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AN ANALYSIS OF PUBLIC AND PRIVATE SECTOR WAGES ALLOWING FOR ENDOGENOUS CHOICES OF BOTH GOVERNMENT AND UNION STATUS

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ABSTRACT

Studies of public/private sector wage differentials typically assume that the government and union status of a worker are exogenous variables. Recently, some studies have relaxed this assumption slightly by allowing the union status to be endogenous. In this paper, we consider a more general selection model in which a worker selects among four labor markets: private/nonunion, private/union, public/nonunion and public/union. A multinomial logit model is estimated to capture this selection decision. Consistent wage equation estimates are then derived using a generalization of the now familiar two-step estimation procedure. Some evidence is found for selection bias in the private/nonunion and the public/union sectors. The pattern of these selection effects produces larger union wage premiums in the public as compared to the private sector. While this is in contrast to the standard findings, the standard errors on the public sector union wage differentials are quite high. In addition, the data indicates that the public/private sector wage differential is largest for federal workers despite the "comparability" process determining their wages.

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I. INTRODUCTION

Over the past several decades, the public sector has experienced both a rapid increase in employment as well as a growing rate of unionization. These developments have focused attention on comparisons between public and private sector labor markets. In this paper we consider three such comparisons. First we examine whether public sector unions have been as successful as private sector unions at generating wage differentials for their members. Second, we examine the level of wages paid in the public and private sector. This comparison is motivated by the fact that federal wages are intended to be set equal to that for "comparable" work in the private sector. We will present evidence both on the size of the public/private sector wage differential for federal workers as well as the magnitude of the differentials for state and local government workers. The final comparison we make between public and private sector labor markets is the differential impact of sex and race on wages. Affirmative action legislation might be expected to have a more significant impact in the public sector. A simple test of this can be carried out by examining the coefficients on sex and race dummy variables in both public and private sector wage equations.

We believe that this study has two basic features which make it better suited than much of the existing literature for making the above comparisons.[1] Nearly all of the recent studies of private sector union wage differentials have been carried out on cross-section or panel data sets of individual workers. In contrast, most of the work on public sector union wage differentials has focused on specific categories of government workers, i.e. public school teachers, and is based on comparisons of union and nonunion contracts. While each methodology has its merits, comparisons between private and public sector union differentials should be carried out with a consistent method of estimation. In this paper, we estimate both types of union differentials using micro data on individual workers.

The second important feature of this study is its handling of potential selection bias due to the endogeneity of both the government and the union status of a worker. Most public sector wage studies using micro data on individual workers control for the government and union status of a worker by including a set of dummy variables for the level of government and a dummy variable for union coverage.[2] Two potential problems exist from this practice. First it assumes that the returns to individual characteristics such as education and experience are the same regardless of government and union status. In fact, these returns may differ significantly by the worker's status. Secondly, individuals choose which of these labor markets to participate in. This raises the possibility of significant selection bias in the coefficients of the wage equations. These problems are handled in this study by estimating four separate wage equations: private/nonunion, private/union, public/nonunion, and public/union. A multinomial logit selection model is estimated in an attempt to correct for any existing selection bias.

Our basic findings can be summarized as follows. Evidence of self-selection was found for the private/nonunion and the public/union labor markets. Workers selecting the private/nonunion sector show indications of having a comparative advantage in that sector. On average, their wages exceed what would be predicted based on their observed characteristics by 4.5

percent. In contrast, workers selecting the public/union sector show evidence of having a comparative disadvantage in that sector. Their wages fall short of what would be predicted based on their observed characteristics by 17.1 percent. While the magnitude of the latter selection effect is quite large, it is subject to large sampling error.

We report both "conditional" and "unconditional" estimates for each wage differential of interest. Both are based on the consistent wage equation coefficient estimates. The unconditional differentials do not include the selection terms when calculating expected wages. The conditional differentials do include average values for the selection terms in the expected wage estimates. Due to the specific nature of the self-selection which exists in two of the four labor markets, these two different types of wage differentials give sharply contrasting results.

The unconditional estimates indicate that public sector unions have increased wages significantly more than unions in the private sector. The aggregate public sector unconditional union wage differential is 31.18 percent. The private sector counterpart is only 18.56 percent. In contrast, the conditional estimates indicate a much lower public sector union wage differential as compared to the private sector. The public and private conditional union wage differential estimates are 3.68 percent and 14.18 percent respectively.

The unconditional and conditional public/private sector wage differentials provide a similar contrasting view. The aggregate unconditional public/private sector wage differential is 16.31 percent whereas the conditional estimate is only 3.80 percent. Despite the comparability process which sets federal wages, the public/private sector wage differentials are largest for this level of government. The unconditional aggregate federal differential is 28.92 percent. The corresponding state and local figures are 9.45 percent and 18.12 percent. The conditional federal differential is 18.84 percent while the conditional state and local differentials are insignificantly different from zero.

Finally, we find some evidence consistent with the view that affirmative action has had a stronger impact in the public sector. The coefficient estimate on the race dummy variable is insignificant in the public/nonunion sector and indicates that whites are paid on average 8.14 percent more in the public/union sector.[3] The coefficient estimates for the private sector indicate that whites are paid on average 13.36 percent more in the private/nonunion sector and 11.25 percent more in the private/union sector. As a consequence, the unconditional public/private sector wage differential is 28.48 percent for non-whites as compared to 13.44 percent for whites.

In the next section we provide a short summary of the empirical literature on both public sector union wage differentials as well as public/private sector wage differentials. In section three we outline the econometric methods used to analyze the data. The construction of the data as well as the empirical findings are presented in the fourth section. A short conclusion summarizes our findings and briefly mentions the directions of our future work on this topic.

II. REVIEW OF THE EMPIRICAL LITERATURE

While unions have a long history in the private sector, they only recently made their debut in the public sector. In 1962, President Kennedy signed executive order 1098 which legitimized collective bargaining in the federal sector for the first time. Federal workers were given the right to join unions and to bargain over working conditions but not wages. A major exception to the constraint against wage bargaining was provided to the Postal workers and employees of federal government authorities i.e. Tennessee Valley Authority.[4] In 1959, Wisconsin became the first state to allow collective bargaining for its public sector workers. By the late 1970's most of the industrialized states had also adopted such statutes.[5] This change in legislative climate toward public sector unions resulted in a rapid acceleration in their growth. From 1964 to 1978, the fraction of federal workers unionized increased from 38.2 percent to 50.2 percent. Over the same time period, the fraction of state and local government (SLG) workers unionized increased from 7.7 percent to 17.4 percent.[6]

Unions representing federal workers typically can not use the bargaining process to generate higher wages for their members. However, they can lobby for wage increases through the political process. In 1962, the Federal Salary Reform Act was signed which established the "comparability" doctrine for workers covered under the General Service (GS) pay system.[7] The wage assigned to a particular job is supposed to reflect the pay rate for comparable work in the private sector. This was intended to provide equitible compensation for federal workers and to allow the federal government to compete with private companies for employees.[8] Each year the Bureau of Labor Statistics (BLS) makes recommendations for pay raises which will maintain this comparability. In nine out of the last thirteen years these suggested pay raises have been altered by through the political process prior to adoption.[9]

The bargaining environment varies across states and localities depending on the specific legislation governing public sector unions. State laws vary with respect to whether unions can bargain over wages, if strikes are allowed, if arbitration is used to settle impasses, what form of arbitration is used, whether union shops are allowed, etc. Most studies of union wage effects at the state and local level have concentrated on controling for this variation in the bargaining environment.[10]

As was mentioned earlier, very few micro studies have been carried out which analyze public sector union wage effects. Baugh and Stone (1982) use Current Population Survey (CPS) data to examine union effects for public school teachers. They find that union membership is associated with increases in annual earnings in the range of 0 to 7 percent. The effect of union membership or membership in an employee association is 12-22 percent. Moore and Raisian (1981) examine noneducation public employees from the CPS and find a union effect of 0 to 18 percent for hourly wages.

