Public Wage Differentials and the Treatment of Occupational Differences

Abstract (Summary)
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[Headnote]
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INTRODUCTION
Public wage setting rests on the broad presumption that earnings in the government should be comparable to those in the private sector. Economists have spent great energy examining whether this presumption in policy is reflected in fact. Much of this work, founded on human capital theory, examines the earnings of individuals across a wide variety of positions holding constant determinants of earnings so as to focus on any remaining government wage premium (for recent reviews see Bender [1999] and Gregory and Borland [1999]). In academic research these examinations have largely supplanted the earlier methodology of comparing the earnings of narrowly defined positions common to the government and the private sector. Yet this earlier methodology lives on in the actual practice of public sector wage setting (Fogel and Lewin, 1974).1 Deciding which approach is more appropriate became known as the debate over "positions" versus "people." The debate remains important because the two approaches continue to yield substantially different conclusions (Belman and Heywood, 1996).

One of the long supposed advantages of the people, or human capital, approach is the ability to compare government and private sector workers doing different jobs, and so avoid the problem of wage surveys which were not able to compare unique occupations (Smith, 1977). Yet despite the genuine contributions of the human capital methodology, the large number of occupations unique to the government still confronts researchers with two imperfect extremes. On the one hand, detailed (three-digit) occupational controls can be included in earnings equations. Moulton (1990) examined the federal earnings differential controlling for literally hundreds of three-digit occupations. Yet the econometrics of this amounts to comparing only occupations common to the public and the private sectors, thereby eliminating the supposed advantage of the "people" approach. On the other hand, the included occupational detail can be limited, typically to the dozen or so major (one-digit) occupations. This, however, allows earnings premiums associated with the underlying detailed occupations to be included in the estimated governmental earnings differential. In short, just as was true of the survey methodology before it, the regression-based, human capital approach must consider the proper treatment of occupations unique to the public sector.

The object of this paper is to measure the relative importance of two explanations for the impact of introducing detailed occupational controls. The first is that such controls account for differences in the distribution of common occupations across the public and private sectors. The second is that detailed controls identify occupations unique to either sector, limiting the estimated differential to that based on occupations common to the public and private sectors. In fact, the latter explanation is more important for the federal differential, while the former is more important for the local government differential.

FRAMING THE ISSUE
The issue of the proper comparison (people or positions) remains highly relevant today, in part because of dramatic differences. A long line of human capital studies-including those of Smith (1977), Quinn (1979), and Gyourko and Tracy (1988)-have shown...
the federal employees earning as much as 20 percent more than their otherwise equivalent private sector counterparts. In contrast, Hartman (1983) contends that, when properly defined positions are compared, federal workers are actually underpaid. Both the periodic surveys by the Bureau of Labor Statistics and external consulting studies confirm Hartman's basic contention (see Krueger [1988] and Belman and Heywood [1996]). Indeed, an external study by Hay Associates suggests that federal pay was 10 percent below that of private sector employees performing similar tasks (again, see Krueger, 1988). These position level studies played a key role in the advent (starting in 1994) of "locality pay," supplements specifically designed to boost federal earnings for those working in high-wage areas.

Such conflicting results are not unique to the federal government. Belman, Franklin, and Heywood (1994) use state wage surveys in Wisconsin to aggregate detailed position level data and estimate a state-level differential of 16.9 percent and a local differential of 9.8 percent. Similarly, Moore and Newman (1991) find a double-digit advantage for employees of the Houston transit authority. Such large positive differentials far exceed the results from nationwide human capital studies, which typically show a near zero differential at the state level and a modest negative differential at the local level (see Moore and Raisian, 1991). Indeed, in human capital-based regression analysis limited to Wisconsin, Belman and Heywood (1995) find a state differential of only 2.3 percent and a local differential of -6.4 percent.

