NBER WORKING PAPER SERIES

THE PERCENT ORGANIZED WAGE (POW) RELATIONSHIP FOR UNION AND FOR NONUNION WORKERS

Richard B. Freeman

James L. Medoff

Working Paper No. 305

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge MA 02138

December 1978

We have benefitted from the comments of K. Abraham, C. Brown, H. G. Lewis and L. Summers. We are most grateful to J. Fay, C. Ichniowski, L. Nelson, M. VanDenburgh, and J. Zax for research assistance. The study has been supported by the National Science Foundation (Grant APR77-16279) and the National Bureau of Economic Research (under its program of research on labor economics). Any opinions expressed are not necessarily shared by the individuals who have aided us, NSF, or NBER.

NBER Working Paper 305
December 1978

The Percent Organized Wage (POW) Relationship for Union and for Nonunion Workers

ABSTRACT

This paper analyses the relation between the percent of workers organized in a product market and the wages received by union workers and by nonunion workers. It argues that the greater is the union coverage of a sector, the lower will be the elasticity of demand for the product of organized firms (since there will be fewer nonunion competitors) and as a result the lower will be the elasticity of demand for union labor and the larger the union wage gains. Estimates of the link between coverage and wages using information on individuals and on establishments shows the expected positive relation for union workers across manufacturing industries. By constrast, nonunion wages in manufacturing appear to be unrelated or only modestly related to the percentage organized. Estimates of the link between the percentage of construction workers unionized in a state and the wages of union and nonunion construction workers reveal relationships similar to those for manufacturing. Overall, the results strongly suggest that the percent organized is an important determinant of union wages and of the union-nonunion wage differential.

Richard B. Freeman and
James L. Medoff
National Bureau of Economic
Research
1050 Massachusetts Avenue
Cambridge, Mass. 02138
(617) 868-3915/3914

The impact of unions on wages is likely to depend on the extent to which they organize workers in the relevant product market.[1] As the organization in a market increases, opportunity for substituting nonunion for union products will be reduced, lowering the elasticity of demand for organized workers and the potential loss of employment for a given wage increase. As a result, the wages of union workers are likely to be higher, all else the same, the greater the percent organized. The wages of nonunion workers may also be influenced by the extent of organization, though the direction of the effect is not clear. On the one hand, union wage gains due to greater coverage may induce increases in nonunion wages because of the threat of organization and/or because of shifts in demand favoring nonunion producers brought about by the increased relative cost of union labor; on the other hand, the supply of labor to nonunion firms may increase as a result of reduced employment in the union sector, which would most likely depress nonunion wages. Whether the threat plus demand effect or the supply effect dominates is an empirical issue. The impact of the percent organized on the union wage differential (the difference between the natural logarithm of union and nonunion wages) depends on the relative magnitudes of the likely positive impact on union wages and positive or negative impact on nonunion wages.

This paper seeks to disentangle the relation between the percent of workers organized in a product market and the wages received by union workers and by nonunion workers.[2] In contrast to most of the union wage literature, which either relates some average of wages in an industry to the percent organized (Lewis, Rosen, Weiss) or which relates the wages of individuals to their membership in a union (Boskin, Johnson and Youmans, Ashenfelter), our analysis examines the impact of the percent organized on the compensation of union labor and nonunion labor taken separately. By relating the wages of unionized workers or establishments to the percent organized, we estimate directly the impact of the extent of collective bargaining coverage on the absolute earnings of union workers as opposed to their earnings relative to comparable workers who are not unionized. By relating nonunion wages to the percent organized, we provide the first direct estimates of the net of the threat, demand, and supply effects on the pay of unorganized labor in a product market.

Two sets of data are used in the analysis: information on individuals from the 1973, 1974, and 1975 May Current Population Surveys (CPS), which contain data on usual weekly earnings and usual weekly hours, union membership status, and key personal characteristics; and information on establishments from the Bureau of Labor Statistics' 1968, 1970 , and 1972 Expenditures for Employee Compensation (EEC) surveys, which contain data on compensation per labor hour, collective bargaining coverage, and some relevant establishment characteristics. The availability of both individual and establishment data, each of which has weaknesses and strengths, provides a valuable check on our findings. Because unionized workers are primarily blue-collar labor, we restrict analyses to production or nonoffice workers. Because of the distinct features of wage-setting in the public sector, we also limit attention to private wage and salary workers. Finally, since the analytic model relates collective bargaining coverage to wages through the product demand curve, we do separate analyses of different sectors, depending on the nature of the product market. We compare coverage and wages across <u>industries</u> in manufacturing, where product markets can be taken to be national, but compare coverage and wages across <u>geographic</u> areas in construction and other sectors, where product markets are local in nature.

The major finding of our research is that in manufacturing the percent organized in a product market has a strong positive association with the wages of union workers, but either no association or a weak positive association with the wages of nonunion workers. As a result, the union-nonunion wage differential in a product market in these sectors is positively related to the extent of organization. For construction, the results appear to be similar, though sensitive to specification. For the small number of industries outside of manufacturing that we have examined, our results are more mixed, but suggest that both union and nonunion wages increase with percent organized, with the union wage differential generally growing with coverage.

The study is divided into four sections. Section I examines the theoretical reasons for expecting the extent of unionization in a product market to affect both union and nonunion wages and the factors likely to influence the direction and magnitude of the effects. Section II describes the data and econometric specification used to estimate the percent organized wage (POW) relation for union and for nonunion workers. Section III presents estimates of the impact of percent organized on the compensation of unionized and nonunionized production workers in manufacturing and construction and considers briefly these relationships in other sectors. Section IV summarizes our findings and offers some suggestions for future research.

I. The POW Relationships

There are two basic reasons for expecting the percent organized in a market to be positively related to the wages of union workers. First, it is likely that the greater is the union coverage of a sector, the lower will be the elasticity of demand for the product of organized firms (since there are fewer nonunion competitors) and, as a consequence, the lower will be the elasticity of derived demand for labor. If, as is most probable, unions are concerned with the number and employment of members, as well as with their wages, they can be expected to press for higher wages in markets with a relatively low labor demand elasticity and, hence, with a relatively high percent covered. Second, a positive POW curve for covered workers might also be observed because unions have located and prospered in sectors with low elasticities of demand for labor. In this scenario, a low elasticity of demand causes a high coverage ratio and, at the same time, allows the union to obtain high wages; the coverage ratio is an indicator of an initially low elasticity of demand for labor rather than a "cause" of a low elasticity. Third, to the extent that unions in heavily organized sectors are able to reduce the substitutability between production labor and other factors, along lines suggested by Freeman and Medoff (1977), a positive coverage-wage relationship may result. In this case the causality is from coverage to the elasticity of substitution (rather than to the product demand elasticity) and then to the elasticity of the demand for labor and the wage.

