How do private equity fees vary across public pensions?*

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Abstract

We document large variation in net-of-fee performance across public pension funds investing in the same private equity fund. In aggregate, these differences imply that the pensions in our sample would have earned $44 billion more – equivalent to $8.50 more per $100 invested – had they each received the best observed terms in their respective funds. There are also large pension-effects in the sense that some pensions systematically pay more fees than others when investing in the same fund. With better terms, the 95th percentile pension would have earned $14.91 more per $100 invested compared to $1.12 for the 5th percentile pension. Pension characteristics such as commitment size, overall size, relationships with fund managers, and governance account for a modest amount of the pension effects, meaning similar pensions consistently pay different fees.

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1 Introduction

Over the last twenty years, state and local defined-benefit pensions have increasingly shifted capital out of traditional asset classes like fixed income and into private-market investment vehicles like private equity and venture capital (Ivashina and Lerner, 2018). This shift has been accompanied by an intense public policy debate over the fees charged by the general partners (GPs) of private-market funds. In response, pensions in states like California, Pennsylvania, and New Jersey have conducted lengthy internal audits of the fees they pay in private equity.[1] While investment costs in private markets are generally known to be large (Gompers and Lerner, 2010; Metrick and Yasuda, 2010; Phalippou et al., 2018), there is no systematic evidence on how they are determined, mainly because fees are privately negotiated, rarely observed, and often not even recorded.[2]

In this paper, we shed some light on the costs public pensions face when investing in private markets using detailed pension-level portfolio data from 1990 to 2018. We overcome the inherent data opacity issues by comparing the net-of-fee cash flows received by different pensions invested in the *same* private-market fund. In other words, we use within-fund variation in net-of-fee returns to assess the degree to which fees vary across investors in the same fund.[3] Figure 1 provides a simple way to visualize this variation in our data. To construct the plot, we compute the standard deviation of returns within each fund based on its latest available data, where returns are measured using the cumulative net-of-fee return on invested capital. The plot then shows the distribution of within-fund return volatility across all funds, broken down by fund vintage year. The figure clearly demonstrates that net-of-fee returns – and thus likely fees – vary across investors in the same fund. We henceforth use variation in net-of-fee returns to gauge within-fund fee variation.

To develop a sense of the economic magnitude of within-fund fee variation, we ask how much

[1] Indeed, CalPERS – the largest U.S. public pension fund with over $350 billion in assets – recently came under scrutiny when it admitted that it did not fully track all of the fees it had paid to the private equity firms in which it invests. CalPERS eventually disclosed that it paid about 700 basis points in annual all-in-costs for its private equity portfolio, which is orders of magnitude larger than the cost of passive investments in public equities.

[2] An investigation by the U.S. Securities and Exchange Commission in June 2020 identified widespread "deficiencies" in the private equity industry related to inadequate disclosure about investment costs.

[3] We define fees as management fees, performance fees, or any other costs that are borne by investors. Our approach to measuring within-fund fee variation is valid if gross-of-fee returns are relatively homogenous across investors in a fund, which we argue is likely to be true for the bulk of investments in our sample.

each pension would have potentially gained had it paid the lowest fees in each of its funds. We then aggregate these potential gains over our analysis sample, which covers $516 billion of investments made by 231 U.S. pension funds into 2,535 private-market funds. According to our estimates, public pensions would have earned $44 billion more on their investments – equivalent to $8.50 more per $100 invested – had each pension received the best ex-post fee contract in its respective funds. This can be naturally interpreted as capital that was redistributed to fund managers or other unobserved investors.[4] Importantly, this estimate is likely a lower bound since we do not observe all of the limited partners (LPs) in a given fund, meaning true fee dispersion and hence potential gains would be larger if other unobserved investors like private endowments or family offices pay lower fees than U.S. public pensions. We do not observe any obvious time trends in aggregate potential gains due to within-fund fee dispersion, though there is meaningful variation across sub-asset classes. In traditional buyout private equity, potential gains due to fee dispersion are $10.37 per $100 invested whereas in venture capital they are $4.86.

We then document the existence of large pension effects, meaning some pensions consistently earn higher net-of-fee returns compared to others when investing in the same fund. Put differently, pensions that underperform in one fund are also more likely to underperform in other funds. More formally, in a simple fixed-effects model, an F-test consistently rejects the null hypothesis of no pension effects. The standard deviation of pension effects on investment performance is over 500 basis points, highlighting the large impact that fees can have on pension-level investment returns. As an alternative way to quantify how differences in fees translate to performance, we compute potential gains from within-fund fee dispersion at the pension level. The 5th percentile pension could have earned $1.12 more per $100 invested in private markets with different contract terms, whereas the 95th percentile pension could have earned $14.91.

The pattern of strong pension effects indicates that some pensions have systematically paid higher fees than other pensions in their respective funds over the course of our sample. There are several economic reasons that fees could vary within the same fund. Broadly speaking, GPs may

[4]Our potential gain estimate assumes that funds generate enough surplus to support this alternative fee schedule. In Section 3.4.2, we argue why this is plausible and provide an alternative estimate that is guaranteed to be feasible.

give fee breaks to investors who lower the cost of raising and managing a fund. For example, some investors may reduce the time it takes to raise capital by sending a positive signal to other investors about the GP's skill, or some investors may be less costly in terms of reporting requirements. Investors could also differ in their willingness to pay fees due to search costs, beliefs and preferences, or financial sophistication. We explore several different mechanisms that might generate pension effects by mapping relative within-fund performance to pension characteristics like size, the investment share in the fund, and the stage at which the pension committed capital to the fund. We find evidence that larger pensions with stronger ties to the GP of a fund tend to outperform other investors in the same fund. These results are consistent with the idea that securing capital from these pensions creates value for GPs, possibly through positive signaling effects, and some of this value flows to these pensions through lower fees. Pensions that have more member representation on their boards also appear to pay lower fees, perhaps because more member representation lowers agency frictions at public pensions (Andonov et al., 2018).

More importantly though, pension characteristics account for only a modest amount of the total pension effects that we find in the data. Using our baseline fixed-effects model, we still find strong statistical evidence of pension effects even after controlling for a wide range of attributes. Strikingly, the distribution of pension effects is virtually identically whether we control for observable characteristics or not. In addition, controlling for pension characteristics does little to change which pensions stand to gain and how much they would gain if they paid similar fees as the best performing investors in our sample. This suggests that two seemingly similar pensions that commit capital at the same time still consistently pay different fees when investing in the same fund.

We then explore a possible mechanism through which fee structures vary in practice. We first provide evidence that investors in the same fund are grouped into tiers in terms of their fee structures, with most funds having two or three tiers of investors. Fee structures in private-market vehicles typically feature a fixed annual management fee and a variable fee (carry) that is based on performance, meaning investors in different tiers may differ along one or both dimensions.[5]

[5]In practice, a fund's true fee structure has more complexities than a simple management or performance fee, but any special fee arrangement can be decomposed naturally into a fixed and variable part.

While we cannot observe the fee schedule for any investor-fund pair, we use our data to infer the average within-fund *range* of carry rates. Intuitively, within-fund differences in carry across investors should be easier to detect for more profitable funds. We indeed find a strong relationship between fund performance and the range of net-of-fee returns within a fund: after controlling for age, within-fund return dispersion is insensitive to fund performance for unprofitable funds and is linearly increasing in performance for profitable funds. Our estimates suggest that carry rates differ by around 10-20 percentage points across investors in the average fund. Thus, carry appears to be an important dimension of how overall fees vary across investors.

In the last part of the paper, we discuss several explanations for our findings. Perhaps the simplest is that observed within-fund variation in net-of-fee returns is not driven by fees, but is instead due to measurement error or bespoke investment structures that may exist for some investors in a fund (e.g., co-investment). To assess the possibility of measurement error, we directly filed Freedom of Information Act (FOIA) requests for a subsample of pensions in our data and found virtually no measurement error in this audit. Moreover, measurement error should bias us against finding strong pension effects, yet we confirm that they are still present when measuring returns using only realized cash flows. We also perform several quality checks of the data and argue that any bias from returns on bespoke structures like co-investment are likely to be minimal. In addition, we find similar amounts of within-fund fee variation in a sample where these structures are rarely employed. Fees therefore appear to be the primary source of within-fund return variation on which our analysis is based.[6]

Investors in the same fund may rationally agree to pay different fees because they differ in their information about manager skill, meaning they have heterogenous expectations about the gross return of the fund. A profit-maximizing GP would then optimally elicit these expectational differences and charge different fees accordingly, perhaps by offering a menu of fees. For rational beliefs to drive our results, some pensions must have rationally chosen fee structures that led to consistently lower returns over several private equity cycles. As discussed above, a related

[6]See Section 3.4 and Section 6.2 for an extensive discussion of these issues.

explanation for fee dispersion is that GPs offer fee breaks to more informed investors in order to attract less informed investors into the fund. This logic extends to the case when there are LP-GP specific synergies. However, while size and relationships do correlate with lower fees, the fact that there are large pension effects after controlling for such characteristics means that they do not fully account for any informational edges, signaling effects, LP-GP synergies or investor search costs that might cause fees to vary within funds.[7]

Another potential interpretation of our findings is that optimization frictions lead some pensions to consistently pay more fees than others even when investing in the same fund. Optimization frictions might arise from agency frictions or a lack of financial sophistication expressed in biased beliefs about gross fund returns or a failure to fully internalize the cost structure of private market investments. These frictions are inherently difficult to identify empirically, though we do have suggestive evidence in this direction: less than 5% of pension investors in our sample have any mention of performance fees or carry on their annual report, despite the fact that we find differences in carry to be an important component of price dispersion. Moreover, there is a growing body of evidence that frictions in labor markets and political considerations distort public pension investment decisions (Dyck et al., 2018; Andonov et al., 2018). Overall, it is hard to imagine that optimization frictions of this kind play no role in explaining why pensions with similar characteristics – and therefore those that in principle should have similar outside options when bargaining, similar preferences, and information – appear to systematically pay different fees when investing in the same private-market fund.

Related Literature There is an active public policy debate about the extent to which investing in private markets enhances the welfare of public pension beneficiaries, who are typically teachers, police, firefighters, and other public servants. A fundamental issue in this debate is how any value that is created by these investment vehicles gets split between investors and investment man-

[7] An analogous argument applies to differences in preferences. Fee dispersion could also arise if the marginal cost of a GP partnering with some pensions is higher than for others. Still, given that we control for several attributes and U.S. public pensions have relatively homogenous reporting requirements, the pension effects that we document seem unlikely to be explained by differences in marginal costs across investors.

agement firms. This is ultimately a question of how fees are determined and whether GPs price discriminate.[8] One way that we contribute to this debate relates to measurement, as the specific contracts between GPs and investors are essentially unobservable. Thus, we view our approach to estimating within-fund differences in fees as a first step in understanding how fees are determined for pensions when they invest in private markets. Our finding of large pension effects implies price dispersion in this setting has important distributional implications for pensions as well.

The fact that pensions systematically pay different fees when investing in the same fund may not be surprising to many, as price dispersion – often a result of price discrimination – is among the first concepts taught in introductory microeconomics courses. Price dispersion is a ubiquitous phenomena, occurring in automobile sales, airline tickets, mortgage markets and the mutual fund industry (e.g., Knetter, 1989; Goldberg, 1996; Allen et al., 2019; Hortaçsu and Syverson, 2004). In the context of private markets, co-investment and other special purposes vehicles have emerged in the last few years as an imperfect way for the general partners of private market funds to differentiate among their investors (Lerner, Mao, Schoar, and Zhang, 2018; Fang, Ivashina, and Lerner, 2015; Braun, Jenkinson, and Schemmerl, 2019). These investment arrangements offer select investors exposure to a different asset mix than the so-called main fund, typically at a reduced cost. Relative to prior work on these alternative fund structures, we argue that price dispersion also occurs more directly through fee contracts in the main fund. The observation that some pensions consistently receive better terms is also consistent with Lerner et al. (2018), who show that GPs offer certain special purposes vehicles only to a select set of investors.

What is perhaps more surprising is that differences in willingness to pay across pensions appear systematic and largely unexplained by easily-observable pension characteristics.[9] This fact is somewhat puzzling, as one might expect that rationally behaving pensions with similar attributes (e.g., size or experience) should should in principle have similar information, expertise, and bar-

[8] There is a closely related debate on the value proposition of private markets. Several studies find that private equity outperforms public equities (Harris et al., 2014; Robinson and Sensoy, 2016; Kaplan and Schoar, 2005), whereas others argue that risk-adjusted returns are zero or negative (Sorensen et al., 2014; Gupta and Van Nieuwerburgh, 2019).

[9] Because pensions with similar attributes are likely to have comparable marginal costs from the perspective of GPs, it seems natural to attribute the price dispersion that we observe to cross-pension differences in willingness to pay.

gaining power. In turn, they should pay similar same fees when investing in the same fund. While it is difficult to unequivocally prove that pensions are not fully optimizing, the notion that they fail to do so on behalf of their beneficiaries is consistent with prior research on agency and labor market frictions at public pensions (Andonov et al., 2017, 2018; Dyck et al., 2018).

Our results also inform theories of fee determination for investors in the same fund. In the benchmark model of Berk and Green (2004), within-fund fee dispersion is nonexistent because investors are assumed to be homogenous in every dimension. In Bernstein and Winter (2012), contracts may vary across investors if some create positive externalities for the fund, perhaps through signaling effects. Our evidence suggests that these externalities must be mostly unrelated to size and several other pension characteristics in order to explain outcomes in private markets.

The paper is organized as follows. In Section 2, we provide background about the data and present a simple accounting framework through which we interpret our results. Section 3 documents how large potential gains from within-fund fee dispersion are in aggregate. We also discuss other potential sources for within-fund variation in net-of-fee returns and conduct several robustness tests to ensure that fees are the primary source of this variation. In Section 4, we show that some pensions systematically, that is across funds, pay higher fees than others when investing in the same fund and measure the extent to which these pension effects can be explained by observables. Section 6.2 evaluates several interpretations for our results and concludes. Additional details and results are available in an online appendix.

2 Data and Empirical Framework

2.1 Institutional details

The focus of this study is public pension investment into private market vehicles, namely private equity (PE). A typical PE fund has two types of investors, the general partner (GP) and the limited partners (LPs). The GP is responsible for investing the fund's total pool of capital and usually contributes about 1 to 5% of their own capital in a fund. Thus, the bulk of the fund's capital

comes from LPs, who are entities like pensions, endowments, and family offices. At the beginning of a fund's lifecycle, GPs secure capital commitments from LPs, after which capital is formally "called" from the LPs and invested by the GP. These investments are then held for several years before they are liquidated and distributed back to the LPs. From start to finish, most funds have a total lifespan of ten to fifteen years.

A legal contract known as a limited partnership agreement (LPA) governs the specific investment terms in a fund, including how the GP will charge fees and allocate costs across LPs in the fund. LPAs are privately negotiated between each GP and LP and are typically not observed by other LPs. The usual fee contract in private market vehicles revolves around two main components: a management fee and a performance fee, the latter of which is also known as carried interest. The standard contract has an annual management fee of 1.5-2% based on the LP's committed capital and a 15-30% performance fee. The performance fees component is typically charged only after the fund has achieved a "preferred" return or hurdle rate of at least 5-8%. After that, during the "the catch up period," the GP gets any positive distributions until it realizes its carried interest (15-30%) on the cumulative distributions up to that point. Any dollar distributions thereafter are split according to the terms set forth in the LPA. The LPA between the LP and the GP also has several other provisions that determine the total cost born by LPs, including but not limited to legal fees, travel expenses, annual meeting expenses, and board member fees.

