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A Submerging Labor Market Institution?
Unions and the Nonwage Aspects of Work

Thomas C. Buchmueller, John E. DiNardo, and Robert G. Valletta

Our understanding of emerging labor market institutions would be incomplete without an understanding of traditional institutions, including apparently diminishing forms such as the American labor union. Our focus in this paper—the effect of labor unions on a variety of nonwage aspects of work—is a small, yet important, aspect of the recent history of American unionism.

The importance of nonwage aspects of jobs to union goals dates back to the origins of the modern labor movement in the United States and other countries. For example, in late nineteenth-century Britain, according to Sidney and Beatrice Webb, "the prospect of securing support in sickness or unemployment [was] a greater inducement [for young men] to join the union . . . than the less obvious advantages to be gained by the trade combination" (Webb and Webb 1897, 158). Similarly, in the United States, to take a prominent example, the American Federation of Labor's resolution that "eight hours shall constitute a legal day's labor from and after May 1st, 1886" was one impetus for the Haymarket Rebellion—an event which is still commemorated in much of the world as International Worker's Day.

Although unions have demonstrated a historical commitment to non-

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wage aspects of jobs, union goals and impacts may have changed as union density and influence have declined. Using data from a variety of databases, we investigate the following questions: How do the nonwage aspects of union jobs differ from those of nonunion jobs? Have these differences changed during the past several decades? We first document and describe differences in hours worked in union and nonunion jobs. We then provide an updated assessment of union impacts on the provision of various fringe benefits, addressed for an earlier period in various studies (most comprehensively by Freeman and Medoff 1984).

7.1 Background

If only a single bit of context is to be highlighted about the role of today's unions, surely it is the decline in union representation. The fraction of workers currently represented by a labor union is at roughly the same level as it was *before* the passage of the National Labor Relations (Wagner) Act in 1935. The Wagner Act declared that it was U.S. policy "to eliminate the causes of certain substantial obstructions to the free flow of commerce and to mitigate and eliminate these obstructions when they have occurred by encouraging the practice and procedure of collective bargaining" (see the National Labor Relations Board website at http://www.nlrb.gov). After a rapid spurt of growth between 1935 and 1945, unionization rates remained roughly level until the late 1950s. Since the 1960s, unionization rates have fallen almost continuously.[1]

Following the publication of Lewis (1963), a comprehensive study of union/nonunion wage gaps, years of subsequent research provided ample quantitative evidence of wage premia associated with unionization. As to benefits, a simple summary of the findings in the review by Freeman and Medoff (1984) is that ceteris paribus, union jobs provide more generous benefits. Current union density, however, is about half as large as it was at that time, which raises the possibility that union impacts on benefit provision may have changed as well.

Some indirect evidence regarding possibly changing union impacts on benefits comes from recent work on the causes of rising wage inequality (Bound and Johnson 1992; Katz and Murphy 1992; Autor and Katz 2000; Levy and Murnane 1992; Card, Lemieux, and Riddell 2003). Unionization has traditionally been an equalizing influence, reducing both inequality *between* groups with different demographic characteristics and inequality *within* demographically homogeneous groups. Unions have typically reduced between inequality by raising the relative pay of groups with low

1. Although it might be tempting to predict continued declines, the certainty of such a forecast should be tempered by the striking similarity of the recent time series path of unionization to that in 1935. In particular, despite years of decline prior to the Wagner Act, during the subsequent ten years the union membership rate doubled.

earnings; for example, unionization rates typically are higher among the less educated. Given this equalizing effect of unions, declining unionization might be expected to increase between inequality, and the evidence bears this out.[2] Because unions also tend to compress pay schedules within groups, a decline in the power of unions to affect wages also should lead to increased within-group inequality. Indeed, deunionization has coincided with an increase in *within* inequality not only because fewer workers are covered but also because the distribution of wages in the union sector increasingly resembles that of the nonunion sector (DiNardo and Lemieux 1997). That is, the ability of unions to minimize differences in wages between demographically similar workers has declined. The changing impact of unions on the wage distribution suggests the possibility of changing union impacts on nonwage aspects of jobs as well.

