WAGE DETERMINATION IN THE UNION AND NONUNION SECTORS

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WAGE differences between the union and nonunion sectors result from sectoral variation in both worker characteristics, such as educational attainment and job experience, and labor market rewards to these characteristics. Wage equations estimated for union and nonunion workers with only a dummy variable indicating union membership explicitly constrain the labor market rewards to all other worker characteristics to be equal in the two sectors. Such a procedure not only fails to address the interesting question of differences in wage determination across sectors but also may yield poor estimates of union-nonunion wage differentials for workers with otherwise similar characteristics.

This study contains estimates of wage equations for white male union and nonunion employees. The authors find that nonunion wages are generally more responsive than union wages to individuals' education and experience and to regional price-level variation. Despite those differences, however, estimates of union-nonunion wage differentials based on these separate equations do not differ greatly from a differential obtained from a union dummy variable in an equation based on combined union and nonunion observations. Union-nonunion differentials vary widely across occupational groups and are generally larger in the lower skilled and more highly unionized occupations. The results for manufacturing, for which additional industry data are available, indicate a negative impact of high concentration ratios on the wages of all workers and a greater impact of establishment size on nonunion than on union wages. Data were drawn from the May 1973 Current Population Survey.

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The objective of this paper is to examine these issues by estimating separate wage equations for union and nonunion white male workers,[1] avoiding the common pitfalls of using a union dummy variable or permitting only partial interaction between unionism and other determinants of wages.[2]

The first section of the paper contains estimates of separate wage equations for individuals in the union and nonunion sectors. The second focuses only on individuals in manufacturing industries for whom more details on the industry in which they work are available.

Basic Wage Equations

Our observations, taken from the May 1973 United States Current Population Survey, are limited to white, non-Spanish

[1]Throughout this paper, all workers who are not members of a union, including those covered by union contracts, are considered to be in the nonunion sector. Orley Ashenfelter, "Union Relative Wage Effects: New Evidence and a Survey of Their Implications for Wage Inflation" (Working Paper 89, Industrial Relations Section, Princeton University), cites a survey indicating about a 90% overlap between "union members" and "workers covered by a collective bargaining contract" for the Current Population Survey, the data source for our equations.

[2]For example, George E. Johnson and Kenwood C. Youmans, "Union Relative Wage Effects by Age and Education," *Industrial and Labor Relations Review*, Vol. 24, No. 2 (January 1971), pp. 171–80, examine union interaction only with education, age, and race; and Michael J. Boskin, "Unions and Relative Wages," *American Economic Review*, Vol. 62, No. 3 (June 1972), pp. 466–72, and Paul M. Ryscavage, "Measuring Union-Nonunion Earnings Differences," *Monthly Labor Review*, Vol. 97, No. 12 (December 1974), pp. 3–10, examine union interaction only with region and occupation.

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males between age 25 and 64 who are employed in the private sector. Focusing on white males circumvents complex interactions among wage rates, unionism, and discrimination.[3] The age boundaries exclude most white males in school or with part-time jobs, for whom reported wage rates often do not reflect market values of accumulated human capital. Excluding public sector employees avoids complications due to the rather different wage structures for private and government employees.[4] However, although our restricted sample reduces wage variation that cannot be explained by our independent variables, it should be noted that results based on this sample may not be applicable to the remaining sectors of the work force.

The dependent variable in our wage equations is the natural logarithm of each individual's usual weekly earnings divided by his usual weekly hours. This variable may include overtime work and thus should not be confused with individual marginal wage rates. The coefficients of the independent variables in our wage equations may be interpreted as the percentage change in the wage rate effected by unit changes in the explanatory variables.

Our most important independent variables are probably education, experience, and experience-squared. These variables have been included in several previous wage-and-earnings studies, with one rigorous rationale for their inclusion being provided by Mincer.[5] Our education variable indicates years of formal schooling between zero and eighteen. One problem with this measurement is that individuals with more than eighteen years of schooling cannot be differentiated from those with exactly eighteen years of education. However, since only .13 percent of the union and 3.24 percent of the nonunion employees reported this maximum level of schooling, we assume that all such individuals have had exactly eighteen years of education. Equations estimated with a dummy variable for eighteen years of schooling and an integer variable for seventeen years or less yield results with virtually identical coefficients to the equations reported here.