Outside observers are rarely privy to the contents of individual LPAs, though a recent 2019 industry survey of GPs does provide some information about the characteristics of these legal contracts.[10] The report details various ways in which different LPs might negotiate different limited partnership agreements within the same fund. For example, nearly half of GPs self-reported that a subset of LPs had amendments to their LPAs that gave them more favorable terms than other investors. The report also notes that its survey statistics likely understate the popularity of these so-called side letters. Conditional on having a side letter with one investor, 60% of GPs report that they offer some LPs a "most favored nation" clause that allows them to observe the LPAs of

[10] See the PE/VC Partnership Agreements Study 2018/2019 conducted by Buyouts Insider.

other investors and choose what they deem to be the best terms. Moreover, roughly one-quarter of the surveyed GPs reported that they charge different LPs different management fees. The report also highlights several important ways in which the calculation of carry may differ across investors or funds (e.g, deal-by-deal carry versus fund-level carry). In many ways, the overarching goal of this study is to quantify how these sorts of contractual differences across LPs ultimately impact investment performance.

2.2 Empirical Framework

We structure our empirical analysis around a general set of accounting identities that relates gross and net-of-fee returns for each investor p in fund f. Let γ_{pft} denote the cumulative gross return at time t for investor p in fund f. We further decompose γ_{pft} into a fund-wide and investor-specific component:

$$\gamma_{pft} = g_{ft} + \varepsilon_{pft}$$

where g_{ft} is the cumulative gross return on fund f at time t that is common to all investors. ε_{pft} is any component of investor p's gross return that is not shared by other investors. In practice, ε_{pft} might reflect co-investment arrangements or specific restrictions that investor p puts on fund f in terms of its investment mandate (e.g., ESG).[11]

Next, define c_{pft} as the cumulative cost at time t that pension p must pay to the GP of fund f. We will define specific elements of the cost structure in later sections, but for now it is useful to think of it as encompassing any contractual feature of the limited partnership agreement that puts a wedge between investor p's gross return and its net-of-fee return. Thus, we define investor p's cumulative net-of-fee return return at time t as follows:

$$\begin{aligned} r_{pft} &= \gamma_{pft} - c_{pft} \\ &= g_{ft} + \varepsilon_{pft} - c_{pft} \end{aligned}$$

[11] For example, some investors might have specific restrictions on investment into tobacco or oil companies. The GP then excludes the non-conforming investments from the portfolio of those specific LPs.

Our identifying assumption for most of the analysis in the paper is that $\varepsilon_{pft} \approx 0$ for all of the pensions that we observe in fund f. This is not to say that ε_{pft} is truly zero, but rather that the net-of-fee returns that we observe are not primarily driven by deviations of investor p's gross performance from that of the other investors. This assumption would be violated if, for instance, cash flows from co-investment are bundled with our data on net-of-fee returns. We discuss when the assumption that $\varepsilon_{pft} \approx 0$ is more or less likely to hold extensively in Section 3.4.3.

Under the assumption that $\varepsilon_{pft} \approx 0$, the net-of-fee return for investor p is

$$r_{pft} = g_{ft} - c_{pft}$$

Furthermore, if we define $\sigma_f(\cdot)$ as the standard deviation operator across investors in fund f, then within-fund variation in net-of-fee returns at any given point in time is driven by variation in costs across investors:

$$\sigma_f(r_{pft}) = \sigma_f(c_{pft})$$

In other words, assuming that all investors that we observe have the same before-cost exposure to fund f, then any variation in net-of-fee returns across investors must be driven by variation in the costs borne by investors in fund f. For this reason, we refer to within-fund return dispersion and within-fund fee (or cost) dispersion interchangeably. In a broad sense, the goal of the paper is to measure the extent of within-fund fee variation and understand its underlying sources. The simple accounting framework laid out above implies that we can do so by studying the net-of-fee performance of different investors in the same fund at a given point in time. After we document variation in r_{pft} across investors, we then analyze how standard contractual features in private markets may give rise to different c_{pft}.

2.3 Data Description

We obtain data from Preqin, a data provider that specializes in alternative assets markets. Preqin's data on private market investments is sourced primarily from legally-required annual reports, Free-

dom of Information Acts (FOIA) requests, and direct contact with investors. The data covers funds from vintage year 1990 onward and contains cash-flow data on LP-level investment into individual funds. The vast majority of investors in our data are U.S. public pension funds (83%) and UK public pension funds (7%). Other investor types in our dataset include public university endowments, government agencies, insurance companies, foundations, and private sector pensions. Throughout the paper, we only use data on U.S. public pensions. The unit of observation is investor-fund-date. We observe the amount of committed capital by the investor in the fund, the amount of capital that has been "called" from the investor (i.e., actual contribution amounts), and the amount of capital that has been distributed back to the investor by the fund. These variables are all reported in cumulative terms, and importantly, are reported *net of all fees*, including both management and performance fees. We also observe the net asset value of each investor's current investments in the fund. For a given investor in a fund, the net asset value reflects the market value of investments that have not been liquidated yet. Together, the contributions, distributions, and remaining net asset values allow us to calculate standard industry variables such as the internal rate of return (which pensions also directly report) as well as the total return multiple on invested capital, defined formally as:

$$r_{pft} \equiv \frac{\text{Market Value}_{pft} + \text{Cumulative Distribution}_{pft}}{\text{Invested Capital}_{pft}}.$$

In practice, this measure of returns is frequently referred to as TVPI or the multiple on invested capital (MOIC) – it is simply the total return received by p from investing in fund f, unadjusted for the timing of cash flows.

The fact that Preqin sources its data primarily through public pensions may cause selection bias in the types of private-market funds that we observe. For example, public pensions may not have access to the same set of private-market funds in which endowments invest. While selection of this kind might bias estimates of, say, private equity returns as an asset class, it should not influence our analysis of *within-fund* variation in returns across public pensions. Nonetheless, because our study is based primarily on within-fund variation in returns, we have taken several steps to ensure the quality of the Preqin data. Perhaps most importantly, we filed direct FOIA requests for data from a

subset of pensions in the Preqin sample. Reassuringly, in the vast majority of cases, the return data from our direct FOIA requests perfectly matches the data in Preqin. The online appendix contains much more detail on the results of this audit.

The online appendix also reports several additional quality checks that we perform on the Preqin data, though we highlight the main ones here. We drop any observations where the vintage of the fund is missing, the fund vintage is before 1990, or the data to compute TVPI is not available. To be conservative we drop co-investments from the data, as co-investments have a different fee and investment structure than a standard fund. We discuss the issue of co-investment and how it might impact our results later in Section 5.2. To avoid any potential accounting issues with younger funds, we also drop observations where the year of the report date is less than one year after the fund's reported close. Given our focus is on U.S. public pensions, we only include observations from U.S. based investors where TVPI is computed in dollars.[12]

We further eliminate all observations within a fund-quarter where the range of TVPI across investors is implausibly large. Our determination of whether a within-fund-quarter TVPI range is implausibly large is based on two standard contractual features of fees in private markets. Building on the accounting framework from Section 2.2, we approximate the cumulative costs borne by investor p in fund f at time t as:

$$c_{pft} \approx m_{pf} \times t + \kappa_{pf} \times \max(g_{ft} - 1, 0) \tag{1}$$

where m_{pf} is the annual management fee paid by investor p to fund f and κ_{pf} is the carry rate. As discussed in Section 2.1, management fees and carry (or profit-sharing) are the two building blocks of fees in private markets. Management fees are paid as a fixed percentage of assets and carry is paid only when fund f is profitable, the latter of which is reflected in the second term above.[13]

[12]Most U.S. institutions report cash flows in dollars, even if the fund that they are invested in is denominated in another currency. There are a few UK and Canadian institutions that report the cash flows of a subset of their funds in dollars. We could potentially expand our investor universe if we allow non-dollar denominated cash flows and convert them into dollars, though this would require additional assumptions on the appropriate timing of exchange rates.

[13]In practice, the way that carry is charged to investors is more complicated than our simple model. For example, it is common that carry is only charged after returns clear a preferred hurdle rate.

With this approximate cost structure in the background, we assume that the maximum *range* of carry rates within a given fund is c_f^{range} and the maximum range in management is m_f^{range}. At each point in time, we then define the maximum allowable range of TVPI as

$$\text{Allowable Range of } TVPI_{ft} = m_f^{range} \times t + c_f^{range} \times \max\left[TVPI_{ft}^{max} - 1, 0\right] \tag{2}$$

where $TVPI_{ft}^{max}$ is the maximum observed TVPI in the fund as of time t. Our bounds on allowable TVPI range are motivated by Equation (1). Specifically, we set the bounds assuming that one investor gets the lowest combination of management and carry and another investor gets the highest combination. Because we do not observe the gross-of-fee return for fund f in our data, we proxy for it using the maximum observed TVPI. Under these assumptions, it is straightforward to show that the range of TPVI within a fund at each point in time is defined as in Equation (2). Despite its limitations, we prefer this economically-driven approach to data screening compared to dropping fund-quarters where the range of TVPI exceeds a fixed threshold. This is because older and more profitable funds should naturally have wider ranges of TVPI than younger, less profitable ones. When implementing our filtering procedure, we set $m_f^{range} = 0.03$ and $c_f^{range} = 0.3$. We also drop fund-quarter tuples where the within-fund range of DVPI, defined as the cumulative distributions divided by invested capital, exceeds the allowable range. Overall, this procedure drops about 7% of distinct fund-quarters in the data, leaving us with a total of just over 300,000 pension-fund-quarter observations. Henceforth, we refer to this as the *master sample*.

Given that our eventual focus will be on within-fund return dispersion, we also create what we call the *core sample* by condensing the master sample as follows. First, we eliminate from the complete dataset any fund-quarter that does not have at least two investors, as this is a necessary requirement to compute within-fund dispersion. Then, for each fund, we choose the quarter with the largest number of investors reporting returns, breaking ties with the latest date. We break ties by taking the latest possible report date because it allows for any differences in management and carry within the fund to play out over the longest possible horizon. By construction, the core

sample is unique at the investor-fund (p, f) level.

2.4 Summary Statistics

Table 1 presents summary statistics for the core sample. There are 231 unique pension funds investing in 2,535 different funds in the core sample. There are 916 general parters (GPs), implying that the average GP has about 3 different funds. Of the 2,535 funds in the core sample, 42% are private equity, 23% are venture capital, 21% are real estate, 11% are private debt, and the remaining are either infrastructure or multi-strategy.

The median fund in the core sample is 7 years old and has 3 different investors. The median investment size is about 30 million dollars, which accounts for about 3 percent of the median fund. The median overall size of the pensions in our data is just under of 10 billion, though we see some extremely large funds that have over 300 billion in assets. Our net-of-fee return measure TVPI shows that the median multiple on invested capital is about 1.3, meaning that investors receive about \$1.3 for each \$1 invested.[14]

3 How large is within-fund dispersion performance?

We begin by measuring within-fund dispersion in net-of-fee returns and assessing how large this dispersion is in aggregate. As our baseline measure of dispersion, we compute how much better off investors would have been had they each earned the best net-of-fee return in their respective funds. We document that the largest potential investor gains are in private equity funds and the smallest potential gains are in venture capital funds. In aggregate, U.S. public pension funds in our sample would have \$8.50 more per \$100 dollar invested if they had received the best returns in the funds in which they invest. Based on the observed investment amounts, this amounts to roughly

[14]It is important to note that TVPI does not account for differences in investment horizon across fund (e.g., older versus newer funds), meaning some of cross-fund variation in TVPI is driven by these age differences. In our analysis in Sections 3 and 4 we control for age differences. In addition, the maximum within-fund range of 1.12 occurred in a 1990s vintage fund whose total return on capital was around 8. Thus, for a fund of that age and profitability, even small differences in contract terms across investors can generate large differences in ex-post performance.

$44 billion dollars of surplus that is either redistributed to other investors or the general partners of private-market vehicles. After presenting our baseline results, we develop a lower bound on the redistribution that occurs due to within-fund performance dispersion and explore the robustness of our potential gain estimates.

3.1 Measuring fee dispersion

3.1.1 Illustrative Example

Because we do not observe the actual contract terms between LPs and GPs (pensions and investment management firms), we instead use ex-post dispersion in net-of-fee returns to detect differences in fees. As discussed in our motivating framework from Section 2.2, this is a reasonable approach under the assumption that investors in the same fund have the same gross return. Intuitively, if two investors in the same fund have the same return before costs, then any difference in their net-of-fee return must arise because they face different costs.

When operationalizing this logic in the data, we use TVPI as our main measure of returns. We prefer this simple measure of returns over internal rates of return (IRRs) because the latter is easier to manipulate. For instance, a recent paper by Andonov et al. (2018) that also uses data from Preqin finds that some pensions selectively omit intermediate cash flows when returns are poor, which makes computing accurate IRRs more challenging. And, in our inspection of the data, the self-reported IRRs often have other issues; for example, it is often the case that reported IRRs are not annualized during the early years of a fund's life, but are annualized in later years. Compared to IRRs, TVPIs are simpler and harder to distort. Of course, comparing TVPIs across investors or a subset of investors in the same fund implicitly assumes that cash flows occur at the same time for that subset. This seems more plausible in the context of private-market investment because of the typical fund structure. With that said, we do our best to explicitly account for the timing of cash flows in our subsequent analysis.

Figure 2 highlights the intuition of our approach using actual data. Panel A of the the figure shows how TVPI evolves for one of the funds in our data. Our data-sharing agreement with Preqin

prevents us from revealing the identity of the fund, so for this example we will refer to it as "Fund I". There are over 20 different investors in Fund I, though for readability we only show TVPIs for three investors, who we will call investors *A*, *B*, and *C*. These three investors have had very different experiences in Fund I. Fifteen years from the fund's close data, Investor *A* has earned $2.45 per dollar invested, *B* has earned $2.59, and *C* has earned $2.82. In other words, despite investing in the same fund, Investor *C* has outperformed *A* by 37 percentage points over the life of the fund. Moreover, at each point in time, Investor *C*'s TVPI in Fund I has always exceeded that of *A*'s. The patterns naturally imply that *C* has a fee contract that dominates *A* in Fund I. We also show the evolution of DVPI for this fund, which is the ratio of cash distributions over invested capital. DVPI does not depend on fund-value reporting standards across pensions or differing fund value estimates. It is simply the ratio of the cumulative cash flow received to the cumulative cash invested. It is clear that investors who outperform in TVPI terms also do so in terms of DVPI.

The preceding example focused on a relatively high–performing fund (e.g., the lowest TVPI is well above 2). In Panel B of Figure 2, we show how TVPI and DVPI evolve for a lower-performing fund in our data ("Fund II"). Fund II has 11 different investors and we focus on two (Investors *X* and *Y*) for the plot to again make it easier to read. After 11 years in Fund II, Investor X has earned $1.43 for each dollar invested whereas Investor Y has earned $1.72. So *Y* has outperformed *X* by 29 percentage points. For the first 18 quarters after close, both investors had nearly identical TVPIs. After that, *Y* consistently outperformed *X* throughout the lifetime of the fund. Thus, Investor *Y* appears to have gotten a better fee contract in Fund II than Investor *X*.[15] As in the previous example, the same investor that does well in terms of TVPI also does well in terms of DVPI, the latter of which reflects actual cash distributions. Overall, these examples illustrate how we use dispersion in within-fund performance to learn about within-fund differences in costs across investors.

[15] In the online appendix, we show that there are similar patterns for both funds when looking at DVPI, which measures returns only using distributed cash flows.

