Industrial & Labor Relations Review

Volume 59, *Issue* 2 2006 *Article* 3

Risk Compensation for Hospital Workers: Evidence from Relative Wages of Janitors

Keith A. Bender* Hosne A. Mridha†

James Peoples‡

*University of Wisconsin,

†North Carolina Central University,

‡University of Wisconsin,

Risk Compensation for Hospital Workers: Evidence from Relative Wages of Janitors

Keith A. Bender, Hosne A. Mridha, and James Peoples

Abstract

This study, using individual worker data on janitors taken from the 1985–2001 Current Population Survey and industry injury data from the Bureau of Labor Statistics, examines whether compensation for working in a high-risk work environment contributes to the relatively high wage rates of hospital janitors. The authors find that when the analysis corrects for risk endogeneity (workers' tendency to sort themselves according to their tolerance for workplace risk), risk compensation increases wage rates by 13.4% for union hospital janitors and is a major source of their wage advantage over nonunion janitors in other industries. Since this risk compensation is only available to union workers, the authors interpret this result as the effect of the union voice mechanism.

KEYWORDS: Risk Compensation for Hospital Workers, Janitors

RISK COMPENSATION FOR HOSPITAL WORKERS: EVIDENCE FROM RELATIVE WAGES OF JANITORS

KEITH A. BENDER, HOSNE A. MRIDHA, and JAMES PEOPLES*

This study, using individual worker data on janitors taken from the 1985–2001 Current Population Survey and industry injury data from the Bureau of Labor Statistics, examines whether compensation for working in a high-risk work environment contributes to the relatively high wage rates of hospital janitors. The authors find that when the analysis corrects for risk endogeneity (workers' tendency to sort themselves according to their tolerance for workplace risk), risk compensation increases wage rates by 13.4% for union hospital janitors and is a major source of their wage advantage over nonunion janitors in other industries. Since this risk compensation is only available to union workers, the authors interpret this result as the effect of the union voice mechanism.

Exploring the question of why one group of workers earns more than another has a long history in economics. The empirical approach used by many studies is to estimate relative wage rates of broad classes of workers—male and female, white and black, union and nonunion, public and private. Less common are studies that examine earnings differentials in a single occupation, which allows researchers to control for particular characteristics of that occupation. This empirical approach may be key if workers in a particular occupation encounter job conditions, such as health risk, that vary across industries.

From this point of view, a potentially interesting occupation to study is that of janitors. Although job conditions may vary across industries for this occupation, job responsibilities are relatively homogeneous, and labor demand and supply should be relatively elastic. Yet, as shown below, there are substantial earnings premia for unionized janitors and janitors who work in hospitals. The question then becomes, why do these premia exist? Are they due to differences in bargaining power, differences in

*Keith A. Bender is Associate Professor in the Department of Economics and in the Graduate Program in Human Resources and Labor Relations at the University of Wisconsin–Milwaukee. Hosne A. Mridha is Adjunct Assistant Professor in the School of Business at North Carolina Central University. James Peoples is Professor of Economics in the Department of Economics at the University of Wisconsin–Milwaukee. The authors are grateful for comments received from participants at the Centre for European Labour Market Research Workshop at the University of Aberdeen, Scotland, and the NEA/AEA Joint Summer Program Conference at the University of Colorado–Denver. They also thank Scott Adams, Daniel Hamermesh, John Heywood, and Robert Sandy for their valuable suggestions.

Stata programs used in analyzing the data are available from the corresponding author: Keith A. Bender, Department of Economics, P.O. Box 413, University of Wisconsin–Milwaukee, Milwaukee, WI 53201.

Industrial and Labor Relations Review, Vol. 59, No. 2 (January 2006).
0019-7939/00/5902 $01.00

productivity, or possibly differences in health risks, particularly in the case of janitors employed by hospitals? Given the high and rapidly increasing cost of health care, understanding the nature of the hospital premium is particularly important. If the cause of the premium is market inefficiencies, the high wages paid to hospital janitors are not economically justified. On the other hand, if the high relative wages are in compensation for higher risk, they may be efficient.

Therefore, this study compares the wage structure of hospital janitors with that of janitors outside the health profession.[1] We focus on two issues. The first is the influence on wage rates of unionization, of health risk, and, importantly, of the interaction of those two factors. The interaction is key since, *a priori*, it is not obvious whether the market sets an accurate price for workplace risk, given the possibility of information asymmetry among employers and employees (Sandy and Elliott 1996). Such an asymmetry may occur because employers can maintain a larger pool of workers at lower wages by understating the workplace risk faced by workers. The lack of accurate information is less likely for union workers than for nonunion workers, because a key objective of unions is to keep members well informed about working conditions.

The second issue is the potential for endogeneity in the relationship between workplace risk and wages. This endogeneity arises from the fact that different workers have different attitudes toward workplace risk and may sort themselves according to their tolerance for risk. Not controlling for this endogeneity may lead to inaccurate estimates of the compensating wage differential for risk.

[1]Past studies examining hospital wage premia have also focused on wage results for a single occupation—nursing (Link 1988; Schumacher and Hirsch 1997; and Leherer et al. 1991). The industry specificity of job responsibilities for nurses, however, makes it difficult to identify suitable comparison groups when examining relative wages across industries.

Table 1. Mean Hourly Earnings in 1983–1984 Dollars. (Sample Populations in Parentheses)

Industry Sector	*Union*	*Nonunion*
Hospital	$9.37 (268)	$8.02 (869)
Other Industries	$9.94 (8,442)	$6.93 (38,084)

Source: 1985–2001 Current Population Survey–Outgoing Rotation Group files.

Wage Determination of Hospital Janitors

Information in Table 1 reveals evidence that wage rates vary greatly depending on union and hospital status. Union hospital janitors received a mean wage premium of 35.2% over nonunion janitors in other industries for the 1985–2001 observation period. This hospital premium closely resembles the union premium for non-hospital janitors, which is 43.4%. The wage premium for union janitors that is attributable to union negotiation advantage, however, possibly differs for individuals employed in the hospital sector and in other industries. High wage rates for union janitors could be partly due to gains from employment in a high-wage service sector, since nonunion hospital janitors also receive high wages. Table 1 reveals that nonunion hospital janitors receive 15.7% higher wages than nonunion, non-hospital workers.

The premium could have any of several sources. One possible source is rent sharing, derived from the relatively inelastic demand for hospital services. Although the highly elastic labor supply curve for janitors makes high wages unlikely, any increases in wages (particularly if brought about by collective bargaining) could be passed on relatively easily to customers of health care services. In addition, many hospitals are public, and the strict profit maximization objectives of private hospitals may not closely correspond to managerial goals in the public sector. Hence,

public hospitals might "pass through" higher pay to customers more readily than their private counterparts. If, furthermore, rent sharing is the cause of the premium, the high relative wage rates will be associated with an inefficiency in the employment of janitors: this rent could be used either to employ more workers in an effort to enhance service quality or to lower operation costs by bringing wages down to competitive rates.

