DECLINING UNIONIZATION: DO FRINGE BENEFITS MATTER?

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INTRODUCTION AND REVIEW

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. Seminal studies by Ashenfelter and Pencavel [1969], Pencavel [1971], and Bain and Elsheikh [1976] have spawned a vast literature in which researchers seek to identify factors responsible for the decline. Factors in the decline include changes in the structure of the economy, management opposition to organized labor, and other socio-political factors.[1]

The structure of compensation also has changed dramatically over time. In 1948 non-wage benefits accounted for nearly 4.5 percent of total compensation; by 1994 that ratio had more than quadrupled, to 19 percent [U.S. Bureau of Economic Analysis, 1998]. Despite this correlation of events, 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 pay, or employer payments for social security and unemployment insurance. The costliest benefits to provide are paid leave, health insurance, and social security contributions [U.S. Bureau of Labor Statistics, 1997; Lettau and Buchmueller, 1999]. The first two are provided voluntarily; social security is mandated by law. Most fringe benefits are paid voluntarily, with mandated benefits constituting roughly only one-third of fringe payments.[2] This study examines whether there is a relationship between benefits and unionization.

Union members receive greater non-wage benefits, on average, than nonunion workers (see Freeman and Medoff [1984]). Moreover, organized labor has bargained and lobbied for welfare improvements, like shorter working hours, compensation for workplace injuries, and unemployment insurance. Many improvements sought

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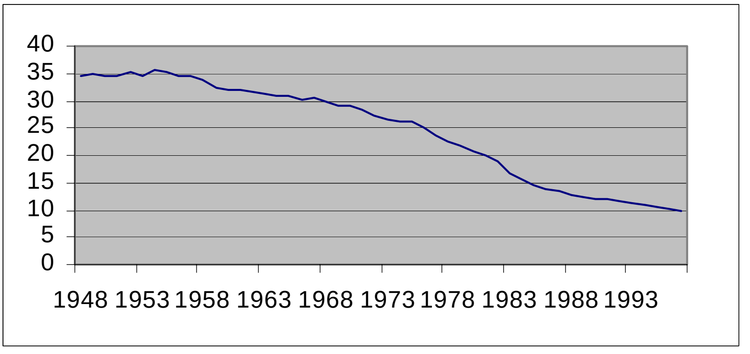
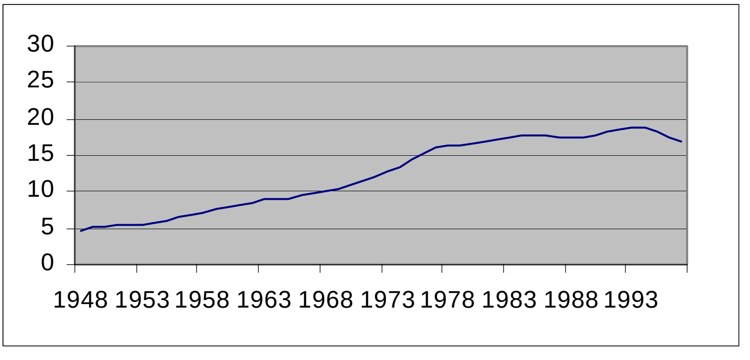
FIGURE 1
Union Density Nationwide: Percentage of Private Non-farm Labor Force that is Unionized



Non-wage Pay: Percentage of Total Compensation



by organized labor, in fact, have been institutionalized by statute. Ironically, this legislative success might undermine support for organized labor.

Some scholars have argued that governmental mandates can replace some union functions. This "substitution hypothesis" was advanced by Neumann and Rissman [1984], who reported that union membership is inversely related to government workplace protections and social welfare programs.[3] According to Bennett and Taylor [2001, p. 261], government-mandated benefits reduce the "scope of issues that may be addressed at the bargaining table." So if legislation makes certain benefits universal, unionization could suffer.[4] In fact, Fiorito [2001] found an inverse relationship between voluntary provision of benefits and workers' intentions to vote for union certification.

