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THE IMPACT OF THE PERCENTAGE ORGANIZED ON UNION AND NONUNION WAGES

Richard B. Freeman and James L. Medoff*

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, the 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 percentage 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 percentage organized on the union wage differential (the difference between the natural logarithms of union and of nonunion wages) depends on the relative magnitudes of the likely positive impact on union wages and the positive or negative impact on nonunion wages.

This paper seeks to disentangle the relation between the percentage of workers organized in a product market and the wages received by union workers and by nonunion workers. In contrast to most of the literature on the union wage effect, which either relates some average of wages in an industry to the percentage organized or which relates the wages of individuals to their membership in a union, our analysis examines the impact of the percentage organized on the compensation of union labor and nonunion labor taken separately.[2] By relating the wages of unionized workers to the percentage covered by collective bargaining in the relevant product market, we estimate directly the extent to which unionized workers in highly organized markets receive higher wages than unionized workers in less organized settings. By relating nonunion wages to the percentage covered, we provide direct estimates of the extent to which, as a result of threat, demand, and supply effects, nonunion workers in highly organized markets receive higher or lower wages than nonunion workers in less organized industries or areas.

Two sets of data are used in the study: information on individuals from the 1973, 1974, and 1975 May Current Population Surveys (CPS), which contain data on usual weekly earnings, 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 Surveys (EEC), which contain data on the components of compensation, labor hours, collective bargaining coverage, and some relevant establishment characteristics. The availability of both individual and establish-

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[1] Throughout our theoretical discussion we refer to a product market, which is the appropriate unit of observation for an analysis of the relationship between percentage organized and wages. However, in the empirical work we focus on either a 3-digit Standard Industrial Classification or Census industry (in the manufacturing analysis) or a Current Population Survey state group (in the construction analysis). 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 percentage organized, see Rosen (1969). For more recent related analyses, see (in alphabetical order) Donsimoni (1978), Hendricks (1975), Kahn (1978), and Lee (1978). For a trenchant treatment of the analysis, see Lewis (1980).

ment 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 our attention to production or nonoffice workers. Because of the distinct features of wage-setting in the public sector, we further limit our focus to private wage and salary workers. Finally, since the analytic model relates collective bargaining coverage to wages through the product demand curve, we let the nature of our analysis depend on the nature of the relevant product markets. We compare union organization and wages across *industries* in manufacturing, where as a first approximation product markets can be taken to be national, but compare organization and wages across *geographic areas* in construction where product markets tend to be local in nature.

The major finding of our research is that in manufacturing the percentage organized in a product market has a strong positive association with the wages of union (production) workers, but either no association or a weak positive association with the wages of nonunion (production) workers. As a result, within manufacturing, the union-nonunion wage differential in a product market is positively related to the extent of organization. For construction, the results appear to be similar, though sensitive to specification.

I. The POW Relationships

There are three basic reasons for expecting the percentage 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 unionized 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 their members, as well as with their members' wages, they can be expected to press for higher wages in markets with a relatively high percentage organized, and hence, with a relatively low labor demand elasticity. Second, a positively sloped percentage organized/wage (POW) curve for covered workers might also be observed because unions have located only in sectors with low elasticities of demand for labor or because, no matter where they have organized, they have survived only where the demand elasticity is low. In these scenarios, a low wage elasticity of demand causes a high percentage organized and, at the same time, allows the union to obtain high wages; the percentage organized 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 the lines suggested by Freeman and Medoff (forthcoming(a)), a positive percentage organized wage relationship may result. In this case the causality is from percentage organized to the elasticity of substitution (rather than to the product demand elasticity) and then to the elasticity of demand for labor and to the wage.

