweep, Food Worker, Manufacturing Worker, Telecommunications	7.7
Professionals	Local private sector professionals and NGO executives	Financial Sector Expert, Doctor, Nurse, Dentist, Private Firm CEO, CIO, Chairman, Pharmacist, Journalist, Media Professional, Architect, NGO Chairman, Owner of Private Firm	23.1

This table reports the immediate prior profession of investment managers (panel A) and the current professions of trustees (panel B). The data are collapsed to the cross-section of public funds. All data are hand-collected.

compensation of pension executives with a prior nonfinance role is $312,911, and that of those who have held a prior nonfinance civil servant role is $198,729.

4.5 Measurement variables for outrage

As noted earlier, we use five measurement variables for outrage, and report summary statistics for these variables in Table 4. The first two measurement variables are reference wages of beneficiaries. First is the median regional income in the municipality (or MSA) where the fund is located, compiled from the agency responsible for collecting and compiling income statistics in each country. We call this variable *Local income*. The average (median) *Local income* is $54,530 ($47,154), translating into a mean *log(Local income*) of 10.86. Second is the wage of the working beneficiaries. We call this variable *Relative wage*. We collect the average wages of working beneficiaries both directly from the annual reports or as a calculation from data on the employee contributions and the reported average rates of contributions (also predominantly from funds' annual reports). The average (median) *Relative wage* is $48,184 ($45,345). In our tests, we follow convention and use the *log(Relative wage)*, which has a mean of 10.74.

The next three measurement variables are based on the proportion of trustees that have a particular occupation that would make them sensitive to outrage. We identify trustee occupation variables from, first, sourcing the names of the trustees from websites and, then, looking up biographical information from CV's on the funds' websites or other web information sources (e.g., Linkedin). Data availability force us to use a single cross-section of data (2011) for trustee biographies. Trustees do not turn over frequently, and our panel is short; thus, this shortcoming is likely not very significant. We assume that the trustee's current profession provides a sense of an appropriate wage level and captures an awareness that a trustee might have toward the beneficiaries' outrage predispositions.

Table 5, panel B, reports six occupational categories. Politicians (those representing the government at large or elected as a politician) account for 6.4% of board seats. Budget civil servants (most commonly, treasurer, revenue commissioner, controller, auditor, and finance directors) account for 34.4%. Other civil servants (clerks, commissioners, public university academics, and legal government officials) account for 13.7%. The remainder of trustees are from three categories: professionals at 23.1% (including finance, medicine, media, NGOs, or other private firms); teachers at 14.7%; and municipal workers at 7.7% (fire workers, librarians, workers at city hospitals, and other such public municipal service occupations that are not internal to the running of the government administration *per se*).

We use the proportion of trustees as teachers, municipal workers, and budget civil servants, as measurement variables for outrage. Teachers and municipal workers are well known as low paid professions and often are the occupations of the beneficiaries, making them especially sensitive to the threat of outrage.

Budget civil servants are involved in the setting of compensation across multiple government agencies, and thus are acutely aware of pay differentials. To further validate this assumption that budget civil servants are sensitive to outrage, we gather information on compensation-related votes from the public records of pension plans. CalSTRS, a large pension plan in California, provides publicly available information on the minutes of meetings and reports votes on compensation by trustee, which we are able to link to trustee occupation types. As summarized in Internet Appendix Table 1, in 9 of the 11 compensation-related motions at CalSTRS over the 2009 to 2019, a budget civil servant voted against higher compensation. In 5 of 11 events, a teacher trustee voted against higher compensation. In one event, a professional voted against higher compensation. In one event, a nonbudget civil servant voted against higher compensation.

4.6 Other political agency variables: Politicization and underfunding

Our empirical methodology allows us to focus on the novelty of our paper vis-à-vis the prior literature—the compensation contract mechanism and the constraints arising from outrage. To account for alternative political agency costs highlighted in the literature, arising from power of politicians in the board, and of underfunding, we introduce two variables to capture these impacts in all our analyses.

First, to capture the strength of politicization of a pension fund we introduce a variable called *Political*. This variable is an equal weighted average of two inputs: whether the board has a political chair (if the chair is appointed by an executive of government, 51% of the sample) and politicians as a percentage of the board (from Table 5). *Political* has a mean of 0.2.

To measure the extent of underfunding pressures, we create an index of two variables. We have data on the funded ratio (the level of assets-to-liabilities), but not for all funds. The other measure of liability strain comes from Andonov, Bauer, and Cremers (2017), who find that funds with a higher age profile of pension beneficiaries have more liability concerns. We construct the average age of pension beneficiaries, using data on the average age of workers and retirees with the fraction of members being retired. Then we construct the *Underfunded index* as the negative of the standardized funded ratio plus the standardized age variable. The *Underfunded index* has correlations of 0.81 with age and of -0.79 with the funded ratio.

5. Results

5.1 Portfolio performance results

5.1.1 CFA extraction of outrage. Table 6 reports our estimation of the CFA-SEM structure depicted in graphical Equation (2). The table reports results from the system estimation using the portfolio net returns as the final outcome variable. The first five columns present the CFA *Outrage* extraction results.

Table 6
Outrage and portfolio performance: Initial results

	Local income	*Relative wages*	*Municipal workers*	*Teachers*	*Budget servants*	*log (compensation)*	Portfolio performance
Outrage (latent)	−0.594***	−0.105***	0.0770***	0.0187***	0.0508***	−1	
	[0.0583]	[0.0206]	[0.0156]	[0.00690]	[0.0137]	[constrained]	
log(compensation) (Endog)							0.00300**
							[0.00151]
Political						−0.1254	−0.00322
						[0.0940]	[0.0053]
Underfunding index (lag)						−0.0332	0.00255
						[0.0348]	[0.00211]
log(size) (lag)						0.296***	−0.000581
						[0.0304]	[0.00213]
Time trend control	Yes	Yes	No	No	No	Yes	No
			Observations	352			
			Number of funds	93			
			Standard deviation of outrage	0.366			
			Pay for standard deviation of outrage	$122,172			

This table reports the coefficients for the structural equation modeling (SEM) system of equations from Equation (2). The latent outrage variable is assumed to have a causal effect on *log(Compensation)*, which, by its turn, affects the outcome variable in the last column. The first five columns report the coefficients for the latent outrage (right-hand side) on each outrage measure (left-hand side). The sixth column uses the adjusted compensation in the left-hand side and constrains the coefficient of the outrage variable to -1 without loss of generality. The seventh column reports the effect of the adjusted compensation on excess portfolio performance over the benchmark. As measurement variables for outrage, we include mean pension workers' wages, median local area income, the percentage of trustees that are *Municipal workers*, the percentage that are *Teachers*, and the percentage that are *Budget civil servants*. *Political* is the average of two variables: a dummy taking a value of one if the chair is appointed by the government; and, the fraction of board members that are appointed by the government. The *Underfunded index* is constructed by taking the mean across the standardized value of one minus the funded ratio and age, following Andonov, Bauer, and Cremers (2017). *log(Size)* is the log of the lagged fund AUM. The number of funds per estimation is indicated below the number of observations. The equations for *Local income*, *Relative wages*, and *log(Compensation)* include a time trend control. We estimate all parameters jointly using GMM, and all standard errors are calculated via bootstrap. $^{*}p<.1$; $^{**}p<.05$; $^{***}p<.01$.

As predicted, the common correlation that maximally explains compensation has a negative relationship to reference wages of beneficiaries and a positive relationship to the proportion of trustees most sensitive to outrage. (Recall that compensation is already adjusted for the finance profession local pay level, so these results are not based on this local mechanical relationship.) All of the measurement variables are significant at the 1% level in their participation with CFA-extracted *Outrage*. Column 6, where *log(Compensation)* is the dependent variable, describes the relationship that gives economical meaning to the CFA factors. This estimation of the CFA factor that maximally predicts compensation takes the covariates *Political, Underfunding*, and *FundSize* into account. That is, we allow these covariates to absorb their correlation with *log(Compensation)* so that it does not load on *Outrage*. Table 6 shows that of these covariates, only *FundSize* has a relationship with the endogenous compensation. In an ordinary least squares (OLS) estimation, the partial R-squared of *FundSize* in explaining compensation is 0.26. The size of a pension fund is clearly important for compensation.