Some benefits, like health insurance, can provide tax breaks for both firms and workers, so both sides have reasons to favor them. But employers may have another incentive to tilt the compensation mix towards fringe benefits. Wages and fringe benefits are suspended during a strike. Wages can be partially replaced, by union strike funds or if workers take temporary jobs elsewhere. But temporary jobs typically provide no benefits, meaning health care coverage is lost during a strike. As health care has become increasingly expensive, workers may be reluctant to jeopardize their coverage by risking prolonged industrial action.[5]

As fringe benefits have become more universal, that may have removed issues from potential collective bargaining. And as fringe benefits have become a more prominent feature of compensation, support for organized labor may have weakened.

This article extends research on the decline of organized labor by analyzing the impact of non-wage benefits on unionization. Specifically, we expand Pencavel's [1971] econometric specification to estimate how the growing role of benefits has affected union density. Thus, our study documents whether the "substitution hypothesis" can be amended. Using time-series observations for 1948-1997, we examine whether unionization nationwide has been influenced by the changing composition of pay. With cross-sectional data for 1983-1996, we also analyze the repercussions for organized labor across states.

BENEFITS AND UNIONIZATION: NATIONWIDE ANALYSIS

Data and Sample

Prior to World War II non-wage benefits were negligible. During the war pay freezes were imposed, so benefits became a means of increasing total compensation. A consistent fringe benefits series for the economy extends back to 1948; the latest information we have 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 (see appendix for all data sources).

The dependent variable is *UNIZ*, the fraction of the private, non-farm labor force that is unionized. Previous empirical studies have used different data to represent this series. 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). 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 (for example, 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.[6] 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.

Specification

We estimate an equation for unionization at time *t* ($UNIZ_t$) using a specification similar to that first proposed by Ashenfelter and Pencavel [1969] and Pencavel [1971], who specified union density as a function of (1) previous unionization, (2) labor market factors, (3) structural factors, and (4) the socio-political environment. We add to this list by including *FRINGE* as an explanatory variable. Specifically, we hypothesize the following relationship for $UNIZ_t$:

$$(1) \quad UNIZ_t = \beta_0 + \beta_1 UNIZ_{t-1} + \beta_2 FRINGE_t + \beta_3 RAISE_t + \beta_4 DEMOCRAT_t + \beta_5 MANUFACTURE_t + \beta_6 \Delta UNFAIR_t + \beta_7 UNEMPLOY_t + \beta_8 CPS_t + \varepsilon_t,$$

where the explanatory variables are as follows:

$UNIZ_{t-1}$:	union density (lagged),
$FRINGE_t$:	share of total compensation accounted for by non-wage benefits,
$RAISE_t$:	percentage change in "wage only" component of pay (1992 dollars) for private, non-farm, production workers,
$DEMOCRAT_t$:	percentage of Democrat members of Congress,
$MANUFACTURE_t$:	share of the civilian labor force in the manufacturing sector,
$\Delta UNFAIR_t$:	change in the number of unfair labor practice allegations filed by firms against organized labor,
$UNEMPLOY_t$:	civilian unemployment rate, and
CPS_t:	0-1 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 [Pencavel 1971; 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 a direct relationship between $UNIZ_t$ and $UNIZ_{t-1}$, meaning $\beta_1 > 0$.

Labor market factors like benefits and pay hikes may influence union membership. To represent these influences, we use as explanatory variables *FRINGE* and *RAISE*. *FRINGE* represents the share of total compensation due to non-wage benefits. *RAISE* is the percentage change in (real) wage for production workers.

If the substitution hypothesis is correct, then as benefits become a more prominent component of compensation, other things equal, union density will decrease. Thus, there would be a negative relationship between *UNIZ* and *FRINGE*, $\beta_2 < 0$.

