

Wages, Pensions, and Public-Private Sector  
Compensation Differentials

by

Philipp Bewerunge, Neuss Associates  
Harvey S. Rosen, Princeton University

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## Abstract

### Wages, Pensions, and Public-Private Sector Compensation Differentials

We use a sample of full-time workers mostly 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 39 percent, taking differences in employee characteristics into account. The differential varies by gender and race. However, there are no statistically discernible differences between state and local workers and their private sector counterparts, *ceteris paribus*. 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 14.3 percent for federal workers, 8.3 percent for state workers, and 8.6 percent for local workers, *ceteris paribus*. Again, there are differences by gender and race. 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.

Philipp Bewerunge  
Neuss Associates  
64 MacDougal Street, Apt. 16  
New York, NY 10012

[pbewerunge@gmail.com](mailto:pbewerunge@gmail.com)

Harvey S. Rosen  
Department of Economics  
Princeton University  
Princeton, NJ 08544

[HSR@Princeton.Edu](mailto:HSR@Princeton.Edu)

## 1. Introduction

The dire economic conditions that arose in the aftermath of the 2008 financial crisis have stimulated a heated debate about the compensation of public employees. The issue has received particular attention because of the severe budgetary constraints facing governments at all levels. For example, commenting on the disturbances in the state of Wisconsin during February of 2011, Governor Mitch Daniels of Indiana argued: “The people who are doing the demonstrating, and their allies...spent that state broke ... These folks have really ruled the roost in Wisconsin, and other places” [Haberman, 2011]. Daniels continued along this vein in another interview, “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]. Confronted with the charge of being overpaid, a paramedic at the same rally remarked, “I drove my Ford Focus [to the capitol building in Madison]. I live in a 950-square-foot condominium!” [Delaney, 2011]. Others say that the uproar over public sector compensation simply stems from envy of government employees’ rightful compensation: “We had the promise of stable retirement [...] People hate to see someone doing better than they are” [Delaney, 2011]. Al

Martinez, president of the Service Employees International Union Local 1107, saw public sector compensation as an arbitrary target for budget cuts: “Why does it have to be on the backs of our employees?... We need to hold the line and maintain what we have” [Cook, 2011].

Putting the inflammatory rhetoric aside, much of the debate boils down to the question of whether public and private sector workers receive about the same compensation. As New Jersey Governor Chris Christie expressed it, “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” [2010]. Exploring this issue 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]. Indeed, if public sector workers have substantially more education than private sector employees, they might be *underpaid* compared to what they would earn in the private sector, a point made by Nobel laureate Paul Krugman [2010]: “[T]hose workers aren’t overpaid. Federal salaries are, on average, somewhat less than those of private-sector workers with equivalent qualifications” [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. Fletcher and Somashekhar [2011] note that, “as a group, public employees generally earn less than comparably educated private-sector employees, but they tend to enjoy far better health-care and retirement benefits.” 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.” Less hyperbolically, Barro and McMahon [2010] note that the defined contribution plan benefits received by most private-sector workers are, on average, “significantly less generous than the pensions provided to government workers.” Public-sector unions argue that such criticism is 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 analysis of 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. 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, and presents summary statistics for the key variables. Section 4 presents the econometric setup and basic results. We find that the hourly wages of federal government workers are about 39 percent more than private sector workers with similar characteristics, *ceteris paribus*. The results differ by gender and race, ranging from about 27 percent for white males to 62 percent for black females. 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 14.3

percent for federal workers, 8.3 percent for state workers, and 8.6 percent for local workers, although these figures mask some differences across demographic groups.

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, we find no evidence that highly educated workers are disadvantaged by working in the public sector. Section 6 concludes with a discussion of the implications of the results for the ongoing public debate on 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 Current Population Survey (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,<sup>1</sup> by using panel data to estimate fixed effects models,<sup>2</sup> and by estimating models on a state-by-state basis to allow for the possibility that labor market institutions, and hence public sector wage differentials, vary across states.<sup>3</sup> 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.<sup>4</sup> 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

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<sup>1</sup> See Quinn [1979], Gyourko and Tracy [1986], Venti [1987], and Moore and Raisin [1991].

<sup>2</sup> See Krueger [1988].

<sup>3</sup> See Belman and Heywood [1995], Schmitt [2010], and Lewis and Galloway [2011].

<sup>4</sup> Hence, the opposite characterization of the literature by Peter Orszag and Paul Krugman (see section 1 above) seems rather curious.