3.1.2 Potential Gains due to Fee Dispersion

Building on the preceding logic, within each fund f we define investor p's potential gain at time t, denoted s_{pft}, as:

$$s_{pft} = r_{ft}^{max} - r_{pft} \tag{3}$$

where r_{ft}^{max} is the maximum TVPI earned by investors in fund f at time t. Intuitively, this is the incremental return that investor p would have earned in fund f if it had received the best return in the fund. For example, suppose that after 10 years pension fund A earned a TVPI of 1.5 in fund f. Further suppose that this is the maximum observed TVPI for the fund. If pension fund B earned a TVPI of 1.4 over the same horizon, then pension fund B's potential gain in fund f would be 0.1. Next, we convert the net-of-fee return gains to dollars by multiplying it by the amount that pension p invested in fund f. Let a_{pft} be the amount that p invested in fund f as of time t. Then,

$$d_{pft} = a_{pft} \times s_{pft} \tag{4}$$

is pension p's potential dollar gain at time t under the counterfactual. Continuing with our previous example, if pension fund B invested \$10 million in fund f, then its potential dollar gain would be \$1 million. This metric implicitly assumes that each fund f generates enough surplus to support this counterfactual fee schedule. We discuss the plausibility of this assumption, as well as other counterfactual exercises in Section 3.4.

Comparison Group for Computing Potential Gains To compute the gain measures s_{pft} and d_{pft} in the data, we need to be precise about the reference group that we use to compute r_{ft}^{max}. In several of our robustness checks, we will consider different ways of defining the reference group. In all cases, we will consider only investors in the same fund f with returns occurring in the same quarter t. In certain cases, we will require further that returns for the comparison group occur in the same month, which we will refer to as return-month restrictions.

Because it is most natural to compare investors who receive distributions from fund f on the

same schedule, in our most stringent tests we will additionally only compare investors who committed capital to fund f at the same time. A limitation of our data is that we do not have well-populated information on the round of fundraising in which investor p committed to fund f. We instead build a proxy using the quarter of the first observed return for investor p in fund f in our data, which we denote by i_p. Then, when computing our gain measures, we only compare investors in fund f at time t who also have the same i_p. We call this an initial-quarter restriction on the comparison group. Similarly, initial-month restrictions on the comparison group require that the month of the first observed return be the same across investors in the comparison group.

3.2 Estimates across vintage-year and asset class

We now provide a sense for the magnitude of potential gains due to within-fund fee dispersion and show how they vary through time and by asset class. For this particular analysis, we use the core sample. As a reminder, for each fund f, the core sample keeps the quarter t where the most number investors report returns, breaking ties with the last observed date. Using the core sample, we then compute potential dollar gains d_{pf} for each investor p in each of their funds f. Here, we suppress time dependencies because the core sample is unique at the investor-fund level.

More formally, let F_v denote the set of funds (and investors in those funds) for vintage year v. Then, the potential gains due to fee differences within funds from vintage year v are given by:

$$G_v = \frac{\sum_{p,f \in F_v} d_{pf}}{\sum_{p,f \in F_v} a_{pf}}.$$

where d_{pf} is defined in Equation 4 and a_{pf} is the amount that p has invested in fund f. When computing potential gains for this subsection, we compare the last observed return of investors in the same fund, conditional on returns occurring in the same quarter. Furthermore, we require that funds be at least 5 years old (i.e., the observed return is at least 5 years after the fund vintage) to ensure that differences in both management and carry have adequate time to impact net-of-fee returns.

Panel A of Figure 3 shows that potential gains typically range between 5% and 12% of dollars invested. The bars corresponding to the left axis plot G_v for each vintage year in our sample. The maroon line in Figure 3 Panel A corresponds to the right axis and shows the number of investor-manager pairs in each vintage year. Unsurprisingly, our underlying sample is less populated for funds in the 1990s, though as risen steadily over time. The peak number of investor-manager pairs occurred in 2006 and 2007 during the pre-crisis boom in private equity. In terms of potential gains, one might expect that within-fund fee differences would trend down through time as the private equity industry matured and investors became more comfortable with the nuances of fund-raising and fee negotiation. However, there are no obvious temporal patterns that immediately stand out from the graph, suggesting that differences in within-fund fees – and hence differences in ex-post performance – have not systematically changed through time.

Within private markets, there are several sub-asset classes like private equity, venture capital, real estate, etc. We can develop a sense of how large fee differences are across these sub-asset classes by computing potential gains within each one. Specifically, let F_y denote the set of fund-investor pairs where the fund is in asset class y. Similar to our other aggregation techniques, we define potential gains in asset class y as:

$$G_y = \frac{\sum_{p,f \in F_y} d_{pf}}{\sum_{p,f \in F_y} a_{pf}}.$$

Panel B of Figure 3 shows plots our estimate of G_y by asset class. The figure reveals fairly large differences across the sub-asset classes in terms of within-fund fee dispersion. In private equity, within-fund fee dispersion is large enough that investors would have $10.37 more per $100 invested had they all received the best ex-post fee in their respective funds. Again, best in this context is defined relative to the other investors that we see in our data. Infrastructure, real estate, and private debt all display relatively similar patterns in terms of the size of within-fund fee dispersion, with gains ranging from $7.48-7.95 per $100 invested. Interestingly, potential gains due to within-fund fee differences are the smallest for venture capital and multi-strategy funds.[16] For venture

[16]There are only two multi-strategy funds in the sample for this analysis.

capital funds, which make up over 20% of the funds in our data, potential gains are $4.86 per $100 invested, less than half of what they are in private equity. In other words, investors in venture capital funds are much more likely to receive similar terms than when investing in private equity.

3.3 Aggregate Estimates for U.S. Pension Funds

Next, we consider how within-fund fee differences have impacted the performance of U.S. pension funds. To aggregate across pensions and funds, we simply sum the dollar shortfall and scale it by the total amount invested in our dataset:

$$G_A = \frac{\sum_{p,f} d_{pf}}{\sum_{p,f} a_{pf}}$$

where the summations are only over U.S. pensions. G_A measures how much better off pensions would be if they received the best return in each of the funds in which they invested. Put differently, G_A captures how much extra surplus U.S. pensions would have captured had they been given the best (ex-post) fee terms in each of their funds.[17]

Table 2 presents estimates of G_A under different restrictions on how we compare investors when computing their potential gains (see Section 3.1.2). In column (1), we consider only funds whose status is liquidated and whose age is at least 10 years.[18] Furthermore, when comparing returns across investors in the same fund, we impose initial-month and month-return restrictions. That is, we require that the investors in the comparison group meet the following criteria: (i) the first month that we observe a return in fund f must be the same across investors; and (ii) the return date that we use to compute return dispersion must be in the same month (as opposed to the same quarter). As previously discussed, these two restrictions are designed to ensure that we only compare investors who invested in fund f at the same time and received cash flows from f on the same schedule.

To see more concretely how these restrictions work in practice, suppose that we observe three

[17] In this context, the best fee terms for fund f are defined based on the other investors that we observe in fund f, including LPs that are not U.S. pensions. Because this is certainly not the full set of investors in the fund, our measure of potential gain is therefore likely to understate the true potential gains.

[18] For a given fund-quarter, we define age as the year of the report date minus the year of the fund's vintage.

different investors $p = A, B, C$ in fund f. Further suppose that the first observation date for both A and B occurs in October 2005 and the first observation date for investor C occurs in November 2005. In this case, we would discard information on C and look for the last observation after 2015 for both A and B, such that the month of the observation is the same. When two such observations exist, we then compute potential gains as described above. If not, then no information from fund f would be included in our aggregation analysis.[19]

Returning to column (1), we estimate 4.69% in potential gains when focusing on liquidated funds with an age of at least 10 years, along with initial-month and return-month restrictions on investors. In column (2), we obtain a similar estimate of 4.55% when we relax the month-restrictions and instead compute within-fund gains by comparing investors whose first and last return quarter is the same. In column (3), we impose no initial date restrictions and compute potential gains for investors in the same fund with same last report quarter. Potential gains are 8.37% for this subset of funds and larger than what we find in columns (1)-(2).

There are several potential reasons why the estimates in column (3) are larger than those in columns (1)-(2). One reason is that liquidated funds in our sample – 60% of which are from 1990s vintages – are more likely to give fee breaks to investors who commit capital in early stages of fund raising. While this would explain why imposing initial date restrictions reduces our estimate of potential gains, we view it as less likely given that initial date restrictions are not as impactful when we analyze non-liquidated funds below. Another possibility is that the funds analyzed in columns (1)-(2) are not representative of the rest of the sample, either because of selection on funds or selection on LPs. With that said, the estimates of nearly 5% in columns (1) and (2) are still fairly large in economic magnitude.

In columns (4)-(6) we broaden the subset of investor-funds over which we compute G_A by considering all funds, not just those that have been liquidated. In all three computations, we impose both initial-month and return-month restrictions. Considering funds that have not yet been

[19] As we explore below, we could require that the first observed return for the pensions that we compare occurs in the same quarter. In this case, then we would include pension C in our return dispersion, provided that A, B, and C have at least one return that occurs in the same quarter after 2015.

liquidated expands our sample substantially, as now we aggregate over at least 2,099 investor-fund pairs, depending on the specification. The difference between columns (4) through (6) is the age restriction that we put on the funds that we consider. Column (4) is more conservative along this dimension, as we only aggregate using data where the fund is at least 10 years old. The age restrictions in columns (5) and (6) are 8 and 4, respectively. In all cases, the estimated G_A ranges from 6.7 to 8.1%.

Columns (7)-(9) mirror the analysis in columns (4)-(6), though instead imposes initial-quarter and return-quarter restrictions when computing within-fund potential gains. This is arguably the most natural way to compare investors in the same fund because it is typical for fund distributions to occur on a quarterly basis; thus, any within-quarter differences in pension reporting are more likely due to differences in reporting practices as opposed to the actual timing of cash flows. The estimates from these columns range from 6.7% to 8.0%, which is similar to columns (4)-(6). Thus, month-restrictions versus quarter-restrictions do not appear to materially impact our results.

Columns (10) through (12) present another set of estimates of aggregate potential gains G_A. In these columns, when computing potential gains, we only require that investors be in the same fund and that their returns occur in the same quarter (plus the usual age restrictions). The potential gain estimates are the largest in this setup, ranging from 8.7% to 9.9%. Compared to the previous estimates, our approach in columns (10) through (12) does not put restrictions on whether investors' first observed return occurs in the same month or quarter. A valid concern here is that we may be comparing investors who do not invest at the same time and thus the timing of their cash flows is unlikely to be economically comparable. On the other hand, just because two pensions do not start reporting returns to Preqin at the same time does not mean that they did not invest at the same time.

Overall, when considering a broad sample of funds (e.g., not only liquidated funds), the average estimate of G_A is 8.5%. In other words, the public pensions in our sample could have captured \$8.50 more in surplus for every \$100 invested had they each received the best fee in their respective funds. In dollar terms, potential gains in our sample are about \$44 billion. This number is based on

aggregate LP contributions of $516 billion.[20] More broadly, Ivashina and Lerner (2018) report that public pensions have shifted over $1 trillion into private market vehicles over the last decade or so. Assuming that our estimates of potential gains apply to this larger sample, then public pensions could have captured nearly $85 billion of extra surplus had they received the same fee contract as other investors in their funds.

It is important to reiterate that our estimate of potential gains is based on an ex-post measure of cost. One way to think about it in ex-ante terms is as follows. Consider a fund that is ten years old, has a gross TVPI of 1.7, and has only two pension investors, A and B. Further suppose that pension A pays an annual management fee of 1.5% and a carry rate of 15%, whereas B pays a 2% management fee and a carry rate of 20%. In this case, A's net-of-fee TVPI after 10 years will be 1.445 (= 1.7 - 0.015$\times$10 - 0.15$\times$0.7) and B's net-of-fee TVPI will be 1.36, both of which are near the average TVPI in our data. Thus, the ex-post difference between investor A and B's net-of-fee return is 0.085, which also corresponds to our estimate of aggregate potential gains.

3.4 Robustness

3.4.1 Alternative Return Measures

DVPI Our analysis thus far has used TVPI to measure returns and hence within-fund fee dispersion. A potential concern is that pension funds differ in their reports of the fund's market value. The fee dispersion we observe may therefore just be due to different reporting standards across pension funds. To address this concern, we use another common measure of returns in private equity called the distributed value to paid-in capital ratio (DVPI). DVPI is the amount of cumulative distributions received by investor p in fund f divided by their cumulative amount of invested capital in the fund. DVPI is thus immune to differences in fund-value reporting standards across pension funds.

Figure 2 shows examples of TVPI and DVPI time paths for different investors in the same fund. For these examples, the performance ranking of investors is consistent across both measures.

[20] Aggregating potential gains using LP commitments of $593 yields potential gains of about $50 billion.

A similar pattern emerges when looking more broadly at the distribution of within-fund DVPI volatility in Panel B of Figure 1. In addition, we compute aggregate potential gains G_A from Section 3.3 using DVPI instead of TVPI. The second row of Table 2 shows that aggregate potential gains based on DVPI are also comparable to those based on TVPI. It therefore seems safe to assume that differences in the reporting of fund net-asset-values do not meaningfully bias our main conclusions.

IRR The internal rate of return (IRR) is another metric that is commonly used in practice to evaluate performance. As discussed in Section 3.1.1, we prefer TVPI to IRR for several reasons (e.g., IRRs are easier to manipulate). For robustness, in the online appendix we recompute potential gains for U.S. pensions using IRRs instead of TVPIs. That is, for each fund f, we define investor p's potential IRR gain as the difference between the maximum observed IRR in the fund and pension p's IRR. We then aggregate over all investors and funds by weighting potential IRR gains by the size of the investment by p in f. In aggregate, we find that U.S. public pensions would have earned 1.65% more in annual IRR had each received the best terms in their respective funds.[21]

3.4.2 Lower Bound on Redistribution from Fee Dispersion

Our counterfactual gain calculation in Section 3.3 makes the following assumption on the size of the surplus of the fund. We assume the fund returned enough surplus such that all pensions invested with this fund could have received the same net-of-fee return per dollar invested as the best performing public pension fund in the fund. This is a plausible assumption because we do not observe all LPs in a given fund, particularly institutions like endowments, sovereign wealth funds, or private pensions. Prior research has found that public pensions underperform these other institutions when selecting private equity GPs (Lerner et al., 2007), so it seems reasonable to think they also do so when investing in the same private-market fund.

With that said, for robustness we consider an alternative measure of fee dispersion that makes

[21] This estimate is based on a combination of hand-reported IRRs by investors in our dataset and manually computed IRRs based on the reported cash flows. See the online appendix for more details.

no assumptions on the unobserved surplus of a given fund. Specifically, we calculate for each pension fund the excess TVPI multiple over the worst performing pension fund:

$$e_{pft} = r_{pft} - r_{ft}^{min},$$

where r_{ft}^{min} is the minimum observed return in fund f at time t. The dollar amount of excess is:

$$d_{pft}^{e} = a_{pft} \times e_{pft}.$$

The dollar amount d_{pft}^{e} is how many more dollars pension fund p received over the pension fund with the worst performance, i.e., highest fee schedule, in the same fund. Critically, this excess gain measure of fee dispersion is budget-feasible because we only consider actual distributions made by the fund to its pension investors. Using the core sample, we calculate the value-weighted average excess gain over the worst performing fund as $E_p = \frac{\sum_{f \in F_p} d_{pf}^{e}}{\sum_{f \in F_p} a_{pf}}$. This denotes the average redistribution of investment gains on private equity funds across public pension funds. The second row of Table 2 presents the results for E_p for different cuts of the core sample. For liquidated funds, the excess gain over the worst performing pension fund is around 2.3%, though in the broader sample it is as high as 4.2%. Assuming that we observe the worst performing investor, these numbers represent a lower bound of excess gains because we only consider redistribution among the pensions that we observe in a given fund.[22]

3.4.3 Does within-fund dispersion in returns reflect differences in fees?

A natural concern could be that our findings are not driven by differences in fees but by specific investor-fund arrangements such as co-investments or investor fund restrictions. Returning to our accounting framework from Section 2.2, the cumulative net-of-fee return of investor p in fund f at

[22] Consistent with Sensoy et al. (2014) and Lerner et al. (2018), public pensions are among the least likely investors to receive preferential treatment from GPs.

time t is:

$$r_{pft} = g_{ft} + \varepsilon_{pft} - c_{pft}$$

where g_{ft} is the cumulative gross return on fund f that is common to all investors, ε_{pft} is any deviation of investor p from the common gross return, and c_{pft} is the cumulative cost paid by investor p to the general partner of fund f. Throughout the paper, we have worked under the assumption that $\varepsilon_{pft} \approx 0$, meaning that variation in r_{pft} across investors in the same fund is mainly driven by variation in c_{pft}. In practice, there are several reasons why this assumption may be less defensible for some investors or some funds. We now discuss two potential issues in more detail and then estimate potential gains from within-fund return dispersion (as in Section 3) on a subsample of the data where these channels are less likely to impact our results.