The conflicting results of human capital and occupation survey studies have generated "hybrid" research blending the two methodologies. At a minimum, most human capital studies include a few very broad occupational categories in the list of earnings determinants. Some studies hold the size of the employer constant (Belman and Heywood, 1990; Moore and Raisian, 1991) while others use alternative treatment of unionization and other attributes of the job (Belman and Heywood, 1993; Linneman and Wachter, 1990). Despite a partial convergence of methodologies, Moulton's (1990) remains the only major attempt to estimate the influence of the most elemental measure of the position in the Current Population Survey, the three-digit occupation.3 He finds that including controls for the hundreds of three-digit occupations yields a far smaller federal wage differential. This smaller federal differential could be justified if the information carried in occupation proxies for working conditions, unmeasured human capital, or effort.4

While investigating the importance of detailed occupations might be agreed upon, there remains a fundamental estimation problem. The narrowing of the federal differential caused by introducing detailed occupations may be the result of accounting for the different sectoral distribution of workers across common occupations (as implied by Moulton). Or, alternatively, the narrowing may be accounted for by the fact that occupations unique to the federal sector earn wages above those predicted by human capital regressions. Thus, what appears to be an abnormally high return to federal employment may be the "high" wages of occupations unique to the federal sector. Indeed, these two possibilities are econometrically inseparable, and accounting for detailed occupations sheds no light on the actual federal differential in this case.

To isolate these points the Current Population Survey is broken down by three-digit occupation. Table 1 examines the six largest occupations for each sector. Only one occupation, secretary, is among the six largest in each sector, and only one additional occupation is common between even two sectors, public administration official. Thus, examining the six largest occupations in each of the four sectors generates 20 different occupations. The largest occupations of the private sector include cook, truck driver, salesperson, and cashier. The largest occupations in the federal sector include mail carrier, postal clerk, auditor, inspector, and compliance officer. The state workforce is dominated by social workers, correctional officers, public administrators, and postsecondary teachers. The workforce of local government is dominated by elementary and secondary school teachers, teachers aides, police officers, and janitors. Also note that of the six largest occupations in the federal sector, only two are common to any extent with the private sector. Of the six largest in the state sector, four are common and, similarly, of the six largest in the local sector, four are common.

Table 2 shows that a substantial share of each sector's workforce is in that sector's six largest detailed occupations (out of some 500 possible occupations). Moving down the diagonal, 16 percent of the private workforce is in its six largest occupations compared with 36 percent of the federal sector, 24 percent of the state sector, and 40 percent of the local sector. While the concentration is greater in the public sectors, the proportion of each sector in the other sectors' six largest occupations presents an important comparison. For instance, while the six largest occupations of the private sector comprise 16 percent of the private workforce, these same occupations comprise less than 5 percent of the federal workforce. More importantly, the six largest federal occupations comprise 36 percent of the federal workforce but only 3 percent of the private workforce. Similarly, the six largest occupations of local government comprise only about four and a half percent of the private workforce. These comparisons make clear the dramatically different occupational compositions of the sectors.

Tables 1 and 2 suggest that a relatively few occupations dominate the public sectors and that many of the occupations important in one sector are either a small proportion of, or absent from, the other. The private sector has 480 distinct occupations, while the federal, state, and local sectors have 342, 353, and 350 occupations, respectively. In our sample, 31 percent of all federal workers are in occupations that lack a private sector occupational counterpart. Of state and local government employees, respectively, 12 and 13 percent are also in occupations absent from the private sector. Of all private sector workers, 13 percent are in occupations that do not exist in the federal sector, 7 percent are in occupations that do not exist in the state sector, and 18 percent are in occupations that lack a counterpart in local government. These numbers make clear that the distribution of workers across occupations shared by the public and private sectors is very different and that many workers are in occupations unique to the public or private sectors. At issue is which of these two differences comes to the fore when estimating public earnings differentials that account for detailed occupation.

The importance of the unique occupations is foreshadowed in Table 3, which compares the means across sectors. As Table 3
shows, those in unique occupations in the public sectors are much more highly educated, older, and many times more likely to be in managerial positions than those in the unique occupations in the private sector. The earnings of those in unique public sector occupations are also much higher; 70 percent higher in the federal sector. This difference is much greater than the difference across shared occupations and suggests that those in unique occupations in the private sector are concentrated among lower-earning occupations while those in unique occupations in the public sector (and certainly the federal sector) are concentrated among higher-earning occupations.

