# The Economic Burden of Pension Shortfalls: Evidence from House Prices*

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**Abstract**

U.S. state pensions are underfunded by trillions of dollars, but their economic burden is unclear. In a model of inefficient taxation, real estate fully reflects the cost of pension shortfalls when it is the only form of immobile capital. We study the effect of pension shortfalls on real estate values at state borders, where labor and physical capital could more easily relocate to a state with a smaller shortfall. Using plausibly exogenous variation driven by pension asset returns, we find that one dollar of pension underfunding reduces house prices near state borders by approximately two dollars. Our estimates imply a deadweight loss associated with addressing pension shortfalls that is consistent with prior research in settings with high returns to public spending and costs of taxation.

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*"Moody's Investors Service estimates state and local pensions have unfunded liabilities of about $4 trillion, roughly equal to the economy of Germany, the world's fourth-largest economy."*
— The Wall Street Journal, July 30, 2018[1]

Underfunded public pensions in the U.S. represent an implicit household liability larger than auto loans, student debt, and credit card balances combined.[2] Despite popular concern about pension shortfalls, little is known about the economic costs associated with addressing them.[3] Several factors complicate the estimation of these costs, including the long maturity of pension liabilities, the uncertain nature of pension asset returns, and the complex politics affecting the size of pension promises and fund contributions.

A model of inefficient taxation motivates our empirical strategy to identify the economic burden of pension shortfalls. In an open economy, where capital and labor are mobile but real estate is not, house prices reflect the total cost of pension shortfalls including any inefficiencies or deadweight loss generated in honoring the obligations (e.g., Oates, 1969; Bradford, 1978; Kotlikoff and Summers, 1987; Harberger, 1995). Conversely, if all forms of capital face a high cost of relocation, then the burden is unclear and the price on any individual asset is unlikely to reflect the true implied cost. Motivated by this insight, we focus our analysis on locations near state borders where immobile factors, such as real estate, should bear the burden and thus reflect the implied cost of pension shortfalls. Since the level of pension underfunding may be affected by local economic conditions and the endogenous response to past contributions and investment decisions, we use

[1] "The Pension Hole for U.S. Cities and States is the Size of Germany's Economy", available from: https://www.wsj.com/articles/the-pension-hole-for-u-s-cities-and-states-is-the-size-of-japans-economy-1532972501.

[2] The state and local pension systems in the U.S. reported $1.378 trillion in unfunded liabilities in fiscal year 2015, but according to Rauh (2016), who discounts cash flows using the Treasury curve to reflect the low risk of pension promises, the accumulated deficit is $3.846 trillion. According to the Federal Reserve Bank of New York, as of the fourth quarter of 2020, outstanding amounts of student loan, auto loan, and credit card debt are $1.55, $1.37, and $0.82 trillion, respectively, totaling to $3.74 trillion.

[3] As evidence that the general public is interested in pension funding, Figure 1 shows a correlation of 0.65 between state pension underfunding and state-level Google search intensity for "public pension" and "pension crisis."

plausibly exogenous variation in pension asset returns or “windfalls” to identify the causal effect of pension shortfalls on house prices.

Our main findings illustrate a consistent pattern: changes in pension funding pass through to house prices in border counties and reflect a deadweight loss of addressing shortfalls. We estimate that the marginal home buyer is willing to pay approximately two dollars for each dollar of additional pension funding per property. This house price pass-through is comparable to the estimated impact of public spending on infrastructure and school salaries (e.g., Cellini, Ferreira, and Rothstein, 2010; Bayer, Blair, and Whaley, 2020).[4] Our theoretical framework shows that this pass-through can be interpreted as a tax multiplier with respect to the funding of future liabilities – for each additional dollar that states will have to raise through future taxes or amenity reductions, households perceive an economic burden of approximately two dollars. Controlling for rental prices slightly decreases our pass-through estimates, which suggests that pension shortfalls affect house prices primarily through the capitalization of future costs and only partially through the current quality of amenities.

We face two major challenges in our empirical analysis. First, where businesses and individuals cannot easily avoid the future taxation associated with pension shortfalls, the burden of taxes will be spread across all assets (i.e., human and physical capital as well as real estate). Thus, we conduct our analysis at state borders, where businesses and individuals face a much lower cost of relocating to a state with a smaller pension shortfall.[5] This allows us to measure the economic cost of pension shortfalls in areas where landowners are expected to bear the brunt of any changes to taxes and the provision of government services.

[4]These estimates reflect spending on potentially “high value” projects. In this paper we study windfalls, so the estimated pass-through reflects the marginal value of the best unfunded projects or the total burden of taxation.

[5]According to Rauh (2016), state pension plans account for \$4.05 trillion (84%) of the \$4.80 trillion total reported pension liability, so our analysis captures most of the U.S. public pension burden. City and county borders would also be natural settings if not for a lack of data on local government pensions, except for the largest municipalities, that precludes our empirical strategy.

This focus on state borders is one of the key features that distinguishes our paper from earlier studies on the relation between pension funding and house prices in individual cities or states (e.g., Epple and Schipper, 1981; Leeds, 1985; Hur, 2008; Albrecht, 2012; MacKay, 2014; Stadelmann and Eichenberger, 2014). These studies do not focus on border areas where real estate is the only immobile asset, so their estimates do not reflect the full economic cost of pension shortfalls. Thus, we answer a fundamentally different question from these earlier papers, using housing markets as a laboratory to measure an economic primitive rather than as the outcome of interest.[6]

Second, pension underfunding is likely to be correlated with omitted variables. If shortfalls were accrued in efforts to improve amenities for state residents, then states with high shortfalls would provide better services than states with low shortfalls. Alternatively, if generous pensions are the result of overinvestment in poorly performing projects, shortfalls may be associated with worse quality of life. Our question requires exogenous variation to identify the causal effect of shortfalls, so we focus on the effect of pension asset returns on house prices. We refine this approach to account for potential "home bias" or "familiarity bias" in pension investments by restricting attention to benchmark returns or unexpected excess returns over these benchmarks.

In our baseline analysis, we compare the pension asset returns in the early part of our sample (2002–2014) with home prices thereafter for properties in county clusters across state borders. We find that increases in raw returns, excess returns, and benchmark returns implied by asset allocations are all associated with increased house prices. To quantify the effect of pension shortfalls, we calculate cumulative dollar pension returns based on 2001 pension assets and find a pass-through of approximately two—for each dollar of pension asset returns, house prices increase by approximately two dollars. Nonparametric analysis of the border discontinuity reveals

[6] In fact, in an analysis of properties in the interior of the state, we find disperse and inconsistent estimates, which is exactly what we would expect. In such settings the relative burden should be split among various forms of capital in a way that depends on their relative mobility and elasticities and is likely to vary across regions.

a clear increase in prices when moving from a low-return state to a high-return state.

Our estimates are robust to a battery of alternative specifications. First, we consider asset returns between 2002 and a property's sale year, instead of using the same return horizon for all houses, and find a similar pass-through of approximately two. The benefit of this approach is that it allows us to include property fixed effects among properties with repeat sales. This alleviates the concern that our findings could be driven by time-invariant factors at the state, local, or property level. Focusing on the sub-sample of repeat sales provides a slightly lower pass-through estimate. This is not surprising as requiring repeat sales on a property moves the average transaction date forward in time, leaving less time on average between 2002 and the sale. Prior work has shown it can take several years for even things like public spending on schools to be fully realized in house prices (e.g., Bayer, Blair, and Whaley, 2020). More importantly, the inclusion of property fixed effects has little effect on the overall pass-through in the repeat sales sample, which suggests that unobservable time-invariant factors are not biasing our estimates.

In supporting our focus on windfalls from pension returns instead of the level of pension shortfalls, we provide novel evidence that pension investment returns crowd out pension fund contributions. Specifically, we show that each dollar of excess pension returns is associated with a 55 cent reduction in contributions. To our knowledge, we are the first to document this mechanism in the academic literature. This finding highlights both the benefits of our empirical strategy, which would be biased by a focus on the level of shortfalls, as well as the political incentives that are likely responsible for the pension underfunding crisis.

To shed light on the economic channels driving our results, we turn our attention to the quality of government amenities and find that higher state-level pension shortfalls correlate with less educational spending and poorer road quality. This raises a question: do our estimates reflect worse current amenities in places with low pension asset returns, or an expectation of future costs that are capitalized in housing prices? To address this, we add time-varying local rental prices to the set of control variables and find slightly attenuated results when measuring windfalls over the entire sample period. While reaffirming that our results are not endogenously related to economic conditions, this test also narrows the interpretation of our pass-through estimates: the economic burden of pension shortfalls primarily reflects expectations about future increases in taxes or decreases in the quality of amenities.[7] Finally, we show that the estimated economic burden is significantly higher in states that currently have higher income tax rates, consistent with these states facing larger distortions associated with increasing tax rates to fund pension obligations in the future.

This paper contributes to the literature on the real effects of public finance. An emerging segment of this literature focuses on the condition of state and local pensions in the United States. Earlier work in this area has focused on the measurement of the pension underfunding (Brown and Wilcox, 2009; Novy-Marx and Rauh, 2011; Novy-Marx and Rauh, 2014), the political economy of pension funding (Brinkman, Coen-Pirani, and Sieg, 2018; Myers, 2021), and the impact of pension funding on municipal borrowing costs (Novy-Marx and Rauh, 2012; Boyer, 2020), the precautionary savings of households (Zhang, 2021), and the economic recovery after the financial crisis (Shoag, 2013). We complement this work by estimating the effect of pension shortfalls on house prices near state borders to quantify the current economic impact of this future burden.

Our theoretical model highlights the analogy between this paper and earlier studies of inefficient taxation. In equilibrium, the effect of an exogenous increase in pension assets is equal to the

[7] An implication of this argument is that marginal home buyers are aware of the condition of local government finances to the extent that they anticipate future tax hikes or reduced service provision. Since prices are based on common signals such as comparable recent property transactions, this does not require that all households are aware but only that marginal ones are. Although we cannot provide direct evidence on how households become informed about pension underfunding, Google search trends (Figure 1) reveal that residents of states with worse pension shortfalls are more interested in the issue. Underprovision of maintenance (e.g., poor roads) might signal information about future conditions even though it provides little disamenity in the present. More directly, homeowners could become informed from news coverage of state fiscal issues.

present value of the tax multiplier, which means the pass-through of pension shortfalls to house prices is comparable to other estimates of the economic burden of raising taxes. Consistent with our estimates, Cellini, Ferreira, and Rothstein (2010) and Bayer, Blair, and Whaley (2020) find economic costs between one and two dollars per dollar of tax revenue in the context of public school spending.

The empirical finding that unfunded pension liabilities are capitalized in house prices is indicative of households' forward-looking behavior. This evidence is also relevant to the applicability of Ricardian equivalence (Barro, 1974). Although our goal is not to test whether Ricardian equivalence holds, our findings are consistent with households internalizing the government's future budget constraint in their optimization problems.

Finally, our results have implications for the political economy of public sector employee compensation. Fitzpatrick (2015) estimates that public school employees in Illinois are willing to pay 20 cents on average for a one dollar increase in the present value of retirement benefits. We find that households perceive substantial deadweight loss associated with addressing pension shortfalls. In combination, these findings question the efficiency of governments promising generous retirement benefits to employees who do not value them and imposing the burden of funding those promises on households who view them as costly.

The remainder of the paper is organized as follows. Section 1 presents a model of tax burden to motivate our empirical analysis. Section 2 describes our data on public pension funding and house transaction prices. Section 3 explains our identification strategy. Section 4 presents the main results. Section 5 concludes.

# 1 Theoretical Framework

In this section, we present a theoretical framework that motivates our research question and identi cation strategy. We rst show that in an open economy, landowners of a state within a country are likely to bear the burden of a tax levied on a domestically mobile factor, motivating our empirical design. We then show that the pass-through of pension shortfalls to house prices is theoretically ambiguous and therefore an empirical question that re ects ine ciency in the public provision of goods and capital raising.

## 1.1 Tax burden in an open economy

Harberger (1962) argues that in a closed economy, the burden of the corporate income tax tends to fall entirely on physical capital. Importantly, in a closed economy, untaxed factors always bear some burden of the tax if the taxed factor's supply (demand) is not perfectly inelastic (elastic).[8] Relaxing the closed-economy assumption, most studies argue that in an open economy, immobile factors bear most, if not all, of the long-run burden of the tax in the economy due to capital mobility across borders.[9] Thus, it is critical for our empirical design to focus on an open-economy setting at state borders to measure the burden of pension shortfalls.