How likely is it that janitors can be effective in collective bargaining? Janitors, in general, provide a ubiquitous service that requires low physical and human capital investment. These occupational characteristics, together with a competitive labor market, suggest that janitors should receive relatively low wages. Despite the negotiation disadvantages associated with representing janitors, such as a likely elastic labor demand curve, the Service Employees International Union (SEIU), which is the major union representative for janitors, has negotiated bargaining contracts that provide substantial wage gains and fringe benefits for its members.[2] The SEIU's success is due in part to its monopoly control over the supply of janitors in several service sectors, such as health care (Erickson et al. 2001). The SEIU's labor market dominance is attributable to its strategy of organizing the few large janitorial firms that clean large buildings and complexes, resulting in a low per-worker cost for organizing workers. For example, in the mid-1990s American Building Maintenance (ABM) and Industrial Service Systems (ISS) employed over 25% of all janitors who cleaned large buildings in the Los Angeles metropolitan area (Erickson et al. 2001). Since hospitals are relatively large employers with high cleanliness standards, they will have relatively large janitorial staffs compared to other industries, meaning that the per-worker cost for unionization may be low in hospitals.

An alternative hypothesis is that the hospital wage premium for janitors reflects risk compensation.[3] Since the job tasks of hospital janitors can include handling hazardous materials and contact with airborne pathogens, hospitals may have to pay janitors an economically efficient compensating differential for industry-related health risk.

There is an absence of research, however, testing the wage effect of workplace risk for this occupation (and, indeed, for the hospital industry).[4] A more common worker sample used in past research on risk compensation is all non-white-collar workers.[5] These studies generally report strong evidence that blue-collar workers receive compensation for workplace risk (Garen 1988; Kniesner and Leeth 1991; Sandy and Elliott 1996; Gunderson and Hyatt 2001).

A complicating factor was identified by Sandy and Elliott (1996), who suggested that the market may fail to accurately account for workplace risk as a wage determinant because firms have an incentive to understate the health risk faced by their employees. In this case, Gunderson and Hyatt (2001) observed that paying wage

[2]Erickson et al. (2001) reported that janitorial members of the SEIU in Los Angeles received an average total compensation of $12.00 an hour, compared to $4.00 for nonunion janitors.

[3]A third possibility comes from the human capital model, which implies that those with more skills or abilities will be paid more. We can control for some differences in productivity, as proxied by education, but controlling for unobserved skill is more difficult, since we do not have a panel of workers. Garen (1988) offered a method for estimating returns to unobserved ability by including the residuals of the first stage regression from the endogeneity correction, but given our formulation of the correction, we are unable to make a similar assumption that these residuals are correlated with unobserved janitorial ability. Sandy and Elliott (1996) and Sandy et al. (2001) also used this approach.

[4]While past research indicates that hospital premia mainly reflect a compensating differential, these studies do not test for the influence of risk compensation (Schumacher and Hirsch 1997).

[5]Past research excludes white-collar occupations from their sample population because these workers usually face working conditions markedly different from those of their blue-collar counterparts (Garen 1988; Gunderson and Hyatt 2001).

rates that do not fully compensate for risk can lead to managers employing an inefficiently large number of workers. One hypothesis from union-voice theory is that unions overcome this potential market failure by providing a mechanism through which members can express a demand for risk compensation (Sandy and Elliott 1996). The negotiation advantage provided by a union voice is especially important to janitors, because nonunion janitors who independently complain about workplace risk are subject to a high probability of job replacement due to the highly elastic labor supply in their occupation.[6]

A larger risk premium for union members is not certain *a priori*.[7] Alternative theory indicates unions might negotiate work assignments that allocate high-risk tasks to low-wage, less senior members, resulting in a pattern of low pay in high-risk jobs. However, it is unclear how unions could sort high-risk from low-risk janitorial assignments in hospitals, where health risk is commonly associated with the entire worksite and is not necessarily limited to job tasks. In such a setting, it seems more likely that unions would concentrate on reducing risks or exerting pressure for compensating differentials (or both).[8] Union voice may thus be an important source of hospital janitors' relatively high wage rates compared to those of nonunion hospital janitors. Large unions such as the SEIU seem especially likely to use their monopoly control over the supply of janitors to effectively negotiate the high wage rates required to compensate members working for high-risk firms.

If these theories are correct, testing the existence of risk compensation for union and nonunion hospital janitors may help identify the role unions play in correcting market imperfections. At a minimum, unions are in a position to inform members of the industry work force about the risk associated with employment in the hospital service sector.

Methodology for Identifying Compensation for Risk

We use several empirical specifications of wage equations to examine hospital and union wage premia among janitors. These specifications vary in the degree to which they account for the possibility of risk compensation, including differences in risk compensation by hospital employment and union status. The initial wage equation is specified as

$$(1)\qquad \mathrm{Ln}(w_j) = \alpha_0 + \alpha_1 \mathbf{X}_j + \alpha_2 \mathrm{NONUNION}_j * \mathrm{HOSP}_j + \alpha_3 \mathrm{UNION}_j * \mathrm{HOSP}_j + \alpha_4 \mathrm{UNION}_j * \mathrm{NONHOSP}_j + e_j,$$

where j indexes individual janitors and w_j is the real hourly earnings of the j^{th} janitor. The vector $\mathbf{X}$ includes the standard demographic and human capital variables used to identify measurable worker characteristics.[9] The mutually exclusive dummy variables NONUNION*HOSP, UNION*HOSP, and

[6]Another hypothesis regarding the effect of union voice is that increased communication makes it possible for workers to more effectively work in a risky environment or develop ways to reduce exposure to risk. This could lead to union workers being more productive than nonunion workers in these environments (see Freeman and Medoff 1984 for a discussion of union voice effects on productivity in more general settings) and, therefore, receiving higher wages. Unfortunately, there is no way to disentangle these two effects of union voice on wages.

[7]Sandy and Elliott (1996) reported that unionization is associated with a lower risk premium than that received by nonunion workers. These authors emphasized, however, that their findings on union risk compensation must be provisional, in part because of collinearity of unionization with other wage determinants included in their wage equation. Other work by these authors (Elliott and Sandy 1998; Sandy et al. 2001) investigated how mismeasured data might affect estimates of the compensating differential.

[8]It is likely that exposure to risk is nonexclusive, that is, it is difficult for hospital janitors to avoid risk, given the possibility of airborne pathogens and working with ill patients. In this respect, risk is often modeled as a "public good." Duncan and Stafford (1980) and Hamermesh and Wolfe (1990) were among the first to view risk in this way.

[9]Yearly time dummies are included to correct for possible changes in the wage determination process over time.

UNION*NONHOSP are of primary interest. NONUNION*HOSP is a dummy equaling one if a nonunion janitor is employed by a hospital. Since the excluded group is nonunion, non-hospital janitors, its estimated coefficient measures the hospital log wage differential for nonunion janitors. UNION*HOSP is a dummy equaling one if a hospital janitor belongs to a union, and its estimated coefficient represents the log wage premium for unionized hospital janitors over their nonunion counterparts in other business sectors. Last, UNION*NONHOSP is a dummy equaling one if a non-hospital janitor belongs to a union. Its estimated coefficient is a measure of the union log wage differential for non-hospital janitors.