Ehrenberg and Schwarz (1983) survey aggregate studies, contract studies, and micro studies of public sector union wage effects. They conclude the following.

> "What is most striking is how small these numbers are! The estimated relative wage differentials associated with union membership or collective bargaining coverage are typically smaller than 10 percent and rarely exceed 20 percent. These estimates are consistently lower than the estimates obtained from private sector studies and they suggest that the relative wage effects of unions have been less in the public sector than in the private sector." (p.10)

Much of the work on public/private sector wage differentials has been carried out by Sharon Smith. We will focus on her 1981 study since it uses data very similar to our own. Smith analyzes individuals from the May 1978 CPS who live in one of 39 selected SMSA's. She focused on these SMSA's since this allowed her to include a metropolitan cost-of-living index calculated by the BLS. Separate equations were estimated using OLS for males and females. Dummy variables were included for union status and for federal, state, and local government status. She reports both nominal and real wage differentials and finds little difference between them.

While the federal comparability legislation covers jobs and not individuals, one might speculate that this legislation would result in smaller federal public/private wage differentials as compared to the differentials at the state and local levels. Smith finds the opposite to be true. The overall real wage differentials by level of government are given below.[11]

Level of Government	Males	Females
Federal	10.5%	22.1%
State	-5.2%	4.7%
Local	9.7%	0.1%

Looking at Smith's estimates for females, we find that the public/private sector wage differential declines as we move from the federal to the state and local levels of government. These figures also show that the federal differential for females is double the male differential. In addition, Smith finds large regional variations in these estimated differentials. Federal differentials for females range from a high of 35 percent in the northeast to a low of 16 percent in the west. In contrast, local female differentials are highest in the west (8.1 percent) and lowest in the northeast (0.8 percent). Finally, since Smith did not interact her union status dummy variable with her government dummy variables, it is impossible to compare public and private sector union wage differentials.

III. EMPIRICAL SPECIFICATION

In this section we provide the details of the two-stage estimation procedure used in our analysis. We start with the assumption that there are four distinct labor markets delineated by their government and union status: private/nonunion, private/union, public/nonunion, and public/union. The "potential" wage for the ith worker in the population in each of these labor markets is given by

$$
(1)\quad
\begin{aligned}
\ln w_{i1} &= X_{i1}\beta_1 + u_{i1} && \text{private/nonunion sector}\\
\ln w_{i2} &= X_{i2}\beta_2 + u_{i2} && \text{private/union sector}\\
\ln w_{i3} &= X_{i3}\beta_3 + u_{i3} && \text{public/nonunion sector}\\
\ln w_{i4} &= X_{i4}\beta_4 + u_{i4} && \text{public/union sector}
\end{aligned}
$$

where

$$
u_{ij} \sim N(0, \sigma_j^2) \qquad \begin{aligned} i &= 1,\dots,N \\ j &= 1,2,3,4 \end{aligned}
$$

The data vectors are indexed by the type of labor market since the private sector equations will contain industry dummy variables while the public sector equations will contain government classification dummy variables. This formulation allows for the potential returns for individual characteristics to vary across each labor market. This is the same as including complete sets of interaction terms between union and government status and the remaining variables in a single wage equation.

Individuals are assumed to select which labor market to participate in by choosing the market which maximizes lifetime expected utility. The ith individual's expected lifetime utility from participating in a particular market is modeled by the following function.

$$(2) \quad I^*_{ij} = Z_i\gamma_j + \eta_{ij} \qquad \begin{array}{l} i = 1,\ldots,N \\ j = 1,2,3,4 \end{array}$$

A few comments are in order concerning the structure of these functions. In the bivariate probit selection models used to estimate union wage differentials it has become customary to make the market choice decision depend in part on the contemporaneous union wage differential.[12] This leads to a reduced form probit which is used to correct for selectivity as well as a structural probit which includes the estimated union wage differential for each individual as a regressor.

We have not followed this tradition for two reasons. First, potential long term participation decisions may not be significantly influenced by contemporaneous wage comparisons. Second, this formulation requires that any variable thought to influence either wage should be included in the market choice decision. This leads to cases of possible spurious correlations entering into the analysis. For example, the percentage of workers unionized in an industry has been found to affect union and nonunion wage rates in that industry. Consequently, this variable is included in the reduced form probit by way of the wage differential term. However, as this variable approaches either zero or one hundred percent, you can predict the individual's market choice decision with increasing accuracy. This result is forced on the problem by the definition of the variable and would hold true even if the wage differential had no theoretical influence on the participation decision.

The individual's choice of markets is based on comparisons of each of these functions. The worker selects the market which yields the highest level of discounted utility.[13] This optimization process is captured by the following labor market indicator function.

$$(3) \quad I_i = j \text{ IFF } I^*_{ij} \geq \underset{k \neq j}{\text{Max}}\, I^*_{ik}$$

Following the formulation in Lee (1982), define the following residual for each individual and labor market.

$$(4) \quad \epsilon_{ij} = \underset{k \neq j}{\text{Max}}\, I^*_{ik} - \eta_{ij} \qquad \begin{array}{l} i = 1,\ldots,n \\ j = 1,2,3,4 \end{array}$$

Substituting for I^*_i from (2) and for ϵ_{ij} from (4) into (3) and rearranging we obtain a reformulation of the labor market indicator function.

$$(5) \quad I_i = j \text{ IFF } \epsilon_{ij} < Z_i \gamma_j \qquad \begin{array}{l} i = 1,\ldots,N \\ j = 1,2,3,4 \end{array}$$

The resulting likelihood function for this selection model depends on the specific distribution of the residuals n_{ij}'s. A straightforward estimation problem exists if we assume that the n_{ij}'s are independently and identically distributed with the type I extreme value distribution with cumulative distribution functions given by

(6) $F(\eta_{ij} < x) = \exp[-\exp(x)]$

As demonstrated in Domenich and McFadden (1975), the implied cumulative distribution functions for the ϵ_{ij}'s are given by

(7) $$\Pr(\epsilon_{ij}<Z_i\gamma_j) \equiv \Pr(I_i=j) = \frac{\exp(Z_i\gamma_j)}{\sum_{k=1}^{4}\exp(Z_i\gamma_j)}$$.

The selection problem, then, is analyzed with a multinomial logit model.[14] In our application, the probability that individual i chooses labor market j depends only on the characteristics of individual i. No market specific characteristics have been included in the selection analysis at this point in time. Estimation requires that we normalize the coefficients in one market to equal zero. With no loss of generality we assume that $\gamma_4 = 0$.