Without loss of generality, in column 6 we normalize the loading of *Outrage* on *log(Compensation)* to be -1, which implies that we are measuring the impact of *Outrage* in *log(Compensation)* units. We provide relevant statistics for assessing the economic impact of the *Outrage* factor at the bottom of the table. The *Outrage* factor variable has a standard deviation of 0.366. This implies that a one-standard-deviation lower *Outrage* translates into higher manager compensation (around the mean) of $302,429. The average fund has 11 trustees, a reference *Local Income* of $54,530 and *Relative Wages* of $48,184. A one-standard-deviation lower outrage is associated with having—all at once—$10,649 more of *Local Income* (60% of its standard deviation), $1,815 more in *Relative Wages* (12% of its standard deviation), a reduction of one-third of a *Municipal worker* (34% of its standard deviation), a reduction of one-twelfth of a *Teacher* (4% of its standard deviation), and a reduction of one-fifth of a *Budget civil servant* (13% of its standard deviation). This all-at-once calculation is an artifact of the CFA technique. In Section 5.3, we instead apply economic magnitude calculations based on observed pension reforms.

5.1.2 Effect of outrage on performance. The final column of Table 6 presents the performance equation results. The dependent variable is excess portfolio performance over the benchmark (i.e., portfolio net returns), where the benchmarks are applied according to portfolio weights at the sub-asset-class level.

In the performance results of Table 6, the final column, *log(Compensation)* now enters as an independent variable. Importantly, when we consider the effect of compensation on performance, the only variation of *log(Compensation)* being used in our GMM procedure is the exogenous variation of *Outrage*, akin to saying "the instrumented effect of" an X variable in an instrumental variable estimation procedure.[18]

We find that *log(Compensation)* explained by *Outrage* has a positive and significant effect on portfolio net returns, with a coefficient of 0.00300. A one-standard-deviation of *Outrage* reduction, passing through *log(Compensation)*, produces a 11-bps greater *Portfolio performance* (0.366*0.00300 = 0.11%).

To assess the dollar impact of outrage on performance, we need a reference firm. Evaluated at the average estimation sample fund size of $92.42 billion in AUM (from Table 2, panel A), a one-standard-deviation reduction in *Outrage* would improve value-added for that pension of $101 million in AUM per year. A typical pension fund is substantially smaller than firms in our estimation sample. For the average fund in our full sample (from Table 2, panel B), a one-standard-deviation reduction in *Outrage* would increase value-added by $49 million per year at the mean, and $15.5 million at the median. For a 25th percentile fund of the representative sample, the benefit from unwinding outrage is $7.8 million in additional AUM per year. Recall that the cost in compensation associated with these gains is $302,429.

[18] Our structural equation modeling does not assume that *log(Compensation)* is exogenous to performance; that is, the OLS orthogonality condition *log(Compensation)*$\perp \varepsilon_P$ is not used in our estimation procedure. This implies that the endogenous variation of *log(Compensation)* cannot be used to estimate the effect of *log(Compensation)* on *PortfolioPerformance*.

Table 7
Robustness results

	Local income	*Relative wages*	*Municipal workers*	*Teachers*	*Budget servants*	*Compensation*	Excess returns portfolio
A. Compensation not adjusted							
Outrage (latent)	−0.660***	−0.114***	0.0202	0.115***	0.0678***	−1	
	[0.123]	[0.0359]	[0.0129]	[0.0399]	[0.0200]	[constrained]	
log(Compensation) (Endog)							0.00470**
							[0.00209]
Observations	352						
B. Compensation adjusted by fund size							
Outrage (latent)	−0.429**	−0.106	0.0192	0.0957*	0.0594***	−1	
	[0.205]	[0.109]	[0.0146]	[0.0521]	[0.0216]	[constrained]	
log(Compensation) (Endog)							0.00304*
							[0.00166]
Observations	352						
C. Excluding funds with 2011 size below 10 billion							
Outrage (latent)	−0.532***	−0.0646***	0.0115*	0.0936***	0.0471***	−1	
	[0.0575]	[0.0227]	[0.00623]	[0.0274]	[0.0136]	[constrained]	
log(Compensation) (Endog)							0.00418**
							[0.00213]
Observations	304						
D. Region FE							
Outrage (latent)	−0.498***	−0.329***	0.00479	0.0441	−0.0285	−0.927***	
	[0.0229]	[0.0643]	[0.0188]	[0.0457]	[0.0518]	[0.151]	
log(Compensation) (Endog)							0.00567*
							[0.00335]
Observations	352						
Region FEs	Yes	Yes	Yes	Yes	Yes	Yes	Yes

(*Continued*)

These portfolio tests with micro-level benchmarking help to control for risks in portfolios. Nonetheless, deviations from the benchmark have distributions varying by asset class, and within-asset-class securities selection varies by risk. In subsequent tests, we consider within-asset-class estimations and estimations calculated on information ratios rather than net returns to show robustness of our portfolio-level findings and to uncover the asset class decomposition of the results.

5.1.3 Robustness. Tables 7 and 8 assess the robustness of these results to potential concerns about the exogeneity assumptions, particularly the first exogeneity assumption. Recall that this first exogeneity condition, the standard latent variable assumption, is that components of the measurement variables that are orthogonal to the CFA outrage factor are uncorrelated with compensation.

5.1.3.1 Compensation adjustment. Our main Table 6 analysis normalizes compensation to the median compensation of finance professionals in the region. We first consider whether the CFA extraction might be picking up the denominator (the regional effect of the finance professional's salary) rather the

Table 7
(*Continued*)

	Local income	*Relative wages*	*Municipal workers*	*Teachers*	*Budget servants*	*Compensation*	Excess returns portfolio
E. Controlling for delegation							
Outrage (latent)	−0.277***	−0.0667	0.0370**	−0.00728	0.0368	−1	
	[0.0615]	[0.0573]	[0.0160]	[0.0238]	[0.0309]	[constrained]	
log(Compensation) (Endog)							0.00977*
							[0.00523]
Delegation index							0.00634
							[0.00961]
Observations	214						
F. Income measurements							
	Localincome	*Relativewages*				*Compensation*	
Outrage (latent)	−0.597***	−0.107**				−1	
	[0.166]	[0.0428]				[constrained]	
log(Compensation) (Endog)							0.00425***
							[0.00163]
Observations	352						
G. Trustee measurements							
Outrage (latent)			0.00554	0.0496***	0.0242***	−1	
			[0.00592]	[0.00860]	[0.00529]	[constrained]	
log(Compensation) (Endog)							0.00270*
							[0.00157]
Observations	352						

This table reports the coefficients for the structural equation modeling (SEM) system estimated under different assumptions. In panel A, we use the log of the CEO/CIO compensation not adjusting for the local level of executive pay. In panel B, we use the log of the CEO/CIO compensation divided by the fund size. In panel C, we exclude all funds that had less than $10 billion in assets in 2011. In panel D, we include region fixed effects for the region of the United States, Canada, Europe, and Oceania. To facilitate the convergence of the GMM estimators, we estimate this model in two steps. First, we estimate the model without fixed effects, as in Table 6. We then use the estimated *Outrage* in the second stage, when we estimate the model with region fixed effects; however, we assume that *Outrage* is observed, not latent. In panel E, we control for the *Delegation index*. In panel F, we exclude the trustee measurement variables. In panel G, we exclude the local income and relative income measurement variables.

numerator (the investment manager’s compensation). If so, the whole nature of the exogeneity assumption changes, and we would not be measuring a fund characteristic, but rather a location effect. In Table 7, panel A, we repeat the Table 6 test using simply the investment manager’s compensation numerator, unnormalized. The results are strikingly similar. *log(Compensation)* explained by *Outrage* continues to have a positive and significant effect on portfolio net returns, with a coefficient of 0.00470.