For a given *FRINGE* rate, suppose there is a bigger raise in take-home pay. The impact on unionization is not clear *a priori*. If workers are collecting bigger wage hikes, for a given split between wage and benefits, they may not be stirred to support unionization. On the other hand, they may desire union representation to protect relatively generous pay scales. Therefore, the influence of *RAISE* on *UNIZ* may be positive or negative.

The remaining explanatory variables control for other economic and socio-political factors. 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]. Over the sample period, the Democratic share of Congress fluctuated between 43 percent and 68 percent, averaging nearly 58 percent. We expect a positive relationship between *UNIZ* and *DEMOCRAT*, $\beta_4 > 0$.

Changes in the structure of the economy may affect union density. When 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, denoted by *MANUFACTURE*, has fallen steadily, from nearly 28 percent to little more than 15.5 percent. We expect decreases in *MANUFACTURE* to inhibit unionization, $\beta_5 > 0$.

The variable Δ*UNFAIR* can be used to reflect changes in the industrial relations climate. Since the late 1940s the number of unfair labor practice claims filed by firms has averaged 2,487 per year, but has varied considerably, from 480 to 5,048. An increase in unfair labor practice filing with the National Labor Relations Board (NLRB) could reflect a managerial tactic to oppose organized labor. More allegations by firms could, however, also be a result of increased union assertiveness. If such assertiveness is seen as a benefit of union membership, the effect on *UNIZ* could be positive. Therefore, the sign for the Δ*UNFAIR* can be used to reflect changes in the industrial relations climate. The sign of the coefficient could be either positive or negative.

Earlier studies account for unemployment (or some measure of employment) as a determinant of union membership (for example, Koeller [1994]). Higher unemployment could lead workers to seek collective power and protection. But when joblessness is more widespread, workers might shun unionization because they are worried about antagonizing employers.

The union density variable *UNIZ* consists of observations from the Troy and Sheflin series (1948-1982) and the CPS (1983-1997). In constructing series for their studies, other researchers have ignored this break. But Figure 1 shows that union density drops following the 1983 break point. 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 included 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, $\beta_8 < 0$.[7]

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].[8] All the series are stationary except the *FRINGE* variable. Figure 1 shows that 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. Increasing $UNIZ_{t-1}$ by 10 points adds 8.0 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. Holding other economic and socio-political factors constant, increasing the share for non-wage benefits reduces union density. Increasing *FRINGE* by 10 points reduces union density by almost 3.1 points. Not only is this estimate statistically 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, approximately 4.42 points of the drop in *UNIZ* (-0.309 x 14.32) is due to the *FRINGE* effect. That is, more than fifteen percent of the decline in unionization can be attributed to the growing influence of benefits.

The *RAISE* variable also exhibits a negative influence on union density. The coefficient estimate implies that increasing a wage hike by one percentage point shaves *UNIZ* by more than one-tenth of a point. For a given *FRINGE* level, bigger pay raises for production workers evidently do not stir up support for organized labor.

Union density is also affected by other factors. The *DEMOCRAT* coefficient is significantly positive, suggesting that union density is affected by the political climate. Changes in the industrial relations climate also influence unionization. The Δ*UNFAIR* coefficient is negative, albeit small, suggesting that there is some managerial opposition that inhibits organized labor. 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 0.7 points lower using the CPS data. The *MANUFACTURE* coefficient is positive as expected, but not significant.

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: a 10-point rise in *FRINGE*, other factors held constant, leads to a 2.8 point drop in union density. For the full fifty-

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 -	7.0163 **	5.9124 **
		(3.9265)	(5.5707)
UNIZ	+	0.8035 **	0.8251 **
(lagged)		(12.5803)	(15.9970)
FRINGE	-	-0.3092 **	-0.2846 **
		(-6.3222)	(-6.9642)
RAISE	+ or -	-0.1075 **	-0.1148 **
		(-2.8318)	(-3.7273)
DEMOCRAT	+	0.0256 ***	0.0269 **
		(1.9061)	(2.7236)
MANUFACTURE	+	0.0421	0.0470
		(0.4861)	(0.5965)
Δ*UNFAIR*	+ or -	-0.0002 ***	-0.0002 **
		(-1.9769)	(-2.7406)
UNEMPLOY	+ or -	-0.0977	-0.1022 ***
		(-1.5030)	(-1.7600)
CPS (=0 thru 1982; =1 otherwise)	-	-0.6831 ***	-0.5173 ***
		(-1.8714)	(-1.9700)
$\bar{R}^2$		0.998	0.998