Co-investment Generally speaking, co-investment structures allow LPs to augment their exposure to a given fund f – what we will call the "main fund"– by allocating additional capital towards a particular deal or set of deals (see Fang et al. (2015) for an in-depth discussion of co-investment). These opportunities are attractive from the perspective of the LPs because they are usually executed at a substantially reduced cost relative to the main fund. To see the way in which co-investment might bias our analysis, suppose there are $p = 1, ...P$ investors in fund f and that investor P is the only one who co-invests. In the main fund, assume that all investors receive the exact same terms. Now suppose that investor P aggregates returns on its co-investment portfolio and the main fund when reporting returns on fund f to Preqin. In this case, the extent to which ε_{Pft} deviates from zero will depend on how much P's co-investment portfolio differs in composition from the main fund. Within-fund variation in r_{pft} across investors will therefore reflect the fact that P earned a different composite gross return than other investors due to co-investment, though we would mistakenly attribute this to differences in cost structures.

While it is difficult to know exactly how much co-investment may distort our analysis, we have several reasons to believe the bias is small. First, and most importantly, it is our understanding that LPs generally list co-investments as a separate line item when reporting performance

to Preqin.[23] As a concrete example, "Fortress Investment Fund IV" and "Fortress Investment Fund IV - co-investment" are classified as two separate funds in our data. Thus, when comparing investors in "Fortress Investment Fund IV", it is unlikely that their net-of-fee returns reflects co-investment. Moreover, for several of the largest LPs in our data, we have manually compared the co-investments that they report on their websites and annual reports against the data from Preqin. In all cases, we have found that cash flows from co-investments are indeed listed separately in the Preqin data for these investors. Because we drop all co-investments in the Preqin data, our assumption that $\varepsilon_{pft} \approx 0$ for all observed investors in our data seems likely to be a reasonable one.[24]

Even in the unlikely case where co-investments are not reported separately in the Preqin data, there is good reason to believe that the degree of bias in our analysis is still small. While public pensions are becoming increasingly more interested in co-investment opportunities, there is some evidence to suggest that co-investment has not been a large part of their private-market portfolios over the majority of our sample (1990-2018). According to data in 2014 from CEM Benchmarking, who provides benchmarking services for thousands of pensions globally, only a small fraction of U.S. public pensions (less than 5%) invest in private equity via co-investments. Furthermore, a 2014 survey by Preqin found that a wide range of institutional investors expressed strong interest in expanding their co-investment capabilities, yet "relatively few LPs are being offered co-investment rights by GPs in the Limited Partnership Agreement".[25] It is natural that co-investment would be less relevant for smaller pensions because it requires the internal infrastructure to evaluate specific portfolio companies and then deploy capital on relatively short notice.

As part of our data quality audit (see Section 2.3 and Internet Appendix Section A.3), we filed direct FOIA requests with 65 pensions and asked them whether they were engaged in any special

[23] We thank Michael Smith and Maeve McHugh at Preqin for many helpful conversations about this issue.

[24] There is of course the possibility that LPs who coinvest receive different fee structures in the main fund. In this case, one would expect that they would pay higher fees in the main fund because those costs are offset by lower fees from coinvesting. If this were true, we should see larger LPs – those that are most likely to coinvest – have relatively *lower* net-of-fee returns in a given fund. As we showed in Section 4.2, the opposite is true in the data.

[25] The CEM report can be found here.. The Preqin survey further states that "there seems to be some contradiction between the attitudes towards and the actual co-investment activity occurring".

investment arrangements such as a side-car deals or co-investments. The vast majority responded that they had no such arrangement. For the few cases that affirmed co-investment arrangements, we were able to verify that these co-investment relationships were reported separately and therefore not included in our analysis.

For larger public pension funds, it also does not appear that co-investments currently dominate their private equity portfolios. For example, CalPERS – the largest U.S. public pension fund – reported in a May 2019 Investment Committee Meeting that it did not start a dedicated co-investment program until 2011 and even that program was suspended in 2016. And, as of 2019, only about 5% of the committed capital in CalPERS' private equity portfolio is dedicated to co-investment. Furthermore, in 2019 CalSTRS – the second largest U.S. public pension fund – was estimated to have roughly 5% of its private equity portfolio in co-investments.[26] Given that the largest U.S. pensions have only modest amounts of co-investment over our sample, it is reasonable to expect that smaller pensions have even less.

To summarize, while co-investment is clearly going to be an important component of public pensions' portfolios going forward, it seems less likely to bias our sample of public pension investment into private markets. Most importantly, co-investment returns appear to be listed separately in the Preqin data and we drop them from our entire analysis. More broadly, for most investors in our data, co-investment is likely to have been a small part of their portfolios over our sample. We also have some survey information from Preqin for the years 2008 to 2017. In the core sample, nearly 90% of the investor-year observations list "No" or did not answer when asked if they co-invest with their GPs. We therefore conclude that co-investment is unlikely to drive the within-fund dispersion in net-of-fee returns that we observe in the data.

Other Investor-Specific Mandates Another reason why ε_{pft} may deviate for some investors in fund f is what we call investor-specific mandates. One prominent example that has boomed in

[26] Based on CalPERS Investment Committee Meeting presentation materials, co-investment was relatively sparse prior to 2011. In response to a direct FOIA request, CalPERS also reported to us that less than 10% of their private equity portfolio's net-asset-value was attributable to co-investment as of 2018. The report on CalSTRS can be found here.

popularity in recent years are so-called environmental, social, and governance (ESG) restrictions. These restrictions mean that investor p in fund f does not allocate capital to specific deals that violate certain ESG criteria (e.g., no investment in firms with a large carbon footprint). More generally, any such investor-specific restriction in fund f implies that investor p's gross return in the fund will necessarily differ from that of other investors.

Unfortunately, we do not have high-quality data on the extent to which pension-specific restrictions exist for the private-market funds in our sample. However, there is some information available on ESG-related restrictions from the National Association of State Retirement Administrators (NASRA), which is an association whose members are directors of a wide array of public pension funds. On their website, NASRA reports that relatively few U.S. pension plans incorporate ESG in their investment process, though some of the larger U.S. pensions have started to do so more in recent years.[27] This is at least suggestive evidence that within-fund return dispersion is probably not driven by investor-specific restrictions for most of the private-market funds in our data. We develop this argument in more detail below.

Potential gains using smaller pensions prior to 2010 In lieu of the preceding discussion, we repeat our computation of potential gains from Section 3.3 using a sample that is less likely to be biased by co-investment and investor-specific mandates. Specifically, we drop large pensions (assets over $100 bn) and private-market funds whose vintage year is after 2009, as co-investment and investor-specific mandates like ESG are a recent trend and are less likely to be relevant for smaller pensions.

Table A2 in the online appendix contains the results of this exercise. Reassuringly, our estimates of aggregate potential gains due to within-fund return dispersion in this restricted sample (~$7.70 per $100 invested) are broadly comparable to those using the all pensions and funds in the core sample (~$8.50 per $100 invested). These findings support our argument that the within-fund return dispersion that we document is most likely due to differences in fee-structures across investors.

[27] See the following link.

4 Do some investors consistently get better terms?

Having documented the aggregate consequences of within-fund performance dispersion, we now explore whether some pensions consistently get better or worse terms when investing in the same fund. Our approach to answering this question also allows us to assess whether any cross-pension differences in fees that we observe in our data are statistically meaningful. In the last part of the section, we map within-fund performance to characteristics like size or relationships, and decompose how much these characteristics can account for observed within-fund performance dispersion.

4.1 Measuring pension-effects

4.1.1 Baseline estimates

To test whether some pensions consistently outperform other investors in their respective funds, we use the following regression:

$$r_{pf} = \alpha_f + \theta_p + \varepsilon_{pf} \tag{5}$$

where r_{pf} is the return of investor p in fund f, α_f is a fund fixed effect, and θ_p is an investor fixed effect. Because the regression contains fund fixed effects, the θ's capture whether some pensions consistently earn above or below the average return in their respective funds. For this reason, we refer to the θ's as pension-effects. Under the null hypothesis of no pension effects, the estimated θ's should not be statistically distinguishable from each other. In other words, if fee contracts in a given fund are randomly assigned to pensions, then we should not be able to reject an F-test that the θ's are jointly equal to each other.

Panel A of Table 3 reports the F-tests and their associated p-values based on the core sample described in Section 2.3. When moving from rows (1) to (3), we conduct the F-test for whether the θ's are jointly equal based on funds that are at least one, four, and eight years old, respectively. In all cases, the estimated F-statistic is large enough that we reject a null of no pension effects with a p-value of less than 0.01.

The standard approach to conducting F-tests like those in Table 3 rely on parametric assump-

tions to test the null of no pension effects. As a robustness check, we run permutation tests where we randomly assign returns to investors in fund f. For each random assignment, we calculate the F-statistic from the test of equality across θ's. We repeat this procedure 1,000 times to generate an empirical distribution of the estimated F-statistics, after which we compute a non-parametric p-value based on where the actual F-statistic falls in this distribution. In the table, we denote the p-values based on these permutation tests as p^*. Reassuringly, we comfortably reject the null of no pension effects even when using these non-parametric p-values.

Though the preceding F-tests provide a statistical sense of the size of pension-effects in our data, they are somewhat silent on the economic magnitude of such effects. To get a better sense of how large the estimated pension-effects are in economic terms, we simply compute the distribution of the estimated pension effects. Before doing so, we use an empirical Bayes procedure to account for the fact that the true distribution of θ's will differ from our estimated distribution because of sampling error (see Chetty et al. (2014) for an example in labor economics). Formally, let θ denote the vector of estimated θ's based on regression (5) and $\widetilde{\theta}$ denote the empirical Bayes estimate. According to Egan, Matvos, and Seru (2018), the two are linked as follows:

$$\widetilde{\theta} = \left(\hat{\theta} - \overline{\theta}\right) \times \frac{F - 1 - 2/(K-1)}{F}$$

where $\overline{\theta}$ is the average of the estimated fixed effect vector $\hat{\theta}$, F is the F-statistic corresponding to the joint test that the elements of $\hat{\theta}$ are equal (reported in Panel A of Table 3), and K is the number of pension effects. Intuitively, the F-statistic is larger when the pension-effects are estimated with more precision, and in turn, the Bayes estimate does not shrink $\hat{\theta}$ as much towards its mean.[28]

Panel B of Table 3 provides summary statistics on the distribution of the pension effects $\widetilde{\theta}$ for different estimation samples. Like with Panel A, the estimation samples only differ in the minimum age of funds that are included when estimating regression (5). Differences in pension effects are economically large, regardless of the age restriction that we impose on funds. For

[28]This particular form of the Bayes estimator relies on the assumption that both the true pension effects and the regression errors are normally distributed (Egan, Matvos, and Seru, 2018). In this case, the procedure can be easily implemented with standard regression outputs.

example, when looking at funds that are at least eight years old, the standard deviation of pension effects is 523 basis points. If we interpret the θ's as a measure of pension quality, then this suggests that *within* a given fund, a one-standard deviation improvement in pension quality leads to a 523 basis point improvement in performance. The tails of the distribution of pension effects are even more pronounced, as moving from the 10th to 90th percentile of quality translates to an increase in within-fund performance of roughly 800 basis points.

To provide some more context for the size of the pension effects, we compute a fund-level measure of performance r_f by taking the median TVPI of $r_{p,f}$ across investors in fund f. In funds of at least eight years of age, moving from the 10th to 90th percentile of pension effects based on *within-fund* performance is roughly equivalent to the difference in performance between the 50th and 40th percentile private equity fund. In other words, within-fund variation in performance is as large in some cases as between-fund variation in performance. More broadly, these results show that some pensions consistently outperform others when investing in the same fund, both in an economically and statistically significant sense.

4.1.2 Robustness using DVPI

One natural concern with our baseline estimates of pension effects is that they rely on TVPI as a measure of returns. For each investor p in fund f, TVPI reflects actual cash distributions that have been received by p as well as the reported net asset value of p's non-liquidated positions in the fund. Consequently, if some pensions consistently report net asset values differently than others, then one might be concerned that this drives our estimate of pension effects. To alleviate these concerns, in the online appendix we repeat our analysis from Section 4.1.1 using DVPI to measure returns instead of TVPI. DVPI is based only cash distributions and is therefore not influenced by how pensions report net asset values. When using DVPI to measure returns, we reassuringly find once again strong evidence of large pension effects.

4.1.3 LP-GP Effects

Research suggests that relationships matter for the performance of private market investments (e.g., Hochberg, Ljungqvist, and Lu (2007)). To obtain further insight into the nature of pension effects and within-fund performance, we investigate whether the relationship between GPs and LPs relate to the fees that LPs pay in a given fund. To this end, we augment regression (5) as follows:

$$r_{pfg} = \alpha_f + \theta_p + \eta_{pg} + \varepsilon_{pfg} \tag{6}$$

where r_{pfg} is the return (TVPI) of investor p in fund f managed by general partner g. As before, α_f are fund effects and θ_p are pension effects. The new term in the regression is η_{pg}, which are LP-GP fixed effects. In this setup, the pension effects θ_p measure the extent to which some pensions outperform others in a given fund and the LP-GP effects η_{pg} measure any incremental outperformance in funds managed by GP g. If, for instance, some investors receive better terms than others in funds managed by a specific set of GPs, then we should reject an F-test of the joint significance of the η's.

Rows (4) through (6) of Panel A in Table 3 reports F-statistics and their associated p-values from testing whether the θ's or the η's are jointly equal. We continue to reject the null of no pension effects, regardless of the conditions that we put on the age of funds. In addition, we reject the null of no LP-GP effects (η's) when using parametric p-values and non-parametric p-values based on the permutation tests described in Section (4.1.1). This evidence suggests that LP-GP relationships are important for explaining why some pensions consistently outperform others even when investing in the same fund. We explore this finding in more detail below.

4.2 Observable Pension Characteristics

Given their importance in determining within-fund performance, we now investigate whether pension effects can be mapped to easily observable characteristics in the data. We do so by replacing

pension fixed effects in regression (5) and with observable characteristics as follows:

$$t_{pf} = \alpha_f + \beta \mathbf{X}_{pf} + \varepsilon_{pf}, \tag{7}$$

where $\mathbf{X}_{pf}$ is a vector of pension characteristics. Again, the fund fixed-effect α_f means we are comparing outcomes of different pensions in the *same* fund.

Observable Characteristics X_{pf} We consider the following set of observable characteristics. For each investor p in fund f, we compute p's share of the total fund as their commitment amount divided by the total fund size. We include each investor's share of the fund to account for potential returns to scale when raising capital. That is, one might expect that GPs might reduce fees for investors that account for a larger fraction of the fund, as this would then free up the GP to focus on optimizing the investment portfolio instead of raising capital.

