# Risk on the Run: Public Pensions, Migration Risk and Pivot to Alternatives

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**Abstract**

How does local debt sustain with a mobile tax-base? This paper provides new evidence on how pension underfunding shapes inter-state migration. Exploiting plausibly exogenous variation in funding gaps driven by relative portfolio performance, I find that widening pension deficits significantly increase net out-migration of households and income flows from underfunded to better-funded states. This mobility erodes the tax base, raising the effective cost of funding pensions and constraining governments' ability to close shortfalls through higher contributions. I further show that states facing higher "migration risk"—those that face larger out-migration when local taxes rise—systematically under-contribute relative to actuarially required levels and shift pension assets toward opaque, higher-yielding alternatives such as private equity and hedge funds. This pattern reveals a novel mechanism by which household mobility exacerbates both fiscal fragility and risk-taking incentives. To interpret these findings, I develop a spatial model of tax-setting and pension investment under endogenous migration. The model highlights a migration–funding–risk feedback loop: governments that fail to internalize out-migration overestimate tax capacity, under-contribute, and induce fiduciaries to pursue riskier portfolios.

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# 1 Introduction

Defined-Benefit (DB) public pension plans across the United States face persistent underfunding, with unfunded liabilities now being the largest off-balance-sheet obligation of state and local governments (Giesecke and Rauh, 2023). Funding ratios remain depressed post–Great Recession, raising concerns about fiscal sustainability. While prior work studies the fiscal, economic, and financial consequences of underfunding (Novy-Marx and Rauh, 2012; Andonov et al., 2017; Anzia, 2022; Aiello et al., 2024), much less is known about how pension shortfalls interact with household mobility. This paper studies (i) how pension funding affects taxpayer migration and (ii) how migration risk conditions states' contribution policies and investment risk-taking by public plans.

When defined-benefit public pension deficits deepen, state and local governments face greater claims on their resources, tightening already constrained budgets. Balanced budget rules require governments to close these gaps by borrowing, cutting expenditures, or raising revenue,[1] since pension obligations are part of the operating budget. Pension shortfalls thus serve as indicators of fiscal health and public service provision: limited fiscal capacity may undermine service quality while imposing higher tax burdens on residents. As gaps in tax liability and amenities between states widen, households and firms may relocate to jurisdictions with lower taxes or better amenities (Moretti and Wilson, 2017). Residents may also anticipate and resist bearing the expected future costs of growing liabilities (Inman, 1982). Indeed, I find that higher underfunding levels are associated with higher corporate tax rates, higher state income tax rates and lower public school expenditures, along with net-migration. Such migration erodes the tax base, further weakening public finances and potentially triggering a vicious cycle in which underfunding and out-migration reinforce each other. This dynamic raises a central question: can pension debt remain sustainable when migration diminishes governments' ability to raise future taxes? The first core question of the paper is: does underfunding causally induce out-migration?

A second phenomenon is the sharp increase in public pension allocations to alternative investments. Since the early 2000s, public pension allocations to private equity, hedge funds, real estate, and other alternatives have expanded substantially, reaching a value-weighted average allocation of 35% of total invested assets in 2023. This shift has been attributed to a range of determinants - political boards (Andonov et al., 2017); evolving discount rate assumptions (Giesecke and Rauh, 2023); beliefs about asset performance (Begenau et al., 2024); and gambling for resurrection (Mittal, 2024). Lucas and Zeldes (2009), Pennacchi and Rastad (2011) and Lu et al. (2019) model optimal risky-asset allocation as a joint problem of minimizing fiscal costs, while investing in risky securities. Novy-Marx and Rauh (2014), draws from the classic "Tiebout" framework (Tiebout, 1956), and posits potential out-migration if states raise contribution to fix underfunding. This suggests an overlooked state variable: migration risk in the joint problem — the ease or difficulty of raising revenue without losing the tax base. This leads to the second core question of this paper: Conditional on funding needs, how do high-migration-risk states contribute to their pensions and how does that influence portfolio choice?

To study the first question, I motivate my baseline empirical design as a location-choice framework for migrant households. This design closely parallels the canonical specification used by Moretti and Wilson (2017) in their exploration of how state-tax differentials induce the location choice of star scientists. My contribution is to posit that tax gaps are a monotonic function of funding gaps between origin and destination states, since balanced budget rules tie underfunding to higher future taxes. However, measuring the pass-through of funding status on migration is riddled with several endogeneity concerns. Underfunded states could represent fiscally profligate states which attract households due to better public services, on the other hand, fiscally conservative states could see greater outflows due to time-varying local economic shocks. There could also be potential measurement error as more underfunded states could understate their liabilities by differing discount rate assumptions. This requires a causal strategy that isolates plausibly exogenous variation in pension funding.

But what explains variation in public pension funding status across states? I run a cross-sectional variance decomposition of public pension funding status with respect to investment performance, changes in contribution policy and changes in actuarial projections. I find that investment performance explains a large portion of the explained variation through time followed by changes in contribution rates. This motivates my instrument construction: I focus on cumulative abnormal performance relative to an ex-ante benchmark as the primary instrument, which is similar to the instrument used in Aiello et al. (2024), and also validate the results with an alternative instrument - public unionization density, as used in Mittal (2024). Variation in public unionization across states likely influences their contribution policy. The exclusion restriction for both instruments is that they affect pension funding status, which pins the state's relative fiscal position and they are orthogonal to other drivers of migration between the origin and destination, after sufficiently controlling for destination-origin pair and time-varying variation.

Therefore, a shock to funding gap influences the present and future gap in taxes or public services that then influences migration outcomes. Using the IRS's County-to-County taxpayer migration data, I incorporate

[1] An additional channel is through reforms that reduce pension benefits or increase employee contributions. Drawing on National Association of State Retirement Administrators (NASRA) reports between 2005 and 2022, I classify reforms into four categories: (i) higher employee contribution rates, (ii) reduced benefits, (iii) plan design changes, and (iv) other adjustments. Details are provided in the appendix.

a rich set of origin-destination and time-varying fixed effects which helps evade confounding variation driving migration patterns. Restricting data to interstate migration, I find that within a origin-destination pair, a one standard deviation higher pension funding gap between the destination and origin counties - which means the destination county's pension funding ratio is 20 percentage points higher than origin - explains 4.4 percentage points higher odds of migration which translates to an annual county-to-county migration probability of 0.176 basis points. For a county the size of Cook County (2 million non-migrant tax returns in 2022), this translates into an a higher annual adjusted gross income (AGI) flow of roughly $3.3 million, raising the bilateral outflow to $43 million. To quantify aggregate effects, I estimate county-level 2-SLS regressions of net out-migration on associated pension funding ratios. I find that a one standard deviation lower funding ratio implies an annual net adjusted gross income outflow of approximately $540 million for a county the size of Cook County.

I also examine the dynamics of migration - how quickly do residents respond when funding gap widens. I show that increasing funding gaps predict higher taxes in the origin relative to destination counties or states with a lag. In the OLS specification I find that when the gap between destination and origin states' funding ratios widens, it increases the odds of out-migrant returns relative to non-migrant returns increases by about 2% over 10 years. After instrumenting increase in funding gap by relative return surprises the odds of out-migration rise by 8% within six years of the gap widening, before tapering to roughly 5% after ten years. This pattern remains when weighting migrants by adjusted gross income. Overall, the evidence suggests that as pension funding deteriorates in origin states relative to destinations, residents increasingly relocate to better-funded regions.

A mobile tax base poses a threat to the sustainability of state and local pension debt. For pension fiduciaries, this raises the question of how contribution policy and investment strategy should adapt when underfunding persists. State governments face a trade-off if increasing contributions through taxes causes out-migration. To assess how much revenue a state can realistically raise through taxation without eroding its tax base, I estimate households' migration response to increases in state and local tax liabilities. States that experience greater net out-migration when taxes rise have more limited fiscal capacity, or higher "migration risk," than states with flat or net in-migration. Measuring migration risk is challenging, however, since observed migration may also reflect other confounding factors. I construct three complementary measures. The first is a state-level regression of net out-migration on changes in income tax rates, controlling for time-varying state characteristics; this provides a transparent, though endogenous, ranking of states' relative migration propensities. The second applies the location-choice framework of Moretti and Wilson (2017). The third exploits plausibly exogenous increases in tax liabilities created by the State and Local Tax (SALT) deduction cap introduced under the Tax Cuts and Jobs Act 2017, which disproportionately affected taxpayers in high-tax states.

All three measures help me rank states by their relative migration risk. I find that states with relatively higher migration risk systematically under-contribute relative to the actuarially required contribution level. The pension funds of these states also have worse funding status. This forms the basis of my second question - how migration risk might influence public pension's portfolio choice. I investigate the heterogeneity of public pension portfolio choice in response to higher migration risk for a given level of underfunding. Across the three measures, pension plans in high-migration risk states respond differently to underfunding: for each one–percentage point increase in deficits, they reallocate an additional 68–141 basis points into alternative asset classes such as private equity, real estate, and hedge funds. This gap persists over time, amounting to 2–7 percentage points higher allocations on average compared with plans in low-risk states. Meanwhile, allocations to fixed income and equities remain flat or decline, consistent with a portfolio re-balancing toward riskier, high-yield assets.

What explains the link between migration risk and pension fund investment policy? Why would states experiencing outflows under-contribute to their pensions, and why should fiduciaries care about migration risk? To address these questions, I develop a two-period spatial model that rationalizes the empirical patterns. In the model, household location choice responds to local amenities and taxes, both of which depend on a state's pension funding status. State governments are represented by a political executive that sets taxes to minimize political costs, subject to pension obligations and exogenous spending, and by a pension fiduciary that determines investment policy. The fiduciary, who is risk-averse, maximizes payoffs proportional to end-of-period funding status by choosing the share of pension assets invested in risky securities. The optimal risky share is decreasing in ex-ante funding levels, state contributions, the sensitivity of funding to performance, and the manager's risk aversion.

Crucially, I assume political executives fail to internalize migration dynamics, which depend on other states' policies and households' idiosyncratic preferences. As a result, underfunded states experience out-migration but overestimate future tax base growth, leading them to under-contribute and implicitly induce fiduciaries to take on greater investment risk. I then contrast this with a social planner equilibrium, where the government fully internalizes both investment risk and the migration response. The reduced-form optimal risky share in this setting is inversely related to the efficiency of public service provision and declines with out-migration. Calibrating the social planner optimum, I show that states take excess risk on their pension portfolios through allocation to alternative asset classes.

This paper thus makes two novel contributions: First, empirically, I show that pension underfunding and

funding gaps between jurisdictions predict migration flows, shrinking the local tax base. Second, I document that states facing higher migration risk systematically under-contribute and reallocate toward riskier, higher-yielding assets such as private equity and hedge funds. In addition, the model formalizes these dynamics in a migration–funding–risk feedback loop, clarifying how pension underfunding ties to migration, contributions and pension portfolio choice.

**Related Literature.** A large literature examines public pension underfunding. Novy-Marx and Rauh (2009) document the post-recession surge in unfunded liabilities; Novy-Marx and Rauh (2012, 2014) study the fiscal imbalance and revenue demands created by pension shortfalls, with Novy-Marx and Rauh (2014) linking prospective gap-filling contributions to Tiebout-style mobility (Tiebout, 1956). Anzia (2022) show that rising pension expenditures reduce public employment, and Aiello et al. (2024) show that investment windfalls capitalize into local property prices. Relative to this work, I focus on long-horizon local effects that manifest through *observed* interstate net migration, using a bilateral location-choice design estimated via two-stage least squares.

On investment policy, Lucas and Zeldes (2009) analyze the trade-off between tax smoothing and risky asset exposure; Andonov et al. (2017) show that higher liability discount rates—chosen by politically influenced boards—spur risk-taking; and Giesecke and Rauh (2023) argue that discount-rate practices understate underfunding. Lu et al. (2019) formalize reach-for-yield behavior in state and local funds; Myers (2023) document excessive risk-taking by cities post-crisis; Begenau et al. (2024) attribute the rise in alternatives to return expectations and experience; and Mittal (2024) find that the most underfunded plans match with smaller private equity funds and earn lower total returns. A complementary literature studies how taxes and local policies shape the geography of economic activity (Suárez Serrato and Zidar, 2016; Giroud and Rauh, 2019; Fajgelbaum et al., 2019; Rauh, 2022; Rauh and Shyu, 2024).

This paper links these strands by showing that pension underfunding shifts migration patterns and that the resulting *migration risk* is a first-order determinant of contribution policy and portfolio tilts toward alternatives.

**Road Map.** Section 2 describes the data. Section 3 presents aggregate trends in state pension plans. Section 4 shows the association of pension funding status to local economic outcomes. Section 5 presents the empirical design and results of the relationship between pension funding and migration. Section 6 develops the state-wise migration risk measure and explores the relationship between migration risk and pension portfolio allocations. Section 7 contains the model. Section 8 concludes.

## 2 Data

The analysis draws on several complementary data sources. The core dataset is the Center for Retirement Research's Public Plans Database (PPD), which reports annual balance-sheet details, asset allocations, and performance metrics for 228 state and local pension systems for the time period of 2001-2023. I use the reported Actuarial Funding Ratio which is the ratio of the pension funds assets and liabilities in the dataset and underfunding is one minus the funding ratio. I get variables on plan-wise employer and employee contribution rates and amounts, investment return assumption, inflation assumption, methods of valuation, amortization and smoothing period, and required contribution rates from this dataset. I also use the plan-wise portfolio allocation data which gives the break up of the pension assets by equity, fixed income, and alternative investments such as private equity, hedge funds, real estate and others.

For migration data, I use Internal Revenue Service's Migration Files which helps track both aggregate as well as county-to-county and state-to-state migration of tax returns and adjusted gross income for the time period 2001-2022. I use the county-to-county outflows to for the baseline regression and get the number of out migrant returns and adjusted gross income (AGI to a different state, and do not use data on migration within a state or to or from abroad. I aggregate to county-level to measure annual net migration by county by subtracting inflows from outflows.

I obtain information on state finances from the Census Bureau's Annual Survey of State and Local Government Finances. I obtain state marginal income-tax rates from NBER's TAXSIM, and state corporate tax rates from the Tax Foundation. Public-school spending data are sourced from the National Center for Education Statistics. Finally, household-level variables—including wages, rents, employment growth, and property-tax liabilities and other miscellaneous variables—are drawn from the American Community Survey and the Current Population Survey. Summary statistics for all variables used in the analysis are reported in Appendix Table 4.

## 3 Stylized Facts

Aggregate funding ratios have trended downwards since the early 2000s. Figure 1 documents a steady decline in the value-weighted funding ratio for state and local pension plans—from full funding in 2001 to below 75 percent by 2023. Panel (b) dis-aggregates the series by administering authority, revealing that city- and school-district plans became markedly more underfunded than their state and county counterparts during the 2010s, although

the gap has narrowed in recent years. Simultaneously public pension allocations have pivoted towards alternative

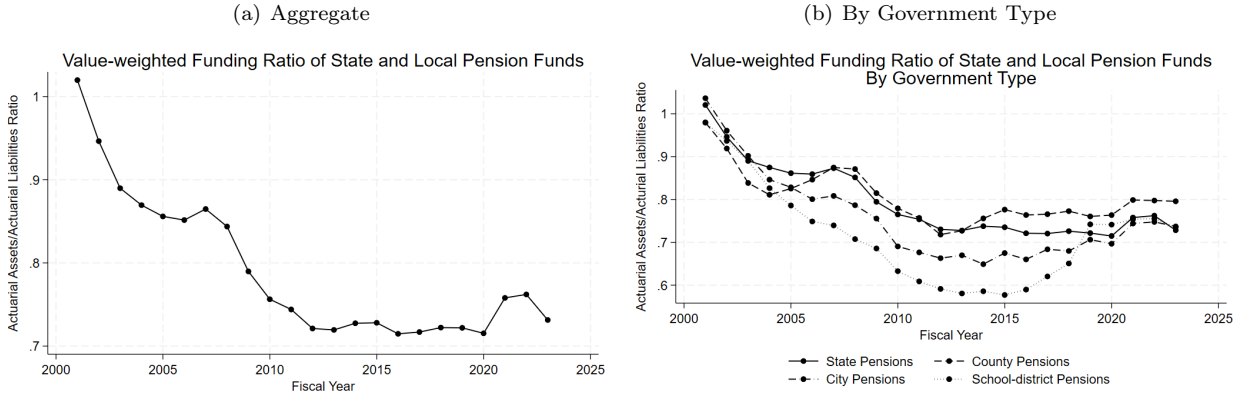



Figure 1: Declining Aggregate Funding Status

These figures plot the value-weighted ratio of pension fund actuarial assets to liabilities. Data comes form Public Pensions Database from the Center for Retirement Research. Panel (a) shows the aggregate funding ratio and panel (b) shows the ratios by government type. Data ranges from 2001-2023.

pension classes such as private equity, hedge funds, real estate and other asset classes from equities and fixed income securities. Figure 2 shows the growing trend towards alternative asset classes since the great recession, with alternative asset classes now accounting for over a quarter of total allocation.

### Variance Decomposition of Pension Underfunding

What explains variation in pension underfunding? Past literature [2] has identified three main determinants of pension underfunding, which are: (i) investment performance (ii) changes in contribution rates and benefits policy (iii) changes in actuarial assumptions such as changes in discount rates, expected rate of return and payroll growth assumptions. To understand what explains the cross-sectional variation in deteriorating pension funding status arising from the three main determinants, we can utilize the following formulation:

$$dU_{it,t-h} = g\big(\Delta \text{ Investment Performance (IP)}_{it}, \Delta\text{Contribution Policy (CP)}_{it}, \Delta\text{Actuarial Projections (AP)}_{it}, \nu_{it}\big)$$

I run year-by-year regressions of cumulative changes in funding status relative to 2001, on cumulative investment return performance, contribution growth and change in normal costs predicted by changes in actuarial assumptions. I then carry out a Shapley decomposition of the explained variation.

Figure 3 documents the variance decomposition. Investment performance remains the biggest drivers of pension funding variation across states initially explain anywhere between 50-75% of variation. Contribution growth has recently become an important explanation, with around one-third of the explained variation rising to about 43% in 2023, which is consistent with the findings in Giesecke and Rauh (2023). Variation in actuarial projections are relatively less influential, even though discount rate assumptions being linked to investment returns can be absorbed by the portion explained by investment performance. Overall, investment performance and contribution policies explain cross-sectional variation in pension funding trends in the data.

## 4 Tracking Pension Funding's Association with State Outcomes

Figure 1 shows that aggregate funding status of public pensions have deteriorated over time since 2001. Since persistent pension underfunding may have a lagged but protracted effect on real outcomes. I adopt a local projection methodology - which tracks effects of the independent variable over long time periods - to trace the lagged impact of pension underfunding. Local projections estimate the response of a variable to an independent variable at each horizon. This is tantamount to running separate regressions run of different leads of the dependent variable on present underfunding. For horizon $h = 0, \ldots, H$:

$$Y_{s,t+h} = \beta_h U_{s,t} + \gamma' X_{s,t} + \delta_s + \tau_{t+h} + \varepsilon_{i,t+h}, \tag{1}$$

where $U_{s,t}$ is underfunding in state $s$ in time period $t-1$, $Y_{s,t+h}$ being the outcome of interest. Each specification also contains the first and second lags of the dependent variable to control for serial correlation over time. $X_{s,t}$

[2] Refer to Giesecke and Rauh (2023) for review

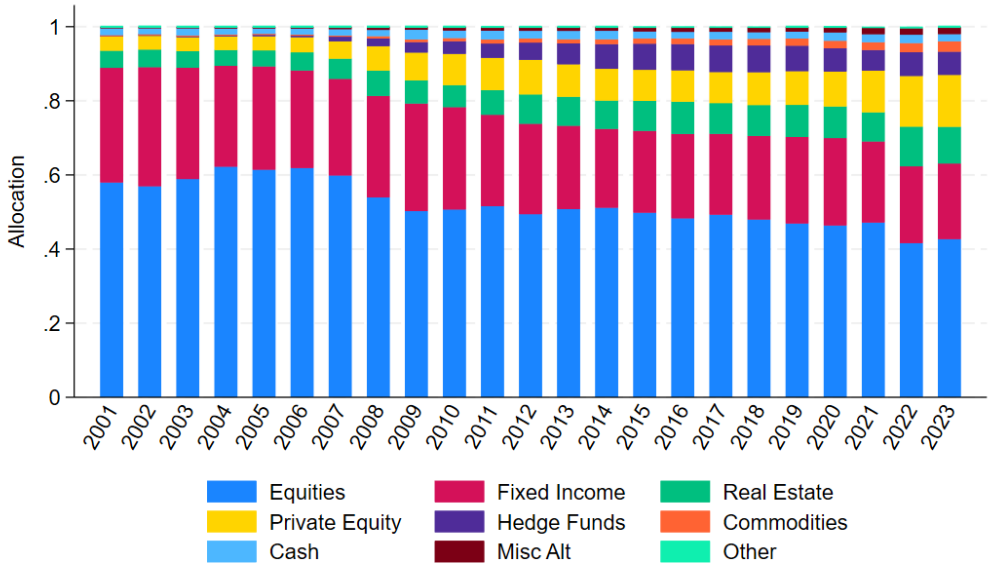



Figure 2: Allocation Trends

This figure plots the aggregate portfolio allocations by asset class for public pension funds. Data comes from Center for Retirement Research's Public Pensions Database and ranges 2001-2023.

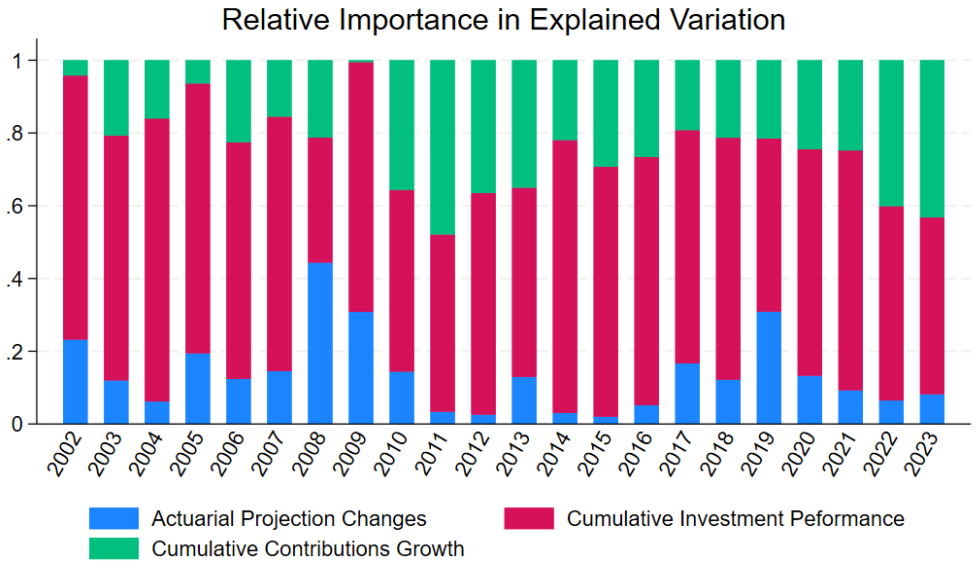



Figure 3: Variance Decomposition

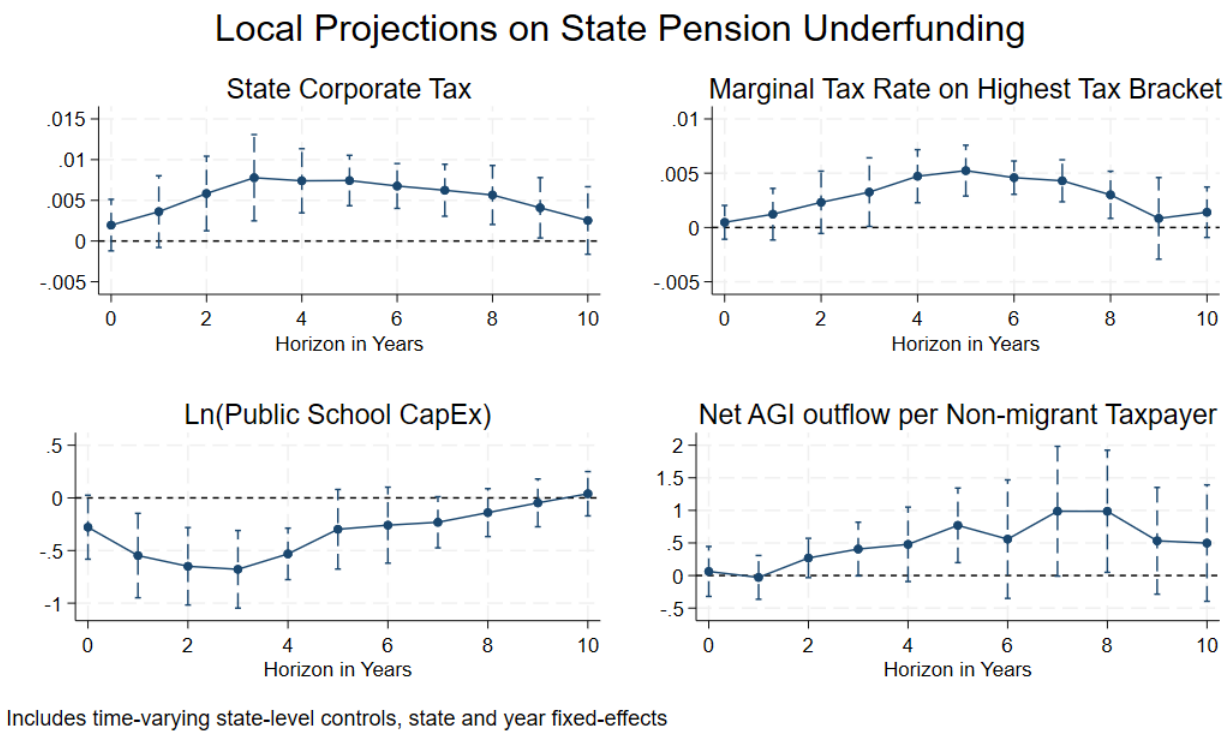



Figure 4: Local Projections - Tracing the impact of state pension underfunding over long horizons

includes state-time varying controls and state and time fixed effects. I include state-wise aggregate pension controls which includes the lagged contribution growth rate and dummy for whether any of the state's constituent pensions changed investment return assumption, inflation assumptions, amortization methods or other actuarial assumptions that could affect total pension liability, in the past year. Residualizing underfunding variable would remove their confounding effect and isolate the impact of underfunding-levels net of growth in contributions or changes in assumptions. Other state level controls include the lagged outstanding debt to GDP ratio which reflects a confounding element in the state's capital budget residualizing the effect of the state being a more heavy borrower in general, state-level employment growth and GDP growth to account for time-varying factors that may effect revenue levels and state's tax base. I also control for the share of republican representatives the state elected to the US house representatives in the lagged election year to control for state's political leaning. I use this empirical design to track the impact on state's corporate tax rates, state's marginal top marginal income tax rate and states' logged total public school capital expenditure. I omit using state employment growth as a control when I use net adjusted gross income (AGI) outflow per non-migrant tax-payer as the dependent variable might be a driver. I additionally use lagged log wage and log state-level average rent variable to control for local economic factors that could impact migration outcomes. I use Driscoll-Kraay standard errors to correct for heteroskedasticity and serial correlation in standard errors.

