IS UNIONIZATION AFFECTED BY FRINGE BENEFITS?

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Abstract

This study examines how non-wage benefits have affected union density in the private sector, using both time-series and cross-sectional samples. As benefits like health care have become increasingly expensive and a more prominent component of compensation, *ceteris paribus*, union density seems adversely affected.

It is well known that private sector unionization in the U.S. has long been declining. After reaching a high mark of 35.6 percent of the civilian workforce in 1954, union density has fallen steadily, to 9.8 percent in 1997. A seminal study by Ashenfelter and Pencavel (1969) has spawned a vast literature, in which researchers seek to identify factors responsible for the decline. For example, empirical studies have shown that factors in the decline include changes in the structure of the economy, management opposition to organized labor, and other socio-political factors.[1] Furthermore, Neumann and Rissman (1984) have argued that as government has institutionalized protections for workers, support for organized labor has been weakened.

The structure of compensation also has changed dramatically. In 1948 non-wage benefits accounted for approximately 4.5 percent of compensation overall; by 1994 that proportion had more than quadrupled, to 19 percent. Despite this coincidence, which is illustrated in Figure 1, there has been no analysis whether the composition of pay has affected organized labor.

Benefits take many forms, for example, health care insurance, pension funding, vacation and sick pay, or mandated employer payments for social security and unemployment insurance. The costliest ones are social security contributions and health care (U.S. Bureau of Labor Statistics, 1997; Lettau and Buchmueller, 1999). Benefits are suspended during work stoppage; they are lost altogether if permanent replacements are hired. Moreover, in recent years firms have been more aggressive in confronting organized labor (for example, see Voos, 1994).

Suppose workers are reluctant to jeopardize health coverage or other benefits. As health care becomes increasingly expensive and pay consists more and more of fringe benefits, perhaps workers are less likely to risk industrial action. If unions lose bargaining leverage as a result, workers may surmise that organized labor has less to offer, thereby undermining support for unions. Furthermore, as benefits have become institutionalized, perhaps workers are less prone to unionize in the first place.

This study extends research on the decline of organized labor by analyzing the impact of non-wage benefits on unionization. We expand Ashenfelter and Pencavel's (1969) econometric specification to estimate how the growing prominence of benefits has affected union density. Using time-series observations for 1948-1997, we examine unionization nationwide. With cross-sectional data for 1983-1996, we also analyze the repercussions for organized labor across states.

Benefits and Unionization: Theoretical Issues

To consider how benefits may affect unionization, consider a simple model of bargaining between a union and employer. Suppose the union's surplus, U, can be written as follows:

$$U = (w - w_0)L, \tag{1}$$

where L is employment, w is the rate of pay (wages and benefits) under a labor contract, and w_0 is "fall-back" pay, that is, what a worker can receive during a labor dispute in the absence of a contract. The firm's surplus, F, is defined as:

$$F = R(L) - wL - \pi_0, \tag{2}$$

where $R(L)$ is the firm's revenue function ($R_L > 0$, $R_{LL} < 0$), and π_0 represents fall-back profit without a labor settlement. Assume the parties bargain to maximize the weighted product of their surpluses, WS:

$$WS = U^{\alpha}F^{1-\alpha}, \tag{3}$$

where the weight α represents the union's leverage, or its bargaining power.

A simple approach is to assume that the union and firm settle on the pay rate w that maximizes the weighted surplus WS; the firm then chooses employment so that $R_L = w$. Such an arrangement is known as "on-the-demand curve" bargaining. It is easily shown, however, that a more efficient outcome occurs when the parties choose contract pay w and employment L to maximize WS, so-called "efficient bargaining."

Whether there is bargaining on-the-demand curve or efficient bargaining, equation (3) can be differentiated to yield optimizing conditions, from which it is easily verified that $\partial w/\partial \alpha > 0$. Other things equal, when the union has less leverage in bargaining, it obtains a smaller share of the "pie" to be divided, meaning lower contract pay. It is also easily seen that $\partial w/\partial w_o > 0$. The lower is expected compensation absent an agreement, the costlier it is for the union to carry on a labor dispute, resulting in lower contract pay.

Fringe benefits are typically suspended during a strike and not easily replaced by individual workers. A primary example is health insurance, which even at group rates might cost a worker several hundred dollars per month (Pauly, 1997, p. 1). It is likely that many workers cannot purchase sufficient insurance to replace lost coverage. Management now often hires replacement workers for strikers, threatens to relocate facilities if a strike continues, or unilaterally implements conditions of employment. Consequently, as benefits become a more prominent part of compensation, organized labor's expected fall-back position is jeopardized.

In a famous formulation, Chamberlain (1955, pp. 80-5) asserted that as the cost of disagreeing with management increases, union bargaining power is diminished. Thus, the more vulnerable expected fall-back position *and* resulting loss of union bargaining leverage combine to diminish w. With a smaller payoff, the benefits of union membership may be reduced, triggering a decline in membership and union density.

Benefits and Unionization: Nationwide Analysis

Data and Sample. Prior to World War II non-wage benefits were negligible. They became more common during the war, when pay freezes were imposed. A consistent fringe benefits series for the economy extends back to 1948; the latest information available is for 1997. Thus, our sample covers the fifty-year period 1948-1997. In our regression analysis, we use *FRINGE* as an explanatory variable, which represents the percentage of total compensation accounted for by non-wage benefits (for details, see "Data Appendix").