† Sources are detailed in the appendix.
** Significant at the 0.05 level.
*** Significant at the 0.10 level.

year sample, non-wage benefits climbed from 4.57 percent of pay to 16.93 percent, a rise of 11.46 points. Thus, 3.27 points of the drop in *UNIZ* (-0.285 x 11.46), nearly twelve percent of the fall in union density, 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 estimates are quite similar (see Table 1). All coefficient signs are the same and the estimated values are of similar magnitude. The *t*-statistics are all similar, with one exception. Using the full data set, it can be seen that *UNIZ* suffers as unemployment increases.[9]

Because the two estimated equations are so alike, the fact that *FRINGE* is nonstationary 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.

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 twelve to seventeen percent of the decline in union density can be attributed to the increased presence of benefits. These findings support the substitution hypothesis that fringe benefits displace union services. 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.

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					

† Sources are detailed in the appendix.
* State with a Right-to-Work law.

BENEFITS AND UNIONIZATION: STATEWIDE ANALYSIS

Overview of the States

Union membership varies from state to state. Where union density is high and unions are strong, organized labor may obtain high levels of compensation, both wages and fringe benefits. In states where organized labor has not been strong, the structure of compensation might have evolved differently. 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 for all data sources). 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. Most fringe benefits are provided voluntarily but others are mandated (like Social Security), so there is a narrower range for *FRINGE*. Values for *FRINGE* vary 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.[10]

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.[11] 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:

$$(2) \quad UNIZ_{it} = \gamma_0 + \gamma_1 UNIZ_{it\text{-}1} + \gamma_2 FRINGE_{it} + \gamma_3 RAISE_{it} + \gamma_4 MFG\ SHARE_{it} + \gamma_5 \Delta UNFAIR_{it} + \gamma_6 RTW_{it} + \gamma_7 UNEMPLOY_{it} + \eta_{it},$$

where *FRINGE* and *RAISE* are defined as before and the other explanatory variables are

$UNIZ_{t\text{-}1}$	union density (lagged),
$MFG\ SHARE_{it}$	state i's share of U.S. manufacturing employment,
$\Delta UNFAIR_{it}$	change in the total number of unfair labor practice allegations filed in the state,
RTW_{it}	0-1 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, to reflect the industrial relations climate we used the number of unfair labor practice allegations filed by firms. These data are not disaggregated by state,

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

† Sources are detailed in the appendix.
* State with a Right-to-Work law.

but we can use the total number of unfair labor practice allegations filed by all parties. As in the national analysis, we have no prior expectation of the sign for the coefficient.

The socio-political environment for organized labor differs across states. Twenty-one states had Right-to-Work laws throughout the 1983-1996 period. 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. Using a standard *F*-test [Balestra 1996, p. 37], we found it appropriate to vary only the *FRINGE* coefficients.[12] Thus, the equation we estimate is

$$(2') \quad UNIZ_{it} = \gamma_0 + \gamma_1 UNIZ_{it\text{-}1} + \gamma_{2i} FRINGE_{it} + \gamma_3\, RAISE_{it} + \gamma_4 MFG\ SHARE_{it}$$

$$+ \gamma_5 \Delta UNFAIR_{it} + \gamma_6 RTW_{it} + \gamma_7 UNEMPLOY_{it} + \eta_{it}.$$

The pooled model represented in equation (2´) 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 heteroskedasticity. Consequently, we estimated the equation using generalized least squares (GLS).

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 11.3 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 9.3 points.

