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ABSTRACT

We use a sample of full-time workers over 50 years of age from the 2004 and 2006 waves of the Health and Retirement Study to investigate whether workers in federal, state, and local government receive more generous wage and pension compensation than private sector workers, ceteris paribus. With respect to hourly remuneration (wages plus employer contributions to defined contribution plans), federal workers earn a premium of about 28 log points, taking differences in employee characteristics into account. However, there are no statistically discernible differences between state and local workers and their private sector counterparts, ceteris paribus. These findings are about the same whether or not indicators of occupation are included in the model. On the other hand, pension wealth accumulation is greater for employees in all three government sectors than for private sector workers, even after taking worker characteristics into account. As a proportion of the hourly private-sector wage, the hourly equivalent public-private differentials are about 17.2 percent, 13.4 percent, and 12.6 percent for federal, state, and local workers, respectively. We find no evidence that highly-educated individuals are penalized by taking jobs in the public sector, either with respect to wages or pension wealth.

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1. Introduction

For the public sector to compete effectively for workers, compensation needs to be comparable to that in the private sector. For many years, both academics and public policy makers have debated whether public sector workers are over- or underpaid relative to their private sector counterparts. This issue has recently received particular attention because of the severe budgetary constraints facing governments at all levels. For example, former Governor Mitch Daniels of Indiana argued, “We have a new privileged class in America…We used to think of government workers as underpaid public servants. Now they are better paid than the people who pay their salaries…Who serves whom here? Is the public sector — as some of us have always thought — there to serve the rest of society? Or is it the other way around?” [Garofalo, 2010]. At the same time, some public employees maintain that they are underpaid. A seventh-grade English teacher demonstrating against Wisconsin Governor Scott Walker’s plan to reduce public sector benefits told the Huffington Post, “I can't get a home loan. I set my thermostat at 62. No cable at my house, no internet…I'm also $36,000 in debt from becoming a teacher” [Delaney, 2011].

New Jersey Governor Chris Christie [2010] captured the essence of the issue nicely when he said, “at some point, there has to be parity between what is happening in the real world and what is happening in the public-sector world." But determining what “parity” means in this context is challenging for two reasons. First, the human capital of public and private sector workers may differ. If, for example, public sector workers have more education than private sector workers, then it is neither surprising nor objectionable that they earn higher wages. This is precisely the argument made by former White House budget director Peter Orszag regarding federal government compensation: "Basically the entire delta between private sector and public sector federal government average pay can be explained by education and experience…while there may be some remaining disparities, I think some of the more dramatic newspaper stories I've seen about that disparity are somewhat misleading" [Tuutti, 2010]. The second reason why public-private sector comparisons are problematic is that compensation consists of more than just wages and salaries. Pension benefits comprise an important part of compensation, so that comparisons of just wages and salaries may be misleading. Indeed, sectoral differences in pension benefits have recently drawn considerable attention. According to Cook [2011], "union chiefs can downplay their pension benefits all they want. The fact is, most of their members have been guaranteed a millionaire's retirement." Public-sector unions argue that such criticism is hyperbolic and unjustified. In the words of Hetty Rosenstein, New Jersey Director of the Communications Workers of America, "there's pension envy because people who are working in the private sector, they're being denied pensions" [Mulvihill, 2011].

In any case, it is clear that a sensible approach to measuring an overall compensation differential requires considering both wages and pension benefits. Further, the analysis of pension differentials needs to take into account that, because pensions are a form of compensation, their magnitude is determined in part by employee characteristics, just as is the case for wages and salaries. Simply examining average pension benefits across sectors is not an appropriate way to estimate sectoral differentials.

As noted below, a great deal of research on public-private sector wage differentials has been done, including careful analysis of micro-level data on individual employees that takes into account differences in human capital endowments. In contrast, there has been little work using micro-level data on differences in pension wealth across sectors. Furthermore, the few attempts to integrate wages and pension benefits into more comprehensive measures of sectoral compensation differentials use different data sets from different time periods to estimate the wage and pension components of the differential. In this paper, we use a single data set and a unified econometric approach to estimate wage and pension wealth differentials between public and private sector workers. Compensation packages, of course, have other components, including employment security, paid vacation, health insurance benefits, and so on.[1] Our analysis focuses on wages and pension benefits because they are of major importance and micro data are relatively accessible.

Section 2 surveys the existing research on public-private sector compensation differentials. Section 3 discusses our dataset, the Health and Retirement Study (HRS), and presents summary statistics for the key variables. A great advantage of the HRS is that it allows us to study wages and pensions within a unified statistical framework. A drawback is that it surveys individuals who are 50 years and older. Such workers are unlikely to be representative of the entire workforce because, for example, the public-private pension differential for older workers may differ from that of their younger counterparts because of changes in public sector pensions that have grandfathered the former group.

[1] In 2006, the year that we study, 70.1 percent of compensation came in the form of wages and salaries. Another 9.5 percent consisted of monetary compensation such as paid vacations and overtime pay. Contributions to defined benefit and defined contribution retirement saving plans were 4.3 percent. For further details on the composition of compensation see http://www.bls.gov/news.release/archives/ecec_06212006.pdf

That said, older workers are not an inconsequential component of the labor force--our tabulations from the 2006 Current Population Survey (CPS) indicate that the proportion of the workforce over 50 years exceeds 30 percent.

Section 4 presents the econometric setup and basic results. We find that the hourly wages of federal government workers are about 28 log points more than private sector workers with similar characteristics, *ceteris paribus*. However, there are generally no statistically discernible differences between state and local workers and their private sector counterparts with similar characteristics. On the other hand, pension wealth accumulation is generally greater for employees in all three government sectors than for private sector workers, even after taking worker characteristics into account. As a proportion of the hourly private-sector wage, the hourly equivalent public-private differentials are 17.2 percent, 13.4 percent, and 12.6 percent for federal, state, and local workers, respectively.

Section 5 presents several alternative econometric specifications to assess the robustness of the results. We find that controlling for occupation leaves our substantive conclusions unaffected. Further, there is no evidence that highly educated workers are disadvantaged by working in the public sector. Section 6 concludes with a discussion of the implications of our results for the ongoing public debate about public-private sector compensation differentials and suggestions for future research.

2. Previous Literature

The relevant literature can be classified into three broad categories: public-private sector wage differentials, pension benefit differentials, and total compensation differentials, which include both wages and pension benefits. We now discuss each of these topics in turn.

2.1 Wage Differentials

The literature on wage differentials between public and private sector employees spans roughly four decades, originating with Smith's [1976a, 1976b, 1977] seminal series of papers. The core of her analysis is the estimation of conventional human capital earnings functions. For example, in Smith [1976b] she uses 1973 CPS data to estimate for each gender a regression of the logarithm of the wage on various worker characteristics such as years of schooling and race, including a series of dichotomous variables indicating whether each individual worked in the federal, state, or local government sectors (the private sector is the omitted category). For males, she finds wage differentials relative to the private sector of 19 percent in federal government and –4.9 percent in local government. The coefficient on the state government variable is statistically insignificant. The differentials for female workers are 31 percent in federal government, 12 percent in state government, and 3.6 percent in local government. In a variation on this approach, in Smith [1977] she estimates separate equations for each sector, in effect relaxing the constraint that coefficients on the right hand side variables are the same across sectors. This modification does not change the qualitative results.

