Compensation in the Nonprofit Sector

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Revised: June 2001

Abstract

This analysis provides an in-depth investigation of the determinants of pay in the nonprofit sector. Data are for 25-55 year olds from the 1994-1988 Current Population Survey Outgoing Rotation Groups. Our econometric results support the hypothesis that compensation is primarily determined in competitive labor markets without “labor donations” to nonprofit employers. One implication is that nonprofit workers receive virtually the same pay as observationally equivalent employees in similar positions with profit-seeking enterprises. We can not rule out the possibility of nonprofit wage penalties or premiums for selected groups; however, the differentials are generally small and competition plays the dominant role in nonprofit wage setting.

* We thank Richard Frank, H.E. Frech III, Dan Rosenbaum, Doug Staiger, and Burton Weisbrod for helpful comments and William Black for research assistance. Ruhm gratefully acknowledges financial support from the National Science Foundation (SES-9876511) and from an NBER faculty fellowship for the study of nonprofit institutions.

Compensation in the Nonprofit Sector

Nonprofit enterprises are an increasingly important part of the American economy. Between 1980 and 1997, the number of nonprofit associations grew 54 percent (Bureau of the Census, 1998) and the fraction of GDP accounted for by them rose from 2.9 to 4.3 percent (Bureau of Economic Analysis, 1998). Nonprofits utilize the majority of volunteer labor and are responsible for a significant proportion of paid employment in some industries. Despite this growing significance, compensation in the nonprofit sector remains poorly understood. There is little question that nonprofit workers earn less than observably similar employees of for-profit firms.[1] However, the distribution of jobs and worker characteristics varies markedly, raising the possibility that the disparities reflect compensating differentials or individual heterogeneity not accounted for in standard earnings regressions.

We address these issues by providing an in-depth investigation of the determinants of pay in the nonprofit sector. Our goal is to ascertain how the earnings of individuals employed by nonprofit enterprises compare to those of identical workers in similar jobs with profit-seeking firms.[2] We use several complementary approaches including: analyzing the size and pattern of the cross-sectional wage differentials (with and without controls for job characteristics), estimating how earnings change when workers shift between nonprofit and for-profit jobs, and examining the disparities in wage levels and growth rates for workers in narrowly defined industries and occupations with a substantial mix of nonprofit and for-profit employment.

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[1] For instance, Preston (1989) indicates that nonprofit managers and professionals earn 18% percent less per hour than their for-profit counterparts, controlling for human capital characteristics and the (one-digit) industry of employment.

[2] Previous researchers (e.g. Krueger, 1988; Moulton, 1990; Belman and Heywood, 1993) have examined whether government workers are "overpaid" relative to private sector employees. We compare workers in nonprofit enterprises to those in profit-seeking firms, rather than government, because there is no reason to assume that government wages are determined by market forces.

Our most important overall finding is that nonprofit workers are paid in competitive labor markets and do not “donate” labor to their employers by accepting lower wages. What this means is that, after controlling for limited set of job characteristics, persons working for nonprofits receive approximately the same pay as they would if employed in equivalent positions by profit-seeking firms. This is true even though nonprofit employees earn an average of 11 percent less than their counterparts with similar observed attributes. The reason for the lower earnings is that nonprofit jobs require fewer hours and are concentrated in a small number of industries that tend to offer relatively low pay but are likely to be desirable places in which to work. Our evidence does not rule out the possibility of nonprofit wage penalties or premiums for selected groups. However, the magnitudes of the differentials are generally small and do not detract from the dominant role of competition in nonprofit wage setting.

1. Relative Earnings in the Nonprofit Sector

Previous research provides several reasons why compensation in nonprofit enterprises might deviate from that in profit-seeking firms. Seminal work by Hansmann (1980), emphasizes that a key feature of nonprofits is that they are barred from distributing net earnings. He and others (e.g. Easley and O’Hara, 1983; Handy and Katz, 1998) argue that nonprofits therefore will be prevalent in markets where the consumer is in a poor position to judge the price, quantity, or quality of services, because this organizational form helps to solve the consumer trust problem resulting from asymmetric information.[3]

The nondistribution constraint provides two reasons why earnings in nonprofit enterprises might *exceed* those in profit-seeking companies. First, managers may have less incentive to hold

[3] Hansmann claims that nonprofits have reduced incentives to raise prices or cut quality, since it is more difficult for the controlling organization to benefit from the resulting increase in profitability. There are other important differences between nonprofit and profit-seeking enterprises. For instance, nonprofits are

down wages since they do not gain from the resulting cost-reductions. This has been called "philanthropic wage-setting" by Feldstein (1971) or "attenuated property rights" by Frech III (1976).[4] Second, nonprofits have less incentive to shirk on quality and so may choose to employ better quality workers.[5] These sources of disparities are distinct. Either may imply greater labor costs in nonprofit enterprises but only the first means that nonprofit workers earn more than they would in identical for-profit jobs and so represents a deviation from competitive labor markets. More generally, since economic models of nonprofits typically involve solving some (possibly restricted) optimization problem, a higher level of compensation is far from assured.[6]

Conversely, some individuals may be willing to "donate" a portion of their paid labor to "socially responsible" nonprofit employers by accepting reduced compensation.[7] The resulting wage gap will be reinforced if nonprofits attract persons placing a relatively high value on institution-specific fringe benefits (such as working conditions) and a low value on money.[8] Also, Lakdawalla and Philipson (1998) postulate that nonprofits will be concentrated in more competitive and less profitable sectors of the economy, where the benefits of choosing the nonprofit form exceed the costs imposed by the nondistribution constraint and other limitations

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frequently exempt from corporate income taxes and receive preferential treatment in state contract procurement processes (Frank and Salkever, 1994).

[4] Feldstein argues that nonprofit hospitals pay relatively high wages due to "philanthropic wage-setting". Frech III emphasizes that "attenuated property rights" reduce the price of nonpecuniary amenities such as pleasant offices and short working hours, resulting in higher production costs.

[5] For instance, in Newhouse's (1970) model of nonprofit hospitals, managers maximize a utility function with quantity and cost as arguments subject to a zero profit constraint. This leads them to choose the lowest cost method of production but to oversupply quality.

[6] For example, a nonprofit that attempts to maximize size or market share still has incentives to minimize labor costs.

[7] A recent study indicates that Cornell University graduates in "socially responsible" occupations or companies were paid substantially less than their counterparts, controlling for sex, curriculum, and grades (Frank, 1996). Nonprofits also employ the vast majority of volunteer labor (Steinberg, 1990).

[8] Rose-Ackerman (1996) argues that "ideologues" may accept lower pay for nonprofit work because they receive greater certainty that their efforts achieve altruistic goals, rather than benefiting stock-holders.

of nonprofit status (e.g. the bar on equity financing). The increased competitiveness implies downwards pressure on wages.[9]

As mentioned, nonprofits may pay relatively low wages if they offer positive compensating differentials such as short work hours or low risk of job loss. Preston (1988) argues that the resulting earnings penalty will be smaller within narrowly defined industries, since organizations engaged in the same activities are likely to generate fairly comparable social benefits and working conditions.[10] Conversely, the generation of social benefits is less likely to be linked to the category of jobs (e.g. clerical workers), implying that the estimated nonprofit differential may be little affected by the addition of occupation controls to an econometric model.[11]

These arguments notwithstanding, we believe that a logical starting point is to hypothesize that nonprofit compensation is determined in competitive labor markets. Competition implies that the marginal worker will be indifferent between identical positions in nonprofit and profit-seeking enterprises. In its absence, some jobs will be rationed and some employers pay more than needed to fulfill their demand for labor. However, competitive markets need *not* require identical levels of pay. As mentioned, wages may deviate if there are compensating differentials or if individuals are willing to donate a portion of their labor to nonprofits. Therefore, we are particularly interested in considering the joint hypothesis of competitive labor markets and the absence of labor donations. The testable prediction is that the

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[9] Weisbrod (1988) believes that nonprofits arise when the government is unable to meet the demand for public goods (e.g. care for the medically indigent). This has no obvious predictions for wage-setting.
[10] Nonprofit workers might even be paid more than others in the same industry due to "philanthropic wage-setting". One implication is that nonprofit premiums are less likely within industries that are extremely competitive.
[11] However, Preston (1989) and Handy and Katz (1998) argue that the nonprofit gap will be greater for managers than blue collar workers, since the latter are further removed from the generation of social

nonprofit wage differential will be eliminated by including sufficient controls for worker and job characteristics.