7.2 Union Effects on Hours of Work

Comparatively little has been written about union/nonunion differences in hours worked for employed individuals. Neither of Lewis's (1963, 1986) comprehensive surveys address the issue at any length; the same is true for more broadly focused summaries of research on labor unions (Freeman and Medoff 1984; Hirsch and Addison 1986), although Freeman and Medoff investigate the cyclical sensitivity of union employment (see Raisian [1979] for a similar focus). Killingsworth's (1983) extensive survey makes no mention of labor supplied in the presence of unionization. Earle and Pencavel (1990) examine the question with the 1978 Current Population Survey (CPS) and time series data. They find conflicting evidence—negative union impacts on full-time hours worked in time series data but positive union effects on total hours worked for some groups (white male laborers and women) in the cross section. In contrast to their cross-section findings, Trejo (1993) found negative union impacts on hours worked using 1985 CPS data. The most comprehensive study of the question across Organization for Economic Cooperation and Development (OECD) countries is Blanchflower (1996) who found, inter alia, that unions reduce total hours of work in the full set of OECD countries analyzed.

The most severe obstacle to such an analysis is the comparatively poor quality of information on hours worked. Moreover, a fuller analysis would incorporate related aspects of time use, such as commuting time. Unfortunately, we are unaware of any data set that allows for a comparison across time in hours worked of the same quality (and quantity) as, say, the wage data in the CPS.

2. See DiNardo, Fortin, and Lemieux (1996); Card (1992); and Freeman (1993) for the United States and Gosling and Machin (1995) for the United Kingdom.

7.2.1 The Economics of Hours Reductions

Before turning to our empirical analysis, it is helpful to motivate why one might see differences in hours worked between the union and nonunion sectors. The textbook monopoly union model generally is silent regarding the effect of unionization on hours worked, treating hours per worker as fixed, and focusing on the number of workers employed. In this type of model the main consequence of unionization is lower employment.

By contrast, in efficient contract models (Brown and Ashenfelter 1986; MaCurdy and Pencavel 1986) the relationship between the union and the firm is oriented around rent sharing. In these models, the union maximizes the joint revenue of the firm and the workers—the wage is merely an artifact of the division of this revenue between the two groups. In appendix A we present a version of the efficient contract model to motivate the possible economic logic of a difference in hours worked between union and nonunion workers. Our intent is not to test an obviously simplistic model of union hours determination.[3] Instead, we use the model to highlight one possible trade-off that employers and union workers face in collective bargaining—the trade-off between a larger number of members splitting smaller shares of the "economic pie" versus fewer members with larger shares. This trade-off arises because increases in labor supplied to firms can come either through an increase in the number of workers employed or an increase in hours worked per employee. As derived in the model, compared to a nonunion setting, an optimal union contract typically results in (1) fewer hours per worker and (2) union members being constrained to work fewer hours than they would like to at the negotiated wage rate (despite being better off as union members than as nonmembers).

In the model, the magnitude of these effects will depend on the extent to which the wage-setting process differs between the union and nonunion settings. The more union wage setting resembles wage setting in the "competitive" sector (the smaller the difference between union member utility and the nonunion alternative), the smaller is the hours differential and the less "constrained" is the union hours outcome. Thus, given the decline in the level of unionization in the United States and the implied decline in union power, it may be reasonable to expect the size of the union effect to decline over time as well.