Estimates of the hospital and union wage differentials from equation (1) are likely to be biased if they reflect a compensation differential for workplace risk. Such bias is of considerable concern when we examine earnings of hospital workers, because workplace risk is a major disamenity associated with patient care. The effect on wages of the compensating differential for risk is captured by three variables—RISK for industry *i*, an interaction of RISK and UNION, and a final interaction of RISK and HOSPITAL. The new specification is

$$(2)\quad \begin{aligned} \mathrm{Ln}(w_j) = {} & \alpha_0' + \alpha_1'\mathbf{X}_j + \\ & \alpha_2'\mathrm{NONUNION}_j*\mathrm{HOSP}_j + \alpha_3'\mathrm{UNION}_j*\mathrm{HOSP}_j + \\ & \alpha_4'\mathrm{UNION}_j*\mathrm{NONHOSP}_j + \alpha_5'\mathrm{RISK}_i + \\ & \alpha_6'(\mathrm{HOSP}_j*\mathrm{RISK}_i) + \alpha_7'(\mathrm{UNION}_j*\mathrm{RISK}_i) + e_j. \end{aligned}$$

The key coefficients in equation (2) are α_5', α_6', and α_7', which together form the compensating differential for risk. The coefficient α_5' measures the risk premium for all workers. The coefficients α_6' and α_7' measure the extra compensation given to hospital and union workers, respectively, for being in risky jobs. The coefficients are interpreted as measuring the portion of the hospital–non-hospital and union-nonunion log wage differentials that are related to risk.[10]

[10]One could also include another interaction of $\mathrm{HOSP}_j*\mathrm{UNION}_j*\mathrm{RISK}_i$ in equation (2). Estimations using this specification never resulted in a statistically significant coefficient on this interaction.

There have been several interpretations of the risk interaction variables in the literature, particularly for the union, risk interaction term. One common interpretation is that it is another source of rents that may accrue to union (and, in this case, hospital) workers. As mentioned above, Sandy and Elliott (1996) interpreted it as a more accurate measure of the compensating wage differential, if firms underreport risk to their workers, that comes from union voice mechanisms. There is also a third interpretation. As Garen (1988) noted, some workers may possess unmeasured skills and abilities that increase their productivity in risky situations. Such individuals may be naturally more careful or adept than other workers in handling risky jobs, or may have received extra training to give them these skills. As mentioned in footnote 6, a source of this increased productivity could also be union voice mechanisms (provided by union representation or even by the firm) that identify strategies to reduce risk. Therefore, the risk interaction terms may be capturing the effects of these voice mechanisms for both union and hospital firms, although there is no empirical way to distinguish between the two effects on wages.

Unfortunately, OLS estimation of the risk coefficients in equation (2) may result in biased coefficients if choice of risk is endogenous. People likely view safety as a normal good, and so increased earnings or high wealth will induce an income effect toward safer jobs (Viscusi 1978; Garen 1988). A two-stage instrumental variables technique is, therefore, employed to make the estimation adjustment for risk endogeneity bias. Since the risk data are industry-based, all variables are averages for the work force of industry *i*. The estimating equation in the first stage is the reduced form equation

$$(3)\quad \mathrm{RISK}_i = \beta_0 + \beta_1\tilde{\mathbf{X}}_i + \beta_2 Z_i + \mu_i,$$

in which the vector $\tilde{\mathbf{X}}_i$ contains the same worker variables (measured at industry averages) presented in equation (1) and Z_i contains a series of industry-level instrumental variables. From this equation, fit-

ted values of industry risk are calculated and used in place of RISK in estimating equation (2).

It should be noted that this methodology differs from the standard method of controlling for endogeneity in this literature (for example, Garen 1988), whereby equation (3) would be estimated using individual-level data for the independent variables, even though the dependent variable is aggregated to the industry level.[11] Methodologically, it is advantageous to estimate such a regression using the same level of aggregation for both dependent and independent variables. Therefore, since the risk rates are calculated at the industry level, they are more closely linked to the work force characteristics than to an individual worker's characteristics. Of course, the optimal method would be to have a measure of workplace risk that is specific to an individual, but the data for such a measure are not available.[12] While the downside to the method we adopt is that the predicted risk values only vary by year and industry rather than by individual, it is consistent with the level of aggregation in the specification of equation (2), where both the dependent and independent variables are measured at the industry level.

The vector of instrumental variables, Z_i, contains variables that influence choice of risk, but not wages, and it includes the industry mean for workers' non-labor income, as well as the percentage of workers in each industry who are home owners, the percentage who have young children living at home, the percentage who are managers, and the percentage who are nonmanagerial white-collar employees.[13] The specification of these variables as instruments models workplace risk as an outcome linked with industries employing workers whose personal characteristics influence their risk preference. For example, workers who have high non-labor income or own their own house (or both) have the financial security to choose a safe workplace. In addition, risk aversion may be higher for some workers, all things equal, and this is controlled for by including a variable measuring the percentage of workers with young children present.

Besides these exogenous controls, the inherent riskiness of the workplace and the extent to which employers monitor such risk is captured by including measures of industry occupation employment shares. For instance, the inclusion of these measures allows for testing whether safe workplaces are staffed more heavily with white-collar workers. Enhanced safety arises because white-collar jobs generally entail less use of dangerous equipment and handling of hazardous materials than do blue-collar jobs. Workplace risk is also influenced by the employment of specific kinds of white-collar workers, such as managers. Because monitoring other workers' activity is one function of managers, for example, employing a relatively large percentage of workers in that white-collar category could promote a safer workplace.

[11]The Garen (1988) methodology includes in the second-stage structural equation a workplace risk variable, the residuals from estimating the risk equation, and the product of the two. These regressors are included to account for bias resulting from worker heterogeneity—that is, we here attempt to identify unmeasured worker productivity. Because we estimate the risk equation (3) at the industry level, the residuals are not person-specific, but industry-specific. Since the residuals do not have the same meaning as in the Garen methodology, we do not include them as regressors in the endogeneity-corrected regressions. This study's correction for risk endogeneity, therefore, more closely resembles that of Viscusi and Moore (1987) and Arnould and Nichols (1983).

[12]An intermediate level of data that would also be preferable is occupational risk by industry (whereby we could identify risk to janitors by industry, rather than by all occupations). However, the BLS does not publish risk data at this level of disaggregation.

[13]The squared and cubed values of non-labor income are also included in the risk equation to pick up any nonlinear relationships between non-labor income and workplace risk. Instrumental variables similar to all those listed in this paragraph can be found in Garen (1988), Siebert and Wei (1994), Sandy and Elliott (1996), Arabsheibani and Marin (2001), and Sandy et al. (2001).

