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Falling Behind? State and Local Government Employee Compensation Since the Great Recession

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Abstract

We examine the relative compensation of state and local government workers compared to their private sector counterparts. Augmenting standard data sources on employee compensation with information on worker benefit provision derived from unpublished Bureau of Labor Statistics data and asset-based methodologies for valuing non-wage benefits, we document a significant erosion in the public sector compensation premium over the past fifteen years. While state and local workers received 13% higher total compensation than private sector workers with similar observable characteristics in 2011, our evidence indicates that this differential had declined to *negative* 3% by 2023. This decline stems from multiple factors: Public sector wages have fallen steeply relative to private sector wages since the Great Recession, public sector pension and retiree health care benefits have declined as a share of compensation, and the historically greater job stability in government employment has eroded. We also document several cross-sectional facts—for example, that: (1) relative public sector compensation has declined much more for workers with at least a college degree, as well as among those without union representation; and (2) the decline in relative pay of public sector workers occurred not only nationally but in every state. These findings have significant implications for public sector recruitment, retention capabilities, and ultimately, for the administrative capacity of state and local governments to effectively implement policies and deliver essential services.

We also document two areas of mismeasurement in official data. First, the use of cash accounting for defined benefit (DB) pensions, as opposed to accrual accounting, significantly skews upward the trend in relative public sector compensation in the Employment Cost Index—a principal source of information on compensation trends in the U.S.—over the past 15 years. Second, the accrual-based measures of public sector defined-benefit pensions in the National Income & Product Accounts understate the value of state and local government annual pension accruals from 2015 to 2021 when real interest rates were quite low.

I. Introduction

State capacity—a government's ability to effectively implement policies and achieve its objectives—has emerged as a focal point in recent policy discussions, particularly in the wake of challenges exposed by the COVID-19 pandemic and the rollout of the Inflation Reduction Act (IRA). Enacted in 2022, the IRA committed substantial resources to clean energy and climate initiatives, yet its implementation encountered significant delays stemming from administrative hurdles and regulatory uncertainties—an illustration that policy ambition can outpace institutional capability. The problem is not confined to the federal level. The water crisis in Flint, Michigan, is at least partly attributable to inadequate technical expertise and oversight capacity within state government (Flint Water Advisory Task Force 2016). A lack of sufficient capacity in state departments of transportation helps explain why public infrastructure costs in the United States far exceed those in peer economies (Liscow, Nober, and Slattery 2023). And while less dramatic, most states undertaking modernization of their unemployment insurance IT systems report that they lack staff with the expertise needed to carry the work through, illustrating how capacity shortfalls can impede even routine government functions (U.S. Government Accountability Office 2023).

These examples have a common element: The quality of the public workforce influences the ability of governments to deliver public goods. Competitive compensation for state and local government workers is therefore central to subnational state capacity, as it directly shapes the recruitment and retention of the skilled personnel on whom effective governance and policy execution depend. At the same time, some researchers argue that state and local employees are overcompensated relative to comparable private-sector workers (e.g., Biggs 2012), suggesting that realigning public pay with market rates could yield meaningful efficiency gains.

Comparisons between state and local government workers and private sector workers are complicated by the very different profile of employees across the sectors; e.g., public sector workers have higher average educational attainment. Moreover, public employees receive a higher share of their compensation in the form of benefits than do their private sector counterparts. Measuring the value of public sector benefits such as defined benefit pensions and retiree health care is methodologically complex, and the standard data sources on worker compensation in the U.S., such as the Current Population Survey (CPS) and Employment Cost Index (ECI), either fail to measure these benefits at all or suffer from notable measurement deficiencies.

In this paper, we measure the total compensation of state and local government (or "public") workers relative to the private sector. In doing so, we make a number of contributions. First, following a surge of work on the topic around 2010, there has been very limited research on relative public sector pay in recent years. We bring the research up to date, accounting for the evolution in the public and private labor markets over roughly the last 15 years. Second, we break down the public sector premia by education, unionization status, and state. Third, we make methodological contributions surrounding the proper measurement of benefit compensation—a major source of disagreement in the past literature. In particular, we utilize accrual-based measurement to value the defined-benefit (DB) pension and retiree health care benefits earned in a given year. In doing so, we document that the principal public source of public sector DB pension accruals, published by the Bureau of Economic Analysis, significantly understates the value of this benefit at times when real interest rates are low. We also develop two

methods to quantify the value of greater job stability—i.e., a lower propensity to be laid off—in the public sector.

We find that since the Great Recession, wages of public sector employees have fallen substantially relative to their private sector peers with the same observable characteristics and in the same occupation. Moreover, the value of defined benefit pensions and retiree health care in the public sector has fallen somewhat relative to wages, and job stability in the public sector has eroded relative to the private sector. Thus, the total compensation of state and local government workers relative to private employees has unambiguously eroded, falling between roughly 10% and 15% since 2011. Moreover, we find that this decline has occurred in all 50 states. These results on the trend in public sector compensation are extremely robust to methodological and data choices.

Our evidence on the level of the public sector compensation differential—as opposed to its trend—is less clear cut, as it is highly dependent on a number of methodological assumptions. That said, we view our evidence on the level of relative pay as conservative in the sense that, in cases where the correct methodological choice is unclear, we have leaned toward choices that boost the level of relative public pay. Accordingly, our results may overstate the level of relative public pay. These results suggest that public sector workers were paid about 13% more than their private sector counterparts with similar observable characteristics as of 2011. Since then, though, this public pay premium has gradually fallen and turned negative; as of 2023, state and local workers were paid about 3% less than their private sector peers when all job benefits are accounted for.

We further document that the trend decline in relative public sector compensation is driven by college-educated employees. This decline in relative pay among college-educated workers is confined to those lacking union representation. As of 2023, we estimate that college-educated workers were paid 13% less than their private sector peers. The decline in relative pay of the highly educated likely has particularly important implications for the capacity of state and local governments.

Our conclusion that relative public sector compensation has declined significantly over the past 15 years stands in stark contrast to the picture provided by the Employment Cost Index—the gold standard for measuring compensation growth in the U.S.—which suggests that relative compensation has been roughly flat over this period. We show that this discrepancy stems from two sources. First, the ECI measures DB pension benefits on a cash basis, while we measure them on an accrual basis. This switch, by itself, induces a downward trend in total relative public sector compensation. The difference is material because state and local governments have sharply increased contributions to pension plans in recent years to address past underfunding, pushing up benefit compensation measured on a cash basis. At the same time, they have reduced the value of pension benefits accruing to current employees—for example, by increasing the employee required pension fund contribution—pushing down benefits measured on an accrual basis. Thus, the two accounting approaches yield very different trends in state and local government benefits. Second, the ECI omits retiree health care and job stability, two components of public sector compensation whose estimated value has also declined since 2010.

One important limitation to our results is that they cannot establish whether or not public employees are earning rents—i.e., if they are "overpaid" or "underpaid"—relative to their private sector peers. This limitation arises from two factors. First, the unobservable quality of the public sector labor force likely reflects relative pay—e.g., as relative pay falls, the unobservable quality of the public workforce may decline. Second, we do not account for, nor can we easily measure, non-compensation aspects of jobs that workers may value (beyond job stability). Both explicit non-compensation amenities, such as flexible scheduling and teleworking policies, and more intangible amenities, such as dignity at work, have been shown to be valued by workers (see Mas 2025 for a comprehensive treatment; see also Collis and Van

Effenterre 2026; Dube, Naidu, and Reich 2022; Mas and Pallais 2017; Wiswall and Zafar 2018; and Maestas et al. 2023). For example, Tsao (2025) compares teachers' pay to their own next-best alternatives—effectively holding both observable and unobservable worker characteristics constant—and finds that teachers in Kentucky were paid more than they would have been paid in their next-best jobs, which tended to be relatively low-paying positions, but that this higher pay was almost fully offset by poorer working conditions, implying little or no rents, particularly for inexperienced teachers.

The declining relative pay of public sector workers that we document is large in magnitude, has occurred over a relatively long period of time, and is very pervasive across the U.S. It is therefore almost certainly associated with important shifts in the public sector labor market. In particular, the results may be associated with several (non-mutually exclusive) labor market changes. The composition of the public sector workforce could be shifting to lower unobservable quality. Rents available to public sector employees could be decreasing. Or the value of non-wage amenities in the public sector could be increasing to offset the reduction in compensation. As discussed below, these outcomes may differ by employee age as older public sector workers tend to face higher costs when transitioning to the private sector due to the structure of pension and retiree health benefits.

Our results point toward a strong need for further research into the precise consequences of the decline in relative public sector pay we document. Although we mostly leave the implications of the decline in relative public sector pay to such future research, we do conclude the paper with a brief discussion of two suggestive pieces of evidence that the decline in relative pay has inhibited the ability of the sector to attract and retain high-quality employees: the difficulty governments had restaffing relative to the private sector following the COVID-19 pandemic, and the eroding labor market landscape for K-12 teachers.

The paper proceeds as follows. Section II provides background information on public and private sector workers, compensation packages in the two sectors, and past work on public-private pay differentials. Section III discusses the data and measurement challenges. Section IV presents our methodology and results. Section V concludes by discussing implications of our findings for public sector recruitment and retention.

II. Background

II.A Employee Characteristics and Compensation in the State and Local Government and Private Sectors

Employee characteristics differ starkly across the state and local and private sectors, as displayed in Table 1. The starkest differences are in educational attainment and occupation: State and local workers are 22 percentage points more likely than private employees to have a four-year college degree or more (row 9) and 22 percentage points more likely to be in a professional or managerial occupation (row 10). Public workers are also more likely to be married and female than are private workers (rows 2 and 3). These differences in characteristics must be accounted for when assessing relative compensation across the sectors.

Table 1
Summary Statistics

row		Private	State and Local Government
	A. Employee Demographics		
1	Age	40.6	44.1
2	Married	0.51	0.60
3	Female	0.47	0.61
4	White	0.60	0.65
5	Black	0.11	0.13
6	Hispanic	0.20	0.15
7	High School	0.26	0.14
8	Some College	0.26	0.22
9	4-year College or More	0.40	0.62
10	Managers and Professionals	0.37	0.59
11	Precision Production, Craft and Repair	0.09	0.03
12	Operators, Fabricators and Laborers	0.13	0.04
13	Union	0.07	0.38
	B. Share of Total Employee Compensation		
14	Wages	0.85	0.70
15	Benefits	0.15	0.30
16	Defined Benefit Pension (accrual basis)	0.01	0.10
17	Retiree Health Care (accrual basis)	.	0.01
18	Other Benefits	0.15	0.19

Source. Panel A: 2024 Current Population Survey; Panel B: 2023 National Income and Product Accounts and author's calculations for defined benefit pensions and retiree health care.
Note. Panel A displays means.

Employees in the two sectors also receive their compensation in different forms. Public workers receive more than one-fourth of their total compensation in the form of benefits, whereas private sector workers receive only around one-sixth (row 15). This difference is mostly explained by defined benefit pension plans, which account for nearly 10% of public sector compensation, but which account for only a very small share of compensation in the private sector (row 16); retiree health care, which is broadly provided in the public sector but is now uncommon in the private sector, also plays a small role (row 17). Other benefits—e.g., health insurance, paid time off, and defined contribution pension plans—account for a somewhat larger share of compensation in the public sector than in the private (row 18).

II.B Past Research

There was a flurry of research on wage differentials between the private and state and local government (S&L) sectors around 2010-2012, driven by the difficult budget conditions confronting these governments in the wake of the financial crisis and Great Recession. These papers generally concur that state and local workers receive lower wages, but higher benefits, than their private sector counterparts. They reach widely divergent conclusions, however, over whether the non-wage compensation is sufficient to result in a positive public sector total compensation premium. The different conclusions generally reflect differences in the methodology used to measure benefits.

Gittleman and Pierce (2012) conducted an unusually comprehensive analysis using the ECI, CPS, and confidential microdata from the BLS's National Compensation Survey (NCS). They found that state and local government workers received higher total compensation than similar private sector workers, with the differential driven primarily by more generous benefits. That said, there are difficulties inherent to using the NCS to measure benefits; in particular, pension benefits suffer from a number of methodological concerns (Biggs and Richwine 2012), retiree health care is generally not captured at all, nor is the value of the greater relative job stability in the public sector. The authors acknowledge these shortcomings and conclude that they are unlikely to overturn their conclusion that state and local workers receive a pay premium. Biggs and Richwine (2014) also find positive public sector compensation premiums on average across states, with a wide range from slightly negative to over 40%. They emphasize the importance of including the value of job security in relative compensation calculations. In contrast, Munnell et al. (2011), accounting for the value of pensions, retiree health care, and job stability, find that state and local workers earn approximately the same as they would in the private sector. Their work emphasizes education as the critical factor—workers with bachelor's degrees or higher face wage penalties in government work.[1] [2]

III. Data and Measurement Challenges

This section describes the main data sources used in the analysis and the measurement challenges that motivate our alternative benefit measures. We first discuss the ECI, CPS, and ECEC, and then summarize the limitations of existing data on DB pensions, retiree health care, and job stability.

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[1] Other notable contributions to this body of work include Allegretto and Keefe (2010), Bender and Heywood (2010), Keefe (2012), Lewis and Galloway (2011), Madland and Bunker (2011), Reilly and Reed (2011), Biggs and Richwine (2012), and Thompson and Schmitt (2010). Important earlier work on public sector compensation premiums include Borjas (2002), Freeman (1987), Katz and Krueger (1991), Krueger (1988), and Poterba and Rueben (1994), and Smith (1977).

[2] We are aware of two contributions—both working papers—on this topic made over roughly the past 10 years. Biggs (2019) examines 1998 and 2017 using BEA data and concludes that public sector relative compensation rose somewhat over this period. Biggs (2022) focuses on Connecticut, but also presents national results using the Census Bureau's American Community Survey and BEA data and finds that in states with the lowest public compensation, private and public compensation are well aligned; in contrast, in the states with the highest public renumeration, state and local government compensation is well above that in the private sector.

III.A The ECI and CPS

To focus on the *trend* over time in relative pay, we primarily use the BLS's Employment Cost Index (ECI)—the principal national economic indicator for measuring compensation *changes* over time. Interpreting wage changes over time is challenging due to the influence of changing worker composition on wages and benefits. The ECI's defining feature is its use of a methodology to isolate changes in compensation costs unrelated to changes in workforce composition.[3] The ECI provides indexes for various pieces of total benefit compensation—e.g., health insurance. However, there are no published tabulations for DB pensions. Fortunately, though, the BLS generously made available unpublished tabulations which provide detailed information on defined benefit pensions.[4]

Although the ECI is very well suited to examining changes in compensation, it fails to provide any information on the *level* of compensation, nor does it allow us to control for worker characteristics when comparing wages across the public and private sectors. To address these shortcomings, we turn to the CPS.

The CPS microdata contains a rich set of information on worker characteristics which allows us to control for the very different compositions of the private and public workforces; these data also allow us to decompose wage differentials by group—e.g., the public sector wage premium for college-educated workers. Another advantage of the CPS is its (limited) longitudinal structure, which we exploit to examine differences across the sectors in job stability. The chief drawback of the CPS, however, is its lack of information on non-wage employee benefits.[5]

Finally, in order to obtain information on the *level* of benefit provision, we turn to the BLS's Employer Costs for Employee Compensation (ECEC). The ECEC is based on the same underlying survey data as the ECI—both are part of the National Compensation Survey (NCS)—and provides information on the dollar value of wages and benefits. The ECI and ECEC have problematic measurement of DB pensions, omit

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[3] The ECI achieves this fixed-composition measure in two steps. First, the ECI uses a fixed-weight approach where compensation changes are measured within detailed cells defined by industry, occupation, and sector (private and state and local government), which are aggregated up using fixed employment weights. Second, the ECI methodology uses a matched-sample approach under which the change in compensation in a cell from time *t-1* to *t* is calculated only using "matched job quotes." Matched job quotes refer to the average compensation costs per hour within an establishment-occupation (i.e., a "job") which are available for both the time *t* and *t-1*. The use of matched job quotes ensures that measured compensation changes do not reflect composition shifts within the detailed industry-occupation cells. E.g., it prevents the entry of a new establishment into the ECI sample at time *t* from influencing measured compensation growth a time *t*.

[4] The ECI DB pension series should be interpreted with care because they do not meet the BLS's standard publication criteria.

[5] The CPS surveys roughly 150,000 individuals each month. Each household is sampled for two four-month periods, with an eight-month gap in between. On the 4th and 8th months in the survey (the outgoing rotations), the CPS collects data on wages. We use the CPS extracts from the Economic Policy Institute (EPI); the EPI Merged Outgoing Rotation Group surveys harmonize data over time on wages and demographic information of workers. We use EPI's wage variable which imputes hourly wages for salaried workers using their reported usual hours worked and adjusts for workers with variable hours. The EPI wage variable also accounts for top-coding. We exclude overtime, tips, and commissions. Following Gittleman and Pierce (2012), we exclude workers for whom wages or hours data are imputed by the Census (about 43,000 observations in the 2024 ORG Data).

retiree health care, and do not consider job stability as it is a job attribute, not a benefit.[6] We therefore turn to alternative sources to measure these job attributes.

III.B Measurement Challenges and Alternative Benefit Sources

The ECI and ECEC measure the value of DB pensions on a cash basis—i.e., by the amount contributed by the government to the pension fund annually. This cash accounting does not reflect the benefits being earned by workers, which requires measuring on an accrual basis—i.e., by calculating the present discounted value of the future pension benefits earned that year by workers, typically referred to as the "normal cost." States and localities have historically failed to fully prefund pension benefits. Moreover, in recent years they have been cutting the level of DB pension benefits while simultaneously increasing pension funding (Lenney, Lutz, Schüle, and Sheiner 2021). For these reasons and others, the value of pension benefits under cash accounting can diverge significantly from the value under accrual accounting. We therefore turn to alternative, non-NCS data sources.

The annual actuarial reports issued by these governments, which provide the pension normal cost for that year, are a possible alternative. However, these reports are themselves problematic because state and local governments, per accounting rules, generally use their expected return on assets as the discount rate. Standard principles of financial valuation, in contrast, suggest the discount rate should reflect the risk profile of the stream of future pension payments, which depends on the probability that the payments will be honored, among other factors. Given the relatively strong legal protections surrounding these payments, it is appropriate to use a discount rate lower than the expected return on pension plan assets (Brown and Wilcox 2009; Novy-Marx and Rauh 2011; Lucas 2012). All else equal, a lower discount rate implies a higher valuation on annual DB pension benefits which will, in turn, increase the measured compensation of public workers relative to the private sector.

To value the compensation attributable to defined benefit pensions, we could in principle turn to the Bureau of Economic Analysis's (BEA) National Income and Product Accounts (NIPA), which restate private and public pension actuarial reports on a common, lower, more appropriate discount rate—the AAA corporate rate—rather than the assumed return on plan assets (Lenze 2013). We do not use the BEA's published normal costs directly, however, because they are excessively insensitive to changes in the discount rate relative to widely used actuarial rules of thumb, and as a result understate normal costs when market real rates are low. Appendix Section B.3 documents this insensitivity and explores its causes in detail.