EMPIRICAL STRATEGY AND INITIAL RESULTS

The two roles of occupation are distinguished by comparing regression estimates of public wage differentials produced with increasingly disaggregate measures of occupation from three closely related samples. The first sample mirrors that used in most past research and consists of all public and private employees without regard to whether they are in unique or common occupations. The successive estimates produced by moving from more aggregate to less aggregate measures of occupation isolate the effect of more precise occupational control but do not, of themselves, differentiate between the two potential roles of occupation in the estimates. The second sample, a sub-sample of the first, is limited to occupations common to the public and private sector. Changes in the estimated public wage differential associated with use of less aggregate occupational controls are then attributable to better control for differences in the distribution of common occupations and are unrelated to the effects of unique occupations. The difference in results from the first and second samples represents the effect of unique occupations. The third sample consists of only individuals in unique occupations. The estimates resulting from this sample are not designed to produce accurate public differentials. Instead, they help indicate the size of the contribution from the unique occupations to typical estimates of the overall public differential. As we emphasize, this contribution may not be attributable to the process of public wage setting and may, alternatively, reflect occupational premiums.

These three samples are constructed separately to estimate public earnings differentials for the federal, state, and local governments. Thus, the models are estimated with individuals from the level of government under study and their private sector counterparts. For example, estimates of the federal government differential are made using federal and private employees but exclude state and local employees.

The sample is drawn from the outgoing rotations of the 1997 and 1999 Current Population Survey (CPS). The sample includes employed, nonagricultural, civilian workers between ages 18 and 65 and is structured so that there is only one observation on an individual. The outgoing rotation tapes were chosen to ensure adequate representation within each narrow occupational category. Using the two years of outgoing rotations yields a federal sample, for instance, of 9504, seven times Moulton's sample of 1328.

As a caveat, the data set used to capture large occupational cells comes with the limitation of few measures of job attributes. Their inclusion could alter the differential. Moreover, estimates of the earnings differential are clearly only one portion of a more general estimate of differences in total compensation. Despite these shortcomings, the most common measure of comparability in the literature is the earnings differential based on human capital regressions such as those that follow.

Log linear hourly earnings equations were estimated by replicating Smith, which are very similar to Moulton. The explanatory variables include detailed measures of: completed education; years of age; age squared; an interaction of the education dummies and age; two dummies for marital status; three dummies for race; union status; part-time status; gender; three regional dummies; residence in an SMSA; and a series of five dummies for urban areas of various size. At each stage the public differential is measured first by a dummy for public status in a single equation and then as the unexplained residual in a two-equation Oaxaca (1973) decomposition. The differentials are estimated four times, each time with increasingly disaggregate occupational controls. The first estimation has no occupational controls. The second adds nine one-digit controls for major occupation. The third estimation disaggregates into 38 controls slightly more fine than two-digit (detailed) occupations. Finally, a fourth specification includes all 509 three-digit occupational controls. The null hypothesis that increasingly detailed controls add nothing to the explanatory power of the regression is tested at each stage by a conventional F-test. For example, in moving from one- to two-digit controls, the appropriate F-test recognizes that the two-digit specification reduces to the one-digit specification if restrictions are imposed that all of the two-digit occupations have the same coefficient within the relevant one-digit occupations. A similar F-test can be constructed for the move from two- to three-digit controls.

The purpose of this article is to distinguish the two potential effects of occupational controls. Toward this end, and to limit the number of parameters the reader must consider, we consciously limit our estimations to what might be considered a prototypical specification. Thus, we do not attempt to control for firm and establishment size effects (Belman and Heywood, 1990) for specific cities and their respective cost of living (Smith, 1981), for how the differential varies across the earning distribution (Bender, 2003; Mueller, 1999), or for the interactive influence of gender. Thus, to take this last example, both men and women are included in a mixed gender sample. Krueger (1988), Moulton (1990), Perloff and Wachter (1984), and Hirsch, Wachter, and Gillula (1999), among others, have also used mixed gender samples. Yet Asher and Popkin (1984), Belman, Heywood, and Voos (2002), and Smith (1976) indicate that the government differentials for women may be magnified by pay discrimination against women in the private sector; this tends to increase the magnitude of the federal differential in mixed-gender samples. Such indications show the underlying tension in the standard of private sector comparability. An exact mimicking of the private sector would include replicating a pattern of gender discrimination causing many to suggest the government should be an "ideal employer" even at the cost of comparability (Belman and Heywood, 1996).