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.<sup>5</sup> 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 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 defined contribution plans. While such a decision was sensible at the time Quinn was writing, this is no longer the case, as defined contribution 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.

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<sup>5</sup> Quinn's calculation eliminates the average sectoral portions of pension wealth that are associated with workers' contributions.

### 2.3 Aggregating Wage and Pension Differentials

Several recent papers have sought to integrate wages with fringe benefits in order to obtain a broad measure of sectoral compensation differentials. 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,<sup>6</sup> 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.<sup>7</sup>

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 firm 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. As in the Ramoni-Perazzi and Bellante study, the adjustment for fringe benefits does not take into account employees’ personal characteristics, thus ignoring the substantial worker heterogeneity within occupation/firm size categories. Other studies following the same general approach are Allegretto and Keefe [2010], Bender and Heywood [2010], Thompson and Schmitt [2010], Cannon [2010] and

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<sup>6</sup> The key variable, sector of employment, is not included in the data set; so in effect, the authors have to impute it.

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

Gittleman and Pierce [2011], 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 consistent multivariate econometric framework based on national micro-level data. Our data allow us to apply the same econometric approach to 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 every two years by the Institute for Social Research at the University of Michigan.<sup>8</sup> Because the HRS tracks everyone in the respondent's household, a number of younger workers are also included. The employment questions in the HRS are asked of the "financial respondent," who is not necessarily the person originally selected for the study and can be under 50. In our analysis sample, about 7.4 percent of the individuals are under 50. Because the HRS is primarily based on older workers, our results might not apply to employees throughout the age distribution.<sup>9</sup> However, in addition to standard demographic data, it has information on sector of employment<sup>10</sup> as well as a

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<sup>8</sup> Quinn's [1979] dataset also includes mainly mature workers. Bureau of Labor Statistics data indicate that the proportion of the workforce over 50 years is about 30 percent (<http://www.bls.gov/cps/cpsaat03.pdf>).

<sup>9</sup> A related phenomenon is that 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 471 government workers we are able to track over time, only 13 individuals made such transitions between 2006 and 2008

<sup>10</sup> 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?"). In the absence of information on sector of employment, some previous studies have imputed it statistically (Ramoni-Perazzi and Bellante [2007]).

comprehensive 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 supplement constructed by Gustman, Steinmeier, and Tabatabai [2010]. Respondents' reports in the HRS of the parameters that determine the value of their pension plans are characterized by some error, but do not appear to be systematically biased (Gustman and Steinmeier [2004]).

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. This leaves us with 3,199 observations. Furthermore, we exclude respondents who, while earning a positive hourly wage, characterized themselves as retired<sup>11</sup>. Finally, we drop workers for whom information is missing for any of our right hand side variables. This leaves us with 3,078 observations, which comprise our

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<sup>11</sup> 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.

basic analysis sample. There are 2,382 respondents in the private sector, 126 in federal government, 271 in state government, and 299 in local government.<sup>12</sup>

Table 1 presents definitions and weighted 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 (42 respondents did not know their primary plan type). Local government workers are the most likely to have a defined benefit plan and private sector workers are the most likely to have a defined contribution 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. Respondents can provide information on employer pension contributions for up to four DC plans as either 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.<sup>13</sup> In

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<sup>12</sup> These figures differ somewhat from those in Table 1 because the computations in Table 1 use only individuals with non-zero sample weights, while these figures include all individuals in the sample.

<sup>13</sup> 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

total, we were able to calculate positive hourly employer contributions for 814 workers out of the 1,278 workers 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<sup>14</sup> from a respondent's most important DB plan and the cash balances in all DC plans on the current job in 2004 (in 2006 dollars). We use the 2004 value because it is the most recent year that GST calculate it for all workers.<sup>15</sup> *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--2,505 of the 3078 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

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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.

<sup>14</sup> 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. While GST adjust the benefits to take into account expected mortality rates, 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.

<sup>15</sup> 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.

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.

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 the private sector and various levels of government.<sup>16</sup> 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, and age. Women account for a large share of the state government workforce (60.4 percent), but smaller shares of the federal workforce (49 percent), local workforce (47.5 percent), and private sector workforce (46.7 percent). Federal and state government have the highest proportions of black workers (14.6 and 14.3 percent, respectively), followed by local government (9.6 percent), and the private sector (8.3 percent). With regards to marital status, state government has the largest proportion of married workers (72.5 percent), while the proportions in federal government (67.1 percent), local government (67.7 percent) and the private sector (66.8 percent) are nearly identical. Workers’ ages are essentially the same across sectors, although one must keep in mind that our data set includes mainly individuals who are 50 years of age or older.