The dependent variable is *UNIZ*, the fraction of the private, non-farm labor force that is unionized. Previous empirical studies have used different series for union density. A popular series has been constructed by Troy and Sheflin (1985), which is based on the financial reports filed by organized labor under the Landrum-Griffin Act. For each union, they divided per capita revenue by the organization's per capita dues rate to compute average annual, full-time, dues-paying members. Summing over all unions yields the overall level of membership.

Unfortunately, the Troy and Sheflin series was discontinued after 1982.

Currently the U.S. Bureau of Labor Statistics (BLS) compiles a union density series using data from the Current Population Survey (CPS).[2] Because of changes in bureaucracy and survey design, the CPS does not provide a consistent unionization series prior to 1959. Moreover, for 1982 the CPS provided no data on union membership.

Some authors have asserted — without elaboration — that the Troy and Sheflin series has fewer inaccuracies than the CPS series (Neumann and Rissman, 1984, p. 178). But what happens when one seeks to analyze unionization after 1982, when the Troy and Sheflin series was discontinued? In that case it has been common to combine the two series.[3] Following the same approach, we combined the Troy and Sheflin and CPS data to construct a *UNIZ* series for 1948-1997. Specifically, we used Troy and Sheflin observations for 1948-1982 and CPS figures for 1983-1997 (see appendix).

Specification. We estimate an equation for unionization (*UNIZ*) using a specification similar to that of Ashenfelter and Pencavel, who specified union density as a function of (1) previous unionization, (2) structural factors, (3) the socio-political environment, and (4) the labor relations climate. We add to this list by including *FRINGE* as an explanatory variable. Specifically, we hypothesize the following relationship for unionization at time t, $UNIZ_t$:

$$UNIZ_t = \beta_0 + \beta_1 UNIZ_{t\text{-}1} + \beta_2 FRINGE_t + \beta_3 REAL\ PAY_t + \beta_4 DEMOCRAT_t + \beta_5 MANUFACTURE_t + \beta_6 WORK\ STOP_t + \beta_7 UNEMPLOY_t + \beta_8 CPS_t + \varepsilon_t, \qquad (4)$$

where the explanatory variables are as follows (see data appendix for sources):

$UNIZ_{t\text{-}1}$:	lagged union density,
$FRINGE_t$:	share of total compensation accounted for by non-wage benefits,
$REAL\ PAY_t$:	average hourly earnings (wages plus benefits) of private, non-farm, production workers (1992 dollars),
$DEMOCRAT_t$:	percentage of Democratic members of Congress,
$MANUFACTURE_t$:	share of the civilian labor force in the manufacturing sector,

WORK STOP$_t$:	number of work stoppages (involving 1,000 or more workers),
UNEMPLOY$_t$:	civilian unemployment rate, and
CPS$_t$:	indicator variable, equals 1 for 1983-1997 (when CPS data are used in the *UNIZ* series).

We follow other researchers in using the lagged value of union density, $UNIZ_{t-1}$, as an explanatory variable (Moore, et al., 1989; Neumann and Rissman, 1984). This variable accounts for possible inertia in union membership. That is, workers who belonged to a union in time period t-1 are likely to be members in time period t, so we expect $\beta_1 > 0$.

Among the explanatory variables, we introduce *FRINGE*. Other things equal — including overall pay — we hypothesize that union density has been adversely affected as *FRINGE* has risen. Thus we expect $\beta_2 < 0$.

The remaining explanatory variables control for other economic and socio-political factors. *REAL PAY* is the average hourly earnings of production workers. Over the fifty-year sample period, real wages plus benefits increased from \$6.80 per hour to \$11.45 (1992 dollars). The impact of pay on union density is not necessarily clear. If production workers earn relatively high pay, they might not seek union membership. Alternatively, they may desire unionization to protect pay scales.

To proxy the political climate, we follow previous studies by including *DEMOCRAT*. Others have argued that a pro-union legislative climate is more likely to emerge the more there is Democratic party representation in Congress (Ashenfelter and Pencavel, 1969; Bain and Elsheikh, 1976). Therefore we expect $\beta_4 > 0$. Over the sample period, the Democratic share of Congress fluctuated between 43 percent and 68 percent, averaging nearly 58 percent.

Changes in the structure of the economy may affect union density. In considering structural change, it has been common to focus on manufacturing, a traditional bastion of organized labor (Lumsden and Petersen, 1975; Stepina and Fiorito, 1986). That sector's share of the civilian labor force, represented by the variable *MANUFACTURE*, has fallen steadily, from nearly 28 percent to little more than 15.5 percent. We expect decreases in *MANUFACTURE* to inhibit unionization, so $\beta_5 > 0$.

The *WORK STOP* variable can be used to reflect the industrial relations climate. Since the late 1940s the number of major work stoppages has averaged 217 per year but has varied considerably, from 29 to 470. An increase in major work stoppages could reflect managerial opposition to union demands, in which case we would expect a negative sign for β_6. More work stoppages, however, could be a result of increased union assertiveness. If such assertiveness is seen as a benefit of union membership, β_6 could be positive. Therefore, we have no *a priori* hypothesis about the sign on the *WORK STOP* coefficient.

Earlier studies account for unemployment (or some measure of employment) as an important determinant of union membership (Koeller, 1994). A higher unemployment rate could lead workers to seek collective power and protection. But when joblessness is more widespread, workers might shun union membership because they are worried about antagonizing employers. Thus, the coefficient on the variable *UNEMPLOY* could be positive or negative.