The experience variable is defined as age minus education minus 6. This variable indicates potential work experience after the completion of formal schooling. It overstates work experience for those individuals who have not held jobs for their entire postschool careers, but this is much less likely to be the case for our sample than for a random sample of American workers.

We expect that the coefficients of the education and experience variables will be positive in both equations and that the coefficient of the experience-squared term will be negative, reflecting diminishing returns to work experience as experience itself increases. We also expect that the effect on wage rates of education and experience will be greater in the nonunion equation than in the union equation if employers in the nonunion sector are more responsive to market forces. This argument is perhaps stronger for the education than for the experience variable, given the importance of seniority and built-in pay increases for union members.[6]

Other independent variables in the equations include marital status, perhaps regarded by many employers as a proxy for such personality traits as stability and responsibility, and veteran status, indicating either training in a given skill or time lost in the civilian labor market. The effect on wage rates of being married is expected to be positive, that of veteran status ambiguous, in both sectors.

Also included is a regional price index,

[3]See Orley Ashenfelter, "Racial Discrimination and Trade Unionism," *Journal of Political Economy*, Vol. 80, No. 3, Part I (May/June 1972), pp. 435–64.

[4]See Sharon P. Smith, "Government Wage Differentials by Sex," *Journal of Human Resources*, Vol. 11, No. 2 (Spring 1976), pp. 185–99.

[5]See Jacob Mincer, *Schooling, Experience, and Earnings* (New York: National Bureau of Economic Research, Columbia University Press, 1974).

[6]Sherwin Rosen, "Trade Union Power, Threat Effects, and the Extent of Organization," *Review of Economic Studies*, Vol. 36(2), No. 106 (April 1969), pp. 185–94, found that wages in the union sector are affected more by age and experience than by educational attainment. He explains this result by the prevalence of seniority systems in which wage rates are based on age and experience rather than education and by the generally relatively low formal educational attainment of older workers.

the construction of which is described in the Appendix. Although we expect the price index coefficient to be positive in both the union and nonunion equations, we expect it to be larger in the nonunion equation, given the tendency for unions to attempt to eliminate wage differentials resulting from geographic differences, especially through centralized bargaining of unions in national product markets. Finally, we have included a set of (two-digit) industry and occupation dummy variables to correct for varying labor demand across labor markets.

Results

Table 1 contains the results from our estimated union and nonunion wage equations, along with a t-statistic indicating the difference between the corresponding coefficients in the two equations.[7] Although not shown here because of space limitations, we also estimated combined wage equations for unionized and nonunionized employees, both with and without a union membership dummy variable.[8] A comparison of those and the Table 1 equations indicates that in neither case can the null hypothesis of equality of the coefficients across equations be accepted. Using the equation with the union dummy, testing the hypothesis of equality for all coefficients except the constant, $F = 18.2 > 1.71 = F_{.01}(30, 1000) > F(34, 12505)$; using the equation without the union dummy, testing the equality of all coefficients including the constant, $\hat{F} = 27.1 > 1.71 = F_{.01}(30, 1000) > F_{.01}(35, 12503)$.

Our predictions of signs and magnitudes of coefficients are strongly supported by the Table 1 regressions. The education coefficients, the partial derivatives of the logarithm of the wage rate with respect to experience, and the price variable are all significantly greater than zero at the one percent level in each equation. Furthermore, the corresponding constructs in each nonunion equation than in the union equation, again at the one percent level.[9] In addition, the coefficients of the variables equation are significantly greater in the indicating marital status are positive in each equation, with the effect stronger for married men with wife absent or deceased. The effect on the nonunion wage is significantly greater than that on the union wage at the one percent level only for those men married with wife present. Veteran status does not appear to affect wage rates significantly in either sector.

In general, unions appear to have the effect of flattening out the wage equation. The union wage-experience profile peaks at about 27 years and, as noted above, is flatter than that of the nonunion sector, which peaks at about 28.5 years. Thus, an additional year of experience for a new worker raises the wage of union workers by about one percent and that of the nonunion workers by about 2.3 percent. For each additional 10 years of experience, the union effect declines by roughly .4 percent and the nonunion effect by .8 percent. The percentage increases in wage rates resulting from an additional year of formal education are 1.8 percent for union workers and 5.1 percent for nonunion workers. The price index coefficient is about one percent in the nonunion equation and roughly half that in the union equation.

