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ABSTRACT

We investigate the impact of union strength on changes in nonunion wages and employment. The prevailing model in this area is the threat model, which predicts that increases in union strength cause increases in nonunion wages and decreases in nonunion employment. In testing the threat model, we are also testing two altematives, the crowding and complements models. In contrast to the prediction of the threat model, decreases in the percent organized (reflecting a declining union threat) are associated with increases in the nonunion wage. Furthermore, increases in union wages appear to decrease, rather than to increase, nonunion wages. Evidence on the determinants of intra-industry variation in nonunion wage premia is somewhat more consistent with the crowding model and is strikingly consistent with the complements model of union and nonunion wage determination. Further evidence on the determinants of intra-industry variation in nonunion employment is consistent with the complements model and the threat model; movements in nonunion industry employment are negatively related to changes in proxies for union strength. Thus, the combined evidence supports the complements model, but neither the threat model nor the crowding model.

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In the union literature of the past several years, considerable evidence has been marshalled to support the fact that union employment and overall union strength have declined dramatically over the 1970s and 1980s, while union wage premiums have increased just as dramatically (see, e.g., Bell, 1989; Linneman, Wachter, and Carter, 1991; Wachter and Carter, 1990, and Neumark, forthcoming). Little attention has been paid, however, to the effect of these developments on the nonunion sectors of the economy. This paper examines these developments, with the goal of providing empirical tests of alternative models of the effects of union sector developments on nonunion wages and employment.

The prevailing model of the impact of unionization on the nonunion sector has been based on the threat model. First tested by Rosen (1969), the threat model predicts that an increase in union power will cause an increase in nonunion wages. The nonunion response occurs because nonunion employers calculate that nonunion workers are less likely to unionize when the nonunion firms are independently responding to union wage increases. Wage developments in the nonunion sector therefore follow, to some extent, wage developments in the union sector; for example, an increase in union wages leads nonunion employers to increase their own wage premium, thereby moving up their demand curve and laying off workers and reducing output.

In the theoretical and empirical analysis of the union threat model in the existing literature, the level of analysis is typically the industry rather than the overall economy. Paralleling this, the union threat model has been offered as an explanation of inter-industry wage differentials in the nonunion sector (e.g., Dickens, 1986; Dickens and Katz, 1987a; Krueger and Summers, 1988). Because one goal of our analysis is to test this explanation of the nonunion wage structure, we also conduct our analysis at the industry level. Since unions or in Dunlop's terms "industrial relations systems" (Dunlop, 1944) are not neatly organized along two- or three-digit industry categories, and because threat effects of any particular union may affect firms in a number of industries ("orbits of coercive comparison" (Ross, 1948)), we conduct our analysis for both one-digit and two-digit industry classifications.

There are two principal alternatives to the threat model: the crowding and complements models. Although our equation specifications are tailored to test the threat model, we also consider the predictions of the crowding and complements models for these specifications, to clarify the interpretation of any rejections of the threat model that may arise.

The "crowding model" is labor supply driven. It is based on the view that nonunion workers are labor market substitutes for union workers. Higher union wages and any resulting lower union employment cause an outward labor supply shift in the nonunion sector. The result is lower wages and higher employment in the nonunion sector. Although nonunion labor demand may also increase if nonunion output is a substitute for union output, the prevailing view is that labor supply effects dominate labor demand effects.

The "complements model," which has received less attention in the literature, focuses on the labor demand ties between union and nonunion labor within the same industry. In this model, there are three sectors: a union sector, a "substitute" nonunion sector that competes with the union sector, and a "complement" nonunion sector that serves as a production complement to both the union and substitute nonunion sectors.

Workers in the complements sector may include nonunion workers employed in the same establishment as union workers (perhaps in other occupations), or nonunion workers in nonunion firms that act as suppliers to the union and substitute nonunion sectors. Lower union sector output would thus reduce the demand for nonunion complements workers, thereby driving down their wages and employment.

The prevailing empirical literature supports, albeit weakly, the threat model. The major papers in the area are cross-sectional tests focusing on nonunion wages as the dependent variable. Little attempt has been made to analyze whether the results support either the crowding or complements models. Nor have the alternative models been tested on time-series data, at the industry or the aggregate level.

In this paper we test the threat model in a pooled cross-section time-series data set, using the period 1973 to 1989, a period during which union wages increased, and union employment decreased dramatically. We also test the predictions generated by the other models of the relationships between the union and nonunion sectors.

Finally, we consider an additional dependent variable. The threat model has typically been analyzed in the context of threat effects on nonunion wages. We extend the analysis by considering the implications of the threat model, as well as the crowding and complements models, for nonunion employment as well. We show that by jointly considering union-nonunion interactions for both wages and employment, we can provide a sharper differentiation among the competing models.

In general, our results are inconsistent with the union threat model and the crowding model, and are largely consistent with the complements model. These conclusions come from a number of empirical tests, and hold at both the one-digit and the two-digit industry level.

Section I discusses the relevant literature and Section II presents the alternative models. The data are described in Section III and the empirical tests performed in Section IV. Section V concludes the paper.

I. Literature Review

A formal model of the threat hypothesis, in which nonunion employers pay above-market-clearing wages to deter unionization, is developed in Dickens (1986). In this model the determinants of the wage that a nonunion employer must pay to avoid unionization are the magnitude of supra-competitive rents (which may or may not exist in the long-run), and the costs of organizing. The large existing literature which ties the percent organized in an industry to the wages of nonunion workers in cross-sectional data (e.g., Freeman and Medoff, 1981; Kahn, 1980; Podgursky, 1986), can be interpreted as testing the validity of this model, on the assumption that the costs of organizing are negatively related to the percent organized or, equivalently, to the percentage of industry employment that is unionized.[1] Dickens and Katz (1987b) interpret their research tying nonunion industry wage premia to measures of market power and profitability (in addition to the percent organized), as conducting indirect tests of the union threat model, since these variables presumably proxy for industry rents; however, as they

[1]Rosen (1969) develops this argument in more detail. He suggests that threat effects on nonunion wages will be largest when the demand for union labor is least elastic, since there are then smaller labor flows to the nonunion sector in response to the union wage increase. This elasticity is likely to be smallest (in absolute value) when the industry is relatively more unionized, since there are then fewer substitution possibilities, so that the percent organized is positively related to threat effects. But Rosen also argues that there may be offsetting factors, perhaps because as the percent organized increases, the remaining nonunion firms are particularly resistant to unionization.

recognize, other theories of non-market-clearing wages would also predict a positive relationship between rents and wages (e.g., Akerlof, 1982; Lindbeck and Snower, 1988). Finally, Montgomery (1989) examines the effects of the percent organized and the union wage premium on employment probabilities of individuals.

The formulation closest to ours is Dickens and Katz (1987a), which conducts an empirical analysis of the union threat explanation of nonunion industry wage premia. They use 1983 annual files of the CPS to estimate individual-level wage regressions, including three-digit Census of Population industry dummy variables.[2] In a second-stage equation they then regress the estimated industry coefficients on a rich set of industry characteristics, to identify the correlates of industry wage premia.[3,4]

Consistent with the threat hypothesis, Dickens and Katz find that the sign of the percent unionized variable in the nonunion wage regressions is generally positive. But they point out that the hypothesis is only weakly supported because the coefficient is not always positive, and is often statistically insignificant. Some of the other variables do support the threat model, particularly the positive coefficient on profitability, which they indicate supports the hypothesis that the higher the potential rents, the higher the wage (Dickens, 1986). Since industry characteristics are highly correlated, however, Dickens

[2]The cross-sectional equation includes the measurable human capital of workers, demographic characteristics of workers, and controls for state of residence, living in an MSA, and occupation.

[3]These include: average wages; average income; average demographic characteristics; unemployment and layoff rates; injury rates; overtime; non-wage compensation; average establishment size; measures of market power, concentration, and profitability; other industry characteristics; and, most directly related to the union threat explanation, the percent of workers in the industry covered by collective bargaining agreements.