The first panel shows the positive association of state corporate tax rates with underfunding. We can see that state corporate tax rates are about 5-10 basis points higher over the time horizon of 1-6 years for a given 10 percent point higher underfunding. The second panel shows that marginal state income tax rates on the top-tax bracket is also higher by about 50 basis points in the four years after a 10 percent point higher underfunding. The third panel shows the impact on public amenities as public school capital expenditures decline by about 5-10 percentage points annually following a 10 percent point decline in funding ratio for over 6 years. Finally the fourth panel shows that the net migration rate measured by the aggregate adjusted gross income (AGI) reported by out-migrants net of in-migrants in the state divided by the total number of tax-payers in the states. The benefit of using this variable is that size-weights out-migrant returns of households and firms. The net migration rises by \$50-100 per-non-migrant taxpayer or equivalent to \$1-2 billion for the state of California, following a 10% point negative shock to the state's pension funding ratio. These correlations, in sum, paint the picture that pension underfunding is a market of state's fiscal health: persistent underfunding predicts higher taxes and worsening public services in the future, which might also be related to the net outflow of the state's tax-base. This poses a cyclical issue - whether states with worse funding status lose their respective tax base due to the issue itself or whether out-migration itself reduces the state's fiscal capacity. To address this cyclicality, I use an empirical design exploiting the difference in funding status between the origin and destination states in the next section.

# 5 Estimating the Pass-through of Funding Status on Migration

A central endogeneity concern in estimating the effect of public pension underfunding on household location decisions is the possible simultaneity of the two phenomena. States with higher underfunding may simply be more fiscally profligate, running persistent deficits. While higher taxes in such states may exert a "push" force on out-migration, these same states may also offer more generous welfare programs or superior public infrastructure, generating an offsetting "pull." Alternatively, optimistic discount rate assumptions may mask the true state of local fiscal health (Giesecke and Rauh, 2023). These concerns highlight the need for an empirical strategy that credibly isolates the impact of pension underfunding from confounders.

To address this, I leverage the IRS migration data, which report annual bilateral flows of tax returns, number of individuals, and adjusted gross income (AGI) between counties. This rich panel allows me to restrict attention to interstate moves and to implement a specification with both origin-time and destination-time fixed effects within each origin-destination pair, thereby controlling for time-varying shocks at both ends of the migration decision, in a two-stage least squares framework with instruments for pension funding status.

## 5.1 Econometric Framework: Funding Gaps and Migration Flows

I use the framework shown in Moretti and Wilson (2017) to conceptually motivate my empirical strategy. An individual's migration decision is motivated by the utility difference between the origin and destination. The indirect utility of an individual who lived in county $o$ in the previous year and moved to county $d$ (in a different state) at time t is as follows:

$$V_{iodt} = \alpha \ln(1 - \tau_{dt}) + \alpha \ln w_{dt} + \gamma \ln r_{dt} + B_d + e_{iodt} - C_{od}$$

Where $w_{dt}$ is the wage in current state, $r_{dt}$ is the rent or housing cost, $\alpha$ is the marginal utility of income, $B_d$ captures location specific amenities, $e_{iodt}$ represents location specific idiosyncratic preferences and $C_{od}$ is the utility cost of moving. Therefore, the utility gain or loss experienced by individual if she moves from state o to state d in year t is a function of the difference in taxes, wages, housing costs, and location-specific amenities between the origin and destination. This simple framework allows us to consider all possible pairs of origin and destination to model the relative bilateral utility gain from moving:

$$V_{iodt} - V_{ioot} = \alpha[\ln(1 - \tau_{dt}) - \ln(1 - \tau_{ot})] + \alpha \ln(w_{dt}/w_{ot}) + \gamma \ln(r_{dt}/r_{ot}) + B_d - B_o - C_{od} + e_{iodt} - e_{ioot}$$

Between all possible county pairs, I test the odds of out-migration from an origin county $o$ to a destination county $d$ in another state as a function of tax, wage, and rent gaps, conditioning on destination–origin fixed effects along with either origin–time or destination–time fixed effects. The innovation of this paper is to extend this framework by identifying migration responses to relative pension funding gaps.

Pension benefits are part of state and local governments' operating budgets, and higher promised liabilities translate into greater expenditures due to balanced budget requirements.[3] When pension assets fall short of covering recurring liabilities, states must finance the gap through their budgets. Empirically, as shown in Figure 4, this manifests in higher tax rates. Consistent with Novy-Marx and Rauh (2014), greater revenue demands accompany higher levels of underfunding. Through the budget constraint, tax rates can therefore be expressed as a decreasing function of the pension funding ratio. Thus, we can substitute $\ln(1 - \tau_{it}) = \ln(g(f_{it}))$, where $g'(f_{it}) > 0$. Appendix Figure 13 illustrates this relationship using dynamic 2SLS regressions of changes in relative funding status, instrumented by portfolio shocks. The results show that widening funding gaps predict widening tax gaps—worse funding status at the origin today implies higher state income tax rates tomorrow. Since $g(\cdot)$ is an increasing function, the qualitative impact of migration from $o$ to $d$ can be recovered by regressing on differences in funding ratios. This provides a reduced-form quantification of the odds of migration driven by pension underfunding. As $g(.)$ is an increasing function of its argument, we can recover the qualitative impact of movement from $o$ to $d$ by regressing the gap in the funding ratios. This can be useful in quantifying the reduced-form odds of migration due to higher levels of funding gap.

$$V_{iodt} - V_{ioot} = \beta[\ln f_{dt} - \ln f_{ot}] + \alpha \ln(w_{dt}/w_{ot}) + \gamma \ln(r_{dt}/r_{ot}) + B_d - B_o - C_{od} + e_{iodt} - e_{ioot}$$

If $e_{iodt}$ follows a Type-1 Extreme Value distribution, then the population shares sorting into migrant versus non-migrant groups would be equivalent to the log odds of migrant relative non-migrant individuals within the origin county at a given point of time. This gives us an econometric specification akin to a gravity equation in spatial economics, with the logit shares of the ratio of migrant and non-migrant returns being the dependent variable:

$$\ln(P_{odt}/P_{oot}) = \beta[\ln f_{dt} - \ln f_{ot}] + \alpha \ln(w_{dt}/w_{ot}) + \gamma \ln(r_{dt}/r_{ot}) + B_d - B_o - C_{od} + e_{iodt} - e_{ioot}$$

[3] https://taxpolicycenter.org/briefing-book/state-and-local-tax/fiscal-federalism-and/what-are-state-balanced

Let $P_{odt}$ denote the number of individuals migrating from state $o$ to state $d$, and $P_{oot}$ the number of non-migrants. Estimating migration directly on funding gaps raises several concerns. First, omitted variable bias may arise if time-varying factors beyond funding gaps, rents, wages, and fixed amenities also drive location choice. Second, measurement error in funding gaps—stemming from heterogeneous actuarial smoothing and liability assumptions—can attenuate estimates. Third, tax-base shrinkage may itself constrain revenue-raising capacity, further biasing OLS interpretations.

To address these concerns, I implement a two-stage least squares design exploiting exogenous variation in funding gaps. The main instrument is constructed from cumulative differences in excess return performance of origin and destination pension funds. This approach parallels Aiello et al. (2024), except that instead of benchmark returns, I use the weighted-average investment return assumption of each state's pension funds as of 2001. For robustness, I also employ an alternative instrument: the logged difference in union density—measured as the ratio of union members to total public employees in each state—similar to Mittal (2024). Results using this specification, reported in the appendix, remain consistent.

$$Z_{odt} = \sum_{h}^{t} [(r_{dh} - r^{a}_{d,2001}) - (r_{ot} - r^{a}_{o,2001})]$$

The exclusion restriction assumes that portfolio return shocks affect state pension balance sheets, which, through balanced-budget rules, translate into changes in taxes, services, and amenities—captured by the funding gap. Fixing actuarial assumptions to an ex-ante baseline ensures that projections of payroll and demographic growth, potentially influenced by migration, do not compromise instrument exogeneity. Conditional on destination–origin fixed effects and origin- or destination-time shocks, cumulative pension return shocks should influence migration only through their effect on relative funding gaps.

To implement the regression, I utilize IRS's county-to-county migration files between years 2000-2022. Each observation is at the origin-county-to-destination-county by year level. I use the public pension data to measure funding gaps. If the local funding status is available through the city or county-level pension fund, I apply that to the relevant county. If neither is available for a given county, I use the funding ratio of the county's state pension plan. I restrict it to county-pairs where the destination is located in a different state. I cluster standard errors by destination-origin county-pair to control for correlation of errors within county-pairs. The results of the regression are reported in table 1.

Panel A reports OLS regressions of log migration odds—measured as the ratio of out-migrant to non-migrant returns from county o to county d in another state—on the pension funding gap between destination and origin counties, controlling for relative rent and wage ratios. Column (1), without fixed effects, shows a positive correlation between funding gaps and migration odds. Adding origin and destination county fixed effects in Column (2) removes this effect. With destination–origin county-pair fixed effects in Column (3), the coefficient regains statistical significance: a one percentage point higher funding gap predicts 0.81 basis points higher odds of out-migration. Including origin-state × year FE (Column 4) shrinks and destabilizes the estimate, while destination-state × year (Column 5) and destination-region × origin-region × year FE (Column 6) flip the coefficient negative, suggesting that controlling for state- or region-level shocks alters the correlation. By contrast, Column (7), with origin-county × year FE, yields a positive and significant effect: a one percentage point higher gap predicts 4.2 basis points higher odds of out-migration, consistent with households moving toward better-funded regions. Column (8), with destination-county × year FE, similarly shows a positive and significant coefficient (1.2 basis points).

Panel B uses logged AGI outflows per non-migrant taxpayer as the dependent variable, capturing tax-base loss. Results are broadly consistent with Panel A, though Columns (4) and (5) reverse signs, indicating a different correlation when outcomes are income-weighted.

Panel C reports the first stage of the 2SLS, using the difference in cumulative excess returns of destination versus origin pension plans—measured relative to each plan's 2001 return assumption—as the instrument. Across all specifications, the instrument is positive and highly significant, with first stage F-stats»40, indicating that past relative investment performance strongly predicts current funding gaps. The second stage therefore captures the local average treatment effect of relative portfolio performance on migration through its impact on funding gaps. Panel D reports the second stage of the 2SLS with log migration odds as the dependent variable and the instrumented funding gap as the main regressor. Column (1), without fixed effects, yields a negative and significant coefficient, suggesting that investment performance gaps predict lower migration odds—but this estimate does not account for aggregate or location-specific factors, limiting its interpretation. With origin and destination county fixed effects (Column 2), the coefficient turns positive: a one percentage point higher funding gap is associated with a 0.41 percentage point increase in migration odds. Adding destination–origin county-pair fixed effects (Column 3) strengthens the effect, with a 10 percentage point higher funding gap predicting a 2 percentage point increase in migration odds.

Controlling for origin-state × year or origin-county × year fixed effects (Columns 4 and 7) renders the coefficients insignificant, indicating that once time-varying origin-specific "push" factors are absorbed, funding

Table 1: Migration Odds vs. Funding Gap

This table reports regressions of county-level migration odds on the pension funding gap between destination and origin counties. Panel A uses the log ratio of out-migrant to non-migrant returns as the dependent variable. Panel B replaces the dependent variable with the log ratio of out-migrant adjusted gross income (AGI) to non-migrant returns. Panel C presents the first-stage of the 2SLS, where the funding gap between destination and origin is instrumented with the difference in cumulative excess returns relative to the 2001 return assumption. Panel D reports the log odds of migration, and Panel E uses the log ratio of out-migrant AGI per non-migrant taxpayer with the instrumented funding gap as the main regressor. Column (1) includes no fixed effects; Column (2) adds origin and destination fixed effects; Columns (3)–(8) include destination × origin county fixed effects. Column (4) further adds origin-state × year FE, Column (5) destination-state × year FE, Column (6) destination-region × origin-region × year FE (using the nine Census regions), Column (7) origin-county × year FE, and Column (8) destination-county × year FE. All specifications control for the log rent and log wage ratios (destination/origin). Standard errors are clustered by destination-origin pair

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
|---|---|---|---|---|---|---|---|---|
| *Panel A: OLS, Dep. Var:* | Ln(Out Migrant Returns/Non-migrant Returns) | | | | | | | |
| Funding Gap | 0.190*** | -0.00194 | 0.00817** | 0.00424 | -0.0156*** | -0.0466*** | 0.0422*** | 0.0118** |
| | (10.14) | (-0.40) | (2.07) | (1.07) | (-3.96) | (-11.22) | (6.62) | (2.04) |
| *Panel B: OLS, Dep. Var:* | Ln(Out Migrant AGI/Non-migrant Returns) | | | | | | | |
| Funding Gap | 0.277*** | 0.00495 | 0.0136** | -0.00362 | 0.0154*** | -0.0405*** | 0.0570*** | 0.0213** |
| | (15.09) | (0.70) | (2.11) | (-0.63) | (2.67) | (-6.65) | (6.31) | (2.53) |
| *Panel C: IV - 1st St., Dep. Var:* | Funding Gap | | | | | | | |
| $Z_r$ | 0.0898*** | 0.0692*** | 0.0897*** | 0.0770*** | 0.0887*** | 0.0775*** | 0.0771*** | 0.0888*** |
| | (21.64) | (41.16) | (39.13) | (56.29) | (55.16) | (36.34) | (54.80) | (53.54) |
| *Panel D: IV - 2nd St., Dep. Var:* | Ln(Out Migrant Returns/Non-migrant Returns) | | | | | | | |
| Funding Gap | -0.310** | 0.408*** | 0.195*** | 0.0322 | 0.303*** | 0.227*** | 0.0324 | 0.354*** |
| | (-2.12) | (7.77) | (6.56) | (0.61) | (6.95) | (6.90) | (0.62) | (8.44) |
| *Panel E: IV - 2nd St., Dep. Var:* | Ln(Out Migrant AGI/Non-migrant Returns) | | | | | | | |
| Funding Gap | -0.372** | 0.568*** | 0.453*** | -0.105 | 0.372*** | 0.146*** | -0.0795 | 0.412*** |
| | (-2.55) | (8.18) | (9.21) | (-1.41) | (6.08) | (3.26) | (-1.11) | (6.96) |
| N | 840217 | 839892 | 818137 | 818137 | 818137 | 818136 | 809319 | 809353 |
| Destination, Origin County-FE | | Y | | | | | | |
| Destination × Origin County-FE | | | Y | Y | Y | Y | Y | Y |
| Origin State × Year-FE | | | | Y | | | | |
| Destination State × Year | | | | | Y | | | |
| Dest. Reg. × Org. Reg. × Yr-FE | | | | | | Y | | |
| Origin County × Year-FE | | | | | | | Y | |
| Destination County × Year-FE | | | | | | | | Y |
| Controls: | Ln(Destination Rent/Origin Rent), Ln(Destination Real Wage/Origin Real Wage) | | | | | | | |

$t$ statistics in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

gaps no longer explain migration within county-pairs. This suggests that underfunding operates through intermediate channels (e.g., taxes, services, amenities), and absorbing those factors removes the direct effect of the funding gap. In contrast, controlling for destination–origin region × year fixed effects (Column 6) restores a positive and significant effect, showing robustness to regional business cycles. Finally, when controlling for destination-state × year or destination-county × year fixed effects (Columns 5 and 8), migration odds are 3–3.5 percentage points higher when origin counties are 10 percentage points more underfunded.

Panel E repeats the analysis using log AGI outflows per non-migrant taxpayer. With origin and destination county fixed effects, a one percentage point higher funding gap predicts a 0.57 percentage point increase in AGI outflows, while within county-pairs (Column 3) the effect is 0.45 percentage points. As in Panel D, origin × year controls (Columns 4 and 7) yield null results, but destination × year and region-pair × year controls (Columns 5–6, 8) show that a one percentage point higher funding gap predicts 0.37–0.41 percentage point higher AGI outflows per taxpayer. This means for Cook county with 2 million non-migrant tax returns in 2022, a higher than baseline outflow of $3.3 million to a destination county with 20% (1 SD) higher funding ratio.

Overall, the results suggest that migration responds to pension underfunding when relative funding differences across origin and destination counties are salient, but the effect attenuates once origin-specific push factors are absorbed. These findings are robust to (i) using the public-sector unionization instrument from Mittal (2024), (ii) truncating the sample in 2019 to exclude pandemic migration years, and (iii) dropping 2014–2016 to address IRS data measurement concerns noted in DeWaard et al. (2022). In the next section, I explore *net* migration from a county.

## 5.2 Measuring Net Migration

Bilateral migration flows provide a lens to study how funding gaps affect migration while leveraging a rich set of fixed effects to absorb confounding variation. I now focus on the reduced-form effect of deteriorating county funding ratios on *net* out-migration (outflows minus inflows). Using IRS county-level migration data, I subtract county-wise inflows from outflows to construct the main outcome variables as the ratios of net outflows of tax returns and adjusted gross income normalized by the number of non-migrant returns in a county in a given year.

To account for local conditions, I include lagged controls for real wages, rents, county GDP growth, state debt-to-revenue, property and state income taxes, property crime incidence, and the political leaning of the state (measured by the lagged share of Republican congressional representatives). This specification tests the reduced-form effect of underfunding on migration while conditioning on key economic, fiscal, and political factors.

Residual confounding variation from state- and county-specific time-varying shocks, as well as potential measurement error, is addressed by instrumenting funding status with cumulative excess returns on state and local defined benefit pension assets relative to their 2001 return assumptions.

$$\tilde{Z}_r = \sum_{h=0}^{t} (r_h - r^a_{2001})$$

The exclusion restriction rests on portfolio shocks only affecting migration through their effect on pension balance sheets, conditional on controlling for county and time fixed effects along with time-varying county and state parameters. The results are presented in table 2.

Column (1) shows a positive and statistically significant relationship between funding ratios and cumulative excess returns relative to 2001 assumptions, confirming instrument relevance. Columns (2)–(3) use net outflows of returns relative to non-migrant returns as the dependent variable, while Columns (4)–(5) use net AGI outflows per non-migrant return. The OLS results in Columns (2) and (4) show no effect. By contrast, the 2SLS estimate in Column (3) indicates that a one–standard deviation decline in funding (14%) increases net outflows of returns by 29 basis points. Column (5) implies that the same decline raises AGI outflows by $265.44 per non-migrant return—equivalent to a $540 million loss for Cook County, which had two million non-migrant returns in 2022.

These results are robust to (i) using the unionization instrument from Mittal (2024), (ii) restricting the sample till 2019 to exclude pandemic-driven migration, and (iii) excluding 2014–2016 to address IRS data concerns raised by DeWaard et al. (2022), and reported in the appendix. I also perform a leave-one-region-out robustness test where I drop data of one of the nine US census bureau regions at a time, to test for whether particular regions drive the results. The results are qualitatively similar and are reported in the appendix. Overall, the evidence suggests that underfunding signals weaker state fiscal health, with the resulting tax and amenity pressures prompting households to relocate from underfunded to better-funded regions.

The next section examines the timing of this channel: does pension underfunding affect household migration immediately, or with a lag?

Table 2: County-wise Net Migration

This table reports regressions of county-level net out-migration on defined benefit pension funding ratios, controlling for local time-varying factors and county and year fixed effects. Column (1) presents the first stage, instrumenting the funding ratio with cumulative excess returns relative to the 2001 investment return assumption $Z_r$ is the cumulative excess return relative to the 2001 investment return assumption. Columns (2) and (3) show the regressions with the net outflow of tax returns relative to non-migrant tax returns as the dependent variable. Columns (4) and (5) shows the net outflow of adjusted gross income relative to non-migrant returns. Columns (2) and (4) show the OLS version of the regressions whereas columns (3) and (5) show the second stage with funding ratio of the county instrumented with the return instrument. Each specification contains the following controls: lagged log rents, real wages, property tax-to-household income ratio, top marginal state income tax rates, state debt to revenue, county-GDP growth and percentage of state's House of Representatives congressmen that belong to the Republican party. Standard errors are clustered at the county-level.

| | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| | Funding Ratio | $\frac{\text{Net Out Migrant Returns}}{\text{Non-Migrant Returns}}$ | | $\frac{\text{Net Out Migrant AGI}}{\text{Non-Migrant Returns}}$ | |
| | First-Stage | OLS | IV | OLS | IV |
| $\tilde{Z}_r$ | 0.0607*** | | | | |
| | (10.22) | | | | |
| Funding Ratio$_{t-1}$ | | -0.000468 | -0.0205** | 0.105 | -1.896** |
| | | (-0.39) | (-2.07) | (1.31) | (-2.45) |
| N | 59403 | 56312 | 56312 | 56312 | 56312 |
| R-Sq | 0.812 | 0.269 | - | 0.425 | - |
| F-Stat | 208.3 | | | | |
| FE | County and Year | | | | |
| Controls | Rents, Wages, S&L Taxes, Debt/Rev, County-growth, Crime, Political | | | | |

$t$ statistics in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

## 5.3 Dynamic Migration response to Pension Funding Gaps

The previous section documents how funding gaps and underfunding predict migration response from the migrants' origin states. In this section, I explore what is the time lag of response in migration when funding gap between two states increases. I use the empirical strategy used in Moretti and Wilson (2017), by using an indicator for the change in the funding gap between the destination and origin counties. The funding gap measure remains the same as in the previous section, but now I regress the *change* in log odds of migration on the change in funding gap indicator over different horizons. I create an indicator variable $D_{odh}$ which is equal to 1 when the funding gap increases, -1 if the funding gap decreases and 0 if the funding gap stays constant. I then estimate the following specification:

$$\ln\left(\frac{Y_{od,t+h}}{P_{oo,t+h}}\right) - \ln\left(\frac{Y_{od,t}}{P_{oo,t}}\right) = \eta_h D_{odh} + \gamma X_{odt} + F_{odt} + \epsilon_{od,t+h}$$

Where $\ln\left(\frac{Y_{od,t+h}}{P_{oo,t+h}}\right)$ is the logged ratio of returns (or AGI) that migrate from county $o$ to county $d$ in year $t+h$ and non-migrant returns that stay in county $o$ in period $t+h$. $X_{odt}$ contains the rent and wage ratio of the origin counties at the time of the change in funding gap. $F_{odt}$ is the region of origin $\times$ region of destination $\times$ year fixed effects, where a region is described as per the 9 US Census Bureau division combining a cluster of states.

As migration and pension funding gap could co-move due to other confounding local and bilateral variation, I also run a two-stage least squares version of the specification where I instrument $D_{odh}$ with the following instrument:

$$Z'_{odt} = (r_{dt} - r^a_{d,2001}) - (r_{ot} - r^a_{o,2001})$$

Where $r_{it}$ is the realized one-year rate of return on pension assets and $r^a_{i,2001}$ is the investment return assumption by the pension actuaries for $i \in \{o, d\}$ as of 2001. The instrument essentially is the difference in return "surprises" between the destination and origin county-pension funds relative to an ex-ante investment return assumption. The exclusion restriction is that the difference in return surprises only affects migration through its impact on pension funding gap between the two counties and is exogenous to factors driving migration, conditional from region-pair-year-wise variation and destination-county rent and real wage ratio. The first stage F-stat >> 40 for each horizon. Standard Errors in both the OLS and 2-SLS specifications are clustered by origin $\times$ destination county pair to control for within county-pair correlation of errors. The results are presented in figure 5. In all specifications I find that log odds of out-migration consistently increases when the funding gap between two counties increases. To interpret magnitudes, in the OLS panel (a) I find that the when destination counties become relatively better funded than origin counties, the odds of out-migration rise by 1% in 5 years and 2% over 10 years. The two stage results in Panel (b) show that change in out-migration odds rise to 6% by the sixth

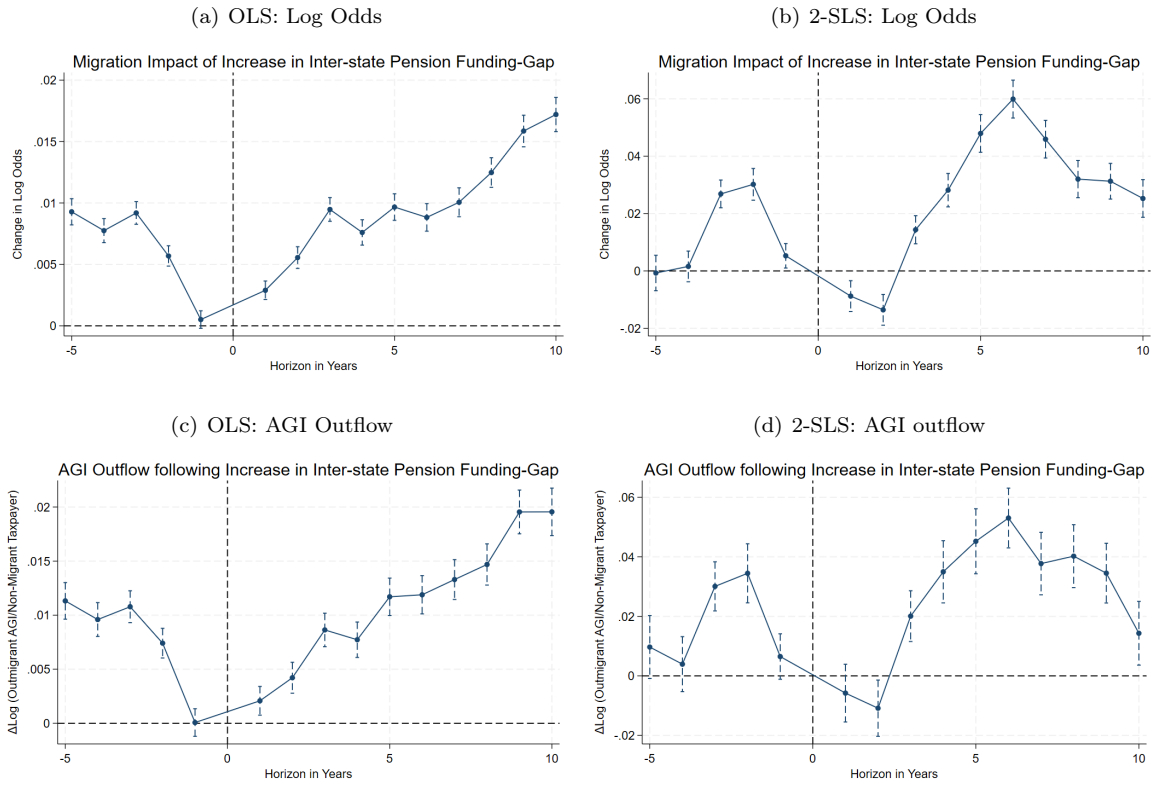



Figure 5: Migration Response to Inter-state Funding Gaps

year but then fall to about 3% by the tenth year. Panels (c) and (d) also capture similar dynamic responses of measured in terms of AGI outflow per non-migrant tax-payers. Thus, in sum: (i) defined benefit public pension funding gaps between two states predict migration flows within the states controlling for various destination and origin specific characteristics, (ii) higher underfunding leads to a loss in the taxable adjusted gross income in states with higher underfunding, and (iii) migration response to increases in funding gaps peaks around 6 years of the the increase in funding gap between the two states - signaling that worse funding could create a fiscal-migratory spiral for states with worse relative funding status. In the next section I investigate how this migratory feedback might affect public plans.

# 6 Migration Risk and Public Pension Risk-taking

How do state governments and public pension respond to higher underfunding? Giesecke and Rauh (2023) document the pervasive lack of sufficient contribution rates due to differences in discounting and other political factors. Combined with political incentives to risk-shift through raising assumed discount rates on pension plans (sww Andonov et al. (2017)), states governments can trade-off making lower contributions with riskier investment strategies. Further, Reach for yield by public pensions has been extensively studied in extant literature such as Lu et al. (2019), Myers (2023), Giesecke and Rauh (2023) and Mittal (2024).

Fully funding obligations would require raising contributions, potentially through higher taxes. This path is complicated by the mobility of the tax base: if households and firms are sensitive to tax changes, raising contributions may prompt out-migration. If states fear raising taxes could drive firms and households to relocate, they may likely reach for yield on their pension investments. This paper adds to that literature by examining how reach for yield behavior varies across states with heterogeneous migration responses. Specifically, it asks: do states invest differently if their local populace are more sensitive to tax changes than other states? I test that by first estimating “migration" risk for each state - what is the net-migration impact of increases in state income taxes. I subsequently explore the heterogeneity of portfolio allocation in response to differential migration risk.

## 6.1 State-wise Migration Risk Measurement

How sensitive a local population is to moving when faced with higher tax burden or deteriorating public services depends on several factors. Work in migration across states model and document various agglomeration and congestion forces that both draw as well as repel people from locating in a particular region (see Foschi et al.