2. Previous Research

Previous studies of nonprofit compensation, summarized in Figure 1, yield ambiguous results. Early examinations (e.g. Johnston and Rudney, 1987; Shackett and Trapani, 1987; Preston, 1989) suggest a large nonprofit wage penalty. However, these are hampered by the lack of information on the type of employer, requiring the researchers to impute nonprofit status,. By contrast, DuMond's (1997) analysis of the 1994-1995 Current Population Survey (CPS), which does identify the type of employer, indicates a considerably smaller (6 to 11 percent) gap in wage levels. Moreover, his fixed-effect estimates, which focus on wage changes occurring when individuals switch between for-profit and nonprofit employment, imply small (0 to 4 percent) and statistically insignificant earnings penalties.[12]

Researchers focusing on narrowly defined industries also obtain equivocal results. Weisbrod (1983) finds that public interest lawyers earn 20 percent less than those in the private sector and believes that this is due to heterogeneity in preferences rather than worker quality. However, Goddeeris (1988), using the same data, claims that the wage penalty reflects personal characteristics and that public interest attorneys earn no less than if they were employed by profit-seeking companies. Borjas, Frech III, and Ginsburg (1983) argue that the relatively high pay observed in nonprofit nursing homes represents rent-sharing due to attenuated property rights. Conversely, Holtmann and Idson (1993) claim the wage premium occurs because nonprofit nursing homes use higher quality labor and that registered nurses could actually earn

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benefits. Easley and O'Hara (1983) similarly suggest that the nondistribution constraint will lead to relatively large pay reductions for nonprofit managers.

more if they switched to for-profit facilities. Preston (1988) shows that federally regulated nonprofit day care centers pay 5 to 10 percent more than for-profit facilities and interprets this as evidence of philanthropic wage-setting. However, she finds no differential for non-federally regulated centers, while Mocan and Viola (1997) show that the statistically insignificant 4 percent overall nonprofit wage premium varies considerably depending on type of ownership and racial composition of the staff.[13]

Most recently, Leete's (2001) examination of data from the 1990 Census indicates that the overall nonprofit differential is eliminated by including detailed controls for industries and occupations. Within 3-digit industries, nonprofit workers are as likely to obtain statistically significant wage premiums as penalties. These conclusions need to be interpreted with caution, however, because the controls for industries and occupations are so extensive (as many as 20,000 industry-occupation interactions in some models) that there is likely to be little variation in the type of employer within many of the narrowly defined industry-occupation cells.[14]

3. **Data**

The analysis below uses data on 25 to 55 year olds from the 1994-1998 Current Population Survey Outgoing Rotation Groups (CPS-ORG).[15] The CPS is a nationally representative survey of roughly 50,000 households. Individuals are interviewed for four

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[12] DuMond's results should be interpreted with caution because it is not clear how he deals with movements into or out of public sector, few respondents switch types of employment over the two year period examined, and because he does not consider the effects of endogenous mobility.
[13] Roomkin and Weisbrod (1999) indicate that there is ambiguity even within industries. Focusing on six top managerial positions in hospitals, they find that nonprofits offer lower compensation in three (chief executive officer, chief operating officer, and top patient care executive) but higher pay in three others (chief financial officer, top human resources executive, and head of nursing services).
[14] There are several other potential methodological problems. First, hourly wages are probably measured with considerable error, since they are estimated from data on annual incomes, weekly work hours, and weeks worked per year. Second, it is not obvious how the analysis treats individuals holding multiple jobs at a point in time or during the year. Third, the demographic characteristics controlled for are unlikely to adequately account for the heterogeneity between nonprofit and for-profit workers.
[15] This age range avoids the special experiences of those making school-to-work or retirement transitions.

months, out of the sample for eight, and then return for four final months. The outgoing rotation groups include persons in the last of each of the four month segments, hereafter referred to as years 1 and 2.[16] Our cross-sectional sample includes data for year 1. The longitudinal analysis refers to individuals for whom information is available in both years 1 and 2 (12 months apart). Not all persons can be matched across years. For instance, individuals are not followed if they change addresses between the surveys. Our match rate of 63 percent is similar to that obtained by other researchers (e.g. MacPherson and Hirsch, 1995) who used slightly less stringent matching criteria. The procedures used to perform the matching are detailed in Appendix A.

The dependent variable is the natural log of weekly wages on the "main" job.[17] Weekly rather than hourly earnings are used because the latter are likely to be measured with greater error. However, most of the regressions directly control for work hours. Respondents report the type of employer and we are primarily interested in comparing persons whose main job is with a private nonprofit organization to those working in for-profit companies. Public sector employees are therefore deleted from the regression analysis (but included when examining how nonprofit employment is distributed across industries and occupations).[18]

The econometric models also control for a quadratic in age and dummy variables for education (high school dropout, high school graduate, some college, college graduate, graduate degree), marital status (currently married, previously married, never married), race (white, black,

[16] The outgoing rotation groups contain supplemental questions on weekly earnings and work hours not included in the regular monthly CPS.
[17] For multiple job-holders, the "main" job is the one at which the person usually works the most hours. If hours are the same at two jobs, it is the position of longest employment.
[18] Persons employed in government jobs in *either* years 1 or 2 are excluded when using the panel sample. In preliminary work, we estimated models that included government workers and directly controlled for public sector employment in the regressions. The resulting nonprofit differentials were similar to those below.

other nonwhite), Hispanic origin, sex, metropolitan area residence, and the survey year. Some specifications add regressors for industry and occupation in the "main" job, as detailed below.

Variable means are similar for the cross-sectional and panel samples (see Appendix Table B.1). The main differences are that the matched individuals earn more and are older, more likely to be married, and to live in metropolitan areas. These disparities probably reflect patterns of mobility and employment stability.

4. Theory and Methods

Assume that the utility (U) an individual receives from employment depends on wages (W), possibly some additional benefit from working in a nonprofit position (N), and other job characteristics (Z). If the utility function is additive separable, this can be expressed as:

(1) $$U = W + cN + dZ,$$

with the wage coefficient normalized to one. Denoting the best available for-profit and nonprofit jobs with the subscripts p and n, utility is maximized by choosing nonprofit employment if

(2) $$W_p + \delta Z_p < W_n + c + dZ_n$$

or

(2') $$W_p - W_n < c + d(Z_n - Z_p),$$

and by working in a for-profit company if the inequality is reversed.

Competitive labor markets imply that the marginal worker is indifferent between the two types of employment. This occurs if:

(3) $$W_p - W_n = c + d(Z_n - Z_p).$$

Equation (3) demonstrates that nonprofit enterprises pay less than profit-seeking firms if either workers are willing to "donate" labor to them ($c > 0$) or they offer superior working conditions ($Z_n - Z_p > 0$). Conversely, equal pay will be received in the two types of employment if there is no labor donation to nonprofits *and* no difference in average working conditions. More generally, controlling for job characteristics, identical wages imply an absence of labor donations (since $E(W_p - W_n | Z_n - Z_p) = c$).