Papers subsequent to Smith's have modified her approach by trying to correct for self-selection of workers into various sectors,[2] by using panel data to estimate fixed effects models,[3] and by estimating models on a state-by-state basis to allow for the possibility that labor market institutions, and hence public-private wage differentials,

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[2] See Quinn [1979], Gyourko and Tracy [1986], Venti [1987], and Moore and Raisin [1991].
[3] See Krueger [1988].

vary across states.[4] A fair way to summarize the findings in this literature is as follows: a robust result, found in almost all the research from Smith's early papers on, is that there is a substantial positive wage differential for federal employees, even after controlling for worker characteristics in the standard way.[5] On the other hand, there is less agreement with respect to state and local government workers. Positive, negative, and zero differentials have all been estimated, sometimes within the same paper.

2.2 Pension Differentials

Data are available on average pension characteristics by sector, including the proportion of plans that are defined benefit (DB), that is, plans in which future benefits are based on a formula involving years of service, age, and so on, and defined contribution (DC), that is, plans in which future benefits are based on a cash balance in a pension account. (See, for example, Clark, Craig, and Ahmed [2010].) However, sophisticated econometric analyses using micro data like those described above for sectoral wage differences are rare. A notable exception is Quinn's [1982] important study of a sample of older workers from the 1969 Retirement History Study. He calculates pension wealth by discounting a stream of expected pension benefit cash flows from age of eligibility through age 100, using standard mortality tables to adjust for the probability of receiving a benefit in any given year.[6] Quinn estimates a regression of his pension wealth measure on dichotomous variables designating federal, state, local, and postal workers, and includes the number of years of job tenure and the final wage rate as control

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[4] See Belman and Heywood [1995], Schmitt [2010], and Lewis and Galloway [2011].
[5] Hence, the opposite characterization of the literature by Peter Orszag (see Section 1 above) and Paul Krugman [2010] seems rather curious.
[6] Quinn's calculation eliminates the average sectoral portions of pension wealth that are associated with workers' contributions.

variables. He estimates pension wealth differentials of 72 percent for federal workers, 80 percent for state workers, and 30 percent for local workers.

A limitation of Quinn's analysis is that he does not control for differences in worker characteristics such as schooling, gender, and race. In addition, his computation of pension wealth excludes DC plans. While such a decision was sensible at the time Quinn was writing, this is no longer the case, as DC plans have become an important component of pension wealth, particularly in the private sector. According to Munnell and Perun [2006], the share of private DC plan assets out of all private pension plan assets increased from 22 percent in 1980 to 50 percent in 2004, and the corresponding share of active participants increased from 35 percent to 70 percent over the same period. As noted below, our analysis takes advantage of a more suitable estimate of pension wealth that encompasses both DB and DC plans.

2.3 Aggregating Wage and Pension Differentials

Some recent research has sought to integrate wages with fringe benefits in order to obtain a broad measure of sectoral compensation differentials. In an important paper using data from the Health and Retirement Study for 1992-2000, Ramoni-Perazzi and Bellante [2007] combine federal, state, and local employees to form a single "public" category, and use the standard approach to estimate a public-private sector wage differential that ranges from 3.5 to 11 percent. They next compute the difference in the average sectoral fringe benefit share of total compensation from Bureau of Labor Statistics data (which are not corrected for differences in the demographic composition of workers across sectors). Adding this figure to their estimated wage differential, Ramoni-

Perazzi and Bellante find a total public-private sector compensation difference that ranges from 6 to 14 percent.[7]

Keefe's [2010] similar approach begins by estimating wage differentials in the conventional fashion (using CPS data for 2009). To incorporate fringe benefits, Keefe marks these wage differentials up by average benefits based on the employee's occupation and establishment size, as reported in the Employer Costs for Employee Compensation Survey. He finds fringe-inclusive differentials of –10.7 percent for state workers and –4.1 percent for local workers. The adjustment for fringe benefits does not take into account employees' personal characteristics. Other studies following the same general approach are Allegretto and Keefe [2010], Bender and Heywood [2010], Thompson and Schmitt [2010], Cannon [2010] and Gittleman and Pierce [2012], all of whom study sub-federal levels of government, and the Congressional Budget Office [2012], which examines federal government compensation. To our knowledge, no analysis of public versus private sector compensation patterns has estimated both wage and pension wealth differentials using a unified multivariate econometric framework based on micro-level data. Our data allow us to apply the same econometric approach for analyzing pension differentials as has been used to study wage differentials.

3. Data

Our analysis sample comes from the 2006 wave of the Health and Retirement Study (HRS), a longitudinal study of Americans aged 50 and over, who are interviewed

[7] This calculation includes fringe benefits other than pensions, such as sick leave, paid vacation, and the value of differences in unemployment probabilities.

every two years by the Institute for Social Research at the University of Michigan.[8] As stressed above, there are reasons to believe that public-private compensation differentials for these older individuals are not typical of the workforce as a whole. But as we also noted, they are an important part of the labor force; more than 30 percent of all workers are over 50 years old. According to our tabulations using the 2006 CPS, the proportions of employees who were 50 years or older in the private, federal, state, and local sectors in 2006 were 29.3 percent, 39.5 percent, 34.8 percent, and 36.5 percent, respectively. Because the HRS is primarily based on older workers, our results might not apply to employees throughout the age distribution.[9] However, in addition to standard demographic data, the HRS has information on sector of employment[10] as well as a rich set of variables on pension benefits, which makes it invaluable for making compensation comparisons that go beyond wage and salaries.

All variables relating to worker characteristics, hourly pay, pension plan types, and employer pension contributions are either included or directly calculated from the original 2006 Core dataset and tracker file. Data on pension wealth are obtained from a

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[8] We use the unrestricted HRS sample, which, unlike the restricted sample, does not include pension plan descriptions provided by employers. As noted by Gustman, Steinmeier and Tabatabai [2010, Chapter 4], there are advantages and disadvantages to each dataset, and for a study like ours, it is not obvious whether one would be better than the other. For example, individuals' self-reports in the unrestricted data are subject to error, especially when it comes to identifying plan type. But in the restricted data, other pension attributes are erroneously reported by employers. Further, the proportion of HRS respondents for whom it is possible to obtain employee information is relatively low in the restricted sample. In the 2005 employer survey, for example, employer-provided plan descriptions were available for only 49 percent of the respondents who reported having a pension (p. 64).

[9] In addition, public employees may leave public sector employment in their 50s and then switch to private sector jobs. However, in our sample, we find little evidence for this phenomenon. Out of 392 government workers with non-zero sample weights whom we are able to track over time, only 6 individuals made such transitions between 2006 and 2008. Relatedly, the propensity to leave the workforce altogether because of health reasons or retirement might differ systematically across sectors, but our data do not allow us to investigate this phenomenon.

[10] Starting in 2006, the HRS clearly identifies government sector affiliation through questions KJ720 ("Are you employed by the government at the federal, state, or local level?") and KJ721 ("Would that be the federal, state, or local government?").

supplement constructed by Gustman, Steinmeier, and Tabatabai [2010]. They note that there is some evidence of errors in reports of pension values, and the accuracy of the reports does not improve over time (p. 213). However, Gustman, Steinmeier and Tabatabai argue that the estimates do not appear to be systematically biased. If the reporting errors differed systematically by sector of employment then this would bias our estimates of pension differentials. There is no evidence with respect to whether pension wealth estimates for public sector employees are more or less accurate than their private sector counterparts.