(2025)). Estimating state-wise migration betas might thus entail controlling for confounding variation that might drive migration other than taxes. Previous work such as Rauh and Shyu (2024), use rich administrative data to estimate the behavioural response of top income earners to increases in effective tax rates. I come up with three approaches to estimate migration responses to taxes by state. The objective of each measure is to create a *relative* sensitivity to changes in tax rates which helps me rank states by their respective out-migration risk.

### 6.1.1 OLS Migration Betas

Akin to deposit betas which measure changes in bank deposits when the Federal Funds Rate changes estimated by Drechsler et al. (2017), I estimate regressions of net migration of individuals from IRS migration data on increase in the statutory tax rates for each state separately, controlling for local real wages, rent, house price index, employment growth, job openings, state GDP growth, crime, state debt to revenue ratio, pupil-teacher ratio and infrastructure quality. The rationale behind putting in the rich set of controls is to proxy for local economic performance and outlook, cost of living, and measures of public amenities that might drive migration response. I measure an increase in marginal tax rate as a positive change in marginal tax rates in any of the state income tax brackets.

For each state, I run the regression:

$$Y_{st+1} = \beta_s \times 1(\Delta\text{Marginal State Income Tax Rate}{>}0)_t + \Gamma X_{st} + \varepsilon_{st}$$

Where $Y_{st+1}$ is the net out-migrant returns relative to non-migrant returns, $\beta_s$ parameter captures the relationship between tax increases and migration after controlling for confounding determinants of migration, for state $s$. I plot the OLS "migration" betas in an ascending order in panel figure 7.

The benefit of using this approach is that it shows the relationship between migration responses to any increase in the state marginal tax rate controlling for other possible drivers of migration by state. The drawback is that this is equivalent to running a time-series regression for each state, which might result in issues of stationarity and, of course, endogeneity. Also several states have consistently zero state income taxes which means dropping them from the estimation. The beta estimates are shown in figure 7 panel (a).

### 6.1.2 Moretti and Wilson (2017) Approach for Each State

Moretti and Wilson (2017) study the impact of state marginal tax rates on the location of star scientists in the United States. They exploit the gap between the tax rates of the destination and origin states to document the migration elasticity to tax changes in the origin or destination states. Just as before, I use IRS county-to-county migration files, restricting to out-of-state migration, and employ the following specification, running it for each state:

$$\ln\left(\frac{P_{od,t}}{P_{oo,t}}\right) = \eta_o(\ln(1-\tau_{d,t}) - \ln(1-\tau_{o,t})) + \eta_{o,c}(\ln(1-\tau^c_{d,t}) - \ln(1-\tau^c_{o,t})) + \gamma_o + \gamma_d + \gamma_{od} + \epsilon_{od,t}$$

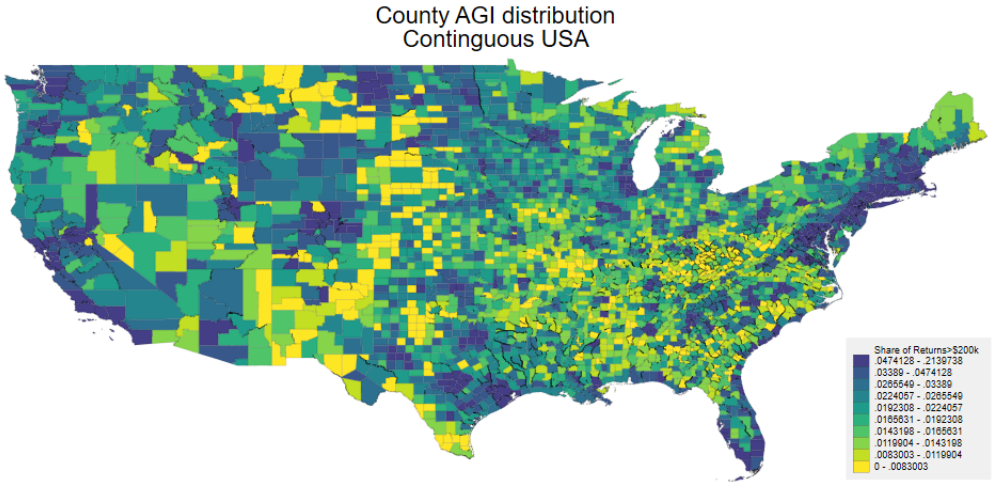
Where $P_{od,t}$ is the number of individuals that migrate from county $o$ to county $d$ (in a different state), while $P_{oo,t}$ is the number of individuals that do not migrate. The right-hand side variables are the logged tax gap, where $\tau_{d,t}$ is the destination top-marginal income tax rate and $\tau_{o,t}$ is the origin counterpart. $\tau^c_{it}$ is the state corporate tax rate of state of county $i$. The specification controls for state of origin, state of destination, and origin-destination pair fixed effects. Standard errors are clustered by destination $\times$ origin pairs. The specification draws from the logit-share based location choice models standard in quantitative spatial models that yield migration elasticity to taxes as $\frac{\partial \ln N_{s,t}}{\partial \ln(1-\tau_{s,t})}$. This approach helps document migration responses of both households and firms in response to location-based tax rates, and Moretti and Wilson (2017) show that the resulting coefficient on the state income tax gap is equivalent to the equilibrium migration elasticity. I run the specification for each state and recover the coefficient on the income tax gap, which I plot in panel (b) of figure 7. The benefit of this approach is that it takes into account the structural parameters driving migration responses to tax gaps between two states and effectively controls for state of origin, destination and destination $\times$ origin fixed characteristics. The issue however is that it still depends on the difference in levels of taxes, and does not necessarily evade confounding factors such as destination-origin-time varying trends that might affect location choices of households and firms.

### 6.1.3 Post-SALT Cap Net-outflows

The third version is to exploit the disproportionate exposure of states with high state and local taxes to the introduction of the SALT deduction cap introduced under the Tax Cuts and Jobs Act of 2017. This policy change limited federal deductibility of state and local taxes upto $10,000, raising effective state tax burden,

especially in states with ex-ante higher tax rates. Figure 7 panel (c) shows the correlation between the increase in federal state and tax liability post-SALT against their top marginal tax rates, estimated using NBER's TAXSIM software. Figure 6 panel (a) shows the spatial distribution of top income tax bracket returns across contiguous United States. This shows a high concentration of top income shares in urban counties of California, Illinois, New England, Texas and Florida. Contrast this with panel (b) which shows the net AGI outflow per tax payer after the introduction of SALT deduction cap - while California, Illinois, and New England show net outflows, whereas Florida and Texas show net-in migration - indicating that migrants moved, from high tax regions to low tax regions (consistent with Rauh (2022)). High tax states see a larger impact on both federal and state tax liabilities of the highest single-filer tax bracket. I then regress net out-migrant individual returns to total returns in the state ratio against the change in tax liability controlling for state-time varying controls and year-fixed effects. I then predict the net cumulative outflow by state predicted by the estimation over the time period of 2018-2019. I stop at 2019 to exclude migration driven by the pandemic.

(a) Top AGI Distribution in 2017




(b) Average Annual Net Outflow

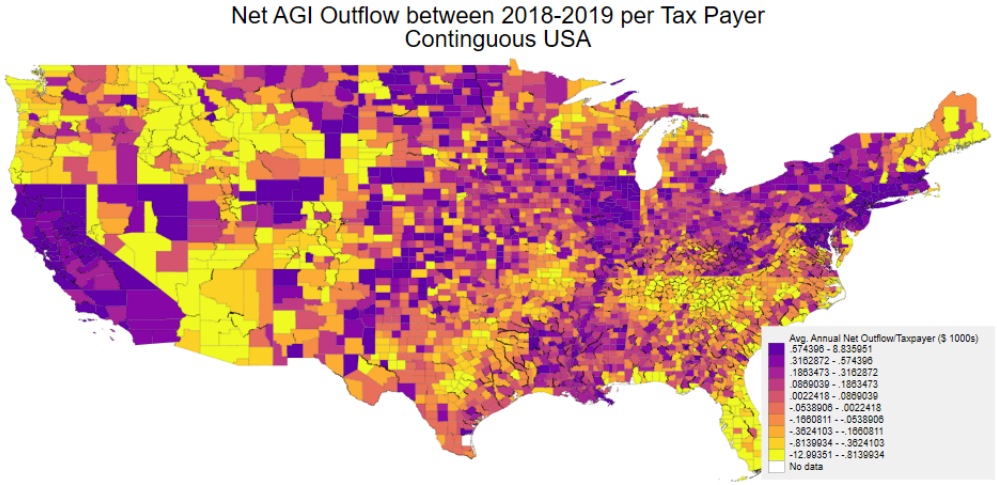



Figure 6: Spatial Distribution of Top-AGI shares and Out-migration

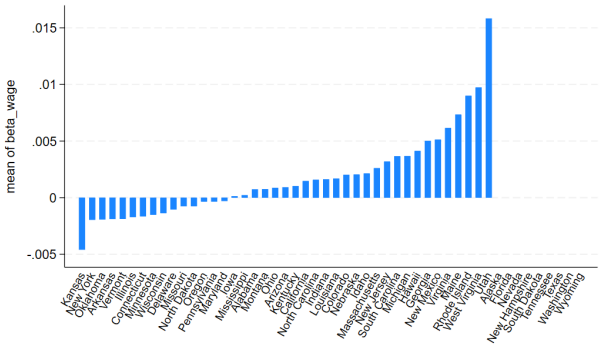
Panel (a) plots the spatial distribution of high AGI tax returns, where each county is colored according to the share of tax returns $>$ \$200k in 2017. Dark blue regions show highest shares and yellow regions are the lowest. Panel (b) plots the average annual net migration measured as the Net AGI outflow per tax payer for the period of 2018-2022. Dark purple region have higher out-migration whereas yellow show net in-migration. Data is from IRS Statistics of Income.

$$Y_{s,t} = \beta \times \Delta\text{Federal and State Tax Liability}_s + \Gamma X_{s,t-1} + \tau_t + \epsilon_{s,t}$$

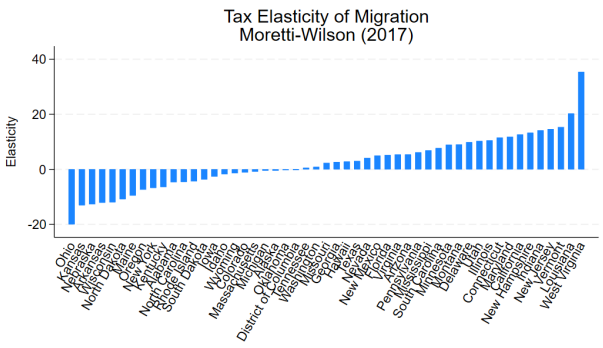
Controls include state-level estimates of average rents, real wages, crime, infrastructure quality, public school quality, state debt-to-revenue and time-fixed effects. Standard errors are clustered by state. The regression shows a positive relationship between tax liability increases and net-outflow rate in appendix table 10. The predicted outflows are displayed in figure 7 panel (d). The benefit of using this measure is that the tax shock is plausibly exogenous to the states. The possible cost is that due to the nature of the tax shock, treatment is proportional to the ex-ante levels of taxes.

The three methods gives us the state-wise relative migration risk measures in response to changes in state income tax rates. All three migration risk measures are positively correlated with each other, with the correlation coefficient between OLS migration beta and Moretti and Wilson (2017) style state-wise elasticity (henceforth MW-2017 elasticity) estimate being 0.26, correlation between the MW-2017 and post-SALT cap predicted outflows being 0.33 and the correlation between the OLS migration beta and post-SALT cap outflows being 0.04.

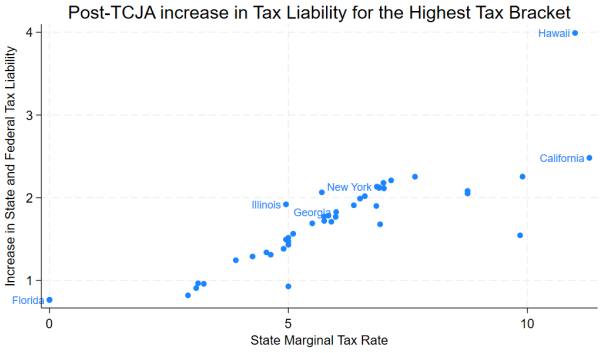
(a) OLS Migration Betas




(b) Moretti-Wilson (2017) Migration Elasticity




(c) Post-SALT cap correlation with State Tax Rates




(d) Post-SALT cap Net Outflows

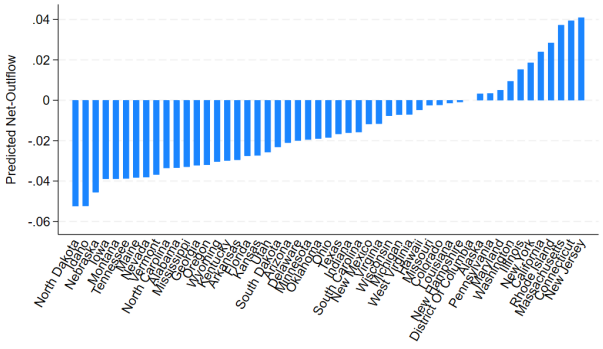



Figure 7: Migration Betas

### 6.1.4 Mapping Migration Risk to Public Pension Portfolio choice

What do migration betas mean for the pension fiduciary? Migration betas are a reflection of the state-wise ability to raise revenue through raising taxes. States which are more likely to lose tax base when they raise taxes, are less likely to fulfill future pension obligations given their unfunded liabilities. To test this, I plot how states with higher out-migration risk (i) contribute to their pension plans relative to the actuarially estimated required contributions (ii) maintain their funding status relative to low-migration risk states. Figure 8 panel (a) shows the value-weighted average percentage of required contributions paid by states classified as above median migration beta estimated as per the Moretti and Wilson (2017) approach (thus "High Migration Beta States") versus the below median states. We can see that there is a considerable gap - high migration beta states under-contribute to their public pensions relative to low migration beta states. In panel (b) I plot their funding status, which also turns out to be lower for the high migration beta states. Therefore, migration betas sort states by their systemic pension under-contribution and subsequent underfunding. This possibly results in fiduciaries taking higher risks on their pension portfolios, which I test next.

I track the allocation patterns of public pension funds with respect to migration risk using pension plan level data. To test heterogeneity I estimate this specification which involves using the interaction between high migration risk and Underfunding ratio (which is simply 1-Funding Ratio) with the following specification:

$$Y_{it+1} = \beta \times 1(\text{High Migration Risk})_s \times \text{Underfunding}_{i,t} + \text{Underfunding}_{i,t} + \Gamma X_{it} + \theta_t + \kappa_i + \varepsilon_{it+1}$$

Where $Y_{it+1}$ is the portfolio allocation between 0 and 1 for a given asset class for pension fund $i$ at time $t+1$. *High Migration Risk* refers to the dummy of whether state $s$ has above median migration response to

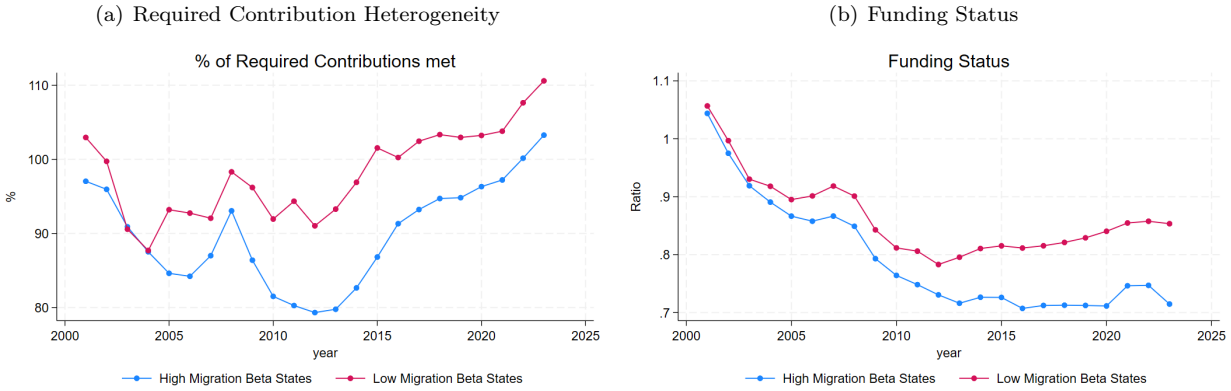



Figure 8: How do migration betas affect public pension fiduciaries?

This figure represents the weighted trends in contribution rates and funding status for states which lie above median migration risk - measured by their Moretti and Wilson (2017)-style migration response to tax changes. Panel (a) represents the plan contributions relative to the actuarially-required level for high-migration risk and low-migration risk states. Panel (b) shows their respective asset-weighted funding ratios.

changes in local tax-liability. Controls include lagged contribution rate growth, and changes in investment return assumption, discount rates, valuation methods or other assumption changes. All specifications include plan and year fixed effects. The coefficient of interest $\beta$ is the relative difference between the allocation to an asset class, between pension funds belonging to high and low migration risk states, for a given level underfunding. Standard errors are clustered at the state $\times$ year level to correct for within state correlation of allocations [4]

The results are displayed in table 3. Columns (1)-(3) track allocations to alternative asset classes such as private equity, real estate, hedge funds, commodities and other miscellaneous asset classes. Columns (4)-(6) track allocations to fixed income securities and (7)-(9) track equity allocations. In Columns (1),(4) and (7), the migration risk measure is the indicator for whether the Moretti and Wilson (2017) migration estimates of state $s$ lies above median. Columns (2),(5) and (8) use the cumulative predicted net-outflows after the tax liability increase due to the SALT deduction cap in 2017. Columns (3), (6) and (9) use the OLS regression betas as the migration response measure to sort states by high versus low risk. What emerges consistently across different measures of migration risk is the relative increase in alternative allocations for high migration risk states. Column (1) predicts a 75 basis point increase in alternative allocations in response to a one percentage point increase in underfunded ratio in states which face a higher than median out-migration response relative to baseline. In columns (2) and (3) that increase is equivalent to 100 basis points and 141 basis points respectively.

Conversely Columns (4), (5), (7) and (8) show that the relative allocation in fixed income or equity allocations are not different from zero for high migration risk states. However columns (6) and (9) show that both equity and fixed income allocations are negative and statistically significant for states with a higher OLS tax beta, suggesting re-balancing towards higher expected return assets.

To measure the dynamic response of migration risk and underfunding, I replicate the specification in a local projections framework. This exercise helps map out allocation patterns over time and how does migration risk contribute to the changes we see over time. I use the following specification for testing this:

$$Y_{it+h} = \beta \times 1(\text{High Migration Risk }) \times \text{Underfunding}_{i,t} + \text{Underfunding}_{i,t} + \Gamma X_{it} + \theta_t + \kappa_i + \varepsilon_{it+h}$$

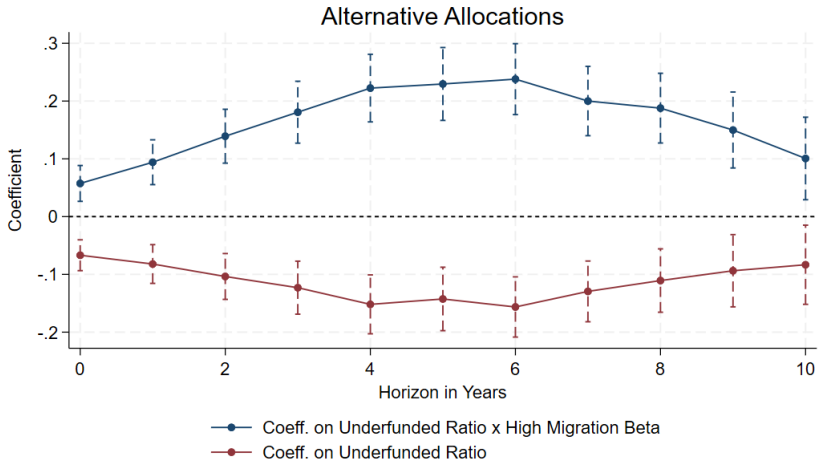
Over horizons $h \in (0, 10)$. $Y_{it}$ is the allocation to asset classes such as equities, fixed income and alternative asset classes. Figure 15 shows the results for three asset classes, with the high migration risk measure being $1(AM\ OLS\ Migration\ Beta)$[5]. The first panel shows that states with high migration risk see the pivot towards alternative allocations, relative to states below the median in migration betas. The allocation goes up to about 200 basis points for a percentage increase in underfunded ratio whereas regions with lower migration risk reduce allocation by 100 basis points. Panel (b) plots the allocation to equities declines by 12 percentage points for funds in high-migration-risk states, on the other hand allocation by lower-migration-risk states in equities rises by 5-10 percentage points. Panel (c) plots the regression coefficients for allocation to fixed-income securities. The coefficient for high migration beta states is consistently negative but statistically not different from zero. Conversely allocation by low migration beta states have a consistently positive, but statistically insignificant coefficient across the horizons. The persistent allocation reflects buy-and-hold behavior which suggests that high-migration risk states are more likely to hold more alternatives through time.

I have therefore shown that (i) pension underfunding correlates with higher tax liabilities and worsening public expenditures (ii) pension funding gap between regions and underfunding by itself thus predicts net-out-

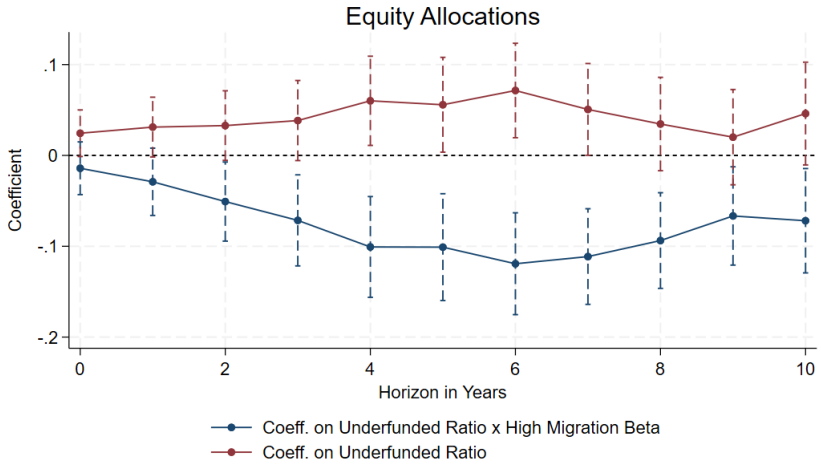
[4] A version with Driscoll-Kraay standard errors is presented in the appendix.

[5] Other migration risk measures are shown in the appendix

(a) Alternative Allocations




(b) Allocation to Equities




(c) Allocation to Fixed Income

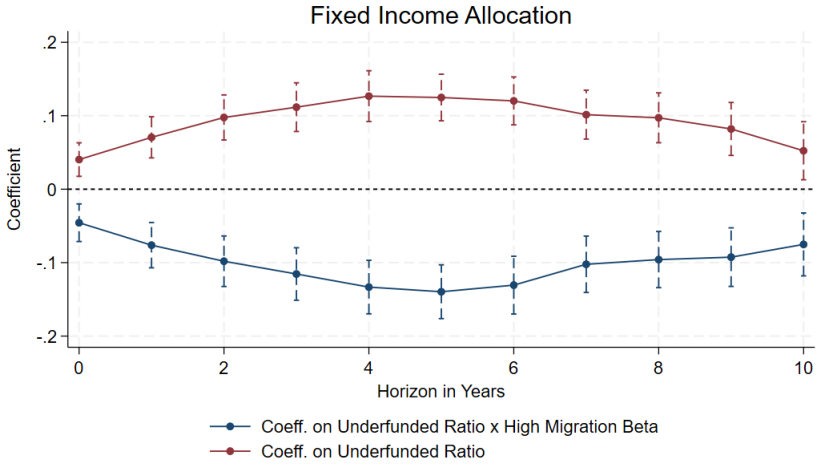



Figure 9: Local Projections: Allocation by Funding Status and Migration Risk

These graphs plot the regression coefficients of the public pension allocations on the interaction of state migration betas and underfunded ratio. The navy-blue line represents the coefficient on the interaction of migration beta and underfunding, while the maroon line represents the coefficient on underfunded ratio. Panel (a) plots the coefficients for alternative allocations (b) plots the coefficients for allocations to equities and (c) plots the coefficient on fixed income securities. The axes represents the leads of the dependent variable in years.

Table 3: Allocation Heterogeneity by Migration Risk

This table presents results from the regressions of public pension portfolio allocations on under-funded ratio (1-Funding Ratio) interacted with the three different measures of migration risk. Columns (1), (2) and (3) use alternative allocations - which is the sum of allocations to private equity, real estate, hedge funds, commodities and other miscellaneous alternative asset classes - as the dependent variable. Columns (4), (5) and (6) use fixed income allocation and columns (7), (8), and (9) use equity allocation as the dependent variable respectively. Columns (1), (4) and (7) show the interaction of the underfunded ratio with above-median OLS migration beta as the main indicator for migration risk. Columns (2), (5) and (8) use the indicator for states with above median Moretti and Wilson (2017) style tax elasticities of migration. Columns (3), (6) and (9) use the dummy for above median post-SALT cap net-outflows in response to tax increases as the high migration risk indicator. All specifications control for pensions contribution growth, changes in actuarial assumptions and plan and year fixed effects. Standard errors are clustered by State × Year pair.

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
|---|---|---|---|---|---|---|---|---|---|
| | Alternative Allocations | | | Fixed Income Allocation | | | Equity Allocations | | |
| Underfunded Ratio (UR) | -0.0654*** | -0.0916*** | -0.113*** | -0.0127 | -0.0257 | 0.0584*** | -0.0615* | -0.0313 | 0.0210 |
| | (0.0234) | (0.0314) | (0.0253) | (0.0223) | (0.0296) | (0.0222) | (0.0356) | (0.0469) | (0.0390) |
| 1(AM Beta Moretti-Wilson (2017)) × UR | 0.0749** | | | 0.0355 | | | 0.0601 | | |
| | (0.0301) | | | (0.0311) | | | (0.0516) | | |
| 1(AM Post SALT-Outmig) × UR | | 0.0996*** | | | 0.0480 | | | 0.000112 | |
| | | (0.0338) | | | (0.0328) | | | (0.0521) | |
| 1(AM OLS Migration Beta)=1 × UR | | | 0.141*** | | | -0.0883*** | | | -0.0866* |
| | | | (0.0295) | | | (0.0303) | | | (0.0516) |
| N | 4251 | 4251 | 4251 | 4251 | 4251 | 4251 | 4251 | 4251 | 4251 |
| R-Sq | 0.728 | 0.728 | 0.730 | 0.577 | 0.577 | 0.578 | 0.602 | 0.601 | 0.602 |
| Pension Plan and Year-FE | Y | Y | Y | Y | Y | Y | Y | Y | Y |

Standard errors in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

migration from worse funded states (iii) states with populations more sensitive to tax changes are more likely to pivot to alternative asset classes[6].

What role does tax migration play in portfolio choice of public pension fiduciaries? What is the channel through which tax mobility induces the trade-off between raising contributions versus reaching for yield on pension assets among state and local governments? What is the welfare maximizing investment policy if a planner that sets both tax rates and acts as fiduciary for the defined benefit pension plans, internalizing the migration impact? To try to rationalize the findings and get address possible mechanisms, I setup a spatial model in the next section.