Given the estimates from the multinomial logit, we can proceed to estimate the wage equation in each market using the generalized two-stage procedure discussed in Lee (1982). While it is possible to jointly estimate all four wage equations, we chose instead to estimate each separately using only observations on individuals who selected that market. Consistent estimation, then, requires that we calculate the conditional expectation for each wage equation.

(8) $$E(\ln w_{ij}| I_i=j) = X_{ij}\beta_j + E(u_{ij}| I_i=j)$$

$$= X_{ij}\beta_j + E(u_{ij}| \epsilon_{ij}<Z_{ij}\gamma_j)$$

The case where u_{ij} and ϵ_{ij} follow a bivariate normal distribution leads to the standard selection bias correction. A similar correction can be derived in our problem by transforming the ϵ_{ij}'s into a standard normal random variable. Let $F_j(x) \equiv \Pr(\epsilon_{ij}<x)$. Then define the following transformed residual.

(9) $\epsilon^*_{ij} \equiv \Phi^{-1}[F_j(x)]$

where

Φ = cumulative distribution function for a standard normal variable variable.

By construction, $\epsilon^*_{ij} \sim N(0,1)$. In addition, we have the following relationship.

(10) $\epsilon_{ij} < Z_i\gamma_j$ IFF $\epsilon^*_{ij} < \Phi^{-1}[F_j(Z_i\gamma_j)]$

We can substitute this inequality into the conditional expectation term in equation (8). This gives us the following regression equation.

(11) $E(\ln w_{ij} | I_i=j) = X_{ij}\beta_j + E\{u_{ij} | \epsilon^*_{ij}<\Phi^{-1}[F_j(Z_i\gamma_j)]\}$

This conditional expectation can be evaluated using standard methods.

(12) $E(\ln w_{ij} | I_i=j) = X_{ij}\beta_j - \sigma_j\rho_j \dfrac{\phi(\Phi^{-1}[F_j(Z_i\gamma_j)])}{F_j(Z_i\gamma_j)}$

where

ρ_j = correlation coefficient between u_{ij} and ϵ^*_{ij}

Consistent estimates for the β's can be obtained by replacing γ_j with our first stage estimates $\hat{\gamma}_j$. This substitution implies that the standard errors for the β's reported by OLS will be biased since they do not account for the sampling variability in the $\hat{\gamma}_j$'s. The "corrected" variance/covariance matrices can be derived following the methods discussed in Lee (1980).[15]

IV. DATA CONSIDERATIONS AND EMPIRICAL FINDINGS

In this section we discuss the criteria used to generate our sample. We also present the multinomial logit estimates, the wage equation estimates, and the various wage differential estimates. The observations used in our analysis were taken from the 1977 May CPS. Individuals were included in the sample if they were in the nonagricultural sector, reported both a usual weekly earnings and a usual weekly hours, were full-time labor force participants, and lived in an SMSA. For individuals living in SMSA's for which the BLS does not report cost-of-living budget data, we chose the budget data for the nearest similar SMSA. Table A1 in the appendix lists the SMSA's included in the study as well as the corresponding budget data used for that SMSA.

Finally, the CPS lists four "industry" classifications for public sector employees: postal, other federal, state, and local. Using detailed occupation data we were able to further disaggregate the local employees into three categories: teachers, police & fire, and other local. However, these classifications only account for 57.4 percent of the total number of government workers. The remaining nonclassified government workers were treated as the left out group in public sector wage equations.[16]

Turn now to the multinomial logit results from the labor market selection process. We chose variables to include in the Z_i vector which we felt would be important in a long term decision over which labor market to participate in. Based on this criteria, the current age of the individual was excluded from the specification.[17] We conjectured that an individual's level of education may be a key determinant of his/her market selection.[18] For this reason, we entered education as a disjointed spline function with break points at high school and college graduation. This allows for more general education effects than would be obtained simply by entering the level of education in a linear fashion. Region dummy variables were also included in the specification. These help to control for such diverse factors such as regional variation in state laws affecting unions, etc.

The breakdown of the sample by labor market is given in Table A2 in the Appendix. Table A3 contains the means and standard deviations for the logistic variables. The logistic coefficient estimates are presented in Table 1. For ease of interpretation, the implied marginal effect of each variable on the probability of selecting each market is given in Table 2.[19]

Examining these marginal effects we find many interesting patterns. The effect of being white is to raise the probability of selecting the private/nonunion sector by over 16 percent. Males have a 21 percent higher probability than females of being in the private/union sector. Veterans are more likely to work in the public sector with the reduction coming entirely from the nonunion side of the private sector. Living in the South or the West reduces the likelihood of being in the private/union sector. These individuals tend to work in either nonunion sector. This pattern is most prominant in the South which may reflect the prevalence of right-to-work laws in these states.[20]

We find that increased levels of education do not reduce the probability of being in a union sector per se. Instead, large investments in education discourage only selecting the private/union sector. For example, graduating from high school reduces the probability of being in this sector by over 8.5 percent. In contrast, while graduating from college reduces the chance of selecting the private/union sector by 3.2 percent it raises the chance of selecting the public/union sector by 3.5 percent. Finally, finishing high school as well as continuing schooling for the first year beyond high school has a positive impact on the incidence of private/nonunion sector employment. However, from the third year beyond high school, continued education diminishes the probability of an individual selecting this sector.

The private and public sector wage equation estimates are given in Table 3 and Table 4. Both OLS and the consistent two-stage estimates are presented. The standard errors for the two-stage estimates have been corrected to account for both the sampling error in the logistic estimates as well as the truncation of the distribution of the residual by the selection process. Rather than deflating the nominal wage by the cost-of-living index (CLI), we include the log CLI as a separate regressor.

Several things are worth noting about these estimates. The variances of the distributions of the residuals are smaller for the union as compared to the nonunion sectors. In addition, the R-square's are smaller in the union sectors. The returns to individual characteristics such as education and experience are also lower in the union sectors.[21] These findings are consistent with the view that unions attempt to standardize wage structures.[22] Evidence is also found suggesting that private sector union contracts also tend to standardize wages across regions. A one percent increase in the metropolitan cost-of-living results in only a one third of a percent increase in the nominal private/union sector wage. In contrast, the coefficients on the log CLI in the other three sectors are all insignificantly different from one.

The data also indicates that workers who select themselves into the private/nonunion sector have a comparative advantage in that sector. Workers who select themselves into the public/union sector, on the other hand, have a comparative disadvantage in that sector. That is, these workers tend to earn less than a worker with similar observable characteristics who is selected at random from the population. Neither of these two selection effects is measured with great precision.