Our empirical implementation first examines the cross-sectional wage equation:

$$W_i = \alpha + X_i\beta + N_i\gamma + \varepsilon_i, \tag{4}$$

where W_i is the natural log of weekly wages for worker i, X is a vector of individual characteristics, N a dummy variable indicating nonprofit employment, and ε is the regression disturbance. The coefficient of interest $\hat{\gamma}$, shows the predicted nonprofit (log) wage differential, controlling for personal but not job characteristics. With an i.i.d. error, $E(\hat{\gamma}) = c + d$, implying that the nonprofit parameter captures the combined effect of labor donations and average differences in working conditions.

Of greater interest are the results obtained by estimating

$$W_i = \alpha + X_i\beta + N_i\gamma + \delta Z_i + \varepsilon_i, \tag{5}$$

where Z is a vector of job characteristics (some combination of work hours, industries, and occupations). If the covariates in Z adequately account for the heterogeneity in working conditions, $E(\hat{\gamma}) = c$ and a coefficient of zero suggests that wages are set competitively without labor donations to nonprofits.

One problem is that if the explanatory variables do not sufficiently control for the selection into nonprofit employment, $cov(N_i\varepsilon_i) \neq 0$ and $\hat{\gamma}$ is biased. For instance, there might be a negative differential because persons with relatively low productivity select into nonprofit jobs. With panel data, first-difference models can sometimes be used to account for these sources of heterogeneity. For example, if the error term for person i at time t is $\varepsilon_{it} = f_i + e_{it}$, where f_i is a fixed-effect and e_{it} an i.i.d. disturbance, the wage equation can be rewritten as:

$$W_{it} = \alpha_t + X_i\beta_t + N_{it}\gamma + \delta Z_{it} + f_i + e_{it}, \tag{6}$$

and selection bias occurs if $cov(N_{it} f_i) \neq 0$. However, the change between period 1 and 2 is:

(7) $$W_{i2} - W_{i1} = (\alpha_2 + X_i\beta_2 + N_{i2}\gamma + Z_{i2}\delta + f_i + e_{i2}) - (\alpha_1 + X_i\beta_1 + N_{i1}\gamma + Z_{i1}\delta + f_i + e_{i1})$$

or

(7') $$\Delta W_i = \alpha + X_i\beta + \Delta N_i\gamma + \Delta Z_i\delta + \Delta e_i,$$

where $\Delta W_i = W_{i2}\text{-}W_{i1}$, $\Delta N_i = N_{i2}\text{-}N_{i1}$, $\Delta Z_i = Z_{i2}\text{-}Z_{i1}$, $\alpha = \alpha_2\text{-}\alpha_1$, $\beta = \beta_2\text{-}\beta_1$, and $\Delta e_i = e_{i2}\text{-}e_{i1}$.[19] Since the fixed-effect has been differenced away, econometric estimates of equation (7') eliminate the bias due to all sources of time-invariant heterogeneity.

The wage change models will yield inconsistent estimates, however, if $cov(\Delta N_i \Delta e_i) \neq 0$. The most important reason this may occur is that job mobility is not random. Instead, workers typically change positions (and sectors of employment) when doing so is expected to raise total compensation. Thus, turnover occurs when $E(\Delta e_i) > 0$, which may induce correlation between ΔN and Δe. Under reasonable conditions, however, this endogenous mobility can be exploited to bound the size of the nonprofit differential.

To show this, it is useful to define two dummy variables. The first, PN, is equal to one for individuals switching from for-profit to nonprofit positions (abbreviated by $P \rightarrow N$); the second, NP, indicates movements from non-profit to for-profit jobs (denoted by $N \rightarrow P$). Since $\Delta N = PN - NP$, (7') can be rewritten as:

(7'') $$\Delta W_i = \alpha + X_i\beta + (PN_i - NP_i)\gamma + \Delta Z_i\delta + \Delta e_i.$$

Relaxing the constraint that PN and –NP have the same coefficient gives the asymmetric wage change equation:

(8) $$\Delta W_i = \alpha + X_i\beta + PN_i\gamma_1 + NP_i\gamma_2 + \Delta Z_i\delta + \Delta e_i.$$

[19] This assumes that the characteristics in X do not change between periods 1 and 2.

Notice that (8) collapses to (7") if $\gamma_2 = -\gamma_1$, and that $\hat{\gamma}_1$ and $-\hat{\gamma}_2$ provide alternative estimates of the nonprofit differential. Importantly, economically motivated turnover implies $\mathrm{cov}(PN_i \Delta e_i) > 0$ and $\mathrm{cov}(NP_i \Delta e_i) > 0$. As a result, both $\hat{\gamma}_1$ and $\hat{\gamma}_2$ are likely to be upwards biased and the nonprofit effects gaps estimated by $\hat{\gamma}_1$ ($-\hat{\gamma}_2$) will be biased upwards (downwards). One testable prediction is that the nonprofit wage penalty should be larger when estimated from $P \rightarrow N$ moves rather than from $N \rightarrow P$ transitions.

An example helps to illustrate why this is so. Assume, *ceteris paribus*, that nonprofits pay 5 percent less than profit-seeking firms. With exogenous turnover, $P \rightarrow N$ ($N \rightarrow P$) mobility would then be associated with 5 percent slower (faster) earnings growth than for workers not changing sectors. Further assume, however, that turnover is endogenous such that switches of the type of employment occur when wage offers 2 percentage points better than average are received. $P \rightarrow N$ moves will therefore lead to 3 percent relative wage decreases and $N \rightarrow P$ mobility to a 7 percent rise in relative earnings, thus bounding the estimated nonprofit differential between 3 and 7 percent.[20]

5. The Distribution of Nonprofit Employment

Tables 1 and 2 summarize how nonprofit employment is distributed across industries and occupations. The first two columns of each table display the industry or occupation composition of all employment and nonprofit jobs. The third and fourth columns indicate the share of employment in the sector accounted for by nonprofit and profit-seeking organizations, with government the residual category. The last column shows average weekly wages in the specified

[20] The symmetric wage change model correctly estimates the nonprofit differential only in the case where $P \rightarrow N$ and $N \rightarrow P$ moves occur with unequal frequency. We do not control for the selection into employment in either year 1 or 2 because of the difficulty in obtaining plausible identifying restrictions. This failure will not cause bias as long as the selection process is similar for persons working in nonprofit and for-profit jobs.

industry or occupation. For example, the first row of Table 1 demonstrates that religious organizations are responsible for 0.7 percent of all employment but 10.4 percent of all nonprofit jobs, that 85 percent of employees in this industry worked for nonprofits, and that persons in this industry were paid an average of $581 per week.

Nonprofit positions are concentrated in eight narrowly defined industries – religious organizations, membership organizations, social services, hospitals, other health services, higher education, nursing/personal care facilities, and primary/secondary education – that account for 21 percent of jobs but 85 percent of nonprofit employment (see Table 1). Fifty-five percent of nonprofit work is located in just three industries – hospitals, social services, and religious organizations – that are responsible for only 8 percent of all jobs. The share of nonprofit employment in these three industries ranges from 34 to 85 percent, compared to an economy-wide average of under 6 percent. By contrast, there is virtually no nonprofit involvement in the personal/business services, transportation/communication/utilities, wholesale/retail trade, agriculture/construction/mining, manufacturing, or public administration sectors. These industries are responsible for two-thirds of employment but just 6 percent of nonprofit jobs. It is noteworthy that, with the exception of education and hospitals, the industries with high shares of nonprofit employment all pay below average wages, and that all of them are widely viewed as engaging in "socially desirable" activities.

Nonprofit work is more dispersed across occupations. Nevertheless, ten occupations – clergy/religion, health professionals, social work, health technicians, health services, educators/librarians, secretaries, other administrative support, managers, and non-health services – account for 86 percent of nonprofit employment versus 54 percent of all jobs (see Table 2).[21]

[21] Health managers are included in the health professional category and education managers in the educator/librarian occupation group. Therefore, managers here refer to those outside of these two fields.

