Compensation in the Nonprofit Sector

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ABSTRACT

We investigate the determinants of pay in the nonprofit sector using data for 25–55 year olds from the 1994–88 Current Population Survey Outgoing Rotation Groups. Our results are consistent with the hypothesis that compensation is primarily determined in competitive markets without ''labor donations'' to nonprofit employers. One implication is that nonprofit workers receive virtually the same wages as observationally equivalent employees in similar positions with profit-seeking enterprises. We cannot rule out the possibility of nonprofit penalties or premiums for selected groups; however, the differentials are generally small and competition appears to play a dominant role in nonprofit wage setting.

I. Introduction

Nonprofit enterprises are an increasingly important part of the American economy. The number of nonprofit associations grew 54 percent between 1980 and 1997 (U.S. Census Bureau 1998, Table 1286) 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 substantial proportion of paid employment in some industries. Despite this growing significance,

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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 through a detailed analysis 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 or occupations with a substantial mix of nonprofit and for-profit employment.

Our results generally support the hypothesis 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 in nonprofits earn approximately the same amount as if they were employed in equivalent positions with profit-seeking firms. This is true even though the wages of nonprofit employees average 11 percent less than those of 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 offer relatively low pay but are probably also desirable places in which to work. Our evidence does not rule out the possibility of wage penalties or premiums for selected groups. However, the magnitudes of the differentials are generally small and do not detract from the dominant role that competition appears to play in setting nonprofit wages.

II. 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 (Easley and O’Hara 1983; Handy and Katz 1998) argue that nonprofits will therefore 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]

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 (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 because there is no reason to assume that government wages are determined by market forces.

3. There are other important differences between nonprofit and profit-seeking enterprises. For instance, nonprofits are frequently exempt from corporate income taxes and receive preferential treatment in state contract procurement processes (Frank and Salkever 1994).

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 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 (Frank 1996).[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 of nonprofit status (such as the bar on equity financing). This increased competitiveness implies downward pressure on wages.[9]

As mentioned, nonprofits may pay relatively low wages because 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 may 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 (clerical workers for example), implying that the estimated nonprofit differential may be little affected by the addition of occupation controls to an econometric model.[11]

4. 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 maximizes market share still has incentives to minimize labor costs.
7. 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.
9. Weisbrod (1988) believes that nonprofits arise when the government is unable to meet the demand for public goods (care for the medically indigent for instance). 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 benefits. Easley and O'Hara (1983) claim the nondistribution constraint may lead to relatively large pay reductions for nonprofit managers.

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

III. Previous Research

Previous studies of nonprofit compensation, summarized in Table 1, yield ambiguous results. Early examinations (Johnston and Rudney 1987; Shackett and Trapani 1987; Preston 1989) suggest a large nonprofit wage penalty but are hampered by the lack of information on the type of employer, requiring the imputation of nonprofit status.

Researchers focusing on narrowly defined industries obtain equivocal findings. Weisbrod (1983) shows that public interest lawyers earn 20 percent less than those in the private sector and believes this is due to heterogeneity in preferences, rather than in worker quality. However, using the same data, Goddeeris (1988) claims the lower wages reflect personal characteristics and that public interest attorneys earn no less than if 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 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. Mocan and Tekin (forthcoming) show that the size of the nonprofit premium in this industry varies considerably with the type of ownership, characteristics of the staff, and hours worked.[12]

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 three-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 occupa-

12. Roomkin and Weisbrod (1999) indicate that there is ambiguity even within industries. Focusing on six top managerial positions in hospitals, they find lower nonprofit 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).

Table 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 percent less per hour than for-profit employees; those in other nonprofits receive an insignificant 1.6 percent 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 percent (11 percent) less per hour than counterparts without (with) controls for industry and occupation. Larger differential for males (19 percent) than females (0 to 5 percent). Gaps shrink to an insignificant 0 to 4 percent 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 percent 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 lawyers (PIL) earn 37 percent 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 percent 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.

Table 1 (*continued*)

Study	Data	Results	Comments
Johnston and Rudney (1987)	1982 Census of service industries	The average annual earnings of nonprofit workers are 21.5 percent less than those employed in for-profit firms.	Hospitals, educational institutions, and religious organizations excluded. No controls for individual characteristics.
Leete (2001)	1990 Census, 5 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.
Mocan and Tekin (forthcoming)	398 child care centers in Calif., Colorado, Conn., and N.C., data from Spring 1993	Nonprofit childcare employees working full-time (part-time) receive a 6 (20) percent hourly wage premium and 8 (10) percent higher total compensation. Considerable variation by type of nonprofit and worker.	Extensive controls for human capital and center characteristics. Discrete factor methods used to control for unobserved heterogeneity. Standard errors not reported.

Preston (1988)	Abt Associates, 1976–77 National Day Care Center Supply Study	Nonprofit weekly wage premium of 5 to 10 percent for childcare workers in federally regulated daycare 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 ≈20 percent 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 percent 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 percent).