In response to these measurement concerns, we calculate our own DB pension normal cost estimates using a multi-step procedure, including the use of microdata on pension plans, a benchmark pension accounting model, and the use of annual, market-based interest rate and expected inflation measures. This methodology is briefly outlined in section IV.B.1 and presented in greater detail in Appendix Section B.

For retiree health care, governments have been required to provide sufficient detail to calculate normal costs only since fiscal year 2018. We use data on a large number of retiree health care plans from Pew

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[6] Job stability, while a valuable job attribute, is not a "benefit" and is therefore not measured in the NCS. Retiree health care is widely provided as a benefit to state and local workers, but is overwhelmingly not prefunded (Pew 2023). Instead, payments are made annually to cover the insurance costs. As these payments are made on behalf of retirees, not current employees, they are not captured in the NCS. Thus, the NCS mostly misses the value of retiree health care.

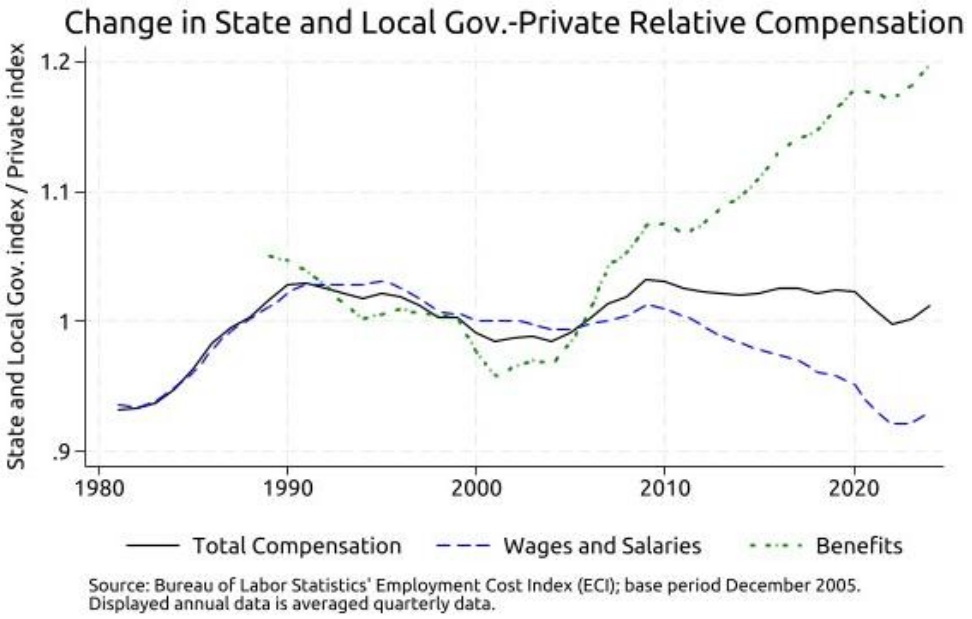
(2023) to calculate the annual normal cost accrual value in fiscal year 2018, and calculate changes over time using data from a smaller subset of plans. The methodology is outlined in section IV.B.2 and presented in greater detail in Appendix Section C.

Finally, we estimate the value of job stability using a two-step procedure. First, we use the longitudinal element of the CPS data to calculate the difference in permanent job loss probabilities between the public and private sectors. We then calculate a dollar value for this difference in job stability using two different approaches—one involving a simple modeling framework and parameterized using the empirical literature on the cost of job loss and another grounded in revealed preference empirical estimates. These two approaches produce remarkably similar values for the greater job stability in the public sector. The job stability methodology is discussed in detail in section IV.B.3.

IV. Results

We start the examination of state and local government compensation differentials with the private sector by examining *trends* in relative compensation over time using the ECI. The black solid line in Figure 1 displays the ratio of the state and local government total compensation index in the ECI to its private sector counterpart; it is indexed to equal 1 in 2005. The series can be interpreted as capturing the percent change in the public pay premia relative to 2005. During the 1980s the public pay premia rose substantially; thereafter, the premia declined gradually through 2005. From 2005 through 2010, the premia reversed as state and local workers gained on their private sector counterparts. ` These results are consistent with past research such as Poterba and Rueben (1994) and Gittleman and Pierce (2011).

Figure 1



Source: Bureau of Labor Statistics' Employment Cost Index (ECI); base period December 2005.
Displayed annual data is averaged quarterly data.

After 2010, relative compensation as measured by the ECI held roughly constant through 2024. That said, the flat relative compensation over this period masks a stark shift in S&L relative compensation toward benefits and away from wages and salaries. In Figure 1, relative wages and benefits are shown by the dashed blue line and dash-dot green line, respectively; they moved broadly in tandem until around 2010. Since then, relative wages have tumbled, while relative benefits have soared. Next, we examine relative wages and benefits in more detail.

IV.A Wages

Relative public sector wages, holding industry, occupation, and "job" fixed, have declined by 8 percentage points through 2024 relative to their 2009 peak according to the ECI. While only a portion of total compensation, wages are likely particularly relevant to attracting new workers, as they are the most salient component of compensation. Moreover, younger workers may place limited value on retirement benefits such as pensions and health care (e.g., Clark and Morrill 2016). More generally, workers of all ages may not fully value benefits such as pensions because of their complexity or other reasons (e.g., Gustman and Steinmeier 2005, Fitzpatrick 2015). Thus, the relative decline in state and local government wages may have important implications for these governments' ability to attract and retain workers regardless of the trend in relative non-wage compensation.

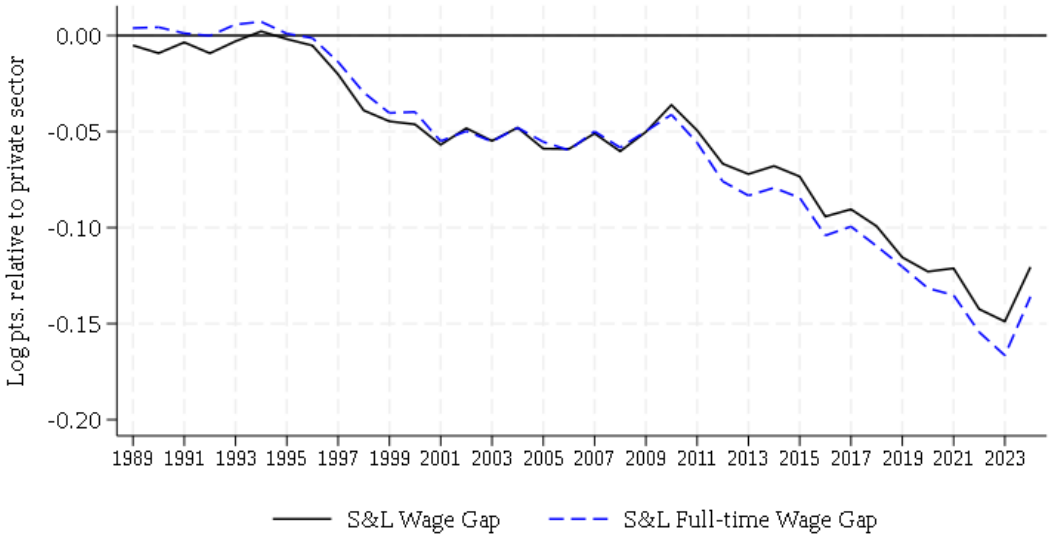
In order to confirm the ECI-based results for the trend in relative public wages, and to shed light on the level of relative wages, we turn to the CPS data. To compare pay of similar workers, we estimate Mincer-style regressions of hourly wages of workers using the CPS from 1989-2024. In particular, following the approach in the past literature—e.g., Katz and Krueger (1991), Poterba and Rueben (1994), and Gittleman and Pierce (2011)—for each year in the sample, we regress the log of the wage on an indicator for whether the worker is a state and local government worker and include the following controls: a quartic in age, four education categories, marital status, Census division interacted with metro status, race interacted with gender, and major occupation. We weight the regression with the CPS survey weights to make the results representative of an average worker.

Figure 2 displays the coefficients on the state and local government indicator. The results indicate that the relative pay of state and local workers leveled off at about 5% lower than comparable private sector workers from the late 1990s through about 2010. After the Great Recession, and coinciding with large deterioration in the budget positions of these governments, state and local government relative worker pay fell substantially. Thus, the evidence from the ECI and CPS paints a broadly similar picture of a steady decline in relative S&L wages since the Great Recession. The similar trend in relative public wages across the two data sources is notable given the very different methodologies used to control for compositional differences across the sectors. In terms of the level of the public wage differential, by 2024 wages of public sector workers were a substantial 12% lower than their comparable private sector counterparts.

We view the above estimates as conservative—i.e., tending toward finding a higher relative public wage premium—as we have omitted control variables often used in the past literature which depress the measured premia. (See Appendix Figure A1, which documents that controlling for firm size or union

status, which we omit from the results displayed in Figure 2, significantly reduces the public wage premium.[7])

Figure 2
S&L government wages relative to private sector



Gap in wages for S&L workers relative to private. Regressions are weighted and control for quartic in age, four education groups, metro status interacted with Census division, race interacted with gender, marital status, and major occupation. Full-time is defined as usually working 35 or more hours per week.

IV.B Benefits

We now turn to relative differences in non-wage compensation. We first correct the three components of compensation that are missing or mismeasured in standard sources—DB pensions, retiree health care, and job stability—and then combine them with wages and other benefits to estimate total relative compensation.

IV.B.1 Defined Benefit Pensions

Unpublished data provided by the BLS and displayed in Figure 3 reveals that the surge in relative public sector benefits displayed in Figure 1 is solely attributable to defined benefit (DB) pension plans. As shown by the blue dashed line, relative benefits are flat once DB pensions are removed from the private and

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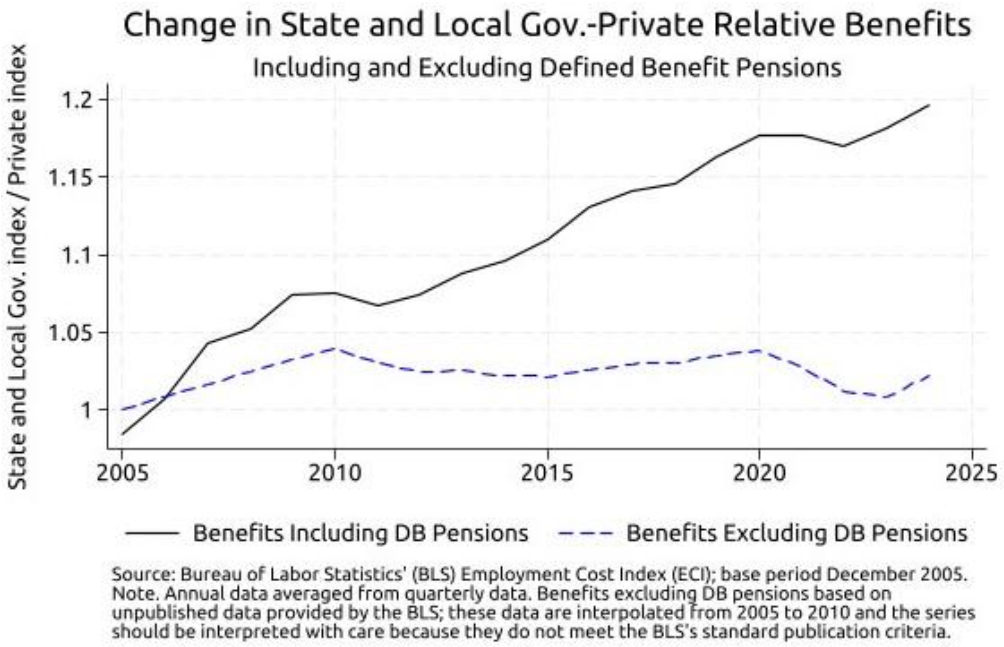
[7] See Gittleman and Pierce (2011) for a detailed discussion of the role of control variables in assessing public-private compensation differentials. We omit union representation from our set of control variables as, in recent years, union representation has become very substantially more common in the state and local government sector than in the private sector. Unionization can thus be viewed primarily as an attribute of the state and local sector which should not be controlled for when comparing compensation across the sectors.

public indexes used to construct relative benefits.[8] The rise in relative pension benefits is surprising because states and localities have been reducing DB pension generosity for both current workers, e.g., by paring back cost-of-living-adjustments and requiring increased employee contributions, and for new workers, e.g., by introducing new, less generous benefit tiers (Abashidze, Clark, and Craig 2021; Aubry and Crawford 2016; Lenney, Lutz, Schüle, and Sheiner 2021). However, as discussed above, the ECI measures DB pension benefits on a cash basis which does not necessarily reflect the annual benefits being earned by workers.

Figure 3



Source: Bureau of Labor Statistics' (BLS) Employment Cost Index (ECI); base period December 2005.
Note. Annual data averaged from quarterly data. Benefits excluding DB pensions based on unpublished data provided by the BLS; these data are interpolated from 2005 to 2010 and the series should be interpreted with care because they do not meet the BLS's standard publication criteria.

We therefore construct an accrual-based measure of DB pension compensation; Appendix B provides the detailed methodology. In brief, our starting point is pension plan microdata from the Public Plans Data (PPD), which include each plan's reported normal cost—i.e., the present discounted value of the future pension benefits earned that year—which reflects its chosen discount and expected inflation rates. We build a simple benchmark actuarial pension model to adjust these normal costs onto a harmonized basis reflecting market-based measures of inflation and real interest rates.[9] Finally, we benchmark our normal cost estimates to the BEA's accrual-based DB pension measure—which captures the universe of state and local pensions—in 2011. We use 2011 because, in that year, our market-based assumptions for real interest rates and expected inflation are extremely close to corresponding assumptions used by the BEA. We then grow the BEA's 2011 normal cost estimate forward and backward in time using an approximated growth factor equal to the sum of the ECI percent change in state and local government

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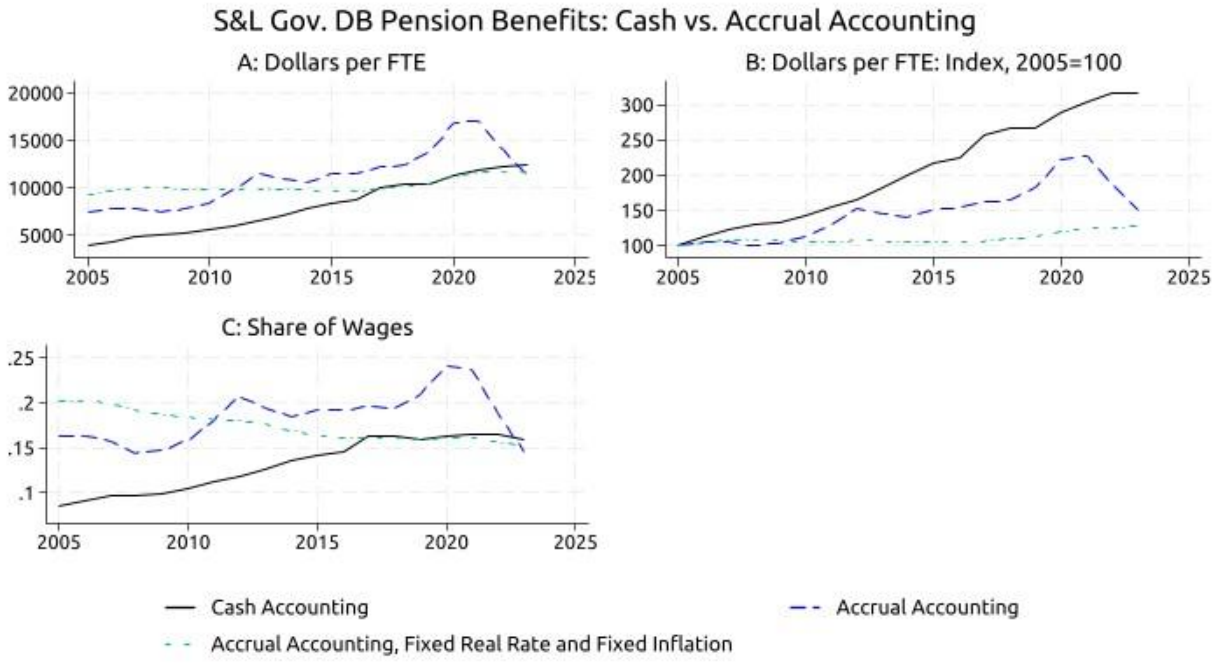
[8] Private DB pensions have become a small share of private sector compensation. As a result, private sector DB benefits do not play an important role in the relative S&L benefits trend displayed in Figure 3.

[9] We use the Moody's AAA nominal corporate bond yield and the Cleveland Federal Reserve's 10-year expected inflation rate.

wages and the percent change in our actuarial model-based normal cost rate—the normal cost expressed as a fraction of annual wages. Finally, to obtain DB pension compensation, we subtract from the total normal cost the employee share of the normal cost—i.e., the annual required employee contribution to the pension fund—and add in pension administrative expenses.

The discussion of these results focuses on the resulting differences between cash and accrual accounting, which are the central issue for the ECI analysis. Figure 4a displays pension benefits per full-time equivalent worker on a cash basis (from the NIPA, black line) and on an accrual basis (our measure, dashed blue line). Benefits on an accrual basis were roughly twice as large as benefits on a cash basis at the start of the sample in 2005. By 2023, though, cash benefits had eclipsed the accrual-based benefits.

Figure 4



Source: Bureau of Economic Analysis's (BEA) National Income and Product Accounts (NIPA). Public Plans Data. Author's calculations. FTE denotes full-time equivalent worker.

Figure 4b displays an indexed version of pension benefits. There is a very stark difference in the growth rates between cash and accrual accounting. As measured by cash accounting (the black solid line), there has been rapid growth in pension benefits. This is consistent with evidence in Lenney, Lutz, Schüle, and Sheiner (2021) that states and localities have significantly increased their contributions to pension funds in recent years. However, such funding can be lower than pension benefit accruals if governments are underfunding pensions or greater than current year accruals if they are paying down legacy liabilities—i.e., addressing past years' underfunding.[10]

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[10] The gap between the accrual and cash benefits in Figure 4a captures the shortfall of funding in that year relative to benefits earned that year (i.e., the normal cost). The gap does not capture the portion of the stock of past unfunded liabilities which standard actuarial methods dictate be amortized in that year.