5.1.3.2 Fund size. Next, we consider the possibility that outrage could relate to not only the absolute level of compensation but also to compensation as a fraction of fund size, as this measures the percentage of their pension the workers are giving up to the manager. In panel B, we revise our compensation measure by expressing adjusted compensation as a ratio to fund size, then repeat the test from Table 6. Again, results are similar with *log(Compensation)* adjusted by fund size explained by Outrage having a positive and significant effect on portfolio net returns, with a coefficient of 0.00304. In the rest of the

Table 8
Instrumental variables results for portfolio returns

	Including *Budget civil servants*		Excluding *Budget civil servants*	
	Equation (1)	Equation (2)	Equation (1)	Equation (2)
Outrage-predicted log(Compensation)		0.00770***		0.00872***
		[0.00260]		[0.00277]
Local income	0.691***		0.789***	
	[0.206]		[0.206]	
Relative wages	0.602*		0.588	
	[0.353]		[0.360]	
Municipal workers	−0.912		−0.734	
	[0.609]		[0.605]	
Teachers	−0.397		−0.229	
	[0.300]		[0.296]	
Budget civil servants	−0.837**			
	[0.361]			
Political	−0.0786	−0.00316**	−0.0522	−0.00271*
	[0.0913]	[0.00131]	[0.0975]	[0.00140]
Underfunding index (lag)	0.0742*	0.000985	0.0738*	0.000704
	[0.0392]	[0.00132]	[0.0385]	[0.00136]
log(Size) (lag)	0.261***	−0.00241**	0.251***	−0.00316**
	[0.0808]	[0.000980]	[0.0837]	[0.00123]
Observations	352	352	352	352

Columns 1–4 report IV/GMM estimates. Columns 1 and 2 report the first stage in which we regress the log compensation on the measurement variables for outrage used as instruments (local income, relative wages, municipal workers, teachers, and budget civil servants) and on controls (*Political*, *Underfunding index*, and *log(Size)*). In columns 3 and 4, we regress portfolio returns on the *Outrage-predicted log(Compensation)* and on controls. $*p<.1$; $**p<.05$; $***p<.01$.

tests, we retain the baseline compensation measure, given the strong micro-foundations for inequality aversion linked to the level of compensation, as a basis for outrage

We also address the possibility of a potential bias from our sample selection criteria to focus on large non-U.S. pension funds with a size above $10 billion in 2011. Given that 2011 is toward the end of the sample, this could result in a bias toward well performing funds. To understand whether this cutoff is somehow driving our results, we perform a simple test and exclude the U.S. funds with 2011 size below $10 billion – around 14% of the sample. If the choice of this cutoff is relevant for the coefficient estimation, one would expect the application of the same cutoff to the U.S. funds to significantly change the coefficients estimates. However, as can be seen in panel C, all numbers are very similar to what was obtained in Table 7.[19]

[19] We also address the possibility that the benefits from greater log adjusted compensation (which captures skill) may be attenuated for larger pension plans. Zhu (2018) is one paper that documents diseconomies at the fund level. Pastor, Stambaugh and Taylor (2015) document diseconomies at the industry level, with insignificant results at the fund level. Diseconomies have not been found in studies of endowment funds (e.g., Brown, Garlappi, and Tiu 2010), the managers of which have greater degrees of freedom than mutual fund managers to reallocate across asset classes. In untabulated tests, we introduce a dummy variable for large funds (top 25th percentile) and then interact this variable with log adjusted compensation. We find fund size and its interaction with log-adjusted compensation are both insignificant.

5.1.3.3 Heterogeneity in management authority, regional effects, and delegation. Board involvement in decision-making has the potential to limit management's ability to earn their skill premium, and board involvement could differ across plans, contributing to differences in plan performance. The test in panel C indirectly helps to address this concern, if an important source of heterogeneity relates to plan size, as we found no difference in results when the sample is restricted to larger plans. The tests to be introduced shortly in panels F and G also provide additional robustness. In this subsection, to more directly address the robustness of our findings to such concerns, we introduce two new tests to capture potential heterogeneity in management authority and explore whether this affects the relationship between outrage and performance.

First, in panel D, we introduce a specification that includes region fixed effects for the US, Canada, Europe, and Oceania. To the extent that differences in ability to earn the skill premium are related to governance rules, there is likely to be commonalities to those rules across regions. To facilitate the convergence of the GMM estimators, we estimate this model in two steps. First, we estimate the model without fixed effects, as in Table 6. We then use the estimated *Outrage* in the second stage, when we estimate the model with region fixed effects, however assuming that *Outrage* is observed, and not latent. We find, as before, *log(Compensation)* explained by *Outrage* has a positive and significant effect on portfolio net returns, with a coefficient of 0.00567.

Next, we account for the fact that in some funds there is greater use of delegated asset management, and we explore whether this heterogeneity in delegation could drive performance, as management teams that rely more on delegated asset management may have greater constraints in their ability to earn their skill premium. For this exploration, we construct a *Delegation index* defined as follows. For each asset class with delegation data available, we normalize the delegated fraction by subtracting the cross-sectional mean and by dividing the result by the cross-sectional standard deviation. We then define the delegation fraction to be the weighted average of the normalized delegation fraction across the asset classes for which we have data available.[20]

Outrage could influence delegation. To the extent that boards are aware of outrage and the constraints outrage imposes, boards may want to reduce these negative consequences. If there is a stronger impact of outrage on compensation for internal than external managers, then one prediction is that outrage-affected plans would use more delegated management. We cannot prove that outrage is greater for internal rather than external managers, but we have good reasons to expect it to be so. It is difficult for beneficiaries and taxpayers to discover the compensation level of outside managers at financial intermediaries. When pension plans disclose their payments to financial intermediaries, they do not report payments to specific named managers at the intermediary. The financial

[20] We use the *Delegation index* as a control in order to maximize the sample of the model estimated in Table 7, panel E, given that our delegation data are very sparse.

intermediaries do not disclose this information. As a result, beneficiaries are unable to identify the specific compensation of individual external managers. In the absence of such information, it is unlikely that inequality aversion will arise.

We introduce two tests using delegation. In Table A.3 in the appendix, we explore outrage and the use of delegated asset management as a dependent variable. For this test, we also include the standard control variables of *Political, Size*, and *LagUnderfunded*, along with year fixed effects. We find a positive and significant coefficient for outrage of 0.387. In terms of economic magnitude, a one-standard-deviation reduction in outrage would shift the delegation fraction downward by 10 percentage points, almost one-third of the mean level of delegation. More importantly, in panel E we introduce the *Delegation index* as a control in the performance equation, allowing us to see whether our outrage effect is simply the outrage variable picking up the role of delegation. We find that delegation is not significant in explaining performance, and our outrage result is robust.[21]

5.1.3.4 Relative income versus trustee measurement variables. Further possibilities for violations of the CFA exogeneity assumption arise if the residual of the common component of a measurement variable is correlated with *log(Compensation)*. Potential stories for such a correlation generally refer to either the relative income measurement variables or the trustee measurement variables. Thus, we reestimate the system using only one set of measurement variables (the relative income ones in panel F) and then the other (the trustee measurement variables in panel G). Recall that, as in an IV estimation, the only portion of compensation used in the performance estimation is that which is CFA-extracted outrage factor. Thus, the measurement variable set is core to the performance results. These measurement variables subgroup tests also help to alleviate omitted variables concerns that the outrage factor is simply spuriously correlated with the heterogeneity in the role of the investment manager in the cross-section, to the extent that the heterogeneity in the role of the investment manager is more likely to be correlated with the trustee measurement variables than with the income measurement variables.

In Table 7, panel F, we find that the *Outrage* induced *log(Compensation)* effect on *Portfolio Performance* has a significant and larger coefficient (0.00425 vs. 0.00300 from Table 6) when we extract *Outrage* using only the relative income variables. In panel G, we just use the trustee measurement variables, and find a significant coefficient for *log(Compensation)* that is 0.00270, only 10% lower than that of Table 6.

[21] We note that because of the incompleteness of reporting on delegation, we can't provide a more complete exploration of the delegation decision.

5.1.3.5 Politicization. Next, we address the potential concern that these measures taken collectively could be related to the degree of politicization of a fund. Andonov, Hochberg, and Rauh (2018) find a significant impact of trustee composition on pension fund performance in private equity, characterizing trustees into nine categories based on both who appointed them and their profession. Their headline result is that the greater the proportion of state officials as trustees, the lower private equity performance. We control for politicization in our estimation, but some mechanical relationship may exist because the sum of all trustees must add to unity. The clearest indication that our results derive from an alternative channel to theirs, is our Table 7, panel F, tests that shows our results carry through even when we exclude all trustee composition variables as measurement variables for outrage.

The trustee composition type closest to the state political category in Andonov, Hochberg, and Rauh (2018) is the category of budget civil servants (political and not), who serve in a finance or budgetary capacity. Although we have shown in Table 7 that this variable cannot be driving our results (because we can estimate the effect with only the relative income variables), we do not want to contaminate any future estimation. Thus, we remove this variable hereafter, although as our examples of CalSTRS showed, the variable captures an important measurement of outrage threat.[22]

5.1.3.6 Instrumental variables Our final robustness specification implements an IV estimation, using a GMM system of two equations, whereby we instrument *log(Compensation)* with the measurement variables. This IV estimation uses the variables as predictors together in a single equation (like any multi-instrument IV), thus using not just the intersection of their correlation with *log(Compensation)*, but the union of all correlations. This is not our preferred estimation technique for two reasons. First, we believe that our CFA approach bears stronger economic intuition, as it directly estimates the latent *Outrage* factor and it enables for the analysis of the effect of *Outrage* shocks, instead of measurement variable shocks. Second, although we think the exclusion restriction of the IV estimator is feasible, it is more difficult to prove. The exclusion restriction is that any correlation of the instruments with *Portfolio excess returns* must come through their correlation with *log(Compensation)*, recalling again that we adjust compensation to local finance salaries.