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-five states and the District of Columbia 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 16.3 points and 11.6 points, respectively. At the other extreme is Alabama, where the influence is a mere 1.7 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: three 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, six 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, seven 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 un-taxed fringe benefits. When *FRINGE* is increased, perhaps workers are pacified, thereby weakening demand for union membership.

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 — Idaho, Iowa, Kansas, Nebraska, and North Dakota — have RTW laws.

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. 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 twelve percent and seventeen percent 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 -1.629 and the state's *FRINGE* values increased by 1.683 points. Consequently, *FRINGE* accounts for approximately 2.7 points (-1.629 x 1.683) of the 11.8 point decrease in union density, more than one-fifth of the drop. As seen in the table, *FRINGE* accounted for anywhere from 4.8 percent of the reduction in union density (Pennsylvania) to more than half of the decline (Colorado, District of Columbia, Maine, Nevada, New Mexico, South Carolina, and Vermont).

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 -	16.752**	Ala.	47	-0.172*	Ida.	46	-0.173
		(6.135)			(-1.911)			(-1.531)
UNIZ	+	0.597**	Alk.	13	-0.802**	Ill.	7	-0.902**
(lagged)		(20.740)			(-4.461)			(-6.361)
RAISE	+ or -	0.006	Arz.	30	-0.502**	Ind.	34	-0.424**
		(0.361)			(-4.822)			(-3.956)
MFG SHARE	+	1.130**	Ark.	43	-0.252**	Iowa	51	-0.009
		(3.589)			(-2.751)			(-0.093)
Δ*UNFAIR*	+ or -	0.001	Cal.	1	-1.629**	Kan.	49	-0.117
		(1.300)			(-6.893)			(-1.316)
RTW	-	-9.366**	Col.	4	-0.962**	Ky.	28	-0.511**
		(-3.386)			(-6.540)			(-4.270)
UNEMPLOY	+ or -	0.317**	Conn.	9	-0.860**	Lou.	44	-0.215**
		(5.282)			(-6.188)			(-2.312)
			Del.	14	-0.771**	Me.	31	-0.482**
					(-4.110)			(-4.026)
			D.C.	6	-0.913**	Md.	26	-0.571**
					(-4.614)			(-4.493)
			Fla.	23	-0.594**	Mass.	3	-0.986**
					(-5.324)			(-7.230)
			Geo.	29	-0.510**	Mich.	25	-0.577**
					(-4.656)			(-4.984)
			Haw.	36	-0.366**	Minn.	11	-0.848**
					(-2.173)			(-5.739)

TABLE 4 (Cont.)
Determinants of Union Density by State, 1983-1996†

	FRINGE coefficients (t-statistics), by state Hypothesis: -							
	Rank			**Rank**			**Rank**	
Miss.	39	-0.338** (-3.435)	N.C.	17	-0.730** (-5.757)	Tenn.	40	-0.315** (-3.372)
Mo.	19	-0.692** (-5.292)	N.D.	50	-0.094 (-0.869)	Tex.	18	-0.703** (-5.443)
Mont.	33	-0.447** (-3.474)	Ohio	21	-0.679** (-5.614)	Utah	35	-0.413** (-4.001)
Neb.	48	-0.161 (-1.536)	Okla.	20	-0.684** (-5.337)	Ver.	15	-0.757** (-5.561)
Nev.	38	-0.345** (-3.152)	Ore.	22	-0.630** (-4.743)	Vir.	37	-0.351** (-3.431)
N.H.	8	-0.877** (-6.271)	Penn.	12	-0.818** (-6.098)	Wash.	32	-0.449** (-3.783)
N.J.	5	-0.948** (-6.119)	R.I.	16	-0.742** (-5.562)	W.V.	41	-0.279** (-2.595)
N.M.	10	-0.857** (-5.851)	S.C.	27	-0.530** (-5.096)	Wisc.	24	-0.588** (-4.836)
N.Y.	2	-1.157** (-6.207)	S.D.	42	-0.255** (-2.443)	Wyo.	45	-0.214* (-1.726)
$\overline{R}^2$		0.981						

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

CONCLUDING REMARKS

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

It appears that expanding fringe benefits "takes issues off the bargaining table," weakening support for unionization. As benefits such as health insurance and pensions have become more prominent, especially in heavily populated areas like California and New York, workers have been less prone to unionize. Or perhaps workers have become more reluctant to risk benefits in industrial action, thereby weakening organized labor. 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.