In this study we focus primary attention on the first of these potential paths between percent organized and higher wages for union workers: the one from coverage to the elasticity of product demand to the elasticity of the derived demand for labor to union wage demands. We control, albeit imperfectly, for the possibility that the union POW schedule is positively sloped because the demand elasticity for labor is innately lower in the union sector, or is made lower by union efforts to restrict input substitution, by either holding fixed for diverse factors that might be expected to affect elasticities (e.g., the industry four-firm concentration ratio, the foreign share of the domestic market, etc.) or by studying the impact of coverage within a given industry. Despite our efforts, however, the possibility that the observed POW relationships for union workers to some extent reflect locational and factor substitution considerations cannot be dismissed.

Formal analysis

The relationship between the extent of organization in a market and potential union wage gains can be discussed formally as follows:[3] Let $\eta_x > 0$ be the elasticity of demand for the final product in the union sector; P be the percent organized/100; α be the fraction of total cost attributable to labor under unionism; $\sigma > 0$ be the elasticity of substitution between unionized labor and other factors; and $\eta_\ell > 0$ be the elasticity of demand for union members with respect to the wage. It is well known that, if the supply of capital is infinitely elastic to a sector,

(1) $\eta_\ell = \alpha\eta_x + (1-\alpha)\sigma$.

Under the highly plausible assumptions that the demand for the products of organized firms is a function of the extent of coverage [$\eta_x = \eta_x(P)$] and that increases in coverage reduce the elasticity of demand (i.e., $\eta_x' < 0$), we get

(2) $d\eta_\ell/dP = \alpha d\eta_x/dP = \alpha\eta_x' < 0$.

The union is assumed to maximize a standard convex utility function in which ln(wage rate) and ln (employment) are arguments:

(3) $U(\ln W, \ln E)$

where W is the wage rate for union members and E is their employment, subject to the demand curve for union labor. Along this curve, $d\ln E = -\eta_\ell d\ln W$. Thus, the wage which maximizes (3) must fulfill the condition

(4) $U_1/U_2 = \eta_\ell$,

where U_1 and U_2 are the partial derivatives of the maximand with respect to lnW and lnE, respectively. Equation (4) requires that the ratio of the marginal utility from an increased wage rate to the marginal utility from additional employment be equated to the elasticity of demand for union labor. To demonstrate the inverse link between W and η_ℓ, let

(5) $U_1/U_2 = \psi(W)$,

and assume that, as seems reasonable, increases in the wage reduce the ratio of the marginal utility of wages to the marginal utility of employment ($\psi' < 0$).[4] Then take the inverse function of ψ:

(6) $W = \psi^{-1}(U_1/U_2) = \psi^{-1}(\eta_\ell)$.

Differentiation of (6) with respect to P yields

(7) $dW/dP = (dW/d\eta_\ell)(d\eta_\ell/dP) = \alpha\eta_x'/\psi' > 0.$

According to (7), the slope of the curve relating union wages to the percent organized depends on labor's share of cost, the effect of coverage on the elasticity of demand for the output of organized firms, and the utility function of the union. As would be expected, the slope depends more on η_x and thus P when labor's share in cost is large than when it is small.

The dependence of the slope of the POW schedule for union workers on the relation between the percent organized and the demand elasticity for the output of unionized firms (η_x') brings the product market to the fore of the analysis. Under reasonably general assumptions about the extent of product market competition, it can be shown that η_x' will, as posited, be negative. Consider, for example, the monopolistic competition situation in which products differ across firms because of either the location of customers in regional or local markets or small differences in the firms' commoditie and where there are N equally sized firms in the industry, each of which has a cross-elasticity of demand of γ with every other firm. Ignoring, for simplicity, the effect of changes in wages and prices on total industry output, the elasticity of demand for the output of the organized sector will depend on the number of firms to whom output can be lost, and thus on the organized share of the market. In the case under consideration, the relation is a simple linear curve, with the elasticity of demand for the output of the union sector falling proportionately to the proportion of enterprises that are organized. For example, if the proportion organized in the relevant product market increases from .3 to .8, the elasticity of demand in the organized sector will decline by $.5N\gamma$.

A negative effect of coverage on the elasticity of product demand, and hence a positive effect on union wage rates, is also to be expected in oligopolistic situations, although the complexity of price-setting precludes any definitive analysis. If all firms in the sector alter prices when the "leading firms" make changes, the key issue may be whether or not the union has organized the leaders. For example, a large change in coverage which fails to alter the union status of the leaders might very well have less impact on the product demand elasticity facing the union sector than a small change that alters the union status of the leading firms. Overall, increases in coverage should increase the union wage rate, but the precise relation cannot be determined.

The effect of greater coverage on union wages in regulated industries is also not easy to ascertain with precision. When an industry is highly organized, regulators may allow roughly offsetting price increases in response to union wage increases; when it is only moderately organized, they may not. In this setting, changes in coverage which lead regulators to base price changes on union wage changes would affect union wages, while those that do not would have no such effect.

The most complex case to consider is that of homogeneous goods produced for a perfectly competitive market, defined as one in which firms with the same U-shaped cost curve can enter freely. In the long-run static equilibrium, unless there were 100 percent union coverage (or equivalently unless nonunion firms paid union wages to forestall organization), the elasticity of demand for output of the organized sector would be infinite, implying that there would be no union wage effect, in the absence of offsetting productivity gains of the type found by Brown and Medoff, Frantz, and Clark. Assuming no threat effects, the POW curve for union workers would be discontinuous, with wage rates unrelated to P until complete coverage (including the automatic coverage of new firms) was achieved, at which point they would jump sharply to a value dependent on the sector labor demand elasticity. However, since going out of business is not instantaneous and a given percent organized can be maintained in the face of competitive disadvantage by organizing new enterprises, the actual relation may be much smoother. Moreover, to the extent that unionized firms are at a competitive disadvantage (compensation differences exceeding any positive union productivity effects), the greater the percent organized the smaller will be the competitive disadvantage to those that are unionized, implying that the rate at which covered establishments go out of business is likely to be lower and the elasticity of demand for union labor smaller the greater the percent organized. This implies an upward sloping POW curve for union workers in the competitive environment posited.