Johnson and Youmans also observe relatively flat union-education and union-age profiles,[10] which they explain in terms of

[7]Let β_u, β_n be the estimated coefficients of the same explanatory variable in the union and nonunion equation and σ_u and σ_n their respective estimated standard errors. Then the t-statistic given in Table 1 is computed as

$$\frac{\beta_u - \beta_n}{\sqrt{\sigma_u^2 + \sigma_n^2}}$$

[8]Detailed results of all the combined equations described in this paper, and of the equations on which Table 2 below is based, can be obtained from the authors on request.

[9]Hypothesis tests regarding the partial derivative of wage rates with respect to experience require estimates of covariances between coefficients of experience and experience-squared, statistics unfortunately not printed out by our computer program. Our remarks in the text, however, will hold at most experience levels unless the magnitudes of these covariances are extraordinarily large. The nonunion experience effect is greater than the union experience effect for all positive experience levels. For the corresponding equations dealing with workers in manufacturing, however, the nonunion experience effect exceeds the union experience effect for levels of experience between 1.9 and 52.3 years.

[10]Johnson and Youmans, "Union Relative Wage Effects by Age and Education."

Table 1. Separate Wage Equations for Union and Nonunion Employees. (Standard errors are in parentheses)

Independent Variables	*Dependent Variable log (hourly wage)* Union	Nonunion	*t-Statistic of Difference*
Education	.01771**	.05070**	−10.771**
	(.00222)	(.00211)	
Experience			
Experience	.00976**	.02272**	−5.587**
	(.00167)	(.00161)	
Experience-squared	−.00018**	−.00040**	5.185**
	(.00003)	(.00003)	
Price index	.00523	.01115	−6.247**
	(.00066)	(.00068)	
Marital status			
Spouse present	.09669**	.21582**	−4.454**
	(.01939)	(.01842)	
Spouse absent	.07578	.14859**	−1.276
	(.04037)	(.04032)	
Widowed	.06159**	.11524**	−1.351
	(.02781)	(.02834)	
Single, never married	—	—	—
Veteran status	−.00603	−.01058	.339
	(.00907)	(.00987)	
Occupation			
Professional workers	.29455**	.36544**	1.633
	(.03149)	(.02986)	
Managers	.16377**	.38286**	−5.517**
	(.02776)	(.02842)	
Sales workers	.06551	.25837**	−3.202**
	(.05175)	(.03080)	
Clerical workers	.05971*	.14681**	−2.115*
	(.02588)	(.03203)	
Craftsmen	.19981**	.21318**	−.414
	(.01735)	(.02728)	
Operatives	.07878**	.05103**	.790
	(.01816)	(.03007)	
Transport equipment workers	.05428**	−.00369	1.463
	(.02082)	(.03371)	
Service workers	−.10337**	−.00119	−2.202*
	(.02974)	(.03563)	
Laborers	—	—	—
Industry			
Mining	.36746**	.54119**	−1.721*
	(.08654)	(.05200)	
Construction	.61119**	.49830**	1.226
	(.08308)	(.03970)	
Manufacturing—nondurables	.23785**	.46514**	−2.475**
	(.08314)	(.03899)	
Manufacturing—durables	.23780**	.49688**	−2.842**
	(.08278)	(.03815)	
Railroad	.36207**	.63607**	−2.319*
	(.08502)	(.08201)	
Other transportation	.45511**	.46091**	−.061
	(.08422)	(.04525)	
Other utilities	.28670**	.59577**	−3.219**
	(.08445)	(.04566)	

Table 1. (Continued).

Independent Variables	Dependent Variable log (hourly wage) Union	Nonunion	t-Statistic of Difference
Wholesale	.28637**	.41775**	−1.374
	(.08666)	(.04039)	
Retail	.25799**	.20195**	.602
	(.08465)	(.03873)	
Finance	.29274**	.43634**	−1.341
	(.09878)	(.04144)	
Business-repair	.24169**	.33518**	−.929
	(.09103)	(.04302)	
Personal services	.26022**	.11579*	1.154
	(.11119)	(.05617)	
Entertainment	.31364**	.21384**	.870
	(.09925)	(.06340)	
Welfare	−.11910	−.47953**	2.349**
	(.14350)	(.05440)	
Hospitals	.11821	.22501**	−.814
	(.11819)	(.05520)	
Medical, except hospitals	.35951*	.33770**	.096
	(.21678)	(.07113)	
Education	.12998	.03052	.850
	(.10548)	(.05043)	
Other professional services	.43615**	.46666**	2.271*
	(.10238)	(.04716)	
Agriculture	—	—	—
Constant	4.84777	3.31053	
R^2	.31042	.39709	
F	57.87279	157.52971	
N	4406	8167	
S.E.E.	.28251	.41935	

*Significant at the .05 level using a one-tailed test.
**Significant at the .01 level using a one-tailed test.

union seniority systems. Under these systems, workers tend to be promoted if they remain with a firm, so there is little incentive for human capital investment. In addition, in unionized firms with seniority systems, employers are more likely to keep an older worker in a responsible position than to promote a younger worker.