7.2.2 The Data

Our analysis of the impact of unions on hours worked relies on two data sources. The Current Population Survey (CPS) is the federal government's

3. Booth (1995), among others, has argued that tests of efficient contract models, such as those employed in Brown and Ashenfelter (1986), are flawed because it is not possible to empirically distinguish between the monopoly union and efficient contracting views using data on employment and wages. She notes that in the absence of such a test, a "pragmatic approach" would rely on studying the texts of collective bargaining agreements instead.

monthly survey of about 60,000 households, which provides information on labor force activities, earnings, and related variables for a rotating cross section designed to be representative of the U.S. population. We use data from the Annual Demographic Supplement (March) to the monthly CPS, which provides information on income and work hours during the complete calendar year preceding the survey. We also use data from the Panel Study of Income Dynamics (PSID), a panel study of 5,000 families initiated in 1968. Each year, the original sample of families and their "splitoffs" (for example, children who leave to form their own households) are reinterviewed; sample attrition has been relatively limited over time, and new families are added when necessary to keep the sample approximately representative of the nation as a whole.[4]

Relative to the March CPS data, the PSID has some unique advantages and disadvantages in regard to estimating the effect of union status on hours worked.

The main advantage of the PSID is its collection of information that provides a complete yearly calendar for each individual; this calendar indicates the total amount of weeks spent at work, on vacation, unemployed, on strike, or away from work due to personal illness or the illness of a family member. Despite some ambiguity in the questions, the response accuracy is enhanced through the survey requirement that the interviewer readminister the questions until the responses add up to fifty-two. This is important given the well-known problems of measurement error in data on hours and weeks worked.[5]

In contrast, the March CPS includes "paid" vacation as part of weeks "worked." This complicates use of the CPS in several ways. Consider a person with two weeks' paid vacation (and no unpaid vacation time). The estimate of annual hours is fifty-two times average hours per week, instead of fifty times average hours per week. Arguably, use of the March CPS leads to systematic underreporting of the hours differential if vacation time is more generous for union workers (as we document in the following, this appears to be the case).

A main disadvantage of the PSID is inconsistency across respondents in their answers to the question regarding "average hours per week"—some respondents appear to base their answer on a "typical" work week rather than a true average. Other problems in the PSID include the comparatively small size of the sample, the lack of information on nonhousehold heads, and the lack of statistical dependence in successive years of data for each panel member.[6]

In the following analyses, we use PSID data for the years 1972–1992 and

4. To enhance our sample's representation of the general U.S. population, we do not include the low-income oversample in the analyses below.

5. See Bound and Krueger (1991) for one illustration.

6. For example, we restrict our analyses to men because most of the detailed information on work time is not available for the "wife" of the household.

CPS data for the years 1983–1997. Although our preference was for an exact match of sample years for both sources, our choice was dictated by data availability constraints in the PSID.[7]

7.2.3 Methodology

An important aspect of the hours data that we analyze in this section is that it is either multimodal (often at 0 and another value) or otherwise distributed in a way that precludes the use of straightforward regression techniques to characterize the conditional distribution. Olson (1998) faced a similar problem in his study of the effects of health insurance on hours worked and found that conventional regression estimates tend to produce misleading results.

Following Olson (1998) and Buchmueller, DiNardo, and Valletta (2002), we apply the reweighting techniques from DiNardo, Fortin, and Lemieux (1996) to adjust the union and nonunion distribution for differences in worker characteristics. That is, we adjust the average hours of a typical worker relative to some base set of characteristics, specifically the characteristics of a typical worker in 1992. This technique, which relies on the estimation of "conditioning weights," allows us to estimate the conditional impact of unionism on hours worked without imposing any misleading distributional assumptions on the data; appendix B describes this method in detail. Stated simply, given our data structure, it is easier to estimate the relationship between union status and related variables and use this relationship to adjust the hours data rather than applying the flexible functional forms needed to directly estimate the conditional impact of union status on hours worked. In the case where our estimated conditioning weights are based on a partition of discrete groups (such as union and nonunion workers), our procedure amounts to calculating simple differences in means for the various groups and then appropriately weighting them to form an overall average. This can be viewed as an application of propensity score weighting as in Rosenbaum and Rubin (1983). We apply this technique to estimate union/nonunion differences for each of the hours measured discussed in this section and listed in tables 7.1–7.5. As discussed in more detail in the appendix, we use a complete set of race dummies, school dummies, marital status dummies, standard metropolitan statistical area (SMSA) dummies, and a cubic in age as explanatory variables for the PSID data. For the CPS data we use five education categories, five age categories, an SMSA dummy, three regional dummies, marital status, and three race categories.