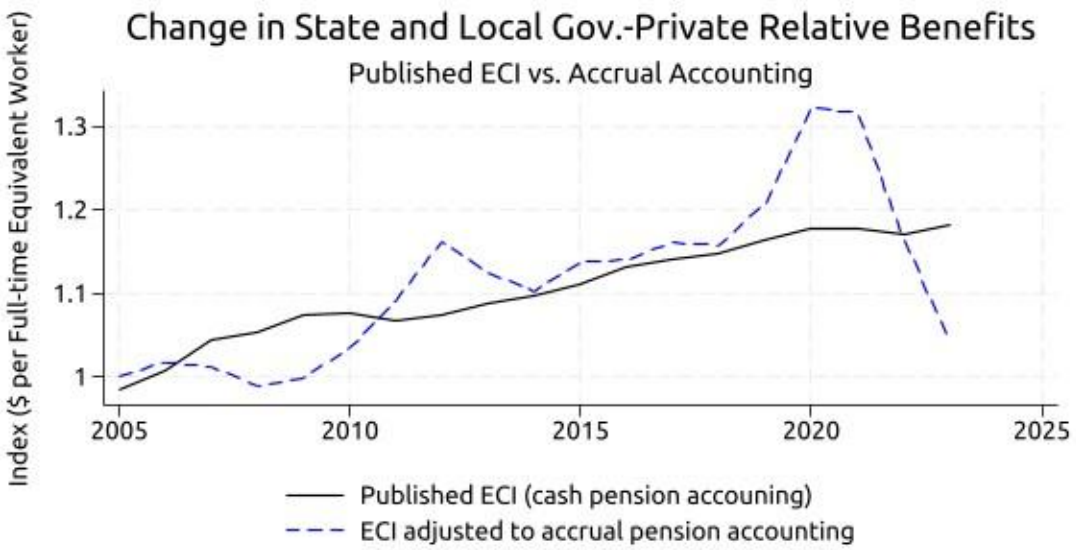
The accrual DB pension measure (the blue dashed line) has also risen, although at a slower pace than the cash measure. The rise in the accrual measure from 2009 to 2021 importantly reflects the gradual fall in real and nominal interest rates over this period. The plunge in normal costs per worker post-2021 reflects the surge in interest rates over that period. These changes in the accrual measure are properly viewed as well-measured benefit changes because they reflect changes in the market price of buying a comparable annuity from a private source. That said, it is unclear if workers perceive such sharp annual swings in benefits driven by financial market prices.

The green lines in Figures 4a-c show the evolution of the accrual value of pension benefits under a counterfactual of constant inflation and interest rates (set equal to their values in fiscal year 2024). As shown in Figure 4a, without changes in interest rates, the nominal value of benefits per FTE would have been fairly flat over the past two decades—meaning that benefits were not keeping pace with wages. This accords well with governments making pension plans less generous over time. Indeed, Figure 4c reveals that pension benefits as a share of wages have been declining when inflation and interest rates are held constant. Pension benefit formulas are almost always a function of wages. Thus, absent change in benefit formulas, benefits would be expected to keep pace with wages.

The dashed blue line in Figure 5 displays the ECI relative total benefit index—the published version of which is displayed on Figure 1—adjusted to be on an accrual basis for DB pensions. The adjusted and unadjusted state and local government relative benefit indexes display similar growth through 2018. The similar growth rates in the two series may appear puzzling given the much more rapid growth in the cash DB pension index relative to the accrual-based index in Figure 4b. However, the more rapid growth in the cash measure is tempered by its relatively smaller size relative to the accrual measure in 2005 shown in Figure 4a; 2005 is the base index year, which defines the aggregation weights used to construct the total benefit index under accrual DB accounting on Figure 5. Stated differently, in aggregating up the indexes in Figure 4b with the index for all other benefits to produce the series on Figure 5, the accrual-based measure receives significantly more weight than the cash-based measure. (See Appendix section D for a detailed treatment of our ECI aggregation methodology, including discussion of the assumptions underlying the aggregation.) Finally, note that the "all other benefits" category includes defined contribution pension plans.

In 2019, the accrual accounting total relative benefit measure in Figure 5 diverges sharply from the cash-based measure, first surging above the cash measure and then tumbling well below it by 2023. By 2023, the difference in total relative benefit growth between the published ECI total benefit index and the accrual DB pension-based index is a substantial 20 percentage points. Thus, compared to the published ECI, correctly accounting for DB pension benefits leads to very different conclusions about relative benefit growth in the public sector.

Figure 5



Source: Bureau of Labor Statistics' (BLS) Employment Cost Index (ECI, base period December 2005) and Bureau of Economic Analysis's (BEA) National Income and Product Accounts (NIPA). Author's calculations.
Note. Annual data averaged from quarterly data. Utilizes unpublished data provided by the BLS; these data are interpolated from 2005 to 2010 and the series should be interpreted with care because they do not meet the BLS's standard publication criteria.

IV.B.2 Retiree Health Care

Access to subsidized health care during retirement is far more common in the state and local sector than in the private sector. In 2024, about 63% of S&L workers age 65 or older had access to retiree health insurance; just 13% of private workers did (Bureau of Labor Statistics, 2024). Moreover, the provision and generosity of retiree health care benefits vary significantly across governments. Some states finance a large share of the cost of health insurance for their retirees, while other states simply allow their retirees to purchase health insurance through their employer plan, providing an implicit subsidy given that retirees are older and have higher expected medical claims. We estimate that the accrual value of retiree health insurance for state and local employees is less than 1% of wages in 36 states but 2% or more of wages in 8 states; it exceeds 4% of wages in Delaware and New Jersey.

As with pensions, the Employment Cost Index (ECI) uses employer contributions to retiree health care trust funds as the measure of the compensation value of employer-provided retiree health benefits. However, since these benefits are predominantly unfunded, the ECI effectively excludes them, failing to capture any changes in their value.[11] Given that most state and local governments offer these benefits to employees, while few private sector employers do, this omission skews relative public sector compensation measurements.

To address this disparity, we estimate both the current value of these benefits and their evolution over time. Appendix Section C details our methodology. Briefly, we begin with Pew (2023), which provides data on 172 state and local retiree health insurance plans for fiscal year 2019. We adjust those data so that

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[11] Less than 10% of OPEB benefits have been prefunded by the states (Pew Charitable Trust 2023).

accrual values are measured using our preferred discount rate rather than the plans' discount rates. Estimating changes over time relative to FY 2019 is difficult because historical data are limited. For 2011 to 2015, we rely on the small subset of public retiree health care plans that published the necessary information; from 2016 on, we use more comprehensive data.[12] Finally, we overestimate the public sector compensation differential due to retiree health care because we assume that private sector workers have no access to retiree health benefits, even though a small share of private sector employees do have access.

Accounting standards require states to use a discount rate that reflects a blend of the expected return on investments and the municipal bond borrowing rate, with the proportion of unfunded liabilities determining the share of the blended discount rate equal to the municipal bond rate. As most plans are unfunded, the applied discount rate is typically the relatively low municipal bond rate—significantly lower than the discount rate used for pension liabilities. But, just like with pensions, the discount rate should correspond to the risk profile of those benefits. Unlike pensions, retiree health insurance is not guaranteed and generally lacks strong legal protections; in most cases state and local governments can modify or terminate these benefits at their discretion.[13] Thus, for our analysis, we discount retiree health benefits using a rate that is 1 percentage point higher than the AAA corporate rate used by BEA to discount future pension benefits, with the higher rate reflecting the much higher uncertainty about whether these benefits will be provided in the future. While we judge default risk as the dominant factor in setting the discount rate for retiree health care, the appropriate rate is a very complex issue with numerous factors to consider—see section 2.4 of Lutz and Sheiner (2014) for a discussion.

As shown in Figure 6a, our findings indicate that, in 2023, the average nominal value of retiree health benefits was approximately $750 per worker in the state and local sector, well below the $1185 in 2011. Figure 6b shows these benefits as a share of wages and also includes a counterfactual in which interest rates are held constant at their 2023 value. Benefits have declined as a share of wages on average (black line), although the low interest rates in 2020 boosted them temporarily. Holding interest rates fixed (dashed blue line), retiree health insurance accruals have been falling as a share of wages since 2015. (We lack data on retiree health care benefits after 2023, so simply assume that benefits as a share of wages would have been constant if not for changes in interest rates.)

This decline likely reflects two factors. First, as shown in Appendix Table 1, many state and local governments have instituted reforms that lower the cost of providing retiree health insurance, either by requiring contributions from employees, tightening eligibility rules, or simply reducing the value of the subsidy. Second, health care costs—particularly the costs of Medicare and Medicare Advantage plans—were significantly lower than expected from 2010 to 2022 (CBO, 2023), further reducing the estimated accrual value of retiree health benefits.

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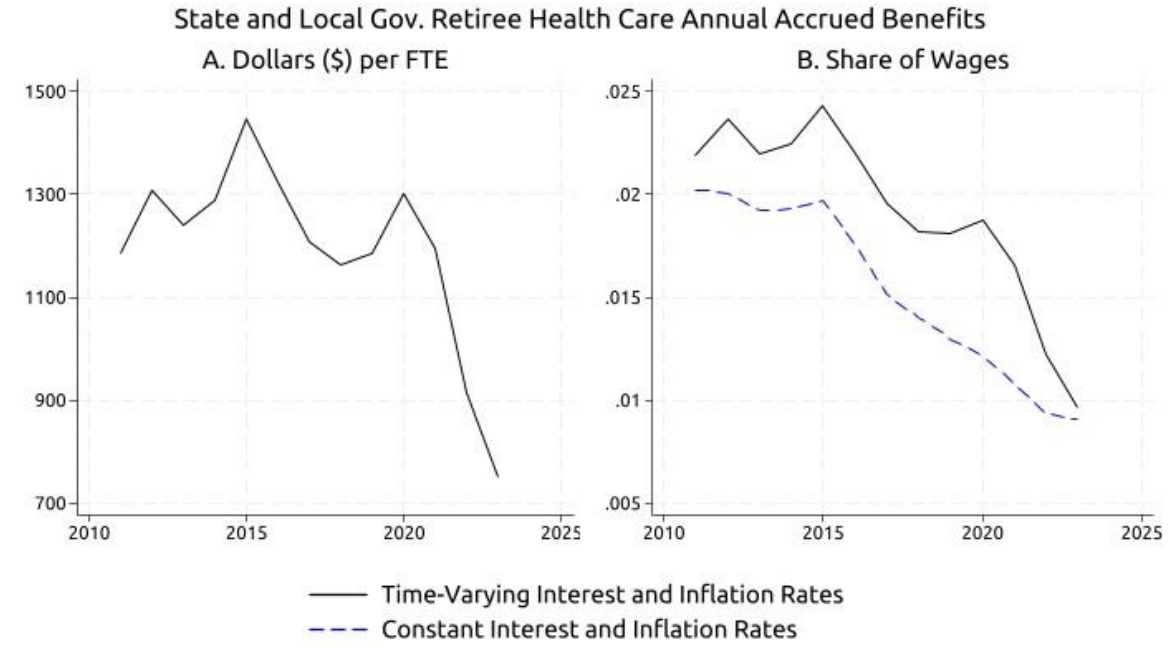
[12] Recent changes in accounting standards, specifically Government Accounting Standards Board (GASB) Statements No. 74 and 75, have mandated that governments report the estimated future cost of retiree health benefits and other non-pension post-employment benefits (OPEB). These data are not always published, however, and are sometimes simply embedded in other financial reports.

[13] For instance, the Texas Employee Retirement System cautions: "The Texas Legislature determines the level of funding for such benefits and has no continuing obligation to provide those benefits beyond each fiscal year" (Employees Retirement System of Texas 2025). More generally, numerous states and localities have reduced the generosity of, or even eliminated, the benefit in the past (Clark, 2009; Franzel and Brown, 2012).

Figure 6



Source: PEW (2023), author's calculations, and BEA. FTE denotes full-time equivalent worker.

IV.B.3 Job Stability

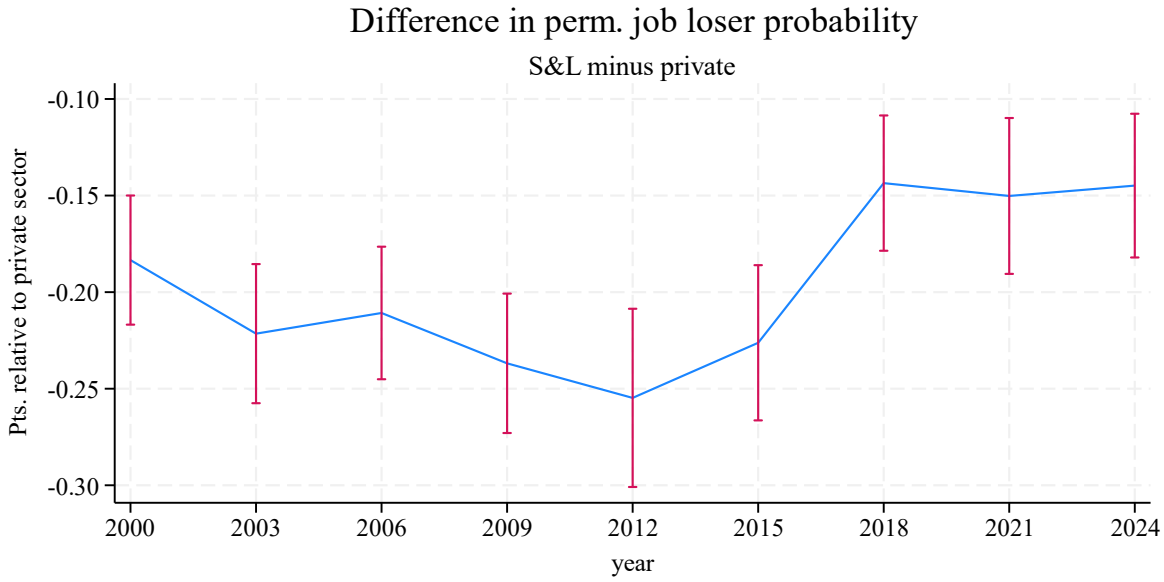
In addition to wages and benefits, jobs are more valuable to workers if they are more stable and secure. Government jobs are typically viewed as less prone to layoff or job loss risk. Our relative compensation estimates above, though, do not capture the value of job stability, as this job attribute is not explicitly captured in any of the data sources we use on job benefits. Moreover, little research has shed light on how relative job stability between the public and private sectors has changed over time.

To estimate the change in relative job stability over time, we estimate linear probability models akin to the wage regressions shown above in section IV.A using longitudinally linked CPS data. In particular, we estimate monthly job loss probabilities among private and state and local government workers. We define job losers as those experiencing a transition from employment to unemployment due to a permanent job loss. The results are qualitatively similar if we examine any type of employment-to-unemployment transition or if we include transitions of those who leave the labor force but report wanting a job. Because job loss probabilities are notably different for school-age and retirement-age workers, we restrict this analysis to those aged 25-54, and control for a similar set of characteristics as in the wage analysis (again omitting union coverage). We pool three years of data together due to the smaller sample sizes when matching data longitudinally.

As shown in Figure 7, we find that the job loss probability of S&L workers relative to similar private-sector workers declined from 2000 through the Great Recession, with the difference falling to 0.25

percentage point per month in 2009—or 3 percentage points per year.[14] However, since 2012, the job loss rate has risen for S&L workers relative to private-sector workers to a gap of about 0.15 percentage point per month—or 1.8 percentage points per year—with S&L workers facing a job loss probability of about 1.4% in the past few years, while comparable private sector workers faced a job loss probability of 3.1%. Thus, the loss in relative job stability of state and local government workers compounded the decline in relative wages and benefits that occurred over roughly the same time period.

Figure 7



U(perm job loser) probability defined as the monthly transition rate from employment to unemployed due to permanent job loss . Gap in monthly non-employment probability is estimated from weighted regressions controlling for quartic in age four education groups, metro status interacted with Census division, race interacted with gender , marital status, and major occupation. Excludes 2020.

To determine how valuable a given degree of job stability is for a worker, we outline a simple model of job loss. We ground the model in the long-running empirical literature which has shown that job losses cause persistent declines in the earnings potential of workers. In particular, we rely on the influential analysis of Davis and von Wachter (2011) who show that job loss causes a roughly 12% decline in the present value of wages for a worker.

Let the present discounted value of a job be represented by the value V_t, equal to the flow of wages, w, from the job discounted by both the real interest rate as well as the probability of losing the job. We assume that the probability of job loss is δ and, consistent with Davis and von Wachter (2011), the cost of the job loss is α percent of the value of the job, $U_{t+1} \equiv (1-\alpha)V_{t+1}$.

The equation for the value of a job can be written

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[14] Because sample sizes are relatively small for longitudinally linked data on state and local government workers, each data point in Figure 7 represents three years of non-overlapping data (for example, the point at the year 2000 in Figure 7 pools data from 1999-2001).

$$V_t = w_t + \beta[\delta U_{t+1} + (1-\delta)V_{t+1}], \quad (1)$$

where $\beta = \frac{1+g}{1+r}$, g is the annual growth in real wages and r is the real discount rate. Plugging in for U_{t+1} and iterating forward for T periods (assumed to be 20 years), where the worker faces the job loss risk each year, the value of a job is given by the expression

$$V_t = w_t \frac{(1-\beta^T(1-\delta\alpha)^T)}{1-\beta(1-\delta\alpha)} \quad (2)$$

The value of the job is therefore declining in the probability of job loss, reflecting a premium for job stability, holding the wage fixed. Therefore, for workers to be indifferent between jobs in the public and private sectors with different job loss rates (δ), the wage must adjust to compensate. In particular, we can totally differentiate the preceding equation which shows the wage increase needed to keep the value of the job fixed after the job loss probability increases by 1 percentage point.

$$\left.\frac{dlnw}{d\delta}\right|_{dV=0} = \left(\frac{\beta\alpha}{1-\beta(1-\delta\alpha)} - \frac{\alpha T\beta^T(1-\delta\alpha)^{T-1}}{(1-\beta^T(1-\delta\alpha)^T)}\right) \quad (3)$$

Using equation (3), the estimate of $\alpha = 12\%$ from Davis and von Wachter (2011), the estimates of δ from the exercise displayed in Figure 7, our annual market-based real discount rate r assumption, and an assumption of 0.02 for real wage growth g, provides us with the estimates of the value of greater job stability shown by the black line in Figure 8. We estimate that the value of better job stability peaked in 2012 at about 3.5% of wages but has declined since to a value of just 2% of wages in 2024. Thus, while S&L workers still enjoy a premium due to job stability which partially offsets lower wages, that wage premium has fallen by more than 1 percentage point from 2010 to 2024.[15]

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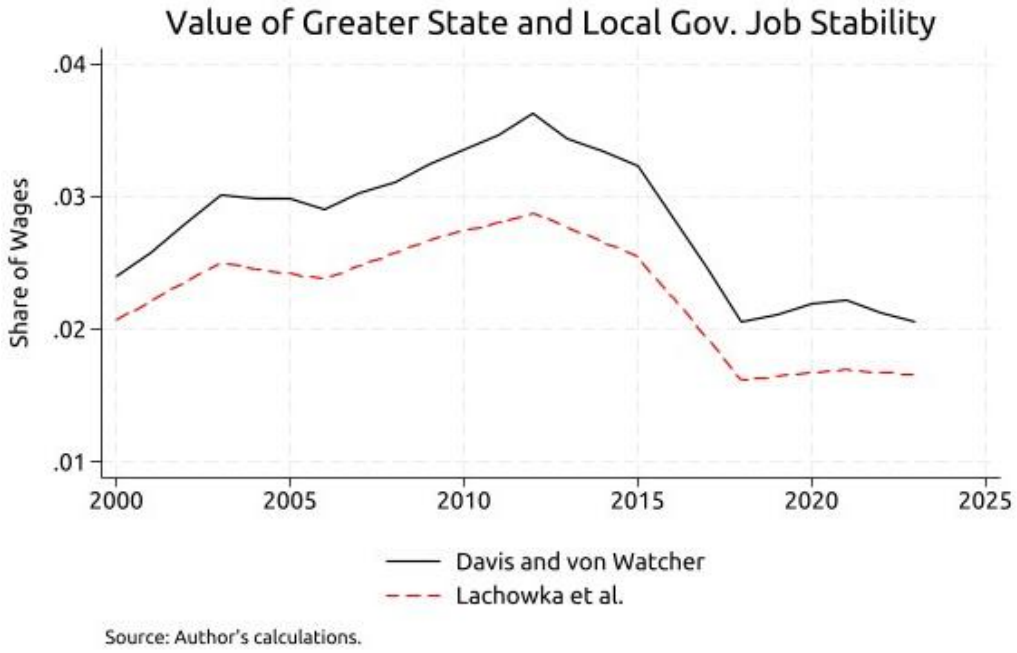
[15] While we view our simple job stability framework as a useful and transparent method to value job stability, we acknowledge there are any number of possible objections. We briefly note two. First, the framework implicitly assumes workers are risk neutral. If we incorporate risk aversion, the certainty equivalent wage premium rises quickly with the degree of risk aversion. That said, the Davis and Von Wachter (2011) empirical estimates are expressed as the PDV of wage loss, not utility; it's unclear that if *U* and *V* were measured in in our framework in utility units, rather than dollars, the Davis and Von Wachter (2011) empirical estimates would remain appropriate. Second, Davis and von Wachter (2011) focus on a particularly costly type of job loss: workers with at least 3 years of tenure who lose their job in a mass layoff event. This type of job loss likely destroys favorable matches and may also occur during downturns in local labor markets. They document that most job destruction does not meet these criteria, because many short-tenure workers lose their jobs (according to the Displaced Worker Survey, workers with less than 3 years of tenure account for about 60% of job displacement) and because not all jobs are destroyed during mass layoffs. Our measure of job loss is broader than that used by Davis and von Wachter (2011) as it includes all permanent job losers regardless of tenure (the monthly CPS does not include data on job tenure) and doesn't condition on mass layoff events. For these reasons, it is likely that the job losses we capture will be less costly than those analyzed by Davis and von Wachter (2011).