The Table 1 regressions also indicate that the union effect on laborers' wages is high relative to that for most occupational groups and that the effect on agricultural workers' wages is high relative to that for most industrial groups. No occupational union-nonunion differential is significantly higher at the 5 percent level than that for the reference occupation, laborers; the differentials for managers and sales, clerical, and service workers are signifiantly lower. Only workers in welfare and other professional services industries have significantly higher industrial union-nonunion wage differentials than employees in the reference industry, agriculture; the differentials for workers in mining, manufacturing, railroads, and other utilities are significantly lower.

The above results must be qualified by the failure of these equations to capture possible effects of the explanatory variables on forms of employee compensation other than money wage rates. If such effects are stronger for union members than others, then the effect of such variables as education and experience on *total* employee compensation will not necessarily be greater for nonunion workers.

Union-Nonunion Differential

The union-nonunion wage differential can be computed in several ways. The simplest is obtained from the coefficient of the union dummy variable in a wage equation identical, except for this dummy variable, to those reported in Table 1 but based on observations on both union and nonunion employees. The antilog of this coefficient (.14734 with standard error .00855) minus one yields an estimated union-nonunion wage differential of 15.87 percent. Alternatively, using the Table 1 equations, we can compute the logarithmic differential of the wage rates assuming first that the union wage structure applies to workers in both sectors and second that the nonunion wage structure applies to all workers. Letting W represent logarithms of individual wage rates, β the vector of estimated coefficients in the wage equations, and Z the vector of explanatory variables, and using superscripts for the union and nonunion wage rates and bars to represent mean values, the first of these measured differentials is

$$\overline{W}^u - \overline{W}^n + \beta^u(\overline{Z}^n - \overline{Z}^u)$$

and the second is

$$\overline{W}^u - \overline{W}^n + \beta^n(\overline{Z}^n - \overline{Z}^u).$$

The logarithmic differential estimated by the first expression is .0888, that of the second .1461. The corresponding percentage union-nonunion wage differentials are 9.29 percent and 15.73 percent. These estimates almost bracket the 15.87 percent differential obtained by simply using a union dummy variable, despite strong evidence of differing wage structures in the union and nonunion sectors.

Equations similar to those reported in Table 1 were estimated for each occupation separately. The independent variables were the same as in the Table 1 regressions except for the omission of the occupation dummies and mild aggregation of the industry variables. Because of the small number of observations in certain industry-occupation cells, only mining, construction, manufacturing (durable and nondurable), railroad, transportation, utility, and wholesale industry dummy variables were included.

The results show that both the wage-education profile and the wage-experience profile are flatter in the union than in the nonunion equation, the former for all occupations except sales and the latter for all except sales, managers, and professional workers. We can reject at the 5 percent level the null hypothesis of equality of coefficients across sectors for each equation except for service, sales, and professional workers.[11]

The estimated union-nonunion differentials for each occupation are presented in Table 2. In all cases but "all occupations" and "sales workers," estimated union-nonunion wage differentials obtained from the separate union and nonunion equations bracket the estimated differential obtained from the union dummy coefficient in a combined equation for the corresponding occupation. And for "all occupations" and "sales workers," the combined equation differential is only slightly greater in absolute value than the differential based on the separate equations and the nonunion structure. The union-nonunion differentials are generally greater in the lower skilled and highly unionized occupations, consistent with the occupation coefficients reported in Table 1.

Workers in Manufacturing

We next estimate wage equations for workers in manufacturing industries by using the four-firm concentration ratio and the percentage of establishments in the industry with more than one hundred employees, rather than industry dummy variables.[12] The United States Census of Manufacturing gives four-firm concentration ratios for four-digit industries, but the Current Population Survey reports three-digit rather than four-digit industry classifications. Therefore we constructed the three-digit values by weighting the four-digit ratios by sales. These national concentration ratios

[11]The hypothesis of equality was rejected at the 5% level for professional workers when the union dummy was excluded, but not when it was included.