Nonunion Wages

The effect of percent unionized on nonunion wages is more complex because there are some factors operating to produce a positive POW relation and others operating to produce a negative POW relation. On the one hand, union wage gains are likely to lead to a reduction of employment in the union sector and a potential increase in the supply of nonunion workers. On the other hand, union-induced increases in cost will shift demand toward nonunion enterprises, raising the demand for nonunion workers. Assuming that capital is immobile or that the nonunion sector is at least as capital intensive as the union sector,[5] and, for the moment, that employees must be employed in either the union or nonunion sector of a given market, the net effect of the two forces on nonunion wage rates must be negative, since the increase in the supply of labor will exceed the increase in the demand. This is because substitution both among inputs and among products operates on the supply side to displace union workers and augment the supply of nonunion labor while only output substitution operates to raise the demand for nonunion products and workers.

Formally, consider a work force divided into two groups, union workers and nonunion workers, and continue to assume that displaced union workers end up employed in the nonunion sector of a market. Then a given percentage increase in the wages of union workers will increase the supply of nonunion workers by $\eta_\ell(L_u/L_N)\cdot 100\% = \eta_\ell(P/1-P)\cdot 100\%$, where L_u = number of union workers in the market, L_n = number of nonunion workers in the market, and η_ℓ = elasticity of demand for union labor. The wage increase will, on the other hand, shift the demand curve for nonunion workers upward by an amount dependent on the increase in the price of the union output relative to the nonunion output and their cross-elasticity of demand. Ignoring for simplicity income effects, the demand for nonunion labor will increase by $\alpha\eta_x(P/1-P)100\%$ for a given percentage increase in wages, where, as stated above, η_x = the elasticity of demand for the output of the union sector. Thus, the wage-induced increase in the supply of workers to the nonunion sector minus the wage-induced increase in the demand for nonunion employees is

(8) $(P/1-P)(\eta_\ell-\alpha\eta_x) = (P/1-P)(1-\alpha)\sigma > 0.$

Hence, the net effect would be to produce a negative relation between the percent organized and nonunion wages.

The conclusion drawn from (8) applies when labor is tied to one product market and is thus not applicable to situations where, because of differences in the geographic locale of organized and nonorganized production, the supply effect is not operative at a product market level. If, as we expect, most displaced union workers do not end up producing the same commodity, the demand effect will most likely dominate, producing a positive POW relation for nonunion workers that does not depend on threat effects.

The likelihood that nonunion wages are affected by unionism independently of the demand and supply effects outlined above creates a more complex situation. When organization is extensive, it is highly possible that nonunion workers will observe higher union wages and seek similar rates of pay. By the "wage relativity" hypothesis, their supply price will increase and their effective supply (hours weighted by productivity) can be expected to fall unless they are given comparable wages.[6] More importantly, perhaps, the probability of organization is likely to increase, creating a pressure for higher wages. In a market subject to threats of organization, the nonunion wage will probably depend on the costs to firms of fighting organization by expending funds in NLRB election campaigns versus the costs of raising wages to reduce the monetary benefits of unionism.

II. Econometric Specification and Data

Estimates of union and nonunion POW relationships were derived by fitting equations which link the wages of union workers and establishments and of nonunion workers and establishments to the percent organized in the relevant 3-digit Census or Standard Industrial Classification (SIC) industry or, for a particular industry, to the percent of the industry's workers who are organized in a relevant geographic area. Given these estimates, we obtain an estimate of the impact of the extent of unionization in a product market on the union-nonunion wage differential (i.e., on the union wage effect) by subtracting the estimated coefficient of percent organized in the nonunion equation from the appropriate estimate in the union equation. Separate regressions have the advantage of providing estimates both of the impact of coverage on the absolute wages of organized labor and of the sign and magnitude of the sum of the demand, supply, and spillover effects on nonunion wages. The major disadvantage is the danger that, if industry controls are incomplete, omitted factors associated with both the fraction unionized and wages, especially those that affect η_{ℓ} independently of the impact of P on η_x, may produce misleading interpretations.

To isolate the effect of percent organized on wages it is necessary in the regression analysis to control for other potentially important factors. Human capital and institutional considerations highlight the need to hold age or experience, skill, and years of schooling fixed. The impact of discrimination and other factors on male-female and black-white differentials must be taken into account. Aspects of residential location, such as the cost of living, should be held constant. Because the prime independent variable in the manufacturing sector analysis, the coverage rate, relates to industries, it is of particular importance to control for other characteristics of an industry--average establishment size, concentration ratio, etc.--which may be correlated with both the percent unionized and the wage in the industry. Formally, the basic regression model for the manufacturing analysis can be written in semi-log form as:

(9) $W_{ij} = aP_j + bWORKER_{ij} + cIND_j + dEST_{ij} + U_{ij}$,

where W_{ij} = ln(wage) of the ith worker(establishment) in industry j

P_j= percent organized in the jth industry/100,

$WORKER_{ij}$ = vector of worker characteristics,

IND_j = vector of industry characteristics,

EST_{ij} = vector of establishment characteristics,

U_{ij} = independently, indentically, and normally distributed residual with mean 0, and where the sample relates to either union or nonunion production (or nonoffice) workers. The particular control variables available in the two data sets under study differ substantively: the Current Population Surveys contain detailed information on the characteristics of individuals but not on the establishments in which they work; the Expenditures for Employee Compensation surveys contain the opposite.

As stated above, the key factor in the estimation of the effect of P on W is the extent to which the relevant industry factors are held fixed. We attack this problem in three ways.

First, in each manufacturing sector regression we include several industry variables measured at the same level of aggregation as P, such as the four-firm concentration ratio, average firm size, and the fraction of domestic consumption that is imported. Inclusion of these variables guarantees that coverage is not "standing in" for their effect on wages.

Second, in the manufacturing analysis we allow for more general industry effects by including 2-digit SIC industry dummies. To the extent that the 2-digit SIC dummies capture some of the differences between more detailed industries, these controls reduce the chance that omitted industry factors bias the estimated coefficient of P.