Nonprofits are virtually absent from the production, sales, laborer, and transportation occupations that provide 36 percent of all employment. There is no evidence of below average pay in occupations with large nonprofit representation.

The econometric estimates below frequently include dummy variables for these eight industries and ten occupations. This contrasts with previous researchers (e.g. Preston, 1989 or DuMond, 1997) who hold constant one-digit industries and occupations, or with Leete (2001) who includes detailed controls for up to 20,000 industry-occupation interactions. Within broad industries or occupations there is substantial heterogeneity in the prevalence of nonprofit activity. Conversely, extremely detailed industry and occupation covariates are likely to absorb much of the "effect" of nonprofit status, since many cells will be dominated by a single class of employer. Our classification system has the advantage of providing a parsimonious but targeted method of accounting for many important differences in job characteristics, while focusing on within-industry (or occupation) variation in the type of employment.

Nonprofit status is reported by survey respondents, raising the possibility of classification error. Undercounting appears particularly likely. A careful analysis by the Hodgkinson, et al. (1996) indicates that nonprofits constituted 6.7 percent of the paid work force in 1994. Conversely, only 5.7 percent of our cross-sectional sample claim this type of employment, suggesting that around 18 percent of nonprofit workers erroneously report for-profit employment.[22] An identical 18 percent error rate is obtained by assuming that 100 percent of employment in religious organization is nonprofit, versus to the 85.1 percent reported by CPS

[22] There is little reason to believe that government or for-profit workers frequently misreport. For example, 100 percent of respondents in public administration report holding government jobs and 99 percent of those in the wholesale/retail trade or manufacturing industries claim for-profit employment.

respondents.[23] Such classification errors are likely to cause the observed wage gaps to be smaller than the actual differentials, since some nonprofit jobs are averaged in with for-profit positions. However, using reasonable assumptions, such misclassification will lead to only a slight understatement of the nonprofit gap.[24]

6. Cross-Sectional Wage Differentials

This section examines cross-sectional wage differentials. Mean weekly earnings for selected industries and occupations are displayed in Table 3. For the full sample, nonprofit workers earn an average of 3 percent less per week than those in profit-seeking firms. However, there is a nonprofit premium within each of the five industries detailed (accounting for more than 70 percent of nonprofit employment), ranging from 9 percent in non-hospital health services to 18 percent in social services. The overall wage gap combined with intra-industry premiums reflects a heavy concentration of nonprofit jobs in poorly paid industries (e.g. social services, religion, and nursing/personal care). There is no corresponding pattern of high nonprofit pay within specific occupations.

Disparities in average earnings may reflect individual heterogeneity, rather than a pay penalty. For instance, nonprofit workers are slightly older and considerably more educated than their counterparts, which is likely to increase relative earnings, but they also work fewer hours and are more often female (see Appendix Table B.1) which will reduce them. A careful

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[23] The 34 percent of hospital workers claiming nonprofit affiliation appears low, given that around 65 percent of acute care hospital beds are in nonprofits. But this industry category also includes heavily for-profit psychiatric hospitals, rehabilitation facilities, and post-acute care hospitals. Hodgkinson, et al. (1996) estimate that hospitals accounted for 33.7 percent of nonprofit employment in 1994, somewhat higher than the proportion in Table 1 (29.9 percent) which covers the 1994-1998 period. However, the hospital share of nonprofit jobs has been trending sharply down over time.

[24] Using a simplified version of the formula derived by Leete (1999), the ratio of the observed to actual gap in log wages (G) is $G = 1 - [\phi\rho/(\phi\rho+(1-\rho))]$, where ρ is the employment share of nonprofits and ϕ is the reporting error rate among nonprofit workers. Assuming that ρ=.067 and ϕ=.18, the observed nonprofit differential will therefore be 98.7 percent as large as the actual gap.

econometric analysis is needed to disentangle these effects from other sources of wage differentials.

Table 4 displays the coefficient on nonprofit status from cross-sectional regressions of equations (4) and (5). The dependent variable is the log of weekly wages in year 1.[25] The first row shows results for the full CPS-ORG sample; the second refers to the panel of individuals observed in both years 1 and 2. As mentioned, public sector employees are deleted from all of the regression samples. In addition to nonprofit status, the econometric specifications control for the survey year, age, marital status, race/ethnicity, education, and metropolitan residence. Weekly work hours and dummy variables for eight industry and ten occupation categories are also frequently included.

Consistent with earlier research (e.g. Preston, 1989), nonprofit employment is associated with an 11 percent wage penalty after controlling for individual attributes but not job characteristics (model a). Accounting for shorter work hours reduces the disparity to around 6 percent (column b) and a slight (1 to 2 percent) premium is predicted when industries are also controlled for (column c) but the occupation covariates have little effect on the expected wage gap (compare columns d versus b and e versus c).[26] Unless otherwise noted, the regressions in the remainder of this analysis always control for work hours, industries, and occupations. When this is done, nonprofit employees are predicted to earn virtually the same wages as their for-profit counterparts (specification e). This result is consistent with earnings being determined in competitive labor markets, where nonprofit jobs pay less because they involve fewer work hours

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[25] Similar results are obtained for year 2.

[26] Information on work hours is missing for around 7 percent of respondents. To avoid excluding these individuals, they are assigned a value of zero hours and a dummy variable for missing hours is included. However, the results are similar when these persons are deleted from the sample.

and are located in industries offering positive compensating differentials. There is no evidence of labor donations based specifically on nonprofit status.

The remainder of Table 4 tests to sensitivity of the findings to changes in the sample or specification. Column (f) restricts the analysis to full-time workers (those employed more than 35 hours per week). This is done because part-time jobs are more common in nonprofit enterprises, raising the possibility of biased estimates due to structural differences in the compensation of full-time and part-time employees.[27] The last two columns provide alternative treatments of top-coded wages, which affect 1.6 percent of for-profit and 1.4 percent of nonprofit workers.[28] Persons with top-coded values are assigned earnings equal to twice the censored amount in model (g) and deleted from the sample in column (h). The estimated nonprofit differential is robust to these changes. Controlling for industry, occupation, and work hours, the earnings of nonprofit workers are predicted to be within 2 percent of those of their for-profit peers in all of these cases. With the exception of model (g) for the cross-sectional sample, the differential is always less than 1 percent and is statistically insignificant.

7. Wage Changes

The estimates in the previous section could be biased the explanatory variables inadequately account for selection into the type of employment. For instance, the cross-sectional models might conceal an earnings penalty if nonprofit workers are of unusually high quality. We next investigate this possibility by estimating first-difference models which, as mentioned, automatically control for all time-invariant sources of heterogeneity.

The first two columns of Table 5 show employment shares in all jobs and nonprofit positions for subgroups stratified by sex, education, and race/ethnicity. Most striking is the

[27] In the cross-sectional sample, 80 percent of nonprofit workers are employed more than 35 hours per week, versus 90 percent of those in profit-seeking enterprises.

disproportionate representation of women and highly educated individuals in nonprofits. The 46 percent of workers who are female hold over 70 percent of nonprofit jobs; the 56 percent with a college education account for 79 percent of nonprofit employment.

The last three columns of the table document average changes in log wages between years 1 and 2. Consistent with endogenous turnover, earnings growth is somewhat faster for persons switching between nonprofit and for-profit jobs than for those remaining in the same sector. More noteworthy is the relatively similar wage growth of persons making $P \rightarrow N$ and $N \rightarrow P$ moves; weekly earnings rise 8.5 percent for the former group versus 7.1 percent for the latter, which further suggests the small size of any nonprofit differential. However, there is some variation across demographic categories. In particular, the faster wage growth after $N \rightarrow P$ than $P \rightarrow N$ transitions raises the possibility of a larger nonprofit penalty for men. Small numbers of minorities in nonprofit employment imply that results for blacks and Hispanics should be interpreted with caution.