Figure 8



Source: Author's calculations.

Lachowska, Mas, Saggio, and Woodbury (2026) offer an entirely different revealed preference approach to valuing the premium from relatively lower job loss probabilities in the public sector. Lachowska et al. follow Sorkin (2018) to infer the relative worker utility from being employed by different firms by ranking firms based on how likely workers are to voluntarily move to that firm vs. leave it. Armed with an estimate of the relative utility of firms, Lachowska et al. can estimate the willingness to pay for the various attributes of the job a firm offers, e.g., for additional hours (the emphasis of their paper) as well as the probability of job loss. For example, their estimates show that a 1% higher wage is associated with 6.9% higher utility and 1 percentage point higher job retention probability—measured as the two-quarter probability of a worker keeping her job—is valued about 1.9 times as much (in utility terms) as a 1% higher wage.[16]

We can infer the job security premium of state and local government workers relative to private-sector workers using Lachowska et al.'s estimate of 1.9 in the ratio of the value of job loss vs. wages. First, we convert the relative difference in the monthly job-loss probabilities shown in Figure 7 into two-quarter differences by multiplying by 6 (to convert to the same two-quarter unit as Lachowska et al.'s); for example, public sector workers had a roughly 1.5 percentage point lower two-quarter job-loss probability in 2009. Multiplying by 1.9 yields a 2.9% wage premium for 2009. These estimates are shown by the red dashed line in Figure 8. While they are somewhat lower on average than our model-based estimates

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[16] These estimates come from Table 4 of Lachowska et al. The coefficient on job security in row 3 is 12.973 and the coefficient on wages in row 2 is 6.890; the ratio of the two—or the relative value in utility placed on a one-percentage-point increase in wages vs. a one-percentage-point lower job separation probability—is 1.9.

(black line), overall they are remarkably comparable despite being derived from an entirely different methodology.[17] [18]

IV.C Comparing Total Compensation

In this subsection, we combine the benefit value measures presented above for DB pensions, retiree health care benefits, and job stability with CPS, ECI, and ECEC data to obtain total relative compensation comparisons.

IV.C.1 Trends in Relative Total Compensation

Figure 9 shows our estimate of the ECI-based trend in public sector total compensation relative to the private sector. To generate the estimate, we aggregate the ECI data on wages and benefits excluding DB pensions with our accrual-based valuation index measures for DB pensions, retiree health care benefits, and job stability.[19] Incorporating accrual accounting for DB pensions and the inclusion of retiree health care benefits and job stability alters the trajectory of relative S&L compensation since 2010. Instead of flat relative compensation over the entire post-2010 period, as displayed using the published ECI data in Figure 1, relative compensation falls gradually from 2011 to 2018, pulled down by declining relative wages. Relative compensation then picks up in 2019 and 2020 as relative benefits surge due to a sharp decline in interest rates, which pushed up the value of public pension and retiree health benefits. Thereafter, relative compensation again falls as a sharp rise in interest rates erodes the value of public pension and health care benefits. All told, total relative compensation in 2023 is nearly 10% beneath its value in 2011.

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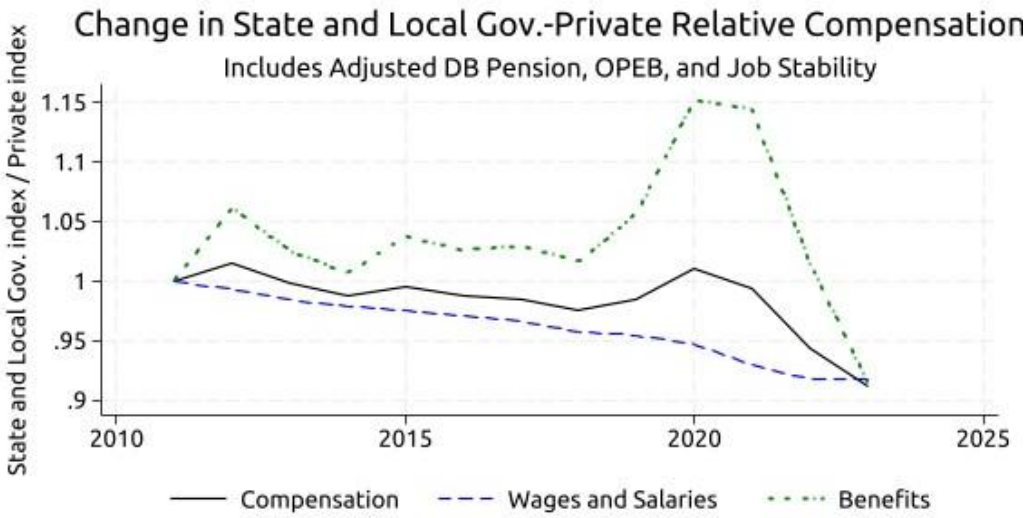
[17] We thank Alex Mas for suggesting this alternative approach for valuing job stability.

[18] Biggs (2011) used data on the premiums charged to federal and private workers for supplemental unemployment insurance premiums and finds that the value of job security was worth 2.4% of wages—similar to the estimates presented here.

[19] The ECI does not capture retiree health benefits, and there is therefore no need to remove it before the aggregation step, as is done with DB pensions.

Figure 9



Source: Bureau of Labor Statistics' (BLS) Employment Cost Index (ECI, base period December 2005) and Bureau of Economic Analysis's (BEA) National Income and Product Accounts (NIPA). Author's calculations.
Note. Annual data averaged from quarterly data. Utilizes unpublished data provided by the BLS; these data should be interpreted with care because they do not meet the BLS's standard publication criteria. Value of job stability included in benefits.

IV.C.2 Level of Relative Total Compensation

We now turn to estimating the absolute *level* of the public relative total compensation premium—which is not possible with the ECI-based approach. Our analysis is based on the following equation:

$$\frac{TC_{SL}}{TC_{Private}} = \frac{W_{SL}}{w_{Private}} * \frac{(1+benshare_{SL})}{(1+benshare_{private})} \quad (4)$$

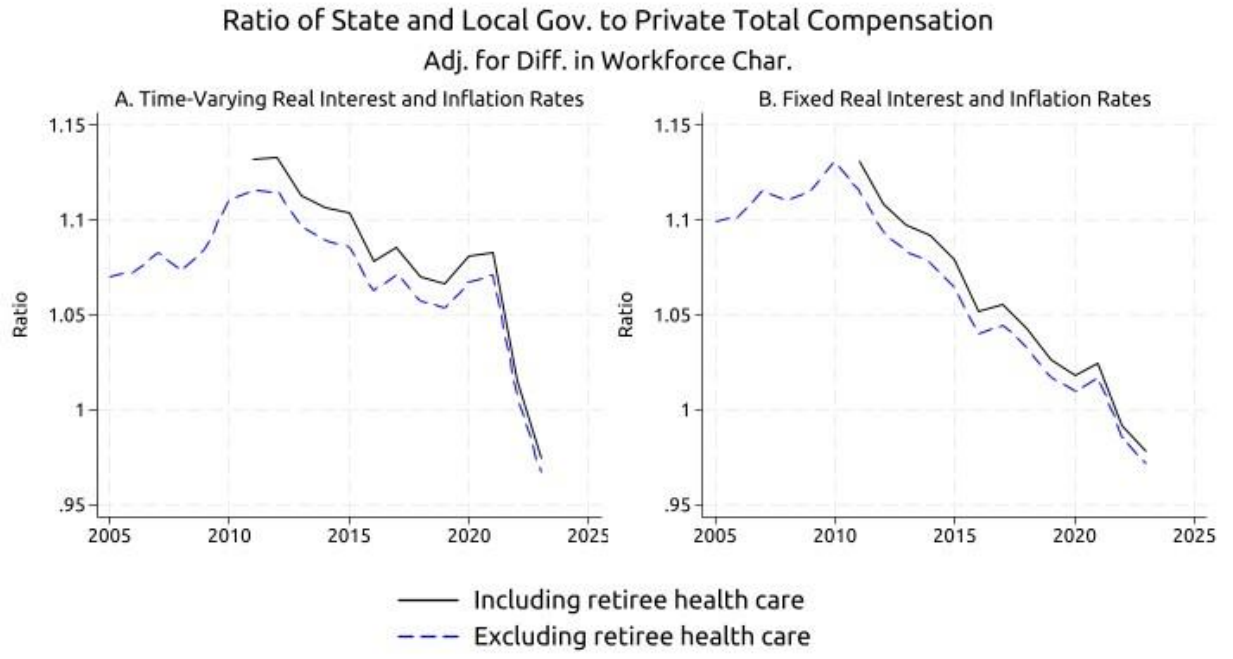
where $\frac{W_{SL}}{w_{Private}}$ is the public sector wage premium and $benshare_{SL}$ ($benshare_{Private}$) is the benefit share—i.e., the value of benefits in the public (private) sector measured as a share of wages. For the wage premium, we use our CPS-based estimate which corrects for differences in the composition of employees across the sectors (Figure 2). The benefit share is simply the sum of our various benefit measures, all expressed as a fraction of wages. For state and local DB pensions, we use our estimate of pension normal costs as a share of NIPA wages and salaries. For private DB pensions, we use the NIPA accrual-based measures of pensions and NIPA wages. For public retiree health care we use our estimates of normal costs for this benefit as a share of NIPA wages and salaries; for private we assume there are no retiree health care benefits (which biases us toward overstating the share of public compensation relative to private). For all other benefits, both public and private, we use hourly benefits as a share of hourly wages

and salaries as measured in the ECEC. Finally, for job stability we use the wage premium estimates displayed in Figure 6b.[20]

Figure 10a presents the relative total public compensation differential expressed as a percent deviation of public compensation from private compensation—i.e., a value of 1 indicates parity in compensation across the sectors. At the onset of the Great Recession, this evidence suggests that state and local workers were paid 13% more than their private sector peers. This differential, however, moves lower over the ensuing years and then plunges with the rise in interest rates—and correspondingly lower present value of relatively generous public pension and retiree health care benefits—such that, as of 2023, the public premium is a negative 3%; i.e., state and local workers were being paid 3% less than their private sector counterparts.

Figure 10b shows our calculations holding interest rates and inflation fixed at their 2024 values. Here we see a different story—a steady decline in compensation in the state and local sector relative to the private sector. While the calculations underlying Figure 10a are clearly correct from a financial accounting perspective, the declining compensation shown in Figure 10b may be more consistent with what workers perceive.

Figure 10



Source: Current Population Survey, Employment Costs for Employee Compensation, National Income and Product Accounts (NIPA), Public Plans Data, and author's calculations.

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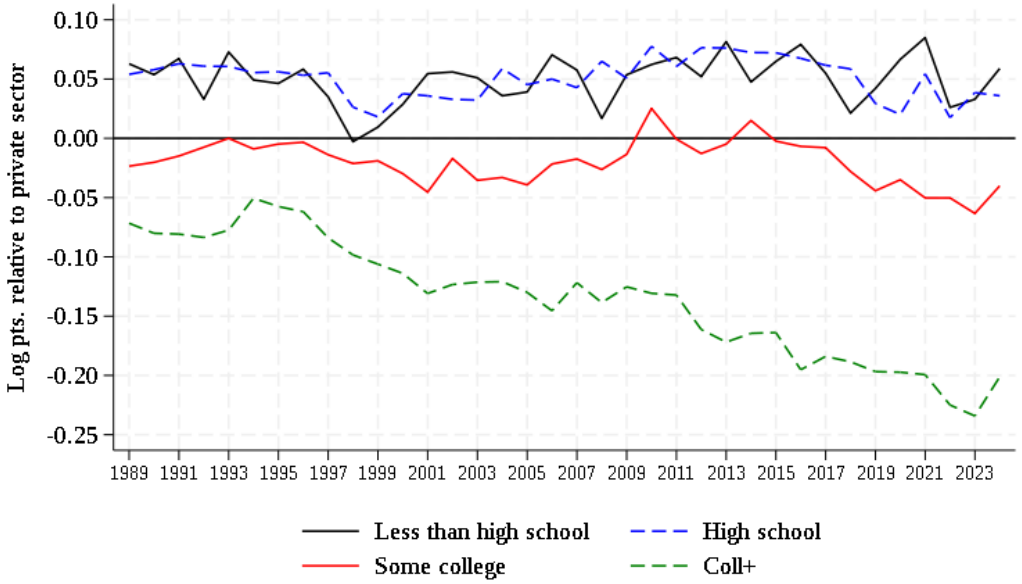
[20] NIPA wages and salaries include supplemental pay and paid leave, whereas the CPS and ECEC categorize these items as benefits. For DB pension and retiree health benefit shares, which are calculated as a percentage of NIPA wages and salaries, we adjust NIPA wages and salaries downward to remove supplemental and paid leave. The adjustment is based on the magnitude of these items relative to wages as measured in the ECEC.

IV.D Relative Compensation by Age and Education

We now turn to an examination of public-private compensation differentials by group to understand the sources of declining relative public compensation and the potential policy implications. We start by returning to the CPS-based wage differentials. The most striking finding, shown in Figure 11, is that the relative decline in wages for public sector workers is entirely concentrated within the most educated group. For those with at least a college degree, the relative wages of S&L workers were about 10 percentage points less than in the private sector in the late-1990s; that gap expanded gradually over time, accelerated after the Great Recession, and reached about negative 20 percentage points in 2024.

Because the most highly educated workers represent nearly two-thirds of state and local workers, the decline in the relative pay among that group explains a large portion of the overall decline in relative pay shown in Figure 2. However, there is both a within-group decline in relative pay for the most highly educated as in Figure 11, as well as a between-group dynamic at play. The latter arose from the fact that over this time period, the educational attainment of workers in both sectors increased. Given that college-educated workers have the largest negative pay differential across educational groups, this compositional shift would increase the negative pay premium in the S&L sector as a whole, even if the college-educated public-private pay differential had not widened. When we hold the educational shares of workers in the S&L sector constant at their 2005 values, we find that roughly half of the overall decline in the S&L pay premium was due to the decline in pay of the most highly educated (the green line in Figure 11) and about half due to the rising educational attainment of workers.[21]

Figure 11
S&L government wages relative to private sector by education



Gap in wages for S&L workers relative to private. Regressions are run separately for each education group and are weighted and control for a quadratic in age, metro status interacted with nine Census divisions, race interacted with gender, marital status, and major occupation.

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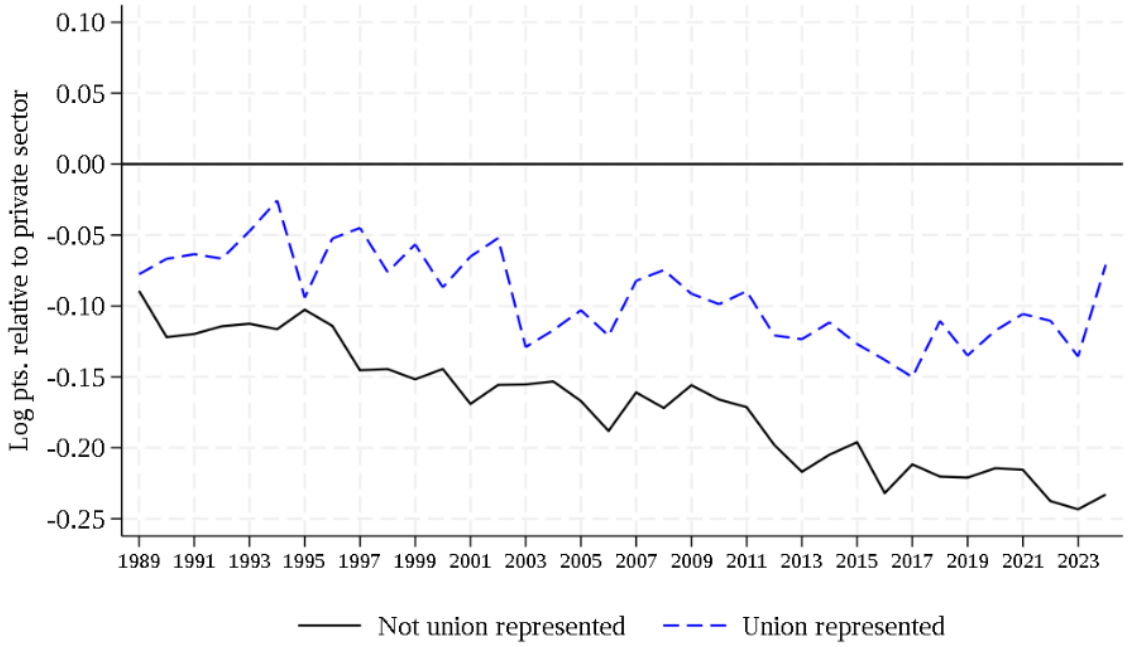
[21] In principle, the decomposition of the change over time in the *relative* pay of two sectors could be calculated in a number of ways which may give somewhat different results. For the purposes of this section, we take the coefficients in Figure 11 as fixed and hold the shares of workers by education in the S&L sector fixed in 2005, allowing the share of workers by education to adjust in the private sector.