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<sup>16</sup> The variables are available for all the observations in the analysis sample.

## 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 when we estimate multivariate models.<sup>17</sup>

### 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 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 [2011, p. 225]).<sup>18</sup> 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 educational attainment, gender, race, marital status, and a quadratic in age.<sup>19</sup> To allow for the possibility that sectoral compensation differentials depend on gender and race, our basic specification also includes interactions between the sectoral variables and indicators for gender (equals one

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<sup>17</sup> The HRS includes a set of sample weights. While we present unweighted regressions here, for the sake of completeness we also estimated weighted regressions. In results available upon request, we find that incorporating the HRS sample weights does not affect our basic conclusions with respect to sectoral differentials.

<sup>18</sup> 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 [2011, 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.

<sup>19</sup> For a subset of our observations, we have years of experience on the current job. 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.



if the respondent is female) and race (equals one if the respondent is black).<sup>20</sup> In effect, this specification constrains all the coefficients, except those associated with 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 generated essentially the same results as the Blinder-Oaxaca approach and is simpler to exposit.

Unlike some previous studies, we do not include occupational variables in our canonical model. As Gittleman and Pierce [2011, 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.<sup>21</sup> However, as shown in Section 5 below, when we include occupation on the right hand side, the substantive results with respect to sectoral differences are largely unchanged.

In a regression with the logarithm of wages on the left hand side, the coefficient on each dichotomous variable is approximately the percentage change in the wage rate associated with this variable. The approximation is close only when the coefficient is relatively “small.” Many of the estimated coefficients are too large for the approximation

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<sup>20</sup> 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 989 observations: only 2,089 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 1,618 private-sector workers in 2006, 3 move into federal government, 5 move into state government, and 6 move into local government in 2008. Out of 83 federal workers in 2006, 1 moves into the private sector, and 1 moves into state government in 2008. Out of 178 state workers in 2006, 7 move into the private sector in 2008. Out of 210 local workers in 2006, 5 move into the private sector, and 1 moves into state government in 2008. This paucity of sectoral transitions would result in imprecise estimates of the sectoral differentials.

<sup>21</sup> 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 [2011, 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.”

to hold. Hence, we report exact computations of the percentage changes rather than relying on the approximation.<sup>22</sup> The standard errors are computed using the delta method.

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 69.1 percent more than private sector employees, state government employees earn 24.6 percent more, and local government employees earn 16.3 percent 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. For convenience, we do not report the coefficients themselves. Rather, we use the regression coefficients to compute the various public-private wage differentials for each demographic group. Consider, for example, the entry of 0.489 (s.e. = 0.0685) for white females in the third row under the heading “Federal.” It is calculated as the sum of the coefficient on the main effect of employment in the federal government (the dichotomous variable *Federal*, which equals one if the respondent is a federal worker) plus the coefficient on the interaction between *Female* and *Federal*. Hence, the entry indicates that white women who work in the federal government earn 48.9 percent more than white women who work in the private sector. Similarly, our estimates imply federal-private sector differentials of 27.2 percent, 40.2 percent, and 61.9 percent, respectively, for white males, black males, and black females, respectively. All of these effects are statistically significant at conventional levels. The overall federal-private

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<sup>22</sup> More precisely, if the estimated coefficient is  $\beta$ , then the percentage change is  $e^{\beta}-1$ .

differential, found by weighting the effects for each demographic group by their respective population proportions<sup>23</sup>, is 38.8 percent.

The figures in column (2) under “State” and “Local” are defined analogously. Unlike the case for federal government, the state and local government differentials are generally rendered 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 differences without taking into account worker characteristics can be misleading—the overall federal differential diminishes substantially, and the state and local differentials become statistically indistinguishable from zero. Our finding of a large and statistically significant federal differential is generally in line with previous results, most notably by Smith [1976a, 1976b, 1977], Venti [1987], Krueger [1988], and Moore and Raisian [1991], although our estimate is somewhat larger than theirs. On the other hand, our results run counter to the assertions made by Orszag ([Tuutti, 2010]) and Krugman [2010] that the positive wage differential of federal workers can be explained by their superior human capital relative to private sector workers.