The union density variable *UNIZ* consists of observations from the Troy and Sheflin series (1948-1982) and the CPS (1983-1997). Other researchers have ignored similar breaks in their union series. But we noticed that union density drops following the 1983 break point (see Figure 1). This fall could coincide with economic and social factors. It also could be that the CPS figures record lower rates of unionization than did Troy and Sheflin. To account for the structural break in the *UNIZ* variable, we also include an indicator variable *CPS* (equals 1 for 1983-1997; 0 otherwise). If the break results in a lower measure of union density, there will be a negative coefficient on the *CPS* variable, $\beta_8 < 0$.

Stationarity. To determine whether simple OLS regression can be used to estimate the union density equation, we must first determine whether the data are stationary. With non-stationary data there might be a problem of spurious correlation: The time-varying nature of the series may — but not necessarily — lead OLS estimates to appear significant when they are merely happenstance (Dhrymes, 1998, pp. 55-71; Kennedy, 1998, pp. 268-69).

We tested the data for stationarity following the procedure outlined by Enders (1995, pp. 256-60).[4] All the series are stationary except the *FRINGE* variable.

Values for *FRINGE* increased between 1948 and 1993, dropping off thereafter. This observation led us to test whether *FRINGE* was stationary for 1948-1993. Focusing only on this subperiod, we found that *FRINGE* indeed was stationary. Therefore we used the shorter data set initially, using OLS regression to estimate a union density equation. We then estimated a separate equation with the full data set to see whether the results of the two analyses are similar.

Empirical Results. Focusing on the truncated sample, 1948-1993, we estimated the union density equation and report the results in Table 1 (left column). Lagged union density has a significantly positive coefficient. According to the estimate, increasing $UNIZ_{t-1}$ by 10 points adds more than 6.6 points to the $UNIZ_t$ measure. Not surprisingly, past membership has a positive impact on current unionization.

Our principal interest is in identifying any *FRINGE* effect. The coefficient on *FRINGE* is negative, as hypothesized, and statistically significant. Controlling for overall compensation, increasing the share of non-wage benefits reduces union density. Increasing *FRINGE* by 10 points reduces union density by more than 4.5 points. Not only is this estimate significant, it is also quite large. To put it into perspective, between 1948 and 1993 *FRINGE* rose from 4.57 percent to 18.89 percent, an increase of 14.32 points. Other things equal, more than 6.44 points of the decline in *UNIZ* (-0.452 x 14.32), more than one-fourth of the drop, can be attributed to the growing influence of benefits.

Our estimates also indicate that it is important to account for the break in the *UNIZ* series. Controlling for other factors, union density is nearly 1.4 points lower using the CPS figures.

For the variables *REAL PAY*, *WORK STOP*, and *UNEMPLOY*, it was not clear *a priori* whether the coefficients would be positive or negative. In all three cases the OLS coefficients are not statistically significant, meaning the estimates are effectively zero. Also, unionization is not significantly affected by increasing the Democratic Party's presence in Congress.

In addition to the OLS regression for the truncated sample, we also estimated an equation for the full data set, 1948-1997 (Table 1, right column). Again we find that *FRINGE* has a statistically negative impact on unionization. All else equal, a 10 point rise in *FRINGE* leads to a 3.5 point drop in union density. For the fifty-year sample, non-wage benefits climbed from 4.57 percent of pay to 16.93 percent, a rise of 11.46 points. Therefore, 4.01 points of the drop in *UNIZ* (-0.35 x 11.46) — more than one-sixth of the fall — is due to the expanded role of fringe benefits.

Comparing the regression results from the full data set with those from the truncated sample, the OLS estimates are quite similar (see Table 1). All coefficient signs are the same, estimated values are of similar magnitude, and the *t*-statistics are similar, with one exception. In the full sample, the *MANUFACTURE* coefficient is both positive and significant. As expected, over the fifty-year sample union density has fallen as the manufacturing share of employment has declined.[5]

Because the two estimated equations are so alike, the fact that *FRINGE* is non-stationary over the full sample period does not appear to lead to spurious regression results. Otherwise, we would not expect the estimated equations to be so similar.

Given total compensation, increasing the share of benefits in the pay package has had a significantly negative impact on unionization nationwide. Controlling for other economic and social factors, anywhere from one-sixth to one-fourth of the decline in union density can be attributed to the increased presence of benefits. Because unionization is not evenly distributed across the country, the *FRINGE* impact may well differ from state to state. Therefore we extend our analysis to consider unionization at the state level.

Benefits and Unionization: State-Level Analysis

Overview of the States. Union membership and bargaining power vary from state to state. Likewise, compensation practices may differ. Where unions are strong, organized labor may obtain high levels of benefits too. Where organized labor has not been strong, on the other hand, employers might use fringe benefits to discourage unionization. Given these possibilities, we examine whether the influence of the benefits-wage mix differs across states.

For each of the fifty states and the District of Columbia, figures are available for union density in manufacturing back to 1983. We use these observations to represent *UNIZ*. Observations for *FRINGE* are available through 1996. Thus, our cross-sectional analysis is for 1983-1996 (see appendix). Mean observations for each state are presented in Table 2.

There are wide variations in manufacturing union density, ranging from a low of 4.9 percent in Arizona to a high of 38.6 percent in West Virginia. There is a narrower range for *FRINGE*; nevertheless, values range from a low of 13.2 percent in Delaware to a high of 21.7 percent in West Virginia. Some of the lowest values for *FRINGE* can be found along the Atlantic coast (Delaware, New Jersey, and New York) while higher values occur in the Midwest (for example, Indiana, Iowa, and Ohio). Curiously, both the highest and lowest *FRINGE* values occur in relatively high union density states.