In Appendix A.2, we provide a simple framework based on Kotliko and Summers (1987) to illustrate this point. There are two factors of production for the single good in the economy: capital and land. Following Harberger (1962), we assume perfect competition and a xed national capital stock that is perfectly mobile within the country. For simplicity, we assume that the factor complementary to capital, here labeled land, is supplied inelastically and is immobile. Since

[8]In Appendix A.1, we present a simple closed-economy framework to illustrate this point.

[9]Notable examples include Bradford (1978), Kotliko and Summers (1987), Mutti and Grubert (1985), Harberger (1995), and Gravelle and Smetters (2001). See Gravelle (2013) for a review of tax burden in general equilibrium.

capital is mobile, rental rates on capital must be equalized across states: a tax imposed on income earned by capital in a state is not fully borne by the capital initially located in the state imposing the tax. In contrast, landowners in the two states are differentially impacted: there is a loss of rental income in the state imposing the tax on capital and a gain in the other state. We summarize the main takeaway of the open-economy model in Proposition 1.

**Proposition 1.** *In an open economy, the immobile factor in a state is likely to bear a significant portion of the burden of a tax that the state levies on a domestically mobile factor.*

*Proof.* See Appendix A.4. □

## 1.2 Pension shortfalls, tax distortions, and property values

The previous section establishes that an open economy is the appropriate setting for our empirical analysis. In this section, we study the capitalization of future pension liabilities in current house prices. Whereas the previous section focused on capital mobility and the elasticity of demand, this section introduces a role for asset prices. The economic burden of a tax is affected by changes in asset prices due to the discounted present values of future tax and public expenditure changes. We argue that the magnitude of the marginal decrease in house prices from an additional dollar of pension shortfalls is theoretically ambiguous and therefore an empirical question.

The model presented here is based on a slight modification of the asset-price approach to tax burden presented in Poterba (1984). The key component of the burden is the price change for existing owner-occupied homes due to the change in the present value of future taxes associated with the asset. The stock of houses is assumed to be fixed in the short run, so the equilibrium rental rate equates the demanded quantity with the existing housing services flow. Denote the market-clearing rental rate by $R(H)$ with $R' < 0$, where $R$ is the inverse demand function for housing services. $R(H)$ represents the marginal benefit of housing services.

Households consume housing services until the marginal value of these services equals their marginal cost. We assume all houses incur depreciation at a constant rate per period, maintenance costs equal to a fraction of the current value, and property taxes at a rate. All households face a marginal income tax rate, can deduct property taxes from taxable income, and can borrow and lend at the nominal interest rate. The cost also includes any capital gain or loss of holding the asset. Let $q_{H;t}$ be the house price at the start of period $t$, so $.q_{H;t+1} * q_{H;t}/$ represents the capital gain or loss during period $t$. In equilibrium, homeowners equalize the marginal cost and marginal bene t of housing services:

$$R.H_t/ = \quad q_{H;t} * .q_{H;t+1} * q_{H;t}/ ; \tag{1}$$

where $r \quad + \quad + .1 * \quad /.r + \quad /$.

Consider a tax on each house that takes the form of a lump-sum payment $T_t$ to cover the unfunded pension liability $L_t$ in period $t$.[10] We assume taxes induce a deadweight loss,

$$T_t = .1 + \quad /L_t ; \tag{2}$$

where $\quad > 0$. This means that to fund each additional dollar of pension liability in period $t$, the state has to raise more than one dollar in taxes. Parameter ects the cost of raising revenues, which we later estimate empirically.

Because houses are durable assets, future taxes can still depress prices today. In each period when the tax is levied, the equilibrium condition (1) becomes

$$R.H_t/ * T_t = \quad q_{H;t} * .q_{H;t+1} * q_{H;t}/ : \tag{3}$$

[10] The tax payment $T_t$ is isomorphic to reducing current amenities to cover the liability $L_t$.

Since $q_{H,t+1}$ is unknown at time $t$, we can solve the price $q_{H,t}$ forward by rewriting (3) as

$$q_{H,t} = \frac{R(H_t) - T_t + q_{H,t+1}}{1+}. \tag{4}$$

Iterating Equation (4) forward and applying the no-bubble condition,[11] the distortionary tax assumption in (2) gives

$$q_{H,t} = \sum_{j=0}^{\infty} \frac{R(H_{t+j})}{(1+\ )^{j+1}} - \sum_{j=0}^{\infty} \frac{(1+\ )L_{t+j}}{(1+\ )^{j+1}}. \tag{5}$$

The second term in Equation (5) is the present value of current and future tax payments to cover pension liabilities. It is clear from (5) that if two states face the same housing demand curves but have different levels of liabilities $L$, all else equal, the one with a higher present value of pension obligations will have lower house prices today.

If the stock of housing is fixed[12] (i.e., $H_{t+j} = H_t$ for all $j$), then from Equation (5) we can determine the impact of an unfunded liability $j$ periods ahead on house prices today:

$$\frac{dq_{H,t}}{dL_{t+j}} = -\frac{1+\ }{(1+\ )^{j+1}} < 0. \tag{6}$$

With reasonable parameter values for income and property tax rates, depreciation, and maintenance costs, the capitalization of future pension liabilities in house prices today can have a magnitude of less or greater than one. It depends on how large the distortion is and how far in the future the tax is imposed. We summarize the main message in Proposition 2.

[11] The transversality (no-bubble) condition in our setting is $\lim_{J\to\infty} \frac{q_{H,t+J}}{(1+\ )^{J+1}} = 0$, which rules out exploding house prices. This condition is consistent with Giglio, Maggiori, and Stroebel (2016), who find no evidence of violations of the transversality condition in the U.K. and Singapore housing markets, even during periods when housing bubbles were thought to be present.

[12] With an endogenous housing stock, changes in future taxes induced by future pension liabilities will also affect current and future investment in housing construction and the stock of housing $\{H_{t+j}\}_{j=0}^{\infty}$. In general, the effect of housing stock adjustments can mitigate the effect of taxes on house prices. See Poterba (1984) for more details.

Proposition 2. The magnitude of the marginal decrease in current house prices from an additional dollar of pension shortfalls is ambiguous; it can be smaller or larger than one.

Proof. See Appendix A.4. □

# 2 Data

## 2.1 State and local public pension plans database

We obtain accounting and actuarial data for state and local pension plans from the Public Plans Database (PPD) from the Center for Retirement Research at Boston College. PPD contains annual plan-level data from 2001 through 2019 for 190 pension plans: 114 administered at the state level and 76 administered locally. This sample covers 95% of public pension membership and assets nationwide.[13] The PPD is updated each spring from data available in the most recent Comprehensive Annual Financial Reports (CAFRs) and Actuarial Valuations (AVs). Intermediate updates may occur when new variables are added or data errors are corrected.

We use the PPD data to calculate the plan-level pension shortfall de ned as the actuarial accrued liabilities less the market value of assets. Actuarial accrued liabilities, measured under traditional Governmental Accounting Standards Board (GASB) 25 standards, are equal to the present value of future bene ts, discounted using the plan's assumed long-term investment return.

## 2.2 Detailed investment data by asset class

The PPD includes detailed annual data on each plan's speci c asset allocations, returns by asset class, and the associated benchmark returns. The asset classes in the PPD are based on the

[13] The PPD sample is carried over from the Public Fund Survey (PFS), which was constructed with an emphasis on the largest state-administered plans in each state, but also includes some large local plans such as New York City ERS and Chicago Teachers. See https://publicplansdata.org/ for more details.

categories reported by plans. We use these data to calculate the cumulative pension plan returns used as instruments for pension shortfalls.[14]

Table 1 reports descriptive statistics on the PPD data. On average across time and funds, the largest asset holdings were equities and fixed income (53% and 28% of total assets, respectively), followed by real estate and private equity (6% and 5% of total assets, respectively). The value of assets is 79% of the actuarial value of liabilities for the mean observation, indicative of substantial underfunding. Appendix Figure A.1 shows that the average ratio of pension assets to liabilities declined from just above 100% in 2001 to 76.4% in 2019, reflecting an increase in underfunding over the period we study.

As discussed in Novy-Marx and Rauh (2011) and Rauh (2016), public pension liabilities are effectively risk-free, so the appropriate discount rate for valuing them is the yield on a zero-coupon Treasury bond with the same duration. To discount pension liabilities using Treasury rates, we need to calculate the duration and convexity of each plan. Unfortunately, the information necessary for this calculation is unavailable in the PPD database prior to changes in pension reporting standards in 2014.[15] Therefore, to adjust the liability discount rate we use the aggregate adjustment factor in Rauh (2016) and inflate unfunded liabilities by a constant factor of 2.86.[16] While we acknowledge this is an imperfect adjustment method, any resulting bias would affect only our analysis of shortfalls and not the analysis that exploits variation in pension asset returns.

[14]The pension return data in the PPD have been used in academic research by Lu, Pritsker, Zlate, Anadu, and Bohn (2019), among others.

[15]Under new GASB 67 guidelines, plans are required to disclose their total pension liabilities (TPL) under alternative scenarios of the discount rate being 100 bps higher ($TPL_{+1}$) and 100 bps lower ($TPL_{-1}$). However, this information is only available starting in fiscal year 2014, when GASB 67 became effective.

[16]In fiscal year 2014, the state and local pension systems in the United States reported aggregate unfunded pension liabilities of \$1.19 trillion under GASB 67. Rauh (2016) applies a correction on a plan-by-plan basis that results in aggregate unfunded accumulated benefits of \$3.41 trillion under Treasury yield discounting. This implies an average adjustment factor of $3.412/1.191 = 2.864$.

## 2.3 Zillow transaction and assessment database

We obtain property-level data from the Zillow Transaction and Assessment Dataset (ZTRAX). ZTRAX is, to the best of our knowledge, the largest national real estate database, with information on more than 374 million detailed public records across 2,750 U.S. counties. It also includes detailed assessor data including property characteristics, geographic information, and valuations on over 200 million parcels in over 3,100 counties. These data have been used by Bernstein, Gustafson, and Lewis (2019), among others.

We filter the Zillow data in three ways. First we retain only residential property transactions for which the price of the transaction is verified by the closing documents as being between the typical home values in the the bottom and top market tiers as calculated at the county-month level by the Zillow Home Value Index (ZHVI).[17] Second, we focus only on single-family residences. Third, in our primary empirical analysis we restrict attention to properties located in counties sharing a border with an adjacent state and are located within 50 miles of the border.

# 3 Empirical Methodology

The objective of this paper is to estimate the economic burden of temporally distant and uncertain public liabilities. We focus on state pension shortfalls because of the growing concern about their magnitude (Rauh, 2016). We examine the impact on real estate values near state borders because according to theory, they should reflect the perceived economic burden of shortfalls.

First, property values provide a parsimonious and direct measure of the perceived present

[17] The ZHVI provides separate time series for the bottom market tier (33rd percentile and below of home values) and for the top market tier (67th percentile and above of home values), representing typical home values in these tiers. We impose an additional floor of $30,000 on the bottom tier and an additional ceiling of $2,000,000 on the top tier to avoid data quality issues. Given that Zillow obtains prices from a variety of third-party sources and anecdotal evidence suggests that these prices are occasionally incorrect, this filter improves the quality of our data.

value of all future costs and benefits for homeowners. Unlike many assets, long-run discount rates in housing tend to be quite low, increasing the plausibility that temporally distant costs could significantly impact current prices (Giglio, Maggiori, and Stroebel, 2014). Also, for many households their home is the largest financial investment, and prices are likely to reflect perceptions when the stakes are highest.

Second, real estate is effectively immobile. As detailed in Section 1, property bears the full brunt of inefficiencies in public capital raising in settings where other capital, consumers, and labor can easily move, such as near state borders. While prior studies have looked at the correlation between pension underfunding and house prices (e.g., Leeds, 1985; Hur, 2008; Albrecht, 2012; MacKay, 2014; Stadelmann and Eichenberger, 2014; Brinkman, Coen-Pirani, and Sieg, 2018), none focus on border regions. We argue that this is critical for properly measuring the economic burden of pension shortfalls. In addition, these earlier studies suffer from endogeneity in the determinants of shortfalls, which preclude a causal interpretation.