The present study focuses primarily on the path from percentage organized 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 possibilities that the union POW schedule is positively sloped because the demand elasticity for labor is lower to start with in the union sector or is made lower by union efforts to restrict input substitution; this is done either by holding fixed for a set of variables capturing potentially relevant industry characteristics or by studying the impact of percentage organized within a given industry. Despite our efforts, however, the possibility that the observed POW relationships for union workers to some extent reflect the locus of unionism and the degree of factor substitution in different settings 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

[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. For similar discussions of the determinants of union wage gains, see the articles by Segal (1964) and Levinson (1967).

in the union sector, P be the percentage organized/100, α be labor's share of total cost in unionized firms, $\sigma > 0$ be the elasticity of substitution between unionized labor and other factors, and $\eta_l > 0$ 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,

$$\eta_l = \alpha\eta_x + (1 - \alpha)\sigma. \tag{1}$$

Under the highly plausible assumption that the demand elasticity for the products of unionized firms is a function of the extent of coverage $[d\eta_x/dP < 0]$, and holding σ constant for simplicity, we get

$$d\eta_l/dP = \alpha d\eta_x/dP = \alpha\eta'_x < 0. \tag{2}$$

The union is assumed to maximize a standard convex utility function

$$U = U(\ln W, \ln E), \tag{3}$$

where W is the wage rate for union members and E is their employment, which together represent a point on the demand curve for union labor. Along this curve, $d \ln E = -\eta_l d \ln W$. Thus, the wage which maximizes (3) must fulfill the condition

$$U_1/U_2 = \eta_l, \tag{4}$$

where U_1 and U_2 are the partial derivatives of the maximand with respect to ln W and ln E, respectively. Since U_1 and U_2 depend only on the wage rate, we can derive the relationship between ln W and η_l by first writing the expression

$$U_1/U_2 = \psi(\ln W) \text{ with } \psi' \text{ assumed } < 0, \tag{5}$$

and then taking its inverse, which gives

$$\ln W = \psi^{-1}(U_1/U_2) = \psi^{-1}(\eta_l). \tag{6}$$

Differentiating (6) we see that

$$\begin{aligned} d \ln W/dP &= (d\psi^{-1}/d\eta_l)(d\eta_l/dP) \\ &= \alpha\eta'_x/\psi' > 0. \end{aligned} \tag{7}$$

The dependence of the slope of the POW schedule for union workers on the relation between the percentage organized and the demand elasticity for the output of organized 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, a product market in which commodities differ across firms because of either the location of customers in regional or local markets or small differences in the firms' commodities 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 example under consideration, the relation is a simple linear curve, with the elasticity of demand for the output of the organized sector falling proportionately to the fraction of enterprises that are organized. For example, if the fraction organized in the relevant product market increases from 0.3 to 0.8, the elasticity of demand in the organized sector will decline by $0.5N\gamma$.

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 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% 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, for example, offsetting productivity gains of the type found by Brown and Medoff (1978), Frantz (1976), and Clark (1978), or quasi-rents due to differences in costs not associated with unionism. Assuming no threat effects and instantaneous market adjustment, 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 demand elasticity for labor in the sector. However, since going out of business is not instantaneous and a given wage effect can be maintained with less than complete coverage by organizing additional enterprises, the actual relationship may be much smoother. Moreover, the greater the percentage organized, the smaller will be the degree of potential nonunion competition faced by union enterprises and, hence, the slower will be the likely speed at which covered establishments go out of business. For this reason, the elasticity of demand for union labor is likely to be smaller the greater the percentage organized. Hence, we can expect an upward sloping POW curve for union workers in the competitive environment posited.

In all of our calculations we take the percentage organized as the independent variable affecting wages. While there is almost always some degree of joint determination of economic variables, we believe this is the most sensible line of causality for two reasons. First, there is no basis in theory or fact (that we know of) for the expectation that wage levels would be positively associated with union organization. Second, even though unions have an incentive to organize sectors where, for some reason, there are substantial rents and, hence, the potential of bargaining for a large wage increase without a reduction in employment, management in these settings is likely to have the resources (as well as the will) to fight unions' efforts.

Nonunion Wages

The effect of percentage organized on nonunion wages in a given product market (with a given standard production function and a given price of capital set by the overall economy) is more complicated because there are some factors operating to produce a negative POW relation and others operating to produce a positive 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 product demand toward nonunion firms, raising the demand for nonunion workers. Assuming for the moment that labor must be employed in either the union or nonunion sector of the product market under analysis (but that capital can move to other product markets), the net effect of the two forces on nonunion wage rates is likely to be negative, since the increase in the supply of labor can be expected to 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 workers.[4]

While the result just stated holds when labor is tied to one product market, it most likely will not apply when, because of differences in the geographic locale of organized and nonorganized production, the supply effect is not operative at a product market level. Since displaced union workers generally do not end up producing the same commodity, the demand effect will usually dominate, putting an upward pressure on nonunion wage rates which does not depend on the "threat" of unionization.