Table 1 (*continued*)

Study	Data	Results	Comments
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 percent 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.

tions 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.[13]

Most similar to the present research is DuMond's (1997) analysis of data from the 1994–1995 Current Population Survey Outgoing Rotation Groups (CPS-ORG). His cross-sectional regressions indicate a nonprofit wage penalty of between 6 and 11 percent. Conversely, fixed-effect estimates, exploiting data on individuals switching between for-profit and nonprofit employment, imply small (0 to 4 percent) and statistically insignificant earnings gaps. Several factors reduce our confidence in these findings. First, it is not clear how movements into or out of public sector are treated. Second, few respondents switch types of employment over the two-year period, decreasing the precision of the estimates. Third, endogenous mobility between sectors is not considered. Fourth, DuMond controls only for broad (one-digit) industries or occupations, which might inadequately account for differences in the job characteristics of nonprofit and for-profit employment. Each of these issues receives attention below.

IV. Data

We use data on 25 to 55 year olds from the 1994–98 Current Population Survey Outgoing Rotation Groups.[14] The CPS is a nationally representative survey of roughly 50,000 households. Individuals are interviewed for four 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 one and two.[15] Our cross-sectional sample includes data for Year 1; the longitudinal sample refers to individuals for whom information is available in both Years 1 and 2 (12 months apart). Not all persons can be matched across time. 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 (such as MacPherson and Hirsch 1995) who used slightly less stringent matching criteria. The matching procedures are detailed in Appendix A.

The dependent variable is the natural log of weekly wages on the ''main'' job.[16] 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

13. Several other potential methodological problems deserve mention. First, hourly wages are probably measured with considerable error. 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.

14. This age range avoids the special experiences of those making school-to-work or retirement transitions.

15. The outgoing rotation groups contain supplemental questions on weekly earnings and work hours not included in the regular monthly CPS.

16. 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.

the regression analysis (but included when examining how nonprofit employment is distributed across industries and occupations).[17]

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, other nonwhite), Hispanic origin, sex, metropolitan area residence, and the survey year. Some specifications add regressors for work hours and the industry and occupation of the ''main'' job.

As shown in Appendix Table B1, variable means are similar for the cross-sectional and panel samples. The main differences are that the second group 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.

V. Theoretical Framework

This section provides a model of nonprofit wage differentials in a competitive labor market without turnover costs and then considers mobility between nonprofit and for-profit jobs. For simplicity, we abstract from many important considerations (labor contracts, for example) influencing the adjustment to a new equilibrium.

Assume the utility (U) that individual i receives from working for employer j at time t depends on wages (W), possibly some additional benefit from holding a nonprofit position (N), and other compensating differentials related to job characteristics or nonwage payments (Z), according to the additive separable function:

$$(1) \quad U_{ijt} = W_{ijt} + cN_{ijt} + dZ_{ijt},$$

where the wage coefficient has been normalized to one.

The wages enterprise j is willing to pay are characterized by:

$$(2) \quad W_{ijt} = G_{it} + \gamma N_{ijt} + \delta Z_{ijt} + e_{ijt},$$

where G represents individual determinants affecting earnings across all employers (such as general human capital); γ is a market differential associated with nonprofit status; δ represents the effect of job characteristics or other compensating differentials; and e is a random variable indicating person-enterprise-time specific determinants of wages such as specific-human capital, worker-firm job matches, idiosyncratic employer payments, or macroeconomic shocks.

With competitive labor markets, no job will be systematically preferred to any other. This occurs if $\gamma = -c$, $\delta = -d$, and $E(e_{ijt} = 0)$, as can be seen by substituting Equation 2 into Equation 1, with the parameter restrictions, to obtain:

$$(1') \quad U_{ijt} = G_{it} + e_{ijt},$$

implying that that $E[U_{ijt}] = G_{it}$ for all j, with $E[.]$ the expectations operator.[18]

17. Persons employed in government jobs in *either* Years 1 or 2 are excluded from the panel sample. In preliminary work, we estimated models that included government workers and directly controlled for public sector employment. Doing so had little effect on the estimated nonprofit differentials.

18. This does not mean that individuals will be indifferent across jobs. For instance, specific-human capital or match quality can vary, implying the position held last period will generally be preferred over others.

Denoting the best available for-profit and nonprofit jobs with the subscripts p and n, utility is maximized by choosing nonprofit employment if

$$(3) \quad W_{ipt} + dZ_{ipt} < W_{int} + c + dZ_{int}$$

or

$$(3') \quad W_{ipt} - W_{int} < c + d(Z_{int} - Z_{ipt}),$$

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

Competition equalizes expected utility across the two types of jobs. Averaging across workers, this occurs when:

$$(4) \quad E[\bar{W}_{pt} - \bar{W}_{nt}] = c + dE[\bar{Z}_{nt} - \bar{Z}_{pt}],$$

where $\bar{W}$ and $\bar{Z}$ represent economy-wide average values of wages or job characteristics.[19] Equation 4 demonstrates that nonprofit enterprises will tend to pay less than profit-seeking firms if workers are willing to ''donate'' labor to them ($c > 0$) or they offer other positive compensating differentials ($\bar{Z}_{nt} - \bar{Z}_{pt} > 0$). Conversely, wage equalization suggests that there is no labor donation to nonprofits *and* no difference in average working conditions, or that the two effects exactly offset each other. Importantly, identical predicted earnings, after controlling for Z_{ijt}, suggests an absence of labor donations.