One final point of interest concerning the wage equation estimates is the comparison of the race coefficients across sectors. Smith (1977, pp. 108, 115) argued that affirmative action may be expected to have its strongest impact in the public sector. The argument is that public sector jobs are highly visible and that public sector agencies are charged with enforcing affirmative action in the private sector. Borjas (1982) found support for this hypothesis at the federal level of government. Using interagency data he found that the extent of racial wage discrimination declined with the size of the minority constituency of an agency as well as with the fraction of its budget devoted to affirmative action compliance in the private sector. Further evidence on this hypothesis is provided by our wage equation estimates. In the public/nonunion sector we find no significant difference in wages based on race. In the public/union sector whites on average earn 8.1

percent more than nonwhites. While this latter difference is statistically significant, it is still much lower in magnitude than the race coefficients in the private sector.[23]

The various wage differentials calculated for this study are presented in Tables 5 through 8. In each table, three types of differentials are reported. The first uses the OLS wage equation estimates and are given for comparison purpose. Two distinct differentials are calculated using the two-stage wage equation estimates. Conditional differentials are listed as specification (2) in each table. These include the selection effects when calculating the expected wage. Unconditional differentials are listed as specification (3) in each table. These exclude the selection effects when calculating the expected wage.[24]

The conditional and unconditional differentials have very different interpretations. The aggregate public sector union differentials will illustrate this distinction. The unconditioned differential corresponds to the following experiment. Take at random from the population an individual with observable characteristics which are the same as the average public sector worker. Since we do not observe this individual's choice of sectors, the union differential reflects only the varying returns for his/her observable characteristics in the two public sectors. The conditional differential corresponds to very different experiment. In this case, we select at random a worker from the public/nonunion sector and a second worker from the public/union sector each with the same observeable characteristics as the average public sector worker. Since we know each individual's choice of sectors, the differential reflects not only the varying returns on their observable characteristics but also the varying returns from their unobservable characteristics which can be inferred from their market selection.

Comparisons of the success of public and private sector unions at generating wage differentials for their members are given in Table 5. The public sector union differentials are calculated using the mean characteristics of public sector workers for the specific category being considered. The private sector union differentials were calculated in a similar fashion.

Two patterns evident in this table will carry over to each of the other sets of wage differentials. The first is that the OLS and the conditional two-stage differentials have virtually the same point estimates and standard errors. This is the direct consequence of the construction of the conditional differentials. Estimating an individual's expected wage using OLS implicitly includes any selection effect which may be present. The estimated expected wages used to calculate the conditional differentials explicitly include these same selection effects. Consequently, it is not surprising that the two methods produce similar results.

The second common feature in these tables is that the unconditional differentials are much greater in magnitude than the conditional differentials. This is a consequence of the specific types of selection effects found in this study and need not occur in general. To see this, consider the public sector union differentials for a moment. Recall that the public/nonunion workers have no significant comparative advantage or disadvantage in their sector while the public/union workers have a comparative disadvantage. In calculating the conditional union differential, the expected union wage is reduced by the average comparative disadvantage of workers in this sector. This acts to lower the resulting union wage differential. In contrast, in calculating the unconditional union differential, the expected union wage is not reduced by the average comparative disadvantage of union workers. Consequently, the resulting union wage differential is much larger. The same pattern occurs in the private sector union differentials because the private/nonunion workers have a comparative advantage in their sector while the private/union workers have no selection effect. The same reasoning applies to the public/private sector wage differentials.

Turn now to the specific results in Table 5. Had we estimated the wage equations using only OLS, then we would have concluded as Ehrenberg and Schwarz do that public sector unions have been unable to generate sizeable wage differentials. As it turns out, the point estimates for the unconditional union differentials are much higher in the public as compared to the private sector. This is an important finding if one believes that the unconditional differential is the conceptually "correct" one to focus on.[25] At issue is what type of experiment the wage differentials are meant to represent. To reiterate our earlier point, the unconditional differentials correspond to the experiment involving drawing at random from the population a worker of a given set of observed characteristics and moving him/her between labor market sectors. In this respect, the unconditional differentials reflect purely the varying returns to a worker's observed characteristics across these sectors.

The figures in Table 5 also point out several contrasting features of public and private sector unions. The regional breakdowns show that the public sector union effects have been the strongest in the northeast and the north central regions while the private sector union effects have been strongest in the south and west.[26] Private sector unions have increased wages by a greater amount for nonwhites as compared to whites and for males as compared to females. Exactly the opposite pattern is found in the case of public sector unions.

In Table 6 we disaggregate the public sector union differentials by level of government and region. Looking first at the aggregate numbers, we find that unconditional differentials are roughly the same across federal, state, and local levels of government. The conditional union differential for local public workers is slightly higher than the corresponding federal or state differential.[27] It is interesting to note that despite the exemption which allows postal worker unions to bargain over wages, the union differential for postal workers is not significantly higher than the differential for other federal workers. Looking at the local level, the unconditional union differentials are roughly the same for teachers, police and firemen, and other local workers. In contrast, police and firemen have lower conditional differentials than either teachers or other local workers.[28] The regional figures for each level of government follow the same basic pattern as noted in Table 5.

Estimates for the public/private sector wage differentials are given in Table 7. These differentials were calculated using the following formulation.

$$(13)\ \Delta W \equiv \frac{\left(\frac{k_3}{k_3+k_4}\right)\left[e^{\bar{X}_3\hat{\beta}_3+\frac{1}{2}\hat{\sigma}_3^2}\right] + \left(\frac{k_4}{k_3+k_4}\right)\left[e^{\bar{X}_4\hat{\beta}_4+\frac{1}{2}\hat{\sigma}_4^2}\right]}{\left(\frac{k_1}{k_1+k_2}\right)\left[e^{\bar{X}\hat{\beta}_1+\frac{1}{2}\hat{\sigma}_1^2}\right] + \left(\frac{k_2}{k_1+k_2}\right)\left[e^{\bar{X}_2\hat{\beta}_2+\frac{1}{2}\hat{\sigma}_2^2}\right]} - 1$$

where

$k_i \equiv$ fraction of workers in the ith sector.

In words, to calculate the expected public (private) sector wage for a hypothetical worker with a given set of individual characteristics we assume that his/her probability of being in the nonunion or union sector is equal to the observed frequency of workers across those two sectors. The expected wage, then, is the weighted average of the two conditional wages in each public (private) sector.[29] The mean individual characteristics used in these calculations were for public sector workers. This was to facilitate comparisons between Table 7 and Table 8. The sample of private sector workers were used only to generate employment shares for the private sector and industry frequency distributions.

The overall unconditional public/private sector differential is estimated to be 16.31 percent. The corresponding conditional estimate is 3.80 percent. The differential is highest in the northeast region and lowest in the north central region. The unconditional differential for nonwhites is more than double the estimate for whites. In addition, the unconditional differential for females is nearly twice the size of the male differential. While Smith (1981) does not report any comparable figures which pool all government workers together, her estimated differentials for male and female federal workers are nearly identical to our overall estimates by sex.

Table 8 contains the breakdown of the public/private sector differentials by level of government. This allows us to check to see if the comparability legislation governing federal wages has resulted in lower differentials than for state or local workers. In fact, both the conditional and unconditional differentials are highest for federal workers. The estimated federal unconditional differential is 28.92 percent as compared to state and local estimates of 9.45 percent and 18.12 percent respectively. This ordering is similar to Smith's findings for male workers. Unlike the case of the union differentials, the public/private differential for postal workers is more than twice that of other federal workers. Looking across the classifications of local workers we see that police and firemen are paid the highest relative to what they might earn in the private sector.[30] The federal differential is highest in the northeast and west. The same regional pattern holds for state and local differentials.

V. CONCLUSION

In this paper we examined both public and private sector union wage differentials as well as public/private sector wage differentials using data drawn from the May 1977 CPS. A general selection model was estimated which indicated that wages for private/nonunion and public/union workers were affected by the selection process. Two conceptually distinct sets of differentials calculated using the consistent wage coefficient estimates were presented and contrasted. The general findings in the literature were consistent with our conditional differential estimates but not with our unconditional estimates. In particular, the unconditional union differentials were much higher in the public as compared to the private sector. Nonwhites were found to have the largest public/private differential of all major categories of workers considered. When we disaggregated by level of government we found that federal workers enjoyed the largest advantage relative to what they would earn in the private sector.