[4]Dickens and Katz rationalize this "two-step" approach, arguing that estimating an individual-level regression with some variables (i.e., industry characteristics) aggregated to the industry level may lead to biased coefficients because the industry aggregates measure the characteristics of the worker's firm with error (Dickens and Ross, 1984).

and Katz note that it is difficult to sort out the independent influence of the large number of industry characteristics that they consider. In addition, because the key percent unionized variable does not provide robust support for the threat model, it is more difficult to interpret the meaning of the profitability variable.[5]

As indirect evidence against the threat model, researchers have cited the fact that the nonunion inter-industry wage structure is quite stable over time and across regions (e.g., Krueger and Summers 1988; Helwege, 1991). According to these researchers, given variation in union wage premia across time or regions, a nonunion wage structure that is highly correlated across time or regions makes it less likely that a strong threat effect exists. However, the variation in union wage premia that might be expected to affect the nonunion wage structure is changes in intra-industry union wage premia across time or regions, not simply changes in overall union wage premia across time or regions. Consequently, the sharp changes in intra-industry union wage premia in the 1970s and 1980s (Linneman, Wachter and Carter, 1990) provide a natural testing ground for the union threat explanation of the nonunion wage structure. In particular, we can ask whether changes have occurred, in nonunion inter-industry wage differentials, that are related to changes in union threats.

[5]Indeed, they cite factor analysis results for industry data suggesting that one factor can account for much of the variation in industry characteristics. A related result is suggested in their paper, in which they show that two principal components can explain about 50 percent of the covariation in their industry data.

II. The Union Threat, Crowding, and Complements Models

Data Set and Dependent Variables

In this paper we construct a panel data set, rather than a cross-sectional data set as in Dickens and Katz (1987a). The dependent variable--the nonunion wage differential or premium by industry and year--is estimated from cross-sectional wage equations using CPS data, including skill and other related control variables. One of the central independent variables, the union wage premium, is formed in the same fashion. The method of construction is discussed in Section III.

The nonunion wage premia (denoted w^{nup}) are then regressed on the proxies for union strength in a second-stage regression, in which we control for other industry characteristics with industry-specific dummy variables. The hypothesis is that the greater the threat effect, the greater the nonunion premium.

In addition, we add a second equation to test the threat model. Besides estimating an equation where the nonunion wage is the dependent variable, we also analyze the threat effect in terms of the nonunion industry employment share (denoted es^{nu}) as a dependent variable.[6] If the union threat increases, it should affect not only the nonunion wage, but also nonunion employment. We define employment shares as a percent of total economy employment rather than industry employment. The reason is twofold. Since the threat model, as well as the crowding and complements models, suggest spillovers between the union and nonunion sectors in the same industry, a within-industry variable would be difficult to interpret. In addition, losses in either

[6]Union and nonunion employment-share variables were estimated directly from the CPS data. Details are provided in Section III.

union or nonunion employment need not be to the other sector in the same industry, as shown in Linneman, Wachter, and Carter (1991).

Independent Variables or Measures of Union Threats

The next step is to determine the variables that can serve as measures of the magnitude of the union threat. The traditional view is to measure union strength by the percent organized (*%org*), with a predicted positive coefficient in an equation for nonunion wage premia (w^{nup}). This is shown diagrammatically in Figure 1, where a decrease in *%org* means a downward shift in the demand for union labor from D^{u} to D^{u1} and an increase in the demand for nonunion labor from D^{nu} to D^{nu1}.[7] The impact of the decline in union employment is to cause the threat locus, which measures nonunion firms' perceived threat from unionization, to decline from T to T^{1}, allowing w^{nup} to decline.[8]

We expand this framework by introducing the union wage premium as another measure of the union threat (as in the model in Dickens, 1986), also with a predicted positive coefficient. In Figure 1, an increase in the union wage (and hence an increase in the union wage premium) from w^{up} to w^{up2} causes a movement up the demand curve

[7]Since neither sector is a market-clearing sector, and both pay above market wages, employment is demand constrained. Hence, the supply curves can be omitted.

[8]The T locus shows the nonunion wage premium that must be paid to deter unionization. It is downward sloping because as es^{nu} is increasing, the threat is decreasing. It shifts when other variables, particularly the union variables described in this section, change.

for union labor D^u. This in turn causes the threat locus, T, to shift to T^2, causing w^{nup} to increase and es^{nu} to decrease.[9]

In addition, we add three additional independent variables that can proxy for the "costs of organizing." We include two election variables: the percentage of National Labor Relations Board (NLRB) representation elections won by unions (denoted *%won*), and the overall number of representation elections (denoted *rep*). We also include the number of unfair labor practice claims against employers (denoted *ulp*). Our hypothesis is that the coefficient(s) on the election variables will be positive in the threat model.[10] Specifically, when union certification elections are low and when unions are losing many of the elections, the relative cost of remaining nonunion decreases, and T shifts down to T^3.

On the other hand, when the number of unfair labor practice charges against management is high, unions themselves are facing increased costs of organizing workers or maintaining their union status. *Ulp* charges are high when management's opposition to unions is high and unions contest the legality of management's activities. Management opposition in this form allows nonunion wages to be lower; an increase in *ulp* shifts the threat locus, T, downwards. Hence, in the threat model the hypothesis is that the coefficient of *ulp* will be negative.

Our data formulation has the advantage that the statistical experiment is conducted on a data set in which there are sharp changes in variables that are plausibly

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[9]Some difficulties posed by using this variable in an equation where the nonunion wage w^{nup} is the dependent variable are discussed below. The union wage variable, however, is pivotal in the equation where the nonunion employment share is the dependent variable.

[10]Notationally, we refer to the two election variables *%won* and *rep*, when discussed jointly, as *elect*.

related to the strength of the union threat effect. The specifications also include industry dummy variables. The intent is to have the dummy variables capture the other industry characteristics considered by Dickens and Katz (1987a), so that we can identify union threat effects from our estimated coefficients. Using industry fixed effects avoids the problem found in cross-sectional studies of multicollinearity among industry characteristics. However, to identify union threat effects, the omitted characteristics that are correlated with the threat variables must be fixed over time. This is the maintained assumption in this paper; given the sharp changes in union strength over the sample period, it may not be far from the truth.

The union strength proxies are treated as exogenous, because we do not believe that there are identifying restrictions available for explicit treatment of the endogeneity question. Denoting the industry fixed effect I, and subscripting the variables by industry (i) and year (t), we estimate equations of the form

$$(1) \qquad w^{nup}_{it} = \alpha + \%org_{it}\beta_1 + w^{up}_{it}\beta_2 + elect_{it}\beta_3 + ulp_{it}\beta_4 + I_i\gamma + \epsilon_{it} \ .$$

Our specification of threat effects on the nonunion employment share parallels the specification of equation (1). All of the variables from (1) are included in (2) with the exception of *%org*. Since the dependent variable is the nonunion employment share, using the percent organized as an independent variable could result in a spurious negative correlation. This is the case even though our dependent variable has total economy employment as the divisor while the independent variable would have industry employment as the divisor.

Denoting the nonunion employment share as es^{nu}, the equation for the nonunion employment share is

$$(2) \quad es^{nu}_{it} = \alpha' + w^{up}_{it}\beta'_2 + elect_{it}\beta'_3 + ulp_{it}\beta'_4 + I_i\gamma' + \epsilon'_{it} \ .$$

Since the nonunion sector is presumed to be acting like the union sector, in response to increases in union strength or the union wage it moves up its labor demand curve, thereby reducing employment. Hence, the threat model implies that the signs of the union threat proxies are reversed in equation (2); the nonunion employment share moves inversely with the nonunion wage premium. The effects on nonunion wages and employment that are implied by the threat model are summarized in the first panel of Table 1.