Next consider economically motivated switches from for-profit employment in period one to nonprofit jobs at time two. If there are no turnover or contracting costs, such moves occur if:

$$(5) \quad U_{ip1} - U_{in1} > 0 > U_{ip2} - U_{in2},$$

where U_{ipt} and U_{int} represent the utility to individual i of the best available for-profit and nonprofit jobs at time t. Rearranging these relationships, this mobility takes place if:

$$(6) \quad U_{in2} - U_{in1} > U_{in2} - U_{ip1} > U_{ip2} - U_{ip1}.^{20}$$

One requirement for this is that the utility of nonprofit employment must rise by a larger amount than the change in profit-seeking firms.

It is useful to distinguish two types of mobility. Utility-enhancing moves reflect unusually good opportunities that pull the worker into the nonprofit sector. In this case, $U_{in2} - U_{ip1} > 0$. Conversely, defensive transitions involve a loss of utility ($U_{in2} - U_{ip1} < 0$) but prevent a still larger reduction (for instance due to loss of specific human capital or a good job match following an involuntary layoff) that would occur if the individual remained in for-profit employment.[21] One implication,

19. Implicitly Equation 4 also requires that the pure utility effect of nonprofit employment (c) is the same across individuals and time periods. This assumption can easily be relaxed.
20. The corresponding condition for nonprofit to for-profit mobility is: $U_{ip2} - U_{ip1} > U_{ip2} - U_{in1} > U_{in2} - U_{in1}$.
21. The situation is analogous for moves from nonprofit to for-profit employment.

relevant in the econometric analysis below, is that $(e_{in2} - e_{ip1})$ will be positive for the first type of mobility and negative for the second.[22]

VI. Empirical Methods

Our empirical implementation begins by examining the cross-sectional wage equation:

$$(7) \quad W_{ijt} = \alpha_t + X_i\beta_t + N_{ijt}\gamma + \varepsilon_{ijt},$$

where W is the natural log of weekly wages, X is a vector of individual characteristics, N is a dummy variable indicating nonprofit employment, and ε is the regression disturbance.[23] The coefficient of primary interest, γ, shows the predicted nonprofit (log) wage differential controlling for personal but not job characteristics. Of greater interest are the results from:

$$(8) \quad W_{ijt} = \alpha_t + X_i\beta_t + N_{ijt}\gamma + \delta Z_{ijt} + \varepsilon_{ijt},$$

where Z is a vector of job characteristics (some combination of work hours, industries, and occupations). If Z adequately accounts for the heterogeneity in compensating differentials, γ will show the ''pure'' effect of nonprofit status, and a zero coefficient suggests that earnings are set competitively without labor donations.

One problem is that if the explanatory variables do not sufficiently control for the selection into nonprofit employment, $\text{cov}(N_{ijt}, \varepsilon_{ijt}) \neq 0$ and the least squares estimate $\hat{\gamma}$ is biased. For instance, a negative differential could occur because individuals with relatively low productivity disproportionately work in nonprofit jobs. With panel data, first-difference models will sometimes account for these sources of heterogeneity. For example, if $\varepsilon_{ijt} = f_i + e_{ijt}$, for f_i an individual fixed-effect and e_{ijt} an i.i.d. disturbance, the wage equation can be rewritten as:

$$(9) \quad W_{ijt} = \alpha_t + X_i\beta_t + N_{ijt}\gamma + \delta Z_{ijt} + f_i + e_{ijt}.$$

The change for person i occurring between Period 1 and 2 then is:

$$(10) \quad \Delta W_i = \alpha + X_i\beta + \Delta N_i\gamma + \Delta Z_i\delta + \Delta e_i,$$

where $\Delta W_i = W_{ij2} - W_{ij1}$, $\Delta N_i = N_{ij2} - N_{ij1}$, $\Delta Z_i = Z_{ij2} - Z_{ij1}$, $\alpha = \alpha_2 - \alpha_1$, $\beta = \beta_2 - \beta_1$, and $\Delta e_i = e_{ij2} - e_{ij1}$. Differencing away the fixed-effect has eliminated the bias due to all sources of time-invariant heterogeneity.