Therefore, we investigate how variation in net pension liabilities per capita, all else equal, translates into variation in property values in regions near state borders. Consider the following border discontinuity design (BDD) regression:

$$PropertyValue_{it} = \ PensionShortfallPerProperty_{st} + \ _{bt} + \omega D_i + \ _l + \ \ddot{} X_{lt} + \ _{it}; \qquad (7)$$

where $PropertyValue_{it}$ is the transaction price of house $i$ and $PensionShortfallPerProperty_{st}$ is the estimated pension shortfall per property in state $s$, in thousands of dollars, in year $t$. $\ _{bt}$ are border county pairs interacted with time fixed effects that allow us to compare properties transacting in physically adjacent regions, just across the state border from each other, in the same time period. This approximates the empirical design suggested by our theoretical framework for an open economy. $D_i$ is the distance to the state border from the property's centroid. If the pension burden is reflected in property values, we would expect prices to jump suddenly at the state border, when shortfalls also jump, even after the inclusion of this distance control. $\alpha_l$ are location fixed effects that capture time-invariant differences by region and region interacted with property characteristics. These include zip code, zip code interacted with property characteristics, and property fixed effects. Therefore, we obtain identification not only from cross-sectional differences across state borders, but from variation in state pension funding status and house prices over time in a border county relative to an adjacent county across the border. Finally, $X_{lt}$ is a vector of time-varying continuous economic controls at the state-year or county-year level.

Figure A.2 illustrates the counties involved in the discontinuity design along with the average shortfall throughout the sample. Our analysis requires sufficient population density to have contemporaneous transactions on either side of the border among comparable property types.

Our theoretical framework suggests that the BDD on shortfalls is an improvement over existing work because of its focus on border regions. However, we still face endogeneity concerns similar to those present in the prior literature. Suppose a state chose to increase local spending on public services instead of funding its pension plans. These sorts of expenditures have been shown to raise property values (e.g., Cellini, Ferreira, and Rothstein, 2010; Bayer, Blair, and Whaley, 2020) and would mechanically increase net pension liabilities per capita. In this case, the estimated pass-through between shortfalls and house prices would understate the economic burden borne by households and could even have the wrong sign. Conversely, if shortfalls are the result of poorly performing expenditures that have negative economic consequences for the state, then the estimated burden may be biased upward.

An ideal empirical setting supplies exogenous, as good as random, shocks to pension shortfalls that allow us to compare real estate transactions before and after the shocks. We therefore focus our analysis on pension asset returns, which cause immediate changes in unfunded pension liabilities driven by factors that are plausibly exogenous to state expenditures. We implement the same empirical design as Equation (7), substituting pension shortfalls with asset performance windfalls:

$$\text{PropertyValue}_{it} = \text{PensionWindfallPerProperty}_{st} + {}_{bt} + !D_i + {}_{l} + X_{lt} + {}_{it}; \qquad (8)$$

where $\text{PensionWindfallPerProperty}_{st}$ is the compounded cumulative return for the pension plans of state $s$ from the beginning of the sample (2002) to the transaction date or interim period of interest (as explained in Section 4.1) multiplied by the assets per property in that state at the beginning of the sample. This can be interpreted as the additional pension assets available per property caused by performance of that state's investment portfolio over that period. The regression coefficient represents the economic burden, which is equal to one plus the deadweight loss, , from our theoretical model in Equation (2). The economic interpretation is consistent with the pass-through in our theoretical motivation because one dollar lower windfall per property implies one dollar of additional pension shortfall per property.

We also consider two-stage least squares (2SLS) designs that recover the economic burden of pension underfunding while alleviating some remaining identification concerns. While our focus on asset returns in border counties reduces many concerns about endogeneity, it is still possible that pension funds' home or familiarity biases (Hochberg and Rauh, 2013) could induce mechanical relations between pension returns and local economic conditions. First, pension managers may buy shares in local firms so that when the local economy does well, both the pension assets and home prices appreciate (home bias). Second, pension managers may overallocate to industries or asset classes that are relatively abundant in a state, inducing a positive correlation between those industries, local economic conditions, and pension returns (familiarity bias). Conversely, pension funds may be used to hedge a state's fundamental risks, resulting in a negative correlation between state economic activity and returns. For example, Texas-based managers with home bias (hedging concerns) might overweight (underweight) both Texan firms and energy-related assets generally.

To alleviate these concerns, we estimate the following 2SLS regression:

$$
\begin{aligned}
PropertyValue_{it} &= \beta\, \widehat{WindfallPerProperty}_{st} + \delta_{bt} + \omega D_i + \gamma_l + \zeta X_{lt} + \epsilon_{it}; \\
WindfallPerProperty_{st} &= \pi\, ExWindfallPerProperty_{st} + \delta_{bt} + \omega D_i + \gamma_l + \zeta X_{lt} + \epsilon_{it}; \qquad (9)
\end{aligned}
$$

where $ExWindfallPerProperty_{st}$ is an instrumental variable that exploits plausibly exogenous variation in pension asset performance. First, we instrument for pension returns using returns in excess of listed benchmarks, which mitigates the familiarity bias concern about the asset category composition of the pension portfolio. However, this first approach leaves open the possibility of home bias where outperformance of local firms drives excess pension returns and provides spoils for the entire state. To alleviate concerns of home bias, we instrument for pension returns using the returns of benchmark assets. To address both concerns simultaneously, we multiply allocations to asset classes that have less potential to be localized (i.e., bonds and equities, and funds that invest in them, rather than commodities, private debt, and real estate) by the relevant benchmark returns from all pensions in the country.[18] In this setup, returns should be unrelated to both local economic conditions and state governance.

[18] Appendix Table A.1 details the asset classes reported in the PPD and delineates which are included in the restricted benchmark return calculations.

# 4 Results

## 4.1 Pension windfalls and property values near state borders

In this section we exploit variation in pension funding coming from windfalls caused by the realized performance of invested pension assets. Our analysis follows the baseline regression in Equation (8), including border county group by year fixed effects that effectively compare the property value at sale of houses in adjacent counties transacting in the same year but in states with different pension windfalls. The group by time fixed effects absorb local trends in economic activity. We control for income per capita at the state level to further alleviate concerns that differential trends in economic activity across the state border affect our estimates. Lastly, we include a continuous measure of distance to the border and a set of fixed effects that controls flexibly for property characteristics.

Within this framework, we begin by using cross-sectional variation in pension asset performance over most of the sample period. In particular, we compare property transaction prices from 2015 to 2018 occurring near state borders where one state had higher pension asset returns from 2002 to 2014 than the other. We focus on this specification for two reasons. First, unless homebuyers are perfectly rational and pay close attention to the evolution of pension funding ratios, short-term variation in asset values is unlikely to impact home prices.[19] Second, to the extent that observable degradation or improvement in public amenities operates as a signal about the financial position of the state government, these effects would likely accumulate over long periods of time.

The pass-through of pension performance to house prices requires that home buyers have some awareness of pension funding. However, it is only necessary that a subset of residents be

[19] Bayer, Blair, and Whaley (2020) find it can take several years for property values to reflect the underprovision of educational public goods.

aware of pension funding for it to have an impact on the housing market equilibrium. In support of this prerequisite, Figure 1 presents Google Trends data showing that internet search volume related to public pensions is higher in states with larger pension shortfalls. In particular, there is a correlation of 0.65 between state level pension shortfalls per household and Google search activity for pension crisis and public pension. States like Illinois, Kentucky, and New Jersey have some of the worst-funded pensions and the most local interest in this issue. This suggests that homeowners are likely aware of the financial problems plaguing their state governments, especially in states with the largest shortfalls.

Table 3 presents formal evidence of how such concerns are reflected in property values. We estimate a BDD that compares house values in adjacent regions just across state borders with varying levels of pension funding caused by pension asset performance from 2002 to 2014. We construct the independent variable of interest as the product of the cumulative pension portfolio return from 2002 to 2014 and the 2001 pension assets per property, which represents the dollar windfall per property. Column (1) reports a positive and statistically significant coefficient of 2.42, which suggests a rise of about two dollars in property values for each dollar of additional pension funding caused by state pension investment outperformance.

Our theoretical framework shows that in a perfectly rational setting with otherwise equivalent circumstances across borders, the coefficient on pension asset returns can be mapped directly to the cost of raising revenues in our model, denoted by For instance, a coefficient of 2.42 suggests that the marginal cost of raising one dollar to pay back pension obligations is $2.42 of total economic burden, including a deadweight loss of $1.42. As discussed in Section 1.2, a pass-through larger than one is not surprising. In a related context, the effect of investment in public education on house prices is also estimated to be large, implying a pass-through in our framework of between one and two (Cellini, Ferreira, and Rothstein, 2010; Bayer, Blair, and Whaley, 2020). In this light, our pass-through estimates are consistent with an underprovision of public goods or services, perhaps driven by severe underfunding of pensions, which is relaxed by higher returns on pension investments.

### 4.2 Addressing identi cation concerns

This section examines potential biases in the estimate presented above. As noted previously, the relative performance of pension assets still has the potential to be endogenously related to state-level outcomes due to familiarity or home bias. We work to alleviate these concerns by restricting variation in pension returns using an instrumental variables framework. An alternative concern with the above approach is that it relies on a single measure of pension windfalls for each state, which could be correlated with unobservable time-invariant state characteristics. We address this concern by constructing a time-varying measure of pension returns and employing property xed e ects. Finally, we present nonparametric estimates of the border discontinuity to ensure the results are not an artifact of our linear regression speci cation.

In the case of familiarity bias, invested asset composition could be driven by familiarity with the sectors prevalent in a region (e.g., timber in Minnesota), inducing a correlation between pension returns and local economic outcomes. Column (2) of Table 3 includes the same sample and control variables as column (1) but incorporates an instrumental variable for the pension windfall in the 2SLS speci cation of Equation (9) using the initial assets per property in 2001 multiplied by the cumulative pension fund performance from 2002 2014 in excess of the mean benchmark performance for each asset class. This restricts variation to relative outperformance within each asset class, rather than variation in allocation across asset classes or sectors. If familiarity bias were driving our results, then using excess returns should eliminate any composition e ect on portfolio returns as long as the benchmarks are well speci ed. Column (2) reports a similar estimate for the economic burden (2.53) that is statistically significant with a strong first stage. This suggests that familiarity bias is unlikely to drive our findings.

However, this still leaves the possibility that home bias could be affecting our estimates. In this case, even within an asset class a pension fund might be more likely to invest in local firms (e.g., Minnesota equities in the Minnesota pension fund). To address this possibility, column (3) takes the pension portfolio composition and applies the benchmark returns of each asset class to calculate implied portfolio returns and reports a similar estimate of the economic burden (2.39). To simultaneously shut down the familiarity channel, in column (4) we collapse the benchmarks into major categories and omit niche asset classes to form our Restricted Benchmarks. Specifically, we restrict attention to assets that have less potential to be localized (i.e., bonds and equities, and funds that invest in them, rather than commodities, private debt, and real estate). Again, we find a similar estimate of the economic burden (2.39), suggesting little evidence of home bias in our primary specification.

One remaining concern with the evidence presented thus far is that it relies on purely cross-sectional variation, so any time-invariant differences across state borders that correlate with pension asset performance could confound identification. To help alleviate this concern, we adjust the independent variable of interest to be the cumulative return between 2002 and the transaction date of the property. This specification allows us to control for unobservable time-invariant confounds, but has a downside relative to our baseline model. Since the sample includes transactions with a shorter window over which pension returns are measured, the regression estimates could be attenuated if it takes time for pension performance to be reflected in property values. This is especially true when we require a house to have repeat sales, which mechanically tilts the sample towards earlier observations.

The first column of Table 4 replicates the regression in column (1) of Table 3 using the rolling measure of cumulative pension returns. This specification yields a positive and significant coefficient of 2.16, quantitatively similar to our baseline estimate. The point estimate is slightly lower in this setup, perhaps reflecting the attenuation bias discussed above.

After establishing similar findings with the rolling measure of cumulative returns, we explore whether time-invariant confounds are biasing our estimates. One possibility is that property values are correlated with 2001 pension assets in a manner unrelated to pension shortfalls (e.g., generous pensions are associated with better or worse public amenities). To address this, we instrument for windfalls using only the public benchmark returns (not multiplied by initial assets per property) from our most restrictive specification in Table 3 (i.e., the first stage is a regression of dollars on returns). Column (2) of Table 4 reports a coefficient estimate based on this approach that is slightly lower (1.83) but similar to column (1).

Next, we restrict attention to properties with repeat sales and add property fixed effects to rule out the possibility that other unobservable time-invariant local factors affect our results. In column (3), we focus on the sub-sample of properties with repeat transactions during our sample period, requiring at least four years between transactions. Unsurprisingly, since this sample allows even less time for property values to reflect pension performance, the coefficient estimates are lower than the full-sample estimates. More importantly, we obtain nearly identical estimates after adding property fixed effects in column (4), which suggests that time-invariant omitted variables at the state, local, and property level do not bias our estimates of the economic burden.

Finally, we show that our findings are driven by neither the construction of windfalls per property nor the functional form of the BDD. In Appendix Tables A.2 and A.3, we present coefficients with the same sign and statistical significance using a simpler specification that focuses on cumulative pension returns without scaling by 2001 pension assets.