It appears that relative wages among the more highly educated have fallen, especially for those who are not covered by unions. Figure 12 shows little relative decline in pay of those represented by a union (the blue dashed line) since 2000, whereas the relative pay of those who are not represented by a union saw pay decline by roughly 10 percentage points, from roughly 2000 to 2024. This result is broadly consistent with Baker, Halberstam, Kroft, Mas, and Messacar (2024), who show that unionization boosts the wages of highly educated public sector workers.

Figure 12
S&L government wages relative to private sector by union status
Restricted to those with at least a college degree



Gap in wages for S&L workers relative to private among those with at least a college degree. Regressions are run separately for each age group and are weighted and control for a quadratic in age, metro status interacted with nine Census divisions, race interacted with gender, marital status, major occupation, and union status.

Finally, Figure 13a-b presents the relative total public compensation differential level—comparable to the results in Figure 10—for workers with a college degree (13a) and for workers with only a high-school education (13b). We assume that the ratio of pension benefits to wages and the value of job stability as a share of wages does not vary by education, but that the ratio of the value of retiree health insurance does—since the cost of the benefit is a fixed amount that is more valuable as a share of wages for lower-wage workers.[22] The results suggest that relative pay of college-educated public workers has declined by about

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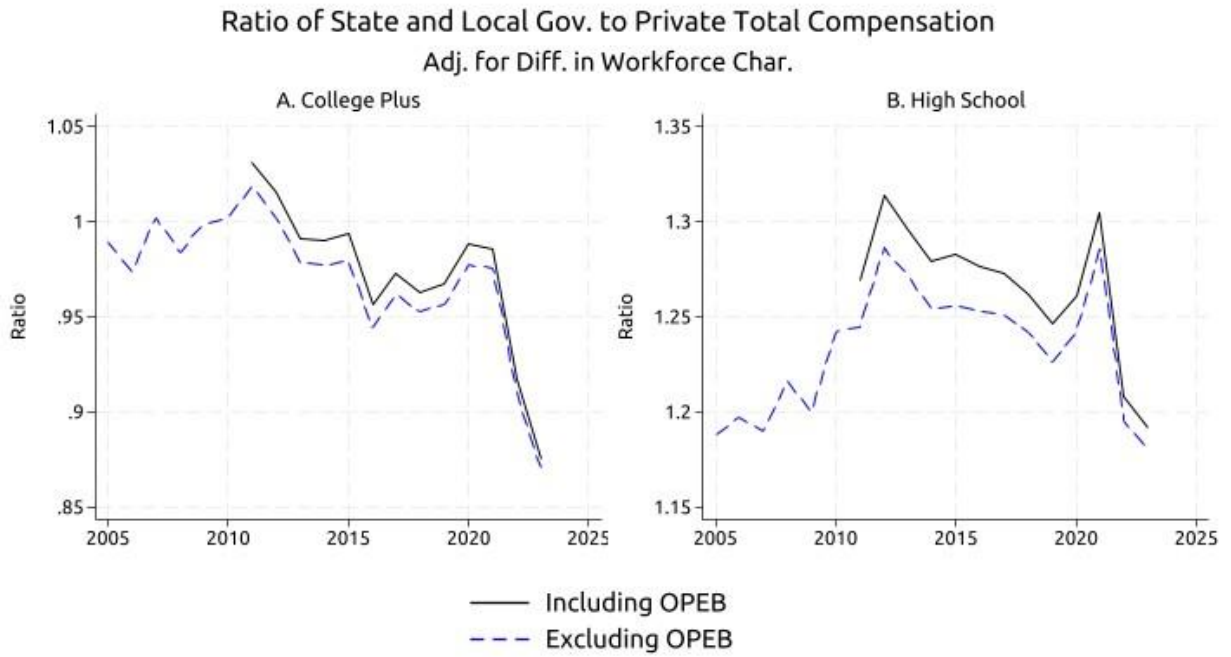
[22] We scale the retiree health care portion of $benshare_{SL}$ in equation (4)—i.e., the aggregate normal cost for retiree health care—by the ratio of the average wage for all workers to the average wage for the given education group. This adjustment ensures workers in all education groups are credited with the same dollar value of benefits. Relative to not making the adjustment, the value of retiree health care is reduced for workers with a college degree and boosted for workers with only a high school degree.

16 percentage points since 2011; in 2023, college plus workers experienced a pay deficit of 13% relative to their private sector counterparts. For high school-educated workers, the decline has been much less steep—about 8 percentage points, and these workers still enjoy a bonus of 19% for working in the S&L sector.

Figure 13



Source: Current Population Survey, Employment Costs for Employee Compensation, National Income and Product Accounts (NIPA), Pew (2023), Public Plans Data, and author's calculations.

IV.E Relative Compensation by State

We conclude our analysis with an examination of the public-private compensation differentials across states. To do so, we use our CPS-based methodology as the CPS provides state-level identifiers. Our methodology for calculating the differentials is generally the same as for the national estimates, but with the data aggregated to the state-year level instead of year.[23]

That said, there are a few important deviations. In order to smooth through year-to-year variability in the public-private wage differential estimates, which can be substantial for smaller states, we use a three-year backward moving average of these estimates. Given that the PPD pension plan data is quite thin for some states—e.g., some have only a single pension plan in the dataset—we eschew our actuarial model approach. Instead, we adjust BEA state-level public pension normal cost estimates to our preferred, market-based inflation and real interest rate assumptions using actuarial rules of thumb. This is an inexact approach, particularly given the methodological problems we have identified with the BEA normal

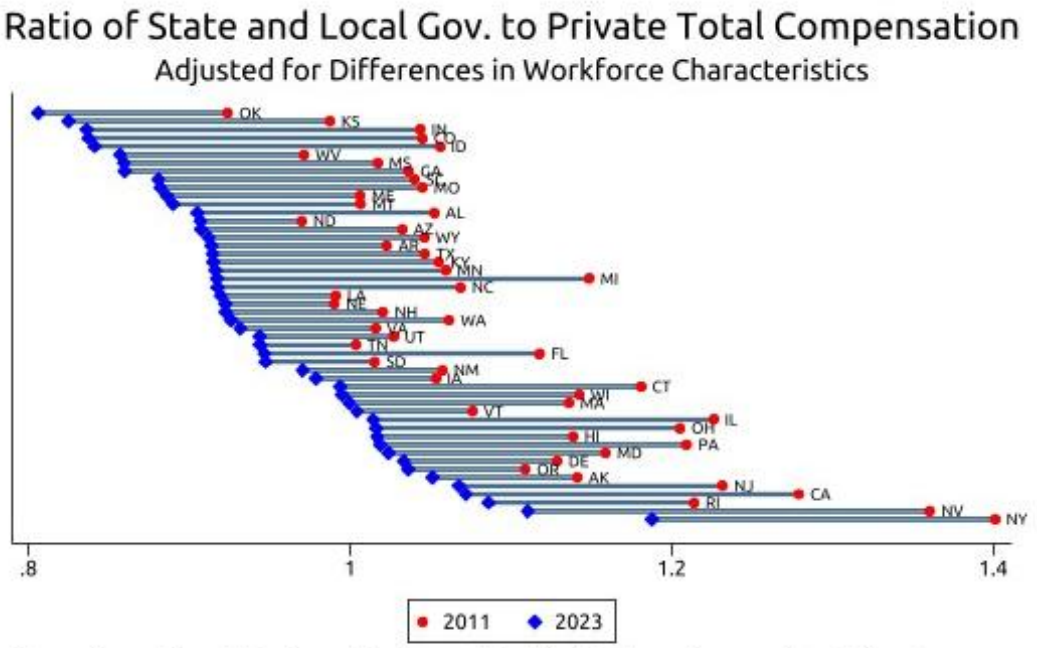
. . .

[23] The NIPA provide measures of state and local government pension normal costs disaggregated by state. For retiree health benefits, we rely on Pew (2023) and state government actuarial reports.

cost estimates. Nonetheless, when these state-level normal costs estimates are aggregated to the national level, they are roughly similar to our more rigorously constructed national-level estimates. (See Appendix section B.6 for more information.) We lack any state-specific information on benefits other than pensions and retiree health care and therefore assign each state the national per-worker values calculated above for other benefits. Fortunately, outside of health insurance coverage, these other benefits tend to vary less across employers than do pension and retiree health care benefits (Biggs 2014). The lack of state-specific information on health insurance coverage, though, is a limitation. Similarly, lacking information on private sector pensions by state, we assign our national-level estimates to each state. As private pension benefits have become a small share of overall private sector compensation, state-specific information on this benefit would not significantly alter our conclusions. Finally, we assign our national relative job stability values to all states.

Figure 14 summarizes the state-level results and yields two primary conclusions. First, there is substantial heterogeneity across the states in relative public sector compensation. In 2023, displayed by the blue dots to the left, it ranged from a high of a 19% premium in New York to a 20% penalty in Oklahoma. A majority of states, 35, have compensation penalties. Second, all states have experienced a decline in relative compensation, and the decline is substantial in most cases. Thus, our firmest national-level conclusion—that relative public sector pay has eroded significantly in recent years—also holds for the more disaggregated state analysis.

Figure 14

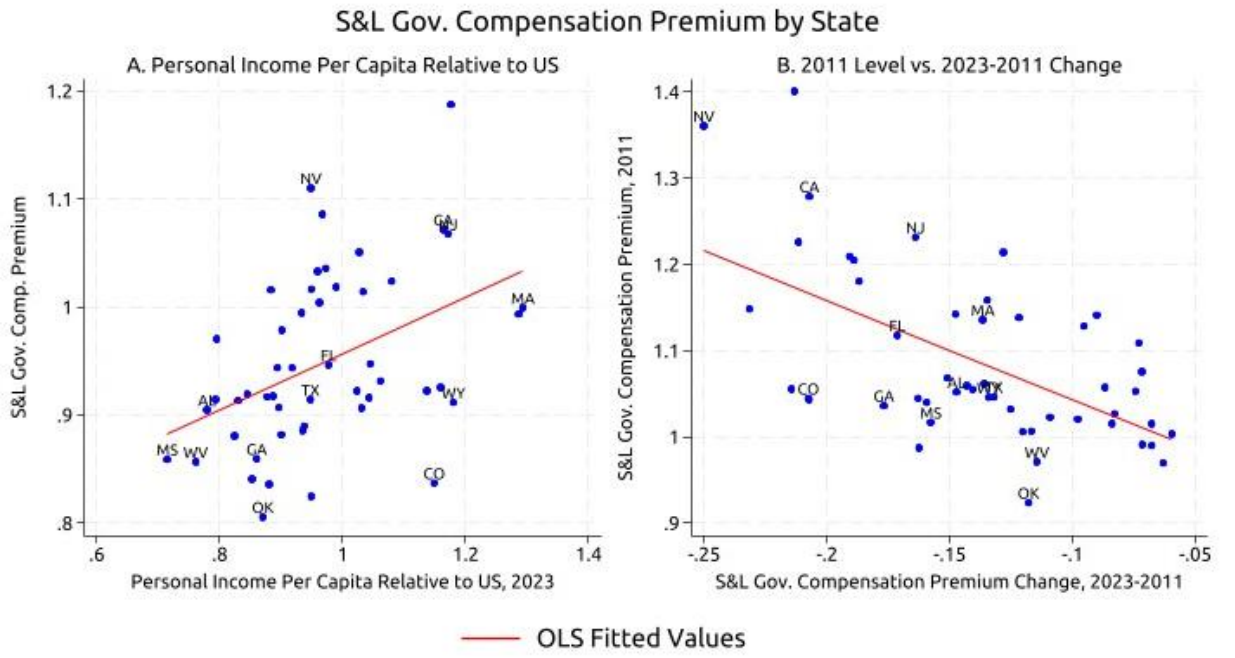


Source: Current Population Survey, Employment Costs for Employee Compensation, National Income and Product Accounts (NIPA), Pew (2023), Public Plans Data, and author's calculations.

Figure 15 unpacks the state results a bit further. Panel a documents that high-income states, measured in terms of personal income per capita, tend to have the highest relative public sector compensation. That

said, there are a few notable exceptions—e.g., Colorado is among the highest income states but has one of the lowest levels of relative public compensation. Panel b reveals a strong relationship between initial 2011 relative pay and the subsequent decline through 2023: States with the largest 2011 relative public sector pay experienced the largest subsequent declines.

Figure 15



Source: Current Population Survey, Employment Costs for Employee Compensation, National Income and Product Accounts (NIPA), Pew (2023), Public Plans Data, and author's calculations.

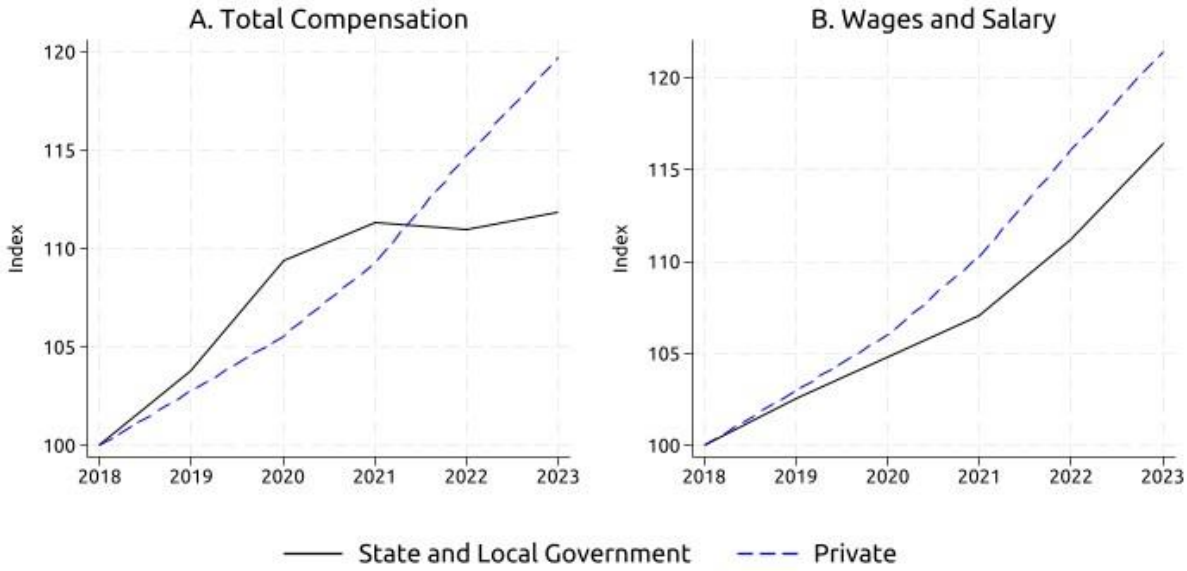
V. Conclusion

Our analysis demonstrates that the compensation premium for state and local government workers relative to comparable private sector employees has eroded significantly over the past 15 years. This relative decline is concentrated among workers with at least a college degree. Although our estimate of the relative level of compensation across the sectors is subject to elevated methodological uncertainty compared to our estimate of the relative trend, we find that college-educated workers' total public sector compensation is now around 13% beneath that in the private sector. Our findings are relevant to the future ability of state and local governments to recruit and retain qualified personnel and, correspondingly, to achieve the administrative capacity to effectively implement policies and deliver essential services. Although a detailed exploration of these implications is beyond the scope of this paper, we offer two suggestive pieces of evidence that the decline in relative compensation may be negatively affecting public sector hiring, as opposed to, for example, decreasing the rents available to public sector employees.

First, as the economy emerged from the COVID pandemic, private sector compensation accelerated, whereas state and local government compensation initially did not. This dynamic can be seen in Figure 16,

which presents both our total compensation measure (panel a) and our wage measure (panel b), both indexed to equal 100 prior to the pandemic in 2018. By 2023, private sector compensation had surged 8 percentage points above state and local compensation; private wages had shot up 5% above state and local wages. Over the same period, state and local payrolls recovered only very slowly from their decline at the onset of the pandemic, despite the budgets of these governments being historically strong (Cashin, Lutz, Peterman, Ratner, and Rodman 2023). In contrast, private sector payrolls recovered more quickly.[24] Thus, it appears that the failure of the public sector to increase compensation in line with the market allowed private firms to outcompete them for workers.

Figure 16
Compensation and Wages Following the Covid Pandemic



Source: Bureau of Labor Statistics' (BLS) Employment Cost Index (ECI) and Bureau of Economic Analysis's (BEA) National Income and Product Accounts (NIPA). Author's calculations.
Note. Annual data averaged from quarterly data and re-indexed to equal 100 in 2018. The left panel utilizes unpublished data provided by the BLS; these data should be interpreted with care because they do not meet the BLS's standard publication criteria.

Second, several aspects of the labor market landscape for K-12 teachers suggest negative impacts of declining relative pay in the public sector. There has been a surge in teacher strikes in recent years with grievances over compensation a key cause (Lyon, Kraft, and Steinberg 2024). There has also been a decline in observable teacher credentials—e.g., the share of novice teachers has increased (Garcia and Weiss 2019). Perhaps most notably, occupational prestige, interest among students, the number of individuals preparing for entry, and on-the-job satisfaction for K-12 teachers all began a sustained decline around 2010 and have now fallen to at or near their lowest level in 50 years (Kraft and Lyon 2024). The

. . .

[24] Private employment exceeded its February 2020 level in April of 2022, whereas state and local government employment did not reach this threshold until October 2023.

decline in relative public sector wages and compensation we document coincides neatly with this decline in the teaching profession and is therefore a potential cause.

While both the COVID era and K-12 teaching episodes are consistent with declining relative public sector pay infringing on the ability of state and local governments to hire and retain high quality workers, they are far from definitive. Future research should aim to estimate the causal connection between relative public sector pay and public sector labor market outcomes.[25]

. . .

[25] Bates and Johnston (2025) suggest that pension compensation does not appear to influence teacher retention in the late stages of teaching careers. More work in this vein, considering a broader set of outcomes and aspects of public sector compensation, is needed.