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 twice as much. Married workers enjoy an

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<sup>23</sup> For these purposes, we use the sample weights included in the HRS.

earnings premium of 7.33 percent, and wage rates follow the usual inverted U-shape with age.<sup>24</sup>

#### 4.2 Hourly Wages and Employer Contributions to DC Plans

We now turn to differences in compensation that come in the form of pensions.<sup>25</sup> Integrating employer contributions to defined contribution plans with wages is fairly 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)<sup>26</sup> 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 defined contribution plans. In principle, one would also want to

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<sup>24</sup> The negative quadratic term starts to dominate the positive term at 58 years.

<sup>25</sup> 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.

<sup>26</sup> 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.

include the accrued value of employer contributions to defined benefit plans, 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. This section examines how total pension wealth depends on sector of employment.

As noted above, just like wages, pension wealth likely depends on an employee's personal characteristics. Gittleman and Pierce [2011, 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, it is the appropriate estimator.<sup>27</sup>

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

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<sup>27</sup> However, when we re-estimated the model using a Tobit estimator, the marginal effects were qualitatively quite similar to those from OLS.

without any additional covariates. Compared to private sector employees, federal workers are 27.3 percentage points more likely to have positive pension wealth, state workers are 17 percentage points more likely, and local workers are 14 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 \$117,511 more pension wealth, state workers have \$76,282 more, and local workers have \$62,795 more.

The setup in columns (3) and (4) is the same as in Table 3, that is, the figures show the sectoral differences by demographic group, computed as the sums of the relevant coefficients from the underlying regressions. The most striking result is that the uncorrected pension wealth differentials from column (2) are reduced yet generally remain substantial and statistically significant once personal characteristics are added to the model. Specifically, column (4) shows federal, state and local pension differentials of \$96,738, \$46,742, and \$54,912, 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 in sector  $s$  the pension wealth differential 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 3 percent.

We illustrate using our results for the overall sample. The mean years of service with the current employer for our 2004 pension wealth subsample<sup>28</sup> are: 12.9 for private sector workers, 17.1 for federal workers, 13.9 for state workers, and 16.2 for local workers. Using these values and the respective “overall” pension wealth differentials from Table 4, the annuitized differentials are \$7,101, \$4,041 and \$4,203 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, 2,195 for federal workers, 2,141 for state workers, and 2,151 for local workers. Therefore, the annuitized pension wealth differentials per hour come to \$3.23, \$1.89 and \$1.95 for federal, state, and local workers, respectively. As a proportion of the hourly private-sector wage, these amount to 14.3 percent, 8.3 percent, and 8.6 percent for federal, state, and local workers, respectively.

#### 4.4 Tentative Conclusions

Summary. Taken together, the results in this section 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, with the exception of black female local workers, who have 22.8 percent higher wages. There is, however, a substantial wage differential for federal workers, *ceteris paribus*, of about 39 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

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<sup>28</sup> 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.

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.<sup>29</sup> 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 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:

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<sup>29</sup> We add the annuitized pension wealth differentials to the differentials from the second rather than the fourth column of Table 3 because the latter include 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.



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, as shown in the Appendix, BLS data allow us to make a rough imputation.

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. The Appendix 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 48.7 percent for federal employees, 6.39 percent for state employees, and 6.33 percent for local employees.

## 5. Alternative Specifications

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.<sup>30</sup> 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 interaction term between *CollegePlus* and *Federal* is negative and statistically significant. The coefficients imply that the differential between highly educated federal and private sector workers is 28.8 percent (s.e. = 5.33 percent); for those with less than college education it is 43.2 percent (s.e. = 6.14 percent). Thus, while the differential is smaller for workers with college or more education, it is still positive and significant. The respective interactions of *STATE* and *LOCAL* with *CollegePlus* are statistically insignificant. Turning now to pension wealth, the interaction terms with *CollegePlus* are positive for each sector, but one cannot reject the hypothesis that they are zero. Taking the results for *WAGE + DC* and pension wealth together, the basic conclusion is that we find no evidence in our data that highly educated workers are disadvantaged by taking work in the public sector.

### 5.2 Controlling for Occupation

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 1,777 observations, considerably fewer than the 3,078 used to estimate our basic model in Table 3. Nevertheless, it is of

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<sup>30</sup> 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.

some interest to see if our results change qualitatively when we augment the model with a set of dichotomous variables designating occupational categories<sup>31</sup>.

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 qualitative similar to their counterparts in Table 3. Once occupation controls are added in column (3), the implied federal differentials remain significant and about the same, and the state and local differentials remain statistically insignificant. In short, our results are generally unchanged when controls for occupation are included.