To illustrate some of the distinctions more clearly, Table 3 lists the five highest- and lowest-ranked states according to unionization (top panel) and benefits (bottom panel). The high *UNIZ* states exhibit above-average values for *FRINGE*, though there is no direct correspondence in rankings. Unionization is least likely in southern and western states with Right-to-Work (RTW) laws. These states also exhibit relatively low *FRINGE* values, but again there is no direct correspondence in the rankings. A simple Spearman rank correlation test confirms that the ranks of *UNIZ* and *FRINGE* are indeed correlated.[6]

Specification. Nationwide, of course, unionization declined over the 1983-1996 sample period. Every state experienced a drop in union density, though there was variation across states.[7] In our regression analysis we use union density in manufacturing in state i at time t, $UNIZ_{it}$, as the dependent variable. We specify an equation similar to that for the national level analysis:

$$UNIZ_{it} = \gamma_0 + \gamma_1 UNIZ_{it\text{-}1} + \gamma_2 FRINGE_{it} + \gamma_3 REAL\ PAY_{it} + \gamma_4 MFG\ SHARE_{it} + \gamma_5 PART\ TIME_{it} + \gamma_6 RTW_{it} + \gamma_7 UNEMPLOY_{it} + \eta_{it}, \quad (5)$$

where *FRINGE* and *REAL PAY* are defined as before and the other explanatory variables are (sources listed in data appendix):

$MFG\ SHARE_{it}$: state *i*'s share of U.S. manufacturing employment,

$PART\ TIME_{it}$: share of part-time workers in state *i*'s manufacturing sector,

RTW_{it}: indicator variable, equals 1 for a state with a Right-to-Work law, and

$UNEMPLOY_{it}$: unemployment rate in state *i*.

Similar to the national level analysis, we hypothesize $\gamma_1 > 0$ and $\gamma_2 < 0$, with no *a priori* expectations for γ_3 or γ_7.

The structure of a state's economy is likely to affect union density. We use *MFG SHARE*, a state's share of U.S. manufacturing employment, to reflect that state's economic profile. Across the states, *MFG SHARE* varies from a low of 0.1 percent to a high of 11 percent. Other things equal, we expect more manufacturing-oriented states to be more unionized, $\gamma_4 > 0$.

At the national level we used major work stoppages as a reflection of the industrial relations climate. Such work stoppage data are not disaggregated by state, so we must consider another measure. Other researchers have considered part-time employment as such an indicator of industrial relations (Hernández, 1995; Riddell, 1993). One way manufacturers might hamper organized labor is to hire part-time workers, who are less likely to unionize or join existing unions. In some states as little as 2 percent of manufacturing workers were part-timers, while in others the share was as high as 24 percent. In the regression we include *PART TIME*, the share of part-time workers in a state's manufacturing sector, expecting $\gamma_5 < 0$.

The socio-political environment for organized labor differs across states. Throughout the 1983-1996 period, twenty-one states had Right-to-Work laws. Such legislation means that workers cannot be compelled to support labor unions. It is well known that union density is lower in Right-to-Work states, so we include the *RTW* indicator for such states, expecting $\gamma_6 < 0$.

With pooled data it is common to estimate parameters with a fixed effects specification. In such a setting a separate intercept term is calculated for each cross section while coefficients on the independent variables are common across all cross sections. But in this case there are two fundamental problems with such an approach. First, the *FRINGE* effect would not be allowed to vary across states, which is what we wish to examine. Second, for the 1983-1996 sample period RTW laws were present in the same twenty-one states, meaning the *RTW* variable would be perfectly collinear with the intercept terms.

Instead of a fixed effects specification, we estimate a model with a common constant and allow the *FRINGE* coefficient to vary across states. We also explored the possibility of letting the other slope coefficients vary. Following the testing procedure outlined by Balestra (1996, p. 37), we found it appropriate to vary only the *FRINGE* coefficients.[8] Thus, the equation we estimate is:

$$UNIZ_{it} = \gamma_0 + \gamma_1 UNIZ_{it-1} + \gamma_{2i} FRINGE_{it} + \gamma_3 REAL\ PAY_{it} + \gamma_4 MFG\ SHARE_{it} + \gamma_5 PART\ TIME_{it} + \gamma_6 RTW_{it} + \gamma_7 UNEMPLOY_{it} + \eta_{it}. \qquad (5')$$

The pooled model represented in equation (5') can be estimated by ordinary least squares regression as long as there are no problems with the disturbances. As always, with cross-sectional analysis we must beware of potential heteroskedasticty. Following the Lagrange multiplier test procedure outlined by Greene (1993, p. 450), we found evidence of hetero-skedasticity.[9] Consequently, we estimated the equation using generalized least squares.

Empirical Results. The GLS coefficient estimates are presented in Table 4. As expected, unionization is positively related to lagged union density. The structure of a state's economy is also significant. The *MFG SHARE* coefficient estimate indicates that a 10 point increase in a state's share of U.S. manufacturing employment adds nearly 11 points to union density in that state. States with relatively high unemployment also tend to be more unionized. The presence of an RTW law reduces *UNIZ* by more than 9.7 points. As was the case nationally, *REAL PAY* does not affect unionization significantly. Although the *PART TIME* coefficient is negative (as expected), it is not significantly different than zero.