We apply this simple form of variation to confirm our main result in a nonparametric border discontinuity design. For each border pair, we determine the state that has the larger pension asset return between 2002 and the sale date of the property and label this a treated state, with $Treated_{st}$ taking a value of 1 for treated states and -1 for non-treated states, restricting attention to properties within 20 miles of the border. We estimate the following regression to obtain a vector of coefficients that reflect the total sales price increase for a house that trades in each one-mile bucket on either side of the border:

$$HousePrice_{it} = \sum_{k=1}^{20} \beta_k Treated_{st} \cdot 1(Miles_i = k) + \alpha_{bt} + \gamma_l + \delta X_{lt} + \epsilon_{it}. \qquad (10)$$

Figure 2 plots the coefficients for five miles on either side of the border. Circular dots represent the $\beta$ coefficient estimates, diamonds are the differences between the treated and untreated coefficients, and lines are the 95% confidence intervals for the differences.

Two distinct patterns are visible. First, for properties very close to the border, we observe a fairly stable premium in states with higher pension returns. Second, as we move across the border there is a sudden jump in the value of the properties in states with higher pension outperformance. This is consistent with our predictions and suggests that our findings are not driven by the functional form assumptions of the BDD.

## 4.3 External validity

Since our analysis restricts attention to a subset of the housing market near state borders, it is worthwhile to assess whether our estimates are likely to apply more generally. As explained above, we focus on state borders because theory suggests that the burden of addressing pension shortfalls should accrue to real estate when labor and physical capital can be relocated to another state at low cost. In contrast to prior work on pensions and house prices, our primitive of interest is the economic burden of pension shortfalls, not a more general average effect on house prices that can be observed across all counties. As we move further away from state borders, the cost of moving other types of capital increases, which disperses the pension burden among other forms of capital and precludes us from making clear predictions about the effect on house prices.

Along these lines, Appendix Table A.4 reports a much smaller, but statistically significant, coefficient when applying our main specification to interior counties.[20] Since we cannot recover the coefficient of interest directly in interior counties, we evaluate whether there is something different about border counties by comparing the observable characteristics of interior and border counties. Our estimates reflect the deadweight loss associated with raising funds or cutting amenities to address pension shortfalls, so we focus our comparison on differences in local government finances and costs of fundraising across these regions. Appendix Table A.5 shows that border counties are similar to interior counties on these dimensions. This analysis uses local government financial data aggregated to the county level by the U.S. Census Bureau for fiscal years 2007 and 2012. We make statistical comparisons for 15 different financial ratios in these two years and find that only four out of 30 differences are statistically significant at the 10% level, none of which hold across both observation years for a given ratio. This suggests that border counties are fairly representative in terms of their financial position.

Nevertheless, to examine whether the observed differences in county characteristics are correlated with the estimated economic burden, Appendix Table A.6 reproduces our main specification using weighted least squares regressions in which the weights are chosen such that border coun-

[20] We use a linear specification in column (1) of Appendix Table A.4 to reveal a statistically significant decline in the coefficient of interest based on distance to the border, suggesting a diffusion of the burden across other forms of capital that precludes identification in interior regions. We also show a larger economic burden when separately estimating effects in border (column 2) relative to interior counties (column 3) with the same specification. For counties internal to a state, we impute the county border group to which it belongs by finding the county border group of the county whose centroid is closest to its own centroid.

ties match interior counties on each characteristic.[21] The results of this approach are identical to those of column (1) in Table 3. In sum, the evidence in Tables A.5 and A.6 suggests that our estimates of the economic burden are likely to apply more generally.

Although modeling the general equilibrium implications of our findings is beyond the scope of this paper, a simple linear aggregation highlights the overall magnitude of the economic burden imposed by pension underfunding. As noted in the introduction, Rauh (2016) estimates that the unfunded portion of U.S. state and local pension promises exceeds $3.8 trillion. Our estimated economic burden of approximately two implies a deadweight loss of approximately one dollar per dollar of shortfall. Since there are about 121 million households in the United States, the 95% confidence interval around the estimate from column (1) of Table 3 corresponds to an average deadweight loss of between $25,611 and $63,422 per household, or between 37% and 92% of median household income.[22]

## 4.4 Crowding out contributions: The shortfall of shortfalls

Although we have motivated the use of a border discontinuity design, we have not fully explained why we use windfalls from variation in pension returns rather than the level of pension shortfalls as the explanatory variable of interest. As a starting point, it is important to note that there is an inverse relation between windfalls and shortfalls that must hold instantaneously. By definition, an additional dollar of assets reduces the net pension shortfall by one dollar. However, at longer horizons the change in the pension funding ratio in response to an exogenous one dollar windfall depends on whether the state reduces pension contributions in response. This crowding

[21] In particular, we follow prior work (e.g., Jacob, Michaely, and Müller, 2018) in using the entropy-balancing method developed by Hainmueller (2012) to obtain weights that would set the weighted average of the border counties to be the same as those in the interior for multiple variables.

[22] Based on 2019 median household income, available from https://www.census.gov/content/dam/Census/library/publications/2020/demo/p60-270.pdf

out between windfalls and contributions would lead observed shortfalls to fall by less than one dollar after a one dollar windfall in equilibrium, since the state responds by contributing less to the pension fund than it otherwise would have.

For direct evidence that the observed pension shortfall is an equilibrium outcome, Appendix Table A.7 shows that pension shortfalls are positively correlated with contributions to the pension system by both the state and its employees. If pension fund outperformance leads to a reduction in contributions and a shift in government spending to value-improving projects, then even a 2SLS regression that instruments for shortfalls would understate the e ects of pension funding. On the other hand, if such expenditures are value-destroying, the same regression would be biased upwards. Ultimately, this is an empirical question that demands variation in pension funding that is una ected by the substitution between pension contributions and local government expenditures and the relative value of those expenditures.

While it does not recover the economic primitive of interest, we can learn something interesting about crowding out and the bene ts of our empirical design by considering windfalls as an instrumental variable for the observed level of pension shortfalls in the following 2SLS regression:

$$
\begin{aligned}
\text{PropertyValue}_{it} &= {}^{I}\,\text{ShortfallPerProperty}_{st} + {}_{bt} + !D_{i} + {}_{l} + \ddot{X}_{lt} + {}_{it}; \\
\text{ShortfallPerProperty}_{st} &= \text{WindfallPerProperty}_{st} + {}_{bt} + D_{i} + {}_{l} + \ddot{X}_{lt} + {}_{it}:
\end{aligned} \qquad (11)
$$

Relating this system of equations to the system in Equation (7), the economic interpretation of the rst-stage regression here is that 1* = 1 * _ represents the crowding out per dollar of windfall. If there is no crowding out, then = 1 and = , and the second-stage estimates are equal whether we use the windfalls or shortfalls as the explanatory variable of interest.

Table 5 presents estimates of Equation (11). Column (2) reports the rst-stage regression, in which the endogenous variable is the observed net shortfall per property and the instrumental variable is windfall per property coming from pension asset returns. The coe cient of 0.45 indicates that each dollar of windfall causes the equilibrium shortfall to fall by about 45 cents. Since the shortfall must fall instantaneously by one dollar, this means that pension contributions are reduced by 55 cents for each dollar of windfall. To our knowledge, this is the rst estimate of the crowding-out e ect of pension asset performance on fund contributions.

While this result is interesting on its own, the comparison between columns (1) and (3) is more important for understanding our empirical strategy. For ease of comparison, column (1) reproduces the same estimate of Equation (8) reported in Table 4, which is based on pension windfalls due to asset returns. Column (3) presents the second-stage coe cient from Equation (11), based on the level of pension shortfalls. The respective coe cient estimates of 2.16 and 4.82 would correspond to vastly di erent implications for the cost of addressing pension underfunding, but the latter estimate is contaminated by the crowding out e ect documented above. Mechanically, the ratio of these estimates is equal to the rst-stage estimate from column (2), which means the bias from using the level of shortfalls in this analysis is increasing in the degree of crowding out. This shows that even if we instrument for the level of shortfalls using plausibly exogenous variation due to windfalls, we obtain an upward-biased estimate of the economic burden because states contribute less to their pension funds when the funds' investments are performing well.

## 4.5 Drivers of the economic burden

### 4.5.1 Current versus future bene ts

In the preceding analysis, we estimate a pass-through of approximately two between the public pension shortfall per property and the value of houses near state borders. Since we measure pension asset returns over long horizons for the properties transacting at the end of the sample, a 16-year period the channel through which pension shortfalls affect house prices is ambiguous. After a long period of disappointing returns, states with large shortfalls may raise taxes or shift the allocation of tax revenues away from public projects to make pension contributions. If this substitution were at play, then the estimated pass-through could reflect a combination of worse current amenities and higher future liabilities.

Columns (1) through (4) of Table 6 attempt to distinguish between the present and future channels by controlling for local rental prices. Renters have the same ability as homeowners to enjoy public amenities, but they are unaffected by the capitalization of future costs and benefits accruing after the term of the rental. If the estimated effect of pension shortfalls on house prices is driven by worse current amenities, then pension shortfalls should also affect rental prices. If instead the effect is driven by expectations about the future costs of addressing shortfalls, then current rental prices should not show the same effect.

Column (1) of Table 6 reproduces our main result and column (2) adds a control for rental prices for similar properties in the same county in the same year.[23] The estimated economic burden is 23% lower after we control for rental prices. According to the logic outlined above, this implies that 23% of the estimate is due to worse current amenities and the other 77% is due to future costs associated with addressing pension shortfalls. The role of amenities is consistent with the evidence on public service provision in Appendix Table A.7.

In contrast, we find virtually no difference in the estimated burden between columns (3) and (4), which add property fixed effects to control for unobservable time-invariant factors. This analysis restricts attention to the sample of repeat sales, which as we note above, allows for less

[23] It is worth pointing out that monthly rental rates are a "bad control" in the estimation of the economic burden. As we argue above, we expect some amount of the burden to pass through into current rental rates, especially over long intervals, which means the coefficient on pension windfalls would change. While the inclusion of rental rates is helpful in distinguishing the effect of windfalls on current and future (dis)amenities, it is unlikely to help us recover a more accurate estimate of the economic burden.

time over which shortfalls can accumulate to impact housing prices. Together with the first two columns, these findings are consistent with pension shortfalls driving changes in current amenities, but these changes occurring relatively slowly. In sum, our findings suggest that the bulk of the perceived burden of pension shortfalls will be realized in the future, but current amenities are deteriorating in the later part of our sample period.

#### 4.5.2 Economic burden and tax distortions

As discussed above, our estimates reflect the perceived deadweight loss associated with improving pension funding through future increases in taxes or reductions in government service provision. The finding of a pass-through larger than one suggests that substantial expected costs would be associated with raising taxes to improve pension funding, or that it might be politically infeasible to do so. According to this interpretation, we should expect to see larger effects of pension shortfalls in states where the distortionary effects of raising additional funds are larger. Income taxes are typically considered to be distortionary, whereas property taxes are less so, especially when levied on the value of land (Mieszkowski and Zodrow, 1989).

Along these lines, column (5) of Table 6 presents evidence of heterogeneity in the estimated pass-through corresponding to the top marginal state income tax rate. In particular, we find that the economic burden is 58 cents larger per dollar of shortfall in a state with a one standard deviation higher income tax rate, consistent with larger distortions of raising tax revenue. This could represent the marginal cost associated with raising funds through income tax hikes, but more likely reflects the high marginal value of a dollar of additional funds in the subset of states that choose to have high marginal income tax rates despite their distortionary effects.

# 5 Conclusion

This paper uses plausibly exogenous variation in state pension funding based on excess asset performance to show that a one dollar increase in the public pension shortfall per property causes a two dollar reduction in property values near state borders. We motivate this research design with a parsimonious theoretical framework showing that due to its relative immobility, real estate on state borders should reflect the economic burden of pension shortfalls. The consistency of our estimates across a variety of instrumental variables specifications supports a causal interpretation. Although we show evidence of reduced public investment in states with larger pension shortfalls, our analysis of rental rates suggests that the house price effects are mostly driven by future costs rather than reductions in current amenities.

Our findings are consistent with models of inefficient taxation or the underprovision of public goods and services. The estimated economic burden of pension shortfalls is comparable to previous estimates of the effect of public spending on infrastructure and public teacher salaries (e.g., Cellini, Ferreira, and Rothstein, 2010; Bayer, Blair, and Whaley, 2020). This highlights the perceived value of projects that could be funded in the absence of large unfunded pension obligations that demand funding in the future.