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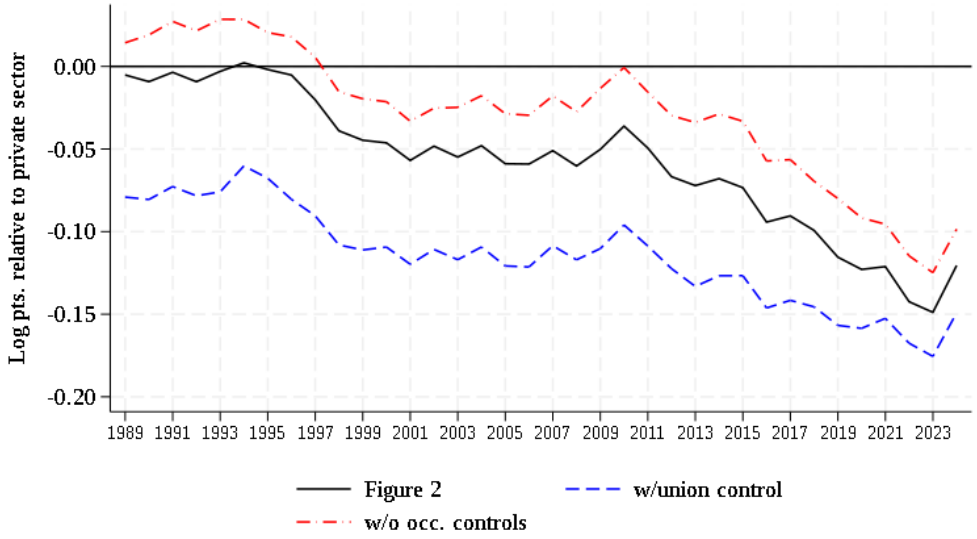
APPENDIX

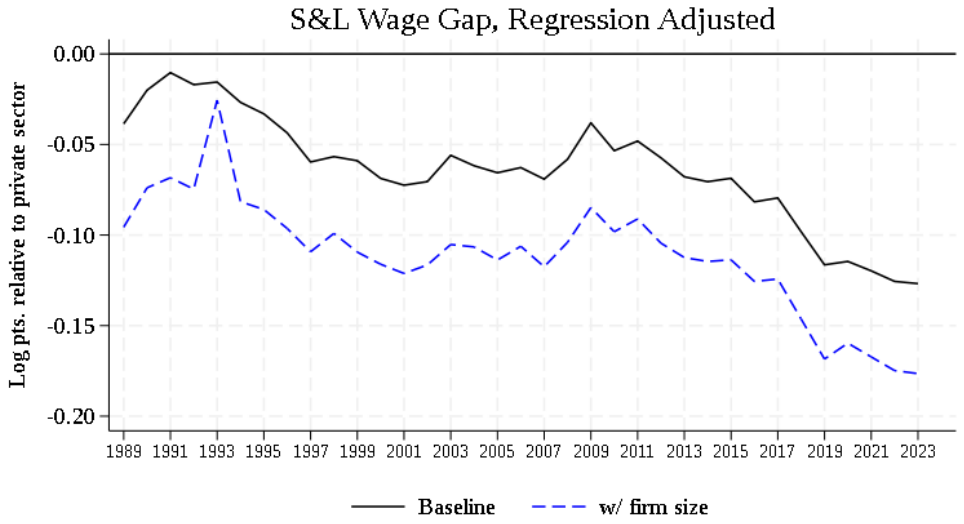
A. Additional Figures and Tables

Figure A1
Panel a



Gap in wages for S&L workers relative to private. Regressions are weighted and control for quartic in age, four education groups, metro status interacted with Census division, race interacted with gender, marital status, and major occupation. Dashed line regression includes baseline controls in Figure 2 plus a control for union status.

Panel b



Gap in wages for S&L workers relative to private. Regressions are weighted and control for quartic in age, four education groups, metro status interacted with Census division, race interacted with gender, marital status, and major occupation. Full-time is defined as usually working 35 or more hours per week.

Appendix Table 1: OPEB Reforms					
	State share of Aggregate Normal Cost	Reform 1	Reform 1 Affects	Reform 2	Reform 2 Affects:
Texas	16%	Reduced Contribution for Workers with less than 20 years of service (State pays 50% for 10-14, 75% for 15-19)	Employees with <5 years of service as of 9/1/2014	Increased Required years of Service from 5 to 10	Employees hired after 9/1/2009
California	15%	State lowered share of premium from 100/90 (retiree/dependent) to 80/80	Employees Hired after 1/1/ 2017 or 1/1/2016, depending on bargaining unit	Began requiring certain actives to contribute; these contributiosn have increased and become more common over time. Also Increased min retirements age from 50 to 52	Employees hired after 1/1/2013
New Jersey	12%	Increased retiree share of premium, with increase largest for those with the largest pensions.	Employees/retirees with less than 20 years of service as of 6/2011		
Illinois	9%	Required retirees to pay 1% of pension benefit if on Medicare, 2% if pre-Medicare as health insurance premium; increased to 2%/4% in 2015.	2014-2015. Then ruled unconstitutional.		
New York	9%	NONE			
North Carolina	6%	Eliminated Retiree Health Insurance - reduced from 50% premium (10-20 years of service) and 100% premium (20+) paid by state	Employees hired after 1/1/2021		
South Carolina	3%	Reduced share of premium paid by employer and increased years of service required to get subsidy	Hired after 5/1/2008		
Massachusetts	3%	Lowered state contribution from 80% to 75% of premium	Employees hired after 7/1/2003		
Georgia	3%	Reduced subsidy for those with less than 30 years of service, with contribution rangins from 15% of premium (10 years 0f service) up to 75% (30 years of service).	Employees with <5 years of service as of 1/1/2012		
Connecticut	3%	Increased # years of contributions from 10 to 15.	Employees hired after 7/1/2017	State required employees to contribute 3% of pay toward retiree health for a period of 10 years	First employees with <5 years of service as of 2009, then expanded to all employees in 2011.
Ohio	2%	Reduced employer contribution to fixed $ amount (based on years of service and Medicare/non-Medicare) rather than share of premium.	2022 all retirees		
Florida	2%	Florida maintained $5 nominal contribution per month per year of service from 2001 to 2023; raised it to $7.50 7/1/2023			
Total	83%				

Note: Share of Aggregate Normal Costs from Pew (2023) on total state and local government retiree health care by state (adjusted to our discount rate), but reforms documented are for the state plans alone.

B. Estimating Defined Benefit Pension Accruals

B.1 Overview

Our goal is to construct an accrual-based measure of state and local government defined benefit (DB) pension compensation that can be used in place of the cash-based pension measure in the ECI. Our approach is as follows. We begin with the Public Plans Database (PPD), maintained by the Center for Retirement Research at Boston College, which provides an annual panel dataset covering the near-universe of state and local government pension plans, or roughly 95% of liabilities. The PPD includes information on plan normal costs and the actuarial assumptions used to calculate them, including the

nominal interest rate and expected inflation. We hand-collect or impute missing observations on normal costs and covered payroll as needed to construct a balanced panel of plans.[26]

Next, we adjust the plan-reported normal costs so that they reflect market-based estimates of interest rates and expected inflation, rather than the assumptions used by individual plans. Section B.2 describes our methodology for making these adjustments. We then aggregate the adjusted plan-level normal cost rates—the normal cost as a share of covered payroll—using plan payroll weights. The resulting adjusted normal cost rate captures changes over time in the value of DB pension accruals as well as changes in the size of covered payroll. We then convert this rate from the fiscal-year basis used by plans to the calendar-year basis used by the NIPA and ECI.[27]

We then calibrate our normal cost estimates by setting them equal to the BEA's measure of state and local government DB pension normal costs in 2011, when our market-based assumptions are close to those used by the BEA (Section B.3 discusses how our assumptions differ from the BEA's in other years). In doing so, we ensure we capture the full universe of state and local pension plans.

We then grow the 2011 value forward and backward in time by the sum of the percent change in the adjusted normal cost rate and the percent change in state and local government wages as measured by the ECI. This approach has the virtue that growth in DB pension benefits caused by wage growth are calculated using the ECI wage measure which, by construction, does not capture wage changes due to compositional change within the public workforce. While the normal cost rate may potentially change over time due to compositional change in the workforce, the scope for such change is relatively limited because there is less variation across worker type in the normal cost rate as compared to other elements of compensation (such as wages or normal costs). Thus, the change over time in our measure of DB pension compensation is less a function of change in workforce composition than would be the case if we simply used annual measures of the normal cost. As discussed in Appendix section D below, this renders our pension compensation measure relatively more appropriate for aggregating together with non-pension elements of the ECI.

Finally, to calculate the value of employer pension compensation, we subtract employees' contributions to the pension plan from the total normal cost and add the plan's administrative expenses, using BEA estimates for both items. This accounting follows the BEA's treatment of implicit compensation from defined benefit pension plans.

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[26] In particular, we hand collect missing payroll numbers for the California Public Employee Retirement Fund, and missing normal costs for the New York State Teachers, Public Employees, and Police and Fire plans, and for New York City Teachers. We impute missing values for the following items: membership, employer contributions, employee contributions, and normal costs. Our imputation proceeds in two steps. First, we use linear interpolation within each pension plan panel. Second, we calculate growth rates by year for three types of pension plans: teacher, public safety, and other. We then apply these growth rates to impute the remaining missing values (i.e., we grow the last non-missing out forward or backward in time).

[27] State and local government fiscal year *g* most commonly runs from July 1 of calendar year *g-1* through June 30 of calendar year *g*. We therefore approximate calendar year *g* using a two-year forward moving average of fiscal year *g* and fiscal year *g+1*. For example, a calendar year 2000 variable is set equal to the average of the fiscal year 2000 and fiscal year 2001 variables. Fiscal year 2000 runs from July 1, 1999 through June 30, 2000 and fiscal year 2001 runs from July 1, 2000 to June 30, 2001. Thus, the numerator of the mean used to approximate calendar year 2000 contains all of calendar year 2000, plus one-half of calendar 1999 and one-half of calendar 2001.

B.2 Adjusting Plans' Reported Normal Costs

As discussed in the main body of this paper, the normal costs reported by the plans are calculated using a higher rate of return than economists view as appropriate given the riskiness of future pension benefits, and so we adjust these reported normal costs upward using the benchmark actuarial model described below. We use the Moody's Seasoned AAA corporate bond yield as our nominal discount rate and the Cleveland Federal Reserve's 10-year expected inflation estimate as our inflation measure. We allow the nominal interest rate, the real interest rate, and expected inflation to all vary annually. As can be seen in Figures A2a, A2b, and A2c, the plans' assumptions, on average, tend to differ from our market-based estimates. The figures also compare our assumptions to the BEA's, a topic we return to in Section B.3.

We use a simple, streamlined pension accounting model to adjust plans' reported normal costs to our preferred assumptions. Specifically, we calculate the normal cost of a benchmark pension plan under our preferred discount rate and inflation assumptions and under the assumptions used by the plans, and then use the ratio of those two values as a multiplicative adjustment factor.

Our model is as follows: Let g be the real growth of wages, r be the real interest rate, ea be the entry age (i.e., the age at which employment with the pension providing government begins), T be the retirement age, m be the mortality rate after retirement, $cola$ be the annual rate of increase in benefits after retirement, $\propto$ be the first-year benefits share of final wages, and π be the rate of inflation. For purposes of exposition, we assume m is constant, but our actual calculations use age-specific mortality rates from the Social Security Administration.

We use the entry age normal actuarial approach to measure the accrual value of pension benefits—the method used by the large majority of the plans and also by the BEA in its pension accrual measure. This method assumes that benefit accruals are a constant share of wages equal to the present value of benefits earned currently and in the future divided by the present value of current and future wages. This is the share of wages that, if invested in a fund that earned the assumed discount rate over an employee's tenure, would be sufficient to fully fund the benefit.

Let

$$\delta = \frac{(1+g)}{1+r}$$

and

$$\theta = \frac{(1-m)(1+cola)}{(1+r)(1+\pi)}$$

Then,

$$PV\ of\ wages = w_{ea} \sum_{t=ea}^{T-1} \delta^{t-ea} = w_{ea} \frac{1-\delta^{T-ea}}{1-\delta}$$

$$PV\ of\ benefits = \propto w_{ea} \delta^{T-ea} \sum_{t=T}^{\infty} \theta^{t-T} = \frac{\propto w_{ea} \delta^{T-ea}}{1-\theta}$$

The normal cost is thus

$$Normal\ Cost = \frac{PV\ of\ Benefits}{PV\ of\ Wages} = \frac{\frac{\propto w_{ea}\delta^{T-ea}}{1-\theta}}{w_{ea}\frac{1-\delta^{T-ea}}{1-\delta}} = \propto \frac{\delta^{T-ea}}{1-\delta^{T-ea}}\frac{1-\delta}{1-\theta}$$

We assume an effective entry age of 36 years and a retirement age of 61, implying 25 years of service at retirement. A retirement age of 61 with 25 years of tenure matches the reported averages in the PPD for the small set of plans that report this information. While typical entry ages are earlier (using the average age of state and local workers with less than 1 year of tenure from the CPS), many workers leave before retirement and get no or reduced benefits. Our representative-worker assumption is meant to capture this attrition in a simple way: Rather than modeling pre-retirement separations and mortality explicitly, we choose an effective entry age that produces average service at retirement. We therefore do not separately apply mortality adjustments before retirement. We use the Social Security actuaries' assumed mortality rates from age 61 on. Wages rise 1 percentage point faster per year than inflation. We discuss our assumption for the COLA below.

With these assumptions, and with a real interest rate of 2.5% and an inflation rate of 2%, our model produces an effective duration with respect to the real discount rate of 23 years—that is, a 1 percentage point increase in the real discount rate lowers the normal cost by around 23%—the value typically cited in the literature (e.g., Munnell and Aubry (2016), Center for Retirement Research (2022), and Winklevoss (1977)). A reduction in the real discount rate raises the normal cost rate because benefit cash flows extend further into the future than wage cash flows, so the present value of benefits increases more than the present value of wages.[28]

Normal costs respond far less to inflation-driven changes in nominal discount rates than to movements in real discount rates. For a pension plan with a COLA fully indexed to inflation and wages that also rise with inflation, a decrease in expected inflation leaves the normal cost unchanged: The present value of wages is unaffected because slower growth of wages is offset by a lower discount rate, and the present value of benefits is similarly unaffected.[29] This can be seen in the formula for the normal cost: θ is unaffected by inflation if $cola = \pi$.

If a plan's COLA does not move with inflation, then a higher discount rate due to higher inflation does lower the normal cost, but the effect is much smaller than for a change in the real discount rate. Higher inflation raises nominal wages and therefore the nominal benefit at retirement, offsetting much of the impact of the higher nominal discount rate on the present value of benefits. At a 2.5% inflation rate and 2.5% real interest rate, we find an effective duration of about 12 years for changes in inflation—that is, a 1 percentage point increase in inflation for a plan with a COLA that is invariant to inflation lowers the normal cost by 12%.

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[28] The size of the effect depends on the duration of the normal cost, which is itself equal to the difference between the duration of benefits and the duration of wages.

[29] E.g., assuming for simplicity that wage growth arises only due to inflation, the present discounted value of wages equals $\sum_{t=1}^{n}\frac{w_1*(1+i)^t}{(1+r)^t(1+i)^t} = \sum_{t=1}^{n}\frac{w_1}{(1+r)^t}$ where time *t=1* is the year of entry into employment, *n* is the year of retirement, *r* is the nominal interest rate, and *i* is the expected inflation rate.

However, many plans do link their COLAs to inflation—the most common COLA is inflation up to a cap, usually 3% (sometimes lower, especially for new hires). For example, Brainard and Brown (2025) examine large state plans and show that a bit less than 50% have inflation-linked COLAs.

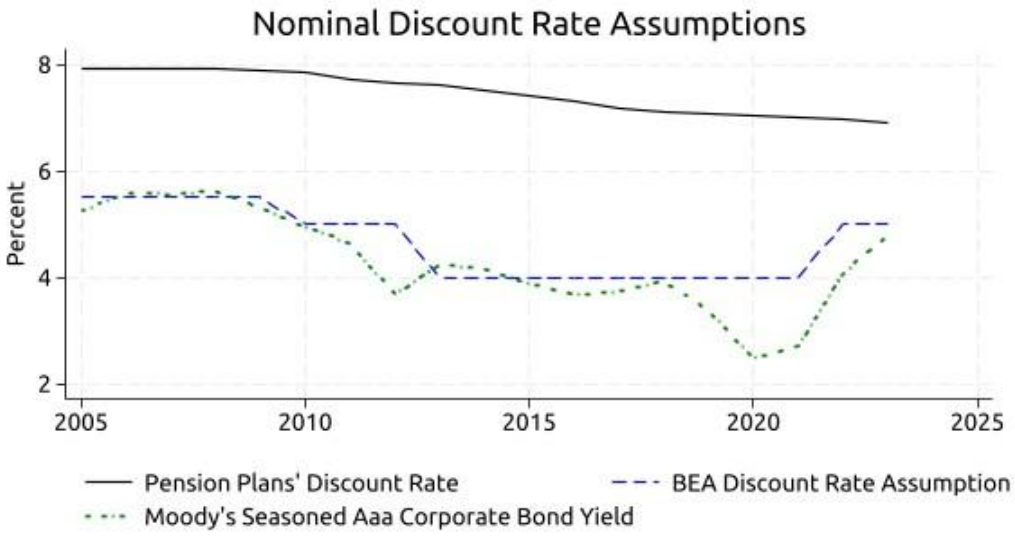
In our simple model, we assume that the COLA is 2.5% when inflation is 2.5% and that it moves 0.5 percentage point for each 1 percentage point deviation in inflation from 2.5%. This gives us a duration with respect to inflation of about 6 years.

In order to adjust the plan data to be aligned with market-based assumptions, we first estimate the normal cost in our pension accounting model under the plan's assumed nominal discount and inflation rates ($Model\ NC_{plan\ rates}$) and under our preferred rates $Model\ NC_{mkt\ rates}$. We adjust the plan's reported normal cost as reported in the PPD ($Normal\ Cost_{plan}$) by the ratio of the two. That is, our adjusted normal cost estimate ($Normal\ Cost_{adj}$) is derived as:

$$Normal\ Cost_{adj} = Normal\ Cost_{plan} * \frac{Model\ NC_{mkt\ rates}}{Model\ NC_{plan\ rates}}$$

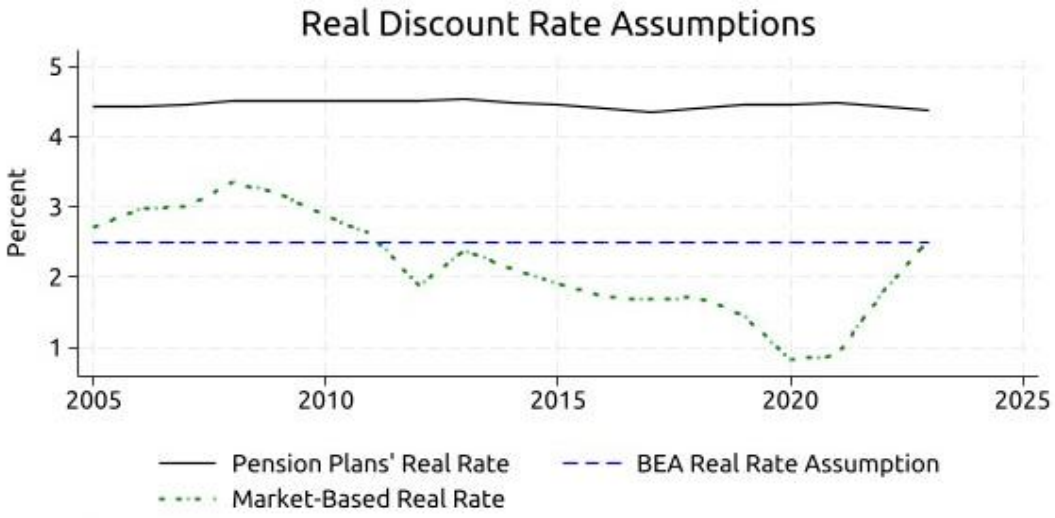
This method allows us to model differences in inflation and real rates simultaneously, rather than using simple rules of thumb based on duration.