Nevertheless, to address a concern that the results might derive from the empirical methodological choice to use CFA as opposed to IV analysis, we repeat the analysis, but take an IV approach using a GMM system of two equations. This estimation offers a final piece of evidence that our results

[22] Table 9, panel A, reports the portfolio-level performance using only four measurement variables. We find that *log(Compensation)* explained by *Outrage* has a positive and significant effect on portfolio net returns, with a coefficient of 0.00368.

Table 9
Outrage, portfolio performance, and asset class performance

A. Portfolio	*Local income*	*Relative wages*	*Municipal workers*	*Teachers*	*Compensation*	Excess returns portfolio
Outrage (latent)	−0.510***	−0.0946***	0.0164*	0.0977*	−1	
	[0.0907]	[0.0359]	[0.00912]	[0.0539]	[constrained]	
log(Compensation) (Endog)						0.00368**
						[0.00156]
Political					0.368*	−0.00552
					[0.1956]	[0.0077]
Underfunding index (lag)					−0.101**	0.00327
					[0.0462]	[0.00228]
log(Size) (lag)					0.261***	−0.00152
					[0.0419]	[0.00224]
Observations	352					

B. Alternatives	*Local income*	*Relative wages*	*Municipal workers*	*Teachers*	*Compensation*	Excess returns alternatives
Outrage (latent)	−0.460***	−0.0726*	0.0161	0.0348**	−1	
	[0.139]	[0.0377]	[0.0103]	[0.0135]	[constrained]	
log(Compensation) (Endog)						0.00902*
						[0.00513]
Political					0.236	−0.0144
					[0.2380]	[0.0230]
Underfunding index (lag)					−0.126**	−0.00442
					[0.0534]	[0.00597]
log(Size) (lag)					0.245***	0.0000572
					[0.0664]	[0.00487]
Observations	260					

(*Continued*)

do not reflect any possible violations to our CFA exogeneity condition. As Table 8 shows, the IV specification produces similar results. *log(Compensation)* explained by *Outrage* has a positive and significant effect on portfolio net returns. The coefficient in this estimate is 0.00770. Our preferred CFA test (in Table 6) provides a more conservative estimate of the impact of outrage.

5.2 Within-asset-class performance results

As noted in the theoretical framework, our model predicts costs for funds that invest in risky asset classes that face a binding outrage constraint –such boards will underperform particularly in the risky asset classes. We thus gain further insight by now turning to tests at the asset-class level of the impact of outrage on performance asset class level. As a performance measure in these tests, we replace portfolio-wide net returns with net returns in alternatives (panel B of Table 9), public equities (panel C) and fixed income (panel D).[23]

[23] The number of observations varies by column because some public funds are not exposed to all of the asset classes, and some funds only report performance at the aggregate portfolio level. We do not report the first equation estimation for each column; they are materially the same as the estimations presented in the first column.

Table 9
(Continued)

C. Equities	*Local income*	*Relative wages*	*Municipal workers*	*Teachers*	*Compensation*	Excess returns equities
Outrage (latent)	−0.552***	−0.0818**	0.0124	0.119**	−1	
	[0.0922]	[0.0345]	[0.00998]	[0.0472]	[constrained]	
log(Compensation) (Endog)						0.00336**
						[0.00137]
Political					0.388**	−0.0135***
					[0.1954]	[0.0049]
Underfunding index (lag)					−0.0977**	0.00251
					[0.0402]	[0.00239]
log(Size) (lag)					0.289***	−0.00199
					[0.0414]	[0.00164]
Observations	302					

D. Fixed income	*Local income*	*Relative wages*	*Municipal workers*	*Teachers*	*Compensation*	Excess returns fixed income
Outrage (latent)	−0.541**	−0.0970*	0.0164	0.0539***	−1	
	[0.242]	[0.0589]	[0.0127]	[0.0207]	[constrained]	
log(Compensation) (Endog)						0.00191
						[0.00622]
Political					0.1998	0.0035
					[0.4020]	[0.0078]
Underfunding index (lag)					−0.00589	0.00683*
					[0.0425]	[0.00412]
log(Size) (lag)					0.273***	0.0019
					[0.0967]	[0.00313]
Observations	260					

This table reports the coefficients for the structural equation modeling (SEM) system of equations. The first five columns report the coefficients for the latent outrage (right-hand side) on each outrage measure (left-hand side). The sixth column uses the adjusted compensation in the left-hand side and constrains the coefficient of the outrage variable to −1 without loss of generality. The seventh column reports the effect of the adjusted compensation on excess portfolio performance over the benchmark. As measurement variables for outrage, we include mean pension workers' wages, median local area income, the percentage of trustees that are *Municipal workers*, the percentage that are *Teachers*, and the percentage that are *Budget civil servants*. *Political* is the average of two variables: a dummy taking a value of one if the chair is appointed by the government and the fraction of board members appointed by the government. The *Underfunded index* is constructed by taking the mean across the standardized value of one minus the funded ratio and age, following Andonov, Bauer, and Cremers (2017). The equations for *Local income*, *Relative wages*, and *log(Compensation)* include a time trend control. We estimate all parameters jointly using GMM, and all standard errors are calculated via bootstrap. $^{*}p<.1$; $^{**}p<.05$; $^{***}p<.01$.

5.2.1 Within-asset-class excess returns results. We find that *Outrage*-induced *log(Compensation)* positively and significantly predicts net returns in alternatives (coefficient of 0.00902) and equities (0.00336), consistent with the model's understanding of how outrage affects skill in realizing returns in risky investment. We find no effect for fixed income. In terms of economic impact, these results imply that a one-standard-deviation lower *Outrage*, passing through $302,429 higher level in compensation, implies 33-bps higher excess returns in *Alternatives* and 12-bps higher excess returns in *Equities*. These results suggest the intuitive inference that skill is most valued in the higher risk of the risky asset classes, consistent with theoretical understanding of active asset management (e.g., Ross 2005).

Regarding the other political agency variables, we find that *Political* has a negative and significant impact on excess returns in equities and a negative insignificant impact in alternatives, where we have less power due to more noisy returns and smaller number of observations. These results are consistent with a pay-to-play interpretation in riskier asset classes, and which has been shown in particular for private equity in Andonov, Hochberg, and Rauh (2018).[24] We find that a one standard devtaiotion increase in the *Political* variable reduces net perfomance in Alternatives by 41 bps. For the sake of comparison, Andonov, Hochberg, and Rauh (2018) find that a one-standard-deviation increase in the fraction of state-appointed board members reduces the IRR in private equity by around 1.05 percentage points.

The only significant impact for *Underfunded* is a positive effect of underfunding in fixed income asset choices. Underfunded pensions have higher performance relative to benchmarks in fixed income choices. When we instead analyze information ratio results in the next section, this result goes away, leading to an intuitive finding that that underfunded pensions take riskier positions within the fixed income asset class, reminiscent of "swing for the fences" in Ang, Chen, and Sundaresan (2013), "gambling for resurrection" in van Binsbergen and Brandt (2015), and the similar risk taking in the presence of underfunded liabilities results in Andonov, Bauer, and Cremers (2017).

These asset-class-level results also help to address concerns of unobserved differences across plans in the importance managers attach to underperforming socially responsible investments. A socially responsible agenda has had its greatest impact on allocations within equities, rather than on allocations within alternatives. The fact that we find similarly strong results at the alternative level suggests that an unobserved socially responsible agenda is not driving our results.

5.2.2 Information ratio results. Higher net returns do not necessarily reflect a higher realized Sharpe ratio if the net return performance arises from taking on increased risk in within-asset-class securities selection. The benchmarking returns that is used to gauge net performance is an ex ante objective provided by the board, but the enacted risk may be different, and this difference may reflect the agency issues studied herein.

Thus, in Table 10, we estimate the effect of *Outrage*-induced *log(Compensation)* on the *Information ratio* (excess returns divided by the realized static tracking error) instead of excess returns. Consistent with prior results, we find at the portfolio level that *Outrage*-induced *log(Compensation)* has a positive effect on the *Information Ratio*. An increase in *log(Compensation)* induced by a one-standard-deviation reduction in *Outrage* increases the overall *Information ratio* by 0.042 (a 19% increase from the average) and the equity *Information ratio* by 0.056 (a 28% increase from the average). The coefficient

[24] The lack of significance in alternatives may arise because our measure of alternatives returns is broad and includes real estate, hedge funds, and infrastructure, in addition to private equity.