7. We assume that the design changes in the CPS questionnaire after 1993 are ignorable in the context of our analyses.

7.2.4 Results

Hours

Our estimates of the union impact on "average weekly hours" (PSID) and "usual weekly hours" (CPS) are presented in tables 7.1 and 7.2. The first column of table 7.1 lists the mean hours for nonunion workers in a given year in the PSID, *calculated as if they had the characteristics of union workers in the 1992 sample.* The second column lists the mean for union workers in a given year, also calculated as if they had the characteristics of union workers in 1992; our estimate for 1992 union workers, therefore, is the unadjusted mean from the data (using the appropriate population weights). The third column in the table lists the differences between the first two columns; under our estimation procedure, this represents the effect on nonunion status on hours worked, conditional on a set of covariates used to estimate the conditioning weights (we include controls for a standard set of individual characteristics; see the end of appendix B for a complete list). The format for the CPS results is similar. Figures 7.1 and 7.2 display the PSID and CPS results (first two columns) graphically.

Considering the PSID estimates first, in almost all years union workers work fewer hours per week than do nonunion workers; during the years 1972–1984, this difference is about three hours per week. After 1984 the differential begins falling, and by 1992 the situation is reversed, with union workers estimated to work about one hour more per week than nonunion workers. However, the standard errors listed in these tables indicate that the union/nonunion differences generally are not statistically significant after 1984.[8]

In table 7.2, we present results based on the CPS data, again using the 1992 union worker as our basis of comparison. The estimates for nonunion workers (appropriately weighted) hover around 42.2 hours per week, and the estimates for union workers are approximately 0.5 hours per week lower. By comparison with the PSID results, the CPS-based estimates of usual weekly hours for nonunion workers are slightly higher in general, although the differential between the two sectors is relatively constant.[9]

The differences between the PSID and CPS estimates could arise from a number of sources. Part of the difference is surely differences in the wording of the questionnaires as well as sampling error due to the comparatively small size of the PSID. Another salient difference is that with both data sets

8. The standard errors listed probably understate the true sampling errors because they were not adjusted for the uncertainty generated by estimation of the conditioning weights nor for the nonindependence of multiple observations per individual in the PSID data.

9. Although problems of sample selectivity and limitations of scope preclude an extensive analysis here, the relationship between union and nonunion workers is reversed when women are included in the sample (Earle and Pencavel 1990).

Table 7.1 **Mean average hours per week for men with characteristics of 1992 union workers (PSID data)**

Year	Nonunion	Union	Difference
1972	45.16	41.95	3.20
	(0.29)	(0.43)	(0.52)
1973	45.43	42.61	2.81
	(0.27)	(0.41)	(0.49)
1974	45.09	41.82	3.26
	(0.26)	(0.34)	(0.43)
1975	44.89	38.82	6.07
	(0.29)	(0.43)	(0.52)
1976	44.67	40.70	3.97
	(0.28)	(0.38)	(0.47)
1977	44.57	39.60	4.98
	(0.27)	(0.44)	(0.51)
1978	44.39	41.54	2.85
	(0.26)	(0.31)	(0.40)
1979	44.17	41.20	2.97
	(0.26)	(0.34)	(0.43)
1980	44.02	40.21	3.81
	(0.26)	(0.37)	(0.45)
1981	44.00	41.27	2.73
	(0.25)	(0.37)	(0.45)
1982	44.31	40.52	3.79
	(0.23)	(0.41)	(0.47)
1983	44.04	40.32	3.72
	(0.23)	(0.39)	(0.46)
1984	43.71	40.73	2.98
	(0.23)	(0.41)	(0.47)
1985	42.99	41.39	1.60
	(0.27)	(0.37)	(0.46)
1986	42.54	42.65	–0.10
	(0.27)	(0.25)	(0.37)
1987	42.57	42.20	0.38
	(0.27)	(0.33)	(0.43)
1988	42.71	41.75	0.96
	(0.26)	(0.36)	(0.45)
1989	42.56	41.44	1.11
	(0.25)	(0.43)	(0.50)
1990	42.31	42.92	–0.61
	(0.25)	(0.35)	(0.43)
1991	41.98	42.68	–0.70
	(0.25)	(0.37)	(0.45)
1992	41.86	42.92	–1.06
	(0.25)	(0.41)	(0.48)