The proper treatment of gender is an important policy issue, as are issues of firm and establishment size, variations in cost of living, and the position within the earnings distribution at which the differential is estimated. Yet settling these issues, and others,
to establish the one true differential is beyond the scope of this paper. Instead, the specifications used in this study involve a set of variables common to most studies of the public sector. Moreover, we emphasize that the basic point of our inquiry is not greatly changed by modest variations in specification.\textsuperscript{11}

The Federal Sector

Table 4 presents estimates of the federal wage differential: The top panel provides estimates of the federal wage differential for the sample of individuals in all occupations; the center panel provides estimates for individuals in the sample of nonunique (common) occupations; and the bottom panel provides estimates for individuals in the sample of unique occupations. The first column presents the federal differential from equations without occupational controls. The upper entry is the estimate from the single-equation system, while the lower entry is the simple arithmetic average of the federal and private base differentials from a Oaxaca decomposition. The following columns present estimates with controls for one-, two-, and three-digit occupations for single-equation (upper) and averaged Oaxaca (lower) methods.

Beginning with estimates for the complete sample (top panel), those without occupational controls indicate substantial overpay-ment for federal workers. The single-equation estimate shows a federal premium of 22.7 percent while the Oaxaca measure is 23.5 percent. The differential based on the single equation using onedigit controls (major occupation), the first entry in the second column of the table, drops to 19.3 percent while the Oaxaca estimate is 18.7 percent. Estimates with two-digit occupation controls are only slightly smaller than those obtained with major occupation. The single-equation federal differential is 18.9 percent while the Oaxaca estimate is 19.1 percent. Three-digit controls, provided in the last column, result in substantially lower estimates of the federal differential with a single equation estimate of 14.2 percent, 4.7 percentage points lower than the estimate obtained with two-digit occupational controls. In each instance the F-statistic associated with adding more detailed occupational controls improves the explanatory power of the equation. Overall, moving from no controls to three-digit controls results in an 8.5 percentage point (22.7 - 14.2) or a 38 percent decline in the estimated federal differential.

Recognizing that most studies use one-digit controls, the move to three-digit controls results in a 5.1 percentage point (19.3 - 14.2) or 26.4 percent decline.

The top panel of Table 4 also provides an initial indication of the difficulties associated with estimation of public wage differentials where there are unique occupations. Although Table 4 includes Oaxaca estimates for the first three equations, those without and with one- and two-digit controls, these are omitted for equations with three-digit occupations because, although calculable, the estimates are misleading. The private sector equation does not have coefficients for three-digit occupations that are unique to the public sector and the public equation does not have coefficients for occupations unique to the private sector. These equations would then be used to estimate the public and private wage for the private sample and for the public sample. Calculating an estimated wage for a sample's own sector (the public wage of public employees, the private wage of private employees) presents no problem, but consider generating the private sector wage for the sample of public workers. Some public employees are in occupations that lack private sector counterparts and there is no estimated return, no coefficient, for their occupation. No matter what their actual occupation, such employees are implicitly assigned to the omitted private sector occupation and are so assigned an occupational return of zero.\textsuperscript{12} Although this might be correct or close to correct for some public employees in unique occupations (e.g., postal letter carriers—see Belman, Heywood, and Voos, 2002), it will generally not be correct individually or on average. The extent of misclassification of occupation and consequent error in projecting such employees' private sector wage can be quite large. Almost one-third of the federal workforce and 13 percent of the private sector workforce would be misclassified. This extent of misclassification and the consequent error in projecting the wage in the alternative sector renders the Oaxaca estimate unreliable. As a consequence, we do not report a three-digit Oaxaca differential in Table 4.\textsuperscript{13}

Although the top panel suggests that better controls for occupation (not surprising) reduce the federal differential, it does not inform us of the degree to which this results from better accounting for different distribution across common detailed occupations and from better controls for occupations unique to one sector. To examine the relative importance of these two sources the methodology was replicated on a sample that excludes observations in three-digit occupations unique to the federal or private sector. If the unique occupations are driving the results, the differentials recovered with one- and two-digit occupational controls should be very similar to the Table 4 estimates using three-digit controls. More than 150 of 509 three-digit occupations were eliminated in the federal estimations because they were unique to one or the other sector.