The results of this study suggest several possible directions for future work in this topic area. The first concerns the list of variables included both in the logit model and in the wage equations. These consisted exclusively of individual characteristics, region dummy variables, and industry and level of government classifications. No state or SMSA specific variables apart from cost-of-living were explored. This is in contrast to the approach taken in the studies based on contract level data. These studies concentrate on state laws governing public unions, the form of local government, etc. It would be interesting to integrate this type of data into our analysis. A second possible topic is to determine whether or not the magnitudes of state and local public/private differentials can explain the adoption in some state of restrictive tax and/or budgetary measures. In addition, we could investigate to what extent these types of legislative measures result in lowering subsequent differentials in that state. These issues can be addressed by estimating this model for years prior to and following the passage of these state laws.

Footnotes

1. See Ehrenberg and Schwarz (1983) for a survey of the public sector labor market literature.

2. See for example the work by Smith (1981). Two noteable exceptions are Venti (1985) who allows for endogeneous choice of government status and Robinson and Thomes who allow for endogeneous choice of union status. See Ehrenberg and Schwarz (1983) for a discussion of selection bias in the contract based studies.

3. We used the natural log of the wage rate as our dependent variable. If β is the coefficient on the race dummy variable, then the implied percentage wage difference can be calculated by $100(e^{\beta}-1)$. See Halvorsen and Palmquist (1980).

4. The rationale given for this exception is that wage increases for these categories of federal workers can be paid for out of charges for their services. This is in contrast to other federal workers whose wages are paid for out of general tax revenue.

5. For a detailed description of these state statues see Freeman and Valletta (1985).

6. See Ehrenberg and Schwarz (1983) for more discussion of these points.

7. Wages set under the Federal Wage System are also intended to reflect private sector rates of compensation. In principal, the comparability doctrine is intended to also apply to Postal workers who can bargain over their wages.

8. See the <u>President's Panel of Federal Compensation</u> (1976).

9. See Venti (1985, pp. 6-7) and Smith (1977) for more discussion of the wage setting process.

10. This is in sharp contrast to studies of union wage effects for private sector employees. These studies concentrate instead on variation in individual and industry characteristics rather than variations in legislative environments. An exception is the work on the effect of right-to-work laws. See Farber (1984).

11. These estimates are derived from the reported coefficients on the government dummy variables for the specifications using the log real wage as the dependent variable. These are reported in Table 1, pp. 91-92. The percent differential is calculated as $100(e^{\beta}-1)$.

12. This tradition was started by the work of Lee (1978).

13. More generally, we observe an individual working in a particular sector only if he/she both desires to be in that sector and can find employment in that sector. Abowd and Farber (1982) augment the market choice functions with firm employment decision functions. They reject the hypothesis that only the market choice functions matter in determining a worker's union status. Given the complexity of our multi-choice model, we have not followed their approach.

14. A summary of this method as well as a comparison to alternative estimation approaches is given in Maddala (1983, pp. 275-278).

15. Details of these derivations are available upon request.

16. In contrast, Smith (1981) eliminates the nonclassified government workers from her sample.

17. Age should be included only if there is systematic movement of individuals between markets as they grow older. We tested for this life-cycle effect but found no evidence of it in the data.

18. An individual's level of education was defined to be the number of completed years of schooling.

19. The noneducation marginal effects are evaluated using the sample means for the noneducation variables and assuming that the individual graduated from high school. The sample mean number of years of education was actually slightly greater than twelve. The education marginal effects are evaluated using the sample means of the noneducation variables.

20. See Farber (1985) for a detailed analysis of the effects of these laws.

21. The only exception is the experience profile for the public/union sector which is slightly steeper than for the public-nonunion sector.

22. See Freeman (1982) for a discussion of this issue.

23. An alternative explanation is that the public sector attracts a more homogeneous group of workers as a result of its standardized pay structure. This implies that the race coefficient is less likely to be picking up unmeasured productivity differences. Similar arguments would apply to unions which also use standardized wage structures. However, notice that the race coefficient in the private/union sector is both large and significant. In addition, within the public sector, the race coefficient is larger in union than nonunion jobs.

24. Standard errors for the estimated differentials are also reported. The standard errors for the two-stage differentials include not only the variances of the coefficients in each market but also the covariance between coefficients in different markets which is induced by the first-stage logit estimation.

25. It should be stressed, though, that the estimates for the public sector union differentials have very large standard errors associated with them. One implication of this is that replicating our study on different years of the CPS data might lead to cases with much lower estimates for these differentials.

26. Recall that we found evidence of standardization of private sector union wages across regions. This will tend to raise the estimated differentials in low cost-of-living areas such as the south.

27. The figures for the state level of government should be interpreted with caution. The data included few state workers since most state capitals are not in the set of SMSA's included in the CPS data.

28. The OLS and conditional two-stage estimates for teachers are consistent with the low end of the range of estimates reported by Baugh and Stone (1982).

29. An alternative approach is to take the characteristics of the worker in question and use them in conjuction with our logit model to predict the probability that this worker would be in each sector. It should also be noted that we did not include sampling variation in the employment shares or the residual variance terms when calculating the standard errors for these wage differentials. Had we opted to use the logistic probabilities in lieu of the employment shares, it would have in principal been possible to include this additional source of sampling variation.

30. This sizeable differential may reflect in large part compensation for the inherent job hazards facing these workers. This illustrates the point that this procedure for estimating wage differentials does not identify what

portion consists of "rent extraction" versus compensation for varying job attributes such as health risk.

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Table 1: Logistic Coefficients

Variable	θ_1	θ_2	θ_3
Intercept	2.7070 (0.4186)	2.1020 (0.4296)	0.3089 (0.5524)
Married	-0.3995 (0.0852)	0.0695 (0.0993)	-0.2578 (0.1043)
White	1.1666 (0.0819)	0.5924 (0.0924)	0.5490 (0.1020)
Male	0.4689 (0.0824)	1.5337 (0.0942)	-0.1955 (0.1035)
Veteran	-0.8700 (0.0858)	-0.7802 (0.0935)	-0.2203 (0.1096)
Education(1)	-0.0988 (0.0431)	-0.1062 (0.0438)	-0.0700 (0.0577)
Education(2)	-1.0698 (0.4044)	-1.5445 (0.4104)	-0.3008 (0.5358)
Education(3)	-0.5288 (0.4122)	-1.3006 (0.4211)	0.1458 (0.5436)
Education(4)	-0.3816 (0.0282)	-0.7947 (0.0406)	-0.1029 (0.0334)
Northeast	-0.6136 (0.0857)	-0.8434 (0.0949)	-0.8402 (0.1185)
South	0.3673 (0.1031)	-0.7473 (0.1173)	1.0428 (0.1209)
West	-0.2396 (0.0911)	-0.6511 (0.1018)	0.0934 (0.1137)
log likelihood	-13532.8		
N	13907		

Note: Asymptotic standard errors are in parentheses.