Finally, our results have implications for the political economy of public sector labor markets and local government finances. In light of prior evidence on public workers' low willingness to pay for retirement benefits (Fitzpatrick, 2015), our findings raise questions about the efficiency of public sector compensation schemes. Why do state governments offer generous retirement benefits to employees who do not value them while imposing the funding burden on households who view it as costly? We quantify the latter dimension of this problem by estimating a deadweight loss associated with addressing pension shortfalls of between $25,611 and $63,422 per household. However, our analysis takes the degree of pension underfunding as given. Though it is beyond the scope of this paper to explain the origins of the public pension crisis, we look forward to future research on this topic.

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Table 1

## Public Plans Data Summary Statistics

Data are from the Public Plans Data (PPD) database provided by the Center for Retirement Research (CRR) at Boston College and are reported at the state-year level. Asset return is average annual portfolio return. Actuarial assets and Actuarial liabilities are ActAssets_GASB and ActLiabilities_GASB in the dataset in millions of dollars. Actuarial funded ratio is given by ActFundedRatio_GASB, which is ActAssets_GASB divided by ActLiabilities_GASB in the dataset. Allocation of pension portfolios to equities, xed-income (FI), real estate (RE), private equity (PE), hedge fund (HF), commodities (Comd), cash, miscellaneous alternative assets (AltMisc), and other assets are shown in percentage terms.

| Variable | Obs. | Mean | Std. Dev | Q5 | Q25 | Q50 | Q75 | Q95 |
|---|---|---|---|---|---|---|---|---|
| Asset return | 624 | 0.056 | 0.076 | -0.086 | 0.010 | 0.060 | 0.115 | 0.163 |
| Actuarial assets (>m) | 624 | 13,947 | 13,469 | 1,445 | 5,135 | 8,612 | 18,082 | 44,358 |
| Actuarial liabilities (>m) | 624 | 17,577 | 15,908 | 2,217 | 6,522 | 11,880 | 24,174 | 52,007 |
| Actuarial funded ratio | 624 | 0.787 | 0.144 | 0.555 | 0.697 | 0.776 | 0.890 | 1.010 |
| Equity share | 624 | 0.529 | 0.094 | 0.360 | 0.472 | 0.538 | 0.598 | 0.663 |
| FI share | 624 | 0.279 | 0.077 | 0.180 | 0.226 | 0.268 | 0.317 | 0.412 |
| RE share | 624 | 0.055 | 0.038 | 0.000 | 0.022 | 0.059 | 0.081 | 0.110 |
| PE share | 624 | 0.053 | 0.051 | 0.000 | 0.007 | 0.041 | 0.082 | 0.146 |
| HF share | 624 | 0.040 | 0.056 | 0.000 | 0.000 | 0.013 | 0.064 | 0.154 |
| Comd share | 624 | 0.014 | 0.022 | 0.000 | 0.000 | 0.000 | 0.021 | 0.064 |
| Cash share | 624 | 0.017 | 0.019 | 0.000 | 0.004 | 0.012 | 0.023 | 0.053 |
| AltMisc share | 624 | 0.009 | 0.026 | 0.000 | 0.000 | 0.000 | 0.000 | 0.076 |
| Other share | 624 | 0.004 | 0.022 | 0.000 | 0.000 | 0.000 | 0.000 | 0.017 |

Table 2

# Housing Transactions Summary Statistics

This table presents summary statistics for our sample of properties that merges ZTRAX (Zillow's Transaction and Assessment Dataset) with state-level annual pension performance/shortfalls, local annual rental rates, and state-level annual income per capita. Rental rates are based on fair market rates for single family residences with the same number of bedrooms as the transaction property (or 3 bedrooms if the number of bedrooms is missing for the transacting property) at the county-year level from the Department of Housing and Urban Development. The sample is restricted to property transactions involving single-family residences in counties sharing a border with an adjacent state that are within 50 miles of that border and have a transaction price between the typical home values in the the bottom and top market tiers as calculated at the county-month level by the ZHVI.

| Variable | Obs. | Mean | Std. Dev | Q5 | Q25 | Q50 | Q75 | Q95 |
|---|---|---|---|---|---|---|---|---|
| Sales Price ($'000s) | 712,505 | 260 | 135 | 105 | 162 | 228 | 325 | 525 |
| Transaction Month | 712,505 | 07_2010 | 54 Mos | 11_2003 | 06_2006 | 06_2010 | 05_2014 | 08_2017 |
| Border Dist (mi) | 712,505 | 16.9 | 10.7 | 2 | 8 | 16 | 26 | 35 |
| | | | | | | | | |
| Building Age (yrs) | 629,379 | 7.7 | 5.5 | 2 | 3 | 6 | 12 | 19 |
| Sq Ft | 603,693 | 1,813 | 793 | 800 | 1,290 | 1,655 | 2,110 | 3,800 |
| Lot Sq Ft | 657,004 | 19,020 | 77,226 | 2,000 | 4,500 | 7,500 | 14,500 | 62,000 |
| # Bedrooms | 492,118 | 3.28 | 0.71 | 2 | 3 | 3 | 4 | 4 |
| # Bathrooms | 554,187 | 4.31 | 1.30 | 2 | 4 | 4 | 5 | 6 |
| | | | | | | | | |
| Shortfall/Prop ($'000s) | 712,505 | 21.28 | 16.10 | 0.54 | 9.19 | 17.68 | 30.71 | 46.84 |
| 02-14 Cum. Port. Ret. (%) | 129,940 | 145 | 23 | 90 | 137 | 141 | 161 | 167 |
| 02-14 Cum. Excess Ret. (%) | 129,940 | 12 | 16 | -2 | 1 | 4 | 28 | 40 |
| '02-14 Cum. Port. Ret. · '01 Assets/Prop ($'000s) | 129,940 | 27.02 | 22.16 | 11.48 | 16.19 | 16.19 | 25.64 | 78.23 |
| | | | | | | | | |
| Rental Price ($) | 712,505 | 1,328 | 345 | 800 | 1,073 | 1,318 | 1,567 | 1,898 |
| State-Year Income PC ($) | 712,505 | 44,607 | 9,195 | 32,996 | 38,103 | 41,862 | 50,035 | 63,598 |

Table 3

# Pension Windfalls and House Prices in Border Counties

This table presents estimates from a state border discontinuity design model where the dependent variable is the sales price, in thousands of dollars, of a residential property. The explanatory variable of interest is based on invested assets' cumulative performance from 2002-2014 in the pension plans associated with the state in which the focal property is located, multiplied by initial assets per property in 2001 to create a measure of additional pension shortfall per property due to asset performance (Windfall). The sample is restricted to property transactions involving single-family residences in counties sharing a border with an adjacent state that are within 50 miles of that border and have a transaction price between the typical home values in the the bottom and top market tiers as calculated at the county-month level by the ZHVI. Fixed e ects for the county border group of the property interacted with the calendar year of the transaction, as well as covariates for the distance to the state border and income per capita at the state-year level, are included throughout. These speci cations also control for property type by including six interacted property characteristic xed e ect cells (square footage of structure, square footage of lot, age of building, number of bedrooms, number of bathrooms, and number of stories). Column (1) is the baseline regression described above where the dependent variable is the sales price (in thousands of dollars), for transactions in the years 2015-2018 and the primary variable of interest is the Windfall per Property over the 2002-2014 period. Column (2) instruments for Windfall using the initial assets per property in 2001 multiplied by the cumulative pension fund performance from 2002-2014 in excess of the benchmark performance for each asset class the fund is invested in. Column (3) instruments for Windfall using the initial assets per property in 2001 multiplied by the cumulative pension fund performance from 2002-2014 that would have occurred based on the fund's asset allocations, had it earned the benchmark performance for each asset class. Column (4) is the same as column (3), but restricts attention to assets that have less potential to be localized (i.e., bonds and equities, and funds that invest in them, rather than commodities, private debt, and real estate). Where applicable, we report the Kleibergen-Paap F-test for weak identi cation. Reported t-statistics in parentheses are heteroskedasticity-robust and clustered at both the zip and transaction month level. ***, **, and * indicate statistical signi cance at the 1%, 5%, and 10% levels, respectively.

| | Sales Price $('000s) | | | |
|---|---|---|---|---|
| | OLS | 2SLS | 2SLS | 2SLS |
| | (1) | (2) | (3) | (4) |
| 2002-2014 Windfall Per Property $('000s) | 2.418***<br>(8.10) | 2.529***<br>(9.23) | 2.391***<br>(7.81) | 2.389***<br>(7.79) |
| Border Distance | X | X | X | X |
| State-Year Income PC | X | X | X | X |
| Border Group-Tran Year FE | X | X | X | X |
| 6 Prop Chars FE | X | X | X | X |
| Instrumental Variable | | Excess Ret. Windfall Per Prop | Bnchm. Ret. Windfall Per Prop | Restr. Bm. Ret. Windfall Per Prop |
| Observations | 129,940 | 129,940 | 129,940 | 129,940 |
| Adj. $R^2$ | 0.813 | | | |
| Weak ID KPF Stat | | 312.5 | 83,047 | 68,209 |

## Table 4

# Rolling Pension Windfall Regressions and Repeat Sales

This table presents estimates from a state border discontinuity design model where the dependent variable is the sales price, in thousands of dollars, of a residential property. The explanatory variable of interest is based on invested assets' cumulative performance in the pension plans associated with the state in which the focal property is located, but only in the years prior to the transaction since 2002, multiplied by the pension assets per property as of 2001 (Windfall). The sample is restricted to property transactions involving single-family residences in counties sharing a border with an adjacent state that are within 50 miles of that border and have a transaction price between the typical home values in the the bottom and top market tiers as calculated at the county-month level by the ZHVI. Fixed e ects for the county border group of the property interacted with the calendar year of the transaction, as well as a covariate for the income per capita at the state-year level, are included. Columns (1) through (3) also include a covariate for the distance to the state border and six interacted property characteristic xed e ect cells that control for property type (square footage of structure, square footage of lot, age of building, number of bedrooms, number of bathrooms, and number of stories). Column (1) is the baseline regression described above. Column (2) instruments for Windfall using the cumulative pension fund performance from 2002 to the sale of the property that would have occured based on the fund's asset allocations, had it earned the benchmark performance for each asset class, but restricting attention to securities that have lessened potential to be localized (i.e., bonds and equities rather than commodities, private debt, real estate) and funds investing in them. Column (3) is the same as column (1) but restricts to properties with repeat sales in the sample. Column (4) is the same as column (3) but replaces the interacted property characteristic xed e ects and the distance to state-border covariate with a property-level xed e ect. In this case, identi cation is based on within-property variation over time coming from repeat sales. Where applicable, we report the Kleibergen-Paap F-test for weak identi cation. Reported t-statistics in parentheses are heteroskedasticity-robust and clustered at both the zip and transaction month level. ***, **, and * indicate statistical signi cance at the 1%, 5%, and 10% levels, respectively.

| | Sales Price $('000s) | | | |
|---|---|---|---|---|
| | OLS | 2SLS | OLS | OLS |
| | (1) | (2) | (3) | (4) |
| 2002-Sale Windfall Per Property $('000s) | 2.162*** | 1.831*** | 1.481*** | 1.425*** |
| | (8.01) | (4.26) | (5.95) | (5.15) |
| Border Distance | X | X | X | |
| State-Year Income PC | X | X | X | X |
| Border Group-Tran Year FE | X | X | X | X |
| 6 Prop Chars FE | X | X | X | |
| Repeat Sales Sample | | | X | X |
| Property FE | | | | X |
| Instrumental Variable | | Restr. Bm. Return | | |
| Observations | 712,505 | 712,505 | 54,882 | 54,882 |
| Adj. $R^2$ | 0.857 | | 0.853 | 0.913 |
| Weak ID KPF Stat | | 43.03 | | |

Table 5

## The Shortfall of Shortfalls

This table presents estimates from a state border discontinuity design model where the dependent variable is the sales price, in thousands of dollars, of a residential property. The explanatory variable of interest is based on invested assets' cumulative performance in the pension plans associated with the state in which the focal property is located, but only in the years prior to the transaction since 2002, multiplied by the pension assets per property as of 2001 (Windfall). The sample is restricted to property transactions involving single-family residences in counties sharing a border with an adjacent state with di erential pension funding that are within 50 miles of that border and have a transaction price between the typical home values in the the bottom and top market tiers as calculated at the county-month level by the ZHVI. Fixed e ects for the county border group of the property interacted with the calendar year of the transaction, as well as covariates for the distance to the state border and the income per capita at the state-year level, are included throughout. These speci cations also control for property type by including six interacted property characteristic  xed e ect cells (square footage of structure, square footage of lot, age of building, number of bedrooms, number of bathrooms, and number of stories). Column (1) is the baseline regression described above and replicates column (1) of Table 4. Column (2) is the  rst stage of the 2SLS regression detailed in Equation (11), where the endogenous variable is the observed net shortfall per property and the instrumental variable is windfall per property coming from pension asset performance. This regression demonstrates the the crowding-out e ect of pension performance on fund contributions. Column (3) is the speci cation in Equation (11) and demonstrates that, because states contribute less to their pensions when they earn high returns, using equilibrium shortfalls leads to a biased estimate of the economic burden, even if shortfalls are instrumented with plausibly exogenous windfalls. Where applicable, we report the Kleibergen-Paap F-test for weak identi cation. Reported t-statistics in parentheses are heteroskedasticity-robust and clustered at both the zip and transaction month level. ***, **, and * indicate statistical signi cance at the 1%, 5%, and 10% levels, respectively.