Third, in the analysis of the nonmanufacturing sector we control for industry factors by examining the effect of coverage on wages across geographic areas within one industry, for which the relevant product markets can be reasonably defined on a regional basis. Analysis within an industry ameliorates both the potential problem of omitted industry factors that might be correlated with coverage and with wage rates and the potential problem of cross-industry differences in the elasticity of labor demand due to technological or product market factors that might bring about POW associations even in the absence of the coverage-product market linkage outlined in Section I. Nevertheless, the geographic analysis does have a potential problem: coverage may be correlated with omitted geographic variables that influence wages, leading to biased estimates of union and nonunion POW effects.

As emphasized in the preceding discussion, an omitted industry variable which is partially correlated with both percent organized and wages will bias estimated POW associations. However, it is important to note that if the omitted variable had the same effect on the wage rates in ln units of union and nonunion workers, and the same partial correlation with percent organized in the union and nonunion samples, the difference between the estimated effects of coverage on union and nonunion wages (i.e., the estimated effect of coverage on the wage differential) would be unbiased. This is because the omitted factor would bias the estimated coverage coefficients in the union and nonunion regressions by the same absolute amount, so that differencing would eliminate the bias term. Whether the unobserved variable has the same effect on the wages of union and nonunion workers and the same partial correlation with percent organized is an issue which cannot, by definition, be answered with the data. However, with a sufficiently large set of industrial controls in the union and nonunion regressions, it is difficult to think of reasons why the two conditions would be grossly violated. As long as the omitted factor has roughly comparable effects on wages in the union and nonunion samples and is similarly correlated with percent covered, differencing ought to provide at least a crude correction for the potential bias. As a result, the analysis is likely to yield better estimates of the impact of percent organized on the union-nonunion wage differential than of the separate effects of percent organized on the wages of union and nonunion workers.

The data

To estimate the effect of percent organized on wages, we have examined two sources of data: the May CPS surveys for 1973-75 and the EEC surveys for 1968, 1970, and 1972.[7] Because the EEC establishment data lack demographic information, we used the CPS responses to calculate the mean values of selected variables for union and nonunion production (blue-collar plus service) workers in each 3-digit Census industry, recoded the variables (as well as possible) to the 3-digit SIC scheme used on the EEC tapes, and added the CPS variables for either the organized or nonorganized sector of a 3-digit SIC industry to each EEC establishment's record in accordance with the establishment's 3-digit SIC industry and the collective-bargaining status of its nonoffice employees. The variables (percent male, percent white, mean schooling completed, mean of age minus schooling completed minus five, and the means of some other potentially relevant characteristics) were derived form weighted counts of private wage and salary production workers represented on the 1973-75 May CPS file.

The principal independent variable in the study, the percent of nonoffice workers organized, was estimated for 3-digit 1967 SIC industries from 1968, 1970, and 1972 EEC survey information regarding whether a majority of each responding establishment's nonoffice employees were covered by union-management agreements; these estimates are described in detail in Freeman and Medoff (1979). State-by-industry percent organized figures for private sector employed wage and salary production workers were derived with 1973-75 May CPS micro-data for the nonmanufacturing industries to be analyzed. It should be noted that the CPS tapes amalgamated states into 29 groups and that the number of observations for the state-by-industry unionization figures were often small.

Variables measuring other industry characteristics that might be correlated with both wage rates and percent organized were added to each of the union and nonunion manufacturing data sets. They are:

- average size of firm, to take account of the well-known positive effect of size on wages (Masters) and the potential positive correlation between unionism and size. This variable was measured as the value of shipments in 1972 in each 3-digit SIC industry divided by the number of firms in the industry in 1972.

- concentration ratio, to take account of the possible impact of concentration on wages (Weiss) and the likelihood that unionization is higher in more concentrated sectors. This variable was measured as a weighted average of the fractions of shipments accounted for by the four leading firms in the 4-digit SIC industries composing a 3-digit SIC industry.

- injury and illness rate, to control for the potential effect of dangerous work conditions on wages and the possibility that this rate is associated with the extent of unionism. The variable is measured as the number of lost workdays due to injury and illness per full-time worker per year.

- the ratio of imports to domestic consumption, to allow for the likely positive association between the elasticity of product demand and the extent of foreign competition. Because only the union wage is expected to depend on the elasticity of final demand for output, the import ratio (and the concentration ratio) should influence union wages but not nonunion wages by influencing this elasticity. The imports ratio was calculated as imports plus import duties divided by the quantity of shipments plus imports plus import duties less domestic exports, using data for the 5- and 7-digit product classes underlying 3-digit product classes.

The industry control variables were constructed with information from several sources and bridged as well as possible to the industrial classifications on the CPS and EEC tapes. The sources of the data and a brief description of the adjustments needed to construct useable variables are given in Appendix A. The actual data and the details of the construction of variables are available on request.

III. Empirical Results

Table 1 presents estimates based on the 1973-75 May CPS surveys of the impact of percent organized in an industry on the usual hourly earnings of manufacturing production workers who are union members and on the earnings of comparable nonmembers. The first and second columns of numbers give the mean and standard deviation of the relevant variables in the union and nonunion samples; the third and fourth columns record for the two samples the regression coefficients and standard errors on percent organized and on four other industry-level variables as well. The other controls in the equation are listed in the table.

The principal result demonstrated in Table 1 is that percent covered has a sizeable positive effect on union but not on nonunion wages, which implies an increasing wage differential with coverage. With the sample of union members, the estimated coefficient on coverage is .168 compared to .003 with the sample of nonmembers. This implies, for example, that the union wage effect in manufacturing industry with 80 percent organized is likely to be about 10 percentage points higher than in an industry with only 20 percent organized.

Table 1 also demonstrates that firm size is positively and significantly associated with wages in both union and nonunion settings; the estimated firm size coefficient is smaller under unionism. The injury and illness rate has the same statistically significant small positive effect on the earnings of union members and nonmembers. The product market variables (i.e., the concentration ratio and the imports to domestic consumption ratio) do not appear to have a systematic impact on the wages of either union or nonunion manufacturing production workers; the finding with the union sample runs counter to our expectation that these factors affect the product demand elasticity, and, hence, the labor demand elasticity, leading to higher wage rates in unionized firms.