A total of 4,759 respondents provide salary and wage information. Respondents are also asked about the number of hours worked in a typical week and the number of weeks worked per year, allowing us to calculate total hours worked per year and hourly pay. This information is available for 4,344 respondents. We drop from the sample the self-employed as well as part-time workers, defined as individuals who worked fewer than 1,500 hours per year.[11] This leaves us with 3,199 observations. Furthermore, we exclude respondents who, while earning a positive hourly wage, earned less than the minimum wage or characterized themselves as retired[12]. Finally, we drop workers for whom information is missing for any of our right hand side variables, leaving us with 2,496 observations, which comprise our basic analysis sample. There are 1,907 respondents in the private sector, 115 in federal government, 225 in state government, and 249 in local government. Compared to the entire sample of 4,759 individuals who

[11] There is, of course, some arbitrariness to the selection of a cutoff for what constitutes full-time work. We experimented with a lower cutoff, 1,250 hours per year, and found that our substantive results were not affected.

[12] Question KJ005 on the 2006 HRS questionnaire reads "Are you working now, temporarily laid off, unemployed and looking for work, disabled and unable to work, retired, a homemaker, or what?" Respondents are allowed to give multiple responses to this question. We excluded all workers who indicated "retired" as one of their responses. There were a total of 333 workers who did so (including part-time workers). Retired individuals were not asked their sector of employment during their working years.

reported wage and salary information, this sample is slightly younger (mean age of 58.2 years versus 59.5) and has fewer black workers (9.0 percent versus 9.7 percent) and female workers (47.8 percent versus 51.4 percent).

Table 1 presents definitions and summary statistics for our dependent variables. *WAGE* is hourly pay in 2006 dollars. Federal and state government employees earn substantially more than either private sector employees or employees in local government. Three dichotomous variables indicate whether a given respondent has a primary (most important) pension plan of a specific type. *DBPlan* indicates a primary defined benefit plan, *DCPlan* indicates a primary defined contribution plan, and *HPlan* indicates a hybrid plan, which has attributes of both DB and DC plans. Respondents who either refused to or did not know how to answer the question about their primary pension plan were excluded from the part of our analysis that involves these variables (37 respondents did not know their primary plan type or refused to identify it). Local government workers are the most likely to have a DB plan and private sector workers are the most likely to have a DC plan.

EmpC is the amount contributed by the employer to all of the worker's DC plans per hour (in 2006 dollars), conditional on it being positive. It includes both the basic contribution and any matching amounts. Respondents can provide information on employer pension contributions for up to four DC plans as either a percentage of pay (in which case *EmpC* is calculated as the reported fraction multiplied by *WAGE*) or as an amount per some unit of time (in which case the conversion to an hourly basis proceeds along the same lines as for *WAGE*). A number of respondents either did not know their employer's contribution or refused to answer the question.[13] In total, we were able to calculate positive hourly employer contributions for 806 workers out of the 1,262 worker**s** who reported having at least one DC plan.

WAGE+DC is defined as the sum of *WAGE* and *EmpC*, that is, wages and salaries plus employers' contributions to DC plans. There is not much difference between the average values of *WAGE+DC* and the figures for *WAGE* in the first row. Differences in DC coverage are not large enough to exert substantial effects on wage and salary differences.

PensionW is pension wealth as calculated by Gustman, Steinmeier, and Tabatabai (GST) [2010]. It is defined as the sum of the present value of future pension benefits based on work to date from a respondent's most important DB plan[14] and the cash balances in all DC plans on the current job in 2004 (in 2006 dollars). More precisely, GST take a worker's expected annual DB payments while in retirement and discount them back to the expected time of retirement.[15] They then further discount these figures back to 2004. The specific discount rate, of course, plays a central role in these

[13] A respondent with several DC plans might only know his employer's contribution to one of them. In this case, *EmpC* would only capture a fraction of the total employer contribution. Given that DC plans are relatively more important in the private sector, the lack of information on employer contributions to DC plans might have different impacts in the public and private sectors. However, we have no way to quantify the possible importance of this phenomenon.

[14] The proportion of workers who have more than one DB plan is relatively small: 2.2 percent in the private sector, 5.5 percent in the federal government, 3.0 percent in the state government, and 4.0 percent in the local government. Gustman, Steinmeier and Tabatabai [2010] do not include the present value of a worker's second DB plan in their calculations because there is a higher probability of confusion among workers with two DB plans who are asked to report plan details (p.337).

[15] Some public employees are not covered by Social Security. They neither contribute to Social Security on their jobs nor collect Social Security benefits. Therefore, their wealth upon entering retirement is lower than their private sector counterparts, but so are the payroll tax contributions they made during their work lives. However, according to Gustman, Steinmeier and Tabatabai [2013, Table 7], the present values of Social Security benefits and taxes experienced by our HRS cohort are virtually identical, so there should be little to no effect on our findings.

calculations. GST [2010] use a discount rate of 5.8 percent, consistent with the figure used by the Social Security Administration in its long-term projections (p.338). [16]

Since the value of a DC plan in 2004 is a cash balance that has been accumulated based on work to date, the value of a DB plan in 2004 also needs to reflect the fraction of its total value that has accrued to it based on work to date. In order to make this adjustment, GST [2010] assume that a given worker will continue to work for his employer up to retirement and multiply the total present value of the DB plan by the number of years of service for the current employer in 2004, divided by the total (anticipated) number of years of service at the time of retirement. Adding the DC cash balances to this adjusted DB present value yields a measure of total pension wealth.

We use the 2004 value because 2004 is the most recent year that GST calculate it for all workers.[17] *PensionW* is available for respondents in our analysis sample who were interviewed in both 2004 and 2006, and worked for the same employer in both years--2,100 of the 2,496 respondents in our analysis sample.

The variable *PositivePW* is one if the individual has positive pension wealth and zero otherwise. According to the figures in Table 1, employees in all levels of government are more likely to have positive pension wealth than private sector workers, and conditional on having pension wealth, their average values are also higher. Both the probability of having positive pension wealth and the conditional mean are greatest for federal government workers.

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[16] Further, they make a mortality adjustment to account for variability in the probability of receiving a benefit. However, there is no correction for the possibility of financial insolvency. Given that both public and private sector entities might renege on part or all of their obligations, it is not clear how this would bias estimates of public-private differentials.

[17] For some observations, GST [2010] impute values using mixed, hot-decking, and replacement methods. See "Imputations for Pension Wealth, Final Version 2.0, December 2006" at https://ssl.isr.umich.edu/hrs/files.php?versid=64 for more detail.

The sectoral differences among the various compensation measures are striking. However, one must be cautious in interpreting them, given that the characteristics of workers might "explain" these differences. Table 2 shows weighted means and standard deviations of the key demographic variables for workers in the private sector and various levels of government. The figures indicate that, indeed, worker characteristics differ substantially across sectors. For example, on average, government employees at all levels have more education than private sector employees, which gives some credence to the notion that the compensation differentials in Table 1 might be due to differences in human capital.

Public and private sector workers also differ with respect to gender, race, marital status, age, and whether they live in highly populated areas.[18] Women account for a large share of the state government workforce (61 percent), but smaller shares of the federal (49 percent), local (47 percent), and private sector workforces (46 percent), respectively. Federal and state government have the highest proportions of black workers (15 and 14 percent, respectively), followed by local government (9.6 percent), and the private sector (8.0 percent). With regards to marital status, state government has the largest proportion of married workers (73 percent), while the proportions in federal government (67 percent), local government (68 percent) and the private sector (67 percent) are nearly identical. Workers' ages are essentially the same across sectors, although one must keep in mind that our data set includes only individuals who are 50 years of age or older.[19] In

[18] The indicators for size of metropolitan area are meant to provide a rough control for differences in the price level across areas. We are grateful to a referee for this suggestion.