[12]The means of these variables for the entire manufacturing sample are 29.57 for concentration ratio and 117.91 for establishment size. The range for establishment size is 0 to 1000.

Table 2. Estimates of Union/Nonunion Wage Differentials.[a]

Occupation	*Percent of Workers in Occupation Unionized*	*Union Structure*	*Nonunion Structure*	*Combined Equation*	*t-Statistic on Union Dummy*
All occupations	35.04	9.29	15.73	15.87	17.23
Craftsmen	47.67	25.54	15.58	19.47	14.75
Operatives	61.84	22.2	15.95	18.64	2.65
Managers	8.70	−12.3	4.35	2.37	5.77
Professional workers	8.05	−6.92	2.86	.51	.11
Transport equipment operatives	57.82	43.99	32.93	38.37	12.14
Clerical workers	27.98	4.52	−2.08	2.31	.71
Laborers	44.98	61.56	33.38	42.21	10.55
Service workers	31.59	19.04	14.41	15.45	2.90
Sales workers	4.64	−.41	−2.96	−4.15	−.54

[a] The union (nonunion) structure column contains estimates based on the assumption that the union (nonunion) equation applies to all workers. The combined equation estimates are based on the coefficient of the union dummy variable in an equation computed from observations on both union and nonunion employees. The t-statistics in the last column are associated with these dummy variable coefficients.

understate the market power of firms in industries with clearly regional or local markets. Employers in concentrated industries may pay higher wages than others if workers in these industries share the monopoly or oligopoly rents obtained by these firms, a situation especially likely to occur when these workers are unionized and therefore have a relatively strong bargaining position. On the other hand, firms in concentrated industries may have substantial financial reserves with which to resist union wage demands, despite the fact that these firms originally may have been easier for unions to organize because of economics of scale.

The Census of Manufacturing also provides establishment-size information. Because concentration ratios and the incidence of large firms are positively correlated, some arguments supporting relationships between wage rates and concentration ratios have also been used in positing relationships between wage rates and establishment size. To the extent that establishment size indicates the ability to achieve economics of scale in providing fringe benefits, establishment size may have negative coefficients in regressions explaining the monetary wage, assuming a zero effect on total employee compensation. To the extent that unions achieve relatively high percentages of fringe benefits in total employee compensation, the establishment size variable may be more strongly negative in the union equation than in the nonunion equation.

The manufacturing wage equations are reported in Table 3. The last column gives the t-statistic for the difference in corresponding coefficients in the union and nonunion equations.

Generally, the results are rather similar to those reported in Table 1. The signs and relative magnitudes of the education, experience, and price coefficients are all the same as in Table 1. The veteran-status coefficients again do not differ significantly from zero and the marital-status variables are again positive, although not always significantly greater than zero at conventional test levels and in no case significantly different across the two equations. Again comparing these two equations with an equation estimated for all manufacturing employees with and without a union dummy variable yields the conclusion that the equations differ significantly across sectors. The null hypothesis of equality of the coefficients is strongly rejected at the one percent level: $\hat{F} = 4.8 > 2.01 = F_{.01}$ (16, 1000) $> F_{.01}$ (18, 2038), using the equation with the union dummy variable; $\hat{F} = 5.3 > 2.01 = F_{.01}$ (16, 1000) $> F_{.01}$ (19, 2036), using the equation without the union dummy variable.

Table 3. Separate Wage Equations for Employees in Manufacturing Industries. (Standard errors are in parentheses)