Table 10
Effect of outrage on information ratio

A. Portfolio	*Local income*	*Relative wages*	*Municipal workers*	*Teachers*	*Compensation*	Information ratio portfolio
Outrage (latent)	−0.538***	−0.0864**	0.0148	0.0931***	−1	
	[0.0873]	[0.0349]	[0.00964]	[0.0247]	[constrained]	
log(Compensation) (Endog)						0.115**
						[0.0583]
Observations	337					
Number of funds	86					

B. Alternatives	*Local income*	*Relative wages*	*Municipal workers*	*Teachers*	*Compensation*	Information ratio alternatives
Outrage (latent)	−0.323***	−0.0692	0.0222	0.0534	−1	
	[0.105]	[0.0923]	[0.0168]	[0.0743]	[constrained]	
log(Compensation) (Endog)						0.106
						[0.123]
Observations	224					

C. Equities	*Local income*	*Relative wages*	*Municipal workers*	*Teachers*	*Compensation*	Information ratio equities
Outrage (latent)	−0.577***	−0.115	0.0111	0.0786**	−1	
	[0.159]	[0.0868]	[0.0427]	[0.0346]	[constrained]	
log(Compensation) (Endog)						0.152*
						[0.0920]
Observations	282					

D. Fixed Income	*Local income*	*Relative wages*	*Municipal workers*	*Teachers*	*Compensation*	Information ratio fixed income
Outrage (latent)	−0.395***	−0.0758	0.0225	0.0318	−1	
	[0.104]	[0.0676]	[0.0184]	[0.0242]	[constrained]	
log(Compensation) (Endog)						0.0114
						[0.122]
Observations	238					

This table reports the coefficients for the structural equation modeling (SEM) system of equations, with most variables defined similarly as in Table 9. The first four columns report the coefficients for the latent outrage (right-hand side) on each outrage measure (left-hand side). The fifth column uses the adjusted compensation in the left-hand side and constrains the coefficient of the outrage variable to -1 without loss of generality. The sixth column reports the effect of the adjusted compensation on the information ratio. The equations for *Local income*, *Relative wages*, and *log(Compensation)* include a time trend control. We estimate all parameters jointly using GMM, and all standard errors are calculated via bootstrap. $^{*}p<.1$; $^{**}p<.05$; $^{***}p<.01$.

estimates for alternatives and fixed income are positive, but statistically insignificant. Our results suggest that the increase in funds net returns induced by a relaxation of outrage does not come from an increase in within-asset-class risk exposure. Instead, it comes from an increase in the ratio of performance to risk, consistent with higher portfolio management skills.

To further explore the robustness of our net return estimations for alternatives and equities to the role of risk, we run the main specification of Table 9 with the inclusion of risk factors as controls. For equities, we follow standard practice and include the market, value, and size factors computed for developed countries, and accessed from Ken French's website. For alternatives, we employ an approach similar to Sharpe (1992) and use indexes as ad hoc factors for broad asset classes. More specifically, we include the returns of (a) S&P Listed Private Equity, (b) Dow Jones U.S. Real Estate, and (c) HFRI Equally Weighted. We chose these indexes because, to the best of our knowledge, they are the only indexes for the asset classes of private equity, real estate, and hedge funds with returns going back as early as 1995, when our sample starts.

Table 11 reports these results. We find the statistical and economic significance are nearly identical to the specification without these additional risk controls. In fact, for the public equities asset class, the risk controls are as insignificant, consistent with the chosen benchmarks capturing such dimensions of risk in portfolios.

Together, these results suggest that the increase in funds net returns induced by a relaxation of outrage comes from an increase in the ratio of performance to risk, consistent with higher portfolio management skills, rather than from an increase in within-asset-class risk exposure.[25]

5.2.3 Asset allocation results. Our theory suggests that a fund exposed to lower outrage can hire a manager with greater managerial skills, who will increase allocations to riskier asset classes. Table 12 tests for the impact of *Outrage*-induced *log(Compensation)* on asset allocation. Here we restrict our analysis to the observations for which we can observe the portfolio weights in all asset classes.

We find a positive and significant impact of *Outrage*-induced *log(Compensation)* on allocation to alternatives (panel A), with a coefficient of 0.0507, consistent with our theory that outrage affects the appetite for risk. In particular, a lower outrage factor implies that the pension may offer a higher pay to risk-taking for the investment manager. The result in alternatives shows that this appetite for risk is particularly strong in the riskiest asset class. What is interesting, in an adding-up of the asset class weights sense, is to see which asset class gives up weight to take on more risk. As panels B and C show, funds with lower outrage and higher compensation trade off more risk in alternatives for less risk in equities, with no effect on fixed income. Note that the negative, significant coefficient for *Outrage*-induced *log(Compensation)* is −0.0476, almost a complete offset to the positive coefficient for alternatives.

In terms of economic magnitudes, the pattern suggests that a pension exposed to a one-standard-deviation lower *Outrage*, passing through $302,429 increase in compensation, implies that portfolio weights shift by 1.85 percentage points toward alternatives and 1.74 percentage points away from equities. For the average pension fund in our sample, the translates into additional $825 million

[25] In further analysis in Internet Appendix Table 3, we address the possibility that our results could come from pension funds strategically choosing low-risk benchmarks that are easy to beat. We construct a synthetic benchmark of each fund, defined in terms of each fund's average weights in each asset class, so its exposure to different asset classes does not change over time. We then construct a benchmark information ratio, defined as the difference between the potentially manipulated benchmark return and the the synthetic benchmark return, all scaled by the static standard deviation of the benchmark net return (calculated using the entire time series available for each fund). We test whether the outrage-induced log compensation affects this benchmark information ratio and find an insignificant impact, consistent with the strategic selection of benchmarks not driving our results.

Table 11
Main performance results, while controlling for risk factors

A. Alternatives	*Local income*	*Relative wages*	*Municipal workers*	*Teachers*	*Compensation*	Excess returns portfolio
Outrage (latent)	−0.460***	−0.0899	0.0172*	0.0319*	−1	
	[0.0839]	[0.0548]	[0.0102]	[0.0165]	[constrained]	
log(Compensation) (Endog)						0.00907*
						[0.00536]
Political					0.122	−0.00574
					[0.107]	[0.0109]
Underfunding index (lag)					−0.132**	−0.000753
					[0.0618]	[0.00598]
log(Size) (lag)					0.244***	0.000418
					[0.0746]	[0.00756]
PE factor						−0.0657*
						[0.0387]
RE factor						−0.0788
						[0.0530]
HF factor						0.368***
						[0.126]
Observations	260					

B. Equities	*Local income*	*Relative wages*	*Municipal workers*	*Teachers*	*Compensation*	Excess returns alternatives
Outrage (latent)	−0.569***	−0.126***	0.0170*	0.0972***	−1	
	[0.0986]	[0.0444]	[0.00968]	[0.0353]	[constrained]	
log(Compensation) (Endog)						0.00322**
						[0.00139]
Political					0.0597	−0.00341
					[0.0999]	[0.00220]
Underfunding index (lag)					0.225***	−0.00158
					[0.0535]	[0.00189]
log(Size) (lag)					−0.0196	0.00257
					[0.0575]	[0.00293]
MKT factor						−0.00000822
						[0.00713]
SMB factor						0.00738
						[0.0221]
HML factor						0.0327
						[0.0221]
Observations	302					

This table reports the coefficients for the structural equation modeling (SEM) system of equations, with variables defined similarly as in Table 9. To capture risk, we include risk factors returns as controls in column 6. For alternatives, we include the returns of (a) S&P Listed Private Equity, (b) Dow Jones U.S. Real Estate, and (c) HFRI Equally Weighted. For equities, we follow standard practice and include the market, value, and size factors computed for developed countries, which we access from Ken French's website. The equations for *Local income*, *Relative wages*, and *log(Compensation)* include a time trend control. We estimate all parameters jointly using GMM, and all standard errors are calculated via bootstrap. $^{*}p<.1$; $^{**}p<.05$; $^{***}p<.01$.

in alternatives and $775 million away from public equities. These numbers may not be economically large relative to the size of the public pension funds in our sample, but suggest a new channel to think about performance and agency in pension management for future exploration.