Notes: Standard errors in parentheses. For description of counterfactual weights see text.

Table 7.2 **Mean average hours per week for men with characteristics of 1997 union workers (CPS data)**

Year	Nonunion	Union	Difference
1983	41.33	40.50	0.83
	(0.04)	(0.04)	(0.06)
1984	41.67	41.03	0.64
	(0.03)	(0.05)	(0.06)
1985	41.82	41.15	0.66
	(0.03)	(0.05)	(0.06)
1986	41.81	41.21	0.60
	(0.03)	(0.05)	(0.06)
1987	42.02	41.37	0.65
	(0.03)	(0.05)	(0.06)
1988	42.05	41.46	0.58
	(0.04)	(0.05)	(0.07)
1989	42.21	41.57	0.63
	(0.04)	(0.06)	(0.07)
1990	42.19	41.47	0.71
	(0.03)	(0.05)	(0.06)
1991	41.88	41.28	0.60
	(0.04)	(0.06)	(0.07)
1992	41.84	41.41	0.43
	(0.04)	(0.06)	(0.07)
1993	41.94	41.61	0.34
	(0.04)	(0.06)	(0.07)
1994	42.15	41.56	0.59
	(0.04)	(0.07)	(0.08)
1995	42.20	41.89	0.32
	(0.04)	(0.07)	(0.08)
1996	42.33	41.76	0.57
	(0.04)	(0.07)	(0.08)
1997	42.30	41.84	0.45
	(0.04)	(0.07)	(0.08)

Notes: Standard errors in parentheses. For description of counterfactual weights see text.

the variability in weekly hours is substantially higher in the nonunion sector. To give the reader some sense of the difference, we display the standard deviation of hours worked from our CPS data in figure 7.3. The variability across workers is fairly constant for the nonunion sector, although it appears to be increasing in the union sector—the relative increase in union sector variance is consistent with other evidence, suggesting that union wage setting is growing more like nonunion wage setting (DiNardo and Lemieux 1997).

Figure 7.4 uses the propensity score (the predicted probability from the logit that we use for our weights) to take a first pass at investigating heterogeneity (e.g., variability) in the union effect across different groups of

Fig. 7.1 Average hours per week for men with characteristics of 1992 union workers PSID data

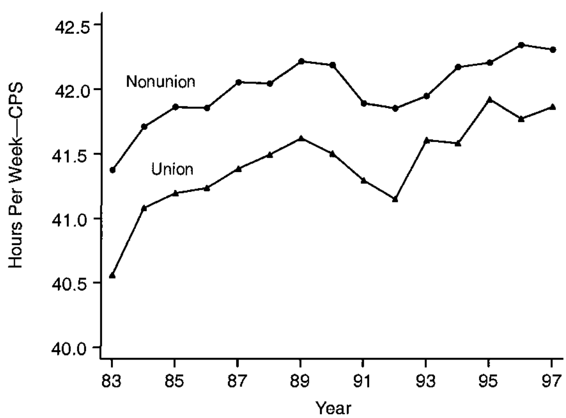

Fig. 7.2 Average hours per week for men with characteristics of 1992 union workers CPS data

male workers in the 1992 CPS cross section. Workers with characteristics "more like" union members have higher values of this propensity score.[10]

10. The actual regression lines are the result of a natural cubic spline with knots at the 33rd and 66th percentile of the observed propensity score. Experimentation with nonparametric estimates and regressograms confirmed the adequacy of the natural cubic spline to fit the data.