### 5.3 Other Worker Characteristics: Intelligence and Risk Aversion

Selection into public sector employment might be correlated with unobserved worker characteristics that are also correlated with compensation. For example, a person's intelligence or risk preferences could influence both her choice of sector and her productivity. This potential problem, of course, is common to all research in this area. Two variables in the 2006 wave of the Health and Retirement Survey allow us to make at least a preliminary assessment of its importance in our data. Specifically, for 341 of our observations, the HRS reports the outcome of a test of "fluid intelligence." It quantifies the respondent's ability to retrieve information correctly by using results from number series questions.<sup>32</sup> In addition, for 290 observations, it has the results from a series of exercises designed to assess the respondent's degree of risk aversion by confronting her

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<sup>31</sup> The definitions of the occupation controls are in the footnote to Table 5.

<sup>32</sup> For further details, see McArdle, Smith and Willis [2009].

with hypothetical choices between a level of certain income and a lottery that involves risk but has a higher expected payoff than the certain income.

To examine whether these attributes might be affecting our sectoral differentials, we begin by estimating our models from Tables 3 and 4 with the observations for which the fluid intelligence measure are available, and then with the observations for which we have the risk aversion score. We then re-estimate the models including the fluid intelligence and risk aversion variables, respectively. Because the samples are only a fraction of the size of those used to estimate our basic models, it is no surprise that the coefficients in the baseline model are imprecisely estimated. That said, neither of these variables has any substantial effect on the sectoral coefficients. (The intelligence score itself has a positive and precisely estimated coefficient; the risk aversion measure is not statistically significant.) Without making too much of it, then, we find no evidence that possible sorting of workers on the basis of intelligence and risk aversion biases our estimates of sectoral compensation differentials.

## 6. Conclusion

We have used a sample of 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 between 26 and 51 percent, depending on their gender and race, *ceteris paribus*. 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 14.3 percent, 8.3 percent and 8.6 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 generally receive higher compensation than private sector workers, and these differentials cannot be explained away by sectoral 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. An important subject for future research is to determine the source of these unexplained differences. As well, 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 in Section 1, 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 12.0, 12.6 and 10.6 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 an important topic for future research.

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## Appendix

This Appendix details the computations to integrate our results on pension wealth differentials with those on wages plus defined contributions.

To compute the proportions of the hour-equivalent pension wealth differentials that are associated with employee compensation, we combine information from several sources. 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. We find that average hourly employer DC contributions are \$0.79 for federal workers, \$0.38 for state workers, and \$0.07 for local workers and the respective employee DC contributions are \$0.83, \$0.55, and \$0.33.

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, BLS data report the average hourly employer cost for state and local workers' DB plans (\$2.27 in 2006)<sup>33</sup> and the average required employee contribution rate for state and local workers' DB plans (6.3 percent in 2007)<sup>34</sup>. 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 percentage of DB plans that require a contribution (77 percent in 2007)<sup>35</sup>, which yields an average employee contribution rate of 4.85 percent across all workers. Since the BLS

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<sup>33</sup> U.S. Bureau of Labor Statistics, *Employer Cost For Employee Compensation - March 2006*, June 2006, p.8.

<sup>34</sup> 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.

<sup>35</sup> 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.

doesn't publish corresponding data for federal workers, we assume that their average required DB contribution rate and their employers' average hourly DB cost are the same. 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 (0.682 in federal government, 0.679 in state government, and 0.767 in local government), we calculate hourly employee DB contributions of \$1.04 for federal workers, \$1.43 for state workers, and \$0.83 for local workers.

Third, we multiply the respective estimated hourly pension wealth differentials computed in Section 4.3 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 of 9.85 percent for federal workers, 6.39 percent for state workers, and 6.33 percent for local workers.

The last step is to add these figures to the overall wage differentials from column (2) of Table 3. As noted in the text, 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 process yields total compensation differentials of 48.65 percent for federal government employees, 6.39 percent for state government employees, and 6.33 percent for local government employees.