TABLE 5
***Impact of FRINGE on UNIZ by State, 1983-1996*†**

State	Δ*UNIZ*, 1983-1996	Δ*UNIZ*, due to *FRINGE*	Share (%)		Δ*UNIZ*, 1983-1996	Δ*UNIZ*, due to *FRINGE*	Share (%)
Alabama*	-9.4	-0.5	5.1	Montana	-19.1	-1.9	10.1
Alaska	-16.0	-3.4	21.1	Nebraska*	-8.5	0.0	0.0
Arizona*	-3.9	-1.5	37.2	Nevada*	-1.2	-0.7	59.9
Arkansas*	-6.2	-0.9	14.7	New Hamp.	-5.7	-2.1	37.3
California	-11.8	-2.7	22.9	New Jersey	-10.8	-1.5	13.7
Colorado	-3.5	-2.4	68.8	New Mexico	-3.0	-3.1	105.7
Connecticut	-17.5	-2.7	15.4	New York	-11.4	-2.5	21.6
Delaware	-11.1	-3.0	26.7	N. Carolina*	-4.4	-2.0	46.5
D.C.	-2.7	-3.0	108.3	N. Dakota*	-9.6	0.0	0.0
Florida*	-3.3	-1.4	42.4	Ohio	-9.6	-1.7	17.3
Georgia*	-8.2	-1.6	19.1	Oklahoma	-8.8	-4.1	46.0
Hawaii	-16.8	-1.1	6.4	Oregon	-9.0	-1.1	11.7
Idaho*	-7.8	0.0	0.0	Pennsylvania	-19.4	-0.9	4.8
Illinois	-8.8	-1.5	17.2	Rhode Is.	-5.9	-1.6	27.4
Indiana	-19.1	-1.2	6.5	S. Carolina*	-1.2	-1.3	105.3
Iowa*	-18.8	0.0	0.0	S. Dakota*	-9.3	-0.9	9.6
Kansas*	-5.5	0.0	0.0	Tennessee*	-6.6	-1.2	18.5
Kentucky	-14.8	-1.6	10.5	Texas*	-7.6	-2.4	32.2
Louisiana*	-5.9	-0.7	12.6	Utah*	-9.5	-1.6	17.3
Maine	-1.2	-2.0	162.6	Vermont	-5.3	-2.8	53.4
Maryland	-7.2	-1.5	20.9	Virginia*	-9.4	-1.1	12.0
Mass.	-16.5	-1.6	10.0	Washington	-6.5	-1.7	26.4
Michigan	-11.3	-1.6	14.3	W. Virginia	-12.1	-1.3	11.1
Minnesota	-5.3	-1.9	36.0	Wisconsin	-12.1	-0.6	5.3
Mississippi*	-8.2	-0.9	11.3	Wyoming*	-2.6	-1.0	39.2
Missouri	-10.4	-1.3	12.3				

† Sources are detailed in the appendix.
* State with a Right-to-Work law.

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 forestall unionization. Perhaps future research will investigate in more detail whether firms indeed have altered the wage-benefits mix strategically.

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.

APPENDIX — *Continued*

Non-wage Benefits. The U.S. Bureau of Economic Analysis (U.S. BEA) reports information for "Total Compensation" in the private sector, which is decomposed as "Wage Only" and "Supplements to Wages" (fringe benefits). To compute *FRINGE*, we simply divided "Supplements to Wages" by "Total Compensation" [U.S. BEA, 1998, pp. 163-66]. *RAISE* is simply the percentage change in the "Wage Only" component of the "Total Compensation" series, deflated to 1992 dollars using the GDP deflator [U.S. BEA, 1998, pp. 159-62].