Data

Three data sources are used to estimate equations (1)–(3). Individual worker data on janitors are taken from the 1985–2001 Current Population Survey–Outgoing Rotation Group files (CPS-ORG). The CPS-ORG files report information on worker characteristics, hourly earnings, and the industry and occupation of employment. This sample is limited to individual respondents who are ages 16 and older, are not self-employed, and report janitor as their occupation (Census Occupation number 453). These selection criteria result in a sample comprising 1,137 hospital janitors and 46,526 janitors in other industries.

Equations (1) and (2) require estimating the determinants of individual-level wages. Worker data taken from the CPS-ORG files and included as regressors in the wage-level equations are respondents' (log of) hourly wages, marital status, race, educational attainment, U.S. region of residence, standard metropolitan statistical area (SMSA) residency status, age, age squared, log of weekly hours worked, full-time/part-time employment status, and private/public sector employment status, as well as year indicator variables.

Besides these wage determinants, equation (2) stresses the importance of controlling for workplace risk. Therefore, following Viscusi and Moore (1987), Garen (1988), Dorman and Hagstom (1998), and others, we employ injury and illness data from the Bureau of Labor Statistics (BLS). These data provide an incidence rate of yearly cases of lost workdays due to injury or illness at the job per 100 full-time workers, by industry.[14] Risk information from this BLS data source is merged with CPS-ORG files to create a data set containing information on individual workers and the industry incidence rate of their employing industries.[15]

A shortcoming associated with the use of the two merged data sources is that a small number of BLS incidence data are not available for every industry for every year, and when they are not, the incidence rates for the most recent year are used. This data problem is not likely to bias the results appreciably, because incident rates change very little in a particular industry in a short time span. Another potential problem with merging BLS and CPS data is that to estimate risk equation (3), industry-level data on worker characteristics are needed to match the level of aggregation for the equation's dependent variable. Individual worker information is aggregated by constructing industry measures using the CPS-ORG files for all the variables that appear in equations (1) and (2).[16] The CPS-ORG files are also used to compute the instrumental variables that measure managers' and other white-collar workers' share of their employing industries' work force. Data on non-labor income, home ownership, and child dependency are taken from the 1985–2001 March CPS files to compute the risk equation's other instruments, since these data are not reported in the CPS-ORG files.

Descriptive statistics for janitors constructed from the matched CPS-ORG and BLS files are presented in Table 2. These numbers reveal two key distinctions be-

[14]The incidence rates are defined as the number of injuries and illnesses or lost workdays per 100 full-time workers calculated as $(N/EH)*200{,}000$, where N is the number of injuries or illnesses or lost workdays, EH is the total hours worked by all employees during the calendar year, and 200,000 is the base for 100 full-time workers (100 workers * 40 hours/week * 50 weeks/year).

[15]A complicating issue associated with the use of BLS incidence data is that they are reported at the 3-digit Standard Industrial Classification (SIC) code, while the CPS-ORG industry codes are based on the detailed Census industry code. However, the BLS publishes a correspondence between the SIC and the Census code. For the most part, these are one-to-one correspondences. Where they are not, employment by 3-digit SIC code was used to weight the incidence rates to obtain an incidence rate for the Census code. The only worker group excluded from the merged data set is individuals employed in public administration. Their exclusion arises because the BLS does not report risk information for this group of workers.

[16]Note that these industry averages are taken for *all* occupations, not just janitors, since the risk rates are also for all occupations.

Table 2. Descriptive Statistics for Janitors by Industry and Union Status.

Independent Variable	*Union Hospital (1)*	*Union Non-Hospital (2)*	*Nonunion Hospital (3)*	*Nonunion Non-Hospital (4)*
Hourly Earnings	$9.370	$9.938	$8.018	$6.925
	(3.722)	(4.075)	(3.451)	(3.606)
Risk Rate	8.649	6.715	8.646	6.050
	(3.310)	(5.347)	(3.327)	(4.241)
SMSA	0.832	0.797	0.703	0.687
Married	0.571	0.639	0.583	0.513
Armed Service Veteran	0.239	0.228	0.224	0.161
Male	0.791	0.756	0.833	0.652
White	0.556	0.754	0.707	0.780
Black	0.347	0.203	0.258	0.169
Other Race	0.097	0.043	0.036	0.051
Elementary	0.123	0.113	0.105	0.144
Some High School	0.500	0.449	0.452	0.454
High School Diploma	0.254	0.304	0.297	0.267
College	0.104	0.111	0.123	0.113
Bachelor Degree	0.007	0.017	0.017	0.017
Graduate Degree	0.011	0.005	0.006	0.005
Northeast Region	0.366	0.378	0.211	0.189
South Region	0.157	0.135	0.307	0.324
North-Central Region	0.254	0.299	0.311	0.269
West Region	0.224	0.188	0.171	0.218
Private Sector	0.541	0.524	0.815	0.798
Age	43.724	44.187	41.849	40.073
	(11.626)	(12.134)	(13.537)	(15.785)
Age Squared	2,046.448	2,099.689	1,934.388	1,855.049
	(1,023.811)	(1,077.951)	(1,171.076)	(1,363.136)
Log(Weekly Hours)	3.651	3.654	3.596	3.320
	(0.203)	(0.193)	(0.302)	(0.638)
Full-Time Worker	0.888	0.889	0.863	0.603
Observations	268	8,442	869	38,084

Notes: Numbers in parentheses under averages are standard deviations of continuous variables. Earnings are in 1983/1984 dollars.

tween hospital employees and the baseline comparison group of nonunion janitors employed in other industries. One distinction is that compared to hospital workers, the observable characteristics of nonunion workers in non-hospital industries are more closely associated with low wage payments. For instance, nonunion, non-hospital janitors are less likely than hospital janitors to be employed full-time, married, male, veterans, and residing in a metropolitan location, and they are, on average, younger than hospital janitors. The other key distinction between hospital and non-hospital workers is the greater workplace risk associated with hospital employment. Hospital employees experience a mean rate of over 8.6 annual cases of workdays lost per 100 full-time workers,[17] which is, on average, 28% higher than the risk rate for workers in other industries.[18]

[17]The mean industry risk rate differs for union and nonunion janitors in hospitals because their sample populations differ over time.

[18]Means of industry-level variables used to estimate the risk equation (3) are available from the authors upon request.

Table 3. Partial Results from Estimating Wage Equations for Janitors.
(Dependent Variable: Log of Hourly Earnings in 1983/1984 Dollars)
(t-Statistics in Parentheses)

	OLS *(1)*	*OLS* *(2)*	*OLS* *(3)*	*Endogeneity* *(4)*	*Endogeneity* *(5)*
Hospital Union	0.242***	0.231***	0.210***	0.247***	0.112*
	(11.017)	(10.527)	(5.844)	(11.277)	(1.753)
Hospital Nonunion	0.099***	0.092***	0.143***	0.107***	0.088
	(8.083)	(7.453)	(4.700)	(8.674)	(1.468)
Non-Hospital Union	0.276***	0.272***	0.217***	0.270***	0.110***
	(58.710)	(57.491)	(29.501)	(56.865)	(8.863)
Injury Rate (Risk)		0.004***	1.3E–03**	0.004***	4.6E–04
		(7.655)	(2.502)	(6.840)	(0.675)
Hospital∗Risk			–0.005		0.002
			(–1.622)		(0.233)
Union∗Risk			0.008***		0.020***
			(9.668)		(13.962)

Notes: Full specification found in Appendix Table A1. The reference group is non-hospital, nonunion janitors.