Table 2: Logistic Marginal Effects

Variable	Private/ Nonunion	Private/ Union	Public/ Nonunion	Public/ Union
Married	-0.0889	0.0728	0.0002	0.0159
	(0.0105)	(0.0136)	(0.0050)	(0.0045)
White	0.1723	-0.0688	-0.0254	-0.0782
	(0.0130)	(0.0109)	(0.0067)	(0.0088)
Male	-0.0899	0.1999	-0.0670	-0.0429
	(0.0106)	(0.0188)	(0.0062)	(0.0056)
Veteran	-0.0861	-0.0131	0.0426	0.0566
	(0.0112)	(0.0096)	(0.0070)	(0.0071)
Junior	-0.0026	-0.0040	0.0010	0.0055
	(0.0042)	(0.0039)	(0.0018)	(0.0027)
Senior	0.0511	-0.0870	0.0314	0.0044
	(0.0153)	(0.0130)	(0.0065)	(0.0075)
1st Yr. College	0.0659	-0.0430	0.0005	-0.0234
	(0.0132)	(0.0119)	(0.0064)	(0.0050)
2nd Yr. College	0.0163	-0.0643	0.0268	0.0212
	(0.0056)	(0.0059)	(0.0012)	(0.0009)
3rd Yr. College	-0.0116	-0.0472	0.0308	0.0280
	(0.0038)	(0.0038)	(0.0020)	(0.0014)
4th Yr. College	-0.0363	-0.0328	0.0336	0.0355
	(0.0038)	(0.0022)	(0.0030)	(0.0025)
1st Yr. Grad.	-0.0560	-0.0218	0.0347	0.0431
	(0.0047)	(0.0013)	(0.0040)	(0.0040)
Northeast	-0.0453	0.0424	-0.0405	0.0434
	(0.0101)	(0.0103)	(0.0044)	(0.0041)
Central	-0.0738	0.1179	-0.0252	-0.0189
	(0.0100)	(0.0124)	(0.0044)	(0.0041)
South	0.1002	-0.1485	0.0762	-0.0279
	(0.0102)	(0.0089)	(0.0068)	(0.0042)
West	-0.0157	-0.0064	0.0145	0.0076
	(0.0102)	(0.0097)	(0.0050)	(0.0047)

Note: Asymptotic standard errors are in parentheses.

Table 3: Private Sector Wage Equation Estimates

Variable	Private/Nonunion OLS	Private/Nonunion Two-Stage	Private/Union OLS	Private/Union Two-Stage
Intercept	0.2726 (0.0448)	0.2126 (0.0585)	1.0273 (0.0814)	1.0236 (0.0840)
Education	0.0538 (0.0018)	0.0545 (0.0019)	0.0275 (0.0026)	0.0270 (0.0039)
Experience	0.0225 (0.0012)	0.0225 (0.0012)	0.0120 (0.0018)	0.0120 (0.0017)
Experience^2	-0.00038 (0.00002)	-0.00038 (0.00002)	-0.00017 (0.00003)	-0.00017 (0.00003)
Male	0.2284 (0.0102)	0.2219 (0.0110)	0.2257 (0.0162)	0.2286 (0.0243)
White	0.1063 (0.0136)	0.1254 (0.0182)	0.1075 (0.0157)	0.1066 (0.0165)
Married	0.1007 (0.0113)	0.0930 (0.0123)	0.0910 (0.0184)	0.0920 (0.0194)
Veteran	0.0942 (0.0117)	0.0846 (0.0132)	0.0550 (0.0136)	0.0550 (0.0135)
Northeast	-0.0938 (0.0140)	-0.0930 (0.0140)	-0.0865 (0.0169)	-0.0871 (0.0172)
South	-0.0093 (0.0125)	-0.0010 (0.0135)	-0.0313 (0.0191)	-0.0344 (0.0273)
West	-0.0346 (0.0119)	-0.0325 (0.0120)	0.0125 (0.0161)	0.0113 (0.0177)
Log CLI	0.9468 (0.0981)	0.9406 (0.0978)	0.3465 (0.1312)	0.3454 (0.1303)
Professional	0.2298 (0.0104)	0.2296 (0.0104)	0.1432 (0.0231)	0.1426 (0.0233)
Selection Parameter		-0.0711 (0.0447)		-0.0062 (0.0389)
σ^2		0.1403		0.0898
	R^2=0.5087 N = 8,232	R^2=0.5089	R^2=0.4263 N = 2,989	R^2=0.4263

Note: Asymptotic standard errors are in parentheses. Thirty-six industry dummy variables are included in each specification.

Table 4: Public Sector Wage Equation Estimates

Variable	Public/Nonunion OLS	Public/Nonunion Two-Stage	Public/Union OLS	Public/Union Two-Stage
Intercept	0.3697 (0.0774)	0.3019 (0.3807)	0.7425 (0.0685)	1.0390 (0.2235)
Education	0.0582 (0.0049)	0.0603 (0.0126)	0.0448 (0.0046)	0.0365 (0.0075)
Experience	0.0186 (0.0030)	0.0186 (0.0030)	0.0202 (0.0025)	0.0203 (0.0025)
Experience^2	-0.00025 (0.00007)	-0.00025 (0.00007)	-0.00028 (0.00005)	-0.00029 (0.00005)
Male	0.1925 (0.0269)	0.1851 (0.0487)	0.1897 (0.0226)	0.2123 (0.0280)
White	0.0110 (0.0277)	0.0076 (0.0335)	0.0352 (0.0210)	0.0782 (0.0375)
Married	0.0794 (0.0280)	0.0799 (0.0280)	0.0092 (0.0239)	-0.0051 (0.0261)
Veteran	0.0508 (0.0301)	0.0558 (0.0403)	-0.0211 (0.0235)	-0.0538 (0.0333)
Northeast	-0.0741 (0.0380)	-0.0776 (0.0425)	-0.0678 (0.0279)	-0.1010 (0.0367)
South	0.0615 (0.0276)	0.0704 (0.0560)	-0.0204 (0.0282)	-0.0072 (0.0298)
West	0.0726 (0.0299)	0.0759 (0.0348)	0.0441 (0.0243)	0.0326 (0.0257)
Log CLI	1.0298 (0.2263)	1.0283 (0.2250)	0.9313 (0.1896)	0.9368 (0.1881)
Professional	0.2444 (0.0272)	0.2451 (0.0274)	0.2025 (0.0284)	0.1968 (0.0285)
Postal	0.3052 (0.0626)	0.3047 (0.0622)	0.1984 (0.0310)	0.2026 (0.0308)
Other Federal	0.2086 (0.0265)	0.2086 (0.0263)	0.1410 (0.0387)	0.1444 (0.0385)
State	0.0662 (0.0453)	0.0659 (0.0450)	0.0100 (0.0486)	0.0136 (0.0482)

Table 4 (Continued)

Variable	Public/Nonunion		Public/Union	
	OLS	Two-Stage	OLS	Two-Stage
Teachers	-0.1096	-0.1095	-0.0223	-0.0260
	(0.0415)	(0.0412)	(0.0271)	(0.0271)
Police & Fire	0.2246	0.2240	0.1953	0.2031
	(0.0520)	(0.0518)	(0.0344)	(0.0345)
Other Local	0.0024	0.0020	-0.0046	0.0000
	(0.0368)	(0.0367)	(0.0350)	(0.0348)
Selection Parameter		-0.0247		-0.1111
		(0.1356)		(0.0798)
σ^2		0.1482		0.0917
	R^2=0.4312	R^2=0.4312	R^2=0.3894	R^2=0.3904
	N = 1,464		N = 1,222	

Note: Asymptotic standard errors are in parentheses.