Table 6 summarizes alternative econometric estimates of the nonprofit differential for the panel sample. The first column refers to wage levels in year 1. The second indicates results from the symmetric wage change model (equation 7). The third and fourth columns display $\hat{\gamma}_1$ and $-\hat{\gamma}_2$, from the asymmetric earnings growth model (equation 8). As discussed, these two coefficients bound the estimated nonprofit gap. All of the regressions control for individual characteristics and the survey year. The levels equations also hold constant work hours, industries, and occupations; the growth models account for changes in these regressors.

The full sample results confirm that nonprofit workers receive virtually the same pay as their for-profit counterparts with equivalent individual and job characteristics. The predicted

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[28] Weekly earnings are top-coded at $1,920 in 1994-1997 and $2,880 in 1998.

differential is –0.4 percent in the cross-sectional regressions, –1.0 percent in the symmetric earnings growth model, and bounded between –0.4 and –1.5 percent in the asymmetric change equations. None of the parameter estimates are statistically significantly different from zero. Thus, the results support the hypothesis of competitive wage determination with no labor donation based specifically on nonprofit status.

There are some differences across demographic groups. Most importantly, men receive around a 3 percent wage penalty for nonprofit work, which may explain why they are under-represented in these jobs.[29] The data also suggest imprecisely measured 2 to 6 percent nonprofit premiums for blacks. Even with this heterogeneity, the scope for labor donations to nonprofit employers appears to be small and competitive labor markets play a dominant role in setting wages.

We further tested the robustness of the results, by utilizing data from Displaced Worker Supplements (DWS) to the 1994, 1996, and 1998 Current Population Surveys to estimate nonprofit differentials for workers involuntary losing jobs due to plant closure, slack work, or position/shift abolished. One advantage of focusing on job losers is that wage changes are more likely to be exogenous (i.e. there is no presumption that $E(\Delta e_i > 0)$), since the mobility is involuntary. The results are consistent with the absence of an overall nonprofit differential. For example, the symmetric wage change regressions predict a statistically insignificant 0.7 percent nonprofit wage premium, with some evidence of more favorable outcomes for women than men.[30]

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[29] Preston (1990) discusses this issue in detail.
[30] Data limitations require that the DWS regressions be estimated without controlling for work hours. Our confidence in the DWS results is further limited by two factors. First the sample size is small (n=4,730) and only about 3 percent of respondents are displaced from nonprofit jobs. Second, changes in the type of employer (i.e. between nonprofit and for-profit) appear to mainly occur after the job loser has looked for but failed to find acceptable work in the original sector. As evidence of this, both P ® N and N ® P

8. Differentials Within Industries and Occupations

We next compare nonprofit workers to their counterparts in the same industry or occupation. Table 7 summarizes the findings for four industries – social services, hospitals, other health services, and nursing/personal care facilities – that account for 54 percent of nonprofit employment but also have substantial involvement by for-profit companies. The regression models are the same as above except that industry controls are excluded and the first-difference sample is limited to persons in the specified industry in both years 1 and 2. The latter restriction avoids confounding the effects of industry mobility with changes in the type of employer.

The findings generally support competitive wage setting without labor donations to nonprofits. For instance, the cross-sectional regressions provide no indication of a nonprofit penalty. Instead, small *premiums* (between 1.1 and 4.5 percent) are predicted in three of the four industries. However, these higher earnings are mostly due to transferable individual characteristics, as evidenced by the statistically insignificant 0 to 1 percent earnings gaps obtained in the symmetric wage change equations for the Social Service, Hospital, and Other Health Service industries. The one exception is that nonprofit nursing/personal care facilities appear to pay a premium of between 3 and 10 percent, raising the possibility of rent-sharing. Nevertheless, the absence of nonprofit differentials in the other three industries suggests that the deviations from competitive wage-setting are limited.

Table 8 displays results for four occupation groups – managers (outside of health and education), health professionals, health technicians/service workers, and administrative support workers – that are responsible for 46 percent of nonprofit employment and have substantial

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transitions are associated with large declines in relative wages (by 9.1 and 13.2 percent respectively) compared to persons not changing sectors.

participation by profit-seeking firms.[31] The regression specifications are identical to Table 7, except that industry rather than occupation covariates are included and the wage change sample is restricted to those in the specified occupation in the two years.

Wages appear to be set competitively and without labor donations to nonprofits in the three non-managerial occupations. For these workers, the gaps in wage levels are of modest size and generally statistically insignificant, and the first-difference regressions indicate that the small cross-sectional disparities are largely due to unobserved transferable characteristics. Conversely, nonprofit managers earn about 7 percent less per week than their for-profit peers, controlling for observables, while the first-difference regressions suggest a wage penalty of between 0 and 8 percent.[32] Thus, the data indicate some scope for labor donations in management occupations.

9. Discussion

Our econometric analysis supports the hypothesis that compensation in the nonprofit sector is primarily determined by competitive labor markets, without explicit labor donations based upon nonprofit status. Weekly wages are an average of 11 percent lower in nonprofit than for-profit jobs, holding constant worker characteristics. However, this disparity is entirely accounted for by a combination of shorter hours and the concentration of nonprofit positions in relatively low paying industries. Thus, nonprofit employees earn virtually the same pay as observationally equivalent individuals with similar positions in profit-seeking enterprises. Further support for competitive wage-setting is obtained by noting that the earnings of persons

[31] These categories correspond to those in Table 2, except that health technicians and service personnel have been combined into a single group, as have secretaries and other administrative support workers.
[32] Managers in the health or education industries were excluded from the manager category to maintain consistency with the groupings used in Table 2. When managers are defined to include these persons, the cross-sectional nonprofit penalty declines to between 4 and 5 percent and the differential from the wage change regressions ranges from –4.3 to 2.6 percent. We thought that the nonprofit penalty for managerial employment might explain the relatively low earnings of male nonprofit workers. The data did not

transitioning between for-profit and nonprofit employment grow at virtually the same rate as those not switching types of employers. This is expected if the (small) cross-sectional differentials result from individual heterogeneity rather than differences in pay-setting.

These findings are consistent with Leete's recent evidence that nonprofit and for-profit workers in sufficiently similar jobs receive equal pay. However, to eliminate the cross-sectional differential, she controlled for as many as 20,000 industry-occupation interactions. By contrast, we can do so by including covariates for just eight narrowly targeted industries. Our first-difference equations further indicate that the result is robust to accounting for all sources of time-invariant heterogeneity. Finally, we carefully consider the consequences of endogenous mobility and show that our asymmetric wage change models provide reasonable upper and lower bounds on the nonprofit differential. The econometric results fit with the model predictions and the bounds on the estimates are generally narrow.

Why are nonprofit jobs are disproportionately located in low-paying industries? The most likely possibility is that these sectors perform "socially desirable" activities (e.g. helping the sick or teaching children), so that employees are willing to work for reduced compensation. This represents a variation of the labor donation hypothesis. However, the key distinction is that individuals accept lower wages not because the employer is a nonprofit, but rather because of the specific goods or services provided. Other potential explanations seem less likely. Wages might be low because nonprofits locate in relatively competitive industries (Lakdawalla and Philipson, 1998) or because disadvantaged groups (such as women or nonwhites) are limited to these sectors.[33] However, for-profit jobs would then be rationed, which is at odds with the evidence

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support this, since the negative nonprofit coefficient was larger (in absolute value) for male than female managers in all specifications that we examined.