[19] In all sectors, most workers are between 50 and 59 years of age. The proportions of workers who are between the ages of 50 and 59 in the private, federal, state, and local sectors are 67.7 percent, 63.2 percent, 67.7 percent, and 72.6 percent, respectively. The proportions who are between the ages of 60 and 69 in the private, federal, state, and local sectors are 29.4 percent, 33.0 percent, 30.6 percent and 24.8, respectively.

this context, an important question is whether our analysis sample is plausibly nationally representative (conditional on the screens used to generate it). The issue is cogent because the 2006 survey does not have many of the respondents included when the HRS was started as a nationally representative data set of individuals who were 51 to 61 years of age in 1992.

This is an empirical issue, and a sensible way to approach it is to apply to a nationally representative data set the same screens that we use to generate our HRS analysis sample, and compare summary statistics of the key variables that are available in both samples. To that end, we took the 2006 CPS, applied the screens described above, computed the relevant (weighted) summary statistics, and compared them to the HRS. The results, reported in the appendix, indicate that the summary statistics for the two samples are quite close. Indeed, for none of the variables can one reject the hypothesis that the respective HRS and CPS means are equal.

4. Econometric Specification and Results

While informative, the figures in Table 1 are only suggestive because they fail to account for the fact that, as documented in Table 2, employee characteristics differ substantially across sectors. This section presents the results from multivariate models.[20]

4.1 Hourly Wages

Basic setup. To begin, we estimate regressions of the log of hourly pay (*LOG(WAGE)*) on worker characteristics and dichotomous variables for sector of

[20] The HRS sample includes features such as stratification, clustering, and differential selection probabilities. All regressions use the HRS respondent sample weights. The HRS contains design-specific sampling error codes in order to facilitate the computation of standard errors. Following the suggestion in the HRS documentation, we use the SVY command set in STATA to incorporate these sampling error codes into the calculation of the standard errors. For further details, see pages 5 and 6 of the HRS document *Sample Evolution: 1992-1998* (http://hrsonline.isr.umich.edu/sitedocs/surveydesign.pdf).

employment. This is sometimes called the "people approach" to analyzing public/private compensation differentials, and is based on the notion that differentials should be computed by comparing the wages of public and private sector workers who have the same characteristics (Gittleman and Pierce [2012, p. 225]).[21] Specifically, we regress the logarithm of the hourly wage on a set of indicators for sector of employment (with the private sector as the excluded category) as well as educational attainment, gender, race, marital status, and a quadratic in age.[22,23] In effect, this specification constrains all the coefficients, except those on the sectoral variables, to be the same across sectors. A more flexible approach, suggested by Blinder [1973] and Oaxaca [1973], allows all the coefficients to vary across sectors. We choose the more constrained model because it generates essentially the same results as the Blinder-Oaxaca approach and is simpler to exposit.

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[21] The alternative approach is to match individuals in the public and private sectors whose job descriptions are comparable. Our data do not permit us to take this approach, and in any case, as Gittleman and Pierce [2012, p. 221] point out, it is problematic because some occupations, such as sales, are almost all in the private sector, while other job categories are almost all public. An appendix table available upon request shows the occupational distribution by sector in our data.

[22] A subset (2,174 workers) of our observations provided responses to the question , "How long have you been with your current employer?" Specifically, we can calculate it for 1,668 workers in the private sector, 92 federal workers, 192 state workers, and 222 local workers. If the 2006 response was not available, we added two years to the 2004 response provided that the worker was still on the same job in 2006 as in 2004. The average tenure by sector is 12.6 in the private sector, 17.7 years in the federal government, 14.8 years in state government, and 17.4 years in local government. However, Altonji and Williams [2005] have argued persuasively that including tenure in an OLS model is inappropriate, because it is positively correlated with the error term in the likely event that individuals with low productivity have high quit and layoff propensities. Similarly, including tenure on the right-hand side in modeling the determinants of pension wealth could also lead to inconsistent estimates—where pensions are more generous, individuals are likely to stay on the job longer, *ceteris paribus*.

[23] In principle, one could take advantage of the panel nature of the data to estimate a fixed effects model. We do not take this tack for two reasons. First, we would lose 802 observations: only 1,694 workers in the 2006 basic sample were re-interviewed and had a positive hourly wage in 2008. Second, in a fixed effects model, the coefficients on the sectoral variables are identified off of worker transitions between employment sectors, of which there are very few between 2006 and 2008. Out of 1,302 private-sector workers in 2006, 2 move into federal government, 5 move into state government, and 5 move into local government in 2008. Out of 75 federal workers in 2006, 1 moves into the private sector, and 1 moves into state government in 2008. Out of 145 state workers in 2006, 2 move into the private sector in 2008. Out of 172 local workers in 2006, 3 move into the private sector. This paucity of sectoral transitions would result in imprecise estimates of the sectoral differentials.

Unlike some previous studies, we do not include occupational variables in our canonical model. As Gittleman and Pierce [2012, p. 227] point out, there is no consensus on this matter. We choose not to include occupation, because occupational choice could very well be jointly determined with wages and it is available only for a subset of our sample.[24] However, as shown in Section 5 below, when we include occupation on the right hand side, the results with respect to sectoral differences are nearly identical.

Results. The estimates are reported in Table 3. In the first column we show the results when only the sectoral variables are included. Substantial positive differentials are present for all government levels: federal government employees earn approximately 37.1 log points more than private sector employees, state government employees earn 16.3 log points more, and local government employees earn 8.6 log points more. All of this is just a different way of characterizing the information from Table 1.

The estimates in column (2) show the results when employee characteristics are taken into account. The federal-private differential falls to 0.278 (s.e. = 0.045), considerably smaller than the corresponding figure in column (1), but still significantly different from zero. In contrast, the state and local government differentials are generally rendered small and statistically insignificant when worker characteristics are included in the model. Taken together, the results in columns (1) and (2) are consistent with the claim that looking at public-private wage differences without taking into account worker characteristics can be misleading. Our finding of a large and statistically significant federal differential is generally in line with previous results, most notably by Smith

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[24] We cannot control for union coverage because the HRS does not include this information in the 2006 wave. In any case, as Gittleman and Pierce [2012, p. 226] argue, controlling for unionization status does not seem appropriate, "because union wage premia probably do not reflect ability differences, and those in the public workforce would not likely take their public sector unionization rates with them if they were to move to the private sector."

[1976a, 1976b, 1977], Venti [1987], Krueger [1988], and Moore and Raisian [1991]. However, our point estimate is somewhat larger than theirs. This raises the question of whether our finding with respect to the federal differential is idiosyncratic to the HRS dataset. To investigate this possibility, we applied to the CPS the same screens as we used to generate our analysis sample, and then replicated the specification in column (2) as closely as we could. The results, reported in the appendix, indicate that the estimated federal differential using the CPS is somewhat smaller than our estimate in Table 3, but one cannot reject the hypothesis that they are equal. Thus, our estimate of the federal hourly earnings differential does not appear to be an artifact of the HRS sample.

Turning to the other covariates listed toward the bottom of column (2), the signs and magnitudes of the coefficients are sensible and in line with the results of previous econometric work on wage determination. Hourly pay increases with educational attainment; for example, compared with workers who have no degrees, those with a four-year college degree earn approximately 64 log points more. Married workers enjoy an earnings premium of about 8 log points, and the wage rate decreases with age throughout the relevant range.[25] Workers in urban and suburban areas have higher hourly wages than those who live in less densely populated areas.