The likelihood that nonunion wages are affected by unionism independently of the demand and supply effects outlined above creates a still

[4] Formally, under our assumptions, the percentage decline in employment in the union sector of the industry is $-(\alpha\eta_X + (1 - \alpha)\sigma)\dot{W}$, where α is labor's share of output in the union sector, η_X is the price elasticity of output demand faced by unionized firms, σ is the elasticity of substitution in the union sector, and $\dot{W}$ is the percentage increase in union wages. Therefore, the percentage increase in the supply of labor to the nonunion sector will be $-(\alpha\eta_X + (1 - \alpha)\sigma)(E^u/E^n)\dot{W}$, where E^u is labor employed in the union sector and E^n is labor employed in the nonunion sector of the industry. Assuming that $\eta_X < \infty$ and that an increase in the union sector output price simply shifts demand to the nonunion sector, the cross-price elasticity of demand for nonunion output with respect to the union price will be equal to the own price elasticity of demand for union output, η_X. Thus, the percentage increase in the demand for labor in the nonunion sector will be $\alpha\eta(E^u/E^n)\dot{W}$. Note that this analysis differs from that offered by Johnson and Mieskowski (1970). In their 2-sector general equilibrium model, the nonunion sector must employ both the displaced labor and the displaced capital, which causes the price of capital, as well as the price of labor, to adjust. In our partial equilibrium analysis, displaced capital may be absorbed by other industries as well as by the nonunion sector of the "small" industry under consideration; hence capital's price is unaffected by changes in that industry.

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. More importantly, perhaps, the probability of organization is likely to increase, which is likely to induce nonunion employers to pay higher wages.

II. Econometric Specification and Data

To isolate the effect of the extent of organization on union wages and on nonunion wages it is necessary in the regression analysis to control for other potentially important factors, such as age or experience, skill, years of schooling, residential location, sex and race. Because the prime independent variable in the manufacturing analysis, the percentage organized, relates to the environment of firms in an industry, it is particularly important to control for other related characteristics—establishment size, concentration ratio, and so on—which may be correlated with both the percentage organized and wage rates. With all of these factors in mind, we write the basic regression model for the manufacturing analysis as

$$W_{ij} = aP_j + UNIT_{ij} + cIND_j + U_{ij}, \qquad (8)$$

where W_{ij} is the ln(*wage*) of the i^{th} worker (establishment) in industry j; P_j is the percentage organized in the j^{th} industry/100; $UNIT_{ij}$ is a vector of worker or establishment characteristics (depending on the data set); IND_j is a vector of industry characteristics; U_{ij} is an independently, identically, and normally distributed residual with mean 0; and where the sample includes 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, a 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 the manufacturing sector regressions we include some industry-level variables, such as the four-firm concentration ratio, average firm size, and the injury rate in the industry. Inclusion of these variables guarantees that P is not "picking up" their effect on wages.

Second, in the manufacturing analysis we allow for more general industry effects by including 2-digit Standard Industrial Classification (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 construction sector we control for industry factors by examining the effect of coverage on wages across geographic areas *within* one industry, whose relevant product markets can be reasonably defined on a regional basis. While analyzing geographic areas within a given industry solves the problem of uncaptured industry factors, this approach also has a potential problem: organization 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 percentage organized and wages will bias estimated POW associations. However, it is important to note that (other variables held constant) if the omitted variable had the same effect on the wage rates (in ln units) of union and nonunion workers and the same relationship with percentage 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 relationship with percentage organized is an issue which cannot, by definition, be answered with the data. However, with a sufficiently large set of industry controls in the union and nonunion regressions, it is difficult to think of reasons why the two necessary conditions for unbiased differences in POW effects would be grossly violated.[5]

The Data

To estimate the effect of percentage organized on wages, we have examined two sources of data: the May CPS for 1973–75 and the EEC for 1968, 1970, and 1972.[6] 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 unorganized 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 (percentage male, percentage white, mean schooling completed, mean of age minus schooling completed minus five, and the means of some other potentially relevant characteristics) were derived from weighted counts of private wage and salary production workers represented on the 1973–75 May CPS file.