# 7 Model: Linking Population Mobility to Public Pension Risk-taking

The previous sections show evidence of a possible trade-off that state and local governments face when addressing their defined benefit pensions: raising contributions through taxes or gambling for resurrection by investing in high-yield assets. To conceptualize the trade-off, I develop a two-period spatial model to test the public pension problem across states. The objectives of this model are (i) to formalize households' location as a function of local taxes and amenities subject to employment and housing market (ii) to study the trade-off faced by state governments between raising taxes and investing in risky securities. To start, there is a continuum of households indexed by i ∈ [0,1] born in states $s \in (1, ..., S)$. Each household draws utility from a homogeneous consumption good and housing, and can choose to locate in any of the states indexed by $s \in (1, ..., S)$ in period two based on the state characteristics in period 1 which gives them the highest indirect utility. I detail household preferences as follows:

## 7.1 Household Preferences

Households draw utility from consumption c and housing h. They supply labor in-elastically at the wage $w_s$

$$U_{is}(c, h, A_s) = B_s \left(\frac{c_{s,t}}{\alpha}\right)^{\alpha} \left(\frac{h_{s,t}}{1-\alpha}\right)^{1-\alpha} + \xi_{ist} \tag{2}$$

subject to:

$$r_{s,t}h_{s,t} + c_{st} = w_{s,t}(1 - \tau_{st}) \tag{3}$$

Where $r_s$ is state-specific rent rent, $h$ is housing consumption and c is a homogeneous consumption good. The price of the homogeneous good is normalized to one. $B_s$ represents location specific amenities. $\xi_{is}$ is

[6]In the appendix section A.2 I test whether pension managers in high migration risk states taken on more risk on their equity portfolios using 13-F data. I do not find evidence of higher portfolio volatility among managers from high-migration risk states. Therefore reach for yield behavior possibly manifests across asset class allocation.

idiosyncratic location preferences distributed Gumbel with dispersion $\sigma_v$ Households' first order conditions give us the consumption and housing demands. I plug them in to derive the location specific payoff for individual $i$:

$$V_{ist} = \underbrace{B_s w_{s,t}(1-\tau_{st}) r_{st}^{\alpha-1}}_{v_s\text{- Location Specifc Payoff}} + \overbrace{\xi_{ist}}^{\text{Idiosyncratic Location Preferences}} \tag{4}$$

Thus individuals' location specific utility go up in terms of local amenities, state-specific income net of taxes and is negatively related to the amount of endogenous rents. Since households are hit with idiosyncratic location preferences that follow type-1 extreme value distribution, we can estimate the probability with which households choose state s.

$$Pr(V_{is} \geq V_{ik} \forall k) = \frac{\exp(v_s/\sigma_v)}{\sum_k \exp(v_k/\sigma_v)} \tag{5}$$

In this model, location decision of households sets the population shares of the states in the next period. Thus we would have:

$$N_{s,t+1} = \frac{\exp(v_{s,t}/\sigma_v)}{\sum_k \exp(v_{k,t}/\sigma_v)} \tag{6}$$

This translates to the the fraction of households that prefer to stay in state $s$ relative to all other states given the realization of tax rates, amenities, wages and rents in state $s$ in the prior period. A spatial equilibrium would require that the ex ante (pre-idiosyncratic shock) expected utility is the same across locations. Since each household chooses locations which give them maximum utility, we need to calculate the expected maximum utility

$$\mathbb{E}(\max\{V_{is}\}_{s=1}^{S}) = \sigma_v \ln\left(\sum_s \exp(v_s/\sigma_v)\right) + C\sigma_v \tag{7}$$

Where C is the Euler Mascheroni constant[7].

## 7.2 Housing Market

To endogenize the cost of housing and introduce a congestion force, I model a simple housing market with housing supply defined as $H_{s,t} = (\bar{H} r_{s,t})^{\phi_s}$ where $\bar{H}$ is a constant housing productivity shifter and $r_{s,t}$ is rents with state-specific supply elasticity $\phi_s > 0$. Landowners make zero profits and face a downward sloping demand curve from the households' decision. Housing-market clearing implies:

$$H_{s,t}^S = N_{s,t} h_{s,t}^* \implies N_{s,t} w_{s,t}(1-\tau_{st})(1-\alpha)/r_s = (\bar{H} r_{s,t})^{\phi_s} \tag{8}$$

This implies that the equilibrium rent in the market is given as

$$r_{s,t} = \left(\frac{N_{s,t} w_{s,t}(1-\tau_{st})(1-\alpha)}{\bar{H}^{\phi_s}}\right)^{1/1+\phi_s} \tag{9}$$

Provided $\phi_s > 0$, rents increase in the population share of a state and wages, while decreasing in the state tax rate and housing productivity $\bar{H}$.

**Proposition 1:** *Housing supply elasticities explain variation in net migration elasticities across states*
Plugging in equation 9 in equation 6, taking logs and differentiating w.r.t $\ln(1-\tau_{s,t})$ we get population location elasticity w.r.t changes in taxes:

$$\frac{\partial \ln N_{s,2}}{\partial \ln(1-\tau_{s,1})} = (1-N_{s,2})\frac{1}{\sigma_v}\frac{\alpha+\phi_s}{1+\phi_s} \approx \frac{1}{\sigma_v}\frac{\alpha+\phi_s}{1+\phi_s} \tag{10}$$

Where the approximation is true for sufficiently large number of states[8]. Since the RHS is increasing in $\phi_s$ location choice (or in-migration) elasticity with respect to tax-rates positively depends on the housing supply elasticity, the share of the consumption good, and inversely proportional to the dispersion of idiosyncratic location preference shocks. Thus migration due to tax changes inversely correlates with housing supply elasticity.

[7] $P(M = \max(V_s)) = \Pi_{s=1} \exp(-\exp(-(x-v_s)/\sigma_v)) = \exp(-\sum_s \exp(-(x-v_s)/\sigma_v)) = \exp(-\exp(-(x-\mu^*)/\sigma_v))$ where $\mu^* = \sigma_v \sum_s \exp(v_s/\sigma_v)$ so $M \sim Gumbel(\mu^*, \sigma_v)$ and $E(M) = \mu + \sigma_v C$

[8] An alternative derivative would be simply $\frac{-\ln N_{s,2}}{\partial \ln \tau_{s,1}} \approx \frac{1}{\sigma_v(1-\tau_s)}\frac{\alpha+\phi_s}{1+\phi_s}$ Thus states that have higher ex-ante tax levels would face a greater out-migration risk.

In this framework, states with relatively inelastic housing supply elasticities, especially in their urban MSAs are more likely to exhibit higher net out-migration when effective tax rates go up. The prediction is empirically supported in figure 10. I plot the predicted net outflow of individuals after the introduction of SALT deduction cap against the population-weighted state-wise housing supply elasticity estimated by Saiz (2010). The negative relationship shows that states that exhibit higher migration risk have inelastic housing supply markets which proxies stronger congestion forces.

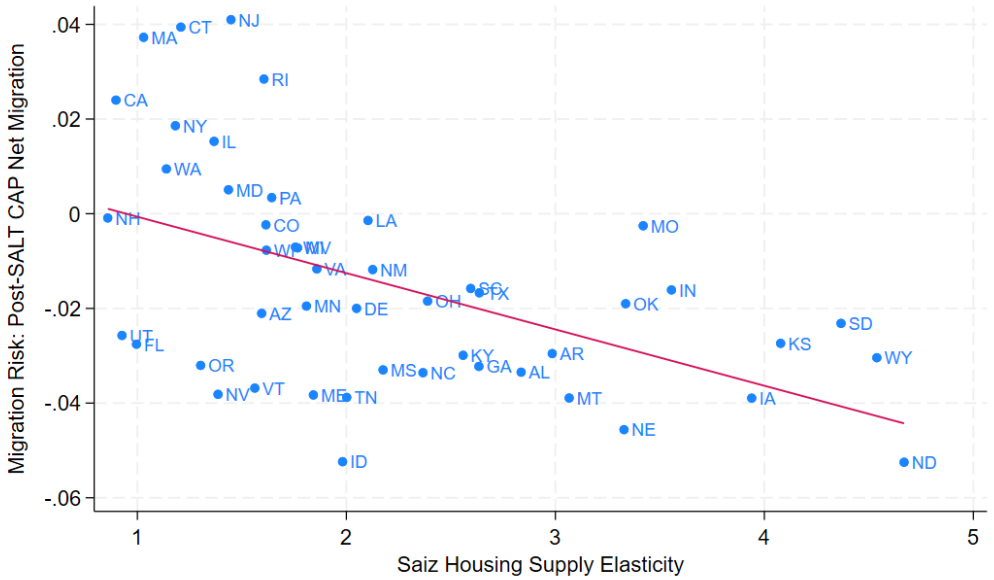



Figure 10: What explains variation in migration elasticities?: Predicted Net Outflows vs. Housing Supply Elasticity

This graph plots the predicted-net outflows due to increase in effective federal and state tax liability after the introduction of SALT deduction cap under the Tax Cuts and Jobs Act (2017) versus the housing supply elasticity estimated by taking a population-weighted average of MSA-level supply elasticities from Saiz (2010).

## 7.3 Firms

Competitive firms with constant returns to scale production are located in each state which use only labor as input to produce the homogeneous consumption good. The representative firm's objective function is:

$$\max_{L} \pi(z_s, l_s) = z_{st} L_{s,t} - w_{s,t} L_{s,t} \tag{11}$$

Where $z_s$ is location specific productivity and $w_s$ is wage. Profit maximization yields $w_{s,t} = z_{s,t}$. In the current version of the model, I shut-off location specific productivity to focus only on variation explained by place-based policies. Since households provide labor in-elastically, labor market clearing $L^D_{s,t} = N_{s,t}$. Therefore, the state's domestic production ends up being $Y_{s,t} = z_{s,t} N_{s,t}$ and therefore state GDP growth has the form:

$$\frac{Y_{s,t+1}}{Y_{s,t}} = \frac{z_{s,t+1}}{z_{s,t}} \times \frac{N_{s,t+1}}{N_{s,t}} = (1 + g_z) \times (1 + g_{s,N}) \tag{12}$$

Therefore, a state's GDP growth, which in this case is equivalent to the state's tax base growth, is the product of aggregate productivity growth and the state's own population growth. Since aggregate population stays fixed in this model across periods, any growth in population reflects relative net migration - states that receive more inflows than outflows would have a growing population and therefore tax base.

## 7.4 State Government Objective

State and local governments set tax rates to spend on state-specific fixed government expenditure plans $X_{s,1}$ and $X_{s,2}$ as well set contributions to finance states' defined benefit public pension plans with legacy liabilities $L_s$ and assets $A_s$. Since payment of retirement benefits form a part of a state's operational budget, states need to meet obligations every period as per their balanced budget requirements. Given the contribution and pension assets, pension fiduciaries - which are often state investment boards in practice - allocate pension assets

to risky and risk-free assets. The timing of the decisions are as follows: State governments, born with an endowment of population $N_{s,1}$ and fixed expenditures $X_{s,1}$ and $X_{s,2}$, simultaneously set tax and contribution rates, given the investment policy of pension fiduciaries, to minimize discounted present value of political costs which are convex in tax rates. Minimizing political costs could be thought of as equivalent to maximizing the probability of staying in power. Households respond to tax rates set in the first period, along with amenities and housing costs, to choose their location for period 2. An important assumption is that state governments do not internalize migration effects when setting tax rates and assume that their tax base would grow at the common, fixed growth rate of $g_z$. The minimization is outlined here:

### 7.4.1 Political Executive's Problem

The political executive chooses $\tau_1, \tau_2 \in [0,1]$ to minimize current and future convex tax-costs $F(\tau)$, subject to the budget-and-feasibility constraints:

$$\min_{\{\tau_{s,1},\tau_{s,2},c_{s,1},c_{s,2}\}} F(\tau_{s,1}) + \beta\mathbb{E}(F(\tau_{s,2})) \tag{13}$$

$$\text{s.t.} \quad \text{(i)} \quad \tau_{s,1} w_s N_{s,1} = X_{s,1} + C_{s,1} \overset{div.\ by\ Y_{s,1}}{\Longrightarrow} \tau_{s,1} = x_{s,1} + c_{s,1} \tag{14}$$

$$\text{(ii)} \quad \tau_{s,2} w_s N_{s,2} = X_{s,2} + C_{s,2} \overset{div.\ by\ Y_{s,2}}{\Longrightarrow} \tau_{s,2} = x_{s,2} + c_{s,2} \tag{15}$$

$$\text{(iii)} \quad C_{s,2} + (A_s + C_{s,1})R_s = L_s \overset{div.\ by\ Y_{s,1}}{\Longrightarrow} c_{s,2}(1+g_s) + (a_s + c_{s,1})R_s = l_s \tag{16}$$

$$\text{Where, } R_s = 1 + r_f + \theta(\mu - r_f + \sigma\epsilon) \quad \& \quad \epsilon \sim N(0,1) \tag{17}$$

Where $c_{s,1}, c_{s,2}$ are pension contributions-state GDP ratio, $\tau_{s,1}, \tau_{s,2}$ are tax rates and $x_{s,1}, x_{s,2}$ are government expenditures-state GDP ratio. Notice that the government budget constraint does not contain borrowings. Since pension contributions and payments are part of the operational budget, I assume that governments borrow in terms of their underfunded pensions. Hence, the intertemporal decision to save from period-one revenues is similar to the setup in Lucas and Zeldes (2009). Optimization gives us the following condition:

$$1 + g_s = \beta\mathbb{E}\left[R_s \frac{F'(\tau_2)}{F'(\tau_1)}\right] \tag{18}$$

This is an Euler equation for tax policy with risky pensions and state tax base growth (in the form of GDP growth). Note that the political executive does not internalize the migration feedback when setting tax rates. This means when the conjectured state growth rate is larger, the political executive expects the future tax base to expand making it optimal to tilt more burden to the future. Alternatively, if asset returns are higher, the state government can contribute less in the first period by substituting contributions for returns.

To see how state-specific growth gets affected by taxes, I can expand equation 12 and re-write it as the following

$$1 + g_s = \frac{z_2 N_{s,2}}{z_1 N_{s,1}} = (1+g_z) \times \frac{N_{s,2}}{N_{s,1}} \tag{19}$$

Where as we have describe before:

$$N_{s,2} = \frac{\exp(B_s w_{s,1}(1-\tau_{s,1}) r_{s,1}^{\alpha-1})}{\sum_k \exp(B_k w_{k,1}(1-\tau_{k,1}) r_{k,1}^{\alpha-1})} \quad \text{where, } r_{s,1} = \left(\frac{N_{s,1}(1-\tau_{s,1}) w_{s,1}(1-\alpha)}{H^\phi}\right)^{1/1+\phi}$$

Thus the component of state growth which is effected by taxes might create dispersion in the *realized* growth rates for states differ from the conjectured growth rates. As states have heterogeneous funding status, states with higher burdens might be forced to set higher relative tax rates even with a common growth assumption across states. Thus I implicitly assume that dispersion in tax rates across states occurs due to (i) heterogeneity of their funding status (ii) heterogeneity of their fixed government expenditure. To link how pension fiduciaries get affected by the political executive's decisions, I model the state pension fiduciary's problem next.

### 7.4.2 Fiduciary's Problem

State governments are essentially playing a dual role: they set the tax rates which indirectly determines the rates of contributions towards their defined benefit pensions and acting as the guarantor of the pension obligations in the second period. Contributions influences fiduciary portfolio choice, which influences returns in the second period. In this section, I develop the fiduciary's problem where they take state's contributions as given and then decide to invest between a risk-free asset with return $1 + r_f$ and a risky asset with expected return $1 + \mu$ and standard deviation $\sigma$.

Let $f_2$ denote the end-period funding ratio of pension assets to liabilities:

$$f_{s,2} = \frac{(a_s + c_{s,1})}{l_s}(1 + r_f + (\mu - r_f)\theta + \theta\sigma\epsilon) \equiv m(\theta, q) + s(\theta, q)\epsilon \tag{20}$$

Where $m = q(1 + r_f + (\mu - r_f)\theta)$, $s = \theta q\sigma$ and $q = \frac{a_s + c_{s,1}}{l_s}$ and $\epsilon \sim N(0, 1)$. $\theta$ is the portion invested in the risky asset. I model the payoff for risk-averse pension fiduciaries with CARA utility as follows:

$$\Pi(f_{s,2}) = k + \xi f_2 \tag{21}$$

Note that the payoff of the pension fiduciary is linear in the end-period funding status. An alternative formulation could be that fiduciaries aim for a threshold funding status $\tilde{f}$ and receive a bonus if funding status ends up being above that threshold, taking the form $\Pi(f_{s,2}) = k + \xi(f_2 - \tilde{f})^+$. This creates an option-like payoff structure which generates stronger risk-shifting response around the threshold, akin to Carpenter (2000). I explore the alternative payoff in the appendix. Assuming a linear pay structure leads to the optimization:

$$\max_{\theta} \mathbb{E}[-\exp(-\gamma(k + \xi f_2)] \tag{22}$$

Which leads to the following optimal risky share:

$$\theta^* = \frac{\mu - r_f}{\gamma\xi q\sigma^2} \tag{23}$$

The political executive takes the fiduciary's optimal investment rule as given and sets contribution rates for the first and second period given the return share. To arrive at a parametric solution I take a functional form for $F(x) = x + \frac{\Omega}{2}x^2$ which is similar to Lucas and Zeldes (2009), where $\Omega$ is a convex cost parameter. The optimization could be condensed into an intertemporal saving decision operating through optimal period-one contributions. The derivation is in the appendix. The optimal period-one contribution is given by:

$$c^*_{s,1} = \frac{(1 + \Omega x_{s,2})(1 + g_s) + \Omega l_s \left(1 - \bar{f}_s(1 + r_f) - \frac{(\mu - r_f)^2}{\gamma\xi\sigma^2}\right) - (x_{s,1}\Omega + 1)(1 + g_s)^2}{\Omega[(1 + g_s)^2 + (1 + r_f)]} \tag{24}$$

An equilibrium for state s $\in \{1, ..., S\}$ is a collection of tax rates, pension contributions, risk-asset share, population shares and rents across the two periods, such that (i) households optimize their utility over consumption, housing and choose the best location (ii) housing market clears, (iii) firms optimize profits w.r.t. labor (iv) political executive's FOC and state budget constraints hold, (v) fiduciaries optimize the portfolio (vi) Combined with the spatial equilibrium in period 2, which translates to state-wise growths multiplied by their initial population shares adding up to zero: $\sum_s^S g_N N_{s,1} = 0$, where $g_N = (N_{s,2}/N_{s,1}) - 1$.

We get the following propositions:

**Proposition 2:** *Holding exogenous spending needs fixed, a higher ex-ante funding ratio $\bar{f}_s = a_s/l_s$ decreases optimal taxes and increases subsequent population share*

The proof is in the appendix. The intuition behind the proposition is that better funding ratio lowers the contributions required to be raised to fulfill period 2 obligations. This lowers the tax rates relative to states with similar exogenous spending needs. Relatively lower taxes induces population in-migration in the low-tax states - resulting in higher relative growth rates. This ties together with how contributions react to expected growth - states with higher perceived growth rates would behave differently than states with lower relative growth. I explore this in the next proposition:

**Proposition 3:** *Governments that over(under)estimate local growth under(over)-contribute to their pensions. Lower(Higher) than required contributions force fiduciaries to take higher(lower) risk.* Notice that optimal contribution equation 24 involves a growth term. States contributions therefore depend on their conjectured growth rate for the second period. In my example states assume a common growth rate - thus contributions across states differ primarily due to pension fundamentals and exogenous non-pension spending needs. The intuition directly comes from the finding that $\frac{\partial c^*_{s,1}}{\partial g_s} < 0$ optimal contribution-rates fall in the growth assumption. The punchline of the proposition is that states' misperception of future growth drives their pension-savings behaviour. States or their political decision-makers cannot predict migration flows, and cannot predict the actions of other states in response to possible migration flows. Thus states that see higher population outflows systematically under-contribute relative to their realized growth rates, which is consistent with what we find in figure 8.

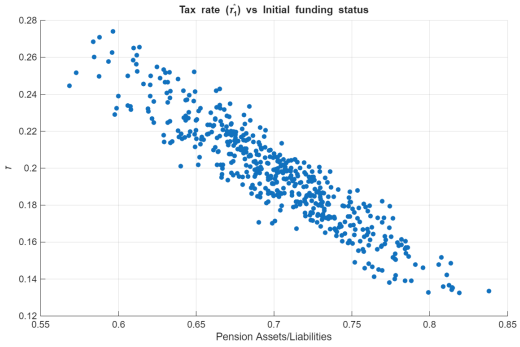
**Corollary:** *If all states assume a common growth rate, states with larger ex-ante funding shortfalls (i) raise taxes and experience net out-migration, yielding growth shortfalls; and (ii) respond by under-contributing and increasing the risky share of pension assets*

To graphically visualize the model's predictions, I simulate the model for 50 states by assuming common parameter values for the political executive's problem, the fiduciary's problem and parameters driving household

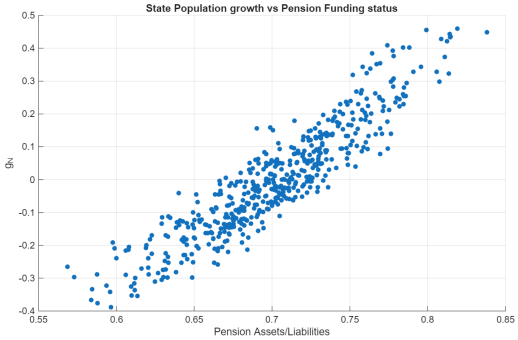
location choice. I assign random pension asset and liabilities to states, with the funding ratio targeted around a 70% cross-sectional average. I then run the simulation by carrying out a fixed-point search algorithm of the model equations, with the details given in the appendix. The results are presented in figure 11. Panel (a) shows that funding status is inversely proportional to the effective first-period tax rate. This follows from the budget rules. Panel (b) shows that population growth is positively related to the initial funding gap. The channel is that worse funded states end up charging higher taxes, which induces out-migration. This is the prediction from proposition 2. Panel (c) shows the risky asset share under the common growth assumption of $g_s = g_z = 0.02$ against the actual population growth $g_N = (N_{s,2}/N_{s,1}) - 1$. The negative relationship supports proposition 3 - states with lower population growth rates tend to have lower contributions and hence higher risky asset shares relative to their actual growth. Panel (d) Plots the graph for risky asset shares $\theta(c_{s,1}(g_z))$ against initial funding ratio $\bar{f} = a_s/l_s$ denoting two groups of states: states with above median (High) and below median (Low) population growth rates. States with worse initial funding tend to have lower population growth rates and higher risky asset share. This supports the essence of proposition 3 and the empirical finding in figure 8 - states with greater out-migration risk tend to under-contribute and take higher risks.

In the appendix, I show an extension where the state government internalizes the migration feedback of tax changes. This results in allocations and contributions being uncorrelated with migration or risky allocation.

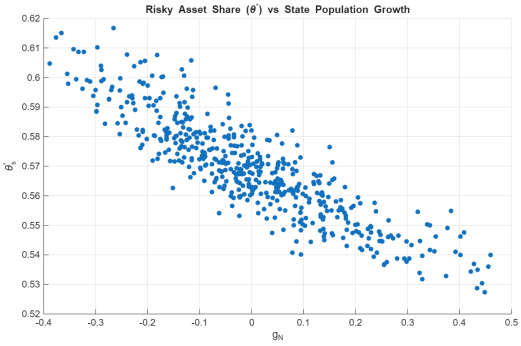
(a) Taxes and Pension Funding Status




(b) Population Growth vs Funding Status




(c) Risky Asset Share and Population Growth




(d) Risky Share vs Funding Ratio by Pop. Growth

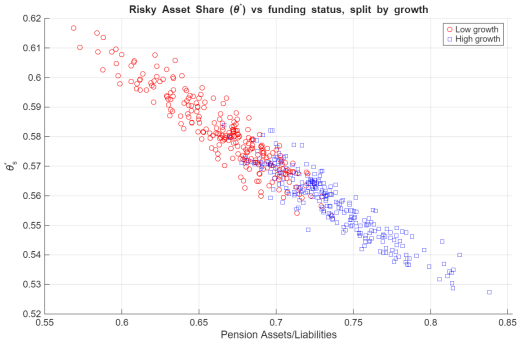



Figure 11: Model Predictions with Simulations

These figures plot simulation results for 50 states drawn 10 times each with $\Omega = 8, \gamma = 4, \xi = 2, \mu - r_f = 0.04, \sigma = 0.18, \sigma_v = 0.5, B_s = B = 3, N_{s,1} = N_1 = 0.02, \alpha = 0.65, \phi = 1 + 0.5 \times e$ with e $\sim N(0,1), w = 1, H = 1$ along with non-pension public spending assigned $x_{s,1}$, and $x_{s,2}$ randomly assigned with small N(0,1) noise around 0.05, and finally pension assets and liabilities r $a_s$ and $l_s$ assigned values with some noise around the initial funding ratio of $\bar{f}_s = a_s/l_s = 0.7$. Panel (a) plots the scatter plot of period-1 taxes against initial pension funding ratio. Panel (b) plots the scatter of state population growth from period-1 to 2 against the initial funding ratio. Panel (c) plots the scatter of risky asset share with growth assumption $g_z$,, $\theta(g_z)$ against the realized population growth rate $g_N$. Panel (d) plots risky asset share $\theta(g_z)$ against the initial funding status, by the realized population growth being above (blue) or below median (red).

# 8 Optimal Policy

What is the wedge between the optimal versus observed pension risk given tax rates and other state-level fundamentals? In the previous section, I rationalize that political costs of raising taxes and myopia leads to over or underinvestment in risky assets across state governments. In this section, I look at the social planner optimum using the local population weighted utility as the objective. The planner internalizes the migration feedback of taxes as well as the investment risk transfer to households in the form of endogenous amenities which depend on pension returns and tax receipts net of pension expenditures and obligations. Pension contributions are set to an exogenous actuarial level, while taxes set the level of non-pension expenditures. Amenities are a function of per-capita tax revenue net of pension expenditures. States vary in their ability to transform revenues to public services, which is reflected by the "Public Service Efficiency" parameter. The rest of the setup remains as before. I calibrate the model using data on state-level taxes, pension assets and liabilities as of 2023. I invert the calibrated model to recover amenity-provision parameters by state and then compare the optimal versus real risky asset share by state.

### 8.0.1 Endogenous amenities

I endogenize location specific amenities $B_s$ as increasing in non-pension spending per capita:

$$B_{s,t} = F(x_s) = \exp(\psi_s(\tau_s - \zeta l_s)) \tag{25}$$

Where $-\psi_s$ is indicative of the state's semi-elasticity of amenities generation for given level of expenditure per capita. This is to capture the efficiency of government's expenditures and a reduced form way to capture differential responses to higher tax rates as some states could charge higher marginal rates but maintain amenities more efficiently. This means the residual resources left to the state after paying for legacy liabilities pins down the level of amenities in the state each period.