Numerous additional experiments were conducted with the CPS data; for example, in one set of regressions similar to those presented, a capital/labor hours variable was included and, in another, the number of occupational dummies was increased from 4 to 18. Under all models fit, the conclusions concerning the impact of percent organized on union wage rates, nonunion wage rates, and the wage differential were the same as those drawn from the regressions presented in Table 1.

Table 2 provides estimates of the impact of percent organized on the hourly compensation (wages plus private fringes) and hourly wages[8] of nonoffice employees in manufacturing establishments in which a majority of the blue-collar workforce was (at the time of the relevant EEC survey) covered by

Table 1: Estimates of the Percent Organized Wage (POW) Relation for Manufacturing Production Workers; 1973-75 May CPS Individual-Level Data

	Mean (S.D.)		Dependent Variable: Ln(Usual Hourly Earnings) Coefficient (S.E.)	
	Union Sample	Nonunion Sample	Union Sample	Nonunion Sample
Independent Variables:				
Industry-Level				
Percent of industry's nonoffice workers covered by collective bargaining/100	.688 (.204)	.547 (.195)	.168 (.027)	.003 (.026)
Ln(Shipments per firm)	1.526 (1.265)	.984 (1.089)	.024 (.006)	.043 (.008)
Four-firm concentration ratio	.430 (.185)	.357 (.151)	-.013 (.031)	.045 (.040)
Imports/Domestic consumption	.069 (.053)	.063 (.051)	-.035 (.065)	.058 (.076)
Injury rate (days per worker)	.729 (.380)	.675 (.439)	.039 (.014)	.039 (.017)
Individual Level*				
Industry dummies (20); state dummies (28); occupation dummies (4); survey dummies (2); SMSA-size dummies (4); marital status dummies (3); sex dummy; race dummy; schooling completed; age-schooling completed-5 and its square; number of dependents; constant	--	--	yes	yes
R^2	--	--	.421	.471
SEE	--	--	.270	.325
N	10,679	11,204	10,679	11,204
Mean(S.D.) of Dependent Variable	1.451 (.354)	1.190 (.445)	--	--

*The four Standard Metropolitan Statistical Area - (SMSA-) size dummies included in the Table 1 regressions indicate whether an individual: lived in an SMSA whose population as of 1970 was ≥1,000,000; <1,000,000 but ≥500,000; <500,000 but still identified on the 1973-75 May CPS tapes (98 SMSA's were identified); lived in an SMSA not identified on the CPS file; or did not live in an SMSA. The marital status dummies included indicate whether the sample member was: married, spouse present; married, spouse absent; widowed or divorced; or never married.

union-management agreements and in comparable establishments in which a majority was not covered.[9] In the case of hourly compensation, the estimated coefficient of proportion organized is .152 for unionized establishments and .045 for similar ones that are non-union. These estimates imply, for example, that the union compensation effect for manufacturing nonoffice employees is about 6 percentage points larger in an industry with 80 percent organized than in one with only 20 percent organized. In the case of wages per hour worked, the estimated union POW effect is .119 and the comparable nonunion effect is .036 These estimated POW relationships indicate that the impact of coverage on the union-nonunion wage differential is smaller than suggested by the CPS data. This divergence might reflect the different unit of observation in the two surveys, the fact that a small number of very large establishments representing 15 percent of total manufacturing employment were excluded from the EEC tapes to preserve confidentiality, definitional differences in various key concepts, and/or differences in the dates of the two surveys. Nevertheless, with the EEC data, the union wage effect in percentage points for manufacturing production workers increases by about .08 of a point for each point increase in the percent organized.

The estimated coefficients on the industry control variables with the EEC samples are for the most part quite imprecise and frequently have the "wrong" sign. The estimates clearly improve when we turn from the variables derived at the industry level to those that exist on an establishment basis. Firm size, measured as ln (nonoffice hours worked in the establishment), has a positive effect on wages (and, similarly, total compensation) with both the union and nonunion samples. A variable equal to the ratio of hours worked by nonoffice employees to hours worked by all employees in the

Table 2: Estimates of the Percent Organized Wage (POW) Relation for Manufacturing Nonoffice Workers; 1968, 1970, and 1972 EEC Establishment-Level Data

	Mean (S.D.)		Coefficient (S.E.) — Dependent Variables: Ln(Hourly Compensation[a])		Coefficient (S.E.) — Dependent Variables: Ln(Hourly Wage[a])	
	Union Sample	Nonunion Sample	Union Sample	Nonunion Sample	Union Sample	Nonunion Sample
Independent Variables:						
Industry Level						
Percent of industry's nonoffice workers covered by collective bargaining/100	.714 (.214)	.510 (.224)	.152 (.036)	.045 (.040)	.119 (.029)	.036 (.039)
Four-firm concentration ratio	.380 (.177)	.339 (.143)	-.049 (.033)	-.086 (.062)	-.057 (.031)	-.092 (.060)
Imports/Domestic consumption	.073 (.068)	.064 (.157)	.055 (.078)	-.299 (.132)	.071 (.074)	-.314 (.129)
Injury rate (days per worker)	.715 (.366)	.638 (.400)	-.002 (.026)	.021 (.041)	.023 (.025)	.019 (.039)
Percent male; percent white; mean schooling completed; mean of age-schooling completed-5 and its square	--	--	yes	yes	yes	yes
Establishment Level						
Ln(nonoffice hours worked)	13.364 (1.534)	12.204 (1.749)	.037 (.003)	.044 (.004)	.031 (.003)	.037 (.004)
Nonoffice hours worked/Total hours worked	.776 (.148)	.811 (.170)	-.406 (.033)	-.395 (.042)	-.379 (.031)	-.353 (.041)
Industry dummies (20); region dummies (3); survey dummies (2); constant	--	--	yes	yes	yes	yes
R^2	--	--	.451	.500	.449	.478
SEE	--	--	.208	.234	.195	.228
N	2,576	1,465	2,576	1,465	2,576	1,465
Mean (S.D.) of Dependent Variable	-- --	-- --	1.511 (.279)	1.189 (.327)	1.427 (.261)	1.149 (.312)

Note: a. In 1972 CPI-deflated dollars.

establishment assumes a negative and very significant estimated coefficient in both the samples. This result suggests either that blue-collar jobs in nonoffice-employee intensive establishments are relatively low skill (e.g., assembly-line jobs), or that office and nonoffice workers are substitutes. Since under the second interpretation the nonoffice labor/total labor variable does not belong in the regression models, we reestimated the Table 2 equations with it excluded. While the conclusions concerning the effect of P on union wages and the wage differential drawn from this estimation are similar to those based on Table 2, the regressions indicate that P has a significant positive effect on the hourly compensation and hourly wages of nonunion as well as union employees.