4.2 Hourly Wages and Employer Contributions to DC Plans

We now turn to differences in compensation that come in the form of pensions.[26] Integrating employer contributions to defined contribution plans with wages is fairly

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[25] The negative quadratic term on *Age* starts to dominate the positive term at 49 years of age.

[26] Recall from Table 1 that, on average, government workers at all levels of government are more likely to have DB plans and less likely to have DC plans than their private sector counterparts. These findings continue to hold in multivariate models that include the demographic variables in Table 3 (results are available upon request). The dramatic differences in the incidence of DB and DC pensions in the public and private sectors cannot be attributed to differences in worker characteristics.

straightforward. (We return later to issues associated with incorporating defined benefit plans.) We create a variable, *WAGE+DC*, which equals the sum of the (hourly) dollar amount of employer DC pension contributions (if any)[27] and hourly pay. The last two columns in Table 3 show the results when we use our standard earnings function to analyze this broader measure of compensation. Comparing column (1) to column (3) suggests that including employer contributions to DC plans has little effect on the raw sectoral coefficients**.** The results after taking demographic characteristics into account (columns (2) and (4)) are similarly very much the same. We conclude that incorporating employers' DC contributions into the analysis does not substantively change the results one obtains by looking at wages only.

4.3 Pension Wealth

So far, our focus has been on *flows* of compensation associated with wages and employer contributions to DC plans. In principle, one would also want to include the incremental accrued value of employer contributions to defined benefit plans during the relevant time period, which is also a flow. Such a variable is not included in the public version of the HRS data. However, a useful *stock* measure, total pension wealth, has been calculated by Gustman, Steinmeier, and Tabatabai [2010]. It is defined as the sum of the present value of future benefits from a worker's most important DB plan on the current job and the cash balances in all DC plans on the current job, based on work to date.

[27] State and local government employees are significantly less likely to receive employer contributions (state workers are 17.4 percentage points and local workers are 19.4 percentage points less likely than private sector employees), but there is no statistically discernible difference between federal employees and their private sector counterparts along this dimension. In results not reported here, we show that apart from the case of federal workers, differences in the likelihood and amount of an employer contribution to a worker's DC plan(s) cannot be attributed to differences in demographic characteristics.

This section examines how total pension wealth depends on sector of employment. As previously noted, just like wages, pension wealth likely depends on an employee's personal characteristics. Gittleman and Pierce [2012, p. 224], for example, conjecture that public sector workers might demand compensation packages skewed towards benefits because they have more education. Hence, a multivariate approach is needed to ascertain the independent effect of sector of employment upon pension wealth. We include the same set of right hand side variables as in the wage regressions. A substantial proportion of individuals in our sample have zero pension wealth (see Table 1). There is some controversy with respect to the best econometric strategy in this situation. Some argue that in the presence of a limited dependent variable, a nonlinear estimator such as probit or Tobit should be employed. However, we follow Angrist and Pischke [2009, p. 94], who note that ordinary least squares provides the best linear approximation to the conditional expectations function, and hence, in a context like ours where censoring is not present, is the appropriate estimator.[28]

Column (1) of Table 4 shows the estimates from a linear probability model of the impact of sector of employment on the probability of having positive pension wealth without any additional covariates. Compared to private sector employees, federal workers are 21.6 percentage points more likely to have positive pension wealth, state workers are 14.8 percentage points more likely, and local workers are 12.9 percentage points more likely. All the coefficients are statistically significant. The corresponding pension wealth differentials are in column (2). Workers in all government sectors have more pension wealth than their private sector counterparts: on average, federal workers have $111,000

[28] However, when we re-estimated the model using a Tobit estimator, the marginal effects were qualitatively quite similar to those from OLS.

more pension wealth, state workers have $89,800 more, and local workers have $80,800 more.

Columns (3) and (4) add in our other covariates. The most striking result is that the uncorrected pension wealth differentials from column (2) are reduced yet remain substantial and statistically significant. Specifically, column (4) shows federal, state and local pension differentials of $93,600, $62,900, and $65,800, respectively. Because wages are a flow variables and pension wealth is a stock variable, the results from Tables 3 and 4 are not directly comparable. To allow comparisons, we begin by computing the hourly annuity value of the pension wealth differential in each government sector. Suppose that the pension wealth differential in sector s is D_s. Suppose further that the average number of years of service for workers in that sector is T_s. Then we compute the annuity value of D_s over T_s years, assuming an interest rate of 5.8 percent, the rate used by Gustman, Steinmeier, and Tabatabai [2010] in their calculations.[29]

The mean years of service with the current employer for our 2004 pension wealth subsample[30] are: 13.1 for private sector workers, 17.1 for federal workers, 14.0 for state workers, and 16.4 for local workers. Using these values and the respective pension wealth differentials from Table 4, the annuitized differentials are $8,778, $6,685 and $6,323 for federal, state, and local workers, respectively. The mean yearly number of hours worked for our pension wealth sample in 2006 is 2,256 for private sector workers, 2,192 for federal workers, 2,148 for state workers, and 2,156 for local workers. Therefore, the

[29] We also did the calculations assuming a lower interest rate, 3 percent, and the results were substantively very similar.

[30] As was mentioned in section 3, some respondents in our sample changed jobs between 2004 and 2006, which makes it impossible to determine whether their pension wealth in 2004 can be associated with a federal, state, local, or private sector job. For this reason, the sample for this analysis includes only those individuals who worked in the same job in 2004 and 2006.

annuitized pension wealth differentials per hour come to $4.00, $3.11 and $2.93 for federal, state, and local workers, respectively. As a proportion of the hourly private-sector wage, these amount to 17.2 percent, 13.4 percent, and 12.6 percent for federal, state, and local workers, respectively.

4.4 Preliminary Conclusions

Summary. Taken together, this section's results suggest several preliminary conclusions. (1) Once differences in worker characteristics are taken into account, there are no statistically significant differences in hourly wages between state and local workers and their private sector counterparts. There is, however, a substantial wage differential for federal workers, *ceteris paribus*, of about 28 percent. (2) The results for hourly wages are essentially unchanged when employer contributions to employee DC plans are included in the measure of hourly compensation. (3) Employees of all levels of government generally have substantially more pension wealth than their private sector counterparts.

Toward a comprehensive measure of compensation differentials. One way to generate comprehensive compensation differentials would be to take the sum of the hourly equivalent pension wealth differentials computed above and the wage differentials from the second column of Table 3.[31] However, doing so could be misleading; seeing why brings us to the next complication in interpreting the pension wealth differentials. Suppose that a given individual's DB pension is funded entirely by her own contributions; the employer puts in nothing. In that case, pension wealth would not

[31] We add the annuitized pension wealth differentials to the differentials from the second rather than the fourth column of Table 3 because the latter embody employer DC contributions, which are already included in the pension wealth variable. In effect, adding the pension differentials to the differentials in log(*WAGE*+*DC*) would be double counting.

represent any additional compensation; it would just be a use to which the individual puts her wages. Hence, it would be inappropriate to add the pension wealth and wage differentials to obtain a total compensation differential.

More generally, as long as a portion of contributions to DB pension plans come from employees, then to some extent double-counting will occur if one simply adds the wage and pension differentials. This observation immediately leads to the question of how much pension wealth is due to employer and to employee contributions, respectively. Our data do not provide an answer. However, using information from several sources, a back-of-the-envelope estimate is possible. The steps in this calculation are as follows:

First, obtain an estimate of the average hourly amount of DC contributions made by employers and employees in each sector. As noted in Section 3 above, this information is available in the HRS data.