### 8.0.2 Two Period Optimization

The objective of the two-period optimization is to pin down optimal policy subject to legacy pension underfunding and mobility constraints. States trade-off between providing higher services through funding amenities which are residual tax revenues after accounting for states' fixed pension contributions. The amount of end-period pension obligation is a function of legacy liabilities net of investment income. Thus, higher risk shares could potentially decrease the need for raising taxes. Each state government sets the tax rates and risky-asset share of the pension investment in order to maximize the discounted sum of expected welfare of its constituents. The timeline is as follows: each state is endowed with an initial population share and legacy pension fundamentals. Each state government then optimizes the expected utility of their states at the beginning of period 1 by setting optimal tax rates for the two periods and the risky asset share, given a mandatory early contribution rule. Given government policies across states, households with idiosyncratic location preferences choose states which give them the highest indirect utility in period 2. In period 2 market returns on pension assets are realized and government has to pay remaining pension liabilities. The optimization can be summarized as:

$$\max_{\{\theta_s, \tau_{s,1}, \tau_2\}} V_{s,1} N_{s,1} + \delta \mathbb{E}(V_{s,2} N_{s,2}) \tag{26}$$

subject to period-by-period budget constraints:

$$\tau_{s,1} = x_{s,1} + \zeta l_s \qquad \tau_{s,2} = x_{s,2} + \frac{1}{1+g_s}\left(l_s - (a_s + \zeta l_s) R_s\right) \tag{27}$$

where $\zeta l_s$ is the actuarial required contribution to pensions, $R_s$ is the gross return on pension investment. Household location choice is given by the previous period's realized utility:

$$N_{s,2} = \frac{\exp(v_{s,1}/\sigma_v)}{\sum_k \exp(v_{k,1}/\sigma_v)} \tag{28}$$

given public service or amenities provision function:

$$B_{s,t} = 1 - \exp(-\psi_s x_{s,t}) \tag{29}$$

Thus, this government is faced with a problem to optimize welfare by providing amenities which are an increasing function of the resources it raises over two periods.

$$\max_{\tau_{s,1},\tau_{s,2},\theta_s} (1-\exp(-\underbrace{\psi_s(\tau_{s,1}-\bar{c}_s)}_{\delta_{s,1}}))\underbrace{K_s(w(1-\tau_{s,1})N_{s,1})^{\varphi_s}}_{v_{s,1}}+$$
$$\left[1-\exp\left(-\underbrace{\psi_s\left(\tau_{s,2}-\frac{l_s}{1+g_s}+\mu_R(\theta_s)\left(\frac{a_s+\bar{c}_s}{1+g_s}\right)-\psi_s\left(\frac{a+\bar{c}_s}{1+g_s}\right)^2\frac{\sigma^2\theta_s^2}{2}\right)}_{\delta_{s,2}}\right)\right]\times\underbrace{K_s((1-\tau_{s,2})wN_{s,2})^{\varphi}}_{v_{s,2}} \tag{30}$$

**First-order conditions**

(i) FOC with respect to the statutory tax rate $\tau_{s,1}$

$$v_{s,1}\left[\frac{\varphi_s}{(1-\tau_{s,1})}-\psi_s\exp(-\delta_{s,1})\right]=\frac{v_{s,2}\epsilon_s\psi_s\exp(-\delta_{s,2})}{\tau_{s,1}}\times\left[\frac{l_s}{1+g_s}-\mu_R(\theta_s)\frac{a_s+\bar{c}_s}{1+g_s}-\psi_s\sigma^2\theta_s^2\left(\frac{a_s+c_s}{1+g_s}\right)^2\right]+$$
$$\frac{\varphi_s\epsilon_s}{\tau_{s,1}}v_{s,2}(1-\exp(-\delta_{s,2})) \tag{31}$$

Where the first term of the LHS reflects the impact of raising labor income tax rates on today's amenities, the second term is the impact on today's disposable income, the third term is the impact on housing rents going down as disposable income goes down. The combined fourth terms relates to change in the state population tomorrow and it's knockdown affect on state's tax base which changes it's amenities, the change in rents as housing markets gets either more crowded or less congested, the third term is documenting the per-capita change in utility of households tomorrow. Rest of the derivation is in the appendix. This FOC leads to possibly multiple solutions for optimal $\tau_{s,1}$ as $N_{s,2}$ is dependent on indirect period 1 utility $V_{s,1}$.

(ii) FOC with respect to the statutory tax rate $\tau_2$

$$\tau_{s,2}=1-\frac{(1-\exp(-\delta_{s,2}))\varphi_s}{\exp(-\delta_{s,2})\psi_s}=1+(1-\exp(\delta_{s,2}))\frac{\varphi_s}{\psi_s} \tag{32}$$

Where the first term captures the impact on expected future amenities due to higher tax rates, the second term captures the negative impact on wages and the third term captures the impact on rents. Since period 2 tax rates only affect period 2 outcomes, it is safe to say that the trade-off is in equating the cost of lower disposal income against higher amenities and lower rents.

(iii) FOC with respect to the portfolio share $\theta_s$

$$\theta_s=\frac{\mu-r_f}{\psi_s\sigma^2}\frac{1+g_s}{a_s+\bar{c}_s} \tag{33}$$

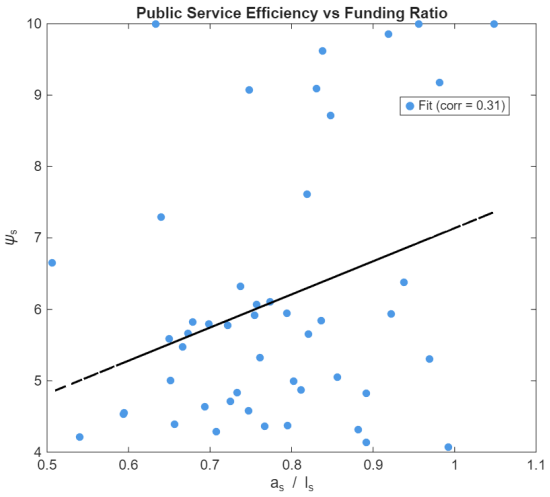
Thus optimal risky share is directly proportional to the excess return to variance ratio of the risky investment but inversely proportional to pension assets per capita and the public service provision efficiency of the state government. As the expression of amenities is similar to a CARA normal utility, we can analogously equate the "risk-aversion" of the social planner as the public service efficiency multiplied by pension assets per capita.

## 8.1 Inversion and Wedge Estimation

I calibrate the model using data from 2023. I use tax-rate data by state using tax data from NBER TAXSIM. I take data on pension assets, liabilities and required contributions from the Public Plans Database. I normalize wage rate to 1. I take housing supply elasticity from Saiz (2010). I use the tax-migration elasticity using the Moretti and Wilson (2017) method by state. I assign an equity risk premia of 6% and volatility of 18%. Consumption share of household expenditures is set to 66%. I take population data by state from US census bureau. Using these data, I obtain the measure of public service efficiency by state by solving for the first-order-conditions of taxes.

Panel (a) of figure 12 shows that public service efficiency parameter is positively correlated with the funding ratio of the state. Though noisy, this in essence captures the structural variation possibly driving the patterns we see in data. States with high public service provision ability are more likely to have better funding status. In other words, pension funding is correlated with a measure of "governance quality". A worse funding status would be costlier in states with lower public service efficiency - as a drop in public spending could result in worse amenities relative to other states. This means a common shock to pension funding, such as a financial crisis, would translate into a heterogeneous impact - states with better public service efficiency could recover

(a) $\psi_s$ vs. $FR$




(b) $\theta_{optim}$ vs. $\theta_{data}$

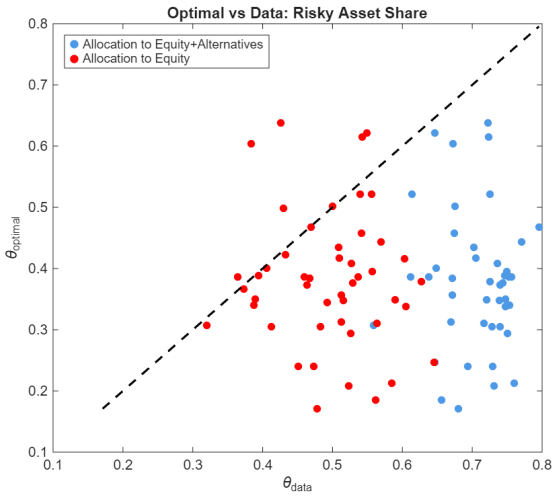



Figure 12: Optimal Policy

These graphs show the relationship between the public service efficiency parameter and funding ratio as well as the wedge between optimal and observed risky-asset shares. Panel (a) plots the public service efficiency parameter $\psi_s$ against the funding ratio of the respective state. Panel (b) plots the model-implied optimum risky asset share against the observed risky asset share in data. The red-dots show the model $\theta$ vs. allocation to public equities observed in the data. The blue-dots show the sum of allocations to public equity and alternative asset classes.

faster while states with worse $\psi_s$ would find it costlier, lose residents to better funded states and thus end up at a worse position.

Assuming their tax-rates correspond to optimal, are states investing optimally? The optimal risky-asset-share formula tells us that states with lower public service efficiency parameters thus should invest more in risky assets. How different are these investment observed in the data versus those implied by the model. Panel (b) of figure 12 plots the observed and model-implied allocations. The scatter in red represents allocation to public equities. We see that about 12 states' public equity allocations are equal or similar to their optimal risky asset shares implied by the model. However, if we account for the additional allocation to alternative asset classes, we find that states allocate much higher than the model-implied optimum. This gap quantifies the excess risk taken by public pension and panel (b) shows that a large part of it comes from allocation to alternative asset classes.

# 9 Discussion and Conclusion

Public sector pension underfunding remains a significant challenge for US state and local governments. This paper traces the link between chronic funding gaps and migration. Persistent shortfalls coincide with higher taxes, reduced public service quality, and other adverse local economic conditions that encourage households and firms to relocate. Governments face a stark choice: raise taxes, under-invest in public services, or pursue riskier investment strategies within their pension portfolios. I show that migration limits fiscal capacity and is associated with states under-contributing to their defined benefit pensions. I then find that states that face higher migration risk pivot towards riskier alternative investments.

I further examine the impact of statutory pension reforms on funding status in Appendix Section A.1. Using data from the National Association of State Retirement Administrators (NASRA)[9] , I collect 180 instances of reforms enacted by state and local pensions between 2005 and 2022. The most common measures involve raising employee contribution rates and reducing pension benefits. On average, funds in states that implement such reforms exhibit improved funding status following their adoption. However, when I explore heterogeneity by migration risk, I find that plans in high-migration-risk states experience smaller improvements in funding ratios relative to other states enacting the same reforms. This pattern underscores the constrained policy space these states face, which may in turn push them toward risk-seeking investment strategies.

That said, there are several promising directions for extending this research, both empirically and theoretically. On the data and measurement side, I could expand beyond the 228 state and local pensions in the Public Plans Database by incorporating GASB 67 statements, and adjust funding status using discount rates

[9] `https://www.nasra.org/pensionreform`

based on risk-free benchmarks rather than reported assumptions (Giesecke and Rauh, 2023). Empirically, my current strategy shows that relative portfolio performance of defined-benefit pensions affects local government budgets, which, in turn, influences household location decisions through higher tax burdens and deteriorating public amenities. To establish a stronger causal link, however, future work would need to compare the effects of local funding and tax shocks against other labor-mobility shocks, and examine how these forces interact.

Theoretically, my model captures how heterogeneous legacy pension obligations force local governments to set heterogeneous tax rates, and households migrate in response to realized tax burdens and amenities, seeking the locations that maximize their utility. Future work could microfound richer mechanisms behind migration. One channel is reduced labor demand in underfunded regions, as firms relocate to better-funded states with lower tax burdens—a pattern well documented in Suárez Serrato and Zidar (2016) and Giroud and Rauh (2019). I do not model or test firm location choice, largely because of limited data relative to the richness of IRS migration flows and the complexities of firms apportioning income across tax jurisdictions (see Suárez Serrato and Zidar (2016)). A deeper analysis could leverage datasets such as the Longitudinal Business Database or the National Establishment Time Series to track both firm relocation and regional business dynamism by funding status.

Another potential channel is expectations: households may migrate not only in response to realized taxes and amenities, but also in anticipation of future fiscal stress. In Appendix Table 15, I provide suggestive but noisy evidence from the Federal Reserve Bank of New York's Survey of Consumer Expectations, which asks respondents whether they expect higher taxes over the next 12 months. I find a negative association between pension funding status and expected tax increases, consistent with households in more underfunded states anticipating heavier tax burdens ahead.

Recent work by Foschi et al. (2025) highlights labor mobility as a central force in shaping regional economic dynamics. Building on this, an important broader question is how migration responses to prolonged pension underfunding influence regional growth and distributional outcomes. Addressing this requires a more comprehensive framework that considers the impacts on all key stakeholders—households, firms, and landowners. Since underfunding effectively signals a local government's fiscal capacity, it constrains the scope of place-based policies and may thus inform the debate over which policy interventions are most conducive to fostering regional economic growth.

The theoretical framework develops a spatial public finance model with a state-level portfolio choice problem for pension fiduciaries, linking public risk-taking to local economic outcomes. A richer structural exercise could estimate key parameters using the Simulated Method of Moments or related techniques, allowing quantification of the welfare costs of public pension risk-taking. Such a framework would also enable counterfactual analyses to evaluate which place-based policies could mitigate the distortionary effects of chronic underfunding.

Therefore, this paper makes two incremental contributions: (i) it links household migration to local pension funding status, and (ii) it documents how pension fund allocations toward riskier asset classes vary with relative migration risk. Together, these findings open the door to broader questions on regional economic dynamics and the consequences of public risk-taking for household outcomes.

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# A Empirical Appendix

Table 4: Summary-Statistics

This table presents the summary statistics of the data used in this paper. The first panel shows state-level data. Tax data comes from NBER Taxsim and the Tax Foundation. Public School Data comes from NCES. Net AGI outflow and migration data comes from the IRS Statistics of Income, Real Wage comes from CPS, Rent comes from ACS, State Finances data comes from Census Bureau, Job openings data comes from Bureau of Labour Statistics, Crime statistics are from the FBI, Road Condition is from Bureau of Transportation. Pension plan level data comes from Center for Retirement Research. 13-F holdings and stock return data comes from LSEG and Pensions' Balance Sheet Data comes from Worldscope. Data ranges 2001-2023.

| | Mean | Median | SD | Min | Max | Count |
|---|---|---|---|---|---|---|
| State-level Data | | | | | | |
| State Marginal Income Tax Rate for AGI >$250k (%) | 5.14 | 5.75 | 2.92 | 0.00 | 12.90 | 1070 |
| State Corporate Tax Rate (0-1) | 0.06 | 0.07 | 0.03 | 0.00 | 0.12 | 1121 |
| Ln(Public School Capital Expenditure) Ln($) | 20.25 | 20.40 | 1.17 | 16.84 | 23.17 | 1100 |
| Net AGI outflow per taxpayer ($ 1000) | -0.08 | 0.00 | 0.84 | -4.86 | 6.25 | 1109 |
| Ln(Real Wage) | 1.05 | 1.03 | 0.21 | 0.41 | 1.69 | 1121 |
| Ln(Rent) | 6.85 | 6.83 | 0.28 | 6.20 | 7.70 | 1121 |
| State Debt/Revenue (0-1) | 0.54 | 0.47 | 0.29 | 0.06 | 2.46 | 1094 |
| State Debt/GDP (0-100) | 68.88 | 62.61 | 35.23 | 8.27 | 203.61 | 1094 |
| Annual Property Tax/Household Income (0-1) | 0.06 | 0.03 | 0.13 | -0.52 | 2.92 | 1121 |
| House Price Index | 201.30 | 188.05 | 57.77 | 100.80 | 574.37 | 1121 |
| Job Openings (%) | 3.62 | 3.30 | 1.31 | 1.60 | 8.90 | 1121 |
| Violent Crimes/100k population | 400.51 | 356.05 | 200.58 | 78.24 | 1637.86 | 1070 |
| Pupil-Teacher Ratio | 15.37 | 14.90 | 2.65 | 10.31 | 24.33 | 1097 |
| Road Condition (0-100) | 80.85 | 83.85 | 15.21 | 3.10 | 99.90 | 1018 |
| State GDP Growth (0-1) | 0.04 | 0.04 | 0.04 | -0.15 | 0.25 | 1121 |
| State Employment Growth (0-1) | 0.01 | 0.01 | 0.02 | -0.11 | 0.08 | 1121 |
| Pension-plan level Data | | | | | | |
| Funded Ratio (PPD) | 0.78 | 0.79 | 0.20 | 0.00 | 2.66 | 5035 |
| 1-year Investment Return | 0.07 | 0.08 | 0.11 | -0.36 | 0.39 | 5128 |
| 1(Assumption Change) | 0.13 | 0.00 | 0.33 | 0.00 | 1.00 | 4837 |
| Contribution Growth (0-1) | 0.16 | 0.05 | 2.07 | -1.00 | 98.54 | 4862 |
| 1-year Investment Return Assumption | 0.08 | 0.08 | 0.01 | 0.03 | 0.09 | 5070 |
| Payroll Growth | 0.03 | 0.03 | 0.08 | -1.00 | 2.50 | 4583 |
| Payroll Growth Assumption | 0.04 | 0.04 | 0.01 | 0.00 | 0.08 | 2629 |
| Allocation to Equities | 0.52 | 0.53 | 0.12 | 0.00 | 1.00 | 4302 |
| Allocation to Fixed Income | 0.26 | 0.25 | 0.09 | 0.00 | 1.00 | 4302 |
| Allocation to Private Equity | 0.06 | 0.05 | 0.07 | 0.00 | 0.43 | 4302 |
| Allocation to Hedge Funds | 0.05 | 0.00 | 0.07 | 0.00 | 0.52 | 4302 |
| Allocation to Commodities | 0.02 | 0.00 | 0.04 | 0.00 | 0.31 | 4302 |
| Allocation to Real Estate | 0.06 | 0.06 | 0.05 | 0.00 | 0.34 | 4302 |
| Pension Fund Cash | 0.02 | 0.01 | 0.02 | 0.00 | 0.37 | 4302 |
| Allocation to Other Assets | 0.00 | 0.00 | 0.02 | 0.00 | 0.47 | 4302 |
| IRS County-to-County Migration Data | | | | | | |
| Out-Migrant Returns/Non-Migrant Returns | 0.0004 | 0.0001 | 0.0014 | 0.0000 | 0.2251 | 866,804 |
| Out-Migrant AGI ($1000s)/Non-Migrant Returns | 0.0209 | 0.0074 | 0.0638 | -5.7001 | 6.7045 | 866,801 |
| Destination-Origin Pension Funding Ratio | 0.0113 | 0.0125 | 0.1985 | -1.7417 | 1.7417 | 880,751 |
| Destination-Origin Cum. Abnorm. Returns | 0.0060 | 0.0000 | 0.3239 | -2.0360 | 2.0360 | 927,008 |
| Destination-Origin Abnormal Returns | 0.0006 | 0.0004 | 0.0642 | -0.4898 | 0.4754 | 880,751 |
| IRS County-level Migration Data | | | | | | |
| $Z_r$ | -0.16 | -0.23 | 0.29 | -0.64 | 1.63 | 67,984 |
| $Z_{union}$ | 0.58 | 0.59 | 0.22 | -0.15 | 1.18 | 64,820 |
| Net Out-Migrant Returns/Non-Migrant Returns | -0.00 | 0.00 | 0.03 | -1.12 | 1.00 | 64,770 |
| Net Out-Migrant AGI/Non-Migrant AGI | -0.14 | -0.01 | 1.22 | -83.92 | 29.46 | 64,770 |

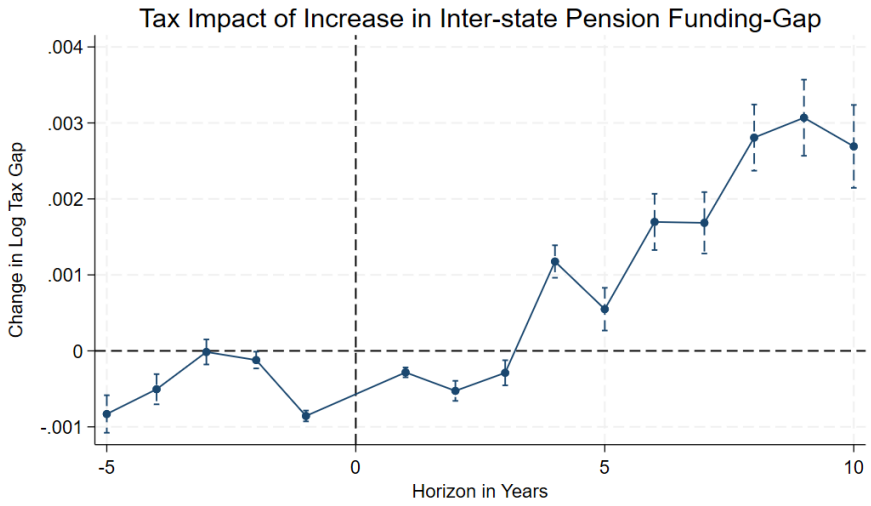



Figure 13: Impact of increase in Funding gap on Tax differential

This graph reports the coefficients of a two-stage least squares regressions of the tax gaps between destination and origin counties on their respective pension funding gaps instrumented by the relative performance of the pensions' investment portfolios. The specification is as follows:

$$\ln\left(\frac{1-\tau_{d,t+h}}{1-\tau_{o,t+h}}\right) - \ln\left(\frac{1-\tau_{d,t}}{1-\tau_{o,t}}\right) = \beta_h D_{odt} + \gamma X_{odt} + F_{odt} + \epsilon_{od,t+h}$$

Where the denotes the change in the ratio of destination and origin tax gaps from periods t, t+h, a higher gap pointing to lower taxes in the destination relative to the origin. The dependent variable is a dummy measuring the change in funding gap today between destination and origin counties: It is 1 if the funding gap widens, -1 if it shrinks and zero otherwise. The dummy is instrumented with the relative surprises measured as $Z_r = (r_{d,t} - r^a_{d,2001}) - (r_{o,t} - r^a_{o,2001})$ where $r_{it}$ is the t-period's investment return and $r^a_{i,2001}$ is the investment return assumption as of 2001. $X_{odt}$ contain ratios of destination-and-origin state debt-to-revenue ratio and state corporate tax ratio. $F_{odt}$ represents US Census Region of Origin $\times$ Region of Destination $\times$ Year fixed effects. All standard errors are clustered at the destination-origin pair level.

Table 5: Migration Odds vs Funding Gap: Public Unionization Instrument

This regression tables shows the results of regressing county-level migration odds against the pension funding gap between destination and origin counties. Panel A shows the OLS regression with logged ratio of out-migrant returns , from the origin county o to destination county-d, to the non-migrant returns of origin county-o as the dependent variable. Panel B has the logged ratio of the out-migrant adjusted growth income to non-migrant returns as the dependent variable. Panel C contains the first-stage of the two-stage least square regression with the funding gap between the destination and origin counties being the endogenous variable and the difference of the ratio of public union members by total government employs (union density) between the destination and origin counties being the instrument. Panel D shows the log odds ratio of migration and Panel E shows the logged out-migrant AGI per non-migrant tax payer as the dependent variable with instrumented funding gap being the main independent variable. Column (1) carries no fixed effects, while column (2) has destination fixed effects and origin fixed effect. Columns (3)-(8) carry destination × origin-county-FE. Column (4) additionally contains origin-state × year fixed effects whereas Column (5) contains regressions with destination state × year fixed effects. Column (6) contains regressions with destination region × origin region × year fixed effects, where region corresponds to one of the nine US Census Bureau regions designated to the county's state. Column (7) contains origin-county × year fixed effects while Column (8) contains destination-county × year fixed effects. All regressions include the logged ratio of destination rents and origin rents and the logged ratio of destination wages and origin wages as controls. Standard errors are clustered by destination-origin pair.

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
|---|---|---|---|---|---|---|---|---|
| *Panel A: OLS, Dep. Var:* | Ln(Out Migrant Returns/Non-migrant Returns) | | | | | | | |
| Funding Gap | 0.190*** | -0.00194 | 0.00817** | 0.00424 | -0.0156*** | -0.0466*** | 0.0422*** | 0.0118** |
| | (10.14) | (-0.40) | (2.07) | (1.07) | (-3.96) | (-11.22) | (6.62) | (2.04) |
| *Panel B: OLS, Dep. Var:* | Ln(Out Migrant AGI/Non-migrant Returns) | | | | | | | |
| Funding Gap | 0.277*** | 0.00495 | 0.0136** | -0.00362 | 0.0154*** | -0.0405*** | 0.0570*** | 0.0213** |
| | (15.09) | (0.70) | (2.11) | (-0.63) | (2.67) | (-6.65) | (6.31) | (2.53) |
| *Panel C: IV - 1st St, Dep. Var:* | Funding Gap | | | | | | | |
| $Z_{union}$ | 0.0650*** | 0.0416*** | 0.0384*** | 0.0408*** | 0.0628*** | 0.00122 | 0.0462*** | 0.0670*** |
| | (64.97) | (12.74) | (11.33) | (14.45) | (22.29) | (0.39) | (16.17) | (23.41) |
| *Panel D: IV - 2nd St, Dep. Var:* | Ln(Out Migrant Returns/Non-migrant Returns) | | | | | | | |
| Funding Gap | 3.266*** | 6.373*** | 7.216*** | 7.303*** | 3.552*** | -107.3 | 7.014*** | 2.986*** |
| | (38.57) | (11.76) | (10.54) | (13.49) | (17.04) | (-0.36) | (15.15) | (17.62) |
| *Panel E: IV - 2nd St, Dep. Var:* | Ln(Out Migrant AGI/Non-migrant Returns) | | | | | | | |
| Funding Gap | 3.190*** | 11.44*** | 12.71*** | 9.073*** | 4.643*** | -126.7 | 8.789*** | 3.860*** |
| | (38.86) | (11.83) | (10.51) | (13.32) | (16.47) | (-0.36) | (14.95) | (16.65) |
| N | 840217 | 839892 | 818137 | 818137 | 818137 | 818136 | 809319 | 809353 |
| Destination, Origin County-FE | | Y | | | | | | |
| Destination × Origin County-FE | | | Y | Y | Y | Y | Y | Y |
| Origin State × Year-FE | | | | Y | | | | |
| Destination State × Year | | | | | Y | | | |
| Dest. Reg. × Org. Reg. × Yr-FE | | | | | | Y | | |
| Origin County × Year-FE | | | | | | | Y | |
| Destination County × Year-FE | | | | | | | | Y |
| Controls: | Ln(Destination Rent/Origin Rent), Ln(Destination Real Wage/Origin Real Wage) | | | | | | | |

$t$ statistics in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 6: Second-stage: Using Public Unionization instrument

This table reports the regression results of the two-stage least square regressions of net migrant returns to non-migrant returns and net adjusted gross income outflow per non-migrant tax payer against state pension underfunding. The first column represents the first-stage results of the aggregate funding ratio in the region (county or state if county-data is missing) against the instrument which is the ratio of union members in the state to the total number of public employees. Columns (2) and (3) show the regressions with the net outflow of tax returns relative to non-migrant tax returns as the dependent variable. Columns (4) and (5) shows the net outflow of adjusted gross income relative to non-migrant returns. Columns (2) and (4) show the OLS version of the regressions whereas columns (3) and (5) show the second stage with funding ratio of the county instrumented with the return instrument. Each specification contains the following controls: lagged log rents, real wages, property tax-to-household income ratio, top marginal state income tax rates, state debt to revenue, county-GDP growth and percentage of state's House of Representatives congressmen that belong to the Republican party. Standard errors are clustered at the county-level.

| | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| | Funding Ratio | $\frac{\text{Net Out Migrant Returns}}{\text{Non-Migrant Returns}}$ | | $\frac{\text{Net Out Migrant AGI}}{\text{Non-Migrant Returns}}$ | |
| | First-Stage | OLS | IV | OLS | IV |
| $Z_{union}$ | -0.0309*** | | | | |
| | (-4.36) | | | | |
| Funding Ratio$_{t-1}$ | | 0.00102 | -0.413 | 0.111 | -26.52* |
| | | (1.05) | (-1.64) | (1.34) | (-1.68) |
| N | 59403 | 56312 | 55245 | 56312 | 55245 |
| R-Sq | 0.804 | 0.268 | | 0.425 | |
| F-Stat | 125.5 | | | | |
| FE | County and Year | | | | |
| Controls | Rents, Wages, S&L Taxes, Debt/Rev, County-growth, Crime, Political | | | | |

$t$ statistics in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 7: Migration Odds vs. Funding Gap: Robustness by Sample period

This table reports regressions of county-level migration odds on pension funding gaps between destination and origin counties. Panel I restricts the sample to 2019 to avoid pandemic migration, and Panel II drops 2014–2016 to address IRS data concerns in DeWaard et al. (2022). Panel A uses the log ratio of out-migrant to non-migrant returns as the dependent variable; Panel B replaces this with the log ratio of out-migrant AGI to non-migrant returns. Panel C reports the first stage of the 2SLS, instrumenting the funding gap with cumulative excess returns (relative to 2001 assumptions) of destination minus origin pension plans. Panel D presents second-stage estimates for log migration odds, and Panel E for log out-migrant AGI per non-migrant return. Column (1) includes no fixed effects; Column (2) adds origin and destination fixed effects; Columns (3)–(8) include destination × origin county-pair FE, with additional specifications: origin-state × year (Col. 4), destination-state × year (Col. 5), destination-region × origin-region × year (Col. 6, using nine Census regions), origin-county × year (Col. 7), and destination-county × year (Col. 8). All regressions control for log rent and wage ratios (destination/origin). Standard errors are clustered by destination–origin pair.