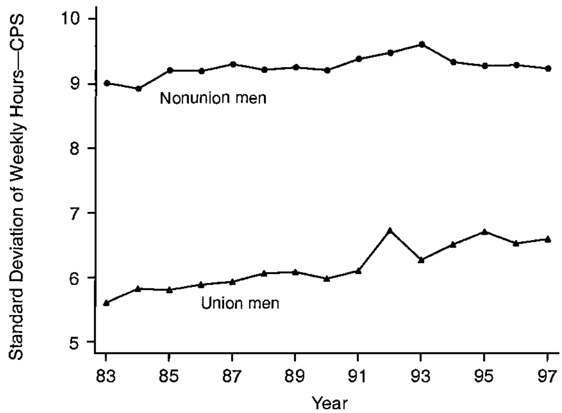

Fig. 7.3 Standard deviation of hours per week men with characteristics of 1992 union workers CPS data

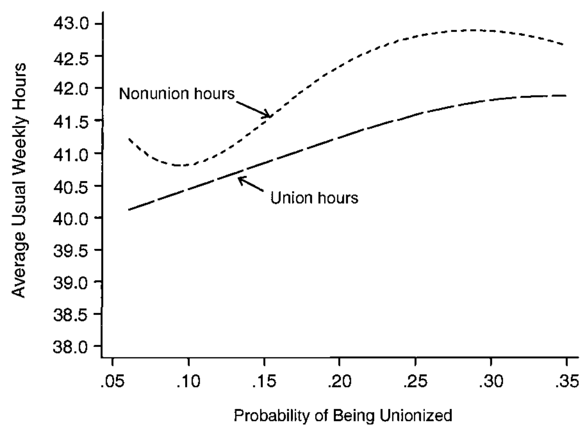

Fig. 7.4 Average weekly hours and union propensity score

Note: Plots are generated with a natural cubic spline with knot points at the 33rd and 66th percentiles of the propensity score.

We restrict our attention in the figure to the domain between the 5th and 95th percentile of propensity scores.[11]

The analysis suggests some variation in the size of the union effect, with the effect being greater for workers who look more like the "typical union worker." On the other hand, the amount of variation explained by the propensity score is tiny. Depending on the precise specification of the pro-

11. Outside this range, the union hours effect is actually reversed; however, our estimates in this range are very unreliable due to the paucity of observations.

pensity score, the total variation in hours explained is typically less than 8 percent. This is not surprising given the fact that the literature on labor supply (see, for example, Killingsworth 1983), which typically conditions on the observed wage in addition to the covariates that we use, has only managed to explain a tiny fraction of the variation in hours worked.

In sum, both the PSID and CPS results suggest that hours are lower in the union sector, although there are some differences in the levels, differences, and trends between the two sets of numbers.

Time Spent Due to Illness (Own or Other)

Based on our PSID data, table 7.3 reports the average number of weeks spent not working due to illness, with the nonunion and union values computed as before (i.e., with the characteristics of union workers in 1992 used as the adjustment base). In general, nonunion workers spend slightly less time away from work due to illness—approximately 1 week per year compared to 1 to 1.4 weeks for union members. This differential is somewhat consistent with our evidence regarding "paid" sick leave, presented in the following (see table 7.10 later in the chapter). While the magnitude of the difference varies from year to year, there is no discernible trend.

Time Spent on Vacation

Recall from our earlier discussion that the treatment of paid vacation time as time spent at work makes the use of the CPS problematic for an analysis of union/nonunion differences: Indeed, as we show later in the chapter (see Table 7.10), there appears to be a difference in the incidence of "paid" vacation time between union and nonunion workers. We therefore use only the PSID—which records all vacation time whether paid or not—for this analysis. These results are listed in table 7.4 (again estimated using our conditioning technique).

Given the relatively small samples, it is noteworthy that the difference between the two groups of workers is fairly constant—union workers typically have one week more vacation time than nonunion workers and, perhaps surprisingly, the level of vacation time in the two groups rose during the years 1972–1992 (albeit slowly).