The center panel in Table 4 reproduces the earlier estimates on the new sample. As in the top panel, the movement to less aggregate occupational controls reduces the federal differential. Movement from no controls to one-digit controls results in a 3.4 percentage point decline in the differential (versus 3.5 percentage points in the sample with all employees) with a similar change in the Oaxaca estimates. Movement to two-digit controls reduces the federal differential by a further 0.6 percentage points (similar to the 0.4 in the upper panel), while further movement to three-digit controls results in a 1 percentage point reduction, substantially smaller than the 4.7 percentage point decline in the upper panel. The F-test of the null that more detailed controls for occupation does not improve the explanatory power of the equation is rejected in each case.

The influence of unique occupations on the federal differential can be obtained by comparing the coefficients in the top and center panels. The estimated differentials from the sample of nonunique occupations are substantially smaller than was obtained with the full sample of occupations save the final, which is identical to the upper-panel estimate. In each instance, removing the unique occupations from the sample causes the federal differential to fall by 3.5 percentage points. These estimates may be used to decompose the proportion of the federal differential attributable to unique occupations. Taking a model with no occupational controls as our base, removing unique occupations from the sample reduces the federal differential
from 22.7 percent to 19.2 percent. Further control to the level of three-digit occupations reduces the differential to 14.2 percent. Of the decline in the federal differential, 41.2 percent is accounted for by unique occupations (3.5/8.5). Most studies however incorporate one-digit controls for occupation, and this provides an alternative base for our calculation. In this instance, the removal of unique occupations causes the estimated federal differential to decline from 19.3 percent to 15.8 percent while further controls for occupation in the sample of nonunique occupations reduces the differential to (again) 14.2 percent. Here unique occupations account for 68.6 percent of the decline in the federal differential (3.5/5.1). More important to those estimating federal differentials, the influence of increasing the detail of the occupational controls beyond one digit is modest in samples of common occupations. The move from a handful of one-digit controls to hundreds of three-digit controls reduces the federal differential by only 1.6 percentage points and accounts for only 10.1 percent of the differential (1.6/15.8).

Only when unique occupations are included in the sample does movement to three-digit controls greatly influence the estimated differential. Thus, the 5.1 percentage point decline in the differential from the one- to the three-digit equation in upper panel is largely a product of the inclusion of unique occupations in the sample. Further, the large effect of moving to three-digit controls in a sample that includes unique occupations results because, econometrically, including three-digit controls excludes unique occupations from the estimation of the public differential. (This is proven analytically in the Appendix.) The equality of the three-digit estimates of the federal (and state and local) differentials in Tables 4 and 5 results because both public differentials are, in effect, estimated on the sample of occupations common to the public and private sectors.

To emphasize the role of unique occupations, the public differential was estimated using a sample of only those workers in unique three-digit occupations. In other words, no workers in the sample come from occupations common to both the public and private sectors. The bottom panel presents the results. As the first column shows, the differentials are extremely large, more than double the estimates for all occupations or for common occupations. The differential is 46.4 percent after adjusting for the explanatory variables, declines to a still very large 36 percent with one-digit controls, and then returns to 44 percent with two-digit controls. It is not possible to estimate a differential with three-digit controls as the federal dummy is perfectly collinear with the occupational controls. These extremely large estimates are built into the typical federal differential, but there is no way to determine whether these large numbers represent a true federal premium, a pattern of occupational premiums, or some combination. With unique occupations, sectoral premia are fundamentally intertwined with occupational differentials. Including these workers in overall estimation, as done in the standard estimations in the literature, implicitly assumes that the observed premium is entirely due to public employment, which seems extreme.

To make this issue concrete, both judges and law enforcement agents are occupations unique to the federal government and each earns more than predicted by a private sector earnings equation. The prevailing methodology includes these "overpayments" as part of the public differential. On those occasions when three-digit controls are included, the "overpayments" become associated with the two occupational dummies that, in essence, remove these workers from the calculation of the public differential. The differential is then estimated only on common occupations. Obviously, some progress might be made if, in an effort to identify appropriate public and private sector comparisons, job characteristics could be compared for unique occupations. This, however, would be exactly how the survey, or position based, approach would begin. The supposed advantage of the people approach was that it avoided the difficulties in making such job comparisons. Yet, as emphasized, avoiding such comparisons implicitly assumes that the large governmental premiums associated with unique positions are, in total, part of the public wage differential.