[5] A priori one cannot readily determine the possible direction of the bias. Longitudinal studies of the union wage effect show smaller differentials than do cross-sectional studies, which suggests that the cross-sectional effects may be upwardly biased due to unobserved personal characteristics. It might be inferred from this fact that our cross-sectional POW effects are biased upward as well. However, we do not believe that this problem can be addressed in an acceptable fashion in the absence of longitudinal data; "correcting" for sample selectivity with a cross-sectional data set like the one we employ by adding an inverse Mills ratio or fitting a system of equations, including a unionization equation, does not appear to yield useful results for analysis of union effects. See Freeman and Medoff (forthcoming (c)).

[6] The CPS and EEC Surveys are described in detail in U.S. Department of Labor (1976a), pp. 5–23 and pp. 175–183, respectively. Note that because of the nature of the CPS and EEC sampling techniques, pooling across years leads to replication of persons or establishments, which could bias our standard errors. However, our work and that of Lewis (1980) yield results comparable to those in the text for the different years taken by themselves. The insensitivity of the results from one year to the next, despite significant changes in macroeconomic conditions, suggests that the POW effects are relatively invariant to economic conditions, at least during the period under analysis.

The principal independent variable in the study, the percentage of nonoffice workers organized, was estimated for 3-digit 1967 SIC manufacturing industries from 1968, 1970, and 1972 EEC 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).

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

—Average size of firm (added only to the CPS records, since the EEC provides information on establishment size), to take account of the well-known positive effect of size on wages (Masters, 1969) and the potential positive correlation between extent of 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. It was calculated with data from U.S. Department of Commerce, table 5.

—Concentration ratio, to take account of the possible impact of concentration on wages (Weiss, 1966) and the likelihood that unionization is higher in more concentrated sectors. This variable was measured as a weighted average of the fractions of shipments in 1972 accounted for by the four leading firms in the 4-digit SIC industries composing a 3-digit SIC industry. It was derived with data from U.S. Department of Commerce, table 5 and with the scaling factors used by Weiss to correct for the market power of firms selling in closed local markets.

—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. A variable, equal to the mean number of lost workdays due to injury and illness per full-time worker in 1972, 1973, and 1974 in each 3-digit SIC industry, was calculated with data from U.S. Department of Labor (1974, 1975, and 1976b).

The industry control variables were bridged as well as possible to the industrial classifications on the CPS and EEC tapes. The details of the construction of variables are available on request, as are the data employed in our analyses.

III. Empirical Results

Table 1 presents estimates, based on the 1973–75 May CPS, of the impact of percentage organized (covered by collective bargaining) 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 means and standard deviations 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 percentage organized and on three other industry-level variables. The other controls in the equation are listed in the table.

The principal result demonstrated in table 1 is that percentage organized has a sizable 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 organization is 0.153 compared to 0.004 with the sample of nonmembers. This implies, for example, that the union wage effect in a manufacturing industry with 80% organized is likely to be about 9 percentage points higher than in an industry with only 20% organized. *At a given set of means*, with 20% organized, the union wage effect is about 5 percentage points; with 80% organized the comparable effect is about 14 points.

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 roughly the same statistically significant positive effect on the earnings of union members and nonmembers. The concentration ratio does not appear to have a

TABLE 1.—ESTIMATES OF THE PERCENTAGE 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				
Percentage of industry's nonoffice workers covered by collective bargaining/100	0.687 (0.208)	0.546 (0.199)	0.153 (0.024)	0.004 (0.024)
ln(*shipments per firm*)	1.526 (1.265)	0.984 (1.089)	0.025 (0.006)	0.044 (0.008)
Four-firm concentration ratio	0.430 (0.185)	0.357 (0.151)	−0.012 (0.031)	0.045 (0.040)
Injury rate (days per worker)	0.729 (0.380)	0.675 (0.439)	0.041 (0.014)	0.037 (0.016)
Individual Level[a]				
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	—	—	0.421	0.471
SEE	—	—	0.270	0.325
N	10,679	11,204	10,679	11,204
Mean(S.D.) of Dependent Variable	1.451 (0.354)	1.190 (0.445)	—	—

[a] 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 SMSAs 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.