The symmetric wage change model described by Equation 10 still yields inconsistent estimates if $\text{cov}(\Delta N_i, \Delta e_i) \neq 0$, as with economically motivated turnover. However, a less constrained version of the first-difference model can bound the nonprofit differential even in this case. Define PN as a dummy variable equal to one for individuals switching from for-profit to nonprofit positions (abbreviated by $P \rightarrow N$) and

22. This can be seen by substituting (1′) into the middle-term in Equation 6 to show that $U_{in2} - U_{ip1}$ is positive (negative) if $e_{in2} - e_{ip1}$ is greater (less) than zero. Mobility for noneconomic reasons (for example, relocation of a spouse) is likely to be largely idiosyncratic, suggesting that $(E[e_{in2} - e_{ip1}] = 0)$.

23. The vector X does not include a time subscript because the individual characteristics we control for generally do not change over time for respondents in the age range analyzed.

NP as a dichotomous indicator for nonprofit to for-profit transitions (denoted by $N \rightarrow P$). Since $\Delta N = PN - NP$, Equation 10 can be rewritten as:

$$(11) \quad \Delta W_i = \alpha + X_i\beta + (PN_i - NP_i)\gamma + \Delta Z_i\delta + \Delta e_i.$$

Allowing PN and $-NP$ to have different coefficients yields the asymmetric wage change equation:

$$(12) \quad \Delta W_i = \alpha + X_i\beta + PN_i\gamma_1 + NP_i\gamma_2 + \Delta Z_i\delta + \Delta e_i,$$

where Equation (12) collapses to (11) if $\gamma_2 = -\gamma_1$, and $\hat{\gamma}_1$ and $-\hat{\gamma}_2$ provide alternative estimates of the nonprofit effect. Economically motivated turnover that is dominated by utility-increasing moves generally implies that cov(PN_i, Δe_i) and cov(NP_i, Δe_i) are positive. In this case, $\hat{\gamma}_1$ and $\hat{\gamma}_2$ are upward-biased and the nonprofit effects gaps estimated by $\hat{\gamma}_1(-\hat{\gamma}_2)$ are biased upward (downward). Conversely, cov(PN_i, Δe_i) and cov(NP_i, Δe_i) are negative for defensive turnover and the direction of bias is reversed.[24]

To illustrate, consider the case where nonprofit jobs, ceteris paribus, pay 5 percent less than for-profit employment. With exogenous turnover, $P \rightarrow N$ ($N \rightarrow P$) transitions will lead to a 5 percent fall (rise) in average earnings, compared to workers not changing sectors.[25] However, with utility-enhancing mobility where wage offers are two percentage points better than average, $P \rightarrow N$ ($N \rightarrow P$) switches result in a 3 (7) percent decrease (rise) in relative earnings, bounding the estimated nonprofit differential between 3 and 7 percent.[26]

There are at least two situations where these estimates may fail to accurately bound the nonprofit gap. First, since job characteristics are relatively crudely controlled for and nonwage compensation (such as fringe benefits) is not examined, the estimated wage effects might not adequately measure differences in total compensation. For example, labor donations to nonprofit employers could be reflected by less generous fringe benefits rather than reduced wages. In this case, the absence of an earnings effect following $P \rightarrow N$ or $N \rightarrow P$ transitions might conceal changes in nonwage compensation. Second, the estimates may be incorrect if mobility results from changes in (unobserved) individual characteristics. For instance, if deterioration (improvement) in health leads to $P \rightarrow N$ ($N \rightarrow P$) transitions, $\hat{\gamma}_1$ and $-\hat{\gamma}_2$ will overstate any nonprofit penalty by failing to attribute the lower nonprofit earnings to the negative productivity effects of poor health.[27] These issues receive further attention below.

24. For the symmetric wage change model to correctly estimate the nonprofit differential, $P \rightarrow N$ and $N \rightarrow P$ moves must occur with equal frequency. Selection into employment in Year 1 or 2 is not explicitly modeled because of the difficulty in obtaining plausible identifying restrictions. This will not cause bias as long as the selection process is similar for persons working in nonprofit and for-profit jobs.
25. This ignores general equilibrium effects that are likely to be small.
26. Similarly, with defensive turnover involving a two percentage point average decline in relative earnings, $P \rightarrow N$ ($N \rightarrow P$) mobility will lead to a 7 (3) percent point decrease (increase) in wages. The nonprofit penalty again will be bounded between 3 and 7 percent but with larger (smaller) differentials now predicted by $P \rightarrow N$ ($N \rightarrow P$) switches.
27. However, in most cases, differences in these time-varying factors will be reflected by cross-sectional wage differentials (for example, nonprofit workers would receive lower average wages due to their poor health).