[33] Industries with high nonprofit shares are often quite competitive. For instance, the four-firm sales concentration ratios in the nursing/personal care facility and social service industries were 14.8, and 7.9

that that highly educated workers (who presumably have the most options) disproportionately select nonprofit employment. By contrast, the education result is consistent with low relative wages reflecting positive compensating differentials.[34]

Some deviations from competitive wage-setting are possible. For instance, a small (2 to 4 percent) nonprofit wage penalty is obtained for males and a less precisely estimated 0 to 8 percent disadvantage for nonprofit managers. Conversely, nonprofit employment in nursing/personal care facilities is associated with a 3 to 10 percent earnings premium. These exceptions show that wage determination is unlikely to be uniform across the entire nonprofit sector and that a single model will probably not capture all of its elements. Nevertheless, the most useful first approximation is that compensation is set competitively and without labor donations that are explicitly linked to nonprofit status. Divergences from competitive wage setting appear to be relatively minor.

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percent in 1992 (Bureau of the Census, 1995). For comparison, Scherer and Ross (1990) indicate four-firm sales concentration ratios of 20 percent or higher for more than 80 percent of U.S. manufacturing industries in 1982. However, the relevant market is likely to be more localized for services than manufacturing, so this comparison may overstate the competitiveness of the sectors dominated by nonprofits. Models emphasizing restricted job availability (e.g. Bergmann, 1974) typically focus on occupations rather than industries and it is not obvious what mechanism might limit access to the latter, given the broad set of occupations they employ.

[34] There is some direct indication of positive compensating differentials for nonprofit jobs. DuMond (1997) finds that pension and health insurance coverage are relatively high and displacement rates relatively low in these positions, although he does not control for individual characteristics or the industry of employment. Gonyea (1999) argues, based on limited evidence, that nonprofit employers may be more sensitive to work-family issues than profit-seeking firms.

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Appendix A: Construction of Longitudinal Sample from the CPS-ORG Files

The Current Population Survey contains household identifiers (ID codes) and record line numbers but not individual codes. Individuals from the same month in consecutive years can potentially be identified using the household ID codes and record line numbers. The household ID represents a permanent residence and does not follow families that relocate. The coding of this variable was changed from 12 to 15 characters in July of 1995, implying that households whose ORG months crossed this date could not be matched. Also, since different states sometimes use the same household ID, state (FIPS) codes are needed to uniquely identify the household.

The following procedure was used to create the ORG matched panel data set. First, individual cross-sectional data sets were created for each of the years 1994 through 1998. The samples were restricted to non-self employed persons working for pay and aged 25 through 55. Second, the five annual data sets were merged and persons in time periods that were not potentially matchable (because of changes in coding the household ID) were deleted. This yielded a sample with 333,134 person-year observations, including 165,516 for year 1 and 167,618 at year 2. The sample was then sorted by year, month, household ID, state, and record line. Several passes of the data were used to limit the sample to cases where consecutive observations were one year apart and had the same calendar month, household ID, state code, and record line. This reduced the sample to 236,122 observations. To further insure that the matched observations referred to the same individual, we deleted cases where there was a change (between year 1 and 2) in sex, race, ethnicity, education, or more than a two year difference in age. (A two year age difference was allowed because the surveys could take place on different days of the month.) Finally, the matched pair was deleted if the first (second) observation was listed as the eighth (fourth) month in sample, rather than the reverse. These restrictions reduced the sample to 103,857 individuals (207,714 person-year observations), corresponding to 62.4 percent of the original sample and 62.7 percent of year 1 observations.

Appendix B

Table B.1:
Variable Means By Sector of Employment for Cross-Sectional and Panel Samples

Variable	Cross-Sectional Sample		Panel Sample	
	For-Profit	Non-Profit	For-Profit	Non-Profit
Weekly Earnings/Hours				
Earnings ($)	573	557	596	572
Work Hours	40.9	38.2	41.2	38.6
Education				
High School Dropout	.116	.039	.103	.033
High School Graduate	.363	.189	.357	.173
Some College	.283	.290	.292	.304
College Graduate	.178	.273	.187	.276
Graduate Degree	.060	.210	.060	.214
Marital Status				
Currently Married	.649	.649	.699	.685
Never Married	.190	.199	.161	.176
Race/Ethnicity				
Black	.110	.110	.100	.096
Other Nonwhite	.046	.035	.043	.036
Hispanic	.107	.045	.094	.044
Other Characteristics				
Age (years)	38.2	39.8	38.8	40.3
Male	.559	.306	.561	.302
Metropolitan Residence	.780	.787	.825	.834
Sample Size	227,018	18,203	74,047	6,143

Note: See notes on Tables 1 and 2. All variable means are computed using CPS sampling weights. The cross-section includes respondents in year 1 of the 1994-1998 Current Population Survey Outgoing Rotation Groups. The panel includes respondents for the same period who could be matched in the fourth months of years 1 and 2. All variables are measured at year 1.

Figure 1: Previous Research on Nonprofit Earnings Differentials

Study	Data	Results	Comments
Borjas Frech III, and Ginsburg (1983)	1973-74 National Nursing Home Survey	Nursing home workers in religious-affiliated nonprofits earn 4% less per than for-profit employees; those in other nonprofits receive an insignificant 1.6% premium. Some evidence of higher wages for homes with more generous Medicaid reimbursement programs.	Many results are statistically insignificant or sensitive to the choice of specifications.
DuMond (1997)	1995 Current Population Survey Outgoing Rotation Groups	Nonprofit workers earn 6% (11%) less per hour than counterparts without (with) controls for industry and occupation. Larger differential for males (19%) than females (0% to 5%). Gaps shrink to a insignificant 0% to 4% in first-difference models. Nonprofit workers have higher pension/ health insurance coverage and lower displacement rates.	Not clear how government workers are treated. Small number of transitions between for-profit and nonprofit employment in wage change models.
Frank (1996)	Cornell Employment Survey and other sources.	Nonprofit differential in annual earnings was –59% for recent Cornell graduates, controlling for sex, GPA, and college curriculum. Other evidence of negative compensating differentials for working for socially responsible employers.	Small and unrepresentative sample in main analysis; few controls.
Goddeeris (1988)	Nationally representative surveys of private and public interest lawyers in 1973/4.	Public Interest (PIL) lawyers earn 37% less than those in private firms but this is entirely due to differences in characteristics. They would earn no less if they switched into the private sector.	Sector definitions differ from Weisbrod (1983). Selection identified by community size, political activities/orientation.
Holtmann and Idson (1993)	Registered nurses in 1985 National Nursing Home Survey	Nonprofits employ higher quality registered nurses. OLS models reveal a 3% hourly wage premium in nonprofit homes and steeper experience/tenure profiles for them. However, selectivity-corrected models indicate that nurses in nonprofits actually earn less than they would if employed in for-profits.	No distinction between government and private nonprofit nursing homes. Identification restrictions of selectivity-corrected models are questionable.
Johnston and Rudney (1987)	1982 Census of Service Industries	The average annual earnings of nonprofit workers are 21.5% less than those employed in for-profit firms.	Hospitals, educational institutions, and religious organizations excluded. No controls for individual characteristics.
Leete (2001)	1990 Census, Five Percent Public Use Microdata Sample	No overall nonprofit wage differential after including detailed controls for industry and occupation. Among specified 3-digit industries with statistically significant nonprofit differentials, positive and negative effects are equally likely.	Estimated hourly wages may be subject to measurement error. Extremely detailed industry-occupation interactions could absorb nonprofit effects.