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 concentration affects the product demand elasticity, and, hence, the labor demand elasticity, leading to higher wage rates in unionized firms.[7]

We have also conducted comparable analyses using a percentage organized variable constructed from CPS data on individuals' union membership status. The results obtained were qualitatively similar to those given in table 1.[8] We focus on the results with the EEC percentage organized measure because we believe coverage is a better indicator of union strength than membership.

Table 2 provides estimates of the impact of percentage organized (covered by collective bargaining) on the compensation (wages plus all voluntary fringes) per hour and "direct payments" (wages plus those voluntary fringes which are paid directly to workers rather than to a fund) per hour 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 union-management agreements and in comparable establishments in which a majority was not covered. In the case of hourly compensation, the estimated coefficient of proportion unionized is 0.157 for covered establishments and 0.047 for ones that are not covered. 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% organized than in one with only 20% organized. In the case of direct payments per hour worked, the estimated union POW effect is 0.125 and the comparable nonunion effect is 0.038.[9] These estimated POW relationships indicate that the impact of industry organization on the union-nonunion direct payments 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% of total manufacturing employment were excluded from the EEC tapes to preserve confidentiality, definitional differences in various key variables, and/or differences in the dates of the two surveys. Nevertheless, with the EEC data, the union direct payments effect in percentage points for production workers in the average manufacturing industry increases by about 0.08 of a point for each point increase in the percentage of the industry's production workers who are unionized. *At a given set of means,* with 20% organized, the estimated union-nonunion direct payments differential is 5 percentage points; at 80% organized the comparable differential is 10 points.

The estimated coefficients on the industry control variables with the EEC samples are for the most part quite imprecise. 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 (*non-office hours worked in the establishment*), has a positive effect on wages (and, similarly,

[7] 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 added, in another the number of occupational dummies was increased from 4 to 57, and in a final one an imports to domestic consumption ratio was included. While the inclusion of the imports variable seemed like the most important change, since it was expected that this variable would be related to an industry's product demand elasticity, the variable did not perform in accordance with the theory set out in section I. However, none of the modifications just described had a meaningful impact on the main results given in table 1. In addition, separate regressions, identical to those in table 1, were fit for white males, black males, white females, and black females. For each group except black males, the conclusions concerning the impact of percentage organized on union wage rates, nonunion wage rates, and the wage differential were roughly the same as those drawn from the regressions presented in table 1. In the case of black males, percentage organized had a small insignificant effect on union as well as nonunion wage rates, for reasons which are not apparent to us.

[8] Regressions otherwise identical to those in table 1 yielded the following coefficients (standard errors) on the extent of membership variable: 0.238 (0.041) with the union sample, and −0.013 (0.036) with the nonunion sample. The CPS membership percentages used in these regressions are discussed at length in Freeman and Medoff (1979).

[9] We also estimated models in which the dependent variable was ln(*straight-time pay per straight-time hour*) whose mean (standard deviation) was 1.403 (0.252) with the union sample and 1.166 (0.296) with the nonunion sample. With this dependent variable, the estimated POW coefficient (standard error) was 0.140 (0.028) for unionized establishments and 0.063 (0.039) for those that are nonunion. While these results are consistent with our primary conclusion that the union-nonunion compensation differential grows substantially with the percentage organized, it should be noted that they indicate a non-trivial positive relationship between the extent of organization and the nonunion hourly straight-time pay component of hourly compensation.