The Crowding Model

The crowding model focuses on the implications of the labor supply spillover from the union sector, where employment is restricted by a wage premium, to a market-clearing nonunion sector. Whereas the nonunion sector "acts like the union sector" in the threat model, the nonunion sector reacts competitively in the crowding model. Thus, the responses of nonunion wages and employment to an increase in the union wage premium, w^{up}, are the reverse of those in the threat model. This, as well as the other differences and similarities between the two theories can be seen by comparing Figures 1 and 2.

Turning to *%org*, a decrease in *%org* causes D^u to shift to D^{u1}, thus causing S^{nu} to shift out to S^{nu1}. The result is a decline in w^{nup}. Note that in the crowding model, the traditional assumption is that labor supply spillovers dominate labor demand spillovers. Hence, although the nonunion demand curve shifts out, the maintained assumption in the model is that the nonunion supply curve shifts out by more. Thus, we show only the latter shift. The resulting decline in w^{nup} is the same as in the threat model, but for a different reason; this is the only variable for which the two theories generate the same prediction in equation (1).

The predicted effects of the NLRB variables are also different from those of the threat model. One has to be careful, however, as to the interpretation given these variables in the crowding model. The NLRB variables are introduced to serve the primary mission of the paper, testing the threat model. With the exception of *%won*, they might not be introduced as likely variables in a crowding model, were we testing that model independently of the threat model. But we want to consider the implications for the coefficients of these variables in equations (1) and (2), under the crowding model (and the complements model), in order to interpret results that may be inconsistent with the threat model.

%won is a useful variable because it measures actual or potential shifts between the union and nonunion sector. For example, a decline in *%won* means that unions are losing elections and thus D^u is shifting inward to D^{u3}, and D^{nu} and S^{nu} are shifting out to D^{nu3} and S^{nu3}, as shown in Figure 2. For w^{nup}, the effect is a wash, but es^{nu} clearly increases. In general *%won* operates in a similar fashion to *%org*, but has a greater demand effect.

We interpret the NLRB variables *rep* and *ulp* as measuring costs of union labor not reflected in w^{up}, and hence operating in a similar fashion to w^{up}, although changes in these variables result in shifts of D^u, rather than shifts along it.[11] The lower the number of union elections and the greater the allegations of unfair labor practices by management, the weaker is the union, and hence the lower are labor costs in the union sector. Hence a decrease in *rep* or an increase in *ulp* causes an outward shift in the demand for union labor, from D^u to D^{u4}, leading nonunion labor supply to shift in to S^{nu4}. These shifts cause es^{nu} to decrease, and thus w^{nup} to increase.

The results for the crowding model are summarized in the second panel of Table 1. In closing, we note that the crowding model is sometimes specified to include a possible shift in labor demand in the nonunion sector in the opposite direction of the employment change in the union sector. Including both the spillover of demand and supply from the union sector, the final effects on w^{nup} and es^{nu} are ambiguous, depending on the magnitudes of the shifts in labor supply and demand. In the existing literature, it is usually assumed that the supply shifts dominate. For example, Kahn (1978) contrasts the threat and crowding models by the different impacts that union wage increases have on nonunion wages (positive for the threat model, and negative for the crowding model). We make the same assumption here.

The Complements Model

Finally, we consider a model in which union and nonunion labor are complements in production. This is a more complex model, since it requires the depiction of two nonunion sectors. It does, however, capture a well recognized multiplier effect, whereby

[11]For example, consider two industries that have identical union wage premia, but the first industry has more representation elections and weaker management opposition to unions. We might expect unions in this industry to extract higher non-wage concessions from management, resulting in a higher cost of union labor.

a decline in union employment will spill over into a decline in nonunion employment that works along with, or as a complement to, union labor.

The model is depicted in Figure 3. The nonunion substitute sector in this figure works in the same manner as did the nonunion sector in Figure 2. The difference is that we postulate the existence of a nonunion sector which follows the level of activity in the union sector.

In this model, a decline in *%org* causes the usual decrease in union employment and increase in demand for nonunion labor that acts as a substitute. In this case, the outward shift of S^{ns} to S^{ns1} causes an increase in es^{nus} and a decline in w^{ns}. Although not drawn, D^{ns} may also shift outward, which would moderate the wage decline. But as in the crowding model, it is assumed that the dominant impact of the union sector on the nonunion substitute sector is supply shifts. More importantly, however, the overall composite wage across the union and nonunion substitute sectors declines. The result is an increase in demand in the complements sector. Hence D^{nc} shifts to D^{nc1}. We assume that developments in the nonunion complement sector dominate the overall nonunion sector. Thus, the prediction is that both es^{nu} and w^{nup} will increase.

Moving on to the w^{up} variable, an increase in w^{up} in the complements model acts as a depressant on activity in the union sector without the same corresponding uptick in nonunion substitute sector employment as was the case with *%org* variable. The result is that D^{nc} declines to D^{nc2}, causing both es^{nu} and w^{nu} to decline.

As noted above *%won* operates in a similar fashion to *%org*. Hence a decrease in *%won* is like a decline in *%org*, causing both es^{nu} and w^{nu} to increase. Similarly *rep* and *ulp* act like w^{up}. Hence a decrease in *rep* and an increase in *ulp* causes both es^{nu} and w^{nup} to increase.

The predicted coefficient signs of the complements model are presented in the third panel of Table 1. Scrutinizing Table 1 reveals the advantage of considering both equation (2) (for employment) as well as equation (1) (for wages) if we are to differentiate among the three models. Equation (1) provides sharply contrasting predictions between the threat model, on the one hand, and the crowding and complements models, on the other; but only the coefficient of *%org*, β_1, differentiates between the crowding and complements models. Equation (2), in contrast, provides sharply contrasting predictions between the crowding model, on the one hand, and the threat and complements models, on the other.

Test of the Models Based on Union Sector Adjustments

We also carry out an additional test, which is a more "direct" test of the extent to which nonunion industry wage premia estimated from cross-section wage regressions measure threat effects within the industry. This test considers the feedback of the size of the threat effect on the responsiveness of union employment to changes in the union wage. Specifically, where threat effects are strong, the decline in union employment in response to an increase in the union wage should be relatively weak. That is, union employment is better protected against increases in union wage premia when the nonunion sector validates the union premium by matching it with an increase in its own premium.

We use the nonunion industry wage premia w^{nup} to index threat effects, and test whether the data are consistent with these premia reflecting threat effects. Denoting the union share of total employment es^{u} and the union wage premium w^{up}, we estimate equations of the form[12]

$$(3) \qquad es^{u}_{it} = \delta + w^{up}_{it}\lambda + w^{nup}_{it}\phi + (w^{up}_{it} \times w^{nup}_{it})\theta + I_i\gamma'' + \eta_{it} \ .$$

In general, union employment should respond negatively to an increase in the union wage, so we should find $\lambda < 0$. But threat effects should temper this response. Consequently, if the nonunion industry wage premia w^{nup} index threat effects, then we should find $\theta > 0$; that is, the decline in union employment in response to an increase in the union wage should be less negative in industries in which there are strong threat effects. The sign of ϕ, the coefficient of w^{nup}, is ambiguous, since w^{nup} can be viewed as the price of another input along with union labor. If nonunion and union labor are complements, then ϕ should be negative; ϕ can also be negative if they are substitutes, but the scale effect of an increase in w^{nup} dominates.

As before, we also consider the predictions of the crowding and complements models for the sign of θ in equation (3). The crowding model yields the same prediction as the threat model for this regression. If union and nonunion labor are substitutes, then the union labor demand response to an increase in the cost of union labor should be smaller if accompanied by an increase in the cost of nonunion labor (i.e., a higher value of w^{nup}), and vice versa. Thus, the crowding model also predicts that $\theta > 0$.

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[12]We use the union share of total employment, rather than the union share of industry employment, to minimize the endogeneity problem. The union share of industry employment is significantly influenced by nonunion industry employment, which may in turn influence the right-hand-side variable w^{nup}.