VII. The Distribution of Nonprofit Employment

Tables 2 and 3 show how nonprofit employment is distributed across industries and occupations. The first two columns of each table display the composition of all jobs and of nonprofit positions. The third and fourth columns indicate the share of employment in the sector accounted for by nonprofit and profit-seeking organizations, with government as the residual category. The last column shows average weekly wages in the specified industry or occupation. For example, the third row of Table 2 demonstrates that religious organizations are responsible for 0.7 percent of all employment but 10.4 percent of 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—accounting for 85 percent of nonprofit employment versus 21 percent of all jobs (see Table 2). Fifty-five percent of nonprofit work is located in just three industries—hospitals, social services, and religious organizations. 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 positions. It is noteworthy that, except for education and hospitals, industries with high nonprofit shares pay below average wages. All of them are also widely viewed as engaging in ''socially desirable'' activities.

Nonprofit work is more dispersed across occupations. Nevertheless, ten of them—clergy/religion, health professionals, social work, health technicians, health services, educators/librarians, secretaries, other administrative support, managers, and nonhealth services—account for 86 percent of nonprofit employment versus 54 percent of all jobs (see Table 3).[28] 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 research holding constant one-digit industries and occupations (Preston 1989; DuMond 1997), or with Leete (2001), who includes detailed controls for up to 20,000 industry-occupation interactions. 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

28. Health managers are included in the health professional category and education managers in the educator/librarian occupation group. ''Other managers'' therefore refer to those outside these two fields.

Table 2
Industry Composition of Nonprofit Employment

Industry	Industry Percent of Employment		Percent of Industry Employment in		Average Weekly Wage ($)
	Overall (a)	Nonprofits (b)	Nonprofits (c)	For-profits (d)	(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 (700–12)	6.9	3.5	2.9	93.9	662
Personal/business services (721–91)	8.3	1.8	1.2	97.8	497
Transport/communication/utilities (400–72)	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 (900–32)	5.9	0.0	0.0	0.0	682

Note: Data are from the 1994–98 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 3
Occupation Composition of Nonprofit Employment

	Occupation Percent of Employment		Percent of Occupation Employment in		
Industry	Overall (a)	Nonprofits (b)	Nonprofits (c)	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 (203–8)	1.6	5.0	17.5	70.5	499
Health service worker (445–7)	2.0	4.4	12.7	72.0	309
Educators/librarians (14, 113–65)	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 administrative support (303–9, 314–89)	12.9	11.0	4.9	70.8	444
Nonhealth service worker (403–44, 448–69)	9.8	7.3	4.3	68.2	358
Farming/fishing/forestry (473–99)	1.4	0.6	2.5	88.5	342
Nonhealth technician (209–35)	2.2	0.9	2.3	82.6	724
Handlers/cleaners/laborers (863–89)	3.5	0.8	1.2	92.4	394
Production/craft/repair (503–799)	18.3	3.4	1.1	94.1	546
Sales (243–85)	9.9	1.8	1.0	97.6	582
Transportation (803–59)	4.4	0.7	0.9	88.1	581

Note: See note on Table 2. The numbers in parentheses refer to three digit census occupations.

job characteristics, while focusing on within-industry (or occupation) variations in the type of employer.

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 holding for-profit jobs.[29] An identical 18 percent error rate is obtained by assuming that 100 percent of employment in religious organization is nonprofit, versus the 85.1 percent reported by CPS respondents.[30] Such classification errors may 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.[31]

VIII. Cross-Sectional Wage Differentials

This section examines cross-sectional wage differentials. Table 4 displays mean weekly earnings for selected industries and occupations. Although nonprofit workers average 3 percent less per week than those in profit-seeking firms, 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 nonhospital 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 (such as social services, religion, and nursing/personal care). There is no corresponding pattern of high nonprofit pay within specific occupations.

Disparities in earnings may reflect individual heterogeneity, rather than differences in nonprofit wage-setting. As shown in Appendix Table B1, nonprofit workers are slightly older and considerably more educated than their counterparts but they also work fewer hours and are more often female. A careful econometric analysis can help disentangle these effects from other sources of pay differentials.

Table 5 displays the coefficient on nonprofit status from cross-sectional estimates

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

30. 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 2 (29.9 percent) which covers the 1994–98 period. However, the hospital share of nonprofit jobs has been trending sharply down over time.

31. 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 $\rho = .067$ and $\phi = .18$, the observed nonprofit differential will therefore be 98.7 percent as large as the actual gap.

Table 4
Average Weekly Wages by Sector of Employment

Industry/Occupation	Weekly Earnings ($) For-Profit	Weekly Earnings ($) Nonprofit	Nonprofit Differential (Percent)
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)
Nonhealth 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 1994–98 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.

of Equations 7 and 8. The dependent variable is the log of weekly wages in Year 1.[32] The first row shows results for the full 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 analysis. 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 vari-

32. Similar results are obtained for Year 2.

Table 5
Econometric Estimates of the Nonprofit Differential in Weekly Wages

Sample/Additional Controls	(a)	(b)	(c)	(d)	(e)	(f)	(g)	(h)
Cross-sectional sample	−0.119 (0.005)	−0.059 (0.004)	0.013 (0.005)	−0.056 (0.004)	−0.013 (0.005)	−0.009 (0.005)	−0.019 (0.005)	−0.007 (0.005)
Panel sample	−0.117 (0.008)	−0.062 (0.007)	0.021 (0.008)	−0.051 (0.007)	−0.004 (0.008)	0.004 (0.008)	−0.010 (0.008)	0.004 (0.008)
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 one and two (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, nonhealth 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 these persons from the analysis. Standard errors are in parentheses.

ables for eight industry and ten occupation categories are also frequently included, as detailed in the bottom six rows of the table.