Given the central role of unique occupations in the federal differential, our conclusion regarding federal earnings differs from that of earlier work. The fact that the federal sector has a greater concentration of three-digit occupations that earn aboveaverage wages in either sector seems far less important. More important is that the unique three-digit occupations in the federal sector pay more than the unique occupations in the private sector, after accounting for the explanatory factors.

This brings us back to the decades old issue of people or positions. Occupations unique to one sector cannot easily be compared using the people approach. The inclusion of a full set of three-digit dummies in a human capital approach appears to allow comparison across these unique occupations but amounts econometrically to limiting the sample to the common occupations. In fact, a sample of observations drawn only from the unique occupations would result in perfect collinearity between the occupational controls and federal status indicator and no federal differential could be estimated. Left unresolved is whether the large differential associated with unique occupations occurs because of legitimate earnings differences associated with the respective unique occupations or because the federal government overpays those occupations.

The State Sector

In contrast with the federal wage differentials, unique occupations increase state differentials only modestly and differences in the distribution of common occupations have a similarly modest effect on the differential. The full sample state differential in the top panel of Table 5 is relatively insensitive to the level of occupational control. The differential is 2.4 percent in the single-equation model with no controls for occupation. Introduction of one-digit controls causes this to decline to 0.0 percent, movement to two-digit controls increases the estimate to 3.3 percent, while control at the three-digit level reduces the state differential to 2.9 percent, similar to the initial estimate. Changes in the Oaxaca differential follow a similar pattern. Increasingly disaggregate control for occupation has a similar effect on the sample of non-unique (shared) occupations as shown in the center panel. The single-equation estimate with no occupational controls is 0.4 percent. This declines to -2.0 percent with one-digit controls, rises to 1.6 percent with two-digit controls, and to 2.9 percent with three-digit controls. Excepting the models with three-digit controls, estimates for the sample of common occupations are 1.6 to 2.0 percentage points below their full sample counterparts. Estimates for the sample of unique occupations are unstable. The single-equation estimates decline with...
less aggregation of occupational controls, from 27.4 percent with no controls, to a nonsignificant 2.0 percent with one-digit controls, to -9.6 percent with two-digit controls, while the Oaxaca estimates increase. This sensitivity to the level of occupational control and method of estimating the differential again points to problems with the measurement of public-wage differentials for unique occupations.

As in the prior case, the results from the first two panels may be marshaled to decompose the two sources of the occupational effect. Using a model without occupational controls for a base, removing unique occupations from the sample reduces the differential from 2.4 to 0.4 percent. Moving to greater occupational control increases the differential by 2.5 percentage points to 2.9 percent. Working from a base of one-digit occupational controls, removing unique occupations reduces the state employee differential from O to -2.0 percent and moving to a model with three digit occupational controls raises the differential to 2.9 percent, a 4.9 percentage point increase. One may conclude that excluding the unique state occupations modestly decreases the state differential while use of more detailed occupational controls modestly increases the state differential. In combination, the two influences largely offset one another (compare the combined sample with no controls, 2.4 percent with the nonunique sample with three-digit controls, 2.9 percent).

The Local Sector

The pattern of change for local employees is distinct from that of the federal or state sectors (Table 6). Again, beginning with the full sample (the top panel), the single equation estimate with no controls indicates a differential of 3.5 percent. Movement to one-digit controls causes the differential to decline to 1.9 percent, but introduction of two- and three-digit controls causes the differential to increase to 7.8 and to 7.0 percent, respectively. When the sample is limited to nonunique occupations the differential is initially 1.5 percent, declining to essentially zero with one-digit controls, but then rising to 6.0 and 6.9 percent with two- and three-digit controls respectively. Estimates for the unique occupations are typically positive for both single-equation and Oaxaca differentials but become unstable as more disaggregate occupational controls are introduced.

Moving from the full sample to the sample of nonunique occupations reduces the public wage differential by 1.8 to 2.0 percentage points, while movement from more to less aggregate controls increases the measured differential by between 6.8 (from a one-digit base) and 5.4 (from a model without occupation controls) percentage points. The increase in the differential associated with different distributions of common occupations is then three to four times larger than the decline associated with eliminating the unique occupations.

Thus, while the federal sector results indicated the importance of removing the unique occupations, the local sector results illustrate the importance of accounting for the differences in distribution of common occupations across sectors. Moreover, while the influence of such accounting is a decrease in the federal differential, it is a large increase in the local differential. These results also suggest that the conclusion that local employees are paid similarly to their private sector comparables is partially due to the occupational controls typically used in these models.