The complements model, however, yields opposite predictions from the threat and crowding models. A high value of the nonunion wage premium corresponds to a high price of a complementary input. Thus, a hike in union labor costs coupled with a high nonunion wage premium results in a larger drop in demand for union labor. That is, the complements model predicts that we should find $\theta < 0$, in contrast to the threat model. These results are summarized in the bottom portion of Table 1.

III. Data on Nonunion Wage Premia, Union Wage Premia, and Union Strength

Our estimates of union and nonunion wage premia come from cross-sectional wage equations estimated for individual years from CPS data. Specifically, we estimated cross-sectional log wage equations for each year 1973-1989, with separate equations for white men, black men, and women.[13] These equations include a union status (membership) dummy variable, and a full set of interactions of all variables, including the industry dummy variables, with the union status dummy variable.[14,15] To focus on competitive market effects of union wages, we exclude government workers from our sample. In addition, we exclude managers, professionals, and technical workers. In part because many of these workers are not covered by the National Labor Relations Act,

[13]1982 is omitted because no data were collected on union membership.

[14]Details of the concordance between 1970 and 1980 SIC codes are available from the authors.

[15]In addition to these variables, we included: nine one-digit occupation dummy variables (with a bridge between the 1970 and 1980 SOC codes); linear and quadratic schooling; linear and quadratic potential experience; dummy variables for four regions; the MSA unemployment rate; dummy variables for three MSA sizes; and dummy variables for married, spouse present, and overtime (based on usual hours worked).

these occupations have low rates of unionization, and changes in their wages and employment are likely to be least affected by developments in the union sector.

The coefficients of the non-interacted industry dummy variables estimate the nonunion industry wage premia.[16] Our estimate of the within-industry union premium is the sum of the coefficients of the industry-union interaction and the union status dummy variable. In all cases we estimated the wage premia for the entire sample by weighting (by industry and union employment) the coefficients estimated from the separate regressions by race and sex.

Our measure of percent organized is constructed from the same CPS data, described immediately above, used in estimating the wage premium variables. The percentage of representation elections won by unions, the number of representation elections held, and the number of unfair labor practice charges against employers are from NLRB Annual Reports.[17]

We constructed a data set at both the one- and two-digit industry levels. Because the NLRB threat proxies are available only at the one-digit level, the two-digit industry data are used only to explore the more limited set of specifications using the union and nonunion industry wage premia and employment shares, and the percent organized.

In Table 2 we provide summary statistics for the one-digit industry data set, reporting mean levels and yearly changes for each of these variables, by industry. The

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[16]Industry wage premia are estimated relative to services for the one-digit analysis, and relative to business services for the two-digit analysis. The regression results are of course insensitive to which industry is omitted. We replicated many of our results defining industry wage premia instead relative to the average nonunion worker (weighting by the representation of workers across industries). The conclusions were unchanged.

[17]Details are provided in the footnotes to Table 2.

yearly changes are related to the data used to estimate the regressions, although fixed-effects are used rather than first differences. The data confirm the now well-known increase in the within-industry union premium (w^{up}) in many industries, coupled with the decline in percent organized (*%org*). For the NLRB variables, the percentage of elections won by unions (*%won*) and the number of representation elections (*rep*) decline for nearly all industries, while the number of unfair labor practice charges (*ulp*) increases for two-thirds of the industries.

In focusing on this time-series cross-sectional data set an interesting question is raised: was the threat effect posed by unions increasing or decreasing during the 1973-1989 period? On the one hand, the threat effect as measured by the percent organized and the other NLRB measures was declining. As the most well-known development in the union sector, the prevailing view is likely to be that the threat effect was declining. On the other hand, the wage differential between union and nonunion workers within industries, which is related to the potential gains from unionization, tended to increase over this period, in manufacturing as well as other industries (Linneman, Wachter, and Carter, 1990, and our Table 2). Since the weight one should give to the opposing factors is unclear, the overall direction of the threat effect over the 1973 to 1989 sample period is indeterminate. But, as we shall show, the question is moot. Although the nonunion sector did respond systematically to developments in the union sector over the sample period, the changes are not those predicted by the threat model.

IV. Empirical Results

Potential Biases

We have identified four sources of bias in the estimates of our equations that may affect the conclusions to which the estimates lead. First, in equation (1) we use w^{up} as a measure of the potential wage gain from unionization, while w^{nup} is the dependent variable. The w^{up} variable is obtained from the wage regression as the sum of the coefficients on the union variable and the union-industry interaction variables. The w^{nup} variable, on the other hand, is the coefficient on the industry dummy variable. To the extent that the covariation between w^{up} and w^{nup} comes from changes in nonunion industry wage premia that do not similarly affect union workers in the industry (perhaps because they are locked into contracts), a negative correlation arises between the dependent variable and the within-industry union premium that does not reflect causal influences running from the latter variable to the former. On the other hand, variation in w^{up} generated by union-specific forces will impact the union coefficient or the union-industry coefficient without creating this spurious negative correlation. Because of this potential problem, when we discuss the empirical results for equation (1), the nonunion wage premium equation, we focus most on the specifications excluding the within-industry union premium, w^{up}.

A second potential problem with the within-industry union premium is that it may reflect the influence of underlying (possibly unobservable) union threat effects, because this measure may reflect the equilibrium after the effects of union threats are assimilated into nonunion wages. For example, the within-industry premium may be large in an industry precisely because unobserved union threat effects are weak, leading to negative bias. Including industry dummy variables should help, however, by controlling for unobserved industry characteristics that are fixed over time. This also leads us to focus on specifications excluding the within-industry union premium.

Third, in specifications excluding w^{up}, the percent organized (*%org*) is a key variable of interest. The estimated coefficient of this variable may, however, be upward biased by endogeneity. Omitted factors that shift up the nonunion premium, w^{nup}, may lead to declines in nonunion employment, and hence increases in *%org*; this generates a positive association between the residual in equation (1) and *%org*. The implication of this is that the results are biased against finding a negative coefficient on *%org*, which, as Table 1 shows, would be evidence in favor of the complements model. Since much of our evidence ultimately points toward the complements model, this potential source of bias only strengthens our results.[18]

Fourth, we have so far ignored the potential effects of industry-specific demand shocks (stemming, perhaps from import competition) that affect both union and nonunion labor. These could generate potentially important biases if, as seems plausible, union wages are less flexible than nonunion wages, with union employment consequently more flexible. In this case a downward industry demand shock, for example, will result in a decrease in the nonunion premium (w^{nup}), a decrease in the nonunion employment share (es^{nu}), an increase in the within-industry union premium (w^{up}), and a decrease in the percent organized (*%org*). Looking at Table 1, we can see that in equation (1), the equation for the nonunion wage premia, this biases the

[18]We do not believe that we have identifying information on the basis of which to correct for the endogeneity bias.

coefficient of w^{up} downward, and the coefficient of the percent organized upward. Together, these biases will tend to favor the crowding model. In equation (2), the equation for the nonunion employment share, this biases the coefficient of w^{up} downwards, creating a bias against the crowding model. To control for industry-specific demand shocks, we introduce the percentage change in the GNP share accounted for by each industry to attempt to control for shifts in industry-specific demand for labor, although we recognize that this variable is not unambiguously exogenous. To account for differences across industries in cyclical fluctuations in this variable, we estimated separate regressions, for each industry, of the GNP share on the aggregate unemployment rate and a post-1976 dummy variable for the change in accounting methods, and used the residual in the regression estimates of equations (1) and (2).[19,20]

Results

In Table 3 we report results from regression estimates of equation (1), in which the nonunion industry wage differential is the dependent variable. In columns (1)-(7) we use the one-digit industry data.[21] In column (1) we present the standard model of the union threat effect. The negative coefficient on the percent organized variable is the opposite of that predicted by the threat model, since the estimated coefficient says

[19]The GNP-share data are from CITIBASE for 1973-1976, and from the Survey of Current Business for the later years.

[20]To explore the role of aggregate demand shocks, we experimented with the inclusion of year dummy variables in our equations, but their coefficients were never statistically significant. Here, in contrast, we are referring to demand shocks that affect specific industries.