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 establishment assumes a negative and very significant estimated coefficient in both the samples.[10] This result suggests either that blue-collar jobs in nonoffice-employee intensive establishments are relatively low skill (e.g., assembly-line jobs) or have relatively desirable nonpecuniary charac-

[10] With the nonoffice hours/total hours variable excluded from the table 2 regressions, the estimated coefficient (standard error) of percentage organized was 0.140 (0.031) in the union hourly compensation regression and 0.082 (0.041) in the comparable nonunion regression. For hourly direct payments, the effect was 0.110 (0.029) with the union sample and 0.069 (0.040) with the nonunion sample. We also experimented with using a percentage union member variable from the CPS as our measure of unionism. The coefficients of this variable, in analogues to the table 2 regressions, were: in the ln(*compensation*) model, 0.192 (0.057) with the union sample and −0.116 (0.086) with the nonunion sample; and in the ln(*hourly direct payments*) model, 0.141 (0.054) with the union sample and −0.111 (0.084) with the nonunion sample. We also fit a number of additional models: 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; two others included the capital-labor hours ratio or the ratio of imports to domestic consumption. The estimated POW relationships under each of these specifications were roughly the same as those indicated by the table 2 results.

TABLE 2.—ESTIMATES OF THE PERCENTAGE ORGANIZED WAGE (POW) RELATION FOR MANUFACTURING NONOFFICE WORKERS: 1968, 1970, AND 1972 EEC ESTABLISHMENT-LEVEL DATA

			Dependent Variables:			
			ln(*Hourly Compensation*[a])		ln(*Hourly Direct Payments*[a])	
	Mean (S.D.)		Coefficient (S.E.)			
	Union Sample	Nonunion Sample	Union Sample	Nonunion Sample	Union Sample	Nonunion Sample
Independent Variables:						
Industry Level						
Percentage of industry's nonoffice workers covered by collective bargaining/100	0.714 (0.214)	0.510 (0.224)	0.157 (0.030)	0.047 (0.040)	0.125 (0.028)	0.038 (0.039)
Four-firm concentration ratio	0.380 (0.177)	0.339 (0.143)	−0.050 (0.033)	−0.096 (0.062)	−0.058 (0.031)	−0.103 (0.060)
Injury rate (days per worker)	0.715 (0.366)	0.638 (0.400)	−0.006 (0.026)	0.038 (0.040)	0.018 (0.024)	0.036 (0.039)
Percentage male; percentage 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)	0.037 (0.003)	0.044 (0.004)	0.031 (0.003)	0.037 (0.004)
Nonoffice hours worked/total hours worked	0.776 (0.148)	0.811 (0.170)	−0.405 (0.033)	−0.400 (0.042)	−0.377 (0.031)	−0.358 (0.041)
Industry dummies (20); region dummies (3); survey dummies (2); constant	—	—	yes	yes	yes	yes
R^2	—	—	0.451	0.498	0.449	0.476
SEE	—	—	0.208	0.234	0.195	0.228
N	2,576	1,465	2,576	1,465	2,576	1,465
Mean (S.D.) of Dependent Variable	—	—	1.511 (0.279)	1.189 (0.327)	1.427 (0.261)	1.149 (0.312)

[a] In 1967 CPI-deflated dollars. Hourly compensation is defined as the ratio of (direct payments to workers (that is, pay for time worked; pay for vacations, holidays, sick leave, and civic and personal leave; severance pay; and nonproduction bonuses) plus employer expenditures for life, accident, and health insurance plans, pension and retirement plans, vacation and holiday funds, severance pay and supplemental unemployment benefits funds, savings and thrift plans, and other private welfare plans) to (total hours paid for minus vacation hours minus holiday hours minus sick leave hours minus civic and personal leave hours). The hourly direct payments variable is constructed by substituting total direct payments to workers for total compensation in the ratio just described.

teristics or that office and nonoffice workers are substitutes. Since under the substitution 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 not very different from 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.

In sum, 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 compensation differential grows substantially with growth in the extent of union organization.

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 percentage 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 (what amounts to) 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 a less inappropriate unit for construction than for other sectors with "local" markets. Limited sample sizes make unionization figures for less aggregate geographic units subject to considerable potential measurement error.