Turning to the influence of benefits, nearly all of the *FRINGE* coefficients are significantly negative. Not only is unionization nationwide affected by the growing role of benefits, in forty-six states union density is negatively related to benefits (Table 4). This *FRINGE* effect is most pronounced for California and New York. For the two most populous states, a 10 point rise in *FRINGE* means decreases in union density of more than 14 points and 10 points, respectively. At the other extreme is Alabama, where the influence is a mere 1.8 points. Although *FRINGE* generally exerts a negative impact on organized labor, there is wide variation across the states.

Examining the *FRINGE* coefficients from Table 4 in more detail, we find other important tendencies. Union-oriented states tend to exhibit a relatively small *FRINGE* effect: Four of the top five ranked union states are in the bottom half of the rankings for *FRINGE* magnitude. In union-oriented states benefits are fairly prominent. Moreover, in such states organized labor is less sensitive to changes in the composition of pay.

Low *UNIZ* states, in contrast, tend to have relatively large *FRINGE* coefficients. Of the ten states with the strongest *FRINGE* effects, seven are in the bottom half of the rankings by union density. In states where organized labor does not have a firm hold, non-wage benefits appear to discourage unionization.

In the ten states where the *FRINGE* effect is most pronounced, eight are in the bottom half of the rankings in terms of benefits. Unionization is most adversely affected in those states where the benefits share has lagged behind the national average. Furthermore, several of those states — California, Massachusetts, New Jersey, and New York — are known for relatively high taxes. In high-wage, high-tax states workers may be happy to see pay tilted more towards untaxed fringe benefits. When *FRINGE* is increased, perhaps workers are pacified, thereby undermining support for organized labor.

Most of the states in which the *FRINGE* effect is weakest are in the south or west. Many of these states have RTW laws. Moreover, states like Alabama, Idaho, Louisiana, and South Dakota have relatively small manufacturing sectors and low taxes. Likewise, states with no significant *FRINGE* effect are similar. All five of them — Iowa, Kansas, Nebraska, North Dakota, and Wyoming — have RTW laws (four of them are even contiguous).

To summarize our findings, the negative impact of *FRINGE* on union density is strongest for high population states with a large manufacturing presence, comparatively low union density, and high taxes. Where there have been relatively high taxes and modest benefits, and no RTW law, employers may use fringe benefits strategically to deter organized labor. In the Right-to-Work states of the south and west, where manufacturing is also less prominent, the influence of benefits is smaller.

In our national level analysis we found that between one-sixth and one-fourth of the decline in union density could be attributed to the *FRINGE* effect. Looking at the states individually, again we find that the impact of *FRINGE* is both significant and large.

For each state we calculated the change in union density that could be attributed to the expanded role of benefits (Table 5). For example, in California *UNIZ* dropped by 11.8 percentage points over the sample period. The *FRINGE* coefficient is -1.462 and the state's *FRINGE* values increased by 1.683 points. Consequently, *FRINGE* accounts for approximately 2.5 points (-1.462 x 1.683) of the 11.8 point decrease in union density, more than one-fifth of the drop. As seen in Table 5, other things equal, *FRINGE* accounted for anywhere from 2 percent of the reduction in union density (Idaho) to more than half of the decline (Colorado, District of Columbia, Maine, Nevada, New Mexico, South Carolina, and Vermont).

Concluding Remarks

The changing nature of compensation evidently has affected union density. In the private sector nationwide, as fringe benefits have become a more prominent component of workers' pay, *ceteris paribus*, union density has declined. Over the fifty-year period 1948-1997, at least one-sixth of the drop in unionization can be attributed to the growing role of non-wage benefits.

Unionization has eroded as workers have been paid more fringe benefits relative to wages. Has organized labor been victimized by past successes in securing fringe benefits? As benefits such as health care and pensions have become more common, especially in heavily populated areas like California and New York, perhaps workers have felt less urgency to unionize. Or perhaps workers have become more reluctant to risk benefits in industrial action, thereby weakening support for unions. An issue for further study would be how union organizing efforts or strike activities have been influenced by fringe benefits.

For the private sector in general and manufacturing in particular, we have established that unionization is related to the composition of pay. Yet the influence of non-wage benefits may vary across industries or even different subsectors of manufacturing. It remains to be seen whether there are similar findings for other sectors.

Not only has the growing prominence of fringe benefits affected unionization overall, but also in virtually every state. In states where unions are relatively strong, organized labor has been less sensitive to changes in the composition of pay. But in states where the share of fringe benefits has lagged behind, especially high tax states, organized labor is particularly sensitive to increases in benefits. In such areas, chiefly those without Right-to-Work laws, employers may have used non-wage benefits to thwart unionization. Perhaps future research will investigate in more detail whether firms indeed have altered the wage-benefits mix strategically.

DATA APPENDIX

National Data

Union Density. *UNIZ* figures for 1948-1982 are available from Troy and Sheflin (1985, pp. A1-A3). Observations for 1983-1997 are from Hirsch and Macpherson (1998, p. 12). For bibliographic citations, see "Data Sources" below.

Non-wage Benefits. To compute *FRINGE*, we divided "Supplements to Wages" by "Total Compensation," both of which are reported by the Bureau of Economic Analysis (U.S. BEA, 1998, pp. 163-66).

Other Explanatory Variables. *REAL PAY* is average hourly earnings (wages plus benefits) of non-farm production workers. Nominal earnings were downloaded from the Bureau of Labor Statistics (U.S. BLS) website (<**www.bls.gov**>, Series ID: EEU00500006). We calculated real earnings (1992 dollars) using the GDP deflator (U.S. BEA, 1998, pp. 159-62).

DEMOCRAT is the percentage of Democratic members of Congress. The number of Congressional Democrats is reported in the *World Almanac* (1999, pp. 89-90).