Independent Variables	*Union*	*Nonunion*	*t-Statistic of Difference*
Education	0.02908**	0.06859**	−6.09850**
	(0.00468)	(0.00448)	
Experience	0.00422	0.02427**	−3.99929**
	(0.00354)	(0.00355)	
Experience squared	−0.00002	−0.00039**	3.73756**
	(0.00007)	(0.00007)	
Married-spouse present	0.11986**	0.14002**	−0.34225
	(0.04030)	(0.04296)	
Married-spouse absent	0.05647	0.07210	−0.12672
	(0.08377)	(0.09053)	
Widowed	0.12439*	0.03251	1.07524
	(0.05761)	(0.06311)	
Veteran status	0.01666	−0.01839	1.23868
	(0.01926)	(0.02073)	
Concentration ratio	−0.00160*	−0.00091	−0.58047
	(0.00081)	(0.00087)	
Establishment size	−0.00012	0.00016	−1.97007*
	(0.00009)	(0.00011)	
Price index	0.00313**	0.00740**	−2.26851*
	(0.00128)	(0.00141)	
Occupation			
Professional workers	0.026995	0.43751**	−1.71829*
	(0.06352)	(0.07399)	
Managers	0.47615**	0.49199**	−0.13548
	(0.09013)	(0.07447)	
Sales workers	0.34333**	0.35711**	−0.08892
	(0.13132)	(0.08229)	
Clerical workers	0.12868*	0.19387**	−0.65510
	(0.06055)	(0.07897)	
Craftsmen	0.22454**	0.27889**	−0.63472
	(0.04652)	(0.07189)	
Operatives	0.12574**	0.12740*	−0.01916
	(0.04615)	(0.07335)	
Transport equipment operatives	0.17043**	0.11710	0.50224
	(0.05627)	(0.09005)	
Service workers	−0.02488	−0.08240	0.46783
	(0.06853)	(0.10208)	
Constant	5.12560	3.88667	
R^2	.19	.47	
F	10.54	65.76	
N	778	1296	
S.E.E.	.244	.350	

*Significant at the .05 level using a one-tailed test.
**Significant at the .01 level using a one-tailed test.

The union-nonunion wage differentials based on the separate equations, −3.43 percent, assuming the union structure applies to all employees, and 6.97 percent, assuming the nonunion structure applies to all employees, bracket the estimate obtained from the coefficient of the union dummy variable in the combined equation, 5.31 percent. These relatively small differentials are consistent with the greater coefficients for manufacturing industry dummy variables in the nonunion as compared with the union equation in Table 1.

The concentration ratio variable is negative in each equation, significantly so at the 5 percent level only in the union equation and not significantly different across equations.[13] These results provide mild support for the ability of firms in concentrated industries to withstand union wage demands and are also consistent with the hypothesis of concentrated firms' monopsony power. The establishment-size variable is negative in the union equation and positive in the nonunion equation, although in neither case significantly different from zero at conventional test levels. However, the nonunion establishment-size coefficient is significantly greater than the union coefficient at the 5 percent level. This is consistent with our hypothesis that both large firms and unions have relatively strong preferences for fringe benefits as compared with monetary wage payments.[14]

Conclusions

The results of this paper clearly indicate differing structures of wage determination in the union and nonunion sectors. Nonunion sector wages are generally more responsive to individual worker levels of education and experience and to regional price-level variation. Despite the greater labor market rewards to these characteristics in the nonunion sector, union-nonunion wage differentials are positive for most occupations, especially so in those that are more highly unionized and less skilled.

Despite the different wage structures in the two sectors, estimated union-nonunion wage differentials obtained from separate union and nonunion equations do not differ greatly from differentials obtained from coefficients of union dummy variables in wage equations incorporating observations on both union and nonunion employees. Consequently, the simpler methodology does not appear to produce seriously biased estimates of union-nonunion differentials despite its masking of sectoral differences in wage determination.

Our estimates of these differentials from both separate and combined equations are generally slightly lower than those obtained by other investigators. Although there are important differences in the analysis and data used in other studies that render comparisons difficult, the most likely explanation for our relatively low estimates is the high unemployment and unanticipated inflation in 1973 relative to the late sixties, when the studies cited above were undertaken. One would expect union bargaining power to be weaker in periods of higher unemployment. In addition, contract-determined union wages are generally not as responsive to unanticipated inflation as are wages in the nonunion sector, although the increased use of cost-of-living escalators should reverse this trend.

As many authors have pointed out, these estimated union-nonunion differentials are imperfect measures of the extent to which unions have raised wages over levels that would have prevailed in the absence of unionism because nonunion wages themselves are affected by the presence of unionism.[15] Nonunion employers may set higher wages in an attempt to prevent their firms from being unionized or simply to compete with union employers in recruiting labor. On the other hand, high union wages and possibly resultant high product prices will tend to reduce employment in the union sectors (and in those nonunion establish-

[13]Sherwin Rosen, "Trade Union Power, Threat Effects, and the Extent of Organization," and Leonard W. Weiss, "Concentration and Labor Earnings," *American Economic Review*, Vol. 56, No. 1 (March 1966), pp. 96–117 also found negative coefficients for union concentration ratio interaction terms in their (combined) wage equations. Our mild negative relationship between concentration ratios and wages differs from the findings of Weiss, who found no relationship, and James A. Dalton and E. J. Ford, Jr., "Concentration and Labor Earnings in Manufacturing and Utilities," *Industrial and Labor Relations Review,* Vol. 31, No.1 (October 1977), pp. 45–60, who found a statistically significant positive relationship between concentration ratios and individual wages. Rosen found a positive but not statistically significant relationship between concentration ratios and mean industry wages. These studies are more compatible with our nonunion than with our union results, especially considering the presumed positive correlation between concentration ratio and establishment size, a variable not included by Dalton and Ford.