We also include two variables that capture the experience of each investor in private markets. A priori, it is plausible to think that skill in fee negotiation improves as investors become more experienced in the nature of private market investment vehicles. Thus, for each investor p and fund f, we include the (log) number of funds in which p has invested in the ten years prior to f's vintage year. We chose a ten year window to allow for turnover on the investment staff responsible for negotiating fees at each pension. The second variable related to experience is specific to each LP-GP pair and is motivated by our finding of LP-GP effects in Section 4.1.3. Specifically, for each GP-LP pair, we count the number of funds that are managed by general partner g in which p has invested. We use the full dataset to compute this measure because we want to capture settings where a GP reduces fees for investor p in fund f in expectation that the investor will invest in future funds raised by the GP.

When there is asymmetric information about its skill, a GP may also reduce fees for investors that will send a positive signal to other potential investors. While it is difficult to capture all of the dimensions through which an investor might be considered a so-called "cornerstone LP", we focus on size and commitment date. We code investor p as "Large" if its total assets under management

are over $100 billion at the time of fund f's launch, a designation that is reserved for a handful of easily recognizable pensions in our data. In addition to the potential signaling effect that they may have on fund raising, large investors are also more likely to possess the ability to deploy large amounts of capital quickly, so size is likely related to the economies to scale in fund raising discussed above. Due to data limitations, we are not able to observe if p is a first-close investor in fund f, but we proxy for it based on whether p's first report date in the master sample occurs before other investors in fund f.

The last set of variables that we include in X_{pf} are related to pension governance.[29] We include board size to account for any potential coordination problems that may cause larger boards to sub-optimally negotiate fee contracts. In addition, Andonov et al. (2018) find that pension boards with more state officials are more likely to make poor investment decisions in private equity, likely due to distortions from political considerations. Motivating by that finding, X_{pf} includes the percent of pension p's board that is made up of elected, as opposed to appointed, members. Intuitively, boards that have more elected officials are more likely to focus only on the beneficiaries of the plan when determining fee schedules with their general partners. For each pension and fund pair (p, f), the total number of board members and the percent of elected members are both measured as of fund f's vintage year.

Results Table 4 reports OLS estimates of equation (7) using the core sample, again with several restrictions on fund age. In all cases, we cluster standard errors by investor-and-vintage year because many of the covariates in X_{pf} are repeated for investor p across all funds in a given vintage year. In addition, for each age restriction, we estimate (7) with and without the governance variables.

Consistent with our finding of LP-GP effects in Section 4.1.3, pensions that have a stronger relationship with a GP – as measured by the number of funds in which they have invested with that GP – tend to outperform other investors in funds managed by the GP. The point estimates on

[29] We are grateful to Josh Rauh for sharing the data on pension board composition from Andonov et al. (2018). The data ends in 2013, so we extend it to 2018 to better match our sample.

LP-GP relationships range from 5.45 to 12.34 depending on the specification and are in most cases statistically significant at conventional levels. These estimates indicate, in a given fund, being invested in a different fund managed by the GP is worth roughly 9 basis points of returns. For additional context, the standard deviation of LP-GP relationships in our data is about 4.

The regression estimates in Table 4 also indicate that being a large investor is an important driver of within-fund performance dispersion. In terms of magnitude, when multiple pensions invest in the same fund, large pensions receive contracts that generate an additional 90 to 159 basis points in returns, depending on the life of the fund. The point estimate on the large-investor indicator is measured with statistical precision, as we can always reject the null of no large-investor effect with 95% confidence. As we discuss further in Section 6.1, there are several reasons why larger pensions might create value for GPs, for instance through signaling effects or returns to scale in fund raising. Our results suggest some of this value is at least partially passed to large pensions through cost reductions.

In terms of governance, we find that pensions with a larger share of elected board members tend to outperform other investors in their respective funds. This elected-board effect is statistically significant with $p < 0.05$ regardless of the fund-age restriction that we impose when estimating equation (7). To get a sense of magnitude, consider funds that are at least eight years in age. For this subsample of the data, a one percentage point increase in the share of elected board members translates to just under 3 basis points of within-fund performance. The 25th percentile pension has no elected members whereas the 75th percentile pension has nearly half of its board elected. Thus, moving from the 25th to 75th percentile pension in terms of elected board members translates to nearly 150 basis points of improved performance, relative to other investors in the same fund.

When looking at the remaining covariates, the share of investor p in fund f, investor p's experience at the time of capital commitment, and board size do not appear to have a statistically meaningful impact on within-fund performance. These findings are perhaps not so surprising, as it is natural to expect that many of these variables are subsumed by size and LP-GP relationships when determining pension fee schedules. Interestingly, the impact of being an early-investor is

generally positive and around 30 basis points in TVPI units, depending on horizon. However, the point estimate is always noisy, which is to be expected given that we do not have comprehensive data on the timing of capital commitments.

4.3 How much of pension-effects are due to observables?

The preceding analysis showed that, within a given fund, pensions that are large, more connected to the GP, and governed by more elected board members outperform other investors in the fund. We now quantify how much these observable characteristics account for the pension effects that we documented in Section 4.1. We do so in two complimentary ways. First, we reestimate pension effects after controlling for observable characteristics, after which we compare the distribution of these characteristic-adjusted pension effects to the raw ones from Section 4.1. Second, we compute potential pension-level gains in an analogous fashion to the aggregate measures from Section 3. We then regress returns on characteristics and reestimate pension-level gains based on the regression residuals. Comparing the raw pension-level gains against the characteristic-adjusted gains provides a sense of how much observable characteristics account for the within-fund return dispersion that we see in the data.

4.3.1 Characteristic-Adjusted Pension Effects

To adjust our baseline pension effects for characteristics, we estimate the following regression:

$$r_{pf} = \alpha_f + \theta_p + \beta X_{pf} + \varepsilon_{pf} \tag{8}$$

where α_f is a fund fixed-effect, θ_p are pension effects, and X_{pf} is the vector of characteristics described in Section 4.2. Because we include X_{pf} in the regression, the estimated pension effects θ_p capture whether some pensions consistently outperform others in their respective funds, even after controlling for observable characteristics. We estimate equation (8) using OLS and conduct an F-test for the null that the θ's are jointly equal to each other. We then apply the empirical Bayes

procedure from Section 4.1 to adjust the θ's for estimation error.

The last three rows of Table 3 summarize the F-tests for the subsample of funds that are at least one, four, and eight years old, respectively. In all cases, even when including observable characteristics, we still consistently reject the null of no pension effects with $p < 0.01$. This is true when we use standard p-values or when we use non-parametric ones generated using the permutation tests described in Section 4.1.1. Thus, from a statistical perspective, the observable characteristics we consider do not account for the strong pension effects are present in the data.

To develop a sense of economic magnitude, Figure 4 plots the distribution of the raw pension effects from Section 4.1 against the characteristic-adjusted pension effects from equation (8). The two curves are virtually indistinguishable from each other, indicating that some pensions consistently outperform others in the same fund by a sizable margin, even after for accounting for attributes like commitment size, overall size, LP-GP relationships, and governance. Put differently, these results imply that two pensions with similar characteristics still consistently pay different costs when investing in the same private-market fund.

4.3.2 Characteristic-Adjusted Potential Gains

Next, we explore a complimentary way to quantify how much pensions observables explain and why some consistently outperform others within the same fund. We begin our analysis by estimating potential gains at the investor level in a similar manner to how we estimated potential gains due to within-fund fee dispersion in Section (3). Formally, let F_p be the set of funds in which investor p invests. We aggregate potential dollar gains to the investor level as follows:

$$G_p = \frac{\sum_{f \in F_p} d_{pf}}{\sum_{f \in F_p} a_{pf}}.$$

As a reminder, d_{pf} measures the dollar gain that investor p would have experienced had it received the best observed return in fund f. a_{pf} is the amount that p invests in fund f. Thus, G_p is simply the sum of investor p's potential gains, scaled by its total investment in all funds.

The green bars in Figure 5 summarize the distribution of G_p across the different U.S. public pensions in our sample.[30] There is considerable heterogeneity across pensions in terms of potential gains due to fee differences. The standard deviation and interquartile range of potential gains are 3.9% and 5.1%, respectively. The extremes of the potential-gains distribution are even more pronounced. For example, the 5th and 95th percentile of G_p are 1.1% and 14.9%, respectively. In other words, the private-market investments of the 5th-percentile pension would be 1.1% larger if it has received the best fee terms in its respective funds, whereas the 95th percentile pension would be 14.9% larger.

To understand how much of the heterogeneity in G_p can be attributed to observable pension characteristics, we obtain the residuals ε_{pf} from the following regression:

$$r_{pf} = \alpha_f + \beta X_{pf} + \varepsilon_{pf},$$

where once again r_{pf} is the TVPI of investor p in fund f in the core sample. The vector of characteristics X_{pf} is the same as before, but we exclude the variables related to governance in order to include as many pensions in the analysis as possible. We have verified in unreported results that we draw similar conclusions when including them. We then define the characteristic-adjusted potential gain of investor p in fund f as:

$$\widetilde{d}_{pf} = (\varepsilon_{pf}^{max} - \varepsilon_{pf}) \times a_{pf}$$

where ε_{pf}^{max} is the maximum regression residual within fund f. The characteristic-adjusted potential gain for investor p, $\widetilde{d}_{pf}$, aggregates as before:

$$\widetilde{G}_p = \frac{\sum_{f \in F_p} \widetilde{d}_{pf}}{\sum_{f \in F_p} a_{pf}}.$$

[30] To ensure reporting of net asset values does not influence our gain measures, we recompute G_p using DVPI instead of TVPI. The rank correlation between G_p computed using TVPI and DVPI is 70%, which suggests that any issues with measuring net asset values do not drive our results.

Intuitively, if observable characteristics account for a large amount within-fund variation in returns, then the within-fund variation in ε_{pf} will be small and so too will $\widetilde{G}_p$. The black outlined bars in Figure 5 overlays the distribution of $\widetilde{G}_p$ on the raw estimates of G_p. The distribution of $\widetilde{G}_p$ does tighten towards its mean to some extent, though it is clear that the distribution of $\widetilde{G}_p$ and G_p largely line up with each other. Thus, consistent with our analysis of characteristic-adjusted pension effects, observable characteristics do not fully explain the wide heterogeneity across pensions in terms of potential gains due to fee differences.

For confidentiality reasons, we are unable to provide estimates of potential gains due to fee-dispersion at the pension level. However, we are able to aggregate them to the state level, which we do in Figure 6. The figure shows the equal-weighted average of characteristic-adjusted gains $\widetilde{G}_p$ by state, which for the reasons mentioned above is essentially the same as raw characteristic gains. It is important to keep in mind that we do not observe all pensions in each state, so the figure paints an incomplete picture in this regards. Overall, though the plot reinforces the finding that pensions differ greatly in terms of how sensitive their returns are to better fee structures. Even after adjusting for characteristics, the top decile of states in potential gains could earn on average \$12.44 per \$100 invested with better terms, whereas the bottom decile could earn on average \$2.32.

5 Fee Dispersion Mechanism

So far, we have documented that investors in the same private market vehicle very often earn different net-of-fee returns. As outlined in our motivating framework from Section 2.2, it is natural to attribute within-fund dispersion in returns to differential fees paid by investors in the same fund. In this section, we provide more details on the nature of the fee dispersion within private market funds. We first present evidence suggesting that funds appear to group their investors into different contracts. That is, the evidence shows that same fund features several distinct clusters of investors' net-of-fee performance. Then, we show that performance fees are an important driver of differences in contract terms among investors.

5.1 Investor tiers

We begin by investigating whether investors in the same fund receive individualized contracts or whether they get clustered into tiers. From the perspective of the GP, it is easy to imagine that tiering investors is advantageous because it avoids any holdup and processing costs associated with implementing many different limited partnership agreements. In our setting, the natural way to detect any such tiering exists is to check for clustering in net-of-fee returns within the same fund. Panel A of Figure 7 provides an example of the clustering we observe for an anonymized fund that we call Fund III. There are over 20 investors in this fund and all have return data that appears in our dataset in the same month and year, which is a loose indication that these investors began contributing capital to Fund III at the same time. The TVPI distribution in the figure clearly shows three distinct clusters. Roughly 70% of the investors in Fund III have a TVPI ratio of around 1.98. The two investors in the best performing cluster both have TVPI ratios close to 2.15, whereas investors in the lowest cluster all earn around 1.7 in net-of-fee cumulative returns. Thus, Fund III appears to tier their pension investors into three groups.

To broaden this analysis across all funds in our sample, we round TVPIs to the third decimal and compute the number of distinct rounded TVPI values in each fund. We observe similar patterns if we instead compute tiers based on investors in the same fund who also have return data that appears in the same year and quarter. Generally speaking, this approach to counting investor tiers is extremely aggressive because two investors who are economically close in terms of returns will nonetheless be treated as belonging to different tiers. With that in mind, Panel B of Figure 7 shows the distribution of tiers measured in this fashion across the funds with at least five investors in the core sample. Even with this aggressive approach to tier classification, we still observe evidence of return clustering, as nearly half of the funds have less than five unique TVPI values.

In Panel C of Figure 7 we employ machine learning techniques to more formally define clusters within a fund. The goal of machine learning clustering methods is to efficiently partition observations into distinct groups. A standard approach to this problem is called k-means clustering, which assigns observations to one of k clusters based on their distance to the clusters. The k clusters are

themselves chosen to minimize the total distance of observations to their respective clusters. We select the optimal number k of clusters based on Silhouette scores, as is common in the machine learning literature. Panel C of Figure 7 shows the distribution of the optimal k across funds in the core sample with at least five investors. As before, we obtain very similar results if we instead analyze returns across investors in the same fund who also have return data starting in the same year and quarter. According to this procedure, the vast majority of funds in our sample have two or three tiers of investors. More broadly, this analysis suggests that investors in the same fund are grouped into tiers in terms of fee structures.

5.2 Fee dispersion via performance fee

As discussed in Section 2.1, there are two main components of the fee structures in private equity: management and performance fees. While we do not explicitly observe either of these specific contract terms, our data does allow us to learn something about how these terms differ across investors in the same fund. To fix ideas, consider a fund f that has only two investors, $p = A$ and B. For the sake of this example, we assume that A and B have identical limited partnership agreements, except for their performance fees. Under this assumption, from Section 2.3 we can write the cumulative net-of-fee returns for both investors at time t as:

$$r_{Aft} = g_{ft} - m_f \times t - \kappa_{Af} \times \max\left(g_{ft} - 1, 0\right)$$
$$r_{Bft} = g_{ft} - m_f \times t - \kappa_{Bf} \times \max\left(g_{ft} - 1, 0\right)$$

where m_f is, by assumption, the common annual management fee for fund f. Our assumption that the two investors pay different rates of carry means that $\kappa_{Af} \neq \kappa_{Bf}$ and without loss of generality we assume that $\kappa_{Bf} > \kappa_{Af}$. Within fund f, the range Δ_{ft} in net-of-fee returns is:

$$\begin{aligned}\Delta_{ft} &\equiv r_{Aft} - r_{Bft} \\ &= (\kappa_{Bf} - \kappa_{Af}) \times \max\left(g_{ft} - 1, 0\right)\end{aligned}$$

Differentiating with respect to g_{ft} yields

$$\frac{\partial \Delta}{\partial g_{ft}} = \begin{cases} 0 & \text{if } g_{ft} < 1 \\ \kappa_{Bf} - \kappa_{Af} & \text{if } g_{ft} \geq 1. \end{cases} \tag{9}$$

Equation (9) implies that we can estimate the range of carry rates within a fund by measuring the elasticity of the within-fund range of returns with respect to the gross level of returns. In other words, within-fund differences in carry can be identified by plotting the within-fund range of net-of-fee returns against the level of fund returns, after controlling for age. The relationship should be flat in the region where the fund is not profitable ($g_{ft} < 1$) because differences in carry do not impact net-of-fee returns. In the region where the fund is profitable ($g_{ft} \geq 1$), the relationship between within-fund range and gross returns should be linear, with the slope revealing differences in fees.