| | Sales Price $('000s) | Shortfall Per Prop $('000s) | Sales Price $('000s) |
|---|---|---|---|
| | OLS | OLS | 2SLS |
| | (1) | (2) | (3) |
| 2002-Sale Windfall Per Property $('000s) | 2.162*** | -0.449*** | |
| | (8.01) | (-7.98) | |
| Shortfall Per Property $('000s) | | | -4.816*** |
| | | | (-5.73) |
| Border Distance | X | X | X |
| State-Year Income PC | X | X | X |
| Border Group-Tran Year FE | X | X | X |
| 6 Prop Chars FE | X | X | X |
| Instrumental Variable | | | Windfall Per Prop |
| Observations | 712,505 | 712,505 | 712,505 |
| Adj. $R^2$ | 0.857 | 0.913 | |
| Weak ID KPF Stat | | | 63.7 |

Table 6

# Pension Windfalls: Drivers of the Economic Burden

This table presents estimates from a state border discontinuity design model where the dependent variable is the sales price, in thousands of dollars, of a residential property. Columns (1) through (4) compare the baseline estimates to the estimates after controlling for time-varying rental rates. Column (5) examines how the e ect of pension funding on house prices varies with the di culty of raising additional funds, as proxied by the highest marginal state income tax rate. The explanatory variable of interest is based on invested assets' cumulative performance in the pension plans associated with the state in which the focal property is located, but only in the years prior to the transaction since 2002, multiplied by the pension assets per property as of 2001 (Windfall). The sample is restricted to property transactions involving single-family residences in counties sharing a border with an adjacent state that are within 50 miles of that border and have a transaction price between the typical home values in the the bottom and top market tiers as calculated at the county-month level by the ZHVI. Fixed e ects for the county border group of the property interacted with the calendar year of the transaction, as well as a covariate for the income per capita at the state-year level, are included throughout. Columns (1) and (2) also include a covariate for the distance to the state border and six interacted property characteristic xed e ect cells that control for property type (square footage of structure, square footage of lot, age of building, number of bedrooms, number of bathrooms, and number of stories). Columns (3) through (5) include a property-level xed e ect to exploit within-property variation over time. Column (1) replicates column (1) of Table 3. Column (2) is the same as column (1) but adds a control for time-varying rental rates. Column (3) replicates column (4) of Table 4. Column (4) is the same as column (3) but adds a control for time-varying rental rates. In column (5), the windfall is interacted with the highest marginal state income tax rate as of 2002 (taxpolicycenter.org), which is standardized at the regression observation level to normalize the interaction e ect. Reported $t$-statistics in parentheses are heteroskedasticity-robust and clustered at both the zip and transaction month level. ***, **, and * indicate statistical signi cance at the 1%, 5%, and 10% levels, respectively.

| | Sales Price $('000s) | | | | |
|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) |
| 2002-2014 Windfall Per Property $('000s) | 2.418*** | 1.853*** | | | |
| | (8.10) | (7.50) | | | |
| 2002-Sale Windfall Per Property $('000s) | | | 1.425*** | 1.353*** | 0.658* |
| | | | (5.15) | (4.87) | (1.66) |
| Est. Mo Rent ($) | | 0.124*** | | 0.0603*** | |
| | | (6.98) | | (2.86) | |
| 2002-Sale Windfall Per Prop $('000s) • Standardized Highest Marginal Income Tax Rate | | | | | 0.582** |
| | | | | | (2.28) |
| Border Distance | X | X | | | |
| State-Year Income PC | X | X | X | X | X |
| Border Group-Tran Year FE | X | X | X | X | X |
| 6 Prop Chars FE | X | X | | | |
| Property FE | | | X | X | X |
| Observations | 129,940 | 129,940 | 54,882 | 54,882 | 54,882 |
| Adj. $R^2$ | 0.813 | 0.814 | 0.913 | 0.914 | 0.913 |

Figure 1. Popular Interest in Public Pensions from Google Search Trends This figure presents evidence on popular interest in the condition of public pensions using data from Google Trends. For each state in our regression sample, depicted in Figure A.2, we obtain monthly series of search trends for the terms "pension crisis" and "public pension" over the period January 2004 to December 2020. Google Trends are computed on a relative basis, so they must be scaled by a common search term to make comparisons across states. We scale the total interest in the two pension-related terms by each state's trend series for the "municipal bond" topic. To match the timing of our estimated house price effects, we take the average ratio of pension search trends to municipal bond search trends from 2015 to 2018, which we plot on the $y$-axis of the figure. The $x$-axis of the figure is the average pension shortfall per property, in thousands of dollars, over the same period. The scatter plot reveals a positive relation between pension underfunding and popular interest in the issue. The corresponding regression coefficient is 0.0034 ($t = 4.03$) and the $R^2$ is 0.43.

Figure 2. Pension Return Discontinuity in House Prices This gure presents nonparametric estimates of a border discontinuity design for house values near the borders of states with di ering pension asset performance between 2002 and the sale date of the property. We plot the coe cients for the ve miles surrounding each border in our sample, with blue dots representing the primary coe cient of interest in Equation (10). Red diamonds denote the di erence between the coe cient estimates for properties in better performing states minus those for equidistant from the border properties in worse performing states. Red lines denote 95% con dence intervals for these estimates.

# Appendix

## A Details of the Model

### A.1 Tax burden in a one-sector closed economy

Consider a closed economy where labor, $L$, and capital, $K$, are used to produce a single good according to a linear homogeneous of degree one production function $F(K, L)$, with $F_L > 0$ and $F_K > 0$. Suppose that capital is inelastically supplied, while labor supply is positively related to the real wage, $W/P$, where $W$ is the wage rate and $P$ is the price of the economy's single good:

$$L = L(W/P). \tag{A.1}$$

The equilibrium wage rate $W$ and the rental rate on capital $r$ are given by the standard first order conditions:

$$F_K(K, L) = r/P; \qquad F_L(K, L) = W/P \tag{A.2}$$

Using market-clearing in the labor market, we have

$$F_L(K, L(W/P)) = W/P. \tag{A.3}$$

Consider the burden of a tax at rate imposed on the elastically-supplied labor, we get

$$PF_L = W(1 + \ ) \tag{A.4}$$

Equating supply and demand for labor in the tax equilibrium and taking the derivative with respect to $\tau$, we find that the percentage change in $W\_P$ from a change in $\tau$, evaluated at $\tau = 0$, is given by

$$\frac{\partial (W\_P)/(W\_P)}{\partial \tau} = \frac{\varepsilon^D}{\varepsilon^S + \varepsilon^D}; \tag{A.5}$$

where $\varepsilon^S$ is the positive elasticity of labor supply, and $\varepsilon^D$ is the negative elasticity of labor demand. The marginal losses of rents to labor, $\frac{\partial (W\_P)}{\partial \tau} L$, and to capital, $\frac{\partial (r\_P)}{\partial \tau} K$, as a ratio of the marginal tax revenue, $(W\_P)L$, can be written as

$$\frac{\frac{\partial (W\_P)}{\partial \tau} L}{(W\_P)L} = \frac{\varepsilon^D}{\varepsilon^S + \varepsilon^D}; \tag{A.6}$$

and

$$\frac{\frac{\partial (r\_P)}{\partial \tau} K}{(W\_P)L} = \frac{\varepsilon^S}{\varepsilon^D + \varepsilon^S}: \tag{A.7}$$

Note that Equations (A.6) and (A.7) sum to 1: the full burden of the tax falls on either capital or labor.

If the supply of labor is perfectly inelastic ($\varepsilon^S = 0$) or labor demand is perfectly elastic ($\varepsilon^D = \infty$), labor bears the full burden of the tax, i.e., the right hand sides of (A.6) and (A.7) are 1 and 0, respectively. At the other extreme, if labor supply is perfectly elastic ($\varepsilon^S = \infty$) or the demand for labor is perfectly inelastic ($\varepsilon^D = 0$), capital bears the full burden of the tax. Importantly, although the tax is imposed on labor, from Equation (A.7), capital always bears some burden of the tax if $\varepsilon^S \neq 0$ and $\varepsilon^D \neq \infty$.

## A.2 Tax burden in an open economy

Suppose there are two bordering states, $A$ and $B$, in the country with production functions $F_A(K)$ and $F_B(K)$ used to produce a common consumption good. Let $K_A$ be the capital in state $A$

and $K_B = \overline{K} - K_A$ be the capital in state $B$, where $\overline{K}$ is the total countrywide capital. If $r$ is the rental rate on capital, and is the tax on capital in state $A$, we have

$$F_A'(K_A) = r + \ ; \qquad F_B'(K_B) = r. \tag{A.8}$$

Using Equation (A.8) and the constraint $K_A + K_B = \overline{K}$, we can show that the change in rents to countrywide capital, $dr\overline{K}$, expressed as a ratio of the marginal tax revenue, $d\ K_A$, evaluated at $= 0$, is given by

$$\frac{(dr/d\ )\overline{K}}{K_A} = -\frac{{}_A\overline{K}}{{}_BK_B + {}_AK_A} \leq 0, \tag{A.9}$$

where $_A$ and $_B$ are the nonnegative demand elasticities for capital in states $A$ and $B$, respectively. If $A$ and $B$ have identical production functions, $F_A(\cdot) = F_B(\cdot)$, then $_A = \ _B$ and $K_A = K_B$ initially. Then the right hand side of Equation (A.9) equals $-1$ and countrywide capital, $\overline{K}$, bears the full marginal burden of the tax in $A$. If the demand for capital in $B$ is perfectly inelastic, $_B = 0$, or is perfectly elastic in $A$, $_A = \infty$, countrywide capital bears more than 100% of the tax. At the opposite extreme, if capital demand is perfectly elastic in $B$ or in perfectly inelastic demand in $A$, $\overline{K}$ bears none of the burden of the tax.

Land rents in $A$ and $B$, denoted $R_A$ and $R_B$, respectively, are given by

$$R_A = F_A(K_A) - (r + \ )K_A; \qquad R_B = F_B(K_B) - rK_B; \tag{A.10}$$

implying[24]

$$\frac{\square\square_\square/\square\square}{\square_\square} = -\frac{\square_\square\square_\square}{\square_\square\square_\square + \square_\square\square_\square} \le 0; \qquad \frac{\square\square_\square/\square\square}{\square_\square} = \frac{\square_\square\square_\square}{\square_\square\square_\square + \square_\square\square_\square} \ge 0. \tag{A.11}$$

The intuition from Equation (A.11) is that landowners in state $\square$ lose rental income, while $\square$'s landowners gain. Note that the three tax burdens in Equations (A.9) and (A.11) sum up to $-1$. With identical production functions, landowners in state $\square$ ($\square$) lose (gain) rents equal to half of the marginal tax revenues. Therefore, in this model, a state within a country is likely to bear a significant portion of the burden of a tax it levies on a domestically mobile factor. Appendix A.3 use the open-economy model to examine the burden of property taxes.

## A.3 Burden of the property tax

We can use the tax burden analysis for in open economy in Section 1.1 to study the burden of a property tax. Property taxes are typically levied on both land and capital, so the burden of the tax can be decomposed into that from taxing land and that from taxing capital. As we saw above, a tax on land rents is fully borne by landowners, while the tax on mobile capital may be shifted. Similarly, the economic burden of taxes depend on the supply and demand elasticities.

Consider the two-state (city) model where $\square_\square$ is the property tax in state/city $\square$ and $\square_\square$ the tax in state/city $\square$. Then assuming $\square_\square(\square) = \square_\square(\square) = \square(\square)$, the capital rental rates in Equation (A.8) become

$$\square'(\square_\square) = \square + \square_\square; \qquad \square'\left(\overline{\square} - \square_\square\right) = \square + \square_\square. \tag{A.12}$$

[24] Differentiating (A.10) with respect to $\square$, we get

$$\frac{\square\square_\square}{\square\square} = \square_\square'(\square_\square)\frac{\square\square_\square}{\square\square} - (\square+\square)\frac{\square\square_\square}{\square\square} - \square_\square\left(1 + \frac{\square\square}{\square\square}\right); \qquad \frac{\square\square_\square}{\square\square} = \square_\square'(\square_\square)\frac{\square\square_\square}{\square\square} - \square\frac{\square\square_\square}{\square\square} - \square_\square\frac{\square\square}{\square\square}.$$

From (A.8), the first two terms in each expression above cancel out and we can use (A.9) to get the expressions in (A.11).