MANUFACTURE, the percentage of the civilian labor force in manufacturing, is private sector manufacturing employment divided by the civilian labor force. The former was downloaded from the BLS (Series ID: LFU11110020000); the latter is reported by the U.S. BLS (2000, p. 166).

The U.S. BLS (1998, p. 70) reports the number of work stoppages involving more than 1,000 workers, *WORK STOP*. We downloaded the civilian unemployment rate, *UNEMPLOY*, from the BLS [Series ID: LFS21000000].

State-Level Data

Union Density. *UNIZ* for 1983-1996 is reported by Hirsch and Macpherson (1994-1998) in the table "Union Membership, Density, Employment, and Earnings in Private Sector Manufacturing by State."

Non-wage Benefits. For 1987 and 1992 *FRINGE* is "Fringe Benefits" divided by "Total Compensation," information recorded by the Bureau of the Census (U.S. BOC) in its *Census of Manufactures* (1987, Table 2; 1992, Table 2-3a).

For 1983-1986, 1989-1991, and 1994-1996 *FRINGE* is calculated with U.S. BOC *Survey of Manufactures* data, specifically, "Supplemental Labor Costs" divided by the sum of "Supplemental Labor Costs" and "Payroll." The supplemental costs (another term for fringe benefits) are reported in Table 3 of the *Surveys*; payroll figures (which reflect wage payments) appear in Table 1. No data have been published for 1988 or 1993. We generated values for those years by interpolating.

Other Explanatory Variables. *REAL PAY* is the average hourly earnings (wages plus benefits) of manufacturing production workers. Nominal figures are listed by Hirsch and Macpherson (1994-1998). We calculated real earnings (1992 dollars) using the GDP deflator.

MFG SHARE is a state's share of U.S. manufacturing employment, reported by the U.S. BOC, Table 1 (1986, 1991, 1996).

Hirsch and Macpherson (1994-1998) list both the share of part-time workers in a state's manufacturing sector, *PART TIME*, and the states with Right-to-Work Laws.

From the BLS, we downloaded the civilian unemployment rate, *UNEMPLOY*, for each state and the District of Columbia (Series IDs: LAUST0x000003).

Data Sources

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NOTES

1. For analyses of unionization and the economy's structure, see Ashnefelter and Pencavel (1969), Bain and Elsheikh (1976), Hernández (1995), Jones (1992), Koeller (1994), Lumsden and Petersen (1975), Moore and Newman (1988), and Stepina and Fiorito (1986). Regarding management opposition to organized labor, see Farber (1990).

2. Figures from this series come from a CPS survey question, asking whether the respondent is a member of a union. The CPS program was initiated in 1940 under the auspices of the Work Projects Administration, transferred to the Census Bureau in 1942, and finally transferred to the BLS in 1959 (Manser, 1998).

3. For example, Jones (1992) spliced the two series in her analysis of structural changes in the labor market. Likewise, Booth (1995, p. 13) used the series interchangeably in her examination of union density over time. Comparing unionization between the U.S. and Canada, Riddell (1993) also combined both series.

4. For details of the stationarity testing procedure, see Edwards (2000). Test results are available from the authors.

5. With time-series data we must beware of potential serial correlation. For both the truncated and full data sets, we used the Breusch-Godfrey Lagrange multiplier test for serial correlation (Godfrey, 1988). In both cases, we could not reject the null hypothesis of no serial correlation. The test results are available on request.

6. For details of the testing procedure, see Daniel and Terrell (1989, pp. 697-99). Test results are available from the authors.

7. In nine states and the District of Columbia *UNIZ* actually increased in the latter years of the sample. Those states are: Colorado, Florida, Maine, Maryland, Missouri, Nevada, New Mexico, South Dakota, and Vermont. Yet in all cases, 1996 union density was lower than for 1983.

8. The Balestra test compares a restricted specification, where the *FRINGE* coefficient is the only one allowed to vary across states, to an unrestricted specification, where other parameter

estimates are allowed to vary. The null hypothesis is that only the *FRINGE* parameter estimates vary by state; the test statistic follows an *F* distribution. We performed hypothesis tests for eight different specifications against the null that only the *FRINGE* parameter estimates vary by state. The eight specifications were as follows:

1 and 2. Vary the slope *MFG SHARE* parameter estimates by state, and then vary them by region (with four regional designations used by the Bureau of Labor Statistics);

3 and 4. Vary the *UNEMPLOY* parameter estimates by state, and then by region;

5 and 6. Vary the *PART TIME* parameter estimates by state, and then by region;

7 and 8. Vary the parameter estimates of *all* the independent variables by state, and then by region. In each case, the calculated *F*-statistic indicated that the null hypothesis could not be rejected. Only the separate *FRINGE* coefficient estimates were statistically distinct from one another. We therefore concluded that varying only the *FRINGE* coefficient was the appropriate procedure. Test results are available from the authors.

9. Test statistics can be supplied on request.

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Table 1

Determinants of Union Density Nationwide, 1948-1997†

Explanatory Variable	Hypothesis	Regression 1: 1948-1993 Coefficient (t-statistic)	Regression 2: 1948-1997 Coefficient (t-statistic)
Constant	+ or –	9.606** (3.671)	7.313** (4.028)
UNIZ (lagged)	+	0.662** (9.399)	0.694** (10.418)
FRINGE	–	-0.452** (-3.754)	-0.345** (-3.555)
REAL PAY	+ or –	0.071 (0.410)	-0.089 (-0.660)
DEMOCRAT	+	0.004 (0.291)	0.009 (0.724)
MANUFACTURE	+	0.112 (1.054)	0.171* (1.917)
WORKSTOP	+ or –	0.001 (1.113)	0.001 (1.115)
UNEMPLOY	+ or –	0.122 (1.371)	0.127 (1.509)
CPS (=0 thru 1982; =1 otherwise)	–	-1.379** (-3.248)	-1.130** (-3.006)
$\bar{R}^2$		0.998	0.998

Note: †Sources are detailed in the "Data Appendix." ** (*) Significant at the 0.05 (0.10) level.