Table 1\*

## Dependent Variable Definitions and Summary Statistics

Variable	Description	Private Sector	Federal Government	State Government	Local Government
<i>WAGE</i>	Hourly pay in 2006 dollars	22.64 (31.00)	31.55 (34.48)	43.51 (216.81)	22.31 (10.56)
<i>DBPlan</i>	1 if respondent has primary defined benefit pension plan	0.201 (0.401)	0.558 (0.499)	0.664 (0.473)	0.737 (0.441)
<i>DCPlan</i>	1 if respondent has primary defined contribution pension plan	0.410 (0.492)	0.299 (0.460)	0.242 (0.429)	0.153 (0.361)
<i>HPlan</i>	1 if respondent has primary hybrid pension plan	0.039 (0.194)	0.085 (0.281)	0.045 (0.207)	0.038 (0.192)
<i>EmpC**</i>	Hourly employer contribution to respondent's DC plans in 2006	2.20 (5.09)	2.11 (4.44)	2.29 (2.66)	0.761 (0.462)
<i>WAGE+DC</i>	Sum of Wage and EmpC	23.35 (31.8)	32.32 (34.61)	43.89 (216.8)	22.39 (10.56)
<i>PensionW**</i>	Sum of pro-rated present value of future benefits from most important DB plan and cash balances in all DC plans in 2004	113,020 (178,969)	206,475 (218,425)	194,983 (264,013)	189,600 (258,862)
<i>PositivePW</i>	1 if pension wealth > 0	0.766 (0.423)	0.947 (0.226)	0.883 (0.322)	0.863 (0.344)

Number of Observations***	2,016	116	231	256
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\* 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
<i>NoDegree</i>	Omitted Category: 1 if the respondent did not complete high school	0.101 (0.301)	0.016 (0.125)	0.034 (0.180)	0.047 (0.212)
<i>GED</i>	1 if the respondent's highest degree is a GED	0.045 (0.207)	0.010 (0.099)	0.032 (0.176)	0.031 (0.172)
<i>HS</i>	1 if the respondent's highest degree is a high school degree	0.512 (0.500)	0.436 (0.498)	0.358 (0.480)	0.404 (0.492)
<i>TwoCollege</i>	1 if the respondent's highest degree is a 2-year college degree	0.081 (0.273)	0.100 (0.301)	0.045 (0.207)	0.071 (0.258)
<i>FourCollege</i>	1 if the respondent's highest degree is a 4-year college degree	0.165 (0.372)	0.188 (0.393)	0.211 (0.409)	0.180 (0.385)
<i>Masters</i>	1 if the respondent's highest degree is a masters degree	0.076 (0.264)	0.200 (0.402)	0.165 (0.372)	0.233 (0.424)
<i>Professional</i>	1 if the respondent's highest degree is a professional degree	0.020 (0.141)	0.050 (0.218)	0.157 (0.365)	0.034 (0.181)
<i>Black</i>	1 if the respondent is black	0.083 (0.275)	0.146 (0.355)	0.143 (0.351)	0.096 (0.296)

<i>Female</i>	1 if the respondent is female	0.467 (0.499)	0.490 (0.502)	0.601 (0.491)	0.475 (0.500)
<i>Married</i>	1 if the respondent is married	0.668 (0.471)	0.671 (0.472)	0.725 (0.447)	0.677 (0.469)
<i>Age</i>	Age of respondent in years	58.26 (4.735)	58.81 (4.589)	58.21 (4.050)	58.10 (4.750)
<i>Age<sup>2</sup></i>	Age squared	3416.9 (584.29)	3480.1 (566.38)	3404.2 (487.9)	3397.7 (583.6)

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\*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 2,016 for the private sector, 116 for federal government, 231 for state government, and 256 for local government. These figures are the same across all variables.



**Table 3\***  
**Hourly Compensation Differentials**

Variable	(1) <i>Log(WAGE)</i>	(2) <i>Log(WAGE)</i>	(3) <i>Log(WAGE+DC)</i>	(4) <i>Log(WAGE+DC)</i>
<i>Federal Overall</i>	0.691*** (0.0810)	0.388*** (0.047)	0.697*** (0.0810)	0.3902*** (0.047)
<i>State Overall</i>	0.246*** (0.0492)	0.0332 (0.0435)	0.240*** (0.0495)	0.0254 (0.0433)
<i>Local Overall</i>	0.163*** (0.0370)	0.0206 (0.0259)	0.145*** (0.0365)	0.0046 (0.0255)
<u><i>Federal</i></u>				
<i>White Males</i>		0.272*** (0.0735)		0.258*** (0.0724)
<i>White Females</i>		0.489*** (0.0685)		0.512*** (0.0721)
<i>Black Males</i>		0.402*** (0.15)		0.367** (0.147)
<i>Black Females</i>		0.619*** (0.193)		0.621*** (0.193)
<u><i>State</i></u>				
<i>White Males</i>		0.00803 (0.0749)		0.00265 (0.0746)
<i>White Females</i>		0.0777* (0.0464)		0.0698 (0.0465)
<i>Black Males</i>		-0.0831 (0.0802)		-0.101 (0.0801)
<i>Black Females</i>		-0.0134 (0.065)		-0.0342 (0.065)
<u><i>Local</i></u>				
<i>White Males</i>		-0.0363 (0.0391)		-0.0550 (0.0384)
<i>White Females</i>		0.0476 (0.0409)		0.0349 (0.0414)