Table 5: Public and Private Sector Union Wage Differentials

Category	Public Sector OLS (1)	Public Sector Two-Stage (2)	Public Sector Two-Stage (3)	Private Sector OLS (1)	Private Sector Two-Stage (2)	Private Sector Two-Stage (3)
Aggregate	3.07	3.68	31.18	14.23	14.18	18.56
	(2.04)	(2.18)	(35.69)	(1.31)	(1.69)	(5.82)
Region:						
Northeast	9.19	9.86	37.30	7.00	6.46	11.06
	(3.63)	(3.91)	(41.35)	(1.76)	(2.53)	(5.57)
North Central	6.89	7.62	37.35	15.04	14.98	19.63
	(3.27)	(3.56)	(39.23)	(1.72)	(2.63)	(5.36)
South	-3.68	-3.18	23.90	19.23	20.10	23.74
	(2.90)	(3.30)	(30.88)	(2.33)	(3.23)	(8.03)
West	2.75	3.29	29.58	18.53	18.33	22.66
	(3.06)	(3.47)	(34.10)	(2.03)	(2.91)	(6.69)
Race:						
White	3.03	3.65	31.94	14.13	14.06	18.18
	(2.13)	(2.29)	(36.57)	(1.36)	(1.77)	(5.88)
Nonwhite	3.01	3.56	27.85	16.50	16.54	22.95
	(3.45)	(4.00)	(32.41)	(2.37)	(3.53)	(6.75)
Sex:						
Male	1.31	2.00	29.79	13.04	12.93	17.89
	(2.35)	(2.53)	(37.25)	(1.38)	(1.91)	(5.15)
Female	4.43	4.89	31.82	10.64	10.67	13.84
	(2.54)	(2.86)	(33.60)	(1.78)	(2.37)	(7.17)

Note: Asymptotic standard errors are in parentheses.

Table 6: Public Sector Union Wage Differentials
By Level of Government

Government	Aggregate			Northeast			North Central			South			West		
	OLS	Two-Stage		OLS	Two-Stage		OLS	Two-Stage		OLS	Two-Stage		OLS	Two-Stage	
	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)
Federal	0.84	1.37	29.11	6.87	7.31	35.17	5.37	5.93	36.31	-5.05	-4.45	22.03	2.25	2.54	29.69
	(3.33)	(3.44)	(35.27)	(4.64)	(4.84)	(42.22)	(4.38)	(4.58)	(40.59)	(3.75)	(4.07)	(30.56)	(4.24)	(4.52)	(35.55)
Postal	-2.45	-2.09	25.12	-0.13	0.03	26.40	0.12	0.30	29.99	-8.84	-7.86	18.24	-3.92	-3.68	22.45
	(6.51)	(6.54)	(38.26)	(7.15)	(7.26)	(41.52)	(7.01)	(7.14)	(40.87)	(6.50)	(6.74)	(32.60)	(6.84)	(6.98)	(35.71)
Other Federal	-3.17	-2.65	23.84	4.08	4.68	31.46	2.11	2.85	31.74	-7.00	-6.49	19.36	-1.03	-0.77	25.22
	(4.24)	(4.33)	(32.80)	(5.51)	(5.68)	(39.92)	(5.26)	(5.44)	(38.49)	(4.40)	(4.65)	(29.74)	(5.07)	(5.29)	(33.91)
State	2.05	1.69	30.07	4.69	4.12	32.22	4.06	3.85	33.76	-3.40	-3.15	25.59	0.60	-0.07	27.00
	(6.58)	(6.55)	(37.17)	(7.23)	(7.27)	(40.92)	(7.18)	(7.25)	(39.22)	(6.79)	(6.94)	(33.11)	(6.89)	(6.98)	(34.57)
Local	6.90	5.24	31.99	11.54	11.62	38.11	7.45	7.41	36.38	2.57	2.05	29.54	3.53	3.44	28.53
	(3.47)	(4.22)	(32.04)	(4.39)	(4.65)	(40.01)	(3.89)	(4.13)	(37.96)	(3.82)	(4.22)	(30.50)	(3.71)	(4.09)	(32.45)
Teachers	11.24	11.62	36.93	13.78	14.60	37.90	14.51	15.26	43.14	5.46	4.75	30.78	10.21	10.81	34.75
	(4.52)	(4.81)	(32.29)	(5.59)	(6.01)	(35.49)	(5.33)	(5.72)	(36.17)	(4.99)	(5.49)	(27.49)	(5.15)	(5.63)	(30.70)
Police & Fire	4.29	4.93	35.46	6.18	6.51	35.98	6.03	6.48	38.71	0.36	1.85	33.41	2.01	2.70	31.46
	(5.77)	(5.82)	(41.83)	(6.61)	(6.74)	(45.05)	(6.26)	(6.41)	(43.70)	(6.18)	(6.46)	(38.93)	(6.07)	(6.27)	(38.29)
Other Local	7.43	8.05	38.53	13.11	13.74	45.29	11.47	11.97	45.60	1.83	2.70	32.37	4.60	5.03	33.02
	(5.28)	(5.33)	(40.49)	(6.42)	(6.58)	(47.82)	(6.16)	(6.29)	(44.85)	(5.55)	(5.81)	(35.02)	(5.58)	(5.80)	(37.27)

Note: Asymptotic standard errors are in parentheses.

Table 7: Public/Private Sector Wage Differentials

Category	OLS	Two-Stage	
	(1)	(2)	(3)
Aggregate	3.48 (1.22)	3.80 (1.44)	16.31 (13.78)
Region:			
Northeast	5.25 (1.73)	5.55 (2.14)	22.70 (12.99)
North Central	-1.45 (1.72)	-1.09 (1.94)	11.02 (13.77)
South	6.19 (1.95)	6.74 (2.55)	14.92 (14.96)
West	6.20 (1.86)	6.31 (2.23)	18.94 (13.64)
Race:			
White	1.25 (1.24)	1.54 (1.48)	13.44 (13.89)
Nonwhite	13.45 (2.23)	13.94 (2.74)	28.48 (13.66)
Sex:			
Male	-0.51 (1.33)	-0.01 (1.55)	12.74 (13.72)
Female	10.06 (1.63)	10.10 (2.09)	22.48 (14.06)

Note: Asymptotic standard errors are in parentheses.