A number of additional models were fit with the EEC data: one controlled for the fraction of production employees (in the union or nonunion sector of the relevant industry) in the five broad occupational groups used in creating the occupation-group dummy variables included in the Table 1 regressions; another held constant the fraction of production employees in the five SMSA-size categories, the fraction in the four marital status groups used in deriving Table 1, and mean number of dependents; another included the capital-labor hours ratio in the appropriate industry. The estimated POW relationships under each of these specifications were roughly the same as those indicated by the Table 2 results.

Thus, analysis of both CPS and EEC data for production workers in manufacturing lends general support to the notion of an upward sloping POW curve in the union sector of a market; reveals either no spillover or a small positive association between coverage and compensation for nonunion workers; and demonstrates that the union-nonunion wage differential grows substantially with growth in the extent of union organization.[10]

Construction Industry

Are the observed POW relationships specific to the manufacturing sector where product markets are likely to be national, or do similar relationships exist in nonmanufacturing industries for which product markets tend not to be nationwide in scope? To address this question we have focused on one major industry characterized by local product markets, construction, examining the relationships between the percent of the industry's workers in a state who are organized and union and nonunion wages. This experiment differs from the one for the manufacturing sector in that it holds technological and market factors fixed by looking within one industry instead of by using industry-level controls in a cross-industry analysis. While states are not the optimal unit for defining product markets in construction, the mobility of construction workers from site to site over a wide geographic area makes states less inappropriate for this sector than for other sectors. Limited sample sizes makes unionization figures for less aggregate geographic units subject to considerable potential measurement error.

As stated earlier, we tabulated from the 1973-75 May CPS tapes the fraction of employed private sector wage and salary construction production workers in each CPS state group who were union members. This variable was added to a construction worker extract from the CPS tapes. Because the CPS amalgamated states into 29 groups, there are just that number of distinct union figures.

Table 3 presents estimates of the effect of construction worker coverage in a state on the wages of union and nonunion construction workers. The primary controls employed in this analysis are: the four production worker occupation group dummies used in the CPS manufacturing sector runs; the region and SMSA-size dummies used for Table 1; and three dummy variables for the sector of the construction industry in which the worker is employed (general building contracting, general contracting, except building, and special trade contracting, with the deleted group of not specified construction). The results lend additional support to the claim that there is a positive POW

Table 3: Estimates of the Percent Organized Wage (POW) Relation for Construction Production Workers; May 1973-75 CPS Individual-Level Data

	Mean (S.D.)		Dependent Variables: Ln(Usual Hourly Earnings) Coefficient (S.E.)		Ln(Usual Weekly Earnings)	
	Union Sample	Nonunion Sample	Union Sample	Nonunion Sample	Union Sample	Nonunion Sample
Independent Variables:						
State Level						
Percent of construction workers in state who are union members/100	.531 (.180)	.386 (.172)	.279 (.056)	.065 (.066)	.180 (.056)	.017 (.076)
Industry Level						
Industry dummies (3); region dummies (3); occupation dummies (4); year dummies (2); SMSA-size dummies (4); marital status dummies (3); sex dummy; race dummy; schooling completed-5 and its square; number of dependents; constant	--	--	yes	yes	yes	yes
R^2	--	--	.275	.285	.262	.304
SEE	--	--	.289	.364	.291	.418
N			2,327	2,825	2,327	2,825
Ln(Usual Hourly Earnings)	1.919 (.337)	1.333 (.428)	--	--	--	--
Ln(Usual Weekly Earnings)	5.599 (.337)	5.002 (.502)	--	--	--	--

relationship for unionized workers. When ln(usual hourly earnings) is the dependent variable, the effect of coverage with the union sample is a sizeable .279; when ln(usual weekly earnings) is on the left, the coefficient on the coverage variable falls sharply to .180, presumably due to the effect of high hourly wages on hours worked. With the nonunion sample, by contrast, the estimated coefficient of the fraction unionized variable in both regressions, while positive, is neither substantially nor significantly greater than zero. The net result is that in construction as in manufacturing the union-nonunion wage differential grows with growth in the percent organized.

The results with the state construction industry sample are potentially sensitive to inclusion of other "state" variables due to the small number of "state" unionization figures (29). In one analysis, we added to the first two regressions presented in Table 3 the mean ln(usual hourly earnings) for male non-construction production workers in the appropriate state and found that the coefficient on percent organized dropped noticeably, to .209 (.064) in the equation for union workers and to -.050 (.083) in the equation for nonunion workers, where the figures in parentheses are standard errors. When a similarly-derived state mean of ln (usual weekly earnings) was added to the third and fourth equations represented in Table 3, the union POW effect fell to .144 (.058) and the nonunion effect to -.075 (.085). Perhaps the safest conclusion to reach is that the union-nonunion wage differential depends positively on the percentage organized. The precise magnitude of the absolute effects of percentage organized on the wages of union and nonunion workers is uncertain due to the small number of unionization observations.

Recent developments in construction suggest that our findings are reasonable in light of changes over time in coverage and the ability of construction unions to extract wage gains. In the 1970's, the percent covered in construction dropped significantly, first in residential and then in heavy construction. While definitive estimates are lacking, the decline may have been as large as 30 percentage points (from 70 percent to 40 percent). Competition with nonunion contractors has significantly reduced the union's ability to obtain large wage gains and in several cases construction unions have agreed to wage decreases or postponed negotiated increases. In Baltimore, Maryland, for example, the Associated General Contractors and Carpenters Union agreed to forego a $1.00 increase and held wages stable from 1975 to 1977 to make union contractors more competitive.[11] Because displaced union workers are likely to be hired by nonunion contractors, the drop in coverage has an important feedback effect on the market. As additional union craftsmen (who probably are more skilled on average than their nonunion peers) find jobs with nonunion employers, the ability of these contractors to compete with those who are unionized for large jobs increases, most likely raising the elasticity of demand for the union contractors' product. At a given union premium, this implies further reductions in union sector employment, a larger pool of relatively skilled former union craftsmen who are available to open-shop contractors, and so forth.