Figure 8 evaluates this prediction in our data. Using the master sample, we compute the range of returns Δ_{ft} within each fund at time t. We also proxy for the gross return in the fund g_{ft} using the maximum observed net-of-fee return at time t, which we denote by $\hat{g}_{ft}$. Figure 8 is a binscatter plot of Δ_{ft} against $\hat{g}_{ft}$, after both have been residualized to the age fund f at time t.

As expected, when $\hat{g}_{ft}$ is less than one there is a flat relationship between the range of within-fund returns and the level of returns. Indeed, in this region, the estimated slope from a regression of one on the other yields a point estimate of $\beta = 0.00$. In the region where $\hat{g}_{ft}$ is greater than one, there is a clear positive and linear relationship between Δ_{ft} and $\hat{g}_{ft}$. The slope estimate is 0.16, with the 95% confidence interval spanning 0.10 to 0.22.[31] Our interpretation of this finding is that carry rates differ by around 16 percentage points across investors for the average fund in our sample. Any option-like features of the limited partnership agreement will influence the point estimate of $\beta = 0.16$, so a broader interpretation is that embedded options in the fee-contract between LPs and GPs are an important driver of within-fund dispersion in net-of-fee returns.

[31] The confidence interval for β is based on standard errors that are double-clustered by fund and vintage year.

6 Interpretation and Conclusion

6.1 Interpretation

Our analysis revolves around variation in net-of-fee returns across investors in the *same* private-market fund. Using this variation, we document strong pension effects in the sense that some pensions consistently earn higher net-of-fee returns than others in the same fund. We now consider several explanations for these facts.

Perhaps the simplest explanation is that the observed within-fund variation in net-of-fee returns is due to measurement error or bespoke investment vehicles (e.g., co-investment). While it is certainly possible that our data contains measurement error, the quality of the data should be reasonably high given that it is mainly based on FOIA requests. As we document in the online appendix, the quality of the data is further confirmed by the audit that we have conducted on a subsample of pensions in the Preqin data.[32] Moreover, measurement error should bias us against finding strong pension effects, yet we do so even when using only realized cash flows (DVPI) to compute returns. In addition, our cash flow data is unlikely to bundle returns from bespoke investment vehicles (see Section 3.4). Consequently, the within-fund variation in returns that we observe is most likely driven by variation in fee structures across investors.

GPs might offer investors different contracts if some are more costly to partner with than others. For example, from the perspective of GPs, public pensions may be more costly investors because they have more stringent reporting requirements than family offices or private endowments. In this case, the GP would offer each pension a contract that reflects their marginal cost. This would clearly generate ex-post fee dispersion. However, we only analyze public pensions and marginal servicing costs seem unlikely to vary within this group because pensions have relatively homogenous reporting and compliance requirements. Plus, to the extent that they do vary across pensions, marginal costs should presumably correlate with size or experience in private markets. Further-

[32] We chose the subsample of pensions based on two criteria. First, we select the set of top and bottom ten pensions based on potential gains from Section 4.3. Second, we select the pensions in the funds with widest range of within-fund variation in returns.

more, we show in the online appendix that there are strong pension effects when we control for size in a more flexible, non-parametric fashion. The fact that we still find large pension effects after controlling for these characteristics cuts against a cost channel and instead suggests that pensions may differ in how they value a fee structure within a private-market fund.

In turn, persistent differences in willingness to pay across pensions can arise for several reasons. One is preferences: if some pensions are more risk averse than others, they may prefer fee schedules with low performance and high management fees, as these would deliver less volatile cash flow streams. While preferences are difficult to measure empirically, in the online appendix we proxy for risk aversion using each pension's portfolio share in cash or cash equivalents. We still find similarly sized pension effects when augmenting our set of controls with this risk aversion proxy. In other words, for preferences to explain the data, it would have to be that are they are mostly orthogonal to these attributes. Furthermore, a preference-based explanation would imply that some pensions in our sample must be content to concede a fairly large amount in terms of ex-post performance.

Pensions may also vary in their willingness to pay because they have different beliefs about gross fund returns, which could arise from differences in information about GP skill. In the context of private markets, it certainly seems reasonable to think that some pensions may be more informed about GP skill than others. Informational asymmetries (or other LP-GP synergies) would also explain why GPs appear to give fee breaks to large pensions with whom they have a strong relationship, as this presumably aids in capital raising by sending a signal to less informed investors.[33] Nonetheless, to explain the data, it would have to be that information edges and signaling effects are largely orthogonal to attributes like size or relationships, as we still observe strong pension-effects after controlling for these characteristics.[34]

One way to generate a disconnect between persistent informational edges and pension size is

[33] In practice, these sorts of fee breaks are often contractually implemented through most favored nation clauses (MFNs), which ensure that special investors will receive the best terms made available to other investors. MFNs also provide their holders with an additional information edge and are not typically given to all LPs.

[34] In the online appendix, we provide some suggestive evidence that pension effects are not driven by return expectations on private equity as an asset class.

as follows. Suppose that the portfolio managers at public pensions differ in their ability to evaluate GPs, but labor market frictions restrict the mobility of talent from smaller, resource-constrained plans to larger plans. Recent estimates of labor market frictions in the context of public pensions suggest that they are rather sizable (Dyck et al., 2018). In this case, the equilibrium distribution of talent at public pensions may be determined by forces that are uncorrelated with pension size (e.g. geographical preferences). While discerning luck from skill is always difficult, this mechanism would account for why some pensions have chosen fee structures that have consistently underperformed over a sample that spans three decades and several private equity cycles.

More generally, optimization frictions could feasibly generate pension effects as well. By optimization frictions, we mean any force that would cause some pensions to consistently choose suboptimal fee structures, including irrational beliefs about gross fund returns, failure to fully internalize the cost structures, or agency frictions. While such frictions are hard to measure cleanly, there are some empirical reasons to doubt that all public pensions have perfectly rational beliefs and choose their fee structures accordingly: less than 5% of the plans in our data have any mention of performance fees or carry on their annual reports, despite the fact that differences in carry are an important component of within-fund fee variation (see Section 5.2).[35] Moreover, political considerations appear to distort the investment decisions of public pensions (Andonov et al., 2018). Overall, it is hard to imagine that these sorts of optimization frictions play no role in explaining why pensions with similar characteristics – and therefore those that in principle should have similar bargaining power, preferences, and information – appear to be systematically willing to pay different fees when investing in the same private-market fund.

Regardless of the precise microfoundation, if pensions differ in their willingness to pay for private-market investments then a profit-maximizing GP will do its best to exploit these differences by price discriminating through fee structures. In practice, willingness to pay might be revealed through negotiations or by offering LPs a menu of fees from which to choose. Our analysis collec-

[35]This statistic is based on the plans for which we could find annual reports in the master sample. Valuing the embedded-options in the fee structure requires estimates of the fund's volatility. Given the lack of consensus on the risk of private equity as an asset class (Harris et al., 2014; Ang et al., 2018), generating a fund-level estimate of volatility and then applying an option-pricing model is a difficult problem for any investor to solve.

tively suggests that price dispersion of this kind is large in the context of private markets and is at least partly driven by pensions that fail to fully optimize when choosing their fee structures.

6.2 Conclusion

We use within-fund variation in net-of-fee returns to show that public pensions investing in the same private-market fund can experience very different returns. In aggregate, within-fund fee dispersion means that pensions could have earned $8.50 more per $100 invested had they each paid the fees of the best observed pension in their respective funds. Moreover, there are consistent winners and losers in the sense that some pensions systematically pay more fees than others even when investing in the same fund (i.e. pension effects). The pension effects that we estimate are large in the cross-section, as some pensions could have earned as much as $15 more per $100 more on their investments over our sample. Size, relationships, and governance account for some of the pension effects, but the majority appear orthogonal to these observable characteristics. We argue that these facts are puzzling from the perspective of several rational models of fee determination and consistent with the idea that public pensions fail to optimize when choosing fees. While this evidence is far from conclusive, it highlights the value of further empirical research on how U.S. public pensions structure fees in private-market investment vehicles.

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Table 1: Summary Statistics for Core Sample

	Min	p25	p50	Mean	p75	Max
Investors per Fund	2	2	3	4	5	30
Years since Inception	1	4	7	8	11	26
Funds per Investor	1	3	11	47	50	504
GPs per Investor	1	2	9	26	31	273
Commitment ($ mm)	0.11	11	27	55	68	1,738
% of total fund commitment	0.02	1.00	2.78	5.07	6.58	99.26
Overall AUM of LP ($bn)	0.01	2.4	8.7	23.2	26.0	326.5
TVPI	0.00	1.08	1.32	1.43	1.62	32.99
Within-Fund TVPI Range	0.00	0.01	0.05	0.09	0.13	1.12
Total Number of Investors	231					
Total Number of Funds	2,535					
Total Number of General Partners	916					

Notes: This table presents summary statistics for the core sample that we use in our analysis. The core sample, which is defined formally in Section 2.3, contains only U.S. public pensions and is unique at the investor-fund (p, f) level. TVPI refers to the total multiple on invested capital for investor p and is defined as the cumulative amount of received distributions plus the remaining net-asset-value, all divided by the cumulative amount of contributed capital.

Table 2: Potential Return Gains Due to Fee Differences

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
TVPI Potential Gain (of % invested)	4.69	4.55	8.55	8.11	8.00	6.69	8.03	7.94	6.71	9.94	9.68	8.66
DVPI Potential Gain (of % invested)	4.51	4.46	8.37	7.52	6.06	4.67	7.42	6.03	4.65	8.99	8.47	6.83
TVPI Excess Gain (of % invested)	2.24	2.25	3.63	3.68	3.70	2.93	3.69	3.72	2.96	4.15	4.18	3.63
Amount Invested ($ bn)	27	29	85	157	186	285	163	193	295	288	335	471
Total N	558	615	2,040	2,099	2,600	4,384	2,238	2,754	4,622	5,396	6,363	9,235
Number of Funds	178	197	597	554	680	1,054	586	713	1,099	1,357	1,589	2,228
Number of Investors	64	68	132	107	125	152	111	126	154	176	190	230
Sample Restrictions:												
Years Since Inception	10	10	10	10	8	4	10	8	4	10	8	4
Fund	×	×	×	×	×	×	×	×	×	×	×	×
First Contribution Month	×			×	×	×						
Last Report Month	×			×	×	×						
First Contribution Quarter		×					×	×	×			
Last Report Quarter		×	×				×	×	×	×	×	×
Liquidated	×	×	×									

Notes: This table presents several measures of within-fund fee dispersion based on the core sample. To generate the first row of the table, we compute the incremental return gain (measured in TVPI) for investor p in fund f if it had earned the best observed return in the fund. The potential dollar return gain is simply the incremental return gain for each investor p multiplied by its contribution to fund f. We aggregate by summing potential dollar gains and then scaling by the total amount of contributed capital across investors. The second row is the same measure using DVPI instead of TVPI to measure returns. The third row is an analogous calculation, but we instead compute each investor's excess gain in fund f as the difference between their actual return (measured in TVPI) and the lowest observed return in the fund. The remaining rows provide summary statistics and details on the subsample that we use to compute these metrics. The subsample restrictions that refer to the first and last contribution dates (either month or quarter) describe how we compute the maximum or minimum observed within-fund return. For example, when the restriction on the first contribution month means that we only compare investors whose first observed contribution month in the data is the same. See Section 3 for a complete discussion.

Table 3: Pension and Pension-GP Effects on Within-Fund Performance

Panel A**:** *Statistical Tests of Pension and Pension-GP Effects*

		Dependent Variable: TVPI (bps)								
Min. Fund		Pension-Effects				Pension-by-GP Effects				
Age	Controls	F	p	p^*	K	F	p	p^*	K	N
1		5.41	<0.01	<0.01	205					10,848
4		5.23	<0.01	<0.01	191					8,493
8		4.13	<0.01	<0.01	158					4,923
1		1.87	<0.01	<0.01	165	1.34	<0.01	<0.01	1,765	6,896
4		1.98	<0.01	<0.01	156	1.41	<0.01	<0.01	1,285	5,013
8		2.13	<0.01	<0.01	117	1.31	<0.01	<0.01	610	2,492
1	×	4.94	<0.01	<0.01	146					8,337
4	×	5.01	<0.01	<0.01	125					6,206
8	×	3.40	<0.01	<0.01	101					3,279

Panel B*: Size Distribution of Estimated Pension Effects*

	TVPI (bps)				
Min. Fund Age	Std. Dev	p60-p40	p80-p20	p90-p10	p95-p5
1	420	136	403	694	1,158
4	471	144	487	770	1,436
8	523	147	485	803	1,501

Notes: Panel A of the table shows regressions of the form $t_{pfg} = \alpha_f + \theta_p + \eta_{p,g} + \beta X_{pfg} + \varepsilon_{pfg}$, where t_{pfg} is the TVPI of pension p in fund f managed by general partner (GP) g. TVPI is defined as the total return on invested capital (market value plus cumulative distributions, scaled by total contribution). All regressions include fund fixed-effects. θ_p is a fixed effect for pension p (pension-effects) and $\eta_{p,g}$ is a pension-by-GP fixed effect. In the last three rows, the we include a vector of control variables X_{pfg} that is defined in Section 4.2. The columns headed with p list the p-values of the F-statistics that test for pension fixed effects. The column headed with $p*$ lists the p-values of the permutation test statistics. We drop singleton fixed effect groups in all regressions. Panel B of the table presents information on the distribution of the estimated pension effects from Panel A (rows 1-3). We use an empirical Bayes procedure to shrink the distribution of the estimated pension effects based on their standard errors. See Section 4.3 for more details.