[14]Albert Rees suggested this interpretation.

[15]See especially H. Gregg Lewis, *Unionism and Relative Wages in the United States* (Chicago: The University of Chicago Press, 1963).

ments where the threats of union organization and competing recruitment are effective) and to increase the supply of labor to the nonunion sector, thus depressing nonunion wages. Union-nonunion wage differentials also allow union employers to ration the presumed excess supply of labor to union jobs by selecting especially well-qualified workers. Part of our estimated union-nonunion differentials may reflect sectoral differences in labor quality not accounted for by our independent variables.

Finally, although the positive correlation generally observed between unionization and wage rates is usually attributed to the positive effect of union membership on wages, it may be the case that the probability of union organization is a positive function of preexisting wage rates. Ashenfelter and Johnson and also Schmidt and Strauss have examined this question with industry and individual observations, respectively, by estimating a simultaneous model including a union membership equation as well as a wage equation.[16] Wages used in the membership equations are current and postunionization rather than the more appropriate preunionization wages. In both cases, the authors found a much lower positive effect of unions on wages than would be obtained from a corresponding single equation model.

Appendix

This appendix contains a discussion of the price variable used in the regressions in the main text. The May 1973 Current Population Survey lists the specific SMSA in which each individual resides for the largest 99 SMSAs in the United States, and otherwise the individual's region (Northeast, North Central, South, or West) and whether or not he resides in an SMSA.[17] The U.S. Department of Labor *Handbook of Labor Statistics 1975* (Washington, D.C.: G.P.O., 1976), Table 145, p. 375, provides annual budget indices at an intermediate living standard for four-person families for 39 SMSAs for autumn 1973. These budget indices may be better than consumer price indices as indicators of the buying power of wages, since they take into consideration regional variation in consumer market baskets. In order to assign budget indices for the remaining SMSAs, we estimated an equation for these 39 SMSAs in which the budget index is the dependent variable and regional dummy variables and SMSA population are the independent variables. We expect the coefficient on SMSA population to be positive, since higher population may indicate increased demand for land and ultimately higher prices for final goods and services. The results are reported in Table 4.

Table 4. Variations in Budget Indices.[a]

Independent Variables	*Coefficients*	*Standard Error*
Constant	102.68298**	1.9445
North Central	−4.58795*	2.1039
West	−6.54861**	2.5214
South	−10.81971**	2.2971
Population	.000000979	0.00034
R^2 = .5567	S.S.R. = 714.612	
N = 39	S.E.E. = 4.65348	

*Significant at the .05 level using a one-tailed test.
**Significant at the .01 level using a one-tailed test.
[a] The price indices for our observations are equal to the original budget indices for the 39 SMSAs incorporated in the above equation, the predicted value of the dependent variable for the remaining SMSAs that were specifically noted in the data set, and the predicted value of the dependent variable for SMSAs not specifically cited when the population observation is set equal to 160,000. The 160,000 figure is roughly the mean of SMSAs ranked 100 to 250 in order of population.

[16] Orley Ashenfelter and George E. Johnson, "Unionism, Relative Wages and Labor Quality in U.S. Manufacturing Industries," *International Economic Review,* Vol. 13, No. 3 (October 1972), pp. 488–508, and Peter Schmidt and Robert P. Strauss, "The Effect of Unions on Earnings and Earnings on Unions: A Mixed Logit Approach," *International Economic Review,* Vol. 17, No. 1 (February 1976), pp. 204–12.

[17] Except in New England, a standard metropolitan statistical area is a county or group of contiguous counties that contains at least one city of 50,000 inhabitants or more. In addition to the county, or counties, containing such a city or cities, contiguous counties are included in a standard metropolitan statistical area if according to certain criteria they are essentially metropolitan in character and socially and economically integrated with the central city. In New England, standard metropolitan statistical areas have been defined on a town rather than a county basis.