Hours Constraints

In table 7.5, we list the percentage of workers who report that they "would like to work more" and those who "would like to work less" (the worker is also free to choose neither more nor less.) During the sample period, two trends are immediately evident for both sets of workers—the percentage of workers who say they would like to work more rises, and the percentage of workers who would like to work less falls. Consistent with the previous tables and with our simple model of union behavior, during the

Table 7.3 **Mean weeks not working due to illness for men with characteristics of 1992 union workers (PSID data)**

Year	Nonunion	Union	Difference
1972	0.92	1.23	–0.31
	(0.06)	(0.11)	(0.12)
1973	0.96	1.06	–0.10
	(0.06)	(0.10)	(0.11)
1974	0.83	1.18	–0.35
	(0.05)	(0.12)	(0.13)
1975	1.12	1.23	–0.11
	(0.07)	(0.08)	(0.10)
1976	1.07	1.36	–0.29
	(0.06)	(0.09)	(0.10)
1977	1.06	1.37	–0.31
	(0.05)	(0.09)	(0.10)
1978	1.15	1.27	–0.12
	(0.06)	(0.09)	(0.10)
1979	1.22	1.24	–0.02
	(0.06)	(0.08)	(0.10)
1980	1.12	1.32	–0.20
	(0.05)	(0.09)	(0.10)
1981	1.15	1.32	–0.17
	(0.05)	(0.09)	(0.10)
1982	1.13	1.26	–0.14
	(0.05)	(0.09)	(0.11)
1983	1.15	1.43	–0.29
	(0.05)	(0.11)	(0.12)
1984	1.17	1.56	–0.39
	(0.05)	(0.13)	(0.14)
1985	1.26	1.22	0.04
	(0.07)	(0.09)	(0.11)
1986	1.09	1.34	–0.25
	(0.06)	(0.12)	(0.13)
1987	1.08	1.33	–0.25
	(0.06)	(0.10)	(0.11)
1988	1.05	1.30	–0.25
	(0.05)	(0.11)	(0.12)
1989	1.02	1.27	–0.25
	(0.05)	(0.10)	(0.12)
1990	1.05	1.18	–0.13
	(0.05)	(0.12)	(0.13)
1991	1.10	1.43	–0.34
	(0.05)	(0.15)	(0.16)
1992	1.15	1.21	–0.06
	(0.05)	(0.14)	(0.15)

Notes: Standard errors in parentheses. For description of counterfactual weights see text.

Table 7.4 **Mean vacation time in weeks for men with characteristics of 1992 union workers (PSID data)**

Year	Nonunion	Union	Difference
1972	1.71	2.17	–0.46
	(0.06)	(0.09)	(0.11)
1973	1.71	2.18	–0.47
	(0.06)	(0.09)	(0.11)
1974	1.77	2.16	–0.39
	(0.06)	(0.09)	(0.11)
1975	1.93	2.37	–0.44
	(0.06)	(0.09)	(0.11)
1976	1.94	2.42	–0.48
	(0.05)	(0.09)	(0.11)
1977	1.99	2.49	–0.50
	(0.05)	(0.10)	(0.11)
1978	1.97	2.55	–0.58
	(0.05)	(0.10)	(0.11)
1979	1.99	2.44	–0.46
	(0.05)	(0.10)	(0.11)
1980	2.03	2.53	–0.50
	(0.05)	(0.10)	(0.11)
1981	2.09	2.46	–0.38
	(0.05)	(0.10)	(0.11)
1982	2.10	2.51	–0.41
	(0.05)	(0.11)	(0.12)
1983	2.09	2.69	–0.59
	(0.05)	(0.12)	(0.13)
1984	2.14	2.56	–0.43
	(0.05)	(0.12)	(0.13)
1985	2.34	2.88	–0.54
	(0.07)	(0.10)	(0.12)
1986	2.35	2.98	–0.63
	(0.07)	(0.12)	(0.13)
1987	2.34	2.96	–0.62
	(0.07)	(0.12)	(0.13)
1988	2.36	2.91	–0.56
	(0.06)	(0.12)	(0.14)
1989	2.40	2.93	–0.53
	(0.06)	(0.12)	(0.14)
1990	2.36	2.87	–0.51
	(0.06)	(0.12)	(0.13)
1991	2.39	2.80	–0.41
	(0.06)	(0.12)	(0.13)
1992	2.31	3.27	–0.96
	(0.05)	(0.15)	(0.15)