Panel I: Sample Ranging 2001-2019

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
|---|---|---|---|---|---|---|---|---|
| *Panel A: OLS, Dep. Var:* | Ln(Out Migrant Returns/Non-migrant Returns) | | | | | | | |
| Funding Gap | 0.208*** | 0.0196*** | 0.0330*** | 0.0621*** | 0.0263*** | -0.00655 | 0.0675*** | 0.0250*** |
| | (10.99) | (4.10) | (8.74) | (9.88) | (4.49) | (-1.61) | (10.82) | (4.50) |
| *Panel B: OLS, Dep. Var:* | Ln(Out Migrant AGI/Non-migrant Returns) | | | | | | | |
| Funding Gap | 0.297*** | 0.0345*** | 0.0471*** | 0.0875*** | 0.0286*** | 0.00237 | 0.0926*** | 0.0259*** |
| | (16.04) | (5.08) | (7.79) | (9.69) | (3.32) | (0.38) | (10.26) | (3.09) |
| *Panel C: IV - 1st St, Dep. Var:* | Funding Gap | | | | | | | |
| $Z_r$ | 0.0895*** | 0.0850*** | 0.113*** | 0.102*** | 0.110*** | 0.0950*** | 0.103*** | 0.110*** |
| | (18.96) | (47.16) | (45.66) | (70.57) | (65.41) | (41.19) | (68.56) | (63.34) |
| *Panel D: IV - 2nd St, Dep. Var:* | Ln(Out Migrant Returns/Non-migrant Returns) | | | | | | | |
| Funding Gap | -0.345** | 0.305*** | 0.102*** | 0.0699* | 0.214*** | 0.215*** | 0.0612 | 0.248*** |
| | (-2.12) | (6.56) | (4.29) | (1.66) | (5.85) | (7.82) | (1.47) | (7.08) |
| *Panel E: IV - 2nd St, Dep. Var:* | Ln(Out Migrant AGI/Non-migrant Returns) | | | | | | | |
| Funding Gap | -0.477*** | 0.338*** | 0.217*** | 0.0637 | 0.182*** | 0.180*** | 0.0856 | 0.202*** |
| | (-2.91) | (5.66) | (5.70) | (1.07) | (3.48) | (4.57) | (1.46) | (4.00) |
| N | 753167 | 752844 | 732063 | 732063 | 732063 | 732062 | 724350 | 724331 |
| Destination, Origin County-FE | | Y | | | | | | |
| Destination × Origin County-FE | | | Y | Y | Y | Y | Y | Y |
| Origin State × Year-FE | | | | Y | | | | |
| Destination State × Year | | | | | Y | | | |
| Dest. Reg. × Org. Reg. × Yr-FE | | | | | | Y | | |
| Origin County × Year-FE | | | | | | | Y | |
| Destination County × Year-FE | | | | | | | | Y |
| Controls: | Ln(Destination Rent/Origin Rent), Ln(Destination Real Wage/Origin Real Wage) | | | | | | | |

Panel II: Sample without 2014-2016

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
|---|---|---|---|---|---|---|---|---|
| *Panel A: OLS, Dep. Var:* | Ln(Out Migrant Returns/Non-migrant Returns) | | | | | | | |
| Funding Gap | 0.188*** | -0.00255 | 0.0111*** | 0.0395*** | 0.0183*** | -0.0483*** | 0.0422*** | 0.0184*** |
| | (10.21) | (-0.52) | (2.72) | (6.13) | (2.98) | (-11.37) | (6.66) | (3.20) |
| *Panel B: OLS, Dep. Var:* | Ln(Out Migrant AGI/Non-migrant Returns) | | | | | | | |
| Funding Gap | 0.271*** | 0.00589 | 0.0180*** | 0.0553*** | 0.0317*** | -0.0417*** | 0.0584*** | 0.0300*** |
| | (15.12) | (0.82) | (2.68) | (5.96) | (3.56) | (-6.65) | (6.39) | (3.49) |
| *Panel C: IV - 1st St, Dep. Var:* | Funding Gap | | | | | | | |
| $Z_r$ | 0.0852*** | 0.0626*** | 0.0832*** | 0.0715*** | 0.0826*** | 0.0708*** | 0.0715*** | 0.0826*** |
| | (20.57) | (36.93) | (36.58) | (51.69) | (50.61) | (33.60) | (50.25) | (49.07) |
| *Panel D: IV - 2nd St, Dep. Var:* | Ln(Out Migrant Returns/Non-migrant Returns) | | | | | | | |
| Funding Gap | -0.360** | 0.467*** | 0.217*** | 0.0398 | 0.310*** | 0.254*** | 0.0411 | 0.365*** |
| | (-2.35) | (8.10) | (6.68) | (0.70) | (6.68) | (6.99) | (0.73) | (8.17) |
| *Panel E: IV - 2nd St, Dep. Var:* | Ln(Out Migrant AGI/Non-migrant Returns) | | | | | | | |
| Funding Gap | -0.410*** | 0.671*** | 0.508*** | -0.125 | 0.406*** | 0.175*** | -0.0929 | 0.450*** |
| N | 773752 | 773428 | 751696 | 751696 | 751696 | 751696 | 744009 | 744003 |
| Destination, Origin County-FE | | Y | | | | | | |
| Destination × Origin County-FE | | | Y | Y | Y | Y | Y | Y |
| Origin State × Year-FE | | | | Y | | | | |
| Destination State × Year | | | | | Y | | | |
| Dest. Reg. × Org. Reg. × Yr-FE | | | | | | Y | | |
| Origin County × Year-FE | | | | | | | Y | |
| Destination County × Year-FE | | | | | | | | Y |
| Controls: | Ln(Destination Rent/Origin Rent), Ln(Destination Real Wage/Origin Real Wage) | | | | | | | |

$t$ statistics in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 8: County-wise Net Migration: Sample Restrictions

This table reports regressions of net-out migration of tax returns from a county on its defined benefit funding ratio, controlling for local time-varying controls and county and time fixed effects. Column (1) shows the first stage where the instrument $Z_r$ is the cumulative excess return relative to the 2001 investment return assumption. Columns (2) and (3) show the regressions with the net outflow of tax returns relative to non-migrant tax returns as the dependent variable. Columns (4) and (5) shows the net outflow of adjusted gross income relative to non-migrant returns. Columns (2) and (4) show the OLS version of the regressions whereas columns (3) and (5) show the second stage with funding ratio of the county instrumented with the return instrument. Each specification contains the following controls: lagged log rents, real wages, property tax-to-household income ratio, top marginal state income tax rates, state debt to revenue, county-GDP growth and percentage of state's House of Representatives congressmen that belong to the Republican party. Panel A restricts the sample to 2001-2019 as robustness to avoid pandemic mobility. Panel B drops 2014-2016 to alleviate methodological concerns in IRS migration data raised by DeWaard et al. (2022). Standard errors are clustered at the county-level.

Panel A: Restricting Sample to 2019

| | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| | Funding Ratio | $\frac{\text{Net Out Migrant Returns}}{\text{Non-Migrant Returns}}$ | | $\frac{\text{Net Out Migrant AGI}}{\text{Non-Migrant Returns}}$ | |
| | First-Stage | OLS | IV | OLS | IV |
| $Z_r$ | 0.0844*** | | | | |
| | (12.38) | | | | |
| Funding Ratio$_{t-1}$ | | 0.00134 | -0.00505 | 0.161*** | -1.268*** |
| | | (1.01) | (-0.56) | (2.74) | (-2.73) |
| N | 50219 | 47960 | 47960 | 47960 | 47960 |
| R-Sq | 0.824 | 0.259 | - | 0.380 | - |
| F-Stat | 166.4 | | | | |
| FE | County and Year | | | | |
| Controls | Rents, Wages, S&L Taxes, Debt/Rev, County-growth, Crime, Political | | | | |

Panel B: Dropping 2014-2016 data

| | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| | Funding Ratio | $\frac{\text{Net Out Migrant Returns}}{\text{Non-Migrant Returns}}$ | | $\frac{\text{Net Out Migrant AGI}}{\text{Non-Migrant Returns}}$ | |
| | First-Stage | OLS | IV | OLS | IV |
| $Z_r$ | 0.0570*** | | | | |
| | (9.89) | | | | |
| Funding Ratio$_{t-1}$ | | -0.000583 | -0.0239** | -0.0220 | -1.635* |
| | | (-0.55) | (-2.44) | (-0.24) | (-1.79) |
| N | 47055 | 43973 | 43973 | 43973 | 43973 |
| R-Sq | 0.798 | 0.368 | | 0.414 | |
| F-Stat | 184.0 | | | | |
| FE | County and Year | | | | |
| Controls | Rents, Wages, S&L Taxes, Debt/Rev, County-growth, Crime, Political | | | | |

$t$ statistics in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 9: Leave-one-out-robustness

This table reports a leave-one-out robustness check of the 2SLS regressions in table 2. Each row is a separate second-stage regression of county-level migration on the funding ratio, instrumented by cumulative abnormal returns relative to the 2001 benchmark, excluding one Census region at a time. Column (1) uses the ratio of net out-migrant returns to non-migrant returns as the dependent variable, while column (2) replaces it with net out-migrant AGI to non-migrant returns.

| Region Left Out | (1) Net Migrant Returns / Non Migrant Returns | (2) Net Migrant AGI / Non Migrant Returns |
|---|---|---|
| East North Central (IL, IN, MI, OH, WI) | -0.0192* | -1.499* |
| | (-1.88) | (-1.82) |
| East South Central (AL, KY, MS, TN) | -0.0298** | -2.952*** |
| | (-2.06) | (-2.67) |
| Middle Atlantic (NJ, NY, PA) | -0.0198 | -1.379 |
| | (-1.49) | (-1.39) |
| Mountain (AZ, CO, ID, MT, NV, NM, UT, WY) | -0.0147 | -2.797*** |
| | (-0.86) | (-2.90) |
| New England (CT, MA, ME, NH, RI, VT) | -0.0196* | -2.173*** |
| | (-1.95) | (-2.82) |
| Pacific (AK, CA, HI, OR, WA) | -0.00900 | -1.020 |
| | (-1.04) | (-1.51) |
| South Atlantic (DE, DC, GA, MD, NC, SC, VA, WV) | -0.0171*** | -1.279** |
| | (-2.66) | (-2.46) |
| West North Central (IA, KS, MN, MO, NE, ND, SD) | -0.0340*** | -1.668 |
| | (-3.45) | (-1.45) |
| West South Central (AS, LA, OK, TX) | -0.0226** | -2.165** |
| | (-2.01) | (-2.50) |

$t$ statistics in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 10: Post-SALT net-outflow

This table estimates the net outflow of individuals following the $10,000 SALT deduction cap. The dependent variable, from IRS migration files, is the ratio of net out-migrants (outflows minus inflows) to total taxpayers in a state. Column (1) regresses this measure on the increase in tax liability for top-AGI taxpayers from NBER's TAXSIM. Columns (2)–(3) interact the Post-SALT Cap dummy (post-2018) with the 2017 top-AGI share. Controls include lagged rents, real wages, house price growth, crime, job openings, debt-to-revenue, and public school capital expenditure. All models have year fixed effects, and column (3) also adds state fixed effects. Panel A covers 2017–2022; Panel B covers 2017–2019.

Panel A: Sample Ranging 2017-2022

| | (1) | (2) | (3) |
|---|---|---|---|
| | Net Out-Migration/Total Tax-Filers | | |
| ΔTax Liability | 0.187** | | |
| | (2.16) | | |
| Share of Filers in Top Income Bracket | | -0.162 | |
| | | (-1.30) | |
| Post-TCJA × Share of Filers in Top Income Bracket | | 0.101*** | 0.0694*** |
| | | (3.47) | (2.84) |
| N | 246 | 246 | 245 |
| R-Sq | 0.279 | 0.293 | 0.921 |
| Within R-Sq | 0.274 | 0.288 | 0.154 |
| FE | Year | Year | State and Year |

Panel B: Sample Ranging 2017-2019

| | (1) | (2) | (3) |
|---|---|---|---|
| | Net Out-Migration/Non-Migrants | | |
| ΔTax Liability | 0.183* | | |
| | (1.70) | | |
| Share of Filers in Top Income Bracket | | -0.188* | |
| | | (-1.70) | |
| Post-TCJA × Share of Filers in Top Income Bracket | | 0.0522** | 0.0378* |
| | | (2.58) | (1.91) |
| N | 148 | 148 | 147 |
| R-Sq | 0.248 | 0.263 | 0.959 |
| Within R-Sq | 0.242 | 0.257 | 0.106 |
| FE | Year | Year | State and Year |

$t$ statistics in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 11: Growing Alternative Allocation with Migration Risk - Table for Figure 15

This table presents the regression results for pension fund allocations in response to migration risk. Data on pension allocations, and pension funding information are from the Public Pension Data. I estimate migration betas as the response of net migration to increases in the top marginal tax rate, controlling for rents, wages and other economic confounders. I control for pension contribution rates, dummies for change in assumptions, and growth in pension payroll along with pension and year fixed effects.

| | Allocation to Alternative Asset Classes | | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|
| | 0 | +1 year | +2 years | +3 years | +4 years | +5 years | +6 years | +7 years | +8 years | +9 years | +10 years |
| Underfunded Ratio$_{t-1}$ | -0.0255** | -0.0231 | -0.0176 | -0.00623 | -0.0177 | -0.0202 | -0.0487* | -0.0543** | -0.0698*** | -0.0908*** | -0.106*** |
| | (-2.33) | (-1.60) | (-0.96) | (-0.29) | (-0.69) | (-0.71) | (-1.75) | (-2.16) | (-2.74) | (-3.34) | (-3.65) |
| High Migration $\beta\times$ Underfunded Ratio$_{t-1}$ | -0.0146 | -0.00748 | -0.0101 | -0.0209 | -0.00183 | 0.0364 | 0.0734** | 0.0953*** | 0.140*** | 0.180*** | 0.175*** |
| | (-0.96) | (-0.38) | (-0.42) | (-0.76) | (-0.06) | (1.08) | (2.21) | (2.92) | (4.40) | (5.45) | (4.99) |
| N | 4237 | 4035 | 3832 | 3635 | 3435 | 3235 | 3035 | 2838 | 2645 | 2450 | 2237 |
| R-Sq | 0.890 | 0.832 | 0.787 | 0.760 | 0.750 | 0.758 | 0.772 | 0.783 | 0.791 | 0.796 | 0.801 |
| Within R-Sq | 0.599 | 0.378 | 0.202 | 0.0887 | 0.0198 | 0.0167 | 0.0326 | 0.0486 | 0.0584 | 0.0560 | 0.0395 |
| Pension Fund and Year FE | Y | Y | Y | Y | Y | Y | Y | Y | Y | Y | Y |
| | Allocation to Equities | | | | | | | | | | |
| Underfunded Ratio$_{t-1}$ | 0.0124 | 0.00334 | -0.00360 | -0.0101 | -0.00286 | 0.00205 | 0.0236 | 0.0339 | 0.0459** | 0.0506** | 0.0694*** |
| | (1.08) | (0.22) | (-0.21) | (-0.50) | (-0.12) | (0.08) | (0.94) | (1.47) | (2.06) | (2.24) | (2.94) |
| High Migration $\beta\times$ Underfunded Ratio$_{t-1}$ | 0.00687 | 0.0204 | 0.00944 | 0.00777 | -0.00409 | -0.0253 | -0.0596** | -0.109*** | -0.139*** | -0.144*** | -0.138*** |
| | (0.48) | (1.10) | (0.44) | (0.31) | (-0.15) | (-0.84) | (-2.07) | (-4.15) | (-5.45) | (-5.50) | (-5.00) |
| N | 3556 | 3365 | 3178 | 2989 | 2805 | 2623 | 2451 | 2273 | 2111 | 1944 | 1763 |
| R-Sq | 0.896 | 0.837 | 0.796 | 0.763 | 0.743 | 0.751 | 0.777 | 0.799 | 0.816 | 0.831 | 0.847 |
| Within R-Sq | 0.652 | 0.439 | 0.279 | 0.143 | 0.0444 | 0.0139 | 0.0212 | 0.0461 | 0.0635 | 0.0580 | 0.0623 |
| Pension Fund and Year FE | Y | Y | Y | Y | Y | Y | Y | Y | Y | Y | Y |
| | Allocation to Fixed Income Securities | | | | | | | | | | |
| Underfunded Ratio$_{t-1}$ | 0.0120 | 0.0322** | 0.0424** | 0.0446** | 0.0473** | 0.0463** | 0.0547*** | 0.0438*** | 0.0426** | 0.0399** | 0.0221 |
| | (1.05) | (2.25) | (2.57) | (2.55) | (2.57) | (2.51) | (2.94) | (2.59) | (2.44) | (2.33) | (1.21) |
| High Migration $\beta\times$ Underfunded Ratio$_{t-1}$ | -0.000617 | -0.0180 | -0.0118 | -0.0119 | -0.0117 | -0.0231 | -0.0375* | -0.0162 | -0.0141 | -0.0345 | -0.0366 |
| | (-0.05) | (-1.11) | (-0.62) | (-0.59) | (-0.56) | (-1.14) | (-1.79) | (-0.80) | (-0.69) | (-1.63) | (-1.62) |
| N | 3556 | 3365 | 3178 | 2989 | 2805 | 2623 | 2451 | 2273 | 2111 | 1944 | 1763 |
| R-Sq | 0.823 | 0.743 | 0.698 | 0.674 | 0.661 | 0.683 | 0.697 | 0.707 | 0.724 | 0.746 | 0.767 |
| Within R-Sq | 0.525 | 0.311 | 0.204 | 0.127 | 0.0632 | 0.0311 | 0.0193 | 0.00892 | 0.00645 | 0.00477 | 0.00943 |
| Pension Fund and Year FE | Y | Y | Y | Y | Y | Y | Y | Y | Y | Y | Y |

$t$ statistics in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

## A.1 Exploring Pension Reforms

Allowin Following a decline in funding status of state pensions, these bodies pursue reforms which include reducing pension benefits, increasing rates of contributions, changing defined-benefit pension plans to direct contribution plans or to a hybrid tier-based system for new hires versus existing employees and retirees. Using Chat-GPT's o1 model, I track 180 pension reforms since 2005-2022 from the National Association of State Retirement Administrators (NASRA) reports, and document who they effect and what they stipulate. On average reforms often bundle several changes to pensions and are also staggered within a state over years.

Table 12: Reform Statistics

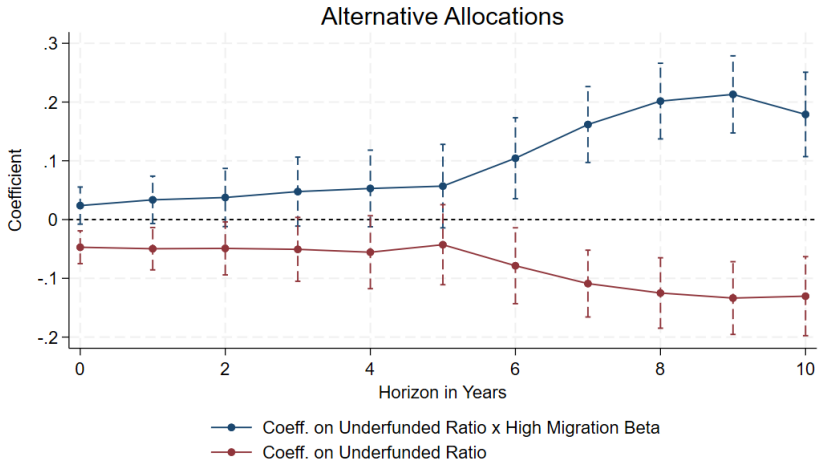
This table documents reform measures undertaken by state and local pensions during the period of 2005-2022 according to the National Association of State Retirement Administrators. Reforms incorporate legislative and executive changes adopted at the state government level. I classify reforms into the following classifications which are not mutually exclusive from each other. The first panel documents who the reform affects, whereas the bottom panel documents the specific policy stipulation of the reform.

| | Share | Count |
|---|---|---|
| **Who do reforms affect?** | | |
| 1(Affects Current Employees) | 0.72 | 180 |
| 1(Affects New Hires) | 0.72 | 180 |
| 1(Affects Retirees) | 0.26 | 180 |
| 1(Affects General Public Employees) | 0.68 | 180 |
| 1(Affects Teachers) | 0.52 | 180 |
| 1(Affects Police Officers) | 0.42 | 180 |
| 1(Affects Firefighters) | 0.39 | 180 |
| **Reform Description** | | |
| 1(Increased Employee Contribution) | 0.58 | 180 |
| 1(Reduced Pension) | 0.58 | 180 |
| 1(Increased Age or Service Requirement) | 0.48 | 180 |
| 1(New Tiers for Future Hires) | 0.42 | 180 |
| 1(Reduced Cost of Living Adjustment) | 0.37 | 180 |
| 1(Change in Final Annual Salary Calculation) | 0.36 | 180 |
| 1(Plan Funding Policy Changes) | 0.33 | 180 |
| 1(Plan Change: DB to DC/Hybrid) | 0.28 | 180 |
| 1(Reduced Benefits Multiplier) | 0.22 | 180 |
| 1(Change in Vesting Requirements) | 0.22 | 180 |
| 1(Change in Plan Design) | 0.21 | 180 |
| 1(COLA Mechanism Changes) | 0.21 | 180 |
| 1(Employer Contribution Changes) | 0.20 | 180 |
| 1(Early Retirement Penalties Change) | 0.19 | 180 |
| 1(Withdrawing Members Employer Contribution Policy) | 0.18 | 180 |
| 1(Retiree Medical/Healthcare Provisions) | 0.13 | 180 |
| 1(Disability Benefits Reforms) | 0.11 | 180 |
| 1(Survivor/Spousal Benefit Changes) | 0.08 | 180 |
| 1(Changes in Service plus Age Formulas) | 0.08 | 180 |
| 1(One-time Infusions or Supplement) | 0.06 | 180 |
| 1(Governance or Oversight Changes) | 0.02 | 180 |
| 1(Decreased Employee Contribution) | 0.01 | 180 |

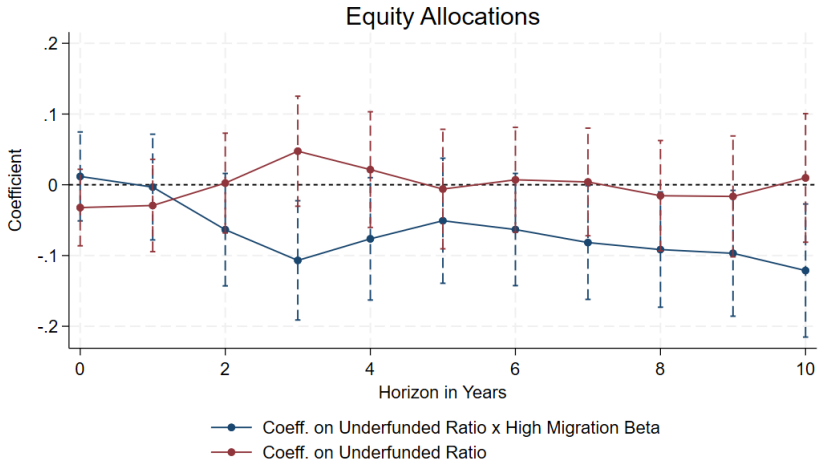
As we can see from table 12, reforms are mostly geared towards current empoyees and new hires, and are less likely to affect retirees. Among the employee-types, reforms are generally geared towards general state-level public employees and less likely for public safety workers such as police or firefighters.

Among the specific reforms measures, increasing employee contributions and reduced pension benefits are the broad categories which appear in 58% of cases. Increased service requirements and new tiers for future or new hires are the additional reforms that appear in over 40% of the cases. Cost of Living Adjustment (COLA) which effectively inflation index the last drawn salary of retirees, feature in over 37% of reforms and the way final annual salary is calculated also change in 36% of the cases. Changes from defined-benefit pension plans, where investment risk lies with the employer (which in this case are the respective adminstering governments), to direct contribution plans where employees bear the burden of investing their pension savings, are less common and happen in only 28% of cases. Legally mandated changes to employer contributions are less common and so

(a) Alternative Allocations




(b) Allocation to Equities




(c) Allocation to Fixed Income

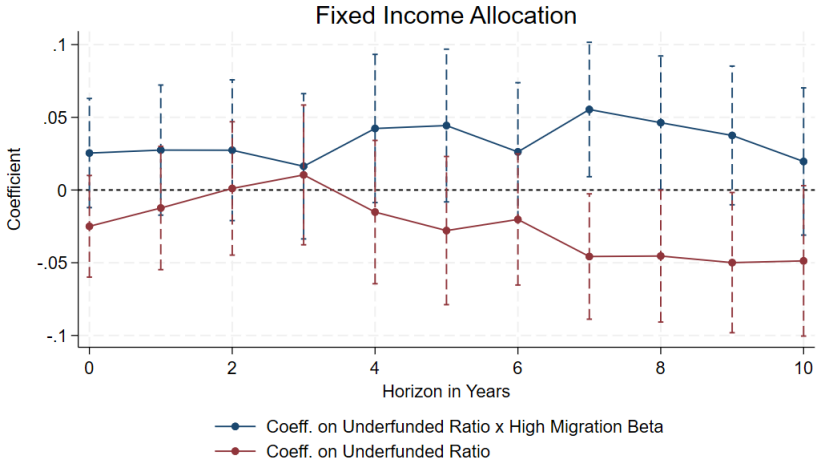
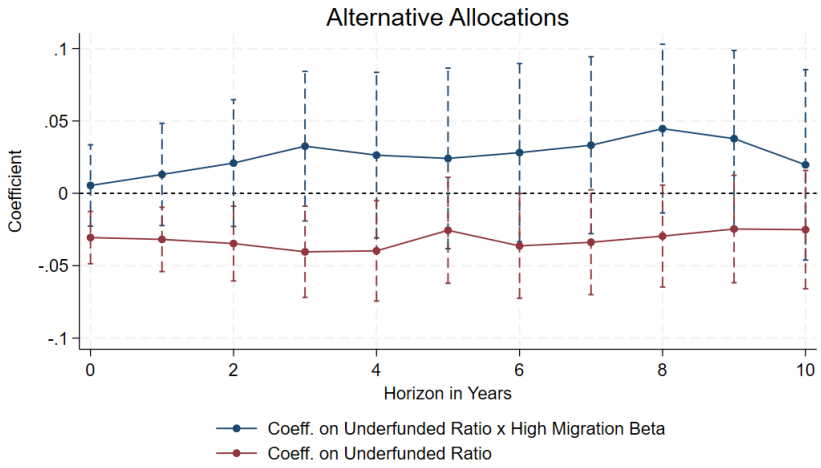


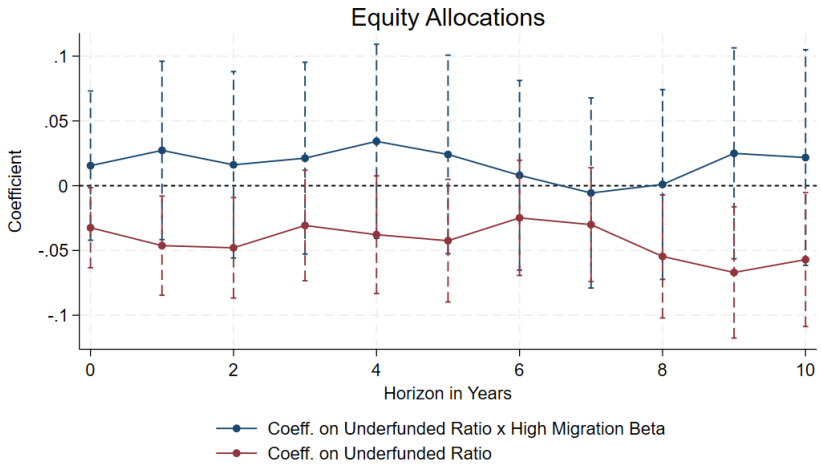

Figure 14: Local Projections: Allocation by Funding Status and Migration Risk - Post-SALT Net Outflows

These graphs plot the regression coefficients of the public pension allocations on the interaction of state migration betas and underfunded ratio. The navy-blue line represents the coefficient on the interaction of migration beta and underfunding, while the maroon line represents the coefficient on underfunded ratio. Panel (a) plots the coefficients for alternative allocations (b) plots the coefficients for allocations to equities and (c) plots the coefficient on fixed income securities. The axes represents the leads of the dependent variable in years.