Figure A2a



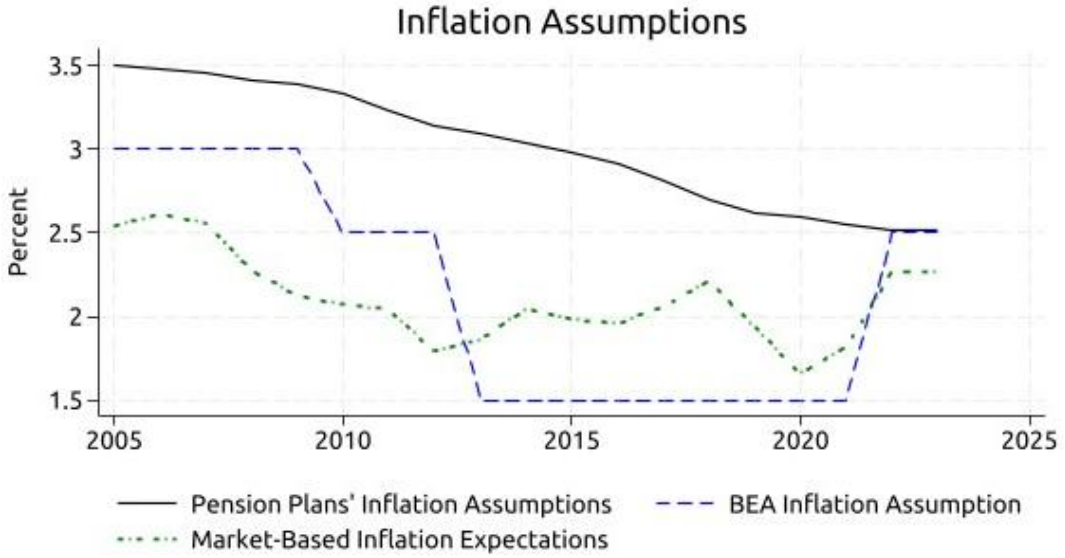
Source. Public Plans Data: Center for Retirement Research at Boston College, MissionSquare Research Institute, National Association of State Retirement Administrators, and the Government Finance Officers Association. Bureau of Economic Analysis (BEA). Moodys Investor Service. Author's calculations.
Note. Pension plans' discount rate assumption is calculated as an annual weighted mean, with weights equal to total plan normal cost.

Figure A2b



Source. Public Plans Data: Center for Retirement Research at Boston College, MissionSquare Research Institute, National Association of State Retirement Administrators, and the Government Finance Officers Association. Bureau of Economic Analysis (BEA). Moodys Investor Service. Cleveland Federal Reserve Bank. Author's calculations.
Note. Pension plans' real rate assumption equals the weighted average of each plan's nominal discount rate minus their assumed inflation, with weights equal to total plan normal cost. The market-based real rate equals Moody's Seasoned Aaa Coporate Bond Yield minus 20-year expected inflation rates.

Figure A2c



Source. Public Plans Data: Center for Retirement Research at Boston College, MissionSquare Research Institute, National Association of State Retirement Administrators, and the Government Finance Officers Association. Bureau of Economic Analysis (BEA). Cleveland Federal Reserve Bank. Author's calculations.
Note. Pension plans' inflation assumption is calculated as an annual weighted mean, with weights equal to total plan normal cost. Market-based Inflation Expectations are 20-year Expected Inflation Rates.

B.3 Comparing Our Normal Cost Measure to That Used by the BEA

The BEA converted to an accrual measure of compensation for defined benefit pension plans in 2013. However, as noted in the main body of this paper, we chose not to use the BEA's published normal costs, for the reasons detailed below.

The BEA methodology, described in Reinsdorf, Lenze, and Rassier (2014) and Lenze (2013), is very similar to ours. In particular, the BEA collects information from the largest pension plans—covering approximately 90% of the universe of plans in terms of active membership and assets (Lenze, 2013)—and uses a prototypical pension plan like ours to adjust plans' reported normal costs to the BEA's preferred discount rate, which they also peg to the AAA corporate rate.[30] That said, the BEA does not reset its discount rate to market rates each year. Instead, its nominal discount rate typically remains fixed for several years and is adjusted only after market rates have moved substantially away from the existing assumption, as shown in Figure A2a.

Our concern with the BEA measure is motivated by the fact that its aggregate normal costs appear too insensitive to changes in discount-rate assumptions. Widely used actuarial rules of thumb suggest that the effective duration of public pension normal costs is equal to 20 to 2—e.g., for every 1 percentage point decrease in the nominal discount rate, the value of public pension normal costs will increase by 20% to 25% (Munnell and Aubry 2016, Center for Retirement Research 2022, and Winklevoss 1977).[31] However, a back-of-the-envelope calculation suggests that when the BEA changes its discount rate assumption, the implied duration is only 14.[32] Consistent with this observation, Biggs (2022) notes that the BEA's state-level pension estimates for fiscal year 2018 are low relative to estimates in pension plan actuarial reports for the small number of state plans that publish normal cost sensitivity analysis.

Our reading of the BEA methodology is that this low sensitivity reflects two related assumptions. First, when decomposing the assumed nominal discount rate into a real rate and an inflation rate, the BEA explicitly fixes the real rate at 2.5% (Lenze 2013); thus, variation in the BEA's nominal discount rate over time solely reflects variation in assumed inflation, defined as the difference between the assumed nominal discount rate and 2.5%. The 2.5% real interest rate assumption was consistent with market-based estimates at the time the BEA introduced accrual accounting for pension liabilities in 2013, but became much less so over time as real rates declined prior to the COVID pandemic (see Appendix Figure A2b). As discussed above in Appendix section B.2, normal costs are significantly more sensitive to changes in the real discount rate than to changes in assumed inflation. Thus, by holding the real rate fixed and treating changes in nominal rates as changes in expected inflation, the BEA methodology makes normal costs too

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[30] BEA's prototypical plan has a bit more detail than ours, including a distribution of wages by tenure and turnover rates.

[31] The duration of a bond (or liability) is the present value–weighted average number of years until payments are made. Effective duration approximates the slope of a bond's value as a function of changes in interest rate. See McCaulay (2013).

[32] Since 2000, the BEA has changed its discount rate assumption four times. These four instances provide an opportunity to very roughly assess the responsiveness of the BEA's pension normal costs to changes in the discount rate. In particular, using the BEA's implied normal cost rate (the normal cost divided by wages and salaries), a detrended percent change is calculated for the year t of a discount rate change as the percent change from year t to year t-1 in the normal cost rate minus the percent change from year t-1 to year t-2. This detrended percent change, in combination with the percentage point change in the discount rate assumption, yields an implied effective duration. The average of the four estimated effective durations is 13.5, far below the 20 to 25 rule of thumb effective duration, but quite close to the effective duration of 12 reported in Reinsdorf, Lenze, and Rassier (2014).

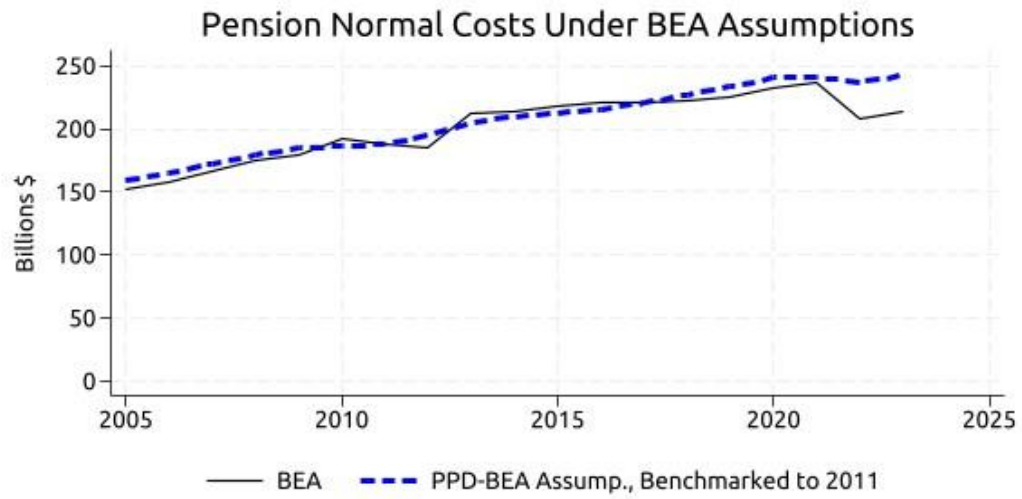
insensitive to changes in market interest rates. In particular, when market real rates are below the BEA's fixed real-rate assumption, normal costs will be understated.

Second, the BEA assumes that plans have fixed COLAs that are independent of inflation. This assumption explains why inflation-driven changes in the BEA's nominal discount rate affect normal costs at all. As discussed above, if COLAs moved one-for-one with inflation, then a change in expected inflation would leave normal costs unchanged: The lower discount rate would be offset by slower wage growth and lower inflation-indexed benefit payments. With fixed COLAs, however, changes in expected inflation do affect normal costs, though much less than changes in real discount rates. Consistent with this logic, Reinsdorf, Lenze, and Rassier (2014) report that the BEA pension model implies an effective duration of 12 with respect to changes in the nominal discount rate—that is, a 1 percentage point change in the nominal rate changes normal costs by about 12%. This duration is close to the 13.5 implied by our back-of-the-envelope calculation using changes in the BEA's assumed nominal discount rate, and to the duration we obtain in our actuarial model when, like the BEA, we assume that COLAs are invariant to inflation. Thus, the relatively low sensitivity of BEA normal costs to changes in market rates appears to reflect the combination of a fixed real-rate assumption and fixed COLAs. In contrast, our actuarial pension model produces a normal-cost effective duration firmly in the range of the standard rule of thumb.

B.4 Validating the Adjustment Method

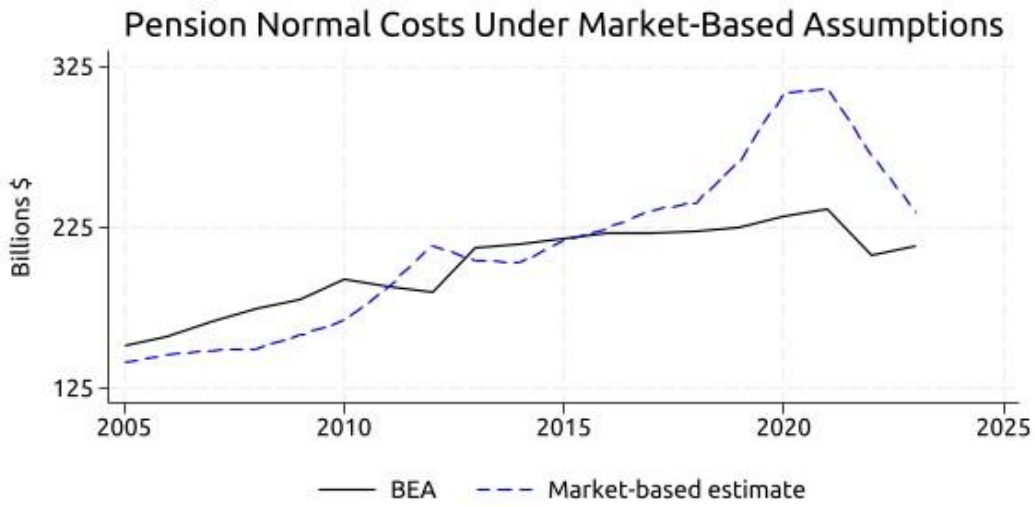
In order to assess the accuracy of the method we have developed for adjusting the plan's normal costs to different interest rate and inflation assumptions, we use our method to adjust the plans' normal costs to be consistent with BEA's assumptions. Figure A3a displays the results of this exercise; the solid black line displays the published BEA normal costs and the dashed blue line displays PPD normal costs harmonized to the BEA interest rate and inflation assumptions using our actuarial model-based approach. The harmonized PPD estimate matches the BEA almost exactly over most of the sample. They diverge somewhat, though, in the final two years. The reasons for this divergence are unclear and, given the complexity of the BEA's pension methodology, would likely be difficult to pin down. Overall, though, we view the exercise as providing assurance that our method for adjusting the PPD normal costs for differing assumptions over interest rates and inflation rates is sound. And the divergence at the end of the sample causes our estimate of pension benefits to exceed the BEA's estimate. A tendency to overstate the growth of these benefits at the end of the sample can be viewed as inducing a conservative bias to our results—i.e., the overstatement pushes against our finding that relative state and local government compensation has been declining.

Figure A3a



Source: Bureau of Economic Analysis's (BEA) National Income and Product Accounts (NIPA). Public Plans Data. 2001-2023: Center for Retirement Research at Boston College, MissionSquare Research Institute, National Association of State Retirement Administrators, and the Government Finance Officers Association. Author's calculations.
Note. PPD-BEA assumption, benchmarked to 2011 line displays the PPD normal cost adjusted to the BEA's assumptions for inflation and real rates using our pension model approach and benchmarked to the BEA in 2011.

Figure A3b



Source: Bureau of Economic Analysis's (BEA) National Income and Product Accounts (NIPA). Public Plans Data. 2001-2023: Center for Retirement Research at Boston College, MissionSquare Research Institute, National Association of State Retirement Administrators, and the Government Finance Officers Association. Author's calculations.
Note. BEA line represents the published normal cost in the NIPA. The Market-based estimate line is our calculation of the aggregate normal cost based on PPD data adjusted to reflect market-based assumptions for inflation and interest rates, and benchmarked to the BEA normal cost in 2011.

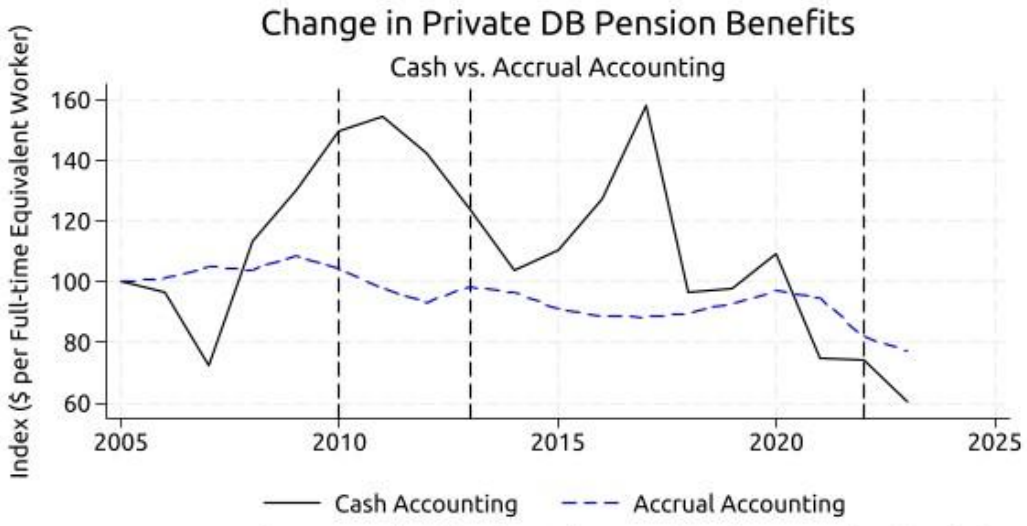
Figure A3b compares our preferred estimate of pension normal costs, which are adjusted to reflect market-based interest and inflation rates, to the BEA's normal cost estimates. Our estimated normal costs

(blue dashed line) are fairly close to the BEA published normal costs (solid black line) from 2005 through 2015. Thereafter, as market real interest rates fell significantly below the BEA's assumption, our market-based normal cost estimate moves substantially above the BEA's normal costs. With the sharp rise in real rates in starting in 2022, our normal costs move back into closer alignment with the BEA's. Thus, relative to using the BEA normal costs, our actuarial model-based estimates neither significantly affect our estimate of the overall decline in relative state and local compensation from 2011 to 2023 nor the level of relative compensation at the end of our sample period. They do, though, produce more accurate estimates from roughly 2015-2022.

B.5 Measuring Private DB Pensions

We measure private DB pensions using the accrual-based measure in the NIPA. While this measure may suffer from some of the deficiencies of the state and local government NIPA measure discussed above, by the start of the sample period in 2005, private DB pensions measured on either a cash or accrual basis had become a small share of private sector benefits. E.g., on an accrual basis as measured in the NIPA they comprised only around 8% of private sector benefits. Hence, they don't play a major role in the evolution of the accrual-based relative benefit series shown in Figure 5. Appendix Figure A4 displays the evolution of accrual-based and cash-based private sector DB pension benefits as measured in the NIPA.

Figure A4



Source: Bureau of Economic Analysis's (BEA) National Income and Product Accounts (NIPA). Author's calculations.
Note. Annual data averaged from quarterly data. Dashed vertical lines denote changes in the discount rate used by the BEA to discount future pension benefits; the discount rate from 2005 to 2009 is 5.5 percent, from 2010 to 2012 it is 5 percent, from 2013 to 2021 it is 4 percent, and from 2022 to present it is 5 percent.

B.6 State-level Public Pension Plan Normal Cost Estimates

Although the PPD pension plan micro data provides good coverage on a national basis, capturing 95% of liabilities, it has limited coverage in some states. Accordingly, we adopt a different approach for producing state-level pension normal cost estimates than the procedure used to produce national estimates and discussed in Appendix sections B.1-B.4. Instead, we adjust BEA state-level public pension normal cost estimates to our preferred, market-based inflation and real interest rate assumptions using standard actuarial rules of thumb. Specifically, we assume that the pension normal costs have an effective duration of 25 with respect to real interest rates and an effective duration of 10 with respect to inflation; e.g., for every 1 percentage point decrease in the real interest rate, the value of public pension normal costs are assumed to increase by 25 percent.

C. Retiree Health Care (OPEB) Estimates

C.1 National-level Estimates

Pew (2023) collects fiscal year (FY) 2019 data on a very large set of OPEB plans, including many local government plans. After adjusting their data for differences in the discount rate (as described below), their aggregate employer share of the OPEB normal cost represents 1.9% of BEA State and Local wages. We use this estimate to calibrate our OPEB estimates.