Table 2

Union Density and Fringe Benefits by State, 1983-1996†

State	*UNIZ* (mean)	State Rank	*FRINGE* (mean)	State Rank	State	*UNIZ* (mean)	State Rank	*FRINGE* (mean)	State Rank
Alabama*	23.23	18	19.38	10	Montana	25.95	12	21.23	2
Alaska	14.55	34	17.95	28	Nebraska*	16.68	29	17.83	31
Arizona*	4.94	51	17.58	35	Nevada*	9.23	44	17.93	29
Arkansas*	16.35	30	19.24	14	New Hamp.	7.99	46	18.12	25
California	15.10	32	17.50	38	New Jersey	24.21	15	14.83	50
Colorado	8.87	45	17.24	40	New Mexico	9.82	43	18.28	23
Connecticut	17.10	28	17.19	41	New York	25.14	13	15.96	47
Delaware	22.17	20	13.23	51	N. Carolina*	5.32	50	16.84	46
D.C.	14.49	35	14.93	49	N. Dakota*	18.94	23	19.08	16
Florida*	7.14	47	17.07	43	Ohio	33.75	4	20.09	7
Georgia*	11.54	39	17.15	42	Oklahoma	18.04	26	18.79	19
Hawaii	31.68	6	18.00	26	Oregon	22.41	19	16.86	45
Idaho*	18.10	25	17.06	44	Pennsylvania	29.82	7	18.63	20
Illinois	26.87	11	17.75	32	Rhode Is.	14.41	37	18.58	21
Indiana	38.37	2	20.43	4	S. Carolina*	5.65	49	17.69	34
Iowa*	27.77	10	20.27	5	S. Dakota*	11.53	40	17.96	27
Kansas*	20.13	21	20.18	6	Tennessee*	17.71	27	19.10	15
Kentucky	29.69	8	19.37	11	Texas*	10.10	42	17.56	36
Louisiana*	18.59	24	19.51	9	Utah*	6.93	48	17.56	37
Maine	23.88	16	21.20	3	Vermont	10.61	41	19.25	13
Maryland	23.71	17	18.87	18	Virginia*	14.57	33	17.85	30
Mass.	15.35	31	17.28	39	Washington	32.68	5	19.26	12
Michigan	37.91	3	19.91	8	W. Virginia	38.56	1	21.70	1
Minnesota	19.99	22	15.51	48	Wisconsin	29.28	9	18.55	22
Mississippi*	14.42	36	18.19	24	Wyoming*	13.38	38	19.05	17
Missouri	24.83	14	17.73	33					

Note: †Sources are detailed in the "Data Appendix." *State with a Right-to-Work law.

Table 3

Union Density and Fringe Benefits: Top and Bottom States†

UNIZ: Top 5 States

State	*UNIZ* (mean)	State Rank	*FRINGE* (mean)	State Rank
W. Virginia	38.56	1	21.70	1
Indiana	38.37	2	20.43	4
Michigan	37.91	3	19.91	8
Ohio	33.75	4	20.09	7
Washington	32.68	5	19.26	12

UNIZ: Bottom 5 States

State	*UNIZ* (mean)	State Rank	*FRINGE* (mean)	State Rank
Arizona*	4.94	51	17.58	35
N. Carolina*	5.32	50	16.84	46
S. Carolina*	5.65	49	17.69	34
Utah*	6.93	48	17.56	37
Florida*	7.14	47	17.07	43

FRINGE: Top 5 States

State	*UNIZ* (mean)	State Rank	*FRINGE* (mean)	State Rank
W. Virginia	38.56	1	21.70	1
Montana	25.95	12	21.23	2
Maine	23.88	16	21.20	3
Indiana	38.37	2	20.43	4
Iowa*	27.77	10	20.27	5

FRINGE: Bottom 5 States

State	*UNIZ* (mean)	State Rank	*FRINGE* (mean)	State Rank
Delaware	22.17	20	13.23	51
New Jersey	24.21	15	14.83	50
D.C.	14.49	35	14.93	49
Minnesota	19.99	22	15.51	48
New York	25.14	13	15.96	47

Note: †Sources are detailed in the "Data Appendix." *State with a Right-to-Work law.