[21]Unless otherwise specified, statements regarding statistical significance refer to five-percent significance levels. For regression coefficients, the statements refer to two-sided tests.

that the greater the union strength in organizing the industry, the lower the wage in the remaining nonunion sector.[22]

In column (2), the NLRB variables are added to the equation. The predictions of the threat model are that the two election variables are positively related to the nonunion wage premium, while the number of unfair labor practice claims is negatively related. The percent organized variable remains largely unchanged, while the NLRB variables present a mixed picture. The coefficients of the elections variables are positive, consistent with the threat model, although neither is statistically significant.[23] The unfair labor practice variable, on the other hand, has the opposite sign from that predicted by the threat model, although it, too, is statistically insignificant. In columns (3) and (4) we replace percent organized with the within-industry union premium, first alone and then along with the NLRB variables. In column (3) the coefficient of the within-industry union premium is negative and strongly significant.
This result rejects the threat model, but the conclusion must be tempered by the potential for a spurious negative correlation for reasons noted above.

When the NLRB variables are added, in column (4), the evidence against the threat model is more one-sided. The coefficients of both election variables are negative,

[22]We recognize, though, that the relationship between the percent organized and the union threat may be ambiguous, as in Rosen (1969). However, for two reasons we have a strong expectation of a positive relationship. First of all, while the difficulty of unionizing may become quite strong once very high levels of unionization are reached, this hardly seems to characterize unionization rates in the U.S. in the sample period. Second, much of the individual-level evidence on the union threat hypothesis looks for and finds positive relationships between individuals' wages and the percent of their industry unionized, although this relationship may be present only for certain demographic groups (Kahn, 1980), or in large firms (Podgursky, 1986), or it may be rather weak statistically (Freeman and Medoff, 1981).

[23]There was no evidence of severe multicollinearity between the two elections variables; in general, the coefficient estimates and standard errors of the election variables were little changed when one or the other was omitted from the equation.

and the number of elections variable is statistically significant. The coefficient of the unfair labor practices variable remains positive but not statistically significant; furthermore, the three NLRB variables are jointly significant. The within-industry union premium coefficient remains negative and significant.

In column (5) we reintroduce the percent organized variable. The results again are almost entirely the opposite of those predicted by the threat model. Both the percent organized and within-industry premium variables have statistically significant negative coefficients. The only NLRB variable that is significant is the unfair labor practice variable, and this also has the positive sign noted above.

In column (6) we add a variable capturing changes in the share of GNP contributed by each industry, to the specification from column (2). As discussed at the beginning of this section, this is intended to control for biases induced by industry-specific demand shocks. The estimated coefficients are little changed. The same is true when we add back in the within-industry union premium, in column (7).

In columns (8)-(10), we report results for the subset of the specifications that can be estimated using data at the two-digit industry level of disaggregation. The greater disaggregation implies that our estimates of industry wage premia and employment shares are estimated from smaller cells, leading to greater measurement error bias and less precise coefficient estimates. The results in columns (8)-(10) nonetheless parallel closely the estimates of the corresponding one-digit specifications. As in the estimates at the one-digit level, the estimates of the *%org* and w^{up} coefficients are negative and statistically significant.

Besides providing strong results against the union threat model, the pattern of results in Table 3 provides mixed support for the crowding model. Remember that in the crowding model, the nonunion sector reacts competitively and, assuming that union and nonunion labor are substitutes, generally in the opposite direction of the union threat model. The crowding model successfully predicts the negative coefficient on w^{up} and the positive coefficient (in some of the specifications) on *ulp*, at the one-digit level, and the corresponding negative coefficient on w^{up}, at the two-digit level. These coefficients were incorrectly predicted by the threat model. In terms of w^{up}, the crowding model predicts that the higher the union wage premium, the more crowding will occur in the nonunion sector, hence the lower is w^{nup}. In terms of *ulp*, the crowding model predicts that the higher the number of unfair labor practice allegations, the weaker is union power, reflecting more active management opposition, and hence, the lower is the cost of union labor. This, in turn, means a higher w^{nup} reflecting the weaker union sector.

The percent organized variable is the one exception to the rule that the crowding model has the opposite prediction from the threat model. For this variable, the direct supply effects are important and probably outweigh the cost of labor interpretation assigned to the other threat variables. Viewed in this fashion, the crowding model, like the threat model, does not predict the negative coefficient on *%org*. That is, the crowding model predicts that the lower the percent organized, the more crowding there is in the nonunion sector, and hence the lower should be w^{nup}. In fact, however, the estimates in Table 3 indicate that lower *%org* leads to higher w^{nup}.

The complements model, however, is more successful in predicting the pattern of results found in Table 3. Remember that in the complements model, nonunion labor that is complementary to union labor follows the level of economic activity in the union sector, and this complementary nonunion labor dominates the nonunion sector overall. The implication is that, with respect to regression equation (1), the complements model has the reverse predictions from the threat model. In the threat model, increases in union power cause the nonunion sector to mimic the union sector, moving up its demand curve and increasing w^{nup}. In the complements model, however, nonunion labor that is a production complement to union labor responds to higher labor costs by decreasing its wage (and reducing its employment share).

Viewed in this fashion, the complements model successfully predicts the negative coefficient on w^{up}; the higher w^{up}, the higher the cost of union labor and thus the lower the level of economic activity in the union sector. The less activity in the nonunion sector that is a complement to the union sector, the lower w^{nup}. Similarly, the complements model correctly predicts the positive sign on the unfair labor practice variable. In terms of *ulp*, the higher the level of management opposition, the lower the cost of union labor and thus the higher w^{nup}. Finally, the negative sign on *%org* is consistent with complementarity between union labor and nonunion labor, overall.

In summary, the result of estimating equation (1) is a set of coefficients that are almost entirely contrary to the predictions of the threat model, are mixed with respect to the predictions of the crowding model, but confirm the predictions of the complements model. This conclusion holds at both the one- and two-digit industry levels.

To gain further evidence on the threat model, and to further distinguish between the alternative models, we next turn to estimates of equation (2), in which the nonunion employment share of total economy employment, es^{nu}, is the dependent variable. Results for these regressions at the one- and two-digit levels are reported in Table 4. This equation has the same specification as equation (1) with the exception that the *%org* variable is excluded. Although the dependent variable, es^{nu}, is a share of total economy employment, it is likely to be spuriously negatively correlated with *%org* (the union share of industry employment).

In each of the specifications estimated in Table 4, all of the coefficients are statistically significant, except for *%won* (and *ulp* in columns (2) and (3)). Focusing on the statistically significant coefficients, with respect to the threat model, the results of Table 4 are largely supportive. Increases in w^{up} cause decreases in es^{nu}, and increases in *ulp* or decreases in *rep* cause increases in es^{nu}. Given the results in Table 3, however, the result with respect to w^{up} is suspect. The reason that es^{nu} should fall in the threat model is that w^{nup} has increased in response to increases in w^{up}. However, we know from Table 3 that this is not supported by the data. Hence, the decrease in es^{nu} as a consequence of increases in w^{up} cannot support the threat model.

For the corresponding analysis at the two-digit level, only one specification can be estimated, the regression of es^{nu} on w^{up} (and the industry dummy variables). The regression is reported in column (5) of Table 4. The coefficient of w^{up} is negative, but is not statistically significant. This is consistent with the one-digit results in the table, although it is not, by itself, informative.

The crowding model is rejected by the results in Table 4. Increases in w^{up} should cause additional crowding in the nonunion sector, hence increasing es^{nu}. But the coefficient on w^{up} is negative. Similarly the crowding model incorrectly signs the *ulp* and representation elections variables in Table 4.

The complements model, however, is supported by the results of Table 4. In the complements model the negative coefficient on the w^{up} variable is predicted. Since increases in w^{up} cause a decline in economic activity in the union sector, they also cause a decrease in economic activity in the nonunion complements sector, hence the decline in es^{nu}. Similarly the positive sign on *ulp*, which is an indirect measure of the cost of union labor, is predicted by the complements model, as is the negative sign on the representation elections variable.