To estimate employer normal costs for OPEB in other years, we have collected data for 26 large plans—mostly state government plans, but some local plans as well—from 20 states.[33] We have data for all our plans for FY 2016 on. On an adjusted basis, the states covered by our plans represent almost 90% of aggregate OPEB cost. We have many fewer plans for earlier years. For example, for 2011-2013, we have data for only 6 large plans.

We adjust the normal costs in the plans for differences between our preferred real discount rate and that used by the plans, and subtract any required employee contributions (which only a minority of plans require) to calculate the employer normal cost. We then convert the normal costs from a fiscal year to a calendar year basis and aggregate them.

We use the changes in the employer normal costs for plans for which we have data to estimate national OPEB normal costs from calendar year 2011 to 2022, weighting each plan's growth rates by their share of the normal costs in our data. For the most recent years (2023 and 2024, for which we have little data), we assume that the plan provisions don't change, but that the normal costs do because of changes in discount rates.

The method we use to adjust the discount rate for OPEB normal costs is very similar to that used for pensions. In particular, we use a simple, streamlined plan to calculate the normal costs of retiree health

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[33] Our data cover the following plans, with the fiscal year start date of our data in parentheses: the New York City Teachers (2010) and New York State Employees plan (2017), the Massachusetts State Employees plan (2010), the Illinois State Employees plan (2010) and Illinois Teachers plan (2011), the New Jersey State plan (2010), the California State (CALPERS) plan (2010), the Pennsylvania State Employees plan (2014), the Texas Teachers (2014) and Texas State Employees plan (2016), the Ohio School Employees (2016), the Ohio Teachers (2016), and the Ohio State Employees plan (2016), and state plans from Delaware (2016), Hawaii (2016), Vermont (2016), Virginia (2016), North Carolina (2016), South Carolina (2016), and West Virginia (2016). Many of the state plans also cover some local employees.

insurance under different discount rate assumptions, and use the ratio of normal costs at our preferred discount rate in our simple plan to the normal costs at the plans' stated discount rate in our simple plan as a multiplicative adjustment factor. Our simple retiree health model differs in three respects from the one used to adjust pensions. First, consistent with data we have collected on OPEB plans, we assume that the cost of retiree health insurance falls by 70% once retirees turn 65 and become Medicare eligible. Second, we assume that health insurance costs rise 1 percentage point more per year than wages—reflecting both increases in health prices and increased health spending as retirees age. The first factor shortens OPEB duration relative to pensions while the second lengthens it. In our estimates, these factors are roughly offsetting, leaving the duration of the OPEB normal cost about the same as the pension normal cost duration. Finally, we assume that health insurance costs fully reflect changes in consumer price inflation; that is, the real increase in health costs over time is invariant to the inflation rate. Stated differently, we assume that when a plan has a higher inflation rate than our preferred rate, they also have higher nominal health care cost growth. This assumption is a sensible one and is also a feature of the model used by many retiree health insurance plans to estimate long-run medical cost inflation (e.g., the Society of Actuaries' Getzen Model of Long Run Medical Cost Trends). As explained for the case of pensions with COLAs fully indexed to inflation above, this assumption means that the inflation rate has no effect on normal costs, and so only real discount rates matter. (Because if wages and health costs rise with inflation, then the real present value of both benefits and wages is invariant to the inflation rate.)

Thus, our method for calculating the employer cost of OPEB benefits under our preferred real discount rate, $Normal\ Cost\ OPEB_{adj}$, is:

$$Normal\ Cost\ OPEB_{adj} = Normal\ Cost\ OPEB_{plan} * \frac{Model\ NC\ OPEB_{mkt\ real\ rate}}{Model\ NC\ OPEB_{plan\ real\ rate}}$$

where $Normal\ Cost\ OPEB_{plan}$ is the plan's normal cost under it's chosen discount rate, $Model\ NC\ OPEB_{mkt\ real\ rate}$ is the model-estimated normal cost under our market-based real discount rate, and $Model\ NC\ OPEB_{plan\ real\ rate}$ is the model-estimated normal cost under the plan's chosen real discount rate.

C.2 Estimates by State

The Pew data include estimates of the employer share of the normal cost by state, which we use for FY 2019. To create time series, we use a combination of methods: For states and years in which we have data on at least one plan, we use changes in the normal costs for those plans to estimate changes for the state as a whole. For years and states for which we have no plan data, we use the national estimates of changes in the normal costs. In collecting plan data, we collected data not only from states that represented a large share of aggregate employer normal costs, but also (where possible) from states that had high employer normal costs as a fraction of wages, so for most of the states where OPEB is a relatively large fraction of wages, we have data on at least one plan. In particular, we have data for 14 of the 18 states for which the average OPEB normal cost is 1% of wages or higher.[34]

. . .

[34] The exceptions are Alaska, Georgia, Louisiana, and Maine.

D. Aggregation of ECI-Based Series Using Alternative Data Sources

The ECI's defining feature is its use of a methodology to isolate changes in compensation costs unrelated to changes in workforce composition. The ECI achieves this fixed-composition measure in two steps. First, the ECI uses a fixed-weight approach where compensation changes are measured within detailed cells defined by industry, occupation, and sector (private and state and local government) then aggregated using fixed employment weights. Second, the ECI methodology uses a matched-sample approach under which the change in compensation from time *t-1* to *t* within a detailed cell is calculated only using "matched job quotes." Matched job quotes refer to the average compensation costs per hour within an establishment-occupation (i.e., a "job") which are available for both time *t* and *t-1*. The use of matched job quotes ensures that measured compensation changes do not reflect composition shifts within the detailed industry-occupation cells. E.g., it prevents the entry of a new establishment into the ECI sample at time *t* from influencing measured compensation growth at time *t*.

Using the terminology of Bureau of Labor Statistics (2022), the percent change in the ECI total benefits index for period *t* can be calculated as

$$\frac{O_t + DB_t}{\tilde{O}_{t-1} + \widetilde{DB}_{t-1}}$$

where DB_t is the "cost weight" for defined benefit pensions at time *t* and O_t is the cost weight for all benefits other than DB pensions. "Cost weights" are defined as the product of the average hourly wage (or hourly benefit cost) times the fixed, base-period employment count and can be thought of as being akin to an hourly wage bill. $\tilde{O}_{t-1}$ and $\widetilde{DB}_{t-1}$ are the prior period cost weights measured at the current time—the cost weights for period *t-1* calculated from job quotes which match those used to calculate $O_t + DB_t$. (Use of matched job quotes ensures that changes in the index do not reflect composition shifts within detailed industry-occupation cells.) Thus, $\tilde{O}_{t-1}$ need not equal O_{t-1}.

The ECI benefit index B_t is then:

$$B_t = B_{t-1} * \frac{O_t + DB_t}{\tilde{O}_{t-1} + \widetilde{DB}_{t-1}}$$

In order to operationalize combining subcomponents of the ECI index (e.g., O_t) with data from alternative sources (e.g., DB_t from our alternative series as shown in Figure 4), we assume that $DB_{t-1} = \widetilde{DB}_{t-1}$. That is, we assume that the difference between $DB_{t-1}$and DB_t does not reflect composition shifts within detailed industry-occupation cells. Furthermore, we must assume that the change in the alternative measure of DB_t over time does not reflect shifts in the composition of industry and occupation. Intuitively, and stated most simply, we must assume that shifts over time in DB_t do not reflect shifts in the composition of jobs.

Making these assumptions, expanding the equation recursively, and noting that the benefit index equals 1 in the base period (B_b=1), the level of the index at time *t* equals

$$B_t = \frac{O_t + DB_t}{O_{t-1} + DB_{t-1}} * \frac{O_{t-1} + DB_{t-1}}{O_{t-2} + DB_{t-2}} * \ldots * \frac{O_{b+2} + DB_{b+2}}{O_{b+1} + DB_{b+1}} * \frac{O_{b+1} + DB_{b+1}}{O_b + DB_b} * B_b = \frac{O_t + DB_t}{O_b + DB_b}$$

where b denotes the base period for measurement, $b<t$. Then we can restate the equation as

$$B_t = \frac{O_t+DB_t}{O_b+DB_b} = \frac{\frac{O_t}{O_b}*O_b+\frac{DB_t}{DB_b}*DB_b}{O_b+DB_b}$$

We obtain $\frac{O_t}{O_b}$ – the other benefits ECI index at time t – directly from the unpublished ECI data. We lack a published measure for $\frac{DB_t}{DB_b}$ – an ECI index for alternative, accrual-based DB pension. To address this, we first approximate DB_t—the defined benefit cost weight—by dividing the aggregate dollar value of our alternative accrual pension measure in year t by the NIPA measure of full-time equivalent employees. (We use the same procedure for retiree health care benefits and the value of job stability.) Second, we calculate $\frac{DB_t}{DB_b}$ as the ratio of the cost weight at time t relative to time b. The growth index $\frac{DB_t}{DB_b}$ provides an approximation of the growth in pension benefits per hour worked.[35] The other benefit cost weight, O_t, is approximated as the dollar value of other benefits per FTE as measured in the NIPA.

The assumption that the change in DB_t over time does not reflect composition shifts across industry-occupation cells or composition shifts within these cells is very unlikely to strictly hold. That said, we nonetheless view the assumption as reasonable for two reasons. First, as discussed above in Appendix section B.1, we calculate the percent change in DB_t over time as the sum of the percent change in state and local government wages as measured in the ECI and the percent change in our actuarial model-based estimate of the normal cost rate—the normal cost expressed as a fraction of annual wages. The ECI-based wage change holds composition fixed. This approach therefore only requires the assumption that shifts over time in the normal cost rate—pension costs as a share of wages—do not reflect shifts in the composition of state and local workers. Normal cost rates are less variable across employees than are normal costs, which vary with differing wages. Thus, our approach is less restrictive than simply assuming that changes over time in overall normal costs are not a function of worker composition shifts.

Second, the available data suggests that compositional shifts are not responsible for the evolution of state and local government compensation over our sample period. As shown in Figures A5a and A5b, there is little difference between the growth of wages and total compensation excluding defined benefit pensions in the ECI, which is based on the National Compensation Survey (NCS) and holds composition fixed, and the ECEC, which is based on the same survey but does not hold composition fixed. This

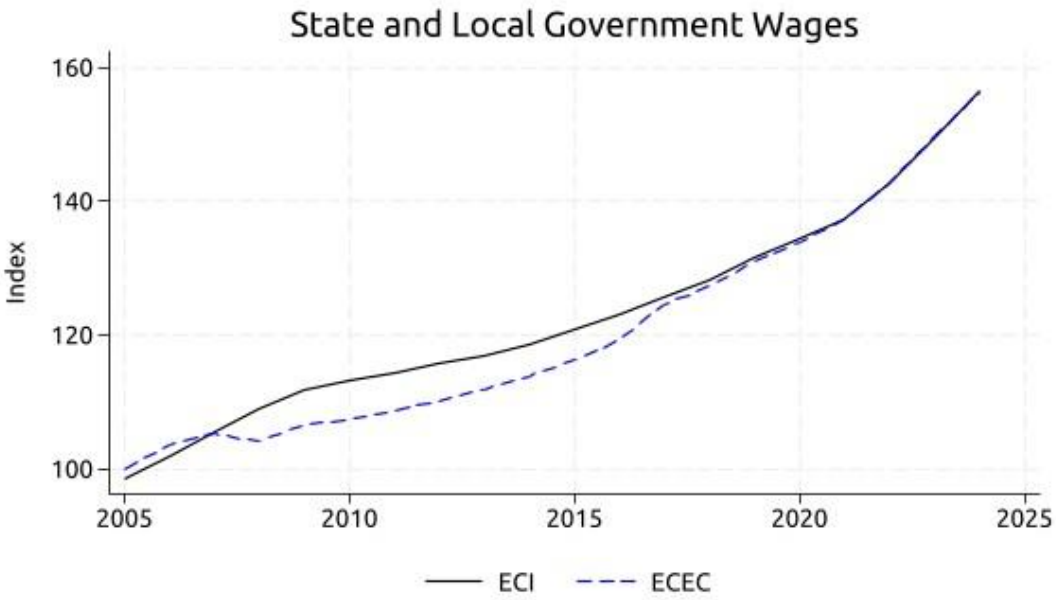
. . .

[35] The ECI measures the growth in employer costs per hour worked. The BEA defines full-time equivalent (FTE) workers as: "Full-time equivalent employees equal the number of employees on full-time schedules plus the number of employees on part-time schedules converted to a full-time basis. The number of full-time equivalent employees in each industry is the product of the total number of employees and the ratio of average weekly hours per employee for all employees to average weekly hours per employee on full-time schedules. An industry's full-time equivalent employment will be less than the number of its employees on full- and part-time schedules, unless it has no part-time employees." Given that the FTE concept is roughly equal to the number of hours worked scaled by the number of hours in a full-time schedule, the indexed version of benefits per hour worked approximately equals the indexed version of benefits per FTE. Put differently, the number of hours in a full-time schedule is assumed constant over time and falls out of the ratio $\frac{DB_t}{DB_b}$.

suggests that there have not been significant changes in job composition in the state and local sector or that such shifts did not induce changes in composition. Moreover, as shown in Appendix Figure A5c, the growth in state and local government defined benefit pension compensation measured on a cash basis in the ECI, ECEC, and NIPA broadly track together, even though neither the ECEC nor the NIPA controls for compositional shifts.[36] As of 2023, the ECI index is 10% lower than the comparable NIPA index, implying a relatively small difference in annual growth rates given the nearly 20-year time horizon and very rapid annual growth in all three series.

Overall, the evidence presented here, while not conclusive, suggests that the ECI-based compensation measures we have adjusted to reflect accrual accounting for defined benefit pensions are unlikely to suffer severe bias due to the inability to fully control for shifts across and within industry-occupation cells.

Figure A5a

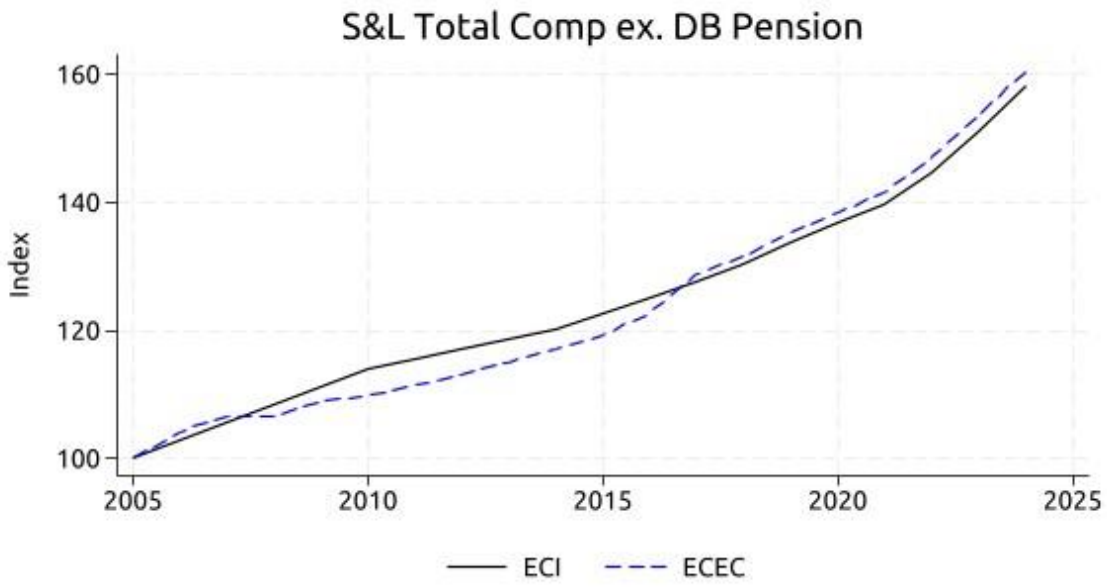


Source: Bureau of Labor Statistics' (BLS) Employment Cost Index (ECI) and Employee Costs for Employee Compensation (ECEC).
Note. Annual data averaged from quarterly data.

. . .

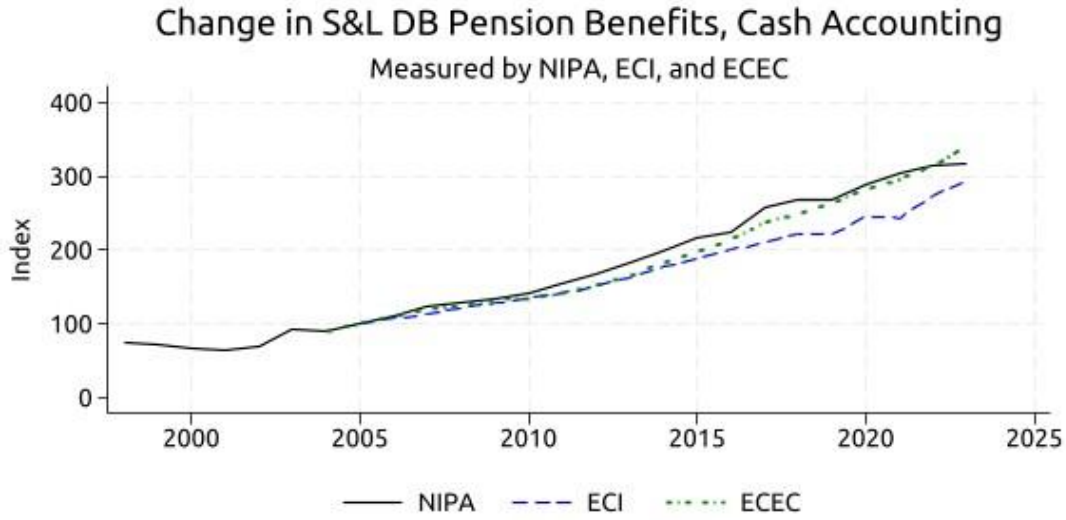
[36] The relevant NIPA measures are not included on Figures A1 and A2 because the NIPA wage measure is broader than the NCS measure—see Nicholson and Powers (2015)—and thus not comparable to the NCS measure.

Figure A5b



Source: Bureau of Labor Statistics' (BLS) Employment Cost Index (ECI) and Employee Costs for Employee Compensation (ECEC).
Note. Annual data averaged from quarterly data. Utilizes unpublished data provided by the BLS; these data are interpolated from 2005 to 2010 and the series should be interpreted with care because they do not meet the BLS's standard publication criteria.

Figure A5c



Source: Bureau of Labor Statistics' (BLS) Employment Cost Index (ECI, base period December 2005) and Employer Costs for Employee Compensation (ECEC). Bureau of Economic Analysis's (BEA) National Income and Product Accounts (NIPA). Author's calculations.
Note. Annual data averaged from quarterly data. ECI series is unpublished data provided by the BLS; these data are interpolated from 2005 to 2010 and the series should be interpreted with care because they do not meet the BLS's standard publication criteria. BEA series is indexed version of defined benefit pension benefits measured in dollars per full-time equivalent worker.

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