*Statistically significant at the 0.10 level; **at the 0.05 level; ***at the 0.01 level.

Wage Regression Results

Table 3 contains the key coefficient estimates from the log wage regressions.[19] Results from the OLS estimation of equation (1) are presented in column (1). These findings suggest that differences in measurable worker characteristics help explain some of the mean wage premium paid to hospital workers.

Yet, even with worker characteristics taken into account, nonunion janitors in hospitals still receive wage rates above those of nonunion janitors in other industries. The estimated coefficient on the hospital nonunion dummy indicates a hospital premium of 10.4% for nonunion janitors[20]—lower than the 15.8% mean hospital premium for nonunion janitors reported in Table 1, but still sizable. Similarly, the wage premia for union hospital and non-hospital workers are smaller than the mean premia, but they are still over 25%.

The next two columns of Table 3 contain various specifications for measuring risk compensation. Column (2) shows that there is a statistically significant, although small, risk premium of 0.4%.[21] Column (3) indicates that the risk premium accrues primarily to union workers, who have an additional risk premium of more than 0.8% compared to a very small risk premium for nonunion, non-hospital workers of 0.13%. The additional pay given for being in a hospital or for being unionized, once the analysis controls for differential returns to risk by hospital and union status,[22] is little changed from column (1).

[19]The complete estimation results are presented in Appendix Table A1.

[20]Log wage differentials are converted to percentages by exponentiating the estimated coefficient, subtracting one, and then multiplying this difference by 100.

[21]Gunderson and Hyatt (2001), Viscusi and Moore (1987), and Garen (1988) also reported a positive and statistically significant estimated coefficient on the nonfatality risk measure when using OLS. Such findings are not universal, as Arabsheibani and Marin (2001) did not find statistical significance when using OLS.

[22]We recognize that the total wage effect of being in the union is a combination of the differential return to risk because a worker is in a union (the

Table 4. Industry-Level Injury Rate Regressions.
(t-Statistics in Parentheses)

Variable	*Coefficient*	*Variable*	*Coefficient*
Non-Labor Income	–0.065*** (–2.636)	% with Some High School	–5.327*** (–2.714)
Non-Labor Income^2	1.1E–03*** (2.714)	% with HS Diploma	–8.858*** (–5.864)
Non-Labor Income^3	–4.3E–06** (–2.483)	% with Some College	–12.004*** (–7.838)
% Own House	0.458 (0.484)	% with Bachelor's	–13.216*** (–7.619)
% Managers	–5.616*** (–6.464)	% with Graduate Degree	–9.093*** (–5.540)
% White-Collar	–3.209*** (–6.999)	% in Northeast Region	–5.335*** (–6.722)
% with Child Present	–0.722 (–0.678)	% in North-Central Region	–7.000*** (–10.329)
% in Union	7.880*** (12.734)	% in West Region	–8.611*** (–11.291)
% Living in SMSA	–3.881*** (–7.296)	% in Private Sector	0.309 (0.692)
% Married	0.039 (0.035)	Average Age	2.327*** (7.724)
% Veteran	–0.968 (–1.451)	Average Age Squared	–0.032*** (–8.457)
% Male	2.609*** (6.039)	% Full-Time	2.209** (2.520)
% White	–1.732 (–0.769)	Constant	–16.897*** (–2.629)
% Black	–0.117 (–0.044)		

Note: Estimation also includes dummy variables indicating the year.
*Statistically significant at the 0.10 level; **at the 0.05 level; ***at the 0.01 level.

Risk endogeneity might explain these relatively large wage premia for hospital and union non-hospital janitors, especially if these groups of workers possess unobserved traits that command high wage rates compared to nonunion non-hospital janitors. A Hausman test presents strong support for the endogeneity of job risk. Using the instruments reported in equation (3), the null hypothesis that risk is exogenous is rejected at the 0.01 level ($F(40,47,622) = 671.54$)). The results from estimating equation (3) are presented in Table 4. Except for industry variables that measure the percentage of an industry's work force with children present at home and the percentage who own homes, the coefficients on the instrumental variables are statistically significant and negative. The findings for these variables suggest that workers with high non-labor income select employment in industries with low injury rates. These findings also indicate that industries employing a higher percentage of managers

coefficient on the UNION*RISK interaction) and the union premium (measured by the coefficient on the union indicator). In this paper, when we refer to the (non–risk-related) union premium we are, therefore, referring to just this latter component of union wages.

and other white-collar workers have lower injury rates.

Estimates of the earnings equation adjusted for risk endogeneity are reported in columns (4) and (5) of Table 3. The results show that these risk adjustments do not alter the estimated coefficients on the key wage determinants (column 4 versus column 2). The estimated coefficient on the industry risk rate remains positive and statistically significant.[23] Union hospital and non-hospital workers still receive a statistically significant and relatively large premium above the baseline comparison group. Estimates from the equation specified in column (4), however, might confound the wage effect of risk endogeneity by not distinguishing risk compensation by union status. Column (5) reports the results when the wage equation includes a UNION-RISK interaction dummy to account for risk compensation varying by union status. Similarly, a HOSPITAL-RISK interaction dummy is included to account for the possibility that risk compensation varies by hospital status. Once risk endogeneity is corrected for, the results (column 5) show that the estimated coefficient on the UNION-RISK dummy is statistically significant and more than twice as large as the OLS estimate for this variable. The risk endogeneity–corrected results indicate that union janitors receive an additional 2% increase in risk compensation for each additional unit increase in the number of work-related injury cases. In contrast to the UNION-RISK compensation findings, the estimated coefficient on the HOSPITAL-RISK dummy lacks statistical significance. Hence, hospital workers receive no additional risk compensation solely because they are employed in a hospital.[24]

The other important effect of allowing risk to vary by union and hospital status in the endogeneity-corrected results is that the coefficients on the union and hospital indicators become much smaller and less statistically significant. There is now no statistically significant difference, *ceteris paribus*, between the pay of nonunion janitors in hospitals and those not in hospitals, and the nonrisk wage premium for union workers in or out of hospitals is cut by more than 50% once risk endogeneity and risk interactions are controlled for. These results clearly show the importance of including those controls in the specification.