CONCLUSION

The treatment of unique occupations has long posed issues for the measurement and realization of public sector comparability. A limitation of private sector occupational surveys has been that they provide little or no information for establishing wages for those occupations unique to the public sector. Regression approaches to comparability apparently resolved this problem by matching individuals by their personal characteristics rather than by position. A limitation of the "people" approach has been that it has not recognized the role that the skills, abilities, and conditions associated with specific occupations play in wage determination. If position as well as people matter in wage determination, adequate control for the effects of position is central to providing useful measures of public wage differentials.

Because of the inability to accurately measure all of the job and personal characteristics that contribute to earnings, disaggregate occupation itself appears an important element of wage formation. Yet its value is limited when estimating public sector earnings differentials. Large shares of government employment are in occupations unique to the public sector. By econometric construction, including three-digit controls for these occupations results in estimated differentials identical to that found when limiting the sample to common occupations. This eliminates a supposed advantage of the human capital approach, the ability to compare workers across sectors doing different jobs. A common "solution" to this quandary has been to estimate models that include no occupational controls or just very broad controls. Yet, this carries the cost of potentially confusing occupational with the sectoral effects. The estimation in this paper suggests that the decline in the federal differential from O to -9.6 percent with two-digit controls, while the Oaxaca estimates increase. This sensitivity to the level of occupational control and method of estimating the differential again points to problems with the measurement of public-wage differentials for unique occupations.

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This finding is generated almost exclusively by federal white-collar workers. The federal differential in a sample of common occupations is very similar when using one-digit and when using three-digit controls.

REFERENCES

[Reference]

Footnote

1 An exception to this is the U.S. Postal Service, which has used differentials derived from human capital regressions in bargaining and interest arbitrations since the early 1980s.

Footnote

2 The pure human capital approach is still found in the literature. For example, Katz and Kreuger (1991) compare returns to education over time in federal and private sectors and include no occupational controls in their earnings equations.

Footnote

3 Although the 509 three-digit occupations are the elemental classification of occupations available in the CPS, it is based on the far less aggregate Dictionary of Occupational Titles, which identifies more than 12,000 distinct occupations.

Footnote

4 For example, the one-digit occupation “protective services” includes not only public uniform services (police, firefighters, and correctional personnel) but also private security guards. It is not until three-digit controls are included that private security guards, who have very different job requirements and duties, are distinguished from public uniformed officers.

Footnote

5 In addition to those identified in Table 2, occupations unique to the public sector include legislator, governmental chief executive public administrators and officials, judge, air traffic controller, supervisor of police and detectives, police and detectives, sheriff and postsecondary home economics teacher.

Footnote

6 In addition to those identified in Table 2, occupations unique to the private sector include funeral director, mining engineer, podiatrist, numeric control tool programmer, automotive salesman, shoe salesman, purchasing agent, lathe and turning setup operative, pattern makers, metal plating machine operators, railroad brake, signal, and switch operators, machine feeders and offbearers.

Footnote

7 We initially deflated the earnings by month using January 1997 as the base but found that the percentage differentials did not change in the second digit after the decimal point for the single-equation estimates. Thus, all estimates are not deflated.

Footnote

8 Moulton (1990) provides far greater detail on location and finds that this detail also serves to shrink the federal differential. To remain focused on the role of occupation we exclude such controls. It is interesting to note that the new “locality pay” signed into law by President Clinton in November of 1993 should serve to reduce the role of location in explaining the federal differential. Moulton estimates also exclude postal workers on the grounds that they are not comparable, although this itself is a matter of some debate and most researchers include them. Indeed, some have examined a specific postal sector/private differential (see Perloff and Wachter, 1984). Our approach will ultimately exclude workers only in those occupations within the postal service that are unique.

Footnote

9 The Oaxaca residual is not calculated for the three-digit level of occupational control as, in the presence of unique occupations, the estimates are influenced by the choice of the base occupational group. At the three-digit level many of the occupational controls are unique to specific sectors. When the “other sector” wage estimates are formed, observations for unique occupations are, by the construction of the Oaxaca, consigned to the base category. For example, if we use secretaries for the private sector base, then the unique federal occupations are implicitly assigned a “secretarial” occupational effect. If the private sector base was lawyers, then the unique occupations would be assigned that occupational effect. Thus, all of the unique occupations from one sector have predicted wages in other sector as if they were in the base detailed occupation. The estimated differential is dependent on the choice of the base occupation and is not unique.