Overall, the results for the employment share regressions suggest that increases in union strength are associated with declines in nonunion industry employment. By itself, this finding is consistent with either the threat model or the complements model, but not the crowding model. *However, only the complements model is consistent with both sets of findings from the employment and wage regressions.* Furthermore, it is reasonable to expect that production complementarities between union and nonunion labor are weaker at a more disaggregated level of analysis, since the disaggregation is more likely to sort the complementary types of labor into different industries. Because of this, we view the two-digit industry results for the wage and employment regressions as particularly compelling evidence in favor of the complements model.

Finally, we turn to estimates of equation (3), intended to test directly whether nonunion industry wage premia are positively related to union threat effects. Table 5 reports results at the one- and two-digit level. To summarize the test, the coefficient of the within-industry union premium should be negative, reflecting movements along the labor demand curve for union labor.[24] If nonunion industry wage premia are positively related to threat effects, then the interaction coefficient should be positive, as strong threat effects diminish the substitutability between union and nonunion workers. A positive interaction coefficient would also be consistent with the crowding model. On the other hand, a negative coefficient on the interaction variable would be consistent with the complements model, indicating that increases in union wages costs coupled with increases in nonunion wage costs lead to relatively larger declines in union employment. The estimates in column (1) are consistent with the complements model. The coefficient of the interaction variable (-.276) is negative and statistically significant. Furthermore, the coefficient of the within-industry union premium is negative, as expected, and the negative coefficient of the nonunion industry wage premium is consistent with complementarity. In column (2) we explore whether the coefficient of the interaction variable is being identified from cross-industry variation in the responsiveness of union employment to union wages (i.e., union labor demand elasticities). To do this we add a set of variables interacting the within-industry union wage premium with a set of industry dummy variables.[25] The signs of the coefficients are unchanged, but the coefficient on the interaction variable is no longer statistically significant. Columns (3)

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[24]We define the nonunion industry wage premium used to construct the interaction variable as deviations about its mean. Consequently, the partial derivative of union employment with respect to the within-industry union wage premium, evaluated at the means, is given by the coefficient of the within-industry union wage premium.

[25]We define the industry dummy variables relative to their sample means so that the within-industry union wage premium coefficient still measures the partial derivative, evaluated at the sample means.

and (4) repeat this analysis at the two-digit level, with the same results; the signs of the point estimates are consistent with the complements model, although the evidence is statistically significant only in column (3).

Overall, the evidence in Table 5 provides additional support for the complements model; as we suggested earlier, we find the evidence in favor of the complements model at the two-digit level particularly compelling. Based on these findings, we can more decisively conclude that the results provide no evidence that nonunion industry wage premia are correlated with the strength of the union threat within industries.

V. Conclusions

In this paper we investigated the impact of union strength on changes in nonunion wages and employment. The prevailing model in this area is the threat model, which predicts that increases in union strength cause increases in nonunion wages and decreases in nonunion employment. In testing the threat model, we are also testing two alternatives, the crowding and complements models.

With respect to nonunion industry wage differentials, the statistical experiment does not ask whether one can explain the "entire" nonunion industry wage structure with the union threat effect model, but only asks whether the threat effect model explains some of the variation in this wage structure over time. The answer appears to be no. In contrast to the prediction of the threat model, decreases in the percent organized (reflecting a declining union threat) are associated with increases in the nonunion wage. Furthermore, increases in union wages appear to decrease, rather than to increase, nonunion wages. Evidence on the determinants of intra-industry variation in nonunion wage premia is somewhat more consistent with the crowding model and is strikingly consistent with the complements model of union and nonunion wage determination. Given that the union threat model does not explain intra-industry variation in nonunion industry wage premia over time, it seems an unlikely candidate for an explanation of the cross-sectional variation in nonunion industry wage premia.

Further evidence on the determinants of intra-industry variation in nonunion employment is consistent with the complements and the threat model; movements in nonunion industry employment are negatively related to changes in proxies for union strength.

Thus, the combined evidence supports the complements model, but neither the threat model nor the crowding model. Evidence from a third test, based on the responsiveness of union employment to union wage changes (again, within industry), strengthens this finding, as it is consistent with the complements model, but neither the threat model nor the crowding model.

In testing the model on a cross-section, time-series data base from 1973-1989, we are also able to explore how the extraordinary changes in the union sector during this period have affected the nonunion sector. In the context of the threat model, increases in the union wage premium in many industries, coupled with sharp declines in union employment shares, raise the possibility that the strength of the union threat in different industries varied little over this period. Thus, the threat model is not necessarily inconsistent with the observed stability of the nonunion industry wage structure over the same period.[26] While our evidence rejects the threat model in favor of the complements model, it still suggests that union employment shares and union wage premia have offsetting influences on nonunion wages; but the effects have the opposite signs from those predicted by the threat model. Thus, it is possible that union sector developments do influence the nonunion industry wage structure, but these influences are masked in the 1970s and 1980s because of offsetting trends in the union sector.

[26]Krueger and Summers (1988) reports correlations around .91 between industry wage differentials estimated using CPS data in 1974, 1979, and 1984.

References

Akerlof, G.A. 1982. "Labor Contracts as Partial Gift Exchange." *Quarterly Journal of Economics* 91(4): 543-69.

Bell, L.A. 1989. "Union Concessions in the 1980s." *Federal Reserve Bank of New York Quarterly Review* Summer: 44-58.

Dickens, W.T. 1986. "Wages, Employment and the Threat of Collective Action by Workers." Mimeograph, University of California, Berkeley.

Dickens, W.T. and L.F. Katz. 1987a. "Inter-Industry Wage Differences and Industry Characteristics." In K. Lang and J.S. Leonard, Eds., *Unemployment and the Structure of Labor Markets* (New York, NY: Basil Blackwell), 48-89.

_____. 1987b. "Inter-Industry Wage Differences and Theories of Wage Determination." NBER Working Paper No. 2271.

Dickens, W.T. and B.A. Ross. 1984. "Consistent Estimation Using Data from More than One Sample." NBER Technical Working Paper No. 33.

Dunlop, John T. 1944. *Wage Determination Under Trade Unions* (New York, NY: Macmillan).

Freeman, R.B. 1986. "The Effect of the Union Wage Differential on Management Opposition and Union Organizing Success." *American Economic Review* 76(2): 92-6.

Freeman, R.B. and J.L. Medoff. 1981. "The Impact of the Percentage Organized on Union and Nonunion Wages." *The Review of Economics and Statistics* LXIII(4): 561-72.

Helwege, J. 1991. "Sectoral Shifts and Interindustry Wage Differentials." Mimeograph, Federal Reserve Board.

Kahn, L.M. 1978. "The Effect of Unions on the Earnings of Nonunion Workers." *Industrial and Labor Relations Review* 31(2): 205-16.

_____. 1980. "Union Spillover Effects on Organized [sic] Labor Markets." *Journal of Human Resources* 15(1): 87-98.

Krueger, A.B. and L.H. Summers. 1988. "Efficiency Wages and the Inter-Industry Wage Structure." *Econometrica* 56(2): 259-94.

Lindbeck, A. and D.J. Snower. 1988. *The Insider-Outsider Theory of Employment and Unemployment* (Cambridge, MA: The MIT Press).

Linneman, P.D., M.L. Wachter, and W.H. Carter. 1990. "Evaluating the Evidence on Union Employment and Wages." *Industrial and Labor Relations Review* 44(1): 34-53.

Montgomery, E. 1989. "Employment and Unemployment Effects of Unions." *Journal of Labor Economics* 7(2): 170-90.

Neumark, D. "Declining Union Strength and Labor Cost Inflation in the 1980s." Forthcoming in *Industrial Relations*.

Podgursky, M. 1986. "Unions, Establishment Size, and Intra-Industry Threat Effects." *Industrial and Labor Relations Review* 39(2): 277-84.