We did weighted counts with the 1973–75 May CPS tapes to determine the percentage of employed private 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 the percentage organized (union members) in a state on the wages of union and nonunion construction workers. The primary controls employed in this analysis are the following: four 1-digit Census production worker occupation group dummies; three region and four SMSA-size dummies; and three dummy variables for the sector of the construction industry in which the worker is employed (General Building Contractors, General Contractors Except Building, and Special Trade Contractors, with the Not Specified construction group deleted). The results lend additional support to the claim that there is a positive POW relationship for unionized workers. When ln(*usual hourly earnings*) is the dependent variable, the effect of coverage with the union sample is a sizable 0.283; when ln(*usual weekly earnings*) is on the left, the coefficient on the coverage variable falls sharply to 0.182. This presumably is due to the effect of high hourly wages on hours worked. With the nonunion sample, by contrast, the estimated coefficient of the fraction organized 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 percentage organized. At a given set of means, with 20% organized the hourly wage effect is 37 percentage points: with 80% organized the comparable effect is 50 points.

The results with the state construction industry sample are potentially sensitive to inclusion of other state (group) variables, presumably due to the smaller number of state unionization figures (29). In one analysis, we added to the first two regressions presented in table 3 the mean of ln(*usual real hourly earnings*) for male nonconstruction production workers in the appropriate state and found that the estimated coefficient (standard error) on percentage organized dropped noticeably, to 0.213 (0.063) in the equation for union workers and to −0.055 (0.083) in the equation for nonunion workers. When a

TABLE 3.—ESTIMATES OF THE PERCENTAGE ORGANIZED WAGE (POW) RELATION FOR CONSTRUCTION PRODUCTION WORKERS; MAY 1973–75 CPS INDIVIDUAL-LEVEL DATA

	Mean (S.D.)		Dependent Variables: Coefficient (S.E.)			
			ln(*Usual Hourly Earnings*)		ln(*Usual Weekly Earnings*)	
	Union Sample	Nonunion Sample	Union Sample	Nonunion Sample	Union Sample	Nonunion Sample
Independent Variables:						
State Level						
Percentage of construction workers in state who are union members/100	0.531 (0.180)	0.386 (0.172)	0.283 (0.056)	0.062 (0.066)	0.182 (0.056)	0.017 (0.076)
Individual Level[a]						
Industry dummies (3); region dummies (3); 1-digit 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	yes	yes
R^2	—	—	0.276	0.286	0.263	0.307
SEE	—	—	0.289	0.364	0.291	0.420
N			2,327	2,825	2,327	2,825
ln(*Usual Hourly Earnings*)	1.919 (0.337)	1.333 (0.428)	—	—	—	—
ln(*Usual Weekly Earnings*)	5.599 (0.337)	5.002 (0.502)	—	—	—	—

[a] See note in table 1 for description of the SMSA-size and marital status dummies.

similarly-derived state mean of ln(*usual real weekly earnings*) was added to the third and fourth equations represented in table 3, the union POW effect fell to 0.147 (0.058) and the nonunion effect to −0.077 (0.085). Perhaps the safest conclusion to reach is that the union-nonunion wage differential in construction depends positively on the percentage organized and that precise POW coefficients for union and nonunion workers cannot be estimated due to the small number of states in what amounts to weighted cross-state regression analysis.[11]

Conclusions

The results presented in this study indicate that in U.S. manufacturing the percentage 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 extent of organization. For industries characterized by local markets, in particular construction, the percentage 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 taken 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 average earnings in an industry (or area) to the extent of unionism in the industry (or area) or relate an individual's or establishment's wage rate to a union dummy variable but not an interaction of this dummy with a product market organization variable.

In our theoretical discussion, we argued that a positive association between percentage organized and the hourly compensation of union

[11] Cross-state analyses were also done for seven other industries for which product markets are more likely to be local than national. While the results were mixed, because, we believe, states (or even SMSAs) do not adequately demarcate the appropriate product market boundaries, the union wage effect was positively related to percentage organized in five of the seven settings.

members was likely to reflect the impact of union coverage in a product market on the availability of substitutes produced in nonunion settings 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. Percentage organized 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 costs (α) as well as on the product demand elasticity (η_x), raising the possibility that our results could reflect the impact of coverage on σ and α or the fact that unions seem to locate and survive where the σ and α 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 relationships we have observed. However, the reader should be reminded that the construction analysis was done for a particular industry. It seems to us much more difficult to explain a *within*-industry positive relationship between percentage 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.

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