Footnote

10 The exact definitions of the occupational controls are available from the authors and follow in straightforward fashion from the three-digit list of occupations associated with the CPS.

Footnote

11 The procedure we outline was replicated for the federal differential separately on men and women. While the exact numbers differ, the pattern we present represents a rough average of those calculated on the separate genders.

Footnote

12 In this instance the occupational base group is secretaries but the problem remains no matter which occupation was taken as the base.

Footnote

13 Oaxaca estimates typically bracket the single-equation estimate.

Footnote

14 As the history of public sector wage determination indicates, obtaining consensus on private occupations that are reasonably similar to unique public sector occupations is quite difficult. For example, the record of the postal service arbitrations of the last 20 years indicates that while the postal unions view parcel truck drivers at UPS and FedEx are appropriate comparables to letter carriers, the postal service advocates newspaper and flyer delivery workers as comparables. This problem is magnified in the national data sets used for regression studies as the occupations of interest are typically mixed in with others that are not suitable. For example, UPS package car drivers have been classified by the Census as Light Truck Drivers (three-digit occupation code 913) and are combined with all drivers of trucks or vans with a capacity of less than 23,000 pounds of gross vehicle weight.

Footnote

15 This finding is generated almost exclusively by federal white-collar workers. The federal differential in a sample of common occupations is very similar when using one-digit and when using three-digit controls.

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[Appendix]
APPENDIX
This appendix demonstrates that the coefficient estimated on a public sector indicator when including three-digit occupational dummies unique to one sector is identical to that estimated with only common occupations.
Consider an earnings regression with a public sector dummy and n three-digit occupational dummies:

\[ Y^*_{i} = \beta_0 + \beta_1 I^*_{g} + \beta_2 I^*_{j} + \epsilon_{i} \]  

where

- \( Y^*_{i} \) is the log of the hourly wage for person i
- \( I^*_{g} \) is a dummy variable indicating a government employee
- \( I^*_{j} \) is a dummy variable indicating person i working in occupation j
- \( \beta_0 \) is a constant
- \( \beta_1 \) is the coefficient for the effect of government employment on wages
- \( \beta_2 \) is the coefficient for the effect of occupation j on wages
- \( \epsilon_{i} \) is a classical error term

Occupations may be unique to the government or to the private sector or may include workers in both sectors. There are J occupations in the data. The vector of estimated coefficients is:

\[ [\beta] = ((1^n)\ast sup -1^n\ast Y) \]  

\[ \text{Eq. 2} \]

where

- n is the number of government employees, n is the number of employees in occupation j and n is the number of government employees in occupation j.

To derive [\beta] we need the first row of (1^n)sup -1^n\ast Y, which can be obtained using Theil's partitioned inverse.

\[ \text{Eq. 1} \]
Which can be simplified to:

With this, the coefficient can be expressed as:

Note that $n_{g}^{Y}g^{Y}$ is the sum of wages of all government workers and that the value of $n_{jg}^{Y}j^{Y}$ depends on the extent to which occupation $j$ has both public and private sector workers.

i) If occupation $j$ has only private sector workers, neither expression in the difference in parentheses of (5) includes wages from this occupation.

ii) If occupation $j$ has only public sector workers, the wages from occupation enter $n_{g}^{Y}g^{Y}$, but are immediately subtracted out as $n_{g}^{Y}g^{Y}j^{Y} = n_{jg}^{Y}j^{Y}j^{Y}$, leaving the expression in (5) unaltered by inclusion of workers from occupation $j$.

iii) Only when an occupation includes both government and private sector employees is $n_{gj}^{Y}j^{Y} > 0$ and $n_{gj}^{Y}j^{Y} = n_{g}^{Y}g^{Y}$ thus altering the estimated value of the coefficient.

Several notes: First, parallel results obtain for the term $k$ such that it is unaltered by inclusion of occupations that are unique to one sector. Second, this presentation has abstracted from explanatory controls other than the occupational and sector indicators. Including them complicates the derivation considerably without changing the basic point.