Second, obtain an estimate of the hourly amount of DB contributions made by employers and employees. The HRS data do not have this information. However, Bureau of Labor Statistics data allow us to make a rough imputation. Specifically, BLS data report the average hourly employer cost for state and local workers' DB plans ($2.27 in 2006)[32] and the average required employee contribution rate for state and local workers' DB plans (6.3 percent in 2007)[33]. Given that the average required employee DB contribution rate is computed only over the set of state and local employees with DB plans requiring an employee contribution, we multiply the 6.3 percent figure by the

[32] U.S. Bureau of Labor Statistics, *Employer Cost for Employee Compensation - March 2006*, June 2006, p.8.

[33] U.S. Bureau of Labor Statistics, *National Compensation Survey: Employee Benefits in State and Local Governments in the United States, September 2007*, March 2008, p.8.

percentage of DB plans that require a contribution (77 percent in 2007)[34], which yields an average employee contribution rate of 4.85 percent across all workers. For federal workers, the website of the Office of Personnel Management indicates that the contribution rate is 7.0, 7.5 or 8.0 percent for both employers and employees; we use 7.5 percent, the midpoint of this range[35]. By multiplying the employee's average required DB contribution rate with the mean hourly wages listed in Table 1 and the average number of DB plans per worker in each sector in our sample, we calculate hourly employee DB contributions of $1.60 for federal workers, $1.55 for state workers, and $0.89 for local workers.

Third, multiply our respective estimated hourly pension wealth differentials by the ratio of the sum of the employer DC and DB contributions to the sum of employer and employee DC and DB contributions. This yields a set of differentials due to employer contributions. An appendix available upon request provides details.

The last step is to add these figures to the wage differentials from column (2) of Table 3. Given that one cannot reject the hypothesis that the state and local differentials are zero, for this purpose we simply treat them as zero. This yields total compensation differentials of 34.2 percent for federal employees[36], 7.49 percent for state employees, and 8.29 percent for local employees.

5. Alternative Specifications

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[34] U.S. Bureau of Labor Statistics, *National Compensation Survey: Employee Benefits in State and Local Governments in the United States, September 2007*, March 2008, p.7.

[35] See http://www.opm.gov/retirement-services/csrs-information/. Employee contributions are mandatory.

[36] For purposes of this calculation, we use the method suggested by Blackburn [2008] to convert the coefficient on the *Federal* variable from Table 3 to a percentage change.

In order to assess the robustness of our results, in this section we estimate several alternate specifications of the basic model.

5.1 Differential Effects by Education

In our model, sectoral compensation differentials are independent of education. This specification runs counter to longstanding concerns that because of inflexibilities in government pay schedules, highly educated public sector workers earn less than their private sector counterparts.[37] We explore this hypothesis by augmenting our basic model with interaction terms between the sectoral dichotomous variables and a binary variable, *CollegePlus,* which equals one if a worker has a four-year college or higher degree, and zero otherwise. We report here only the key results. In the equation for *WAGE* + DC, the interactions of the sectoral dichotomous variables with *CollegePlus* are all statistically insignificant. For pension wealth, the interaction term between *STATE* and *CollegePlus* is positive and significant at $75,820 (s.e. = $27,906), while the other two sectoral interaction terms are insignificant. Thus, we find no evidence in our data that highly educated workers are disadvantaged by working in the public sector.

This specification constrains the coefficients on all the other right hand side variables to be the same across education groups. As an alternative approach, we relax this assumption and estimate the model for *WAGE + DC* separately for each education group. In no education group is there a negative and statistically significant coefficient on either the federal or state sectoral variable. In the very highest education group, there is a statistically significant negative coefficient on the local variable, but we are not inclined to make too much of a single anomalous estimate. In short, our basic conclusion that

[37] See Congressional Budget Office [2002, 2012] and Blue Mass Group [2011] for discussions of this phenomenon at federal and sub-federal levels of government, respectively.

there is no penalty for highly educated workers who work in the public sector is robust to this change in the specification of the model.[38]

5.2 Controlling for Occupation and Firm Size

As noted above, it is controversial whether occupation should be included in models of public sector compensation differentials. We chose not to include controls for occupation, because of concerns that it might be jointly determined with compensation. In addition, in our data, occupation is reported for only 1,453 observations, considerably fewer than the 2,496 used to estimate our basic model in Table 3. Nevertheless, it is of some interest to see if our results change qualitatively when we augment the model with a set of dichotomous variables designating occupational categories.[39] Table 5 displays the results for the logarithm of (*WAGE+DC)*. In columns (1) and (2) we establish a baseline by using the smaller sample to estimate the models from the last two columns of Table 3, and then add the occupation variables in column (3). Only the coefficients on the sectoral variables are reported. The coefficients in columns (1) and (2), while having somewhat different magnitudes, are qualitatively similar to their counterparts in Table 3. Once occupation controls are added in column (3), the implied federal differentials falls slightly and remains statistically significant. The point estimate, 0.265, is virtually the same as the corresponding figure in Table 3, 0.277. The state and local differentials

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[38] To assess the robustness of this result, we also estimated this specification using the CPS data set described in the Appendix. The result was the same. There is no education group for which the coefficient on *Federal* is negative. The *State* coefficient is negative and significant for individuals whose highest educational attainment is a master's degree, but returns to insignificant for those who have a doctorate or professional degree. There is no education group for which the coefficient on *Local* is negative and significant.

[39] An appendix table available upon request provides a list of the occupation categories that are available in the HRS, and the proportion of employees in each category.

remain statistically insignificant, also just as they were in Table 3. In short, our substantive results are essentially unchanged when controls for occupation are included.[40]

Another variable that sometimes appears in empirical studies of compensation, especially with respect to pensions, is firm size. We are inclined to agree with Gittleman and Pierce [2012, p. 226] who argue that including firm size in compensation regressions is not appropriate because it is not a skill-related factor that an individual can transfer from job to job. However, as a check on the robustness of our findings, we estimated the pension wealth regression using the 1,716 observations for which we have information on firm size (indicators for whether the firm is small, medium, or large)[41]. Consistent with previous studies, we find that pension wealth increases with firm size, and the coefficients are statistically significant. With the size of firm variables included, the sectoral differentials are $87,200, $61,900, and $65,600, respectively, for federal, state, and local government, and all are statistically significant at the 0.01 or smaller level. The story is essentially the same as in our basic model in Table 4—large and statistically significant positive differentials for all sectors.

6. Conclusion

We have used a sample of older full-time workers from the 2006 Health and Retirement Study to study wage and pension compensation differentials between government and private sector employees. With respect to hourly wages (including employee contributions to defined contribution plans), federal workers earn a premium of

[40] In the same way, the coefficients on the sectoral variables in our pension wealth equation (column (4) of Table 4) are not substantively changed when we augment it with the occupation variables. Moulton [1990], using CPS data from 1988, finds that including occupation variables has a substantive effect on the public-private differential.