Other Nonmanufacturing Industries

Because product markets in most other nonmanufacturing industries are likely to encompass much narrower areas than states, it is difficult to analyze the within-industry cross-market relationship between percent organized and union and nonunion usual hourly earning for these sectors. An appropriate market for, say, retail trade, is likely to be a city or Standard Metropolitan Statistical Area, for which reliable unionization figures at the industry level are unavailable. Despite the weakness of the experiment, however, some information about POW curves outside the construction and manufacturing sectors can be obtained by analyzing the effect of within-industry cross-state unionization on union and nonunion earnings in industries in other sectors which are characterized by local product markets.

Accordingly we estimated the effect of the percentage organized in a state-industry cell on the usual hourly earnings of union and nonunion production (blue collar plus service) workers in the following 3-digit 1970 Census industries: eating and drinking places; grocery stores; real estate, including real estate-insurance-law offices; auto repair and related services; hotels and motels; miscellaneous entertainment and recreation services; and hospitals.

The results of these within-industry analyses were quite mixed. For instantce, in "auto repairs and related services" (where P = .074), the estimated coefficient (standard error) in a ln(usual hourly earnings) regression identical to those in Table 3 was .746 (.749) for union production workers and .375 (.251) for the comparable nonunion sample, while in "eating and drinking places" (where P = .091), the estimate was .403 (.319) with the union sample and .615 (.132) with the non-union sample. Although the estimated percent organized coefficient was positive for both union and nonunion workers in all seven industries, it was statistically significant at the .05 level less than half of the time. The unweighted mean estimated proportion organized effect was .484 with the union samples and .313 with the nonunion samples. The union wage premium grew with organization in five out of seven industries.

These findings suggest that the union-nonunion wage differential outside of manufacturing and construction is positively related to the percent organized and raise the possibility that nonunion as well as union wages may be higher where coverage is high than where it is low. Whether these tentative conclusions will stand up on more disaggregate analyses of POW relations in industries characterized by local markets remains to be seen.

Conclusions

The results presented in this study indicate that in U.S. manufacturing the percent organized in a product market has a strong positive association with the wages of union members but not with the wages of nonmembers, making the union wage differential a positive function of the extant of unionization. For industries characterized by local markets, the percent of an industry's workers in a geographic area who are organized appears also to raise the union wage differential, but the effects on the wages of union and nonunion workers separately are less clear. Thus, the findings suggest that the percentage organized is an important determinant of "union power," a conclusion that could not be drawn from standard union wage studies, which either relate earnings to a union dummy variable but not a union coverage variable (with a file of data on individuals) or relate average earnings in an industry to coverage (with a file of data on industries). In addition, they suggest that traditional studies of the union-nonunion wage differential in manufacturing, which have been interpreted as providing information solely on the impact of unions on relative wages, also provide information on the impact of unions on absolute wages, an issue of considerable concern (see Lewis, pp. 1-2, p. 16). That is, since coverage appears to have a nonnegative effect on nonunion wages in manufacturing, it can be inferred that traditional estimates of the union wage effect give lower bounds for the impact of unions on the absolute wages of covered workers.

In our theoretical discussion, we argued that a positive association between percent organized and the hourly compensation of union members was likely to reflect the impact of union coverage in a product market on the availability of nonunion substitutes and thus on the derived demand elasticity for union workers; a lower elasticity of labor demand would lower the cost in terms of lost membership for a given wage increase, leading unions to make larger wage demands. This is not, however, the sole possible explanation of the observed regression results. Coverage could be positively related to both the demand elasticity for labor and union wage gains without being related to the demand elasticity for union-made products. This is because the elasticity of labor demand depends on the elasticity of substitution (σ) and labor's share of cost (α) as well as on the product demand elasticity (η_x), raising the possibility that the results could reflect the impact of coverage on σ or α. While it is difficult to argue *a priori* that increased coverage in a product market *causes* a reduction in σ or change in α that lowers η_ℓ, it is quite possible that unions locate and survive where labor's share of costs and the elasticity of substitution are such as to produce an innately small η_ℓ. Unfortunately, it is not possible to sort out the relative importance of the determinants of η_ℓ -- σ, α, and η_x -- in bringing about the strong positive POW relationship we have observed. However, the reader should be reminded that the construction industry analysis (and the other nonmanufacturing analyses) was (were) done for a particular industry. It seems to us much more difficult to explain a *within* industry positive relationship between percent organized and the union-nonunion wage differential in terms of cross-state variation in the technological parameters σ and α than in terms of cross-state differences in the demand elasticity for the relevant locally-traded product.

Finally, it should be stressed that this paper has taken only a first step toward providing a more refined analysis of union wage effects in terms of the market conditions likely to influence union wage gains. There are several potentially fruitful and important directions for union wage studies to go: direct evidence concerning the impact of P on η_x and of η_x on η_ℓ can be sought;

other aspects of the collective bargaining relation, in addition to coverage, such as those studied by Donsimoni, can be used to obtain a richer picture of where unions have the greatest ability to raise wages; studies of the distribution of union organizing resources and the stated logic of those who seek to extend coverage will shed light on the degree to which representatives of the labor movement believe that P affects the wage differential through its effect on the elasticity of demand for labor, and, finally, analysis of the compensation policies of nonunion firms in industries where there appears to be a positive nonunion POW relationship will facilitate a decomposition of the effect of coverage on nonunion wages into the portion associated with the threat of unionization, the portion attributable to increased demand for nonunion products, and the portion due to an increased pool of available labor. Taken together, these studies would help explain the nature of union power.

Footnotes

[1]Throughout our theoretical discussion we refer to a product market, which is the appropriate unit of observation for an analysis of POW relationships. However, in the empirical work we focus on either a 3-digit Standard Industrial Classification or Census Industry or a Current Population Survey state group. Unfortunately, the data used do not permit a closer correspondence between the theoretical and empirical parts of our study.

[2]For an early attempt to disentangle this relation, using average wages in an industry and percent organized, see Rosen.

[3]While the model to be presented is quite simplistic in that it abstracts from union-management bargaining and other factors likely to determine wage rates under unionism, it does capture the critical fact that, all else the same, unions can increase wages at a lower price in terms of lost employment where the product demand elasticity is low.

[4]This will be the case as long as neither W nor E is a Giffen good.

[5]See Johnson and Mieskowski for a two sector model which emphasizes the importance of induced changes in capital-labor ratios in the union and nonunion sectors.