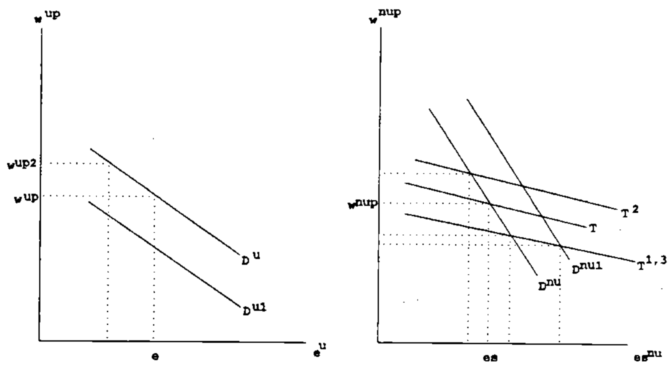
Rosen, S. 1969. "Trade Union Power, Threat Effects and the Extent of Organization." *Review of Economic Studies* 36(2), No. 106: 185-96.

Ross, Arthur M. 1948. *Trade Union Wage Policy* (Berkeley, CA: University of California Press).

Wachter, M.L. and W.H. Carter. 1989. "Norm Shifts in Union Wages: Will 1989 Be a Replay of 1969?" *Brookings Papers on Economic Activity* 2: 233-64.

Figure 1

Union Threat Hypothesis



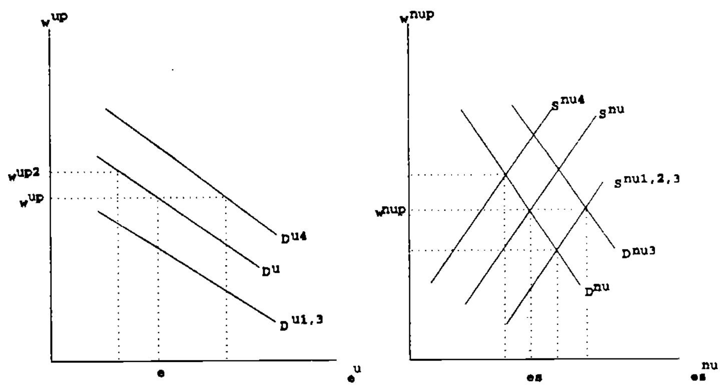
Union Sector

Nonunion Sector

Change	Effect	w^{nup}	es^{nu}
1. Decrease in %org ($D^{u} \rightarrow D^{u1}$, $D^{nu} \rightarrow D^{nu1}$)	$T \rightarrow T^{1}$	Decreases	Increases (by assumption)
2. Increase in w^{u} to w^{u2}	$T \rightarrow T^{2}$	Increases	Decreases
3. Decrease in %won, rep, increase in ulp	$T \rightarrow T^{3}$	Decreases	Increases

Figure 2

Crowding (Supply) Alternative



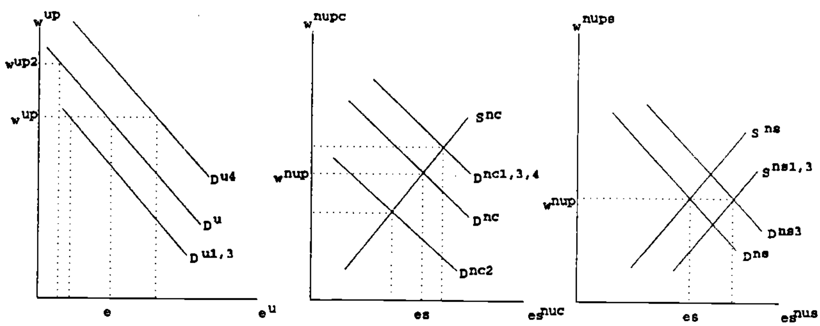
Union Sector　　　　Nonunion Sector

Change	Effect	w^{nu}	es^{nu}
1. Decrease in %org	$D^{u} \rightarrow D^{u1}$, $S^{nu} \rightarrow S^{nu1}$	Decreases	Increases (by assumption)
2. Increase in w^{u} to w^{u2}	$S^{nu} \rightarrow S^{nu2}$	Decreases	Increases
3. Decrease in %won	$D^{u} \rightarrow D^{u3}$, $D^{nu} \rightarrow D^{nu3}$, $S^{nu} \rightarrow S^{nu3}$	No change	Increases
4. Decrease in rep, increase in ulp	$D^{u} \rightarrow D^{u4}$, $S^{nu} \rightarrow S^{nu4}$	Increases	Decreases

Note: We assume shifts in S^{nu} dominate shifts in D^{nu}, and therefore generally do not show the latter.

Figure 3

Complements (Demand) Alternative



Union Sector | Nonunion Sector (Comp.) | Nonunion Sector (Subst.)

Change	Effect	w^{nc}	es^{nc}
1. Decrease in %org	$D^{u} \to D^{u1}$, $S^{ns} \to S^{ns1}$, $D^{nc} \to D^{nc1}$	Increases	Increases (by assumption)
2. Increase in w^{u} to w^{u2}	$D^{nc} \to D^{nc2}$	Decreases	Decreases
3. Decrease in %won	$D^{u} \to D^{u3}$, $D^{nc} \to D^{nc3}$, $S^{ns} \to S^{ns3}$, $D^{ns} \to D^{ns3}$	Increases	Increases
4. Decrease in rep, increase in ulp	$D^{u} \to D^{u4}$, $D^{nc} \to D^{nc4}$	Increases	Increases

Note: We assume that movements in wages and employment of nonunion complement labor dominate the nonunion sector. We assume shifts in D^{ns} dominate shifts in S^{ns}, and therefore do not show the latter.

Table 1

Summary of Coefficient Predictions from Alternative Models

Equations 1 and 2:		Nonunion wage (1)	Nonunion employment (2)
		Threat Model	
%org		+	not included
w^{up}		+	−
%won		+	−
rep		+	−
ulp		−	+
		Crowding Model	
%org		+	not included
w^{up}		−	+
%won		no change	−
rep		−	+
ulp		+	−
		Complements Model	
%org		−	not included
w^{up}		−	−
%won		−	−
rep		−	−
ulp		+	+
Equation 3:		**Union employment (3)**	
		Threat Model	
w^{up}	λ	−	
$w^{up} \times w^{nup}$	θ	+	
		Crowding Model	
w^{up}	λ	−	
$w^{up} \times w^{nup}$	θ	+	
		Complements Model	
w^{up}	λ	−	
$w^{up} \times w^{nup}$	θ	−	

Table 2

Descriptive Statistics by One-Digit Industry[1]

	Levels							
	Nonunion wage premium (1)	Within-industry union wage premium (2)	Nonunion employment share[2] (3)	Union employment share[2] (4)	Percent organized (5)	Percent elections won by unions (6)	Representation elections[3] (7)	Unfair labor practice charges against employers[3] (8)
Industry:								
Construction	.18 (.01)	.36 (.01)	.05 (.001)	.03 (.002)	.35 (.02)	.49 (.01)	.05 (.01)	.19 (.01)
Mining	.34 (.02)	.16 (.02)	.01 (.0004)	.01 (.0005)	.40 (.03)	.43 (.02)	.01 (.001)	.04 (.004)
Manufacturing-durable	.17 (.01)	.17 (.01)	.12 (.001)	.08 (.01)	.40 (.02)	.43 (.01)	.20 (.02)	.63 (.03)
Manufacturing-non-durable	.12 (.01)	.15 (.02)	.09 (.001)	.04 (.003)	.33 (.01)	.42 (.01)	.13 (.02)	.37 (.02)
Transportation, communications, and public utilities	.22 (.02)	.19 (.02)	.04 (.001)	.05 (.002)	.52 (.02)	.48 (.01)	.11 (.01)	.36 (.01)
Wholesale trade	.15 (.01)	.15 (.02)	.05 (.002)	.01 (.0004)	.13 (.01)	.42 (.01)	.06 (.01)	.13 (.01)
Retail trade	-.06 (.01)	.29 (.01)	.13 (.01)	.02 (.001)	.12 (.01)	.41 (.01)	.10 (.01)	.25 (.01)
Finance, insurance, and real estate	.12 (.02)	.05 (.03)	.07 (.003)	.003 (.0001)	.05 (.002)	.53 (.02)	.02 (.001)	.03 (.003)
Services	...	.14 (.01)	.17 (.004)	.03 (.001)	.16 (.004)	.54 (.01)	.17 (.01)	.39 (.02)