If $\square_{\square} = \square_{\square}$, capital rental rate $\square$ declines by the full amount of the tax, and capital bears the full burden of the property taxes levied on capital. If instead $\square_{\square} > \square_{\square}$, then $\square_{\square} - \square_{\square}$ will reduce land rents in $\square$ and increase rents in $\square$. In this case, depending on differences in capital demand elasticities, capital will bear the differential tax in part, in full, or more than in full. To see this, we can replace $\square$ by $\square_{\square} - \square_{\square}$ and $\square$ by $\square + \square_{\square}$ in Equations (A.8) and (A.10).

## A.4 Proofs

### A.4.1 Proof of Proposition 1

*Proof.* The proof directly follows from the discussion in Appendix A.2. Equation (A.11) implies that landowners in state $\square$ imposing the tax on capital within its border lose rental income, while state $\square$'s landowners gain. In this model, the immobile factor (land) in a state is likely to bear a significant portion of the burden of a tax the state levies on a domestically mobile factor. □

### A.4.2 Proof of Proposition 2

*Proof.* From Equation (6), the magnitude of the marginal decline in current house prices ($\square_{\square,\square}$) from an additional dollar of pension shortfall $\square$ periods ahesd ($\square_{\square+\square}$) depends on how large the distortion $\square$ is and low far in the future the tax is imposed.

With reasonable parameter values for income and property tax rates, depreciation, and maintenance costs, the capitalization of future pension liabilities in house prices today can have a magnitude of less or greater than one. □

## Table A.1

# Asset Class Detail

The PPD provides detailed breakdowns of the various asset classes invested in by public pensions. This table reports summary statistics for the allocations of the 616 state-year pension plan observations available. The average allocation and the standard deviation of the allocation across pension years are reported, as well as the percent of state-years that had a non-zero allocation to that asset class (short positions are also reported and accounted for in the below). Also reported is whether the asset class is included in our Restricted Benchmark measure. See https://publicplansdata.org/wp-content/uploads/2013/12/Investment-Codebook.xlsx for definitions of Asset Classes.

| Asset Class | Obs. | Average Allocation | Std. Dev Allocation | Percent of State-Years with non-zero Allocation | Included in Restricted Benchmark |
|---|---|---|---|---|---|
| AbsRtrn | 616 | 0.0079 | 0.0212 | 0.2419 | Yes |
| AltInflation | 616 | 0.0009 | 0.0057 | 0.0357 | Yes |
| AltMisc | 616 | 0.0094 | 0.0277 | 0.1932 | Yes |
| Cash | 616 | 0.0169 | 0.0214 | 0.8506 | Yes |
| Commod | 616 | 0.0023 | 0.0092 | 0.1899 | No |
| CoveredCall | 616 | 0.0000 | 0.0001 | 0.0065 | Yes |
| CreditOpp | 616 | 0.0052 | 0.0216 | 0.0990 | Yes |
| DistrssedDebt | 616 | 0.0000 | 0.0004 | 0.0032 | No |
| EQCore | 616 | 0.0002 | 0.0025 | 0.0065 | Yes |
| EQDomesticLarge | 616 | 0.0200 | 0.0599 | 0.2565 | Yes |
| EQDomesticMid | 616 | 0.0006 | 0.0031 | 0.0503 | Yes |
| EQDomesticMisc | 616 | 0.2519 | 0.1696 | 0.8506 | Yes |
| EQDomesticSmall | 616 | 0.0074 | 0.0246 | 0.2565 | Yes |
| EQGlobal | 616 | 0.0083 | 0.0340 | 0.1802 | Yes |
| EQGlobalGrowth | 616 | 0.0000 | 0.0006 | 0.0065 | Yes |
| EQIntlActv | 616 | 0.0001 | 0.0016 | 0.0097 | Yes |
| EQIntlDev | 616 | 0.0125 | 0.0397 | 0.1380 | Yes |
| EQIntlEmerg | 616 | 0.0071 | 0.0193 | 0.2143 | Yes |
| EQIntlMisc | 616 | 0.1225 | 0.0840 | 0.8669 | Yes |
| EQIntlPass | 616 | 0.0008 | 0.0079 | 0.0114 | Yes |
| EQLarge | 616 | 0.0002 | 0.0038 | 0.0016 | Yes |
| EQMicro | 616 | 0.0001 | 0.0011 | 0.0081 | Yes |
| EQMisc | 616 | 0.1002 | 0.1924 | 0.3669 | Yes |
| EQPrivate | 616 | 0.0570 | 0.0575 | 0.7938 | Yes |
| EQSecLend | 616 | 0.0004 | 0.0021 | 0.0568 | Yes |
| EQSmall | 616 | 0.0001 | 0.0008 | 0.0065 | Yes |
| Farm | 616 | 0.0000 | 0.0004 | 0.0114 | No |
| FIAlt | 616 | 0.0109 | 0.0638 | 0.0422 | Yes |
| FIBelowInvestGrd | 616 | 0.0005 | 0.0050 | 0.0097 | Yes |
| FIConv | 616 | 0.0005 | 0.0037 | 0.0471 | Yes |
| FICore | 616 | 0.0171 | 0.0447 | 0.2419 | Yes |
| FICorpBonds | 616 | 0.0008 | 0.0059 | 0.0455 | Yes |
| FIDomestic | 616 | 0.0384 | 0.0960 | 0.3425 | Yes |
| FIEmerg | 616 | 0.0023 | 0.0102 | 0.0763 | Yes |
| FIETI | 616 | 0.0000 | 0.0001 | 0.0146 | Yes |

| Asset Class | Obs. | Average Fund Allocation | Std. Dev Fund Allocation | Percent of State-Years with non-zero Allocation | Included in Restricted Benchmark |
|---|---|---|---|---|---|
| FIFundsFunds | 616 | 0.0000 | 0.0000 | 0.0081 | Yes |
| FIGlobal | 616 | 0.0022 | 0.0134 | 0.0909 | Yes |
| FIHighYield | 616 | 0.0062 | 0.0151 | 0.2403 | Yes |
| FIIntl | 616 | 0.0040 | 0.0144 | 0.1981 | Yes |
| FIInvestGrd | 616 | 0.0035 | 0.0233 | 0.0471 | Yes |
| FILoans | 616 | 0.0001 | 0.0014 | 0.0211 | Yes |
| FIMisc | 616 | 0.1772 | 0.1230 | 0.7825 | Yes |
| FIMortgage | 616 | 0.0011 | 0.0058 | 0.0974 | Yes |
| FINominal | 616 | 0.0001 | 0.0011 | 0.0081 | Yes |
| FIOpp | 616 | 0.0001 | 0.0006 | 0.0227 | Yes |
| FIStructured | 616 | 0.0001 | 0.0012 | 0.0130 | Yes |
| FITIPS | 616 | 0.0096 | 0.0294 | 0.2630 | Yes |
| FITreasury | 616 | 0.0006 | 0.0108 | 0.0227 | Yes |
| GTAA | 616 | 0.0047 | 0.0229 | 0.1315 | No |
| Hedge | 616 | 0.0098 | 0.0258 | 0.2890 | Yes |
| HedgeEQ | 616 | 0.0008 | 0.0069 | 0.0519 | Yes |
| Infrast | 616 | 0.0012 | 0.0066 | 0.1185 | No |
| MLP | 616 | 0.0010 | 0.0049 | 0.0909 | No |
| MultiClass | 616 | 0.0037 | 0.0121 | 0.1510 | No |
| NatResources | 616 | 0.0004 | 0.0036 | 0.0146 | No |
| Opp | 616 | 0.0009 | 0.0048 | 0.1445 | No |
| OppDebt | 616 | 0.0005 | 0.0047 | 0.0146 | Yes |
| OppEQ | 616 | 0.0002 | 0.0014 | 0.0162 | Yes |
| Other | 616 | 0.0020 | 0.0072 | 0.7192 | Yes |
| PrivateDebt | 616 | 0.0011 | 0.0069 | 0.0519 | No |
| PrivatePlacement | 616 | 0.0009 | 0.0053 | 0.0519 | No |
| PrivRealEstate | 616 | 0.0008 | 0.0061 | 0.0633 | No |
| RealAssets | 616 | 0.0051 | 0.0155 | 0.2159 | No |
| RECore | 616 | 0.0002 | 0.0032 | 0.0049 | No |
| REIT | 616 | 0.0004 | 0.0019 | 0.0844 | Yes |
| RelativeRtrn | 616 | 0.0000 | 0.0001 | 0.0162 | Yes |
| REMisc | 616 | 0.0534 | 0.0385 | 0.8377 | Yes |
| RENonCore | 616 | 0.0002 | 0.0029 | 0.0049 | No |
| RiskParity | 616 | 0.0017 | 0.0112 | 0.0568 | Yes |
| Timber | 616 | 0.0019 | 0.0072 | 0.1347 | No |

**Table A.2**

## 2002-2014 Pension Returns and 2015-2018 House Prices

This table presents estimates from a state border discontinuity design model where the dependent variable is the logarithm of the sales price, in thousands of dollars, of a residential property. The explanatory variable of interest is based on invested assets' cumulative performance over the period 2002-2014 in the pension plans associated with the state in which the focal property is located. The sample is restricted to property transactions involving single-family residences in counties sharing a border with an adjacent state that are within 50 miles of that border and have a transaction price between the typical home values in the the bottom and top market tiers as calculated at the county-month level by the ZHVI. Fixed effects for the county border group of the property interacted with the calendar year of the transaction, as well as covariates for the distance to the state border and the income per capita at the state-year level, are included throughout. These specifications also control for property type by including six interacted property characteristic fixed effect cells (square footage of structure, square footage of lot, age of building, number of bedrooms, number of bathrooms, and number of stories). Column (1) is the baseline regression described above where the dependent variable is the logarithm of the sales price, in thousands of dollars, for transactions in the years 2015-2018 and the primary variable of interest is the cumulative pension fund performance from 2002-2014. Column (2) is the same as column (1), but replaces the primary variable of interest with the cumulative pension fund performance from 2002-2014 in excess of the benchmark performance for each asset class the fund is invested in. Column (3) is the same as column (1), but replaces the primary variable of interest with the cumulative pension fund performance from 2002-2014 that would have occurred based on the fund's asset allocations, had it earned the benchmark performance for each asset class. Column (4) is the same as column (3), but restricts attention to assets that have less potential to be localized (i.e., bonds and equities, and funds that invest in them, rather than commodities, private debt, and real estate). Reported $t$-statistics in parentheses are heteroskedasticity-robust and clustered at both the zip and transaction month level. ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

| | Log Sales Price $('000s) | | | |
|---|---|---|---|---|
| | (1) | (2) | (3) | (4) |
| 2002-2014 Cum. Port. Ret. | 0.323*** | | | |
| | (6.87) | | | |
| 2002-2014 Cum. Excess Ret. | | 0.604*** | | |
| | | (4.93) | | |
| 2002-2014 Cum. BenchMk Ret. | | | 0.311*** | |
| | | | (4.26) | |
| 2002-2014 Cum. (Restr.) BenchMk Ret. | | | | 0.304*** |
| | | | | (4.07) |
| Border Distance | X | X | X | X |
| State-Year Income PC | X | X | X | X |
| Border Group-Tran Year FE | X | X | X | X |
| 6 Prop Chars FE | X | X | X | X |
| Observations | 129,940 | 129,940 | 129,940 | 129,940 |
| Adj. $R^2$ | 0.864 | 0.863 | 0.863 | 0.863 |