Table 4

Determinants of Union Density by State, 1983-1996†

Explanatory Variable	Hypothesis	Coefficient (*t*-statistic)	*FRINGE* coefficients (*t*-statistics), by state Hypothesis: –	Rank			Rank	
Constant	+ or –	18.279** (5.564)	Ala.	47	-0.179** (-1.984)	Ida.	46	-0.185* (-1.659)
UNIZ (lagged)	+	0.598** (20.908)	Alk.	14	-0.782** (-4.328)	Ill.	11	-0.846** (-6.279)
REAL PAY	+ or –	-0.037 (-0.321)	Arz.	30	-0.501** (-4.849)	Ind.	34	-0.414** (-3.804)
MFG SHARE	+	1.071** (3.33)	Ark.	42	-0.269** (-2.913)	Iowa	51	-0.011 (-0.908)
PART TIME	–	-0.093 (-1.558)	Cal.	1	-1.462** (-7.176)	Kan.	49	-0.127 (-1.434)
RTW	–	-9.733** (-3.461)	Col.	3	-0.974** (-6.582)	Ky.	28	-0.526** (-4.257)
UNEMPLOY	+ or –	0.330** (5.472)	Conn.	9	-0.862** (-6.212)	Lou.	44	-0.216** (-2.335)
			Del.	12	-0.820** (-4.363)	Me.	29	-0.512** (-4.112)
			D.C.	5	-0.912** (-4.652)	Md.	24	-0.580** (-4.542)
			Fla.	25	-0.567** (-5.146)	Mass.	4	-0.962** (-7.104)
			Geo.	31	-0.499** (-4.557)	Mich.	27	-0.534** (-4.880)
			Haw.	36	-0.361** (-2.109)	Minn.	10	-0.856** (-5.721)

Table 4 (continued)

Determinants of Union Density by State, 1983-1996†

FRINGE coefficients (*t*-statistics), by state
Hypothesis: −

	Rank			Rank			Rank	
Miss.	37	-0.356** (-3.579)	N.C.	17	-0.725** (-5.695)	Tenn.	40	-0.310** (-3.313)
Mo.	19	-0.688** (-5.142)	N.D.	50	-0.086 (-0.791)	Tex.	20	-0.668** (-5.238)
Mont.	32	-0.458** (-3.483)	Ohio	22	-0.632** (-5.481)	Utah	35	-0.408** (-3.998)
Neb.	48	-0.165 (-1.583)	Okla.	18	-0.714** (-5.412)	Ver.	13	-0.787** (-5.706)
Nev.	39	-0.345** (-3.174)	Ore.	21	-0.644** (-4.753)	Vir.	38	-0.348** (-3.412)
N.H.	7	-0.912** (-6.481)	Penn.	16	-0.758** (-5.990)	Wash.	33	-0.451** (-3.790)
N.J.	6	-0.912** (-6.005)	R.I.	15	-0.766** (-5.587)	W.V.	41	-0.301** (-2.701)
N.M.	8	-0.875** (-5.835)	S.C.	26	-0.545** (-5.259)	Wisc.	23	-0.590** (-4.721)
N.Y.	2	-1.022** (-6.614)	S.D.	43	-0.266** (-2.561)	Wyo	45	-0.196 (-1.581)
$\bar{R}^2$		0.979						

Note: †Sources are detailed in the "Data Appendix." ** (*) Significant at the 0.05 (0.10) level.

Table 5

Impact of FRINGE *on* UNIZ *by State, 1983-1996*†

State	Δ*UNIZ*, 1983-1996	Δ*UNIZ*, due to *FRINGE*	Share (%)	State	Δ*UNIZ*, 1983-1996	Δ*UNIZ*, due to *FRINGE*	Share (%)
Alabama*	-9.4	-0.5	5.3	Montana	-19.1	-2.0	10.5
Alaska	-16.0	-3.3	20.6	Nebraska*	-8.5	0.0	0.0
Arizona*	-3.9	-1.5	38.5	Nevada*	-1.2	-0.7	58.3
Arkansas*	-6.2	-1.0	16.1	New Hamp.	-5.7	-2.2	38.6
California	-11.8	-2.5	21.2	New Jersey	-10.8	-1.4	13.0
Colorado	-3.5	-2.4	68.6	New Mexico	-3.0	-3.2	106.7
Connecticut	-17.5	-2.7	15.4	New York	-11.4	-2.2	19.3
Delaware	-11.1	-3.2	28.8	N. Carolina*	-4.4	-2.0	45.5
D.C.	-2.7	-2.9	107.4	N. Dakota*	-9.6	0.0	0.0
Florida*	-3.3	-1.3	39.4	Ohio	-9.6	-1.6	16.7
Georgia*	-8.2	-1.5	18.3	Oklahoma	-8.8	-4.3	48.9
Hawaii	-16.8	-1.1	6.5	Oregon	-9.0	-1.1	12.2
Idaho*	-7.8	-0.2	2.6	Pennsylvania	-19.4	-0.9	4.6
Illinois	-8.8	-1.4	15.9	Rhode Is.	-5.9	-1.7	28.8
Indiana	-19.1	-1.2	6.3	S. Carolina*	-1.2	-1.3	108.3
Iowa*	-18.8	0.0	0.0	S. Dakota*	-9.3	-0.9	9.7
Kansas*	-5.5	0.0	0.0	Tennessee*	-6.6	-1.2	18.2
Kentucky	-14.8	-1.6	10.8	Texas*	-7.6	-2.3	30.3
Louisiana*	-5.9	-0.8	13.6	Utah*	-9.5	-1.6	16.8
Maine	-1.2	-2.1	175.0	Vermont	-5.3	-2.9	54.7
Maryland	-7.2	-1.5	20.8	Virginia*	-9.4	-1.1	11.7
Mass.	-16.5	-1.6	9.7	Washington	-6.5	-1.7	26.2
Michigan	-11.3	-1.5	13.3	W. Virginia	-12.1	-1.4	11.6
Minnesota	-5.3	-1.9	35.8	Wisconsin	-12.1	-0.7	5.8
Mississippi*	-8.2	-1.0	12.2	Wyoming*	-2.6	0.0	0.0
Missouri	-10.4	-1.3	12.5				

Note: †Sources are detailed in the "Data Appendix." *State with a Right-to-Work law.

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