While the results in Table 3 report the changes in the marginal effects of the key wage determinants that occur when we change empirical methodologies, they do not give a sense of how the variables' wage effects change in magnitude in each subsample. To get a sense of the relative size of these wage determinants, we decompose in Table 5 the average wage, while holding the wage structure (the coefficients) constant for each subsample. The coefficients used in this table are the same as in Table 3 (and Appendix Table A1), columns (3) and (5) for the OLS and endogeneity specifications, respectively.[25] Therefore, log wage differences between samples in the "base" column, which includes all covariates except risk, union status, hospital affiliation, and risk interactions, come from the differences in average

[23]The wage-injury risk results derived from using the 2SLS procedure resemble the findings of Garen (1988). In general, however, wage-injury risk findings are mixed when using this estimation technique. For example, Gunderson and Hyatt (2001) reported a lack of statistical significance, while Arabsheibani and Marin (2001) found a negative and statistically significant wage-injury risk association.

[24]Several alternative specifications of the risk equation were used to estimate the fitted value of industry injury rates. These specifications omit different combinations of the instrumental variables. Estimating the wage equation with the different risk-fitted values, however, does not markedly alter the wage-risk associations presented in this study. The wage findings that use the various risk-fitted values are available from the authors upon request.

[25]The only difference is that we re-specify the hospital and union variables. Rather than having four mutually exclusive variables summarizing the four different union and hospital combinations, we include dummy variables for union status, hospital affiliation, and an interaction between the two dummy variables. This allows us to disaggregate the effects of hospital and unionization on the wages of union, hospital janitors.

Table 5. Log Wage Structure Decomposition.

Sample	*Base*	*Linear Risk*	*Hospital*	*Union*	*Risk Interaction*
		OLS Specification			
Nonunion, Non-Hospital	1.949	1.957			
		(0.8%)			
Nonunion, Hospital	2.084	2.096	2.238		2.195
		(1.2%)	(15.3%)		(–4.2%)
Union, Non-Hospital	2.134	2.143		2.360	2.413
		(0.9%)		(24.2%)	(5.4%)
Union, Hospital (1)	1.896	1.908	2.050	2.118	2.144
		(1.2%)	(15.3%)	(7.0%)	(2.6%)
Union, Hospital (2)	1.896	1.908	2.118	2.124	2.144
		(1.2%)	(–0.6%)	(24.1%)	(2.6%)
		Endogeneity Specification			
Nonunion, Non-Hospital	1.959	1.962			
		(0.3%)			
Nonunion, Hospital	2.096	2.099	2.186		2.197
		(0.3%)	(9.1%)		(1.1%)
Union, Non-Hospital	2.146	2.150		2.258	2.422
		(0.4%)		(11.4%)	(17.8%)
Union, Hospital (1)	1.904	1.907	1.994	2.016	2.142
		(0.3%)	(9.1%)	(2.2%)	(13.4%)
Union, Hospital (2)	1.904	1.907	2.016	2.014	2.142
		(0.3%)	(0.2%)	(11.3%)	(13.4%)

Notes: See text for a full explanation. Numbers not in parentheses are log wages. Numbers in parentheses are log wage percentage changes, calculated as $(\exp(\Delta \ln w)-1)*100$, from the previous column. Because of interaction terms, the "Union, Hospital" wage structure decomposition can be calculated in two ways. The first way is analogous to the other decompositions. The second way adds union first and then the hospital interaction terms, so the percentage changes for the "Union" column are between "Union" and "Linear Risk," for the "Hospital" column are between "Hospital" and "Union," and for "Risk Interaction" are between "Risk Interaction" and "Hospital."

characteristics between the samples.[26] The rest of the columns show the log wage rates (and the percentage changes in parentheses) for risk, union status, hospital affiliation, and risk interaction terms.[27] Differences between the "Linear Risk" and "base" columns give the wage rate increase due to the risk premium and the amount of risk faced by workers on average, or the compensating wage differential. The "Hospi-

[26]For example, the average, "base" log hourly wage for nonunion, non-hospital workers is 1.949 (approximately, $7.02) using the OLS coefficients reported in column (3) of Appendix Table A1, while the average, "base" log wage for union, non-hospital workers is 2.134 ($8.45). Interestingly, union, hospital workers have the lowest "base" wage of all groups ($6.66), indicating that the observable characteristics for this group generate relatively low wages.

[27]There are two decompositions for the union, hospital group, since the ordering of adding hospital or union to the wages of this group may give different results. (This is akin to the index problem in the Oaxaca-Blinder-style decompositions.) The first decomposition mirrors the order that is present in the other samples, that is, it adds in the hospital effect first and then the union (and union-hospital interaction) effect. On the other hand, the second decomposition adds in the union effect first and then the hospital (and union-hospital interaction). This means that the calculation of the percentage changes differs from that in the other decomposition. The percentage change in the "Union" column is between "Union" and "Linear Risk." The percentage change in the "Hospital" column is between "Hospital" and "Union," while the percentage change in the "Risk Interaction" column is between "Risk Interaction" and "Hospital."

tal" column shows the increase in wage rates due to a non–risk-related hospital premium, while the "Union" column gives the non–risk-related union premium. Finally, the "Risk Interaction" column records the extra compensation for risk generated by differential risk premia due to being employed in a hospital or union.

In general, the OLS results show that the most important contributors to wages are working in a hospital and union affiliation. Increased compensation for risk is small, between 0.8% and 1.2% for risk in general, and under 5.5% for the risk interaction terms. On the other hand, nonunion, hospital workers receive 15.3% more and union, non-hospital workers receive 24.2% more than the base and linear risk components of their wage. These premia are similar to what union, hospital workers get for both union and hospital affiliation. The conclusions based on the OLS results, therefore, are that there is a very little compensating wage differential for risk (either linearly or in the interactions) and that workers in unions, in hospitals, or in both unions and hospitals are likely receiving rents.

As with Table 3, Table 5 shows that endogeneity correction is important. Now risk, particularly in the interactions with union and hospital status, becomes the predominant determinant of wage rates over the base wage rate. For each group, the linear risk compensation contributes very little to wage rates. The extra pay given for being employed in a hospital now increases wages only 9.1% for hospital workers, compared to the 15.3% in the OLS results. Likewise, the non-risk union premium contributes a smaller amount to wages—just over 11%, compared to 24% in the OLS results, for union, non-hospital workers. Both of these results show that these potential measures of "rent" are much smaller when endogeneity is controlled. The most important part of the wage structure now is the risk interaction term. While it increases wage rates by 1.1% for nonunion, hospital workers, it increases the wage rates of union, non-hospital workers and union, hospital workers by 17.8% and 13.4%, respectively.

Again as mentioned above, this extra risk premium could be interpreted in several ways. First, it could represent another form of union rent. This possibility seems unlikely, however. Negotiating such large risk premia, which generate large wage differentials across workers, would be at odds with unions' traditional efforts to level wages. Alternatively, a union voice interpretation would be that collective bargaining, by countering the market power possessed by large employers, generates a compensating wage differential that these employers otherwise would be able to withhold (Sandy and Elliott 1996). However, this seems unlikely. Occupational safety and health regulations with which firms, particularly hospitals, must comply would make it difficult for firms to take advantage of any informational asymmetries between them and their workers.