5.3 Implied economic impacts of policies to insulate from outrage

In evaluating economic impacts, we have followed standard practice in CFA analysis and considered the impact of a one-standard-deviation change in

Table 12
Effect of outrage on asset class weights

A. Alternatives

	Local income	*Relative wages*	*Municipal workers*	*Teachers*	*Compensation*	Weight alternatives
Outrage (latent)	−0.501***	−0.0718	0.0199	0.0402**	−1	
	[0.0958]	[0.0495]	[0.0154]	[0.0197]	[constrained]	
log(Compensation) (Endog)						0.0507***
						[0.00975]
Observations	210					

B. Equities

	Local income	*Relative wages*	*Municipal workers*	*Teachers*	*Compensation*	Weight equities
Outrage (latent)	−0.541***	−0.0645	0.00876	0.0442**	−1	
	[0.133]	[0.0511]	[0.0189]	[0.0207]	[constrained]	
log(Compensation) (Endog)						−0.0476***
						[0.0133]
Observations	210					

C. Fixed income

	Local income	*Relative wages*	*Municipal workers*	*Teachers*	*Compensation*	Weight fixed income
Outrage (latent)	−0.535***	−0.0757	0.0154	0.0476*	−1	
	[0.177]	[0.0549]	[0.0189]	[0.0278]	[constrained]	
log(Compensation) (Endog)						−0.00225
						[0.00773]
Observations	210					

This table reports the coefficients for the structural equation modeling (SEM) system of equations, with most variables defines similarly as in Table 9. The first four columns report the coefficients for the latent outrage (right-hand side) on each outrage measure (left-hand side). The fifth column uses the adjusted compensation in the left-hand side and constrains the coefficient of the outrage variable to −1, without loss of generality. The last column reports the effect of the adjusted compensation on portfolio weights. The equations for *Local income*, *Relative wages*, and *log(Compensation)* include a time trend control. We estimate all parameters jointly using GMM, and all standard errors are calculated via bootstrap. *$p<.1$; **$p<.05$; ***$p<.01$.

outrage on performance, implicitly making a combination of changes in reference wages, and changes in the proportion of trustees on the board that are particularly sensitive to outrage. For policy purposes, here we provide a more *ad hoc* analysis, looking first in isolation at the unit sensitivity to the measurement variables for outrage, and then considering the potential impact of policy reforms.

We consider the linear projection of the *Outrage* variable into the measurement variables of outrage. This linear projection is given by

$$Outrage_{it} = \underset{(0.021)}{-1.538} \times LocalIncome_{it} \underset{(0.014)}{-0.168} \times RelativeWages_{it} \underset{(0.040)}{+0.153} \times MunicipalWorkers_{it} \underset{(0.024)}{+0.477} \times Teachers_{it} \underset{(0.036)}{+0.031} \times BudgetCivilServants_{it}. \tag{16}$$

The negative coefficient for *Local income* suggests beneficiaries living in poorer communities, where the gap between local wages and high wages for investment managers is likely to be largest, stand to lose in weaker pension plan performance. This makes clear the point mentioned in the introduction: outrage costs are not distributed equally, but fall harder on *Main Street* communities, which already exhibit lower local wages and greater income inequality. The negative coefficient for *Relative wages* suggests that the lower the income of beneficiaries, the more they stand to lose in pension plan performance as a result of outrage.

Examples of proposed governance reforms of pension plans include exclusively focusing on reducing political appointees on the board. We consider depoliticization in our context to be the replacement of replacing budget civil servants with trustees that have no more than average inequality aversion. Alternatively, one can consider broader reforms to move toward a skills-based board and away from a constituency-based board. With the average board having one budget civil servant, one teacher, and half of a municipal worker, removing these board members and replacing them with trustees with average level of inequality aversion would imply a 6% reduction in outrage, projecting to a 2-bps improvement in excess performance over the benchmark. These benefits would have to be considered along with greater challenges of representation and accountability arising from a skills-based board.

Aside from performance consequence, there may be additional implications if boards were able to be insulted from outrage concerns. As noted above, and shown in Table A.3 in the appendix, there is a positive relationship between outrage and the use of delegation (that had no impact on performance in Table 7), that could derive from more outrage for pay for internal than external managers. The estimates in Table A.3 in the appendix suggest that if reforms reduce outrage, the use of external managers would be reduced.

6. Conclusion

The paper introduces a model where trustees of public pension funds consider the threat of private costs from outrage arising from inequality aversion. This concern leads to an equilibrium with trustees hiring investment managers with lower skills, which in turn creates distortions in portfolio allocation and weaker performance in the risky asset classes.

We test these predictions using a hand-collected global panel data set that includes information on investment manager compensation and measurement variables for Outrage. We find that outrage impacts fund performance and hence beneficiary welfare. A one-standard-deviation reduction in *Outrage* would increase portfolio returns in excess of benchmark returns by 11 bps per year. For an average (median) fund this would translate into \$49 (\$15.5) million per year in greater value-added at the costs of an increase in compensation of \$302,429. The costs that the fear of outrage creates for pension performance are particularly important in areas where finance salaries are much larger than the average income of local residents. Such areas may be more readily prone to outrage, but also are areas in which the local wage earners have little slack to support faltering pension systems.

Although it is beyond our scope to consider all possible ways to insulate the board from outrage pressures, we have a few ideas. Modifying or clarifying risk and profit-sharing arrangements so that beneficiaries' expected benefits become more closely tied to the performance of the fund could increase salience to the importance of quality investment management. In this situation, it would be easier to garner support for governance reforms that could insulate plans from outrage, and deliver performance improvements. Our paper projects gains, but also limits, from an exclusive focus on depoliticization of trustees. Our paper projects more significant returns from broader governance reforms, such as a skills-based board, that would reduce all trustee types particularly sensitive to outrage.

Reducing the transparency of compensation arrangements also may be considered. While this is a crude way to insulate board members from outrage pressures, it is likely to be imperfect. Board members likely fear that compensation arrangements will be eventually released or leaked, leading to much of the same behavior. Therefore, as long as the board members remain exposed to outrage concerns, the same problems will emerge. The added advantage of transparency is that it also reduces the likelihood of pay-to-play arrangements and other political frictions.

Table A.1
Compensation outrage anecdotes in the media

1. Oregon	"Unspoken, but also politically inconvenient is the compensation to attract talent from the private sector. The state's existing investment officers are some of the best paid public employees, making an average of $200,000 a year. But Treasury officials quietly complain that staff is underpaid by industry standards, and bristle about having to explain and get approval from the Legislature to release performance-based pay each year." *Source*: Sickinger (2013).
2. CalPERS	"Our compensation is just too low," board member Richard Costigan said in May. "We're not attracting quality candidates. The quality candidates who want to come here are negatively impacted by the salary levels." *Source*: Ashton (2018).
3. Kentucky Retirement System	"We've got our issues here and it's hard enough attracting applicants," Thielen said, referencing KRS's status as one of the worst-funded pensions in the country. Thielen, who announced his intention to retire last year, has already had to stay on longer than planned due to a lack of qualified applicants for his position. . . . As for the provisions regarding fund personnel and their compensation, Thielen said the bill would "create significant problems for us attracting and retaining staff." While KRS links employee compensation to performance, the bill would require adoption of the government's tenure-based pay structure. *Source*: White (2016).
4. New York Teachers' Retirement System	Depoliticizing, professionalizing, and streamlining the management of our pension funds will enhance investment returns and reduce pension costs . . . . The proposal calls for the investment entity to be staffed by experienced industry professionals and for compensation packages to attract those investment professionals . . . . A Chief Investment Officer will lead the new investment management entity. *Source*: *Targeted News Service* (2011).
5. Missouri State Employees Retirement System	"Dahl, chief investment officer for the Missouri State Employees Retirement System, will receive a $125,155 cash bonus this summer and up to that amount in deferred compensation, payable in two years. In effect, he could double his $250,309 salary.... The payments, originally scheduled for February, are slated to go out in June, a delay designed to avoid public scrutiny amid legislative budget-cutting. It's a politically sticky subject, because Gov. Jay Nixon and legislators are considering cutting thousands of government jobs, services for the disabled and college scholarships among many other things. Senate Appropriations Committee Vice Chairman Kurt Schaefer, R-Columbia, was surprised Thursday to learn of the bonuses. "Now is not the time for anyone to be getting a state-funded bonus," said Schaefer . . . . Nixon, who last year called MOSERS bonuses "unconscionable," said Thursday that the bonus system is on the way out, thanks to his appointees to the board of trustees." *Source:* Young (2010).
6. Florida SBA	The Florida State Board of Administration (SBA) has bumped the annual paycheck of CIO Ash Williams up to $367,500 from $325,000. Williams, who oversees a team managing $176.4 billion in pension and endowment assets, has not had a pay raise since 2008, and in line with SBA rules, does not receive incentives, Dennis Mackee, a spokesman for the fund, told MMI. Public CIO compensation has been a hot-button topic in the industry. According to industry insiders, a freshly minted MBA graduate starts out in the private sector earning at least $300,000 a year. The typical public fund CIO earns about $200,000-350,000 annually. *Source*: Lim (2014).
7. New Mexico SIC	The New Mexico SIC has been in the market for a fixed-income director to oversee a $4 billion credit portfolio . . . "The council is seeking to find a qualified credit portfolio manager, which is difficult under the current budgetary constraints.... New Mexico's portfolio managers currently command approximately $100,000-120,000 in annual compensation. Market practitioners estimate that the state needs to offer at least $150,000 to fill the position . . . New Mexico's compensatory challenge highlights a tricky dance public funds must perform to persuade state legislatures to grant investment staff compensation levels that are higher than other public employees. "Pay scales in public plans tend to reflect the pay scales for the state bureaucracy. A public plan is looked at as just another state agency," said Charles Skorina of recruitment firm Skorina & Co., which specializes in recruiting for asset management firms and endowments and foundations. Asset management and E&F executives generally command two to four times more compensation than public pension peers in similar positions. *Source:* Lim (2013).