Consistent with earlier research (Preston 1989 for example), nonprofit employment is associated with an 11 percent wage penalty after holding constant 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).[33] Unless noted, the regressions in the remainder of this analysis include covariates for work hours, industries, and occupations (Specification E). When this is done, nonprofit employees are predicted to earn virtually the same wages as their for-profit counterparts. This result is consistent with earnings being determined in competitive labor

33. 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. The results are similar when these persons are deleted from the sample.

markets, where nonprofit jobs pay less because they require fewer hours and are located in industries offering positive compensating differentials. Conversely, there is no evidence of labor donations based specifically on nonprofit status.

The remainder of Table 5 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.[34] 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.[35] 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 statistically insignificant.

IX. Wage Changes

The first two columns of Table 6 detail employment shares in all jobs and in nonprofit positions for subgroups stratified by sex, education, and race/ethnicity. Most striking is the disproportionate representation of women and highly educated individuals in nonprofits. Females represent 46 percent of the labor force but hold over 70 percent of nonprofit positions; 56 percent of the sample is college educated but these individuals account for 79 percent of nonprofit employment.

The last three columns of the table document average changes in log wages occurring between Years 1 and 2. Earnings growth is somewhat faster for persons switching between nonprofit and for-profit jobs than for those remaining in the same sector, suggesting the importance of utility-increasing mobility. More noteworthy is the relatively similar growth for persons making $P \rightarrow N$ and $N \rightarrow P$ transitions. Weekly earnings rise 0.082 log points (8.5 percent) for the former group versus 0.069 log points (7.1 percent) for the latter, which again hints at the small size of any nonprofit differential. There is some variation across demographic categories. In particular, faster wage growth after $N \rightarrow P$ than $P \rightarrow N$ moves raises the possibility of a larger nonprofit penalty for men. Small numbers of minorities in nonprofit employment imply that the results for these groups should be interpreted with caution.

Table 7 summarizes alternative econometric estimates of the nonprofit differential using data for the panel sample. The first column refers to cross-sectional regressions of wage levels in year one; the second indicates results from the symmetric wage

34. In the cross-sectional sample, 20 percent of nonprofit workers are employed fewer than 35 hours per week, versus 10 percent of those in profit-seeking enterprises.

35. Weekly earnings are top-coded at $1,920 in 1994–97 and $2,880 in 1998.

Table 6
Employment Shares and Wage Changes for Different Demographic Groups

	Share of Employment In		Wage Change Between Years 1 and 2		
Group	All Jobs	Nonprofit Jobs	All	$P \rightarrow N$ Transitions	$N \rightarrow P$ Transitions
All	1.00	1.00	0.063 [79,600]	0.069 [1,874]	0.082 [2,184]
Males	0.542	0.298	0.061 [42,227]	0.048 [523]	0.104 [640]
Females	0.458	0.702	0.066 [37,373]	0.077 [1,351]	0.073 [1,544]
No college	0.443	0.210	0.059 [34,853]	0.075 [493]	0.085 [524]
Attended college	0.557	0.790	0.067 [44,765]	0.067 [1,381]	0.081 [1,660]
Whites	0.858	0.868	0.063 [69,527]	0.057 [1,596]	0.085 [1,866]
Blacks	0.099	0.099	0.069 [6,443]	0.116 [199]	0.072 [222]
Hispanics	0.090	0.043	0.065 [5,791]	−0.017 [69]	0.124 [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.

change model (Equation 10); the third and fourth columns display $\hat{\gamma}_1$ and $-\hat{\gamma}_2$ from the asymmetric earnings growth model (Equation 12). 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 variables.

The full sample results provide further evidence that nonprofit workers receive virtually the same pay as their for-profit counterparts with equivalent individual and job characteristics. The predicted nonprofit differential is −0.4 percent in the cross-sectional regression, −1.0 percent in the symmetric earnings growth model, and bounded between −0.4 and −1.5 percent in the asymmetric wage change equation. None of these parameter estimates differ significantly from zero. There are some disparities across demographic groups. Most importantly, men receive roughly a 3 percent wage penalty for nonprofit work, possibly explaining why they hold these