[41] Firms with fewer than 100 employees are classified as small, between 100 and 500, medium, and more than 500 employees, large.

about 28 log points. However, no statistically significant differences for state and local workers emerge once employee characteristics are taken into account. Workers at all levels of government accumulate more pension wealth than private sector workers, even after holding employee characteristics constant. Specifically, as a proportion of the hourly private-sector wage, the hourly equivalent public-private differentials are about 17.2 percent, 13.4 percent and 12.6 percent at the federal, state, and local levels, respectively. Our results offer some support to both sides in the rather noisy public debate over public sector compensation. The argument that taking worker characteristics into account can make a big difference in public-private comparisons is correct. However, those who assert that examining wages alone can be misleading are also right. Once pension benefits are taken into account, employees at all levels of government receive higher compensation than private sector workers, and these differentials cannot be explained away by differences in worker characteristics.

As usual, one must be cautious in interpreting compensation differences. They might be due to compensating differentials associated with unobserved job characteristics, or they might arise from successful rent-seeking. Also, to the extent that public sector employment contracts are more back-loaded than those in the private sector, this could contribute to the presence of public-private differentials. An important subject for future research is to determine the source of these unexplained differences. Further, it would be interesting to apply the methods used in this paper to a sample that includes a larger proportion of younger workers, if a data set with all the requisite variables were to become available.

As mentioned at the outset, we have focused on wage and pension differentials because comparable data on other dimensions of the compensation package are not readily available. Ideally, future work in this area should incorporate other components of compensation, particularly employer-provided health insurance benefits. The HRS does include information on *whether* an individual has health insurance coverage. In results available upon request, we find that government workers are in fact significantly more likely to have employer-provided health insurance than their private sector counterparts with the same characteristics. The differentials are 10.2, 10.0 and 8.1 percentage points for federal, state, and local workers, respectively. However, there is no information on the cost to employers of providing this benefit. Integrating information on the value of employee-provided health insurance with wage and pension differentials within a micro data framework is another important topic for future research.

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Appendix 1

Comparing the HRS and CPS

As detailed in the text, we apply a set of screens to the HRS data in order to obtain our analysis sample. The purpose of this appendix is to investigate whether our HRS analysis sample is statistically similar to a sample generated by applying the same screens to the CPS.

Table A1 exhibits the means of a set of variables that are common to both data sets.[42] Column (1) has the HRS means along with their standard deviations, and column (2) has analogous calculations for the CPS. A glance at the table suggests that the means are quite similar. Indeed, for none of the variables can one reject the hypothesis that the respective HRS and CPS means are equal.

A related question arises in the context of our estimated wage differentials. As shown in Table 3 in the text, once personal characteristics are taken into account, the state and local differentials are small and statistically insignificant, a result not unlike that found in a number of earlier papers. Our federal differential is 0.278 and statistically significant. While a positive and statistically significant federal differential is a common finding, the magnitude of our estimate is somewhat larger than in many papers. One possible explanation for the difference is that the HRS sample is somehow non-representative. Another is that previous papers have not applied the same screens to generate their analysis samples—they generally look at the entire population, while we focus on workers over 50. To investigate this issue, we applied the same screens to the

[42] All means are computed using sample weights. The labeling of the education categories in the two data sets is not exactly the same. We assume that what the CPS refers to as "bachelor degree" is the same as the "4-year college degree" in the HRS, and that "professional degree" in the CPS includes doctoral degrees.

CPS as we did to the HRS, and used this sample to replicate as closely as possible our basic wage regression from column (2) of Table 3. The key results are as follows: The state differential in the CPS is -0.0112 (s.e. = 0.0360) and the local differential is 0.00626 (s.e. = 0.0291). Like their counterparts in Table 3, they are small and statistically insignificant. The federal differential is a statistically significant 0.170 (s.e. = 0.0406). While smaller than the estimate in Table 3, one cannot reject the hypothesis that they are equal-- the chi-square statistic for a test of equality of these two estimates is 2.17, while the critical value at the 0.05 level is 3.84. In short, there is no statistically discernible difference between our HRS estimate of the federal-private differential and an estimate generated by the CPS.

Table 1*

Dependent Variable Definitions and Summary Statistics

Variable	Description	Private Sector	Federal Government	State Government	Local Government
WAGE	Hourly pay in 2006 Dollars	23.22 (31.40)	31.56 (34.54)	44.09 (218.49)	22.70 (10.56)
DBPlan	1 if respondent has primary defined benefit pension plan	0.206 (0.405)	0.556 (0.499)	0.668 (0.472)	0.744 (0.437)
DCPlan	1 if respondent has primary defined contribution pension plan	0.420 (0.494)	0.300 (0.460)	0.241 (0.429)	0.154 (0.361)
HPlan	1 if respondent has primary hybrid pension plan	0.041 (0.197)	0.086 (0.281)	0.045 (0.208)	0.039 (0.194)
*EmpC***	Hourly employer contribution to respondent's DC plans in 2006	2.22 (5.12)	2.11 (4.44)	2.29 (2.66)	0.78 (0.45)
WAGE+DC	Sum of Wage and EmpC	23.95 (32.2)	32.33 (34.67)	44.48 (218.5)	22.78 (10.36)
*PensionW***	Sum of pro-rated present value of future benefits from most important DB plan and cash balances in all DC plans in 2004	113,515 (179,497)	205,682 (218,432)	196,930 (264,745)	190,786 (260,060)
PositivePW	1 if pension wealth > 0	0.416 (0.423)	0.946 (0.226)	0.884 (0.321)	0.867 (0.341)
Number of Observations***		1,907	115	225	249

* This table shows means and standard deviations for dependent variables by employment sector, weighted by 2006 respondent weights. Variables that are measured in dollars (*WAGE, EmpC, WAGE+DC,* and *PensionW*) are in 2006 dollars.
** Means and standard deviations for *PensionW* and *EmpC* are computed over positive observations.*** These are the maximal number of respondents with non-zero sample weights in each sector. However, not every question is answered by all respondents; therefore, the number of observations varies slightly across variables.

Table 2*
Variable Definitions and Summary Statistics

Variable	Description	Private Sector	Federal Government	State Government	Local Government
NoDegree	Omitted Category: 1 if the respondent did not complete high school	0.090 (0.286)	0.016 (0.125)	0.031 (0.173)	0.043 (0.202)
GED	1 if the respondent's highest degree is a GED	0.044 (0.204)	0.010 (0.099)	0.032 (0.177)	0.030 (0.172)
HS	1 if the respondent's highest degree is a high school degree	0.517 (0.500)	0.438 (0.498)	0.359 (0.481)	0.397 (0.490)
TwoCollege	1 if the respondent's highest degree is a 2-year college degree	0.082 (0.274)	0.100 (0.301)	0.045 (0.209)	0.073 (0.260)
FourCollege	1 if the respondent's highest degree is a 4-year college degree	0.170 (0.376)	0.189 (0.393)	0.206 (0.406)	0.185 (0.389)
Masters	1 if the respondent's highest degree is a masters degree	0.078 (0.268)	0.198 (0.400)	0.167 (0.374)	0.239 (0.427)
Professional	1 if the respondent's highest degree is a professional degree	0.021 (0.142)	0.050 (0.219)	0.159 (0.367)	0.035 (0.183)
Black	1 if the respondent is black	0.080 (0.272)	0.146 (0.355)	0.138 (0.346)	0.096 (0.295)

Female	1 if the respondent is female	0.464 (0.499)	0.492 (0.502)	0.605 (0.490)	0.468 (0.500)
Married	1 if the respondent is married	0.674 (0.469)	0.670 (0.472)	0.726 (0.447)	0.678 (0.468)
Age	Age of respondent in years	58.17 (4.596)	58.82 (4.600)	58.21 (4.048)	58.10 (4.780)
Age^2	Age squared	3404.4 (564.13)	3480.8 (567.15)	3404.6 (487.69)	3397.8 (587.67)
Urban	1 if the respondent lives in metro area of 1 million or more population	0.452 (0.498)	0.416 (0.495)	0.340 (0.475)	0.436 (0.497)
Suburban	1 if the respondent lives in metro area of 250,000 to 1 million population	0.266 (0.442)	0.311 (0.465)	0.250 (0.432)	0.210 (0.408)
Ex-Urban	Omitted Category: 1 if the respondent lives in area of less than 250,000 population	0.282 (0.450)	0.273 (0.447)	0.413 (0.493)	0.354 (0.479)

*This table shows means and standard deviations for independent variables by employment sector, weighted by 2006 respondent weights. The maximal number of respondents with non-zero sample weights is 1,907 for the private sector, 115 for federal government, 225 for state government, and 249 for local government. These figures are the same across all variables.