A second union voice interpretation is that unions may offer a mechanism for their members to voice their unhappiness with risk in the workplace and to demand safeguards to minimize it, increasing their relative productivity in such settings. Therefore, the union-risk interaction could be a measure of returns to the increased productivity because of this voice mechanism, although the data needed to test this explicitly do not exist.[28]

Finally, the results from the endogeneity-corrected models in Table 3 allow for calculating how much a firm would have to pay janitors for a small incremental increase in workplace risk. Following Thaler and Rosen (1976), the formula used to calculate janitors' implicit value of workplace injury is

$$100 \times (\partial(\ln w) / \partial \text{RISK}) \times \bar{w} \times 2000,$$

where w and RISK are defined as in equations (1)–(3), and $\bar{w}$ depicts the mean hourly

[28]In either union voice interpretation, it is unlikely that there would be significant spillovers to nonunion settings, both because janitorial unions likely organize an entire hospital and because the risks in hospitals are likely very locality-specific. For example, the risks in a central city hospital might be different in kind from those in a rural or suburban location.

Table 6. Value of Workplace Injury.

Sample	*Coefficient(s)*	*Mean Wage*	*Value*
Nonunion, Non-Hospital	0.0004550	$6.93	$630.63
Nonunion, Hospital	0.0028903	8.02	4,636.04
Union, Non-Hospital	0.0201545	9.94	40,067.15
Union, Hospital	0.0225898	9.37	42,333.29

Note: The equation to calculate the value comes from Thaler and Rosen (1976): $100*(\delta(\ln w)/\delta \text{RISK})*w*2000$.

industry wage rate for janitors.[29] The results from these values of workplace injury calculations are reported in Table 6. These findings suggest that the additional cost to firms associated with a one unit increase in workplace risk to janitors ranges from a low of $630.63 for nonunion janitors in nonhospital sectors to a high of $42,233.29 for union hospital janitors, which is consistent with health risk valuations estimated in the literature (for example, Viscusi 1993). The wide dispersion comes from the union-risk interaction term and, if the union voice interpretation of this term is correct, indicates the effect of either the minimization of information asymmetries or the extra productivity from union voice (or both) on the valuation of workplace injuries.

Conclusion

The highly inelastic demand for hospital care can support inefficient wage compensation for union workers who are employed in this service sector. Paying high wage rates is a major concern to hospital management partly because of increasing public concern over the rising cost of hospital services. This study has examined whether the high worker compensation in this service sector could be attributed to greater unionization or to the risky work environment. We confined our investigation to the wages of janitors because focusing on workers in a single occupation minimizes interindustry wage distortion attributable to differences in job responsibilities, such as risk, across industries.

The findings reveal a statistically significant hospital wage advantage for union and nonunion hospital janitors over nonunion janitors in other industries. The findings on nonunion hospital janitors suggest that the observable characteristics of this group of individuals contribute to their high wages. In addition, the marginal effect of hospital employment adds markedly to their wages, whereas risk compensation does not contribute noticeably. OLS findings on union hospital janitors suggest that a large share of their hospital premium is due to union rent seeking, as shown by the large coefficient on the UNION*HOSPITAL variable. Furthermore, the relatively small compensation for risk, shown by the coefficient on the UNION*RISK interaction variable, suggests that unions are only modestly effective in extracting wage concessions from firms with more risky jobs. However, once risk is endogenized, the results show less rent seeking behavior and an increased risk compensation for union hospital janitors. Indeed, it is this union risk compensation that explains much of the wage advantage of union hospital workers. These findings are consistent with theories of union voice.

This study's evidence on marginal wage effects does not flatly reject the possibility of rent sharing. However, compensation for employment in a high-risk workplace provides a compelling explanation for the wage advantage of hospital workers. For instance, competitive wages should prevail for nonunion hospital janitors given their elastic labor demand schedule. The mar-

[29]Note that the BLS reports the nonfatal injury and illness rate per 100 workers for a 2,000-hour work period.

ginal wage effect of hospital employment, then, is more likely consistent with this study's hypotheses concerning union negotiation strength and, particularly, union voice. In contrast, the monopoly control that janitors' unions command does provide union hospital janitors the opportunity to share rent. Yet, the wage effect associated with hospital employment and union membership does not contribute to wages that are markedly different from those of nonunion, non-hospital janitors. Union wages for hospital janitors exceed these competitive levels only when the analysis also includes the marginal effect from risk compensation. Overall, decomposing the wage structure of hospital janitors provides evidence that the wage advantage for these workers does not imply an inefficient allocation of hospital resources. Rather, unions are seemingly able to use voice either to gain better information about the riskiness of hospital jobs or to make members more adept in risky situations, both of which generate extra compensation for union members.

Appendix Table A1
Full Wage Regression Results: Various Specifications.
(Dependent Variable: Log of Hourly Earnings in 1983/1984 Dollars)
(t-statistics in parentheses)

Independent Variable	*OLS (1)*	*OLS (2)*	*OLS (3)*	*Endogeneity (4)*	*Endogeneity (5)*
Hospital Union	0.242***	0.231***	0.210***	0.247***	0.112*
	(11.017)	(10.527)	(5.844)	(11.277)	(1.753)
Hospital Nonunion	0.099***	0.092***	0.143***	0.107***	0.088
	(8.083)	(7.453)	(4.700)	(8.674)	(1.468)
Non-Hospital Union	0.276***	0.272***	0.217***	0.270***	0.110***
	(58.710)	(57.491)	(29.501)	(56.865)	(8.863)
Injury Rate (R)		0.004***	1.3E–03**	0.004***	4.6E–04
		(7.655)	(2.502)	(6.840)	(0.675)
Hospital*R			–0.005		0.002
			(–1.622)		(0.233)
Union*R			0.008***		0.020***
			(9.668)		(13.962)
SMSA	0.049***	0.052***	0.052***	0.051***	0.052***
	(13.215)	(13.763)	(13.900)	(13.662)	(14.005)
Married	0.087***	0.088***	0.087***	0.088***	0.087***
	(23.698)	(23.738)	(23.678)	(23.782)	(23.562)
Veteran	0.020***	0.019***	0.019***	0.020***	0.018***
	(4.079)	(4.003)	(3.851)	(4.086)	(3.798)
Male	0.086***	0.084***	0.084***	0.085***	0.086***
	(22.414)	(21.930)	(22.002)	(22.306)	(22.382)
White	0.013*	0.013*	0.014*	0.013*	0.012
	(1.671)	(1.744)	(1.800)	(1.715)	(1.563)
Black	–0.050***	–0.049***	–0.049***	–0.050***	–0.050***
	(–5.702)	(–5.646)	(–5.620)	(–5.720)	(–5.780)
Some High School	0.094***	0.094***	0.094***	0.095***	0.094***
	(17.614)	(17.663)	(17.651)	(17.823)	(17.594)
HS Diploma	0.150***	0.151***	0.151***	0.152***	0.151***
	(25.579)	(25.806)	(25.825)	(25.934)	(25.790)
Some College	0.177***	0.179***	0.179***	0.180***	0.179***
	(25.522)	(25.783)	(25.800)	(25.929)	(25.801)
Bachelor's Degree	0.181***	0.183***	0.184***	0.185***	0.183***
	(13.540)	(13.668)	(13.742)	(13.796)	(13.728)
Graduate Degree	0.144***	0.148***	0.149***	0.150***	0.149***
	(6.274)	(6.449)	(6.493)	(6.519)	(6.511)
Northeast Region	0.086***	0.088***	0.089***	0.087***	0.091***
	(17.522)	(17.921)	(18.273)	(17.814)	(18.641)
North-Central Region	0.047***	0.046***	0.045***	0.046***	0.045***
	(10.239)	(10.141)	(9.906)	(10.078)	(9.789)
West Region	0.078***	0.079***	0.079***	0.079***	0.080***
	(15.610)	(15.866)	(15.946)	(15.761)	(16.023)
Private Sector	–0.023***	–0.032***	–0.036***	–0.028***	–0.037***
	(–5.827)	(–7.813)	(–8.558)	(–6.970)	(–9.092)
Age	0.024***	0.024***	0.024***	0.025***	0.024***
	(36.454)	(36.428)	(36.399)	(36.523)	(36.399)
Age Squared	–2.5E–04***	–2.5E–04***	–2.5E–04***	–2.5E–04***	–2.5E–04***
	(–32.597)	(–32.497)	(–32.496)	(–32.538)	(–32.521)
Log of Hours	–0.095***	–0.098***	–0.097***	–0.098***	–0.097***
	(–23.833)	(–24.373)	(–24.155)	(–24.468)	(–24.101)
Full-Time	0.206***	0.206***	0.205***	0.207***	0.205***
	(41.699)	(41.680)	(41.525)	(41.897)	(41.492)
Constant	1.022***	1.012***	1.027***	1.000***	1.035***
	(51.622)	(51.030)	(51.677)	(49.904)	(51.352)
Adjusted R-Squared	0.359	0.360	0.361	0.360	0.362
F-Statistic	704.12***	688.4***	658.45***	687.93***	661.78***