Table 7
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	−0.004	−0.010	−0.004	−0.015
	(0.008)	(0.007)	(0.011)	(0.010)
Males	−0.030	−0.034	−0.025	−0.040
	(0.015)	(0.013)	(0.020)	(0.018)
Females	0.010	0.001	0.004	−0.001
	(0.009)	(0.008)	(0.012)	(0.012)
No college	−0.017	0.012	0.031	−0.006
	(0.016)	(0.014)	(0.020)	(0.019)
Attended college	0.002	−0.017	−0.015	−0.018
	(0.009)	(0.018)	(0.013)	(0.012)
Whites	−0.009	−0.018	−0.013	−0.022
	(0.009)	(0.008)	(0.011)	(0.010)
Blacks	0.057	0.030	0.045	0.016
	(0.025)	(0.025)	(0.036)	(0.035)
Hispanics	0.021	−0.030	−0.037	−0.023
	(0.040)	(0.037)	(0.056)	(0.056)

Note: See notes on Table 5. 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: $\Delta W_i = \alpha + X_i\beta + \Delta N_i\gamma + \Delta Z_i\delta + \Delta e_i$; the table displays $\hat{\gamma}$. The asymmetric specifications are: $\Delta W_i = \alpha + X_i\beta + PN_i\gamma_1 + NP_i\gamma_2 + \Delta Z_i\delta + \Delta e_i$, 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 $-\hat{\gamma}_2$.

jobs relatively infrequently (Preston 1990). The data also suggest an imprecisely measured 2 to 6 percent nonprofit premiums for blacks. Even noting this heterogeneity, the evidence suggests that the scope for labor donations to nonprofit employers is generally small and that competitive labor markets may play a dominant role in setting wages.[36]

36. We also estimated models using data from the Displaced Worker Supplements to the 1994, 1996, and 1998 Current Population Surveys for persons losing jobs due to plant closure, slack work, or position/shift abolished. The results again indicate the virtual absence of an overall nonprofit wage differential, with estimates from the asymmetric wage change models pointing to an anticipated dominant role of defensive turnover for this group.

X. Differentials Within Industries and Occupations

We next examine nonprofit differentials within specific industries or occupations. Table 8 summarizes the results 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 panel sample is limited to persons in the specified industry in both years one and two. The latter restriction avoids confounding the effect of industry mobility with that of changes in nonprofit status.

The findings are again generally consistent with the hypothesis of competitive wage setting without labor donations to nonprofits. 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. These higher earnings are mostly due to transferable individual characteristics, however, as evidenced by the statistically insignificant 0 to 1 percent earnings differentials obtained in the

Table 8
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	0.015	0.025	−0.003	0.044
	(0.015)	(0.008)	(0.017)	(0.016)
Panel sample	0.011	0.047	−0.039	0.034
	(0.028)	(0.012)	(0.028)	(0.029)
Wage changes for industry stayers				
Symmetric	0.000	0.002	0.010	0.059
	(0.033)	(0.012)	(0.042)	(0.032)
Asymmetric				
$P \rightarrow N$	0.054	0.009	0.016	0.092
	(0.052)	(0.019)	(0.066)	(0.049)
$N \rightarrow P$	−0.045	−0.005	0.005	0.031
	(0.047)	(0.018)	(0.058)	(0.046)

Notes: See notes on Tables 5 and 7. The wage levels are calculated for Year 1. Wage 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 9
Nonprofit Earnings Differentials for Specific Occupations

Sample/Procedure	Managers (Not Health/ Education)	Health Professionals	Health Technicians/ Services	Administrative Support
Wage levels in Year 1				
Cross-sectional sample	−0.074	−0.004	0.020	0.022
	(0.014)	(0.010)	(0.013)	(0.010)
Panel sample	−0.069	0.005	0.055	0.023
	(0.022)	(0.016)	(0.022)	(0.017)
Wage changes for occupation stayers				
Symmetric	−0.047	0.003	0.011	0.007
	(0.025)	(0.016)	(0.024)	(0.018)
Asymmetric				
$P \rightarrow N$	0.004	0.013	0.013	0.031
	(0.041)	(0.025)	(0.036)	(0.028)
$N \rightarrow P$	−0.079	−0.006	0.009	−0.013
	(0.032)	(0.023)	(0.034)	(0.026)

Notes: See notes on Tables 5 and 7. The wage levels are calculated for Year 1. Wage 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.

symmetric wage change equations for the social service, hospital, and other health service industries. One exception is that nonprofit workers in nursing/personal care facilities receive a pay premium estimated at between 3 and 10 percent, raising the possibility of rent-sharing. The bounds on the nonprofit differential obtained from the asymmetric first-difference model are also reasonably wide for Social Service workers, although not significantly different from zero.

Table 9 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 sizeable participation by profit-seeking firms.[37] The regression specifications are identical to Table 8, 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.

The results suggest small or nonexistent nonprofit differentials for the three nonmanagerial occupations, generally ranging between −1 and 3 percent and usually

37. These categories correspond to those in Table 3, except that health technicians and service personnel have been combined into a single group, as have secretaries and other administrative support workers.

statistically insignificant, providing further support for competitive wage-setting without labor donations. However, nonprofit managers earn about 7 percent less than their for-profit peers, controlling for observables, with the wage-change regressions bounding the penalty between 0 and 8 percent.[38] This indicates some scope for labor donations by managers.