(*Continued*)

Table A.1
(Continued)

8. Qsuper, Australia	Brad Holzberger, chief investment officer of the $54 billion QSuper retirement fund was the highest paid executive in the not-for-profit superannuation sector last year, taking home $1.2 million. . . . Mark Delaney, who oversees the investment portfolio of the $78 billion AustralianSuper fund . . . . was paid $971,000. Ian Silk, the boss of AustralianSuper, the largest not-for-profit fund in the country, was paid $700,000. The salaries are modest compared with the remuneration packages of fund managers, whose services are bought by super funds. The highest paid executive director at Platinum Asset Management, which has $24 billion under management, is Philip Howard, the finance director, who was paid $3.6 million last year. Fund managers can earn up to $10 million a year. *Source:* Patten (2014).
9. Qsuper, Australia	Superannuation chiefs managing the nest eggs of Queensland public servants are receiving fat-cat bonuses while members are facing delays in getting advice. *Source:* Viellaris (2014).

This table presents nine anecdotes of media outrage concerning the compensation of public fund investment managers.

Table A.2
Variable definitions

Variable	Definition	Source
	Compensation, portfolio choice, and performance variables	
Investment manager compensation	The maximum compensation of the fund's investment managers, including CEOs and CIOs	Hand-collected from annual reports, public filings, newspapers, Freedom of Information requests
Portfolio allocation	Portfolio weights in each of three asset class—alternatives (real estate, private equity, hedge funds, infrastructure), public equity, and fixed income—as percentage of the total	Center for Retirement Research (CRR), CEM Benchmarking and annual reports
Return	Realized returns in each asset class and for the overall portfolio	Center for Retirement Research (CRR), CEM Benchmarking and annual reports
Benchmark return	We use benchmarks as reported by Boston College Centre for Retirement Research or CEM. Benchmarks are chosen by pension trustees. Most funds report for each asset class multiple subasset classes. The asset-class-level benchmark is a weighted sum of these sub-asset-class benchmarks with weights set at the beginning of the reporting period. CEM subjects the reported benchmarks to additional checks for validity. A visual inspection of this information indicates the benchmarks capture dimensions of risk differences across and within asset classes	Center for Retirement Research (CRR), CEM Benchmarking
Tracking error	A single observation by fund for each asset class and the portfolio, calculated as the time-series average of the squared residuals from a regression of the pension fund returns on the benchmark returns, with no constant	Center for Retirement Research (CRR), CEM Benchmarking and annual reports
Portfolio delegation	Fraction of assets managed via delegation in each asset class	CEM Benchmarking

(*Continued*)

Table A.2
(*Continued*)

Variable	Definition	Source
	Political agency variables	
Municipal workers	The fraction of trustees that are workers providing basic services to city residents, usually through city government	From annual reports. Professional designation based on biographies and web sources, such as LinkedIn
Teacher	The fraction of trustees that are workers providing basic services to teachers or education administrators	From annual reports. Professional designation based on biographies and web sources, such as LinkedIn
Budget civil servant	The fraction of trustees that are civil servant in finance service to the government	From annual reports. Professional designation based on biographies and web sources, such as LinkedIn
Regional income	Logarithm of the local household income within the smallest region available (MSAs for the United States)	Regional income reported by National statistical offices (Census Bureau in the United States)
Worker wage	Logarithm of the average wage of the constituents of the pension fund	Hand-collected from annual reports. If not reported, we estimate based on working employee contributions and reported contribution rates as a percentage of salary
Political board	A dummy variable for the chair being appointed by either government executives or ministries or serving in the role *ex officio* because of his or her executive government position	Collected from pension fund charters and annual reports
Underfunded index	The negative of the standardized funded ratio plus the standardized age variable	Center for Retirement Research (CRR), CEM Benchmarking, annual reports, funds' current and cached websites, direct requests to the funds

This table defines the main variables used in this paper and lists their data sources.

Table A.3
Outrage and delegation

	Fraction of delegated assets		
Outrage	0.290**	0.322**	0.387**
	[0.147]	[0.147]	[0.151]
Political		−0.254**	−0.189*
		[0.108]	[0.105]
Underfunding index (lag)		0.0256	0.0398
		[0.0560]	[0.0580]
log(Size) (lag)		−0.106*	−0.108*
		[0.0589]	[0.0613]
Year FE	No	No	Yes
Observations	843	802	802

In this table, we regress the fraction of assets managed through delegation on the estimated *Outrage* factor and on several controls. Political is the average between a dummy taking a value of one if the chair is appointed by the government and the fraction of board members that are appointed by the government. The *Underfunded index* is constructed by taking the mean across the standardized value of one minus the funded ratio and age, following Andonov, Bauer, and Cremers (2017). *log(Size)* is the log of the lagged fund AUM.

Appendix A. Model Solution

In this appendix, we prove that the optimal contract for the manager is indeed that provided in Equation (10).

A.1 Optimal Contract

First, we assume that the manager with skill s is hired, and then we calculate the optimal contact offered by the board of trustees.

We can clearly assume that $\boldsymbol{b}=(\mathbf{1}-\boldsymbol{a})\boldsymbol{\kappa}$, given that financial and political returns are perfectly exchangeable in our model, which implies that the board would always offer the same fraction of political and of financial returns to the portfolio manager.

To find the optimal value of the risk-sharing parameter a, note that the objective function of the portfolio manager, given in Equation (8), simplifies to

$$r_f+(1-a)\boldsymbol{w}^{\mathsf{T}}B(s)-\frac{1}{2}\lambda(1-a)^2\boldsymbol{w}^{\mathsf{T}}\Sigma\boldsymbol{w}, \tag{A.1}$$

where $\boldsymbol{w}$ is the vector of portfolio weights, Σ is the covariance matrix of returns, and $B(s)$ is the vector $B(s)=(s\varphi_{MV}, s\varphi_P+\kappa)^{\mathsf{T}}$.