<i>Black Males</i>		0.144 (0.0923)		0.123 (0.0925)
<i>Black Females</i>		0.228*** (0.0883)		0.213** (0.089)
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<i>GED</i>		0.159*** (0.0497)		0.176*** (0.0534)
<i>HS</i>		0.408*** (0.0360)		0.420*** (0.0370)
<i>TwoCollege</i>		0.644*** (0.0646)		0.665*** (0.0665)
<i>FourCollege</i>		1.002*** (0.0678)		1.01*** (0.0691)
<i>Masters</i>		1.488*** (0.0954)		1.52*** (0.0982)
<i>Professional</i>		1.959*** (0.179)		2.09*** (0.192)
<i>Married</i>		0.0733*** (0.0214)		0.0786*** (0.0218)
<i>Age</i>		0.0705*** (0.0154)		0.0733*** (0.0155)
<i>Age<sup>2</sup></i>		-0.000608*** (0.000127)		-0.000632*** (0.000128)
<i>Constant</i>	2.8*** (0.012)	0.604 (0.402)	2.821*** (0.0123)	0.543 (0.543)
Observations	3,078	3,078	3,078	3,078

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\* Columns (2) and (4) summarize the results from regressions which include, *inter alia.*, interactions between the sectoral variables and indicators for gender (equals one if the respondent is female) and race (equals one if the respondent is black). The entries are computed from the regression coefficients, and show the various overall public-private wage differentials as well as the differentials by demographic group. For example, the entry 0.272 for white males under *Federal* in column (2) indicates that white males in the federal government have hourly wages that are 27.2 percent higher than wages for white males in the private sector. Other 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 ( <i>Pension Wealth &gt;0</i> )	(2) Amount of <i>Pension Wealth</i>	(3) Pr ( <i>Pension Wealth &gt;0</i> )	(4) <i>Pension Wealth</i>
<i>Federal</i>	0.273*** (0.0340)	117,511*** (19,314)	0.181*** (0.0293)	96,738*** (21,810)
<i>State</i>	0.170*** (0.0298)	76,282*** (15,311)	0.0926*** (0.0283)	46,742** (16,370)
<i>Local</i>	0.140*** (0.0292)	62,795*** (14,393)	0.111*** (0.0267)	54,912*** (15,255)
<u><i>Federal</i></u>				
<i>White Males</i>			0.165*** (0.0395)	114,667*** (37,398)
<i>White Females</i>			0.0184*** (0.0442)	83,152*** (21,005)
<i>Black Males</i>			0.246*** (0.0773)	84,395** (36,405)
<i>Black Females</i>			0.264*** (0.0712)	52,880* (29,272)
<u><i>State</i></u>				
<i>White Males</i>			0.0613 (0.0433)	55,765** (29,067)
<i>White Females</i>			0.0986*** (0.0354)	41,001*** (15,586)
<i>Black Males</i>			0.206*** (0.0657)	34,792 (33,914)
<i>Black Females</i>			0.243*** (0.0558)	20,029 (26,710)
<u><i>Local</i></u>				
<i>White Males</i>			0.0865** (0.0411)	86,080*** (28,326)

<i>White Females</i>			0.135*** (0.0352)	15,976 (13,433)
<i>Black Males</i>			0.0982 (0.0837)	113,632* (63,456)
<i>Black Females</i>			0.147** (0.0716)	43,528 (43,119)
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<i>GED</i>			0.157*** (0.0567)	6,078 (8,286)
<i>HS</i>			0.203*** (0.0319)	33,373*** (5,258)
<i>TwoCollege</i>			0.252*** (0.0446)	50,985** (21,568)
<i>FourCollege</i>			0.288*** (0.0361)	65,632*** (10,113)
<i>Masters</i>			0.322*** (0.0386)	111,970*** (15,818)
<i>Professional</i>			0.346*** (0.0465)	202,368*** (34,611)
<i>Married</i>			0.0206 (0.0205)	13,213** (5,448)
<i>Age</i>			0.286*** (0.0192)	37,294*** (3,683)
<i>Age<sup>2</sup></i>			-0.00238*** (0.000164)	-322.3*** (60.7)
<i>Constant</i>	0.589*** (0.0114)	60,660*** (3,312)	-8.080*** (0.560)	-1,035,000*** (105,454)
Observations	2,505	2,505	2,505	2,505