**Table A.3**

## Rolling Pension Returns and House Prices

This table presents estimates from a state border discontinuity design model where the dependent variable is the logarithm of the sales price, in thousands of dollars, of a residential property. The explanatory variable of interest is based on invested assets' cumulative performance in the pension plans associated with the state in which the focal property is located, but only in the years prior to the transaction. The sample is restricted to property transactions involving single-family residences in counties sharing a border with an adjacent state that are within 50 miles of that border and have a transaction price between the typical home values in the the bottom and top market tiers as calculated at the county-month level by the ZHVI. Fixed effects for the county border group of the property interacted with the calendar year of the transaction, as well as covariates for the distance to the state border and the income per capita at the state-year level, are included throughout. These specifications also control for property type by including six interacted property characteristic fixed effect cells (square footage of structure, square footage of lot, age of building, number of bedrooms, number of bathrooms, and number of stories). Column (1) is the baseline regression described above where the dependent variable is the natural logarithm of the sales prices, in thousands of dollars, that transact in a given year and the primary variable of interest is the cumulative pension fund performance from 2002 until the year prior to that particular transaction. Column (2) is the same as column (1), but replaces the primary variable of interest with the cumulative pension fund performance from 2002 until the year prior to that particular transaction in excess of the benchmark performance for each asset class the fund is invested in. Column (3) is the same as column (1), but replaces the primary variable of interest with the cumulative pension fund performance from 2002 until the year prior to that particular transaction that would have occurred based on the fund's asset allocations, had it earned the benchmark performance for each asset class. Column (4) is the same as column (3), but restricts attention to assets that have less potential to be localized (i.e., bonds and equities, and funds that invest in them, rather than commodities, private debt, and real estate). Reported $t$-statistics in parentheses are heteroskedasticity-robust and clustered at both the zip and transaction month level. ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

| | Log Sales Price $('000s) | | | |
|---|---|---|---|---|
| | (1) | (2) | (3) | (4) |
| 2002-Sale Cum. Port. Ret. | 0.222*** | | | |
| | (4.67) | | | |
| 2002-Sale Cum. Excess Ret. | | 0.363*** | | |
| | | (5.04) | | |
| 2002-Sale Cum. BenchMk Ret. | | | 0.173*** | |
| | | | (2.70) | |
| 2002-Sale Cum. (Restr.) BenchMk Ret. | | | | 0.167*** |
| | | | | (2.61) |
| Border Distance | X | X | X | X |
| State-Year Income PC | X | X | X | X |
| Border Group-Tran Year FE | X | X | X | X |
| 6 Prop Chars FE | X | X | X | X |
| Observations | 712,505 | 712,505 | 712,505 | 712,505 |
| Adj. $R^2$ | 0.879 | 0.878 | 0.878 | 0.878 |

**Table A.4**

## House Prices and Pension Windfalls: Border vs. Interior Counties

This table presents estimates from a state border discontinuity design model where the dependent variable is the sales price, in thousands of dollars, of a residential property that transacted in 2015-2018. The explanatory variable of interest is based on invested assets' cumulative performance in the pension plans associated with the state in which the focal property is located from 2002-2014, multiplied by initial assets per property in 2001 to create a measure of additional pension shortfall per property due to asset performance (Windfall). The sample is restricted to property transactions involving single-family residences that have a transaction price between the typical home values in the the bottom and top market tiers as calculated at the county-month level by the ZHVI. Fixed effects for the county border group of the property interacted with the calendar year of the transaction, as well as covariates for the distance to the state border (column (1) only) and income per capita at the state-year level, are included throughout. These specifications also control for property type by including six interacted property characteristic fixed effect cells (square footage of structure, square footage of lot, age of building, number of bedrooms, number of bathrooms, and number of stories). Columns (1) and (2) restrict the sample to properties located in counties sharing a border with an adjacent state that are within 50 miles of that border. Column (3) restricts the sample to properties that do not meet the definition of being in a border county (i.e., only counties in the interior of the state). Column (2) differs from column (1) of Table 3 only in the exclusion of a measure of distance to the state border. Reported $t$-statistics in parentheses are heteroskedasticity-robust and clustered at both the zip and transaction month level. ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

| | Sales Price $('000s) | | |
|---|---|---|---|
| | (1) | (2) | (3) |
| 2002-2014 Windfall Per Property $('000s) | 3.662*** | 2.510*** | 0.920*** |
| | (6.65) | (8.46) | (3.82) |
| 2002-2014 Windfall Per Prop $('000s) × Border Distance (mi) | -0.113*** | | |
| | (-2.81) | | |
| Border Distance | X | | |
| State-Year Income PC | X | X | X |
| Border Group-Tran Year FE | X | X | X |
| 6 Prop Chars FE | X | X | X |
| Sample | Border | Border | Interior |
| Observations | 129,940 | 129,940 | 769,004 |
| Adj. $R^2$ | 0.816 | 0.813 | 0.717 |

**Table A.5**

# County-Level Municipal Finances: Border vs. Interior Counties

This table presents county-level regressions of various financial outcomes on an indicator for whether the county is on a state border. The sample includes counties in states that qualify for our regression sample, depicted in Figure A.2. These specifications include state fixed effects to account for differences in financial ratios across states. Information regarding the finances of local governments (counties, cities, and other local municipalities) is aggregated to the county level by the U.S. Census Bureau and available for the years 2007 and 2012. We estimate separate regressions for these two reporting years. The estimates suggest that border counties are comparable to counties on the interior of their state with respect to the financial health of local governments. Reported $t$-statistics in parentheses are heteroskedasticity-robust. ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

| | Border Relative To Interior | | | Border Relative To Interior | |
|---|---|---|---|---|---|
| Variable | 2007 | 2012 | Variable | 2007 | 2012 |
| Total Revenues | 0.04 | 0.17* | Total Expenditures | 0.01 | 0.19* |
| Per Capita | (0.13) | (1.87) | Per Capita | (0.05) | (1.96) |
| Revenues From Federal Govt | 0.00 | 0.01 | Capital Expenditures | -0.03 | 0.01 |
| Per Capita | (0.11) | (0.96) | Per Capita | (-0.87) | (0.31) |
| Revenues From State Govt | 0.04 | 0.07*** | Education Expenditures | -0.06 | 0.00 |
| Per Capita | (0.50) | (3.34) | Per Capita | (-0.64) | (-0.11) |
| Total Taxes | -0.06 | -0.03 | Safety Expenditures | -0.01 | 0.00 |
| Per Capita | (-0.62) | (-0.78) | Per Capita | (-0.31) | (0.56) |
| Property Taxes | -0.05 | -0.04 | Utility Expenditures | 0.08 | 0.05 |
| Per Capita | (-0.75) | (-1.20) | Per Capita | (1.31) | (0.92) |
| Sales Taxes | 0.00 | 0.00 | Short-Term Debt | -0.01 | 0.00 |
| Per Capita | (0.09) | (0.62) | Per Capita | (-0.95) | (0.53) |
| Income Taxes | -0.02 | 0.00 | Long-Term Debt | 0.82 | 0.65* |
| Per Capita | (-0.85) | (0.29) | Per Capita | (1.39) | (1.87) |
| Other Taxes | 0.00 | 0.00 | | | |
| Per Capita | (0.12) | (0.43) | | | |

## Table A.6

# Pension Windfalls and House Prices in Border Counties Weighted to Match Interior Counties

This table presents estimates from a weighted least squares state border discontinuity design model where the dependent variable is based on the sales price, in thousands of dollars, of a residential property. The explanatory variable of interest is based on invested assets' cumulative performance in the pension plans associated with the state in which the focal property is located from 2002-2014, multiplied by initial assets per property in 2001 to create a measure of additional pension shortfall per property due to asset performance (Windfall). The sample is restricted to property transactions involving single-family residences in counties sharing a border with an adjacent state that are within 50 miles of that border and have a transaction price between the typical home values in the the bottom and top market tiers as calculated at the county-month level by the ZHVI. Fixed effects for the county border group of the property interacted with the calendar year of the transaction, as well as covariates for the distance to the state border and the income per capita at the state-year level, are included throughout. These specifications also control for property type by including six interacted property characteristic fixed effect cells (square footage of structure, square footage of lot, age of building, number of bedrooms, number of bathrooms, and number of stories). These specifications are similar to that of column (1) in Table 3, but have been re-weighted such that these border counties match interior counties on the specified dimension(s). Columns (1)-(4) use weights chosen to match the four variables in Table A.5 with statistically significant differences between border and interior counties. Column (5) uses weights chosen to match all four variables jointly. Reported $t$-statistics in parentheses are heteroskedasticity-robust and clustered at both the zip and transaction month level. ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

| | Sales Price $('000s) | | | | |
|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) |
| 2002-2014 Windfall Per Property $('000s) | 2.413*** | 2.414*** | 2.414*** | 2.414*** | 2.407*** |
| | (8.27) | (8.35) | (8.30) | (8.30) | (8.43) |
| Border Distance | X | X | X | X | X |
| State-Year Income PC | X | X | X | X | X |
| Border Group-Tran Year FE | X | X | X | X | X |
| 6 Prop Chars FE | X | X | X | X | X |
| 2012 Balance Variable(s) | Total Revenues, PC | Revenues From State Govt, PC | Total Expenditures, PC | Long-Term Debt, PC | Cols. (1)-(4) |
| Observations | 129,940 | 129,940 | 129,940 | 129,940 | 129,940 |
| Adj. $R^2$ | 0.820 | 0.822 | 0.821 | 0.821 | 0.825 |

## Table A.7

# State Responses to Shortfalls

This table presents regressions of various economic outcomes on lagged state pension shortfalls. Observations are at the state-year level. Column (1) regresses employer pension contributions per property on the prior year's state-level pension shortfall per property after including state fixed effects. Columns (2-5) are the same as column (1), but the dependent variables are employee pension contributions per property, secondary education appropriation per property, and annual changes in the percentages of rural and urban roads in poor condition, respectively. Reported $t$-statistics in parentheses are heteroskedasticity-robust and clustered at the state level. ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

| | Employer Pension Contribution | Employee Pension Contribution | Secondary Education Appropriation | Change in Percent of Rural Roads in Poor Condition | Change in Percent of Urban Roads in Poor Condition |
|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) |
| Lagged Shortfall Per Property | 0.0213*** | 0.00379*** | -0.00251*** | 0.0161* | 0.0160* |
| | (6.12) | (5.39) | (-2.73) | (1.88) | (1.85) |
| State FE | X | X | X | X | X |
| Observations | 806 | 806 | 450 | 383 | 393 |
| Adj. $R^2$ | 0.606 | 0.802 | 0.942 | 0.046 | -0.045 |

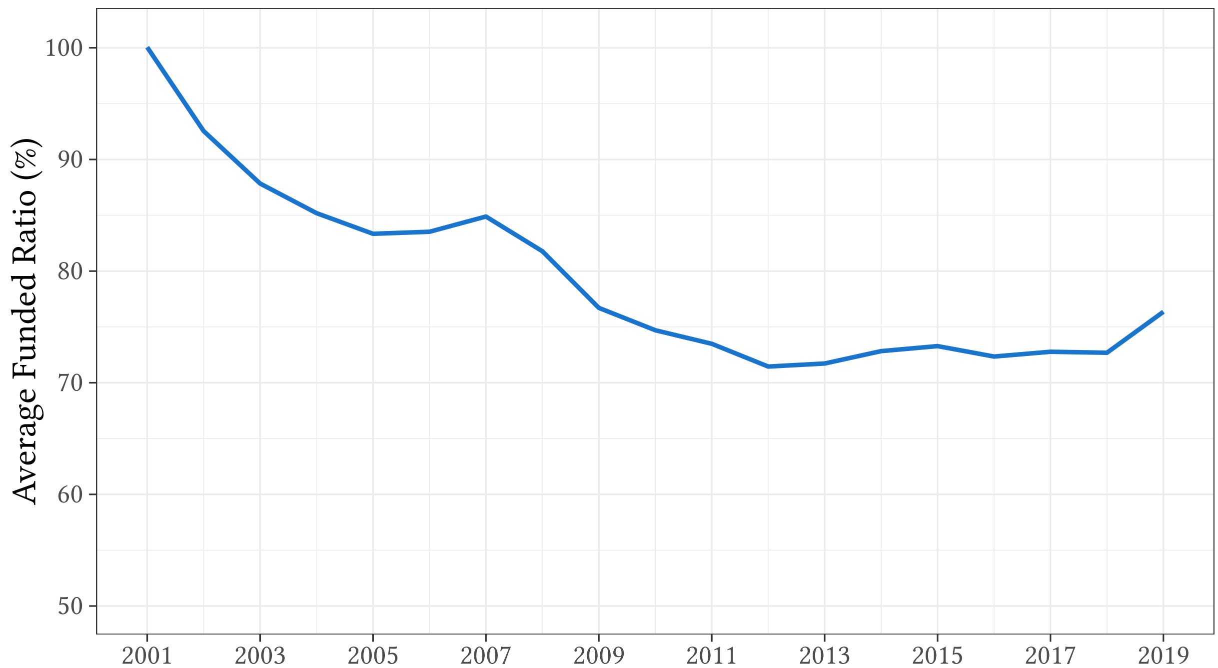


**Figure A.1. Average Funded Ratio** This figure presents the time-series of average ratio of pension assets to liabilities, the actuarial funded ratio, at the state-year level for the Public Plans Data (PPD) database provide by the Center for Retirement Research (CRR) at Boston College. Actuarial funded ratio is given by ActFundedRatio_GASB, which is ActAssets_GASB divided by ActLiabilities_GASB in the database.