The optimal response that maximizes (A.1) is given by

$$\boldsymbol{w}=(1-a)^{-1}\lambda^{-1}\Sigma^{-1}B(s). \tag{A.2}$$

Now, we can write the board's objective function, given in Equation (6), as follows:

$$r_f+\boldsymbol{w}^{\mathsf{T}}aB(s)-c-\frac{1}{2}\lambda_{board}a^2\boldsymbol{w}^{\mathsf{T}}\Sigma\boldsymbol{w}. \tag{A.3}$$

Let $\nu=\frac{a}{1-a}$. Basic algebra shows that (A.3) is proportional to

$$\nu-\frac{1}{2}\frac{\lambda_{board}}{\lambda}\nu^2. \tag{A.4}$$

We can then determine the υ that maximizes (A.4), which is $\nu=\frac{\lambda_{board}}{\lambda}$. This implies that the optimal a is given by

$$a^*=\frac{\lambda}{\lambda+\lambda_{board}}. \tag{A.5}$$

A.2 Optimal Manager Quality

By plugging the optimal contract, given in Equation (10), into the board objective function, given in Equation (6), we find the following indirect utility function:

$$V_{board}(s)=r_f+\frac{1}{2\bar{\lambda}}B(s)^{\mathsf{T}}\Sigma^{-1}B(s)-O(s), \tag{A.6}$$

where $\bar{\lambda}=(\lambda^{-1}+\lambda_{board}^{-1})^{-1}$. The underlying first order condition for the choice of the optimal managerial skill is

$$B(s)^{\mathsf{T}}\Sigma^{-1}\boldsymbol{\varphi}=O'(s), \tag{A.7}$$

where $\boldsymbol{\varphi}=(\varphi_{MV},\varphi_P)^{\mathsf{T}}$. It's easy to see that this implies the following condition on the marginal payment to managers:

$$\frac{\left(\sigma_P^2\varphi_{MV}^2-2\rho\sigma_P\sigma_{MV}\varphi_{MV}\varphi_P+\sigma_{MV}^2\varphi_P^2\right)s+\left(\sigma_{MV}^2\varphi_P-\rho\sigma_P\sigma_{MV}\varphi_{MV}\right)\kappa}{\bar{\lambda}\sigma_P^2\sigma_M V^2(1-\rho^2)} \tag{A.8}$$

Appendix B. Comparative Statics Computations

Table 1 lays out model predictions by showing the comparative statics of how manager skill, portfolio weights, and returns change in the model with changes in political agency variables. In this appendix, we derive these predictions.

First, we consider the case when the outrage constraint is not binding, and after that we compare the derivatives of the binding and not-binding cases.

B.1 Partial Derivatives of Manager Skill

If the outrage constraint is not binding, then the optimal manager skill s^* maximizes the ex ante utility function of the board $V_{board}(s)$, which can be written as:

$$V_{board}(s) = \frac{1}{2\bar{\lambda}} B(s)^{\mathsf{T}} \Sigma^{-1} B(s) - O(s), \tag{B.1}$$

where Σ is the covariance matrix of returns, $O(s)$ is the outside option for a manager with quality s, and $B(s)$ is a vector defined by $B(s) = (s\varphi_{MV}, s\varphi_P + \kappa)^{\mathsf{T}}$. It's easy to see that we can write the underlying first order condition as

$$\bar{\lambda}^{-1} \boldsymbol{\varphi}^{\mathsf{T}} \Sigma^{-1} \left[s\boldsymbol{\varphi} + \kappa \boldsymbol{e}_2 \right] = O'(s), \tag{B.2}$$

where $\varphi = (\varphi_{MV}, \varphi_P)^{\mathsf{T}}$ and $\boldsymbol{e}_2 = (0,1)^{\mathsf{T}}$. Differentiating (B.2) with respect to the political return κ, we get

$$\left[O''\left(s^*\right) - \bar{\lambda}^{-1} \boldsymbol{\varphi}^{\mathsf{T}} \Sigma^{-1} \boldsymbol{\varphi} \right] \frac{\partial s}{\partial \kappa} = \bar{\lambda}^{-1} \boldsymbol{\varphi}^{\mathsf{T}} \Sigma^{-1} \boldsymbol{e}_2. \tag{B.3}$$

The term $\left[O''(s^*) - \bar{\lambda}^{-1} \boldsymbol{\varphi}^{\mathsf{T}} \Sigma^{-1} \boldsymbol{\varphi} \right]$ is positive by the concavity of the objective function at the maximum, while the term $\left[\bar{\lambda}^{-1} \boldsymbol{\varphi}^{\mathsf{T}} \Sigma^{-1} \boldsymbol{e}_2 \right]$ is negative if the Sharpe ratio of the mean-variance efficient securities is sufficiently larger than the Sharpe ratio of the political assets. This implies that

$$\frac{\partial s}{\partial \boldsymbol{\kappa}} < 0. \tag{B.4}$$

Now, differentiating (B.2) with respect to the political return $\bar{\lambda}$, we get

$$\left[O''\left(s^*\right) - \bar{\lambda}^{-1} \boldsymbol{\varphi}^{\mathsf{T}} \Sigma^{-1} \boldsymbol{\varphi} \right] \frac{\partial s}{\partial \bar{\lambda}} = -\bar{\lambda}^{-1} O'(s). \tag{B.5}$$

The term $\left[O''(s^*) - \bar{\lambda}^{-1} \boldsymbol{\varphi}^{\mathsf{T}} \Sigma^{-1} \boldsymbol{\varphi} \right]$ is positive, while the term $\left[-\bar{\lambda}^{-1} O'(s) \right]$ is negative, which implies that

$$\frac{\partial s}{\partial \bar{\lambda}} < 0. \tag{B.6}$$

B.2 Partial Derivatives of Portfolio Weights

The vector of portfolio weights will be given by

$$\boldsymbol{w} = \bar{\lambda}^{-1} \Sigma^{-1} \left[s\boldsymbol{\varphi} + \kappa \boldsymbol{e}_2 \right]. \tag{B.7}$$

Differentiating (B.7) with respect to $\boldsymbol{\kappa}$, we get

$$\frac{\partial \boldsymbol{w}}{\partial \kappa} = \bar{\lambda}^{-1} \{\det(\Sigma)\}^{-1} \begin{bmatrix} \sigma_{MV} \sigma_P^2 \left(\frac{\varphi_{MV}}{\sigma_{MV}} - \frac{\varphi_P}{\sigma_P} \right) \frac{\partial s}{\partial \kappa} - \rho \sigma_{MV} \sigma_P \\ \sigma_{MV}^2 - \sigma_{MV}^2 \sigma_P \left(\frac{\varphi_{MV}}{\sigma_{MV}} - \frac{\varphi_P}{\sigma_P} \right) \frac{\partial s}{\partial \kappa} \end{bmatrix},$$

from which it follows that

$$\frac{\partial w_{MV}}{\partial \kappa} < 0, \quad \frac{\partial w_P}{\partial \kappa} > 0. \tag{B.8}$$

Similar algebra shows that (a) the investment in fixed income is increasing in the risk aversion and (b) the investment in the mean-variance efficient security is decreasing in the risk aversion.

B.3 Partial Derivatives of Returns

In our model, the asset class expected returns ($E[R_{MV}]$ and $E[R_P]$) are proportional to the manager skill s. This implies that

$$\frac{\partial E[R_{MV}]}{\partial \kappa}<0,\ \frac{\partial E[R_{MV}]}{\partial \bar{\lambda}}<0,\ \frac{\partial E[R_{MV}]}{\partial \kappa}<0,\ \frac{\partial E[R_{MV}]}{\partial \bar{\lambda}}<0. \tag{B.9}$$

Now, the total portfolio expected return ($E[R]$) will be given by

$$E[R]=r_f+s\boldsymbol{\varphi}^{\mathsf{T}}\boldsymbol{w}\ =r_f+s\bar{\lambda}^{-1}\boldsymbol{\varphi}^{\mathsf{T}}\Sigma^{-1}\left[s\boldsymbol{\varphi}+\kappa\boldsymbol{e}_2\right]=r_f+sO'(s). \tag{B.10}$$

Differentiating (B.10) with respect to $\bar{\lambda}$, we arrive at

$$\frac{\partial E[R^e]}{\partial \bar{\lambda}}=\frac{\partial s}{\partial \bar{\lambda}}O'(s)+s\frac{\partial s}{\partial \bar{\lambda}}O''(s)<0. \tag{B.11}$$

B.4 Partial Derivatives with Respect to Outrage

If s^{free} denotes the optimal manager skill without outrage constraints, then the optimal manager skill is given by

$$s=\min\left\{s^{free},s^{outrage}\right\}. \tag{B.12}$$

This implies that

$$\frac{\partial s}{\partial s^{outrage}}=\begin{cases}0 & \text{if}\ \ s<s^{outrage}\\ 1 & \text{if}\ \ s>s^{outrage}\end{cases}. \tag{B.13}$$

Therefore,

$$\frac{\partial E[R_{MV}]}{\partial s^{outrage}}=\frac{\partial E[R_{MV}]}{\partial s}\frac{\partial s}{\partial s^{outrage}}\begin{cases}0 & \text{if}\ \ s<s^{outrage}\\ 1 & \text{if}\ \ s>s^{outrage}\end{cases}. \tag{B.14}$$

$$\frac{\partial w_{MV}}{\partial s^{outrage}}=\frac{\partial w_{MV}}{\partial s}\frac{\partial s}{\partial s^{outrage}}\begin{cases}0 & \text{if}\ \ s<s^{outrage}\\ 1 & \text{if}\ \ s>s^{outrage}\end{cases}. \tag{B.15}$$

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