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\* 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. Columns (3) and (4) summarize the results from regressions which include, *inter alia.*, interactions between the sectoral variables and indicators for gender (equals one if the respondent is female) and race (equals one if the respondent is black). The entries

are computed from the regression coefficients, and show the various overall public-private wage differentials as well as the differentials by demographic group. For example, the entry 114,667 for white males under *Federal* in column (4) indicates that white males in the federal government have pension wealth that is \$114,667 higher than pension wealth for white males in the private sector. Other 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)
<i>Federal</i>	0.852*** (0.133)	0.436*** (0.0664)	0.357*** (0.0586)
<i>State</i>	0.290*** (0.0686)	0.0369 (0.0509)	0.079 (0.0543)
<i>Local</i>	0.130*** (0.0503)	0.000679 (0.0369)	0.0628 (0.042)
<u><i>Federal</i></u>			
<i>White Males</i>		0.218** (0.0979)	0.220** (0.0890)
<i>White Females</i>		0.652*** (0.119)	0.475*** (0.0948)
<i>Black Males</i>		0.325 (0.221)	0.376 (0.232)
<i>Black Females</i>		0.759** (0.322)	0.631** (0.314)
<u><i>State</i></u>			
<i>White Males</i>		0.0162 (0.0647)	0.0655 (0.0667)
<i>White Females</i>		0.0791 (0.817)	0.127 (0.0836)
<i>Black Males</i>		-0.0878 (0.115)	-0.107 (0.0967)

<i>Black Females</i>		-0.0249 (0.0935)	-0.0463 (0.0905)
<u>Local</u>			
<i>White Males</i>		-0.0596 (0.0575)	0.0306 (0.0649)
<i>White Females</i>		0.0419 (0.0598)	0.0809 (0.0548)
<i>Black Males</i>		0.0657 (0.122)	0.118 (0.104)
<i>Black Females</i>		0.167 (0.116)	0.168* (0.0977)
Observations	1,777	1,777	1,777

\* This table augments the model of Table 3 with a set of dichotomous variables for occupation. Columns (1) through (3) report OLS estimates of the percentage effects on the sum of hourly wages and employer DC contributions. The number of observations differs from that of the basic sample (3,078) because not every respondent provided information about occupation type. Columns (2) and (3) summarize the results from regressions which include, *inter alia.*, interactions between the sectoral variables and indicators for gender (equals one if the respondent is female) and race (equals one if the respondent is black). The entries are computed from the regression coefficients, and show the various overall public-private wage differentials as well as the differentials by demographic group. For example, the entry 0.220 for white males under *Federal* in column (3) indicates that white males in the federal government have hourly wages that are 22.2 percent higher than wages for white males in the private sector. Other right-hand side variables are the same as in Table 3, but not reported here. The occupation variables included in column (3), but not shown for the sake of brevity, are defined as follows: Business Operations Specialists (Occ2); Financial Specialists (Occ3); Computer and Mathematical Occupations (Occ4); Architecture and Engineering Occupations (Occ5); Life, Physical, and Social Science Occupations (Occ6); Community and Social Services Occupations (Occ7); Legal Occupations (Occ8); Education, Training, and Library Occupations (Occ9); Arts, Design, Entertainment, Sports, and Media Occupations (Occ10); Healthcare Practitioners and Technical Occupations (Occ11); Healthcare Support Occupations (Occ12); Protective Service Occupations (Occ13); Food Preparation and Serving Occupations (Occ14); Building and Grounds Cleaning and Maintenance Occupations (Occ15); Personal Care and Service Occupations (Occ16); Sales Occupations (Occ17); Office and Administrative Support Occupations (Occ18); Farming, Fishing, and Forestry Occupations (Occ19); Construction Trades (Occ20); Extraction Workers (Occ21); Installation, Maintenance, and Repair Workers (Occ22); Production Occupations (Occ23); Transportation and Material Moving Occupations (Occ24); Military Specific Occupations (Occ25). Management Occupations (Occ1) is the omitted category. 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.