[6]For a discussion of the relativity hypothesis see Dunlop and Hicks, pp. 66-72.

[7]The CPS survey and EEC surveys are described in detail in U.S. Department of Labor, pp. 5-23 and pp. 175-183, respectively.

[8]Hourly compensation was defined as the ratio of the quantity total gross payroll plus employer expenditures for: life, accident, and health insurance plans; pension and retirement plans; vacation and holiday funds; severance pay and supplemental unemployment benefit funds; savings and thrift plans; and other private welfare plans to the quantity total hours

paid for minus vacation hours minus holiday hours minus sick leave hours minus civic and personal leave hours. The hourly wage variable was constructed by substituting total gross payroll for total compensation in the ratio just described.

[9] With the nonoffice labor/total labor variable excluded from the Table 2 regressions the estimated coefficient (standard error) of percent organized was .136 (.032) in the union hourly compensation regression and .079 (.041) in the comparable nonunion regression. For hourly wages, the effect was .116 (.030) with the union sample and .067 (.039) with the nonunion sample.

[10] Efforts at estimating the shape (second derivatives) of the various POW curves yielded imprecise results, which varied with the model estimated, presumably because of the collinearity between percent organized and its square.

[11] See R.D. Cochran, in which letters between the contractors and the local are presented.

Boskin, M.J., "Unions and Relative Real Wages," American Economic Review, June 1972, 62, 466-472.

Brown, C. and J. Medoff, "Trade Unions in the Production Process," Journal of Political Economy, June 1978, 86, 355-378.

Clark, K. "Unions and Productivity in the Cement Industry," Ph.D. Thesis, Harvard University, 1978.

Cochran, R.D., "The Construction Industry in Baltimore: Open Versus Closed Shop", mimeographed Economics 1650 term paper, Harvard University, 1978. A shorter version appears as "The Construction Industry in Baltimore," Harvard College Economist, Cambridge, MA (October 1978).

Donsimoni, M.P., "An Analysis of Union Power: The Role of the Structure and Conduct of the Labor Movement," Ph.D. thesis, Harvard University, 1978.

Dunlop, J.T., The Eighth Monash Lecture, 1974, "Inflation and Income Policy: The Political Economy of Recent U.S. Experience," Monash University, Australia, September 9, 1974.

Frantz, J., "The Impact of Trade Unions on Production in the Wooden Household Furniture Industry," Senior Honors Thesis, Harvard College, 1976.

Freeman, R.B., "The Effect of Unions on Fringe Benefits," National Bureau of Economic Research, June 1978.

Freeman, R.B. and J. L. Medoff, "Substitution Between Production Labor and Other Inputs in Unionized and Non-Unionized Manufacturing," Review of Economics and Statistics (forthcoming).

Freeman, R.B. and J.L. Medoff, "New Estimates of Private Sector Unionism in the United States," Industrial and Labor Relations Review, January 1979.

Hicks, J.R., The Crisis in Keynesian Economics (New York: Basic Books, 1974).

Johnson, G. and K. Youmans, "Union Relative Wage Effects by Age and Education," Industrial and Labor Relations Review, Jan. 1971, 24, 171-179.

Lewis, H.G., Unionism and Relative Wages in the United States (Chicago, University of Chicago Press, 1963).

Masters, S., "Wages and Plant Size: An Interindustry Analysis," Review of Economics and Statistics, Aug. 1969, 51, 341-345.

Rosen, S., "Trade Union Power, Threat Effects, and the Extent of Organization," Review of Economic Studies, April 1969, 36, 185-196.

U.S. Department of Labor, Bureau of Labor Statistics, BLS Handbook of Methods for Surveys and Studies, Bulletin 1910, (Washington, D.C.: U.S. Government Printing Office, 1976.

Weiss, L.W., "Concentration and Labor Earnings," American Economic Review, March 1966, 56, 96-117.

</ref_section>

APPENDIX A

Derivation of Industry Variables

1. EEC Collective Bargaining Coverage and CPS Union Membership Percentages

For a detailed discussion of these estimates see Freeman and Medoff (1979).

2. Concentration Ratios

The concentration ratios, which give the fraction of industry shipments accounted for by the four leading firms in the industry in 1972, were taken from the U.S. Department of Commerce, Bureau of the Census, 1972 Census of Manufactures, Special Report Series, Concentration Ratios in Manufacturing, October 1973, Table 5. These ratios, which were presented for 4-digit 1972 SIC industries, were first bridged to 4-digit 1967 SIC industries; then, using shipments to derive industry weights, the ratios were aggregated to the 3-digit 1967 SIC level and bridged to 3-digit 1970 Census industries for use with the CPS data set. Some ratios were adjusted with the factors used by Weiss, to account for the market power of firms selling to closed local markets.

3. Firm Size

The average firm size variable was calculated as ln(value of industry shipments/number of firms in the industry) for 1972. Value of shipments and number of firms were taken from the same source as the concentration ratios.

4. Imports Ratio

The imports ratio, which gives the ratio of the value of imports to the value of total domestic consumption in 1971, was calculated as:

$$\frac{IMP + DTY}{(SHP - EXP + IMP + DTY)}$$

where IMP is the value in the foreign country of imports for U.S. consumption,

DTY is the calculated import duty,

SHP is the value of manufactures' shipments,

and EXP is the value at port of exports of domestic merchandise.

The data used in creating this variable were taken from the U.S. Bureau of the Census, U.S. Commodity Exports and Imports as Related to Output, 1971 and 1970, January 1974, Table 2C. These values were available at the level of 5- and 7-digit product classes. Only the classes for which all values were available were used in computing the ratios for 3-digit groups.

5. Injury and Illness Rates

Lost workdays incidence rates (per 100 full-time private sector workers) for 1972-74 were taken from the U.S. Department of Labor, Bureau of Labor Statistics, Occupational Injuries and Illnesses by Industry, 1972, Bulletin 1830, 1974; Occupational Injuries by Industry, 1973, Bulletin 1874, 1975; and Occupational Injuries and Illnesses by Industry, 1974, Bulletin 1932, 1976.

The incidence rates (taken from Table 2 in each Bulletin) were bridged to 3-digit 1970 Census industries using average annual employment figures (from Table 1 in the Bulletins) to derive weights. Simple averages of the 1972 to 1974 incidence rates multiplied by 100 were used in the regressions.