Other Explanatory Variables. *DEMOCRAT* is the percentage of Democratic members of Congress. The number of Congressional Democrats is reported in Famighetti [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].

Data on unfair labor practice allegations come from the annual reports of the National Labor Relations Board [1948-1997]. 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 *Annual 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 were published for 1988 or 1993. We generated values for those years by interpolating.

Other Explanatory Variables. *RAISE* is calculated from the same data series as *FRINGE* using the wage-only components. 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 Census Bureau in Table 1 of its annual *Surveys* [U.S. BOC, 1986, 1991, 1996].

Hirsch and Macpherson [1994-1998] list the states with Right-to-Work Laws.

UNFAIR represents the total of all unfair labor practice allegations filed by all parties with the NLRB. The data come from NLRB annual reports.

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

APPENDIX — *Continued*

Data Sources

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NOTES

This research was funded partially by a Campbell R. McConnell Fellowship. We are grateful to Barry T. Hirsch and Leo Troy for advice and information concerning data. We thank two anonymous referees for their comments. Thanks also to John Anderson, Craig MacPhee, James McClure, Meghan Millea, David Rosenbaum, and Cary Thorp for helpful comments and suggestions. An earlier version of this study was presented at IZA in Bonn. Econometric results, including those of all diagnostic testing procedures, are available from Wayne Edwards at the University of Alaska Anchorage, Economics Dept. RH307F, 3211 Providence Drive, Anchorage, AK 99508.

1. For discussion of findings and citations, see the symposium "The Future of Private Sector Unions in the United States," published in the *Journal of Labor Research*; particularly, see Fiorito [2001], Kaufman [2001], Kleiner [2001], Lipset and Katchanovski [2001], Potter [2001], and Troy [2001].
2. According to U.S. Department of Labor classifications, mandated benefits consist of Social Security, state and federal unemployment insurance, and workers' compensation. Voluntary benefits consist of paid leave, health insurance and other insurance, retirement and savings, and supplemental pay (meaning all other types of non-wage compensation).
3. Neumann and Rissman's "hypothesis" formalizes an earlier observation by Reder [1951] that some functions performed by labor unions had been taken over by the state. Others have questioned whether there is empirical evidence to support the substitution hypothesis, for example, Chaison and Rose [1991] and Stepina and Fiorito [1986].
4. A similar point has been made by Jacoby [1994].
5. Even at group rates, it can cost a worker several hundred dollars per month to buy health insurance [Pauly, 1997, p. 1]. So in the event of a strike, it is likely that many workers could not purchase sufficient insurance to replace lost coverage. Their worries may be compounded because firms can hire replacement workers for strikers, threaten to relocate facilities if a strike continues, or unilaterally implement conditions of employment, all actions which have become more common over time. Of course, if employers were to try to cut existing benefits, then workers may be willing to strike to prevent such a roll back.
6. 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.
7. Farber and Western [2001] argued that the CPS measure of union membership is biased and they suggested a correction factor for the bias. In our sample, CPS figures are used for only a small portion of the UNIZ series (1983-1997). Nevertheless, we also estimated a union density equation

using the Farber and Western correction factor on the CPS portion of the UNIZ series. None of our findings was affected.

8. For details of the stationarity testing procedure, see Edwards [2000].
9. 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.
10. For details of the testing procedure, see Daniel and Terrell [1989, pp. 697-99].
11. 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.
12. The *F*-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 ten different specifications against the null that only the *FRINGE* parameter estimates vary by state. The ten 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 Δ*UNFAIR* parameter estimates by state, and then by region;
    7 and 8. Vary the *RAISE* parameter estimates by state, and then by region;
    9 and 10. 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. So we concluded that it was appropriate to vary only the *FRINGE* coefficient.

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