Table 3*
Hourly Compensation

Variable	(1) *Log(WAGE)*	(2) *Log(WAGE)*	(3) *Log(WAGE+DC)*	(4) *Log(WAGE+DC)*
Federal	0.371***	0.278***	0.373***	0.277***
	(0.052)	(0.045)	(0.0523)	(0.0464)
State	0.163***	0.0559	0.152***	0.0379
	(0.046)	(0.0414)	(0.0466)	(0.0418)
Local	0.086**	-0.0134	0.0673*	-0.0339
	(0.037)	(0.0333)	(0.0372)	(0.0338)
GED		0.122***		0.138**
		(0.0394)		(0.0431)
HS		0.318***		0.328***
		(0.0279)		(0.0283)
TwoCollege		0.425***		0.435***
		(0.0455)		(0.0461)
FourCollege		0.639***		0.649***
		(0.0434)		(0.0443)
Masters		0.819***		0.835***
		(0.0525)		(0.054)
Professional		1.000***		1.05***
		(0.0525)		(0.067)
Black		-0.183***		-0.180***
		(0.0314)		(0.034)
Female		-0.190***		-0.190***
		(0.0197)		(0.0194)
Married		0.0758**		0.0789**
		(0.0181)		(0.0182)
Age		0.0174		0.0219
		(0.066)		(0.0659)
Age^2		-0.000177		-0.000216
		(0.000537)		(0.000537)
Urban		0.239***		0.241***
		(0.0311)		(0.0307)
Suburban		0.0828***		0.0904***
		(0.0304)		(0.0306)
Constant	2.93***	2.048	2.956***	1.927
	(0.018)	(2.017)	(0.0183)	(2.015)
Observations	2,496	2,496	2,496	2,496

* Estimation is by ordinary least squares. The right-hand side variables are defined in Table 2. Standard errors are in parentheses. A (***) indicates that the variable is statistically significant at the 1 percent level, a (**) at the 5 percent level, and a (*) at the 10 percent level.

Table 4*
Pension Wealth Differentials

Variable	**(1)** Pr (*Pension Wealth* >0)	**(2)** Amount of *Pension Wealth*	**(3)** Pr (*Pension Wealth* >0)	**(4)** *Pension Wealth*
Federal	0.216*** (0.0314)	111,342*** (26,162)	0.178*** (0.030)	93,636*** (27,297)
State	0.148*** (0.0375)	89,757*** (21,729)	0.0995*** (0.0374)	62,913*** (18,122)
Local	0.129*** (0.028)	80,809*** (16,009)	0.102*** (0.0295)	65,775*** (15,478)
GED			0.0991 (0.0771)	13,292 (14,801)
HS			0.204*** (0.0377)	43,629*** (9,036)
TwoCollege			0.188*** (0.0499)	61,696** (25,981)
FourCollege			0.284*** (0.0433)	71,561*** (15,009)
Masters			0.317*** (0.0482)	125,315*** (18,771)
Professional			0.312*** (0.0642)	206,395*** (44,560)
Black			-0.0386 (0.0323)	3,626 (15,674)
Female			-0.00337 (0.0214)	-46,354*** (6,355)
Married			0.0721*** (0.0252)	29,812*** (8,274)
Age			0.179*** (0.0282)	26,957** (11,427)
Age^2			-0.00157*** (0.000228)	-247.0*** (92.9)
Urban			0.0199 (0.0214)	32,398*** (10,174)
Suburban			0.0356 (0.024)	17,717* (9,764)
Constant	0.727*** (0.0139)	82,490*** (5,639)	-4.604*** (0.864)	-720,121* (349,683)
Observations	2,100	2,100	2,100	2,100

* Columns (1) and (3) provide estimates of the marginal effects on the probability that the respondent has positive pension wealth, based on linear probability models. Columns (2) and (4) are the corresponding OLS estimates of the effects on the amount of pension wealth. The right-hand side variables are defined in Table 2. Standard errors are in parentheses. A (***) indicates that the variable is statistically significant at the 1 percent level, a (**) at the 5 percent level, and a (*) at the 10 percent level.

Table 5*
Hourly Wages plus DC Contributions, Controlling for Occupation*

(Dependent Variable is Log of (*WAGE+DC*))

Variable	**(1)** (no occupation controls; no other covariates)	**(2)** (no occupation controls; same covariates as in Table 3)	**(3)** (occupation controls; same covariates as in Table 3**)**
Federal	0.485***	0.302***	0.265***
	(0.0713)	(0.0647)	(0.0618)
State	0.197***	0.0398	0.0821
	(0.0518)	(0.0511)	(0.0503)
Local	0.0739***	-0.0389	0.0316
	(0.0513)	(0.0498)	(0.0480)
Observations	1,453	1,453	1,453

* This table augments the model of Table 3 with a set of dichotomous variables for occupation. A list of the occupation categories is in Table A2. The number of observations differs from that of the basic sample (2,946) because not every respondent provided information about occupation. The other right-hand side variables are the same as in Table 3, but not reported here. Standard errors are in parentheses. A (***) indicates that the variable is statistically significant at the 1 percent level, a (**) at the 5 percent level, and a (*) at the 10 percent level.

Table A1*
Selected Summary Statistics from the HRS and CPS*

Variable	Description	(1) HRS Sample	(2) CPS Sample
WAGE	Hourly pay in 2006 dollars	25.37 (71.14)	21.93 (12.81)
NoDegree	Omitted Category: 1 if the respondent did not complete high school	0.077 (0.266)	0.079 (0.269)
FourCollege	1 if the respondent's highest degree is a 4-year college degree	0.175 (0.380)	0.193 (0.395)
Masters	1 if the respondent's highest degree is a masters degree	0.107 (0.309)	0.104 (0.305)
Professional	1 if the respondent's highest degree is a professional degree	0.035 (0.185)	0.047 (0.212)
Black	1 if the respondent is black	0.090 (0.286)	0.096 (0.295)
Female	1 if the respondent is female	0.478 (0.500)	0.464 (0.499)
Married	1 if the respondent is married	0.679 (0.467)	0.701 (0.458)
Age	Age of respondent in years	58.19 (4.57)	56.42 (5.32)
Age^2	Age squared	3407.01 (560.29)	3212.1 (637.57)

*Column (1) reproduces the weighted means and standard deviations of selected variables in our HRS analysis sample from Tables 1 and 2. Column (2) exhibits means and standard deviations computed from a sample generated by applying the same screens to the 2006 CPS. The number of respondents with non-zero respondent weights is 2,496 in the HRS sample and 3,000 in the CPS sample.