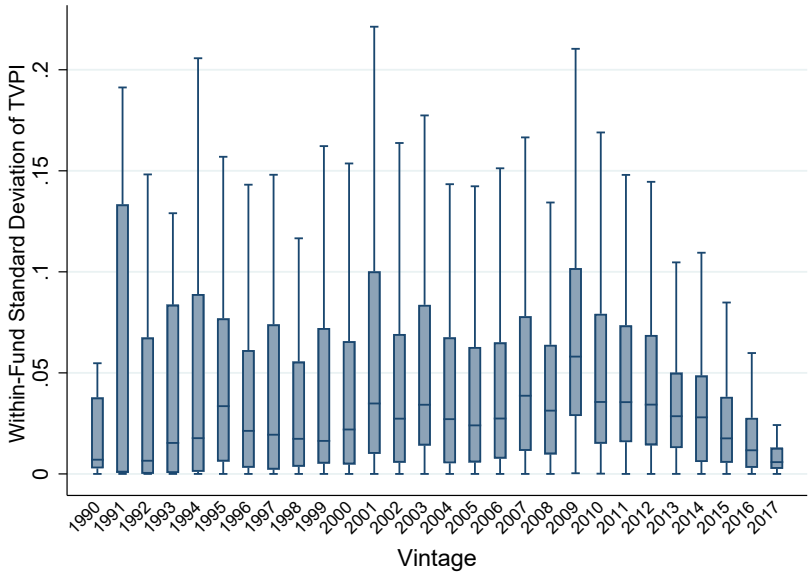
Table 4: Within-Fund Performance and Investor Characteristics

	Dependent. Variable: TVPI (bps)					
	(1)	(2)	(3)	(4)	(5)	(6)
Percent Invested in Fund	-1.63	-1.44	-1.83	-1.58	-2.52	-2.54
	(-0.83)	(-0.62)	(-0.81)	(-0.57)	(-0.92)	(-0.73)
Experience	1.67	0.09	-2.82	-5.85	-15.42	-44.33
	(0.13)	(0.01)	(-0.18)	(-0.30)	(-0.66)	(-1.54)
LP-GP Pairs (Full Sample)	7.39**	10.17**	8.02*	12.34**	5.45	9.07
	(2.17)	(2.70)	(1.84)	(2.41)	(0.85)	(1.09)
Large Investor	116.13**	89.74**	142.54**	110.89**	158.93**	158.54**
	(3.32)	(2.36)	(3.40)	(2.36)	(2.57)	(2.30)
Early-Stage Investor	30.21	22.13	36.86	28.16	18.10	-2.23
	(1.55)	(0.99)	(1.56)	(0.99)	(0.54)	(-0.05)
Board Size		-0.45		-3.41		-9.53**
		(-0.19)		(-1.18)		(-2.27)
Percent of Board Elected		1.43**		2.01**		2.76**
		(2.93)		(3.18)		(2.89)
Adjusted R^2	0.99	0.99	0.99	0.99	0.99	0.99
R^2-Within	0.00	0.01	0.00	0.01	0.00	0.01
Min. Investment Period (yrs)	1	1	4	4	8	8
Mean TVPI Shortfall (bps)	589	593	677	702	722	771
N	10,879	8,360	8,532	6,245	4,948	3,294

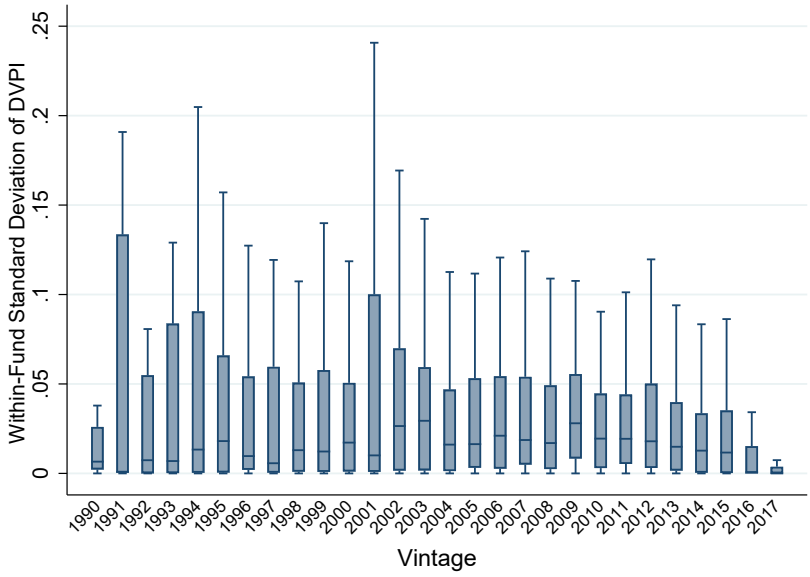
Notes: The table shows regressions of the form $t_{i,f} = \alpha_f + \beta' X_{i,f} + \varepsilon_{i,f}$, where $t_{i,f}$ is the TVPI of investor i in fund f. TVPI is defined as the total return on invested capital (market value plus cumulative distributions, scaled by total contribution). All regressions include fund fixed-effects and are clustered by investor×fund-vintage. The covariates are defined fully in Section 4.2. t-statistics are in parentheses. * indicates a $p < 0.1$ and ** indicates $p < 0.05$.

Figure 1: Within-Fund Standard Deviation of Net-of-fee Returns, by Fund Vintage

Panel A: Returns based on Realized Distributions and Remaining Net-Asset-Value (TVPI)

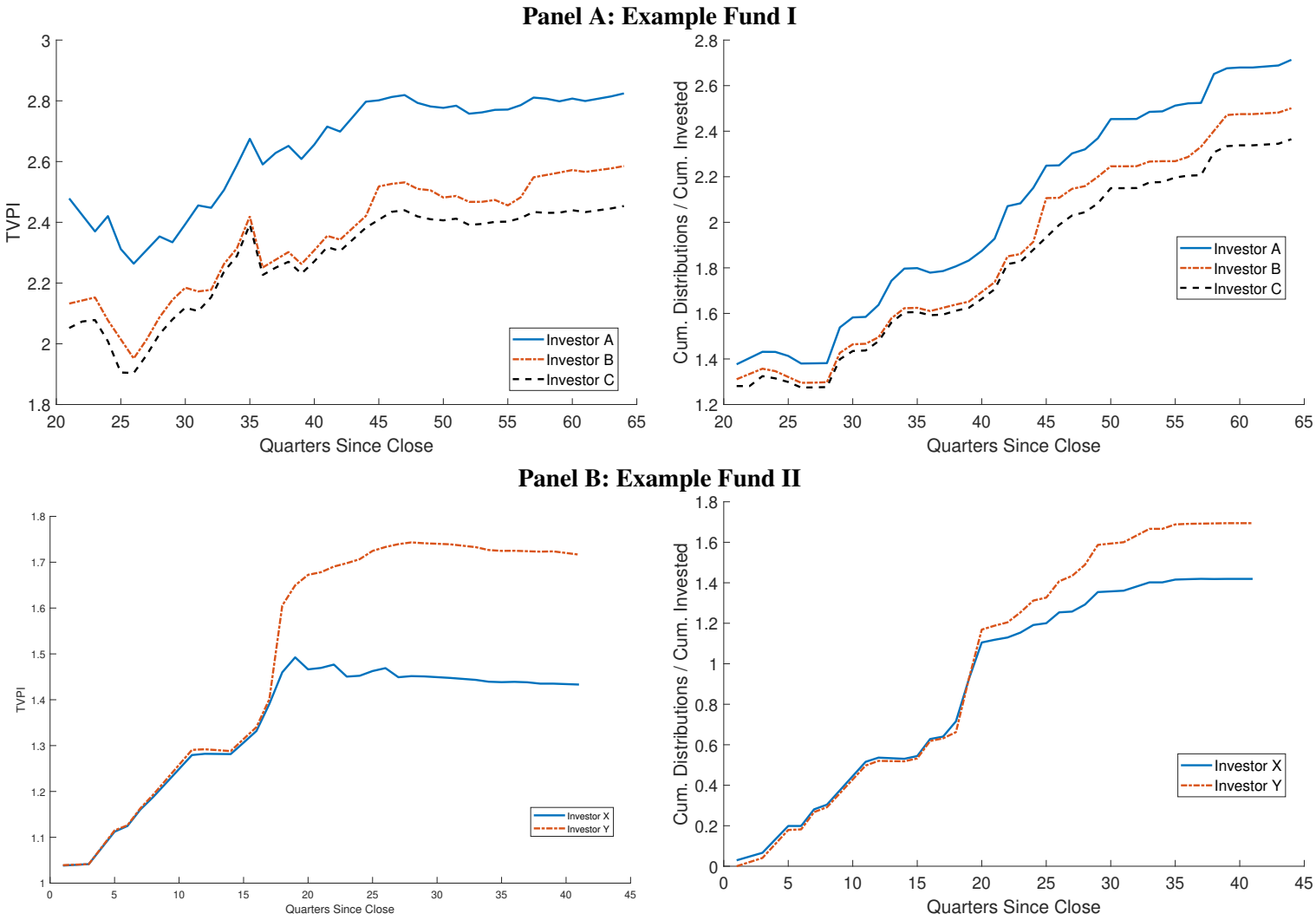


Panel B: Returns based only on Realized Distributions (DVPI)



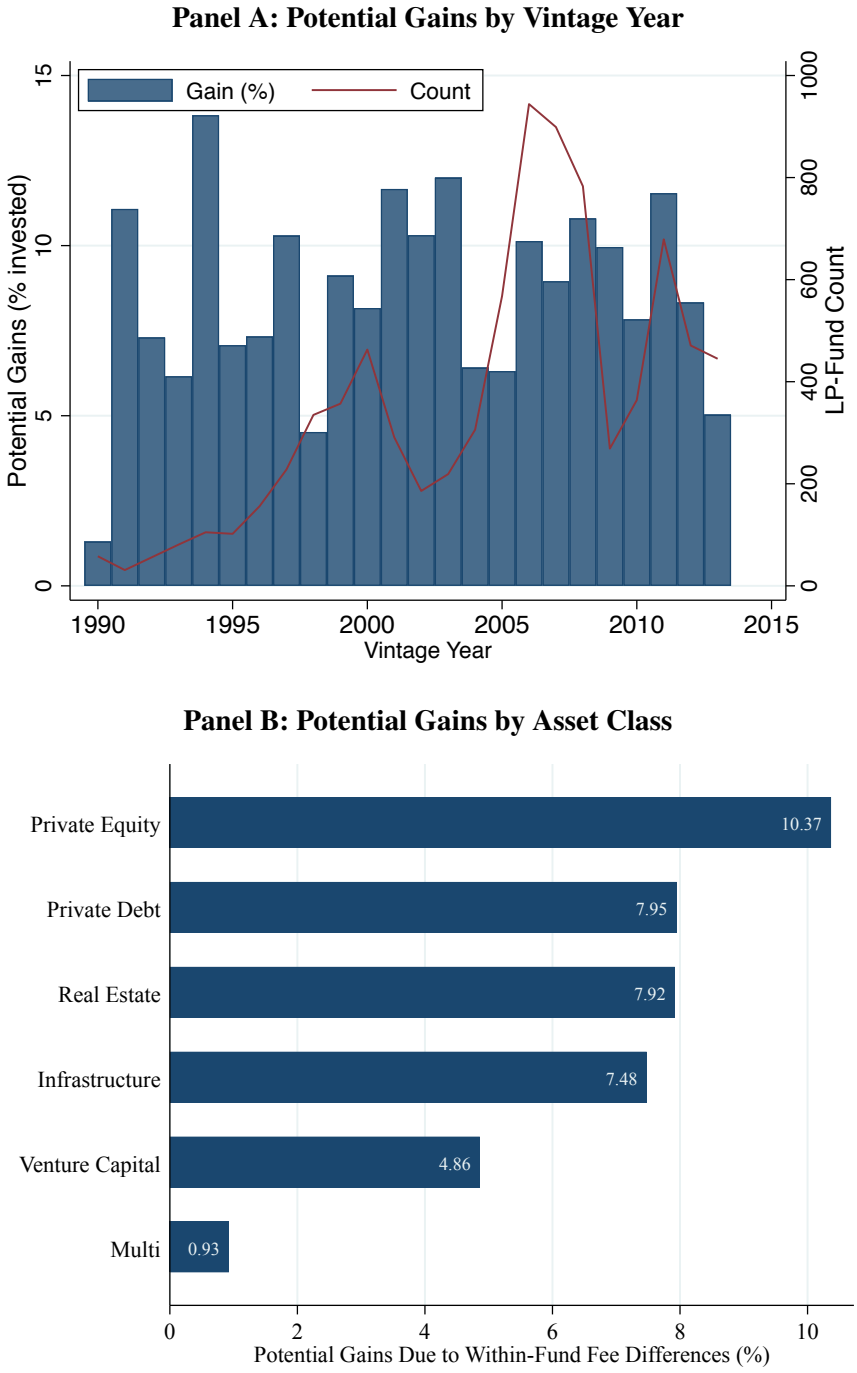
Notes: This plot shows the distribution of within-fund return dispersion across funds. For each fund, we compute the standard deviation of returns across investors, σ_f. The plot shows the distribution of σ_f broken down by the vintage year of f. See Section 2.3 for more details on our sample construction. Panel A and B of the figure show σ_f when measuring returns using TVPI and DVPI, respectively. TVPI is defined as the cumulative amount of distributions plus any remaining net-asset-value, scaled by the cumulative amount of contributions. DVPI is defined as the cumulative amount of distributions, scaled by the cumulative amount of contributions.

Figure 2: Examples of Within-Fund TVPI and DVPI Evolution



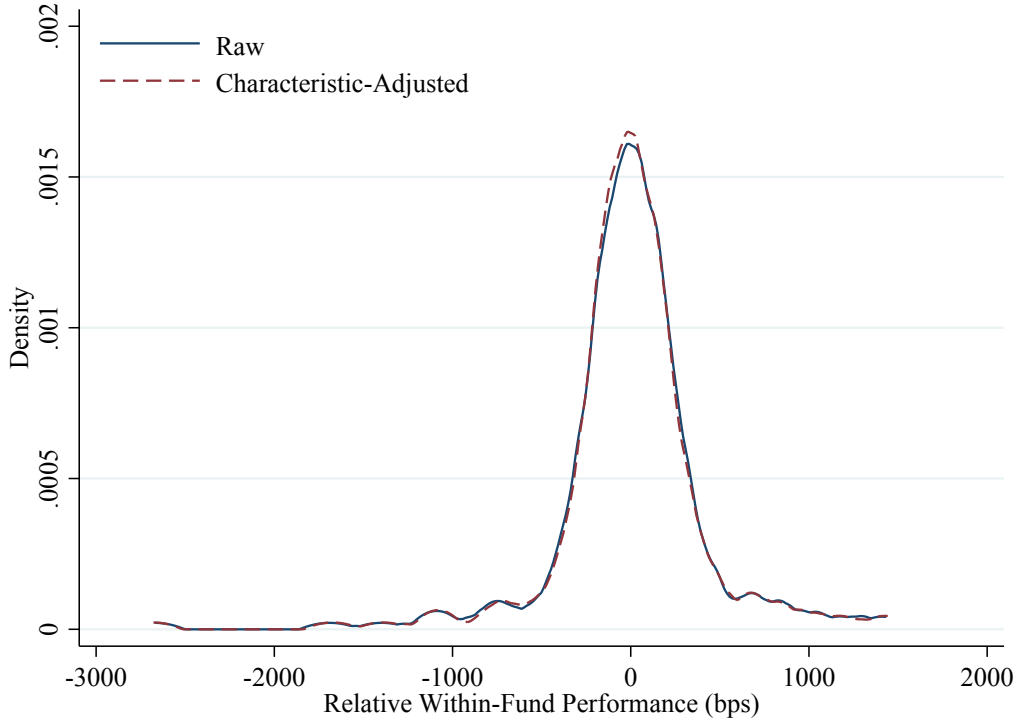
Notes: This plot shows sample TVPI (sum of cash distribution and fund value over invested capital) and DVPI (cash distribution over invested capital) paths for several investors in two different funds in our data. We are not able to identify individual funds or investors per our data-sharing agreement with Preqin, so we refer to the two funds in the plot as Fund I and Fund II. At each point in time t for a given investor p in fund f, is defined as the cumulative distributions received by p as of time t plus the remaining net-asset value of p's investment in fund f, scaled by the cumulative amount of contributions made by p into fund f as of time t.

Figure 3: Potential Gains Due to Fee Differences, by Vintage and Asset Type



Notes: This plot shows potential gains due to fee differences, broken down by vintage years and asset class. We compute the potential dollar gain for investor p in fund f by computing how much more in dollars investor p would have earned if it had received the best observed net-of-fee return in fund f. When aggregating across investor-fund pairs, we scale the dollar gains by the amount of capital contributed by the aggregation group. Thus, the potential gains can be interpreted as the % increase per $1 invested. In Panel A, we aggregate over funds of a given vintage. In Panel B, we aggregate over funds of a given asset type. Data from Preqin spans 1990 to 2019 and does not cover all state and local pensions. For both figures, we only consider investor-fund observations where the observed return is at least 5 years after the fund vintage in order to ensure adequate time has passed to observe fee differences. See Sections 3.1 and 3.2 for more details.