XI. Discussion

Our econometric results suggest that compensation in the nonprofit sector is primarily determined in competitive labor markets, without explicit labor donations based upon nonprofit status. Weekly wages average 11 percent less in nonprofit than for-profit jobs, holding constant worker characteristics, but this is almost entirely due to shorter hours and the concentration of these positions in relatively low-paying industries. As a result, nonprofit employees earn virtually the same amount as observationally equivalent individuals in similar positions with profit-seeking enterprises. The wage growth of persons making $P \rightarrow N$ or $N \rightarrow P$ transitions are also generally similar, further hinting at the small size of any overall nonprofit differential.

Why are nonprofit jobs disproportionately located in low-paying industries? The most likely reason is that these sectors perform ''socially desirable'' activities (such as helping the sick or teaching children), so that employees are willing to accept decreased compensation. This represents a variation of the labor donation hypothesis. However, the key distinction is that the reduced wages reflect the specific goods and services provided, rather than because of the nonprofit status of the employer. 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.[39] However, for-profit jobs would then be rationed, which is at odds with the evidence that highly educated workers (who presumably have the most options) disproportionately select nonprofit employment.

Our findings are subject to two caveats. First, (unobserved) time-varying individual factors might be correlated with movements into or out of the nonprofit sector.

38. Managers in the health or education industries were excluded from the manager category to maintain consistency with the groupings used in Table 3. 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 tested whether the nonprofit penalty for managerial employment explains the relatively low earnings of male nonprofit workers. The nonprofit penalty was estimated to be larger for male than female managers in all specifications examined, providing no support for this possibility.

39. 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 percent in 1992 (U.S. Census Bureau 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 (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.

The first-difference models only fully control for fixed-characteristics and so might not adequately account for these changes. However, the resulting bias is likely to be fairly small since such heterogeneity would generally also lead to a substantial cross-sectional differential, which is not observed in the data. Second, this analysis focuses on earnings and so could miss other components of compensation. This represents a useful topic for future research, particularly since limited available evidence hints at the possibility of larger nonwage benefits in nonprofit than for-profit jobs.[40] This raises the possibility that our results modestly overstate (underestimate) the size of any nonprofit penalty (premium).

Notwithstanding these qualifications, the findings largely support the hypothesis of competitive wage setting, without labor donations explicitly linked to nonprofit status. However, some deviations are possible. For instance, we uncover a small (2 to 4 percent) nonprofit wage penalty for males and a less precisely estimated 0 to 8 percent disadvantage for managers. Conversely, employment in nursing/personal care facilities is associated with a 3 to 10 percent earnings premium. These exceptions suggest 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.

Appendix A

Construction of Longitudinal Sample from the CPS-ORG Files

Individuals from the same month in consecutive years of the Current Population Survey can potentially be identified using available information on household ID codes and record line numbers. Because different states sometimes use the same household ID, state (FIPS) codes are also needed to uniquely identify the household. Two restrictions should be noted. First, the household ID represents a permanent residence and so does not follow families that relocate. Second, 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.

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 nonself-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 of 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

40. DuMond (1997) finds that pension and health insurance coverage are relatively high and displacement rates relatively low in nonprofit 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. Mocan and Tekin (forthcoming) find relatively high levels of nonwage compensation for childcare workers employed full-time in nonprofit enterprises, but not for corresponding part-time workers.

code, and record line. This reduced the sample to 236,122 observations. To further ensure 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, and 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 the 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 B1
Variable Means by Sector of Employment for Cross-Sectional and Panel Samples

	Cross-Sectional Sample		Panel Sample	
Variable	For-Profit	Nonprofit	For-Profit	Nonprofit
Weekly earnings/hours				
Earnings ($)	573	557	596	572
Work hours	40.9	38.2	41.2	38.6
Education				
High school dropout	0.116	0.039	0.103	0.033
High school graduate	0.363	0.189	0.357	0.173
Some college	0.283	0.290	0.292	0.304
College graduate	0.178	0.273	0.187	0.276
Graduate degree	0.060	0.210	0.060	0.214
Marital status				
Currently married	0.649	0.649	0.699	0.685
Never married	0.190	0.199	0.161	0.176
Race/ethnicity				
Black	0.110	0.110	0.100	0.096
Other nonwhite	0.046	0.035	0.043	0.036
Hispanic	0.107	0.045	0.094	0.044
Other characteristics				
Age (years)	38.2	39.8	38.8	40.3
Male	0.559	0.306	0.561	0.302
Metropolitan residence	0.780	0.787	0.825	0.834
Sample size	227,018	18,203	74,047	6,143

Note: See notes on Tables 2 and 3. All variable means are computed using CPS sampling weights. The cross-section includes respondents in Year 1 of the 1994–98 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.

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