(a) Alternative Allocations




(b) Allocation to Equities




(c) Allocation to Fixed Income

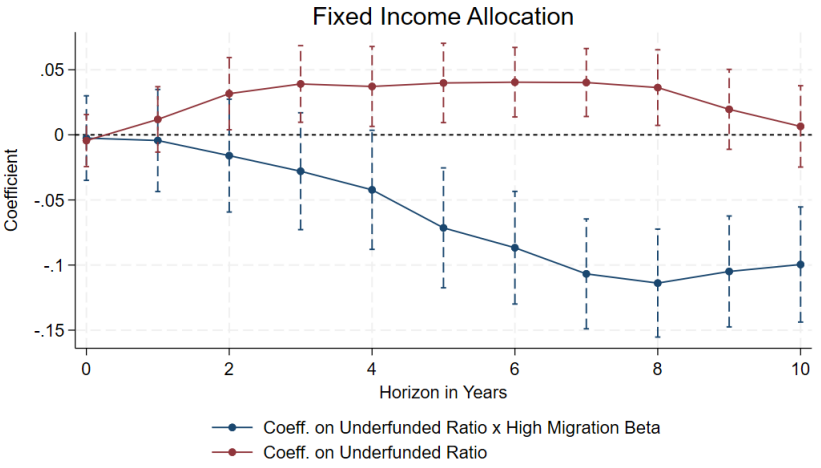



Figure 15: Local Projections: Allocation by Funding Status and Migration Risk - Moretti and Wilson (2017) Betas

These graphs plot the regression coefficients of the public pension allocations on the interaction of state migration betas and underfunded ratio. The navy-blue line represents the coefficient on the interaction of migration beta and underfunding, while the maroon line represents the coefficient on underfunded ratio. Panel (a) plots the coefficients for alternative allocations (b) plots the coefficients for allocations to equities and (c) plots the coefficient on fixed income securities. The axes represents the leads of the dependent variable in years.

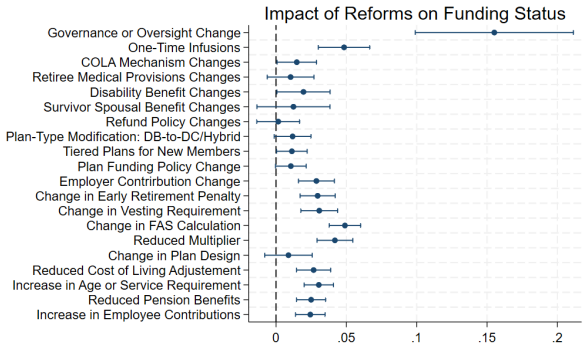
are other direct reforms to the pension plans themselves.

To test how different pension reforms affect funding status I run the following regression:

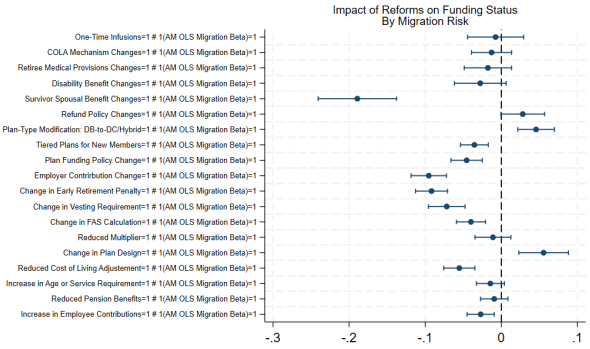
$$\text{Funding Ratio}_{i,t+1} = \beta \times 1(\text{Reform}_j)_{it} + \gamma X_{it}\tau_t + \eta_i + \epsilon_{it+1}$$

Where the dependent variable is the pension fund's funding ratio and the independent variable is the indicator for any reform enacted by the state and local governments. I control for indicators for changes in actuarial assumptions, lagged investment annual investment returns and payroll growth, along with pension plan and year fixed effects. I run this regression for each reform measure that I extract. Panel (a) Reports the baseline regression which show the relationship between reforms and funding status. In general, for after most reforms, except governance and oversight changes, funding status improves between 1-5 percentage points within each plan. Governance and oversight changes reflect reforms enacted by the state of Indiana, which has considerably improved its funding status by 17 percentage points after enacting such reforms.

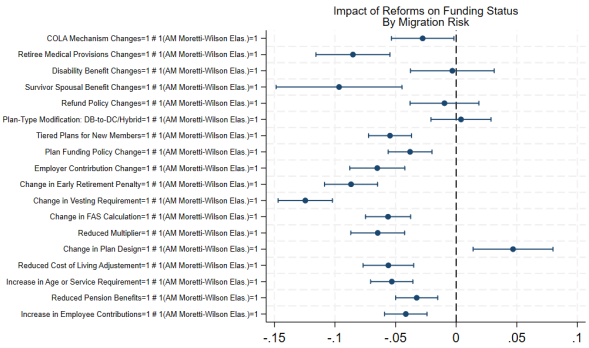
(a) Baseline




(b) By OLS Migration Betas




(c) By Moretti and Wilson (2017) Elasticity




(d) By Post-SALT cap Net Outflows

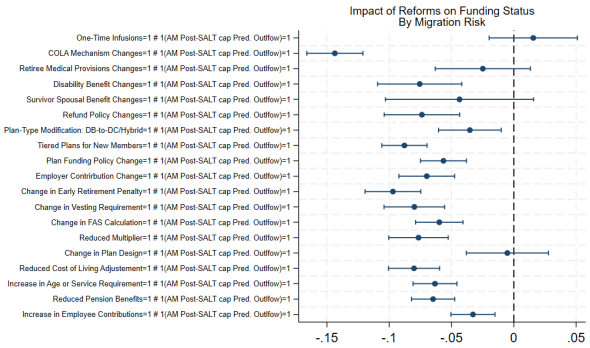



Figure 16: How do Reforms Impact Funding Status?

To test how pension funds in states which face higher migration risk fare when they enact pension reforms, I run interaction regressions across the three different measures.

$$\text{Funding Ratio}_{i,t+1} = \beta \times 1(\text{Reform}_j)_{it} \times 1(\text{High Migration Risk})_{s(i)} + \gamma X_{it}\tau_t + \eta_i + \epsilon_{it+1}$$

Panels (c), (d) and (e) present the results. Across all panels, I find that the coefficient on the interaction is negative and significant for states with higher migration betas - which suggests that pension funds located in these states see a relatively poorer funding performance when reforms are undertaken. For instance, while the baseline regression suggests a boost in the funding ratio of about 2 percentage points when a reform such as increasing employee contributions is enacted, panel (c) suggests that pension funds in states which lie in the above-median category of migration risk measured through the Moretti and Wilson (2017) effect, see a funding ratio lower by 4 percentage points. Overall this paints the picture, that most commonly undertaken reforms have a muted effect in states which face higher migration outflows.

## A.2 Risk-taking within Asset Classes: Evidence from 13-F Pensions

As equities still remain the largest investment class for public pensions, I first document how public pension fund managers in 13-F differ from private pension managers. Using the holdings data of large institutional investors from SEC-13-F filings, I use the list of pension fund managers identified by Koijen and Yogo (2019) to track portfolio performance. I merge the data with public pensions data. I identify 28 public pension managers[10] and 22 private pension fund managers[11] for which I was able to map balance sheet data from Worldscope, out of the 114 pension fund managers in the 13-F managers file.

| | Mean | Median | SD | Min | Max | Count |
|---|---|---|---|---|---|---|
| 13-F Pensions Balance Sheets | | | | | | |
| Funded Ratio | 0.81 | 0.81 | 0.18 | 0.00 | 4.27 | 2877 |
| 1(Public Pension) | 0.73 | 1.00 | 0.44 | 0.00 | 1.00 | 3089 |
| Ln(Pension Assets) | 19.03 | 18.22 | 2.90 | 7.14 | 25.41 | 2898 |
| Portfolio Return | 0.09 | 0.04 | 0.47 | -0.53 | 15.44 | 2482 |
| Annual Portfolio Volatility | 0.16 | 0.13 | 0.12 | 0.01 | 0.88 | 383 |
| Public Pension Portfolio Return | 0.09 | 0.04 | 0.48 | -0.38 | 15.44 | 1804 |
| Private Pension Portfolio Return | 0.10 | 0.04 | 0.46 | -0.53 | 5.88 | 678 |
| Number of Stocks | 1232.90 | 1008.00 | 960.90 | 1.00 | 4497.00 | 2482 |
| Number of Stocks of Public Pensions | 1575.78 | 1550.00 | 890.28 | 1.00 | 4497.00 | 1804 |
| Number of Stocks of Private Pensions | 320.57 | 137.00 | 354.97 | 1.00 | 1564.00 | 678 |

Table 13: Summary Statistics for 13-F Pensions

I then plot funding status by pension manager type in figure 17 panel (a). We can see that both private and public pension managers see a decline in funding ratios post-2008 recession. However, we can see a divergence in funding status, since 2016, where private pension funds have recovered near full funding in 2022, but public pensions' funding status remains relatively flat. One caveat is that the gap could be due mechanical difference in discounting, where private pensions could anchor their discount rates to the prevailing policy rates, while public pensions have used past or expected investment returns as discount rates, which may decouple funding status with interest rate cycles, and also understate the level of underfunding for public pensions (Giesecke and Rauh (2023)). Panel (b) plots the quarterly value-weighted portfolio returns of the public and private pensions since 2007Q4. Panel (c) tracks the annual portfolio volatility. There does not seem to be a remarkable difference in group-wise performance. Those this may arise due to the sampling bias - large 13-F public pensions are not qualitatively different in their investment performance relative to their private counterparts. I therefore look within the portfolio risk heterogeneity of public pension managers by their state's tax migration risk group.

### A.2.1 Portfolio Risk Heterogeneity in Public Pensoins by Migration Risk

Do public pensions based in states with high tax mobility carry more risk on their equity portfolios? In order to test this, I regress the annual portfolio volatility of public pension managers against their underfunded ratio and migration risk measure. Similar to table 3, the I use the three measures of migration response to tax increases interacted with the underfunded ratio of the pension fund.

$$Y_{it+1} = \beta \times 1(\text{High Migration Risk }) \times \text{Underfunding}_{i,t} + \text{Underfunding}_{i,t} + \Gamma X_{it} + \theta_t + \kappa_i + \varepsilon_{it+1}$$

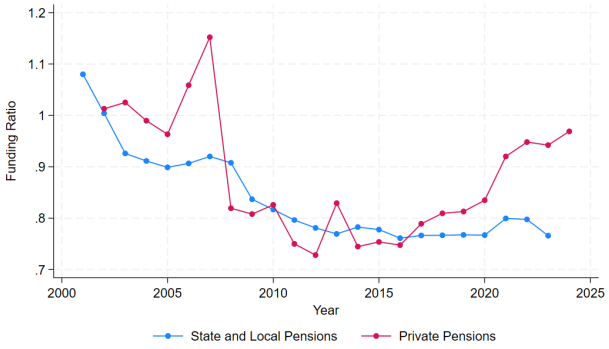
Where controls include dummies for changes in pension plan's actuarial assumptions and contribution growth, along with manager and year fixed effects. The results are displayed in table 14

Across all columns underfunded ratio is positively correlated with portfolio volatility. We can see that in column (1), the interaction with OLS state tax beta produces a negative coefficient for the same level of underfunded relative to baseline - meaning managers of states sorted by the OLS migration beta risk cut equity portfolio volatility as they move to alternative asset classes. Columns (2) and (3) however suggest an opposite result where the coefficient on the interaction terms are positive and weakly significant for column (3). This reflects the different rankings accorded to states by different beta estimates. However given, the statistically insignificant results, high migration risk do not have a different equity portfolio volatility relative to their low

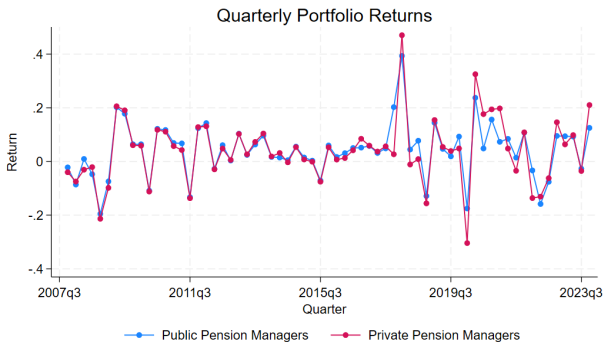
[10] Public pension managers could refer to the state investment boards which manage the investment of multiple constituent state and local level pension funds. The 25 public pension managers represent 41 public pension funds in the public plans data.

[11] These, for example, are United States Steel and Carnegie Pension Fund, Dow' Employees Pension or Sears Holdings Pension Fund which consists of frozen defined benefit, defined contribution or multi-employer retirement benefits, and non-contributory defined benefit pension plans

(a) Funding Status by Manager Type




(b) Quarterly Portfolio Returns




(c) Riskiness of Investments

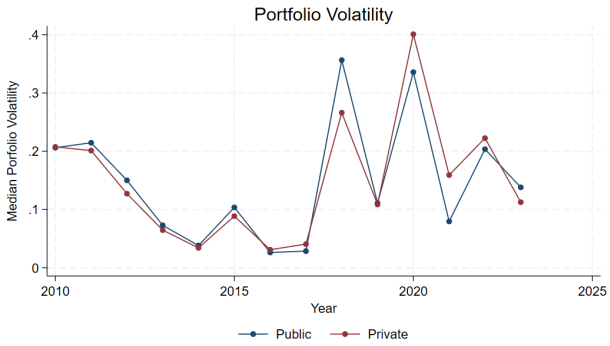



Figure 17: Public vs. Private Pension Managers

These figures plot the relative performance of public pensions relative to private pension fund managers identified in 13-F and Worldscope data. Panel (a) plats the assets-to-liabilities ratios. Public pension assets and liabilities come from the Public Plans Data from Center for Retirement Research. I measure private pension assets as the ratio of their actuarial assets and accumulated benefit obligations made available through Worldscope. Panel (b) plots the value-weighted portfolio returns of public and private managers from 13-F. Panel (c) plots the median annual portfolio volatility of the respective manager groups.

## Table 14: Equity Portfolio Volatility

This table shows regressions of annual portfolio volatility in the 13-F equity investments of public pension managers matched using the list of pension funds given by Koijen and Yogo (2019) on underfunding (1-funding ratio) interacted with a migration risk indicator. Column (1) uses the 1(Above Median OLS Migration Beta) indicator, column (2) uses the 1(Above Median Moretti and Wilson (2017) Elasticity) and column (3) uses 1(AM Post-SALT Cap Predicted Outflows) as the migration risk indicators respectively. All specifications include controls for lagged manager size, contribution growth and dummy for actuarial assumption changes along with manager and year fixed effects.

| | (1) | (2) | (3) |
|---|---|---|---|
| | Annual Portfolio Volatility | | |
| Underfunded Ratio$_{t-1}$ (UR) | 0.360* | 0.0710 | 0.126* |
| | (1.81) | (0.89) | (1.89) |
| 1(AM OLS Mig. Beta) × L.(lastnm) UR | -0.325* | | |
| | (-1.78) | | |
| 1(AM Moretti and Wilson (2017) Elas) × UR | | 0.183 | |
| | | (1.04) | |
| 1(AM Post-SALT CAP Pred. Outflows) × UR | | | 0.0352 |
| | | | (0.29) |
| N | 305 | 305 | 305 |
| R-Sq | 0.864 | 0.860 | 0.859 |
| FE | Plan and Year | | |

$t$ statistics in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

migration risk counterparts. Notwithstanding the sampling bias, this means that public pensions in states with higher migration risk are more likely to pivot to higher risk asset across asset classes than within.

Table 15: Evidence from Survey of Consumer Expectations

This regressions reports results from the Federal Reserve Bank of New York's Survey of Consumer Expectations. Columns (1), (2) and (3) have an indicator for whether the survey respondent thinks that the their federal, state and local tax liability would go up in the next 12 months. Columns (4), (5) and (6) ask how much tax rates are about to change in the next 12 months, in percentage points. The main independent variable is the respondent's state funding ratio. Controls include Log Real Wage, Log Rent, Log Top-Marginal Income Tax Rate, (%) of Republican congressman from the state of the respondent, and the State's Debt-to-Revenue ratio. Columns (1) and (4) do not absorb any fixed effects. Columns (2) and (5) control for education category, age category, region, household income category and year fixed effects. Columns (3) and (6) include all aforementioned fixed effects interacted with year. Standard errors are clustered by state.

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | 1(Tax Increase in Next 12 months) | | | Exp Pct Tax Inc in Next 12 months) | | |
| Funding Ratio | -0.0158 | -0.0734* | -0.0669* | -0.574 | -1.589 | -1.662* |
| | (0.0461) | (0.0366) | (0.0384) | (1.186) | (1.028) | (0.963) |
| N | 24804 | 24479 | 24479 | 24804 | 24479 | 24479 |
| R-Sq | 0.000738 | 0.0344 | 0.0422 | 0.000906 | 0.00682 | 0.0228 |
| Educ, Reg, Inc, Age, Year-FE | N | Y | N | N | Y | Y |
| Educ-Yr, Reg-Yr, Inc-Yr, Age-Yr FE | N | N | Y | N | N | Y |
| Controls | Wage, Rent, Inc Tax, Pol Leaning, Debt/Rev | | | | | |

Standard errors in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

# B Model Appendix

# Proofs

## B.1 Proof of Proposition 2

Assuming that the government conjectures a growth rate of $g_z$, from the optimal contribution equation 24, we can differentiate beginning of period contributions w.r.t $\bar{f}_s$ that yields:

$$\frac{\partial c^*_{s,1}}{\partial \bar{f}_s} = -\frac{l_s(1+r_f)}{(1+g_z)^2+(1+r_f)} < 0 \tag{34}$$

From the budget rules $\tau^*_{s,1} = x_{s,1} + c^*_{s,1}$, this implies lower taxes.

Population growth in the second period is dependent on period one taxes. Differentiating the population share equation 6 we get

$$\frac{\partial N_{s,2}}{\partial \tau_{s,1}} = \frac{1}{\sigma_v} N_{s,2}(1-N_{s,2}) \frac{\partial v_{s,1}}{\tau_{s,1}} = -\frac{N_{s,2}(1-N_{s,2})}{\sigma_v} B_s (N_{s,1} H^{-\phi_s}(1-\tau_{s,1}))^{(\alpha-1/1+\phi_s)} \frac{\alpha+\phi_s}{1+\phi_s} w^{\phi_s+\alpha/1+\phi_s} \frac{\partial c^*_{s,1}}{\partial \bar{f}_s} > 0 \tag{35}$$

Which is positive and proves that population share in period 2 is increasing in the.

## B.2 Proof of Proposition 3

The derivative of optimal period-one contributions w.r.t. conjectured growth $g_s$ is given by:

$$\begin{aligned}\frac{\partial c^*_{s,1}}{\partial g_s} = & \frac{1+\Omega x_{s,2} - 2(1+\Omega x_{s,1})(1+g_s)}{\Omega[(1+g_s)^2+1+r_f]} \\ & - \frac{(1+\Omega x_{s,2})(1+g_s) + \Omega l_s \left(1 - \bar{f}_s(1+r_f) - \frac{(\mu-r_f)^2}{\gamma\xi\sigma^2}\right) - (1+x_{s,1}\Omega)(1+g_s)^2}{(\Omega[(1+g_s)^2+1+r_f])^2} \times 2\Omega(1+g_s)\end{aligned}$$

Assuming $g_s < 1$ and $x_{s,2} - 2x_s, 1 < 1$ as $x_s$ is government expenditure to state GDP, we van sign the derivative as $\frac{\partial c_{s,1}}{\partial g_s} < 0$. Therefore states that over-estimate growth would under-contribute relative to actual growth rates. Under-contribution leads to higher risk-taking by pension fiduciaries through equation 23.

If state governments conjecture a constant $g_z$, states with relatively worse funding status would fix higher taxes (Proposition 2) and hence see a population growth lower than better funded counterparts. Thus worse funded states would have $g_s < g_z$. This implies $c_{s,1}(g_s) > c_{s,1}(g_z)$. Through equation 23 this implies a higher risky investment share as $\theta(c_{c,1}(g_z)) > \theta(c_{s,1}(g_s))$ $\square$.

## B.3 Derivation: Optimal period 1 contributions:

Taking optimal investment rule (equation 23) from the fiduciary as given we can get the following formulation:

$$\begin{aligned}\min_{c_{s,1}} \; & (c_{s,1}+x_{s,1}) + \frac{\Omega}{2}(c_{s,1}+x_{s,1})^2 + \beta \left[\frac{l_s - (a_s+c_{s,1})(1+r_f) - \frac{(\mu-r_f)^2 l_s}{\gamma\xi\sigma^2}}{(1+g_s)} + x_{s,2}\right] \\ & + \frac{\beta\Omega}{2}\frac{(\mu-r_f)^2 l_s^2}{\sigma^2\xi^2\gamma^2(1+g_s)^2} + \frac{\beta\Omega}{2}\left[\frac{l_s - (a_s+c_{s,1})(1+r_f) - \frac{(\mu-r_f)^2 l_s}{\gamma\xi\sigma^2}}{(1+g_s)} + x_{s,2}\right]^2\end{aligned}$$

FOC:

$$0 = 1 + \Omega(c_{s,1}+x_{s,1}) - \beta\frac{1+r_f}{1+g_s} - \beta\Omega\left[\frac{l_s - (a_s+c_{s,1})(1+r_f) - \frac{(\mu-r_f)^2 l_s}{\gamma\xi\sigma^2}}{(1+g_s)} + x_{s,2}\right]\frac{1+r_f}{(1+g_s)}$$

Taking $\beta(1+r_f) = 1$, we can re-arrange to get:

$$(1+g_s)(1+\Omega(c_{s,1}+x_{s,1})) = 1 + \Omega\left[\frac{l_s - (a_s+c_{s,1})(1+r_f) - \frac{(\mu-r_f)^2 l_s}{\gamma\xi\sigma^2}}{(1+g_s)} + x_{s,2}\right]$$

Re-arranging gives the following:

$$c^*_{s,1} = \frac{(1+g_s)(1+\Omega x_{s,2}) - (1+g_s)^2(1+\Omega x_{s,1}) + \Omega l_s \left[1 - \bar{f}(1+r_f) - \frac{(\mu - r_f)^2}{\sigma^2 \xi \gamma}\right]}{\Omega[(1+g_s)^2 + 1 + r_f]}$$

## B.4 Simulation Algorithm

I solve for state-level contributions and growth rates $\{c_{s,1},\, g_s\}_{s=1}^{S}$ that jointly satisfy the per-state equilibrium conditions

$$R_s(c_{s,1}, g_s;\, \text{parameters}) = 0, \qquad s = 1, \ldots, S.$$

Starting from feasible guesses $(c_{s,1}^{(0)}, g_s^{(0)})$, the routine alternates by block:

1. **Contribution step.** Holding $g_s^{(t)}$ fixed, solve the scalar equation

$$R_s\Big(c, g_s^{(t)}\Big) = 0$$

for each state $s$ to obtain $c_{s,1}^{(t+1)}$ (within simple bounds $[\underline{c}, \overline{c}]$).

2. **Growth step.** Holding $c_{s,1}^{(t+1)}$ fixed, solve

$$R_s\Big(c_{s,1}^{(t+1)}, g\Big) = 0$$

for each state $s$ to obtain $g_s^{(t+1)}$ (within bounds $[\underline{g}, \overline{g}]$).

After the $g$-update, I enforce aggregate feasibility via the linear constraint

$$\sum_{s=1}^{S} g_s^{(t+1)} N_s = 0 \quad \Longleftrightarrow \quad \sum_{s=1}^{S} (1 + g_s^{(t+1)}) N_s = 1,$$

along with the same box bounds. I iterate these two steps until the maximum absolute residual falls below a tight tolerance,

$$\max_s \left| R_s(c_{s,1}^{(t+1)}, g_s^{(t+1)}) \right| < \texttt{tol}.$$

Post-convergence objects are then computed from the fixed point, e.g.

$$\theta_s \;=\; \frac{(\mu - r_f)\, l_s}{\big(a_s + c_{s,1}\big)\, \sigma^2\, \xi\, \gamma}.$$

Conceptually, this is an alternating fixed-point routine that reduces a high-dimensional system to a sequence of independent one-dimensional solves at the state level.

## B.5 Extension: Growth-internalizing Political Executive

If the government internalizes population mobility response of taxes (taking $g_z = 0$ without loss of generality), the resulting first-order conditions would be as follows:

$$\min_{c_{s,1}} \quad F(c_{s,1} + x_{s,1}) + \beta \mathbb{E}\left[F\left(\overbrace{\frac{l_s - (a_s + c_{s,1})R_s}{(1+g_s)}}^{\text{Inter-temporal Tax Smoothing}} + x_{s,2}\right)\right]$$

$$\Longrightarrow \quad \underbrace{F'(\tau_1)}_{\text{Current Tax Cost}} + \overbrace{\beta \mathbb{E}\left[\frac{R_s}{(1+g_s)} F'(\tau_2)\right]} - \underbrace{\beta \mathbb{E}\left[\frac{l_s - (a_s + c_{s,1})R_s}{(1+g_s)^2} \frac{\partial g_s}{\partial \tau_1} F'(\tau_2)\right]}_{\text{Migration Feedback}} = 0$$

$$\implies 1+\Omega\left[\frac{l_s-(a_s+c_{s,1})(1+r_f)}{1+g_s}+\frac{(\mu-r_f)^2 l_s}{\sigma^2\gamma\xi(1+g_s)}+x_{s,2}\right]=$$

$$(1+g_s)\Big(1-\Omega(c_{s,1}+x_{s,1})\Big)+\beta B_s\frac{\phi_s+\alpha}{\sigma_v(1+\phi_s)}\cdot w^{\frac{\phi_s+\alpha}{1+\phi_s}}\cdot\left(\frac{(1-c_{s,1}-x_{s,1})N_{s,1}(1-\alpha)}{H_s^{\phi}}\right)^{\frac{\alpha-1}{\phi_s+1}}\times$$

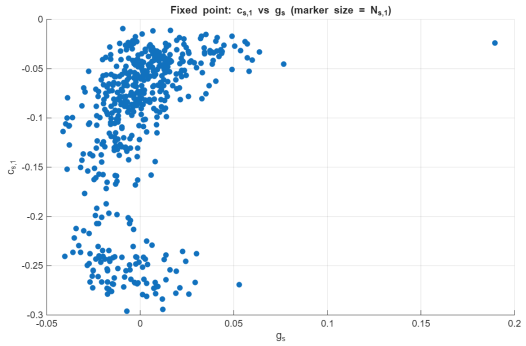
$$\left[l_s-(a_s+c_{s,1})(1+r_f)-\frac{(\mu-r_f)}{\sigma^2\gamma\xi}+\frac{\Omega l_s^2(\mu-r_f)^2}{(1+g_s)\sigma^2\gamma^2\xi^2}\right]+\dots$$

$$-\Omega\left[\frac{l_s-(a_s+c_{s,1})(1+r_f)}{1+g_s}-\frac{(\mu-r_f)^2 l_s}{\sigma^2\gamma\xi(1+g_s)}+x_{s,2}\right]\times$$

$$\left[l_s-(a_s+c_{s,1})(1+r_f)-\frac{(\mu-r_f)^2}{\sigma^2\gamma\xi}\right]$$

Using fixed-point algorithm I get the following simulation result in figure 18. I find that the relationship between first-period contributions and growth rates dis-associates as policymakers internalize the effects of migration within their optimization. This is exemplified in simulation graph (a) which uses the same parameters as the main estimation, except adding the migration feedback into the fixed-point update. The second panel (b) shows the lack of relationship between risky asset investment shares and growth rates.