Table 8: Public/Private Sector Wage Differentials
By Level of Government

Government	Aggregate			Northeast			North Central			South			West		
	OLS	Two-Stage		OLS	Two-Stage		OLS	Two-Stage		OLS	Two-Stage		OLS	Two-Stage	
	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)	(1)	(2)	(3)
Federal	18.22	18.84	28.92	21.06	21.54	40.45	12.27	12.75	23.82	20.24	21.09	26.78	23.01	23.44	34.36
	(2.17)	(2.37)	(18.06)	(2.69)	(3.04)	(16.00)	(2.60)	(2.89)	(17.91)	(2.80)	(3.40)	(18.94)	(2.95)	(3.29)	(18.61)
Postal	22.47	23.19	47.88	25.12	25.45	53.72	18.79	19.19	42.11	21.63	23.52	42.51	27.37	28.07	54.05
	(3.24)	(3.40)	(17.17)	(3.52)	(3.92)	(17.84)	(3.59)	(3.77)	(17.40)	(4.38)	(4.72)	(16.73)	(3.81)	(4.14)	(17.75)
Other Federal	17.61	18.15	22.75	17.35	17.79	27.61	8.60	8.98	12.76	20.23	20.97	24.91	21.38	21.68	26.12
	(2.52)	(2.76)	(20.52)	(3.25)	(3.50)	(19.40)	(3.07)	(3.25)	(21.35)	(2.99)	(3.60)	(19.81)	(3.48)	(3.80)	(22.47)
State	0.96	1.34	9.45	2.76	3.20	18.65	-4.21	-3.88	-1.96	0.49	0.82	6.74	5.79	6.03	11.86
	(3.45)	(3.51)	(16.25)	(3.70)	(3.89)	(13.84)	(4.18)	(4.29)	(20.08)	(4.10)	(4.45)	(15.70)	(4.18)	(4.36)	(18.06)
Local	2.60	2.01	18.12	4.43	4.76	23.66	-2.69	-2.29	13.74	-1.23	-1.26	12.09	3.99	3.95	18.91
	(1.78)	(2.27)	(11.45)	(1.94)	(2.40)	(12.46)	(1.84)	(2.09)	(12.59)	(2.10)	(2.66)	(11.48)	(2.03)	(2.44)	(12.06)
Teachers	-2.80	-2.92	14.30	0.75	1.06	18.96	-6.42	-5.96	11.50	-5.36	-6.28	9.08	1.04	0.69	18.12
	(1.79)	(2.24)	(10.89)	(2.20)	(2.85)	(11.02)	(2.10)	(2.46)	(11.85)	(2.41)	(3.13)	(9.92)	(2.31)	(2.93)	(11.00)
Police & Fire	13.56	14.37	33.46	16.41	16.83	42.49	8.67	9.20	26.18	15.24	17.19	32.83	17.47	18.31	33.79
	(3.00)	(3.16)	(16.28)	(3.37)	(3.72)	(16.94)	(3.21)	(3.40)	(16.29)	(3.85)	(4.21)	(17.24)	(3.67)	(3.96)	(16.64)
Other Local	-2.73	-2.26	7.85	0.04	0.47	14.16	-7.53	-7.04	3.18	-2.14	-1.42	6.38	0.17	0.47	9.72
	(2.54)	(2.61)	(14.95)	(2.92)	(3.11)	(15.09)	(2.73)	(2.84)	(14.72)	(3.06)	(3.39)	(14.60)	(2.96)	(3.16)	(15.29)

Note: Asymptotic standard errors are in parentheses.

Table A1: List of Cities Included in the Sample by Region

Region	City	Budget Data Used
Northeast	New York	*
	Philadelphia	*
	Boston	*
	Pittsburgh	*
	Newark	New York
	Buffalo	*
	Rochester	Buffalo
	Nassau-Suffolk	New York
	Patterson-Clifton-Passaic	New York
	Albany-Troy	Buffalo
North Central	Chicago	*
	Detroit	*
	St. Louis	*
	Cleveland	*
	Milwaukee	*
	Cincinnati	*
	Kansas City	*
	Columbus	Dayton
	Akron	Cleveland
	Minneapolis-St. Paul	*
	Indianapolis	*
	Gary	Chicago
South	Baltimore	*
	Houston	*
	Atlanta	*
	Miami	Orlando
	New Orleans	Baton Rouge
	Birmingham	Nashville
	Washington, D.C.	*
	Dallas	*
	Tampa-St. Petersburg	Orlando
	Ft. Worth	Dallas
	Norfolk	Washington, D.C.
	Greensboro-Winston-Highpoint	Durham
West	Los Angeles-Long Beach	*
	San Francisco-Oakland	*
	Seattle-Everett	*
	San Diego	*
	Denver	*
	San Jose	San Francisco-Oakland
	Portland	Seattle-Everett
	Sacramento	Bakersfield
	San Bernardino-Riverside-Ontario	Los Angeles-Long Beach
	Anaheim-Santa Ana-Garden Grove	Los Angeles-Long Beach

Table A2: Breakdown of Sample by Labor Market

	Nonunion	Union	
Private	8,232	2,989	11,221
Public:	1,464	1,222	2,686
Federal:	408	187	595
Postal	42	122	164
Other Federal	366	65	431
State:	84	39	123
Local:	321	503	824
Teachers	117	321	438
Police & Fire	70	102	172
Other Local	134	80	214
Non Classified	651	493	1,114
	9,696	4,211	13,907

Table A3: Means and Standard Deviations of Variables Used in the Logit Analysis

Variable	Mean	Standard Deviation
Married	0.7908	0.4068
White	0.8638	0.3430
Male	0.6174	0.4860
Veteran	0.2842	0.4510
Education	12.7131	2.9523
Northeast	0.2583	0.4377
South	0.2242	0.4171
West	0.2466	0.4310

Table A4: Means and Standard Deviations of the Wage Equation Varibles

Variable	Private/ Nonunion	Private/ Union	Public/ Nonunion	Public/ Union
Log Wage	1.6221	1.8136	1.7848	1.8865
	(0.5290)	(0.3957)	(0.5097)	(0.3662)
Education	12.8124	11.2603	14.0936	13.9435
	(2.8521)	(2.6311)	(2.8284)	(2.9784)
Experience	17.1199	22.0545	17.7356	20.8633
	(13.3788)	(13.8700)	(12.8620)	(12.8946)
$Experience^2$	472.0609	678.7136	479.8723	601.4149
	(608.6446)	(685.6099)	(569.5679)	(624.5284)
Male	0.5776	0.7892	0.5157	0.5876
	(0.4940)	(0.4079)	(0.4999)	(0.4925)
White	0.8938	0.8351	0.8265	0.7766
	(0.3081)	(0.3712)	(0.3788)	(0.4167)
Married	0.7609	0.8535	0.7930	0.8355
	(0.4265)	(0.3537)	(0.4053)	(0.3709)
Veteran	0.2430	0.3593	0.2971	0.3625
	(0.4289)	(0.4799)	(0.4572)	(0.4809)
Professional	0.3216	0.0753	0.4720	0.4591
	(0.4671)	(0.2639)	(0.4994)	(0.4985)
Log CLI	-0.0035	0.0061	-0.0131	0.0177
	(0.0633)	(0.0582)	(0.0555)	(0.0639)
Northeast	0.2511	0.2854	0.1510	0.3691
	(0.4337)	(0.4517)	(0.3581)	(0.4827)
South	0.2418	0.1325	0.3723	0.1522
	(0.4282)	(0.3391)	(0.4836)	(0.3594)
West	0.2208	0.2041	0.2377	0.2398
	(0.4148)	(0.4031)	(0.4258)	(0.4271)
Selection Parameter	0.6210	1.1575	1.5872	1.6899
	(0.1659)	(0.3698)	(0.3429)	(0.3302)
Postal			0.0287	0.0998
			(0.1670)	(0.2999)
Other Federal			0.2500	0.0532
			(0.4332)	(0.2245)