Note: Estimation also includes year dummy variables.
*Statistically significant at the 0.10 level; **at the 0.05 level; ***at the 0.01 level.

REFERENCES

Arabsheibani, G. Reza, and Alan Marin. 2001. "Self-Selection Bias with a Continuous Variable: Potential Pitfall in a Common Procedure." *Applied Economics,* Vol. 33, No. 15 (December), pp. 1903–10.

Arnould, Richard J., and Len M. Nichols. 1983. "Wage-Risk Premiums and Workers' Compensation: A Refinement of Estimates of Compensating Wage Differential." *Journal of Political Economy,* Vol. 91, No. 2 (April), pp. 332–40.

Dorman, Peter, and Paul Hagstrom. 1998. "Wage Compensation for Dangerous Work Revisited." *Industrial and Labor Relations Review,* Vol. 52, No. 1 (October), pp. 116–35.

Duncan, Greg J., and Frank P. Stafford. 1980. "Do Union Members Receive Compensating Wage Differentials?" *American Economic Review,* Vol. 70, No. 3 (June), pp. 355–71.

Elliott, Robert F., and Robert Sandy. 1998. "Adam Smith May Have Been Right after All: A New Approach to the Analysis of Compensating Differentials." *Economics Letters,* Vol. 59, No. 1 (April), pp. 127–31.

Erickson, Christopher L., Catherine L. Fisk, Ruth Milkman, Daniel J.B. Mitchell, and Kent Wong. 2001. "Justice for Janitors in Los Angeles and Beyond: A New Form of Unionism in the 21st Century?" Paper presented at the 23rd Annual Middlebury Economics Conference.

Freeman, Richard B., and James L. Medoff. 1984. *What Do Unions Do?* New York: Basic Books.

Garen, John. 1988. "Compensating Differentials and the Endogeneity of Job Riskiness." *Review of Economics and Statistics,* Vol. 70, No. 1 (February), pp. 9–16.

Gunderson, Morley, and Douglas Hyatt. 2001. "Workplace Risks and Wages: Canadian Evidence from Alternative Models." *Canadian Journal of Economics,* Vol. 34, No. 2 (May), pp. 377–95.

Hamermesh, Daniel S., and John R. Wolfe. 1990. "Compensating Wage Differentials and the Duration of Wage Loss." *Journal of Labor Economics,* Vol. 8, No. 1 (January), pp. 175–97.

Kniesner, Thomas J., and John D. Leeth. 1991. "Compensating Differentials for Fatal Injury Risk in Australia, Japan, and the United States." *Journal of Risk and Uncertainty,* Vol. 4, No. 1 (January), pp. 75–90.

Lehrer, Evelyn L., William D. White, and Wendy B. Young. 1991. "The Three Avenues to a Registered Nurse License: A Comparative Analysis." *Journal of Human Resources,* Vol. 26, No. 2 (Spring), pp. 262–79.

Link, Charles R. 1988. "Returns to Nursing Education, 1970–1984." *Journal of Human Resources,* Vol. 23, No. 3 (Summer), pp. 372–87.

Sandy, Robert, and Robert F. Elliot. 1996. "Unions and Risk: Their Impact on the Level of Compensation for Fatal Risk." *Economica,* Vol. 63, No. 250 (May), pp. 291–309.

Sandy, Robert, Robert F. Elliott, W. Stanley Siebert, and Xiangdong Wei. 2001. "Measurement Error and the Effects of Unions on the Compensating Differentials for Fatal Workplace Risks." *Journal of Risk and Uncertainty,* Vol. 23, No. 1 (July), pp. 33–56.

Schumacher, Edward J., and Barry T. Hirsch. 1997. "Compensating Differentials and Unmeasured Ability in the Labor Market for Nurses: Why Do Hospitals Pay More?" *Industrial and Labor Relations Review,* Vol. 50, No. 4 (July), pp. 557–79.

Siebert, W. Stanley, and Xiangdong Wei. 1994. "Compensating Wage Differentials for Workplace Accidents: Evidence for Union and Non-Union Workers in the UK." *Journal of Risk and Uncertainty,* Vol. 9, No. 1 (July), pp. 61–76.

Thaler, Richard, and Sherwin Rosen. 1976. "The Value of Saving a Life: Evidence from the Labor Market." In Nestor E. Terleckyj, ed., *Household Production and Consumption*. New York: Columbia University Press.

Viscusi, W. Kip. 1978. "Wealth Effects and Earnings Premiums for Job Hazards." *Review of Economics and Statistics,* Vol. 60, No. 3 (August), pp. 408–16.

____. 1993. "The Value of Risks to Life and Health." *Journal of Economic Literature,* Vol. 31, No. 4 (December), pp. 1912–46.

Viscusi, W. Kip, and Michael J. Moore. 1987. "Workers' Compensation: Wage Effects, Benefits Inadequacies, and the Value of Health Losses." *Review of Economics and Statistics,* Vol. 69, No. 2 (May), pp. 249–61.