Figure 1 (continued)

Mocan and Viola (1997)	398 child care centers in CA, CO, CT, and NC	Nonprofit child care workers receive a statistically insignificant 4% premium in hourly wages; considerable variation by type of nonprofit and worker.	Extensive controls for human capital and center characteristics.
Preston (1988)	Abt Associates, 1976-77 National Day Care Center Supply Study	Nonprofit weekly wage premium of 5% to 10% for child care workers in federally regulated day care centers; no difference for other centers. Results consistent with the former being less competitive and able to pay rents to workers.	Center characteristics, labor quality, parental participation, and donations controlled for. Some differences across center types could persist.
Preston (1989)	1990 Survey of Job Characteristics (SJC); May 1979 Current Population Survey (CPS)	OLS results for SJC imply negative nonprofit differential of $\approx 20\%$ for managers/professionals, no effect for clerical workers; larger negative effects for both groups in CPS. Selectivity-corrected results sensitive to model specification. CPS wage change regressions indicate no differential for clerical workers, statistically insignificant 10% premium for managers and professionals. For-profit workers more often have pensions, health insurance.	SJC sample is small ($n \approx 300$). Exclusion restrictions are questionable for selectivity-corrected estimates. Nonprofit status inferred (not observed) in CPS data.
Roomkin and Weisbrod (1999)	Hay Management Consultants, 1992 Hospital Compensation Survey	Nonprofit hospitals offer higher base salaries but lower bonus payments to six top managerial positions. Total compensation is higher in three positions and lower in three.	Job complexity and hospital characteristics controlled for; individual characteristics are not. Low response rate (19%).
Shackett and Tapani (1987)	National Longitudinal Surveys of Young Men and Young Women	Compared to private nonregulated industries, the nonprofit wage differential is 11, 0, -14, and –8 percent for white females, black females, white males, and black males.	Nonprofit status not observed; instead it is assumed to include all persons in hospital and educational services industries.
Weisbrod (1983)	Same as Goddeeris (1988)	PIL lawyers earn 20% less annually than if employed in private sector. These attorneys are aware of the negative earnings effects and expect them to be permanent. Differences in preferences consistent with type of employment.	Small sample size (53 PIL lawyers); PIL lawyers may not be representative of other attorneys in nonprofits. Work hours and fringe benefits not controlled for.

Table 1: Industry Composition of Nonprofit Employment

Industry	Industry % of Employment: Overall (a)	Industry % of Employment: Non-profits (b)	% of Industry Employment in: Non-profits (c)	% of Industry Employment in: For-profits (d)	Average Weekly Wage ($) (e)
All Industries	100.0	100.0	5.7	76.4	581
"The Nonprofit Sector"	20.8	84.5	23.2	33.4	549
Religious Organizations (880)	0.7	10.4	85.1	14.9	437
Membership Organizations (881)	0.3	3.7	65.4	34.6	557
Social Services (861-871)	2.2	15.1	39.6	37.1	409
Hospitals (831)	5.1	29.9	33.5	49.9	599
Other Health Services (840)	1.8	5.1	16.3	67.6	544
Higher Education (850-860)	2.5	8.1	18.3	21.8	607
Nursing/Personal Care Facilities (832)	1.5	3.5	13.4	76.8	373
Primary/Secondary Education (842)	6.7	8.6	7.3	7.1	585
"The Rest of the Economy"	79.2	15.5	1.1	87.7	589
Finance/Insurance/Real Estate (700-712)	6.9	3.5	2.9	93.9	662
Personal/Business Services (721-791)	8.3	1.8	1.2	97.8	497
Transport/Communication/Utilities (400-472)	8.3	1.5	1.0	79.5	669
Wholesale/Retail Trade (500-691)	17.0	1.3	0.4	99.0	470
Agriculture/Mining/Construction (10-60)	7.4	0.5	0.4	92.4	569
Manufacturing (100-392)	19.0	1.1	0.3	99.2	629
Public Administration (900-932)	5.9	0.0	0.0	0.0	682

Note: Data are from the 1994-1998 Current Population Survey Outgoing Rotation Groups for persons in their fourth interview month (n=301,208); means are calculated using CPS sampling weights. The numbers in parentheses refer to three digit census industries.

Table 2: Occupation Composition of Nonprofit Employment

Industry	Occupation % of Employment: Overall (a)	Occupation % of Employment: Non-profits (b)	% of Occupation Employment in: Non-profits (c)	% of Occupation Employment in: For-profits (d)	Average Weekly Wage ($) (e)
All Occupations	100.0	100.0	5.7	76.4	581
Clergy/Religious Workers (176,177)	0.4	5.4	84.2	15.5	570
Health Professional (15, 83-106)	4.2	18.5	25.4	58.2	763
Social Worker (174)	0.8	3.1	23.7	17.2	564
Health Technician (203-208)	1.6	5.0	17.5	70.5	499
Health Service Worker (445-447)	2.0	4.4	12.7	72.0	309
Educators/Librarians (14,113-165)	6.4	13.7	12.3	16.8	678
Secretaries (313)	2.8	5.1	10.5	66.5	401
Other Managers (4-13, 16-37)	13.2	12.7	5.5	79.6	824
Other Admin. Support (303-309, 314-389)	12.9	11.0	4.9	70.8	444
Non-Health Service Worker (403-444, 448-469)	9.8	7.3	4.3	68.2	358
Farming/Fishing/Forestry (473-499)	1.4	0.6	2.5	88.5	342
Non-Health Technician (209-235)	2.2	0.9	2.3	82.6	724
Handlers/Cleaners/Laborers (863-889)	3.5	0.8	1.2	92.4	394
Production/Craft/Repair (503-799)	18.3	3.4	1.1	94.1	546
Sales (243-285)	9.9	1.8	1.0	97.6	582
Transportation (803-859)	4.4	0.7	0.9	88.1	581

Note: See note on Table 1. The numbers in parentheses refer to three digit census occupations.

Table 3:
Average Weekly Wages By Sector of Employment

Industry/Occupation	Weekly Earnings ($)		Non-Profit Differential (%)
	For-Profit	Nonprofit	
Full Sample	573	557	-2.7
	(1)	(3)	(0.5)
Industry			
Social Services	359	422	17.5
	(5)	(6)	(2.2)
Hospitals	572	636	11.2
	(4)	(5)	(1.1)
Other Health Services	539	586	8.8
	(6)	(11)	(2.6)
Education	547	602	10.0
	(7)	(7)	(1.8)
Nursing/Personal Care Facilities	360	412	14.4
	(4)	(9)	(2.8)
Occupation			
Health	566	652	15.2
	(3)	(6)	(1.2)
Educator	556	594	6.8
	(7)	(8)	(1.9)
Administrative Support	432	381	-11.8
	(1)	(4)	(1.0)
Other Managers	844	701	-16.9
	(3)	(9)	(1.3)
Non-Health Service Workers	292	274	-6.2
	(1)	(5)	(1.9)

Note: The table shows average weekly earnings on the main job for respondents in year 1 of the 1994-1998 Current Population Survey Outgoing Rotation Groups, with standard errors in parentheses. The nonprofit differential shows the percentage difference in weekly wages compared to persons in the same industry or occupation holding for-profit jobs. The education industry includes primary, secondary, and higher education. Health occupations include professionals, technicians, and service workers. Administrative support includes secretaries and other administrative support occupations.