Figure 4: Distribution of Estimated Pension Effects

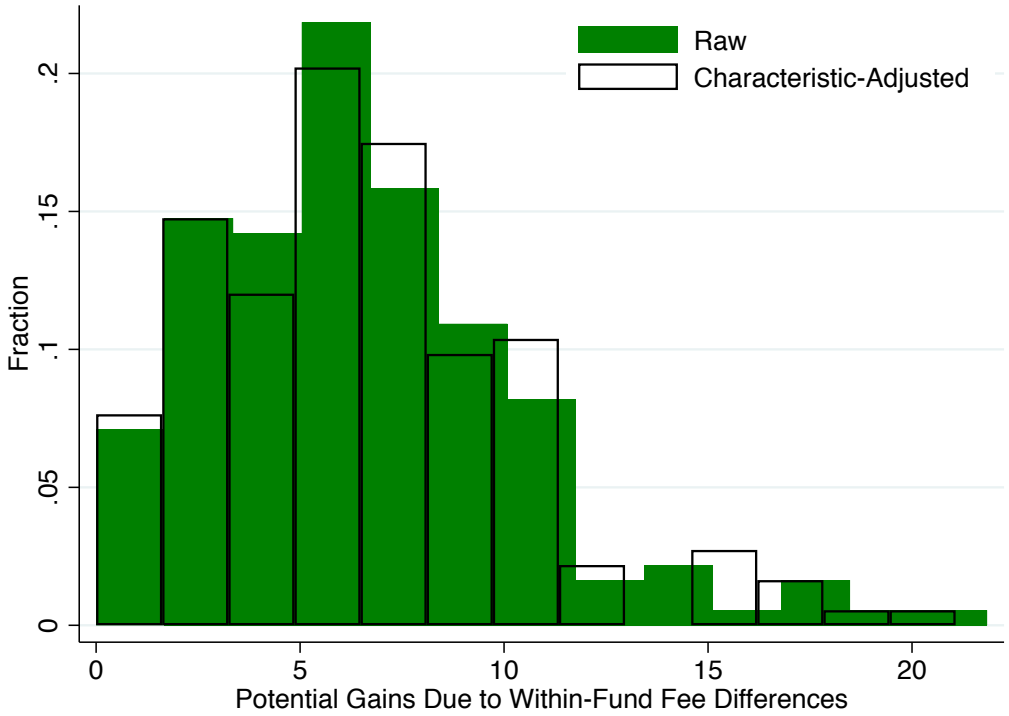


Notes: This figure shows the distribution of estimated pension-effects on within-fund performance. Pension effects are based on the following regression:

$$r_{pf} = \alpha_f + \eta_p + \varepsilon_{pf}$$

where r_{pf} is the TVPI of investor p in fund f. α_f are fund fixed effects and η_p are pension fixed effects. The figure plots the distribution of the estimated η's, after accounting for estimation error using an empirical Bayes procedure. The blue solid line shows the estimated pension effects with no other covariates and the red-dashed line shows their distribution after adjusting for pension-level characteristics. We adjust the pension effects for characteristics by including size, the number of relationships between the pension and the general partner of fund f, the number of investments to-date by pension p as of the year of fund f, the percent of fund f accounted for by investor p, and a dummy variable proxying for whether p is a first-close investor in fund f. See Section 4.3 for a more detailed description of the empirical Bayes procedure and the characteristic-adjustment. Data from Preqin spans 1990 to 2019 and does not cover all state and local pensions.

Figure 5: The Distribution of Potential Gains Across U.S. Pensions

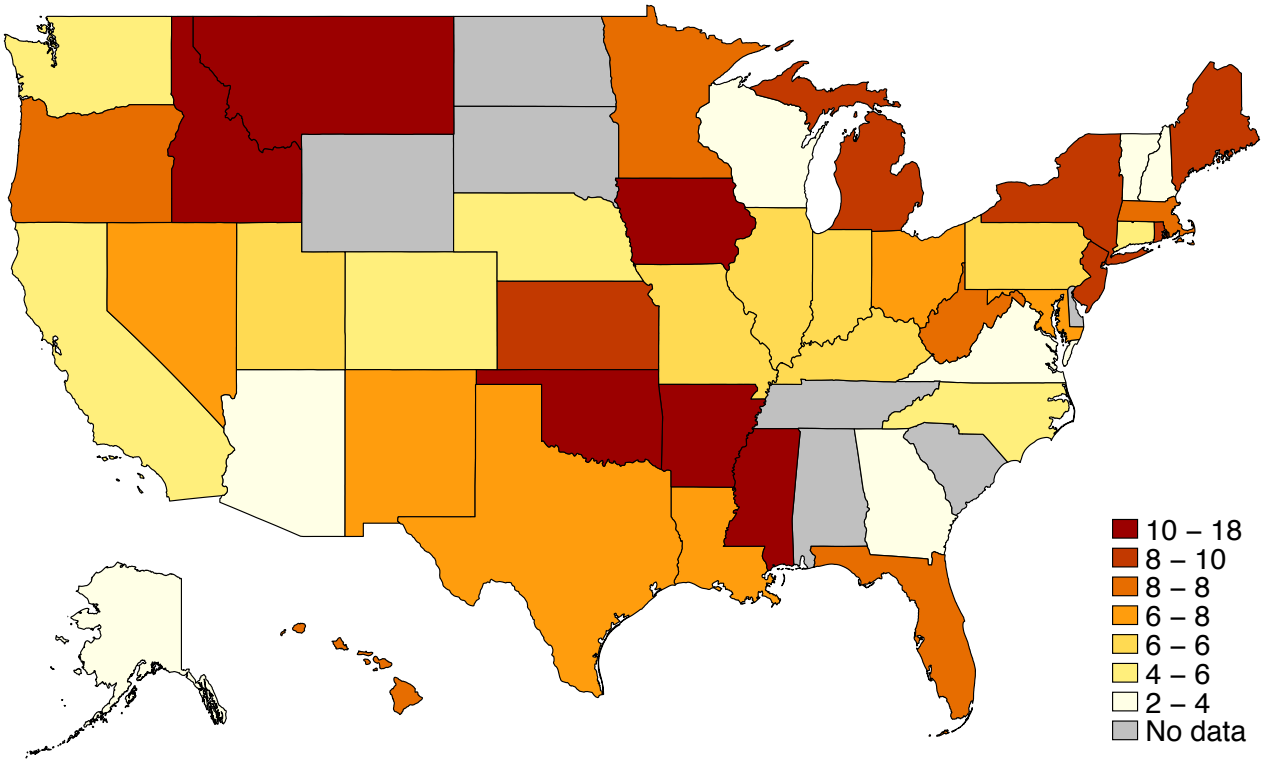


Notes: This figure shows the distribution of potential return gains due to within-fund dispersion in contracts across pensions. Within a fund f, we compute the potential return gain of investor p as the difference between the maximum observed TVPI in fund f and investor p's actual return in the fund. We then compute potential dollar gains for p in f by scaling the potential gain by the amount invested. To aggregate potential gains for investor p, we then sum up potential dollar gains across all of its funds and divide it by the total amount of invested capital by p in our data. We call this measure G_p and the plot shows the distribution of G_p across investors, both in raw terms and adjusted for characteristics. To adjust for characteristics, we run the following regression:

$$r_{pf} = \alpha_f + \beta X_{p,f} + \varepsilon_{pf}$$

where r_{pf} is the TVPI for investor p in fund f and X_{pf} is a vector of characteristic . We use the residuals ε_{pf} to compute our gain measure G_p and then overlay them on the plot. See Section 4.3 for more details. Pensions must have invested in at least 3 different funds in order to be included in the graph. Data from Preqin spans 1990 to 2019 and does not cover all state and local pensions.

Figure 6: Characteristic-Adjusted Potential Gains by State

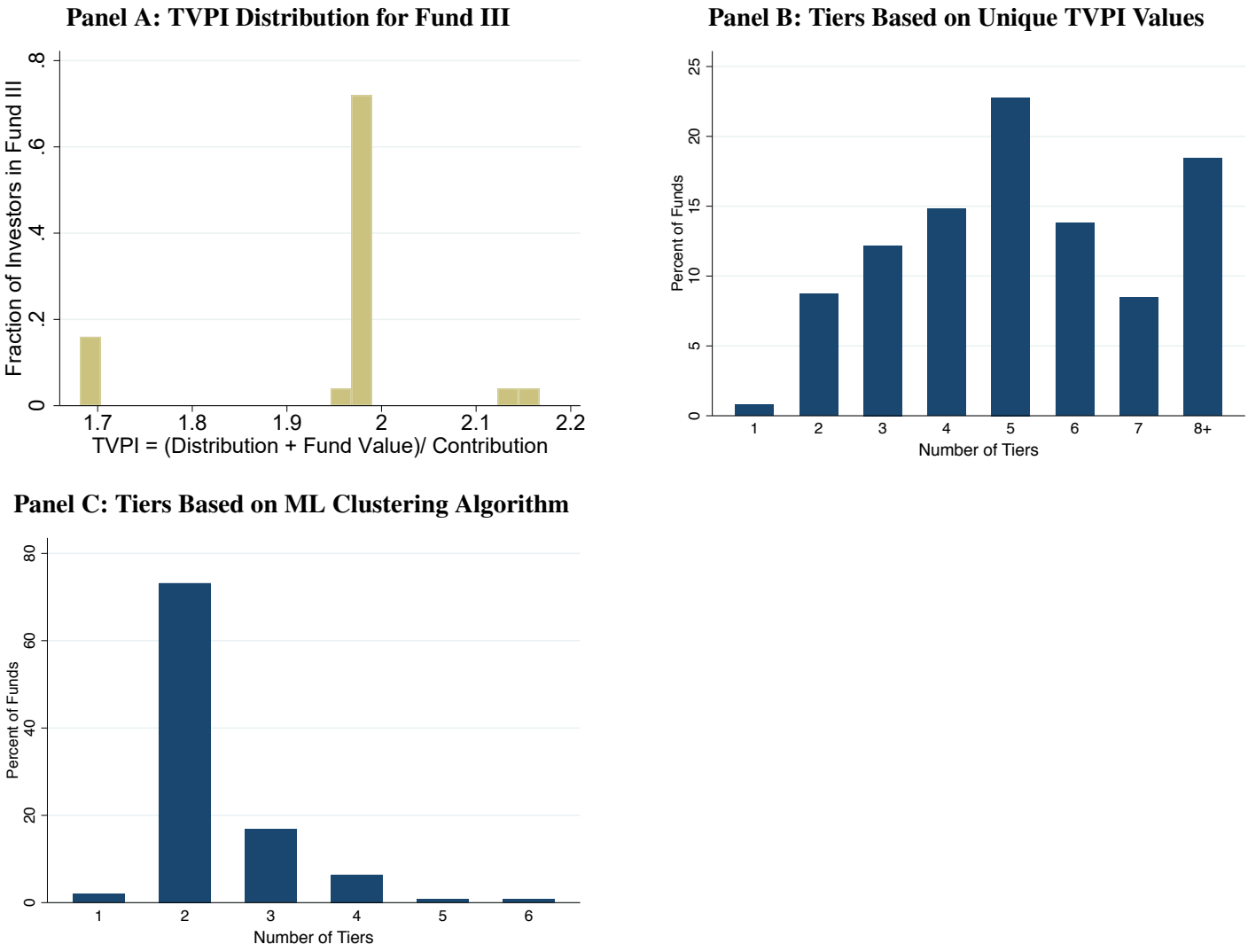


Notes: This figure shows the distribution of potential return gains due to within-fund dispersion in contracts across pensions, averaged at the state level. Within a fund f, we compute the potential return gain of investor p as the difference between the maximum observed TVPI in fund f and investor p's actual return in the fund. We then compute potential dollar gains for p in f by scaling the potential gain by the amount invested. To aggregate potential gains for investor p, we then sum up potential dollar gains across all of its funds and divide it by the total amount of invested capital by p in our data. We call this measure G_p and the plot shows the distribution of G_p across investors, both in raw terms and adjusted for characteristics. To adjust for characteristics, we run the following regression:

$$r_{pf} = \alpha_f + \beta X_{p,f} + \varepsilon_{pf}$$

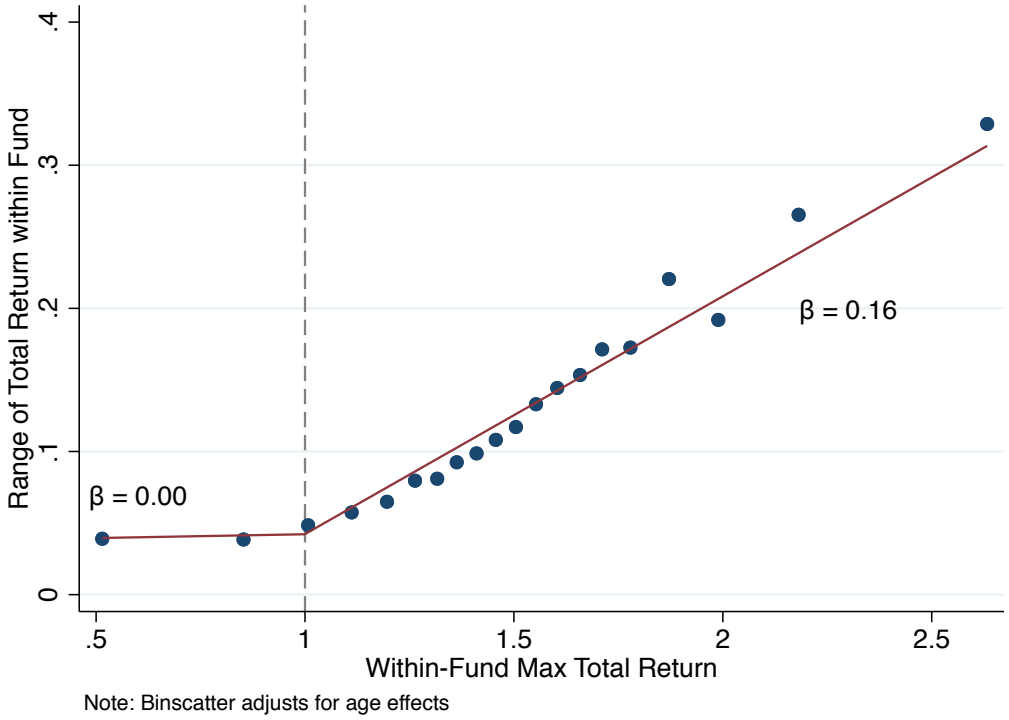
where r_{pf} is the TVPI for investor p in fund f and X_{pf} is a vector of characteristics. We compute potential gains using the residuals ε_{pf} and then take an equal-weighted average to aggregate to the state level. See Section 4.3 for more details. Pensions must have invested in at least 3 different funds in order to be included in the graph. Data from Preqin spans 1990 to 2019 and does not cover all state and local pensions.

Figure 7: Investor Tiering



Notes: Panel A of the figure shows the distribution of TVPI for twenty investors in an anonymized fund (Fund III) in the core sample. Panel B of the figure shows the distribution of tiers across all funds in the core sample, where tiers are defined as the unique number of TVPIs (rounded to three decimals). In Panel C, we instead classify tiers using a machine learning clustering algorithm (see Section 5.1 for more details). Funds must have at least five investors to be included in Panel B or Panel C. Data from Preqin spans 1990 to 2019 and does not cover all state and local pensions.

Figure 8: Within-Fund Dispersion and Fund Returns



Notes: This figure plots the within-fund range of returns against the maximum observed return in a given fund f at each point in time. For each fund f and date t, we compute the range of returns within the fund (Δ_{ft}) against the maximum observed return in the fund ($\hat{g}_{ft}$). The figure is a binscatter of Δ_{ft} against $\hat{g}_{ft}$, after controlling for vintage-year dummies and a linear adjustment for age. To create the binscatter, we separately regress Δ_{ft} and $\hat{g}_{ft}$ on fixed-effects for fund f's vintage year and the number of years since the fund's inception. We then divide the $\hat{g}_{ft}$-residuals into 20 equal-sized groups and plot the means of the Δ_{ft}-residuals within each bin against the means of the $\hat{g}_{ft}$-residuals. We add back the sample means to both the x and y-variables to facilitate interpretation of the units. The solid line in the graph shows the best linear fit estimated on the underlying data using OLS, where we allow the slope to vary depending on whether $\hat{g}_{ft}$ is above or below 1. Data from Preqin spans 1990 to 2019 and does not cover all state and local pensions.