(a) Contributions vs Growth




(b) Risky Asset Share vs Growth

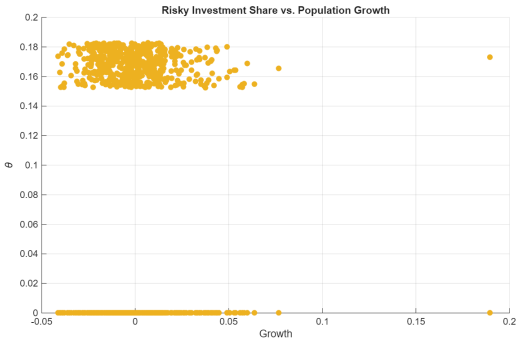



Figure 18: Model Predictions: When Executive Internalizes Migration

## B.6 Extension: Threshold based Payoff

The objective of this exercise is to model a fiduciary pay-off structure where fiduciaries target a threshold funding status in the end-period. Let's call this funding ratio as $\tilde{f}$. The fiduciary's payoff thus looks as follows:

$$\Pi(F)=k+\xi(f-\tilde{f})^{+} \tag{36}$$

Where $\xi$ is compensation linked to funding status remaining above a prespecified funding status level $\tilde{f}$. The payoff thus resembles a call option in the pension funding status. I assume a CARA utility form for pension fund managers which leads to the following optimization:

$$\max_{\theta}\mathbb{E}[-\exp(-\gamma(k+\xi(f-\tilde{f})^{+})] \tag{37}$$

Using the log normal form of the expectations, we get:

$$\max_{\theta}-\exp(-\gamma k)\left[\Phi(\kappa)+\exp(\gamma\xi(\tilde{f}-m)+\frac{1}{2}\gamma^2\xi^2 s^2)[1-\Phi(\kappa+\gamma\xi s)]\right] \tag{38}$$

Where $\kappa = \frac{\tilde{f}-m}{s}$ and $\Phi$ is the standard normal CDF.
The FOCs produce

$$\phi(\kappa)\left[\frac{\mu - r_f}{\theta\sigma} + \frac{\tilde{f}-m}{\theta^2 q\sigma}\right] = \exp\left(\gamma\xi(\tilde{f}-m) + \frac{1}{2}\gamma^2\xi^2 s^2\right) \times \\ \left[(\gamma^2\xi^2 q^2\sigma^2\theta - \gamma\xi q(\mu - r_f))(1 - \Phi(\kappa + \gamma\xi s) + \phi(\kappa + \gamma\xi s)\left(\frac{\mu - r_f}{\theta\sigma} + \frac{\tilde{f}-m}{\theta^2 q\sigma} - \gamma\xi q\sigma\right)\right] \tag{39}$$

I can solve for the optimal investment problem as a fixed point and then compare the optimal policy for the linear pay structure against the threshold pay structure. The result is presented in figure 19. As we can see,

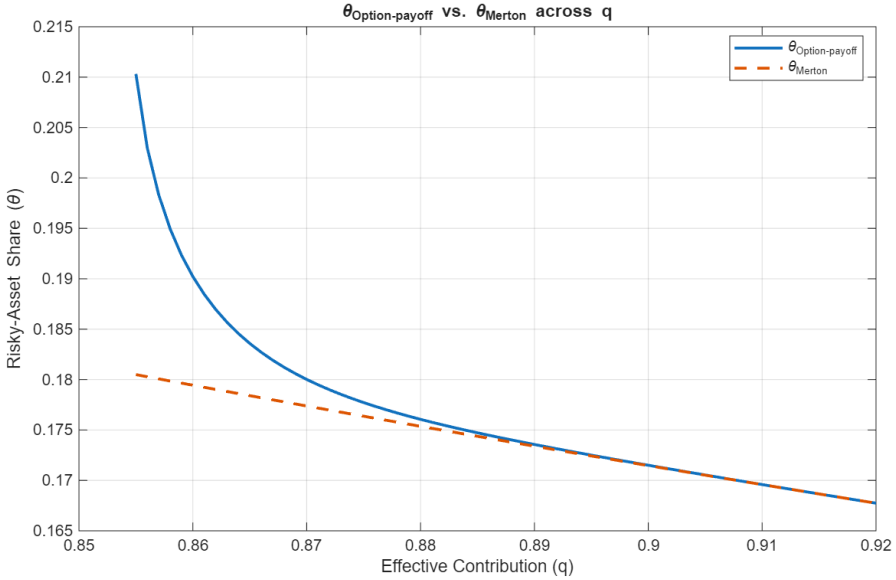



Figure 19: Investment Policy under Option-like payoff vs. Linear pay

This figure shows how optimal risky asset share varies under a linear payoff structure versus an option-like payoff structure where the fiduciaries get rewarded if the end of period funding is above a threshold $\tilde{f}$. I set the threshold at $\tilde{f} = 0.8$ and vary effective contributions which are $q = (a_s + c_{s,1})/l_s$ across a grid. I set parameters $\xi = 2, \gamma = 4, \mu - r_f = 0.04, \sigma = 0.18$. Option-payoff shares are shown in solid blue line, whereas the linear payoff shares are dashed red lines.

risky asset shares are very similar when effective contributions $q$ are considerably above the threshold of 0.8. Once effective contributions reach below 0.88, the risky asset share in the option payoff exponentially diverges from the risky share under the linear payoff.

## B.7 Derivation of Optimal Policy

The social planner's objective is given by:

$$\max_{\{\theta_s, \tau_{s,1}, \tau_2\}} V_{s,1}N_{s,1} + \delta\mathbb{E}(V_{s,2}N_{s,2}) \tag{40}$$

subject to period-by-period budget constraints:

$$\tau_{s,1}w_s N_{s,1} = X_{s,1} + \zeta_s L_s \qquad \tau_{s,2}w_s N_{s,2} = X_{s,2} + L_s - (A_s + \zeta_s L_s)(1 + r_f + \theta_{s,t}(\mu - r_f) + \theta_s\sigma\epsilon_t) \tag{41}$$

and endogenous household location shares:

$$N_{s,2} = \frac{\exp(v_{s,1}/\sigma_v)}{\sum_k \exp(v_{k,1}/\sigma_v)} \tag{42}$$

given public service or amenities provision function:

$$B_{s,t} = F(X_{s,t}/N_{s,t}) = \bar{B}_s(1 - \exp(-\psi_s(G_{s,t}/N_{s,t}))) \tag{43}$$

Thus this government is faced with a problem to optimize welfare by providing amenities which are an increasing function of the resources it raises over two periods.

**First-order conditions**

(i) FOC with respect to the statutory tax rate $\tau_{s,1}$

$$0 = \frac{\partial B_{s,1}}{\partial \tau_{s,1}} w_{s,1}\,(1-\tau_{s,1})\,r_{s,1}^{\alpha-1}\,N_{s,1} \;-\; B_{s,1}\,w_{s,1}\,\alpha\,r_{s,1}^{\alpha-1}\,N_{s,1} \;-\; B_{s,1}\,w_{s,1}\,(1-\tau_{s,1})\,(1-\alpha)\,r_{1,s}^{\alpha-2}\,\frac{\partial r_{1,s}}{\partial \tau_{s,1}} N_{s,1}$$
$$+\;\left[\frac{\partial E[B_{s,2}]}{\partial N_{s,2}}\,w_{s,2}\,(1-\tau_{s,2})\,r_{2,s}^{\alpha-1}\,N_{s,2} \;+\; \mathbb{E}[B_{s,2}]\,w_{s,2}\,(1-\tau_{s,2})(\alpha-1)\,\frac{\partial r_{2,s}}{\partial N_{s,2}}\,r_{2,s}^{\alpha-2}\,N_{s,2} \;+\; \mathbb{E}[V_2]\right]\delta\frac{\partial N_{s,2}}{\partial \tau_{s,1}}. \tag{44}$$

Where the first term of the LHS reflects the impact of raising labor income tax rates on today's amenities, the second term is the impact on today's disposable income, the third term is the impact on housing rents going down as disposable income goes down. The combined fourth terms relates to change in the state population tomorrow and it's knockdown affect on state's tax base which changes it's amenities, the change in rents as housing markets gets either more crowded or less congested, the third term is documenting the per-capita change in utility of households tomorrow. Taking the derivative of amenities w.r.t. taxes results in the folowing:

$$\begin{aligned} B_{s,1} \;&=\; \bar{B}_s\left[1-\exp\big(-\psi_s\,(\tau_{s,1}w_{s,1}-\zeta_s\,\frac{L_s}{N_{1,s}})\big)\right],\\ \Longrightarrow \frac{\partial B_{s,1}}{\partial \tau_{s,1}} \;&=\; \bar{B}_s\,\exp\Big[-\psi_s\,(\tau_{s,1}w_{s,1}-\zeta_s\,\frac{L_s}{N_{1,s}})\Big]\,w_{s,1}\psi_s \;=(\bar{B}_s-B_{s,1})w_{s,1}\psi_s \end{aligned}$$

Derivative of the housing–market rent $r_{s,1}$

$$\begin{aligned} r_{s,1} \;&=\; \left[\frac{N_{s,1}\,w\,(1-\tau_{s,1})(1-\alpha)}{H^{\phi_s}}\right]^{\frac{1}{1+\phi_s}}\\ \Longrightarrow\quad \frac{\partial r_{s,1}}{\partial \tau_{s,1}} \;&=\; -\,\frac{1}{1+\phi_s}\left[\frac{N_{s,1}\,w_{s,1}(1-\alpha)}{H^{\phi_s}}\right]^{\frac{1}{1+\phi_s}}(1-\tau_{s,1})^{-\frac{\phi_s}{1+\phi_s}} \;=\; \frac{-r_{1,s}}{(1+\phi_s)(1-\tau_{s,1})} \end{aligned} \tag{45}$$

Given log-normality, we can work with the parametric form of the expected value of $B_{s,2}$

$$\begin{aligned} \mathbb{E}[B_{s,2}] =\; & \bar{B}_s\left\{1-\mathbb{E}\Big[\exp\big(-\psi_s\,(\tau_{s,2}w_{s,2}+\frac{1}{N_{s,2}}(\underbrace{A_s+\zeta_s L_s}_{\tilde{A}_s})\,(1+r_f+\theta_s\,(\mu-r_f+\sigma\,\epsilon)-\,L_s)\big)\Big]\right\}\\ =\; & \bar{B}_s\left[1-\exp\Big(-\psi_s\big(\tau_{s,2}w_{s,2}+\tilde{A}_s\,(1+r_f+\theta_s(\mu-r_f))-L_s\big)/N_{s,2} \;+\; \tfrac{\psi_s(\tilde{A}_s)^2\sigma^2\theta_s^2}{2\,N_{s,2}^2}\Big)\right]. \\ =\; & \bar{B}_s\left[1-\exp\left(-\psi_s K_{s,1}+\psi_s^2\frac{K_{s,2}^2}{2}\right)\right] \end{aligned} \tag{46}$$

Where,

$$K_{s,1} \;=\; \tau_2\,w_{s,2} \;+\; \frac{\tilde{A}_s\,(1+r_f+\theta_s\,(\mu-r_f)) \;-\; L_s}{N_{s,2}}, \qquad K_{s,2} \;=\; \frac{\tilde{A}_s\,\theta_s\,\sigma}{N_{s,2}}.$$

Derivative of $\mathbb{E}[B_{s,2}]$ with respect to $N_{s,2}$

$$\frac{\partial \mathbb{E}[B_{s,2}]}{\partial N_{s,2}} \;=\; -\bar{B}_s\,\exp\Big[-\psi_s\,K_{s,1}+\tfrac{\psi_s^2K_{s,2}^2}{2}\Big]\Big[-\psi_s\,\frac{\partial K_{s,1}}{\partial N_{s,2}} \;+\; \psi_s^2\,K_{s,2}\,\frac{\partial K_{s,2}}{\partial N_{s,2}}\Big].$$

Derivatives of $K_{s,1}$ and $K_{s,2}$

$$\frac{\partial K_{s,1}}{\partial N_{s,2}} \;=\; -\frac{1}{N_{s,2}^2}\left[\tilde{A}_s\,(1+r_f+\theta_s(\mu-r_f))-L_s\right] \;=\; -\frac{K_{s,1}-\tau_2\,w_{s,2}}{N_{s,2}},$$

$$\frac{\partial K_{s,2}}{\partial N_{s,2}} \;=\; -\frac{A_s\,\theta_s\,\sigma}{N_{s,2}^2} \;=\; -\frac{K_{s,2}}{N_{s,2}}.$$

Plugging them back in

$$\frac{\partial \mathbb{E}[B_{s,2}]}{\partial N_{s,2}} \;=\; \bar{B}_s\,\exp\Big[-\psi_s\,K_{s,1} \;+\; \tfrac{\psi_s^2K_{s,2}^2}{2}\Big]\,\Big[-\psi_s\,\frac{K_{s,1}-\tau_{s,2}w_{s,2}}{N_{s,2}} \;+\; \psi_s^2\,\frac{K_{s,2}^2}{N_{s,2}}\Big]. \tag{47}$$

Derivative of the housing rent $r_{2,s}$ w.r.t. $N_{s,2}$

$$\frac{\partial r_{2,s}}{\partial N_{s,2}} \;=\; \frac{1}{1+\phi_s}\, N_{s,2}^{-\frac{\phi_s}{1+\phi_s}} \left[\frac{w\,(1-\tau_{s,2})(1-\alpha)}{H^{\phi_s}}\right]^{\frac{1}{1+\phi_s}}. \tag{48}$$

Expected period-2 indirect utility function

$$\mathbb{E}[V_2] \;=\; \bar{B}_s \left[1-\exp\!\big(-\psi_s\,(K_{s,1}-\tfrac{\psi_s K_{s,2}^2}{2})\big)\right] w_{s,2}\,(1-\tau_{s,2})\, r_{2,s}^{\alpha-1}. \tag{49}$$

Population response in period 2 to a change in tax rate $\tau_{s,1}$

$$\begin{aligned}
\frac{\partial N_{s,2}}{\partial \tau_{s,1}} \;&=\; \sigma_v^{-1} N_{s,2}\left(1-N_{s,2}\right)\frac{\partial v_1}{\partial \tau_{s,1}} = \\
&\sigma_v^{-1} N_{s,2}(1-N_{s,2})\left[\frac{\partial B_{s,1}}{\partial \tau_{s,1}}\, w_{s,1}\,(1-\tau_{s,1})\, r_{s,1}^{\alpha-1} \;-\; B_{s,1}\, w_{s,1}\, r_{s,1}^{\alpha-1} \;-\; B_{s,1}\, w_{s,1}\,(1-\tau_{s,1})\,(1-\alpha)\, r_{1,s}^{\alpha-2}\,\frac{\partial r_{1,s}}{\partial \tau_{s,1}}\right] \\
&= \sigma_v^{-1} N_{s,2}(1-N_{s,2})\Big[(\bar{B}_s - B_{s,1}) w_{s,1}^2 \psi_s (1-\tau_{s,1}) r_{1,s}^{\alpha-1} - B_{s,1} w_{s,1} r_{1,s}^{\alpha-1} \\
&+ B_{s,1} w_{s,1}(1-\tau_{s,1})(1-\alpha)\frac{r_{1,s}^{\alpha-1}}{(1+\phi_s)(1-\tau_{s,1})}\Big]
\end{aligned} \tag{50}$$

Substituting into the main first order conditions we get:

$$\begin{aligned}
0 = \Big[&(\bar{B}_s - B_{s,1}) w_{s,1}^2 \psi_s (1-\tau_{s,1}) r_{1,s}^{\alpha-1} N_{s,1} \\
&-\; B_{s,1}\, w_{s,1}\, r_{s,1}^{\alpha-1} N_{s,1} \\
&+\; B_{s,1}\, w_{s,1}(1-\tau_{s,1})(1-\alpha)\left(\tfrac{r_{s,1}^{\alpha-1}}{(1-\tau_{s,1})(1+\phi_s)}\right) N_{s,1} \\
&+\; \Big\{(\bar{B}_s - \mathbb{E}[B_{s,2}])\Big[-\psi_s\,\tfrac{K_{s,1}-\tau_{s,2}w_{s,2}}{N_{s,2}} + \psi_s^2\,\tfrac{K_{s,2}^2}{N_{s,2}}\Big]\Big\}\, w_{s,2}(1-\tau_{s,2})\, r_{2,s}^{\alpha-1} N_{s,2} \\
&-\; \mathbb{E}[B_{s,2}]\, w_{s,2}(1-\tau_{s,2})(1-\alpha)\left(\tfrac{r_{2,s}^{\alpha-1}}{1+\phi_s}\,\tfrac{1}{N_{s,2}}\right) N_{s,2} \\
&+\; \mathbb{E}[B_{s,2}]\, w_{s,2}(1-\tau_{s,2})\, r_{2,s}^{\alpha-1}\Big\}\delta\sigma_v^{-1}\left[N_{s,2}(1-N_{s,2})\right] \\
&\times\; \Big[(\bar{B}_s - B_{s,1}) w_{s,1}^2 \psi_s (1-\tau_{s,1}) r_{1,s}^{\alpha-1} \\
&-\; B_{s,1}\, w_{s,1}\, r_{s,1}^{\alpha-1} \\
&+\; B_{s,1}\, w_{s,1}(1-\tau_{s,1})(1-\alpha)\left(\tfrac{r_{s,1}^{\alpha-1}}{(1-\tau_{s,1})(1+\phi_s)}\right)\Big].
\end{aligned}$$

Re-arranging and simplifying:

$$\begin{aligned}
\implies 0 = &\overbrace{\Big[(\bar{B}_s - B_{s,1}) w_{s,1}^2 \psi_s (1-\tau_{s,1}) r_{1,s}^{\alpha-1} - B_{s,1}\, w_{s,1}\, r_{s,1}^{\alpha-1} + B_{s,1}\, w_{s,1}(1-\alpha)\,\tfrac{r_{s,1}^{\alpha-1}}{1+\phi_s}\Big]}^{T_1} \\
\times\; &\underbrace{\Big[N_{s,1} + \delta\sigma_v^{-1} N_{s,2}(1-N_{s,2})\Big((\bar{B}_s - \mathbb{E}[B_{s,2}])\Big[-\psi_s\,(K_{s,1}-\tau_{s,2}w_{s,2}) + \psi_s^2\,K_{s,2}^2\Big]\, w_{s,2}(1-\tau_{s,2})\, r_{2,s}^{\alpha-1} \\
&\text{- E[B}_{s,2}]\, w_{s,2}(1-\tau_{s,2})(1-\alpha)\,\tfrac{r_{2,s}^{\alpha-1}}{1+\phi_s} + \mathbb{E}[B_{s,2}]\, w_{s,2}(1-\tau_{s,2})\, r_{2,s}^{\alpha-1}\Big)\Big]}_{T_2}
\end{aligned} \tag{51}$$

This ends us giving us multiple roots as either $T_1 = 0$ or $T_2 = 0$. One candidate is simply $T_1 = 0$

$$\begin{aligned}
&(\bar{B}_s - B_{s,1}) w_{s,1}^2 \psi_s (1-\tau_{s,1}) r_{1,s}^{\alpha-1} - B_{s,1}\, w_{s,1}\, r_{s,1}^{\alpha-1} + B_{s,1}\, w_{s,1}(1-\alpha)\,\tfrac{r_{s,1}^{\alpha-1}}{1+\phi_s} = 0 \\
&\implies \frac{\bar{B}_s - B_{s,1}}{B_{s,1}}(1-\tau_{s,1}) = \frac{\phi_s + \alpha}{w_{s,1}\psi_s(1+\phi_s)} \implies G(\tau_{s,1}) = \frac{\phi_s + \alpha}{w_{s,1}\psi_s(1+\phi_s)}
\end{aligned} \tag{52}$$

Which has a solution given the functional form of $B_{s,1}$. Since this amounts to essentially $\frac{\partial v_1}{\partial \tau_{s,1}} = 0$, hence the next period population doesn't respond to period 1 taxes. The other candidate solution is from solving the non-linear equation:

$$\frac{\sigma_v N_{s,1}}{\delta N_{s,2}(1-N_{s,2})} = \mathbb{E}[B_{s,2}(N_{s,2})] w_{s,2}(1-\tau_2) r_{2,s}(N_{s,2})^{\alpha-1} \left( \frac{\phi_s + \alpha}{1+\phi_s} - \frac{\bar{B}_s - \mathbb{E}(B_{s,2}(N_{s,2}))}{E(B_{s,2}(N_{s,2}))} K_2(N_{s,2}) \right), \quad (53)$$

$$\text{where } K_2(N_{s,2}) = -\psi_s \left( K_{s,1}(N_{s,2}) - \tau_{s,2} w_{s,2} \right) + \psi_s^2 K_{s,2}(N)^2$$

which is a non-linear function of $N_{s,2}$ which itself is a function of $v_1$ and hence $\tau_{s,1}$. The LHS is a U-shaped function of $N_{s,2}$ in the feasible ranges while the shape of the RHS depends on the parameter values. This opens up the possibility of multiple possible $N_{s,2}$ which correspond to multiple $\tau_{s,1}$ since

(ii) FOC with respect to the statutory tax rate $\tau_2$

$$0 = \frac{\partial \mathbb{E}[B_{s,2}]}{\partial \tau_{s,2}} \, w_{s,2} \, (1-\tau_{s,2}) \, r_{2,s}^{\alpha-1} \; - \; B_{s,2} \, w_{s,2} \, r_{2,s}^{\alpha-1} - \; (1-\alpha) \, B_{s,2} \, w_{s,2} \, (1-\tau_{s,2}) \, r_{2,s}^{\alpha-2} \, \frac{\partial r_{2,s}}{\partial \tau_{s,2}} \, . \quad (54)$$

Derivative of $\mathbb{E}[B_{s,2}]$ w.r.t. $\tau_2$

$$\frac{\partial \mathbb{E}[B_{s,2}]}{\partial \tau_{s,2}} \; = \; \psi_s \, \bar{B}_s \, \exp\Big[ -\psi_s \, K_{s,1} \; + \; \tfrac{\psi_s^2 K_{s,2}^2}{2} \Big] \, \frac{\partial K_{s,1}}{\partial \tau_{s,2}}, \quad (55)$$

where

$$K_{s,1} = \tau_{s,2} w_{s,2} + \frac{\tilde{A}_s \, (1 + r_{1,s} + \theta_s(\mu - r_f)) - L_s}{N_{s,2}} \implies \frac{\partial \mathbb{E}(A)_2}{\partial \tau_{s,2}} = \; w_{s,2} \psi_s \bar{B}_s \, \exp\Big[ -\psi_s \, K_{s,1} \; + \; \tfrac{\psi_s^2 K_{s,2}^2}{2} \Big].$$

Derivative of the housing return $r_{2,s}$ w.r.t. $\tau_2$

$$\frac{\partial r_{2,s}}{\partial \tau_{s,2}} \; = \; - \, \frac{1}{1+\phi_s} \left[ \frac{N_{s,2} \, w_{s,2} (1-\alpha)}{H^{\phi_s}} \right]^{\frac{1}{1+\phi_s}} (1-\tau_{s,2})^{-\frac{\phi_s}{1+\phi_s}} = \frac{-r_{2,s}}{(1+\phi_s)(1-\tau_2)}. \quad (56)$$

Plugging the terms into the FOC, we get

$$\implies 0 = \; w_2^2 \psi_s \bar{B}_s \, \exp\Big[ -\psi_s \, K_{s,1} \; + \; \tfrac{\psi_s^2 K_{s,2}^2}{2} \Big] (1-\tau_2) r_{2,s}^{\alpha-1} - \mathbb{E}[B_{s,2}] w_{s,2} r_{2,s}^{\alpha-1} + (1-\alpha) \mathbb{E}[B_{s,2}] w_{s,2} \frac{r_{2,s}^{\alpha-1}}{1+\phi_s}$$

$$\implies \frac{\phi_s + \alpha}{(1+\phi_s) w_{s,2} \psi_s} = \frac{(\bar{B}_s - \mathbb{E}(B_{s,2}))}{\mathbb{E}(B_{s,2})} (1-\tau_2) \implies \frac{\phi_s + \alpha}{(1+\phi_s) w_{s,2} \psi_s} = (1-\tau_2) \frac{\exp(-\psi_s K_{s,1} + \psi_s^2 K_{s,2}^2/2)}{1 - \exp(-\psi_s K_{s,1} + \psi_s^2 K_{s,2}^2/2)} \overset{Say}{=} G_2(\tau_2) \quad (57)$$

Which given the functional form will have a unique solution.
(iii) FOC with respect to the portfolio share $\theta_s$

$$0 \; = \; \frac{\partial \mathbb{E}[B_{s,2}]}{\partial \theta_s} \, w_{s,2} (1-\tau_2) r_{2,s}^{\alpha-1} N_{s,2} \, . \quad (58)$$

Using

$$\frac{\partial \mathbb{E}[B_{s,2}]}{\partial \theta_s} = -\bar{B}_s \, \exp\Big( -\psi_s \, K_{s,1} + \tfrac{\psi_s^2 K_{s,2}^2}{2} \Big) \Big[ -\psi_s \, \frac{\partial K_{s,1}}{\partial \theta_s} + \psi_s^2 \, K_{s,2} \, \frac{\partial K_{s,2}}{\partial \theta_s} \Big],$$

we get

$$0 = - \, \bar{B}_s \, \exp\Big( -\psi_s \, K_{s,1} + \tfrac{\psi_s^2 K_{s,2}^2}{2} \Big) \Big[ -\psi_s \, \frac{\partial K_{s,1}}{\partial \theta_s} + \psi_s^2 \, K_{s,2} \, \frac{\partial K_{s,2}}{\partial \theta_s} \Big], \quad (59)$$

which implies

$$\psi_s \, \frac{\partial K_{s,1}}{\partial \theta_s} \; = \; \psi_s^2 \, K_{s,2} \, \frac{\partial K_{s,2}}{\partial \theta_s} \quad \implies \quad \frac{\partial K_{s,1}}{\partial \theta_s} = \psi_s \, K_{s,2} \, \frac{\partial K_{s,2}}{\partial \theta_s}.$$

Since

$$K_{s,1} = \tau_{s,2} w_{s,2} + \frac{\tilde{A}_s \, (1 + r_{1,s} + \theta_s(\mu - r_f)) - L_s}{N_{s,2}}, \qquad K_{s,2} = \frac{\tilde{A}_s \, \sigma \, \theta_s}{N_{s,2}},$$

we have

$$\frac{\partial K_{s,1}}{\partial \theta_s} = \frac{\tilde{A}_s \, (\mu - r_f)}{N_{s,2}}, \qquad \frac{\partial K_{s,2}}{\partial \theta_s} = \frac{\tilde{A}_s \, \sigma}{N_{s,2}}.$$

$$\boxed{\theta_s \; = \; \frac{(\mu - r_f) \, N_{s,2}}{\tilde{A}_s \psi_s \, \sigma^2} \equiv \frac{(\mu - r_f)(1+g_N)}{(\bar{f}_s + \zeta_s) l_s \psi_s z_s \sigma^2}}$$

Thus optimal risky share is directly proportional to the excess return to variance ratio of the risky investment but inversely proportional to pension assets per capita and the public service provision efficiency of the state government. The second equality follows from $N_{s,2} = (1+g_N) N_{s,1}$, $a_s = \bar{f}_s l_s$ and $a_s + \zeta l_s = (A_s + \zeta_s L_s)/Y_{s,1} = (A_s + \zeta_s L_s)/(z_s N_{s,1})$.