Table 4:
Econometric Estimates of the Nonprofit Differential In Weekly Wages

Sample/Additional Controls	**(a)**	**(b)**	**(c)**	**(d)**	**(e)**	**(f)**	**(g)**	**(h)**
Cross-Sectional Sample	-.119 (.005)	-.059 (.004)	.013 (.005)	-.056 (.004)	-.013 (.005)	-.009 (.005)	-.019 (.005)	-.007 (.005)
Panel Sample	-.117 (.008)	-.062 (.007)	.021 (.008)	-.051 (.007)	-.004 (.008)	.004 (.008)	-.010 (.008)	.004 (.008)
Weekly Work Hours		x	x	x	x	x	x	x
Industries			x		x	x	x	x
Occupations				x	x	x	x	x
Full-Time Workers Only						x		
Top-Coded Earnings Doubled							x	
Top-Coded Earnings Deleted								x

Note: The table shows the coefficient on a dummy variable indicating nonprofit status from regressions where the dependent variable is the natural log of weekly earnings in year 1. Persons working in the public sector are excluded from the analysis. The first row shows results for the CPS-ORG cross-sectional sample (n=243,674); the second row refers to the panel with matched observations in years 1 and 2 (n=79,600). The equations also control for age and age squared, marital status (currently married and never married), race/ethnicity (black, Hispanic), education (high school graduate, some college, college graduate, post-graduate education), metropolitan residence and the survey year. Additional covariates are sometimes held constant, as detailed in the bottom panel, including weekly work hours, eight industry categories (hospitals, other health services, nursing/personal care facilities, social services, religious organizations, membership organizations, primary/secondary education, higher education), and ten occupation categories (health professionals, health technicians, health service workers, social workers, clergy/religious workers, educator/ librarians, secretaries, other administrative support, other managers, non-health service workers). Model (f) restricts the sample to persons working at least 35 hours per week. In column (g), workers with right-censored earnings are assumed to receive twice the top-coded amount, while column (h) excludes from these persons from the analysis. Standard errors are in parentheses.

Table 5:
Employment Shares and Wage Changes for Different Demographic Groups

Group	Share of Employment In:		Wage Change Between Years 1 and 2		
	All Jobs	Nonprofit Jobs	All	$P \rightarrow N$ Transitions	$N \rightarrow P$ Transitions
All	1.00	1.00	.063 [79,600]	.069 [1,874]	.082 [2,184]
Males	.542	.298	.061 [42,227]	.048 [523]	.104 [640]
Females	.458	.702	.066 [37,373]	.077 [1,351]	.073 [1,544]
No College	.443	.210	.059 [34,853]	.075 [493]	.085 [524]
Attended College	.557	.790	.067 [44,765]	.067 [1,381]	.081 [1,660]
Whites	.858	.868	.063 [69,527]	.057 [1,596]	.085 [1,866]
Blacks	.099	.099	.069 [6,443]	.116 [199]	.072 [222]
Hispanics	.090	.043	.065 [5,791]	-.017 [69]	.124 [87]

Note: The first two columns indicate the percentage of overall or nonprofit employment held by members of the specified group in year 1, calculated using CPS sampling weights. The last three columns show the average change in log wages, between years 1 and 2, for all sample members and for persons transitioning the nonprofit and for-profit sectors, over the two years. The sample consists of ORG respondents matched in years 1 and 2. Sampling weights are used in all calculations. Persons working in government jobs in either year are excluded. Sample sizes are shown in brackets.

Table 6:
Alternative Estimates of Nonprofit Earnings Differential Using Panel Sample

Group	**Wage Level in Year 1**	**Symmetric Wage Change**	**Asymmetric Wage Change**	
			$P \rightarrow N$	$N \rightarrow P$
All	-.004	-.010	-.004	-.015
	(.008)	(.007)	(.011)	(.010)
Males	-.030	-.034	-.025	-.040
	(.015)	(.013)	(.020)	(.018)
Females	.010	.001	.004	-.001
	(.009)	(.008)	(.012)	(.012)
No College	-.017	.012	.031	-.006
	(.016)	(.014)	(.020)	(.019)
Attended College	.002	-.017	-.015	-.018
	(.009)	(.018)	(.013)	(.012)
Whites	-.009	-.018	-.013	-.022
	(.009)	(.008)	(.011)	(.010)
Blacks	.057	.030	.045	.016
	(.025)	(.025)	(.036)	(.035)
Hispanics	.021	-.030	-.037	-.023
	(.040)	(.037)	(.056)	(.056)

Note: See notes on table 4. All specifications include controls for age and age squared, marital status, race/ethnicity, education, metropolitan residence and the survey year. The wage level regressions also control for weekly work hours and the eight industry and ten occupation categories. The first difference models control for changes (between years 1 and 2) in work hours and in the eight industries and ten occupations. The wage level models show the nonprofit differential from estimates of: $W_i = \alpha + X_i\beta + N_i\gamma + \delta Z_i + \varepsilon_i$, where N_i is a dummy variable indicating whether respondent i works for a nonprofit employer in year 1. The symmetric wage change equations take the form: $\Delta W_i = \alpha + X_i\beta + \Delta N_i\gamma + \Delta Z_i\delta + \Delta\varepsilon_i$; the table displays $\hat{\gamma}$. The asymmetric specifications are: $\Delta W_i = \alpha + X_i\beta + PN_i\gamma_1 + NP_i\gamma_2 + \Delta Z_i\delta + \Delta\varepsilon_i$, where PN_i, (NP_i) is a dummy variable indicating movement for-profit to nonprofit (nonprofit to for-profit) employment between years 1 and 2. In this case, the nonprofit differentials are estimated by $\hat{\gamma}_1$ and $-\hat{\gamma}_2$.

Table 7:
Nonprofit Earnings Differentials for Specific Industries

Sample/Procedure	Social Services	Hospitals	Other Health Services	Nursing/ Personal Care Facilities
Wage Levels in Year 1				
Cross-Sectional Sample	.015	.025	-.003	.044
	(.015)	(.008)	(.017)	(.016)
Panel Sample	.011	.047	-.039	.034
	(.028)	(.012)	(.028)	(.029)
Wage Changes for Industry Stayers				
Symmetric	.000	.002	.010	.059
	(.033)	(.012)	(.042)	(.032)
Asymmetric				
$P \rightarrow N$	.054	.009	.016	.092
	(.052)	(.019)	(.066)	(.049)
$N \rightarrow P$	-.045	-.005	.005	.031
	(.047)	(.018)	(.058)	(.046)

Notes: See notes on Tables 4 and 6. The wage levels are calculated for year 1. Wages changes refer to the panel data set for persons remaining in the same industry in years 1 and 2. All specifications include controls for age and age squared, marital status, race/ethnicity, education, metropolitan residence, the survey year, and levels or changes in weekly work hours and ten occupation categories. Sample sizes are 5,155, 13,085, 4,469, and 4,231 (1,541, 4,685, 1,487, and 1,322) for social services, hospitals, other health services, and nursing/personal care facilities in the wage level regressions for the cross-sectional (panel) sample. Corresponding sample sizes for industry stayers in the wage change equations are 1,051, 4,063, 848, and 971.

Table 8:
Nonprofit Earnings Differentials for Specific Occupations

Sample/Procedure	Managers (Not Health/ Education)	Health Professionals	Health Technicians/ Services	Adminis-trative Support
Wage Levels in Year 1				
Cross-Sectional Sample	-.074 (.014)	-.004 (.010)	.020 (.013)	.022 (.010)
Panel Sample	-.069 (.022)	.005 (.016)	.055 (.022)	.023 (.017)
Wage Changes for Occupation Stayers				
Symmetric	-.047 (.025)	.003 (.016)	.011 (.024)	.007 (.018)
Asymmetric				
$P \rightarrow N$	.004 (.041)	.013 (.025)	.013 (.036)	.031 (.028)
$N \rightarrow P$	-.079 (.032)	-.006 (.023)	.009 (.034)	-.013 (.026)

Notes: See notes on Tables 4 and 6. The wage levels are calculated for year 1. Wages changes refer to the panel data set for persons remaining in the same occupation in years 1 and 2. All specifications include controls for age and age squared, marital status, race/ethnicity, education, metropolitan residence, the survey year, and levels or changes in weekly work hours and eight industry categories. Sample sizes are 33,273, 10,813, 9,462, and 35,971 (11,532, 3,879, 3,015, and 12,158) for managers, health professionals, health technicians/service workers, and secretaries/administrative support workers. Corresponding sample sizes for occupation stayers in the wage change equations are 7,896, 3,223, 2,168, and 7,863.