Table 2 (continued)

Annual Changes

	Nonunion wage premium (1)	Within-industry union wage premium (2)	Nonunion employment share[2] (3)	Union employment share[2] (4)	Percent organized (5)	Percent elections won by unions (6)	Representation elections[3] (7)	Unfair labor practice charges against employers[3] (8)
Industry:								
Construction	.005	-.007	.001	-.001	-.012	.002	.007	.005
	(.006)	(.010)	(.001)	(.001)	(.009)	(.014)	(.005)	(.007)
Mining	.012	-.004	.0002	-.0002	-.021	-.012	-.0002	.004
	(.031)	(.037)	(.0003)	(.0003)	(.011)	(.020)	(.001)	(.003)
Manufacturing-durable	.003	.003	-.00003	-.005	-.014	-.005	-.016	-.002
	(.012)	(.022)	(.002)	(.003)	(.005)	(.005)	(.009)	(.020)
Manufacturing-non-durable	.003	.001	-.00005	-.002	-.011	-.003	-.010	-.0001
	(.013)	(.025)	(.001)	(.001)	(.005)	(.009)	(.006)	(.014)
Transportation, communications, and public utilities	.001	.003	.001	-.001	-.012	-.007	-.003	.005
	(.019)	(.027)	(.001)	(.001)	(.009)	(.009)	(.005)	(.009)
Wholesale trade	.002	.003	.001	-.0002	-.006	-.009	-.004	.003
	(.018)	(.031)	(.001)	(.0001)	(.005)	(.013)	(.004)	(.007)
Retail trade	.0003	.004	.003	-.001	-.006	-.006	-.007	-.002
	(.007)	(.014)	(.002)	(.0004)	(.003)	(.009)	(.003)	(.007)
Finance, insurance, and real estate	.005	-.002	.001	-.00003	-.001	.009	-.0007	.001
	(.021)	(.058)	(.004)	(.002)	(.002)	(.019)	(.001)	(.002)
Services	...	.003	.003	.0003	-.001	-.002	-.0003	.014
		(.017)	(.003)	(.001)	(.004)	(.020)	(.009)	(.011)

1. Means are reported, with standard errors in parentheses.
2. Share of total employment.
3. The representation elections variable is the number of cases received in the fiscal year stemming from petitions filed by a labor organization or employee seeking an election for determination of a collective bargaining representative. The number of elections is divided by 10,000. The unfair labor practice variable is the number of charges filed by labor organizations or employees, against employers. The number of charges is divided by 10,000.

Table 3

Nonunion Wage Regressions, 1973-1989, Weighted Least Squares
Dependent Variable: Nonunion Industry Wage Premium[1]

	One-Digit Results							Two-Digit Results		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Percent organized	-.273 (.040)	-.328 (.046)	...	...	-.248 (.038)	-.327 (.048)	-.236 (.039)	-.107 (.023)	...	-.106 (.019)
Percentage of elections won by unions	...	.019 (.039)	...	-.034 (.035)	.011 (.031)	.019 (.040)	.014 (.032)	...	...	...
Number of representation elections/10,000	...	.106 (.063)	...	-.126 (.055)	.004 (.052)	.104 (.069)	-.024 (.057)	...	...	...
Unfair labor practice charges against employers /10,000	...	.046 (.044)	...	.045 (.040)	.079 (.035)	.046 (.044)	.077 (.035)	...	...	...
Within-industry union premium	...	...	-.452 (.052)	-.476 (.051)	-.404 (.046)	...	-.411 (.046)	...	-.377 (.029)	-.376 (.028)
GNP share[2]	...	...	...	...	...	-.044 (.569)	-.558 (.454)	...	...	...
$\bar{R}^2$	.964	.965	.969	.971	.978	.965	.978	.938	.953	.956

1. Standard errors are reported in parentheses. There are 144 observations for the one-digit analysis, and 480 observations for the two-digit analysis. Industry dummy variables are included. Observations are weighted by number of observations for year from which micro-level coefficients were estimated. The within-industry union and nonunion premiums are estimated from log wage regressions using the outgoing rotation group annual files of the CPS for 1983-1989, and the May files for 1973-1981. Regressions were estimated separately by race and sex, omitting government workers, managers and professionals, and technical workers. The premiums were then calculated as weighted averages of the premiums estimated from the separate wage regressions. Other variables included in the individual-level wage regressions were: one-digit industry dummy variables (construction, mining, durable manufacturing, nondurable manufacturing, transportation, finance, insurance and real estate, retail trade, wholesale trade, and services); nine one-digit occupation dummy variables (with a bridge between the 1970 and 1980 SOC codes); linear and quadratic schooling; linear and quadratic potential experience; dummy variables for four regions; the MSA unemployment rate; and dummy variables for three MSA sizes; and dummy variables for married, spouse present, overtime (based on usual hours worked).

2. This is the residual from a regression estimated for each industry of the GNP share of output produced by the industry on an intercept, the aggregate civilian unemployment rate, and a post-1976 dummy variable to capture the change in accounting methods used in the numbers reported in the Survey of Current Business.

Table 4

Nonunion Employment Share Regressions, 1973-1989, Weighted Least Squares
Dependent Variable: Nonunion Industry Employment/Total Employment[1]

	One-Digit Results				Two-Digit Results
	(1)	(2)	(3)	(4)	(5)
Within-industry union premium	-.033 (.017)	...	-.046 (.016)	-.040 (.015)	-.008 (.006)
Percentage of elections won by unions	...	.019 (.011)	.019 (.011)	.011 (.010)	...
Number of representation elections/10,000	...	-.082 (.017)	-.090 (.017)	-.057 (.016)	...
Unfair labor practice charges against employers /10,000	...	.017 (.013)	.022 (.013)	.022 (.011)	...
GNP share[2]	...	...	...	.823 (.143)	...
$\bar{R}^2$	.979	.982	.983	.986	.953

1. Standard errors are reported in parentheses. There are 144 observations for the one-digit analysis, and 480 observations for the two-digit analysis. Industry dummy variables are included. Observations are weighted by number of observations for year from which micro-level coefficients were estimated. See footnotes to Table 3 for more details.

Table 5

Employment Share Response Regressions, 1973-1989,
Weighted Least Squares
Dependent Variable: Union Share of Total Employment[1]

	One-Digit Results		Two-Digit Results	
	(1)	(2)	(3)	(4)
Within-industry union wage premium (w^{up})	-.078 (.025)	-.022 (.071)	-.007 (.003)	.002 (.017)
Nonunion industry wage premium (w^{nup})	-.084 (.036)	-.095 (.052)	-.011 (.004)	-.010 (.005)
Within-industry union wage premium (w^{up}) × nonunion industry wage premium (w^{nup})[2]	-.276 (.135)	-.186 (.185)	-.039 (.015)	-.030 (.024)
Includes industry dummy × within-industry union wage premium interactions	No	Yes	No	Yes
$\bar{R}^2$	.872	.874	.805	.817

1. Standard errors are reported in parentheses. Industry dummy variables are included. There are 144 observations for the one-digit analysis, and 480 observations for the two-digit analysis. Observations are weighted by number of observations for year from which micro-level coefficients were estimated. See footnotes to Table 3 for more details.
2. This is included as deviations about its mean, so that the coefficient of the within-industry union wage premium measures the partial derivative of the employment share variable with respect to the within-industry union premium, evaluated at the means.