Notes: Standard errors in parentheses. For description of counterfactual weights see text.

Table 7.5 **Percent saying "Would like to work more" or "Would like to work less" for men with characteristics of 1992 union workers**

	Would like to work more			Would like to work less		
Year	Nonunion	Union	Difference	Nonunion	Union	Difference
1972	0.17	0.24	–0.06	0.08	0.06	0.03
	(0.01)	(0.02)	(0.02)	(0.01)	(0.01)	(0.01)
1973	0.15	0.26	–0.11	0.10	0.07	0.03
	(0.01)	(0.02)	(0.02)	(0.01)	(0.01)	(0.01)
1974	0.15	0.25	–0.10	0.08	0.07	0.01
	(0.01)	(0.02)	(0.02)	(0.01)	(0.01)	(0.01)
1975	0.17	0.34	–0.16	0.07	0.09	–0.02
	(0.01)	(0.02)	(0.02)	(0.01)	(0.01)	(0.01)
1976	0.25	0.36	–0.12	0.05	0.06	–0.01
	(0.01)	(0.02)	(0.02)	(0.01)	(0.01)	(0.01)
1977	0.24	0.37	–0.13	0.08	0.10	–0.02
	(0.01)	(0.02)	(0.02)	(0.01)	(0.01)	(0.01)
1978	0.18	0.30	–0.12	0.09	0.18	–0.08
	(0.01)	(0.02)	(0.02)	(0.01)	(0.02)	(0.02)
1979	0.20	0.34	–0.15	0.06	0.12	–0.05
	(0.01)	(0.02)	(0.02)	(0.01)	(0.01)	(0.01)
1980	0.22	0.34	–0.12	0.07	0.06	0.01
	(0.01)	(0.02)	(0.02)	(0.01)	(0.01)	(0.01)
1981	0.28	0.36	–0.08	0.03	0.06	–0.02
	(0.01)	(0.02)	(0.02)	(0.00)	(0.01)	(0.01)
1982	0.26	0.39	–0.12	0.04	0.05	–0.01
	(0.01)	(0.02)	(0.02)	(0.00)	(0.01)	(0.01)
1983	0.30	0.39	–0.08	0.05	0.03	0.02
	(0.01)	(0.02)	(0.02)	(0.00)	(0.01)	(0.01)
1984	0.27	0.33	–0.05	0.06	0.08	–0.02
	(0.01)	(0.02)	(0.02)	(0.01)	(0.01)	(0.01)
1985	0.27	0.30	–0.03	0.05	0.03	0.02
	(0.01)	(0.02)	(0.02)	(0.00)	(0.01)	(0.01)
1986	0.31	0.31	–0.00	0.03	0.04	–0.01
	(0.01)	(0.02)	(0.02)	(0.00)	(0.01)	(0.01)
1987	0.31	0.31	0.00	0.04	0.02	0.02
	(0.01)	(0.02)	(0.02)	(0.00)	(0.01)	(0.01)
1988	0.29	0.32	–0.03	0.05	0.02	0.03
	(0.01)	(0.02)	(0.02)	(0.00)	(0.01)	(0.01)
1989	0.30	0.34	–0.05	0.05	0.02	0.03
	(0.01)	(0.02)	(0.02)	(0.00)	(0.01)	(0.01)
1990	0.32	0.31	0.01	0.04	0.02	0.02
	(0.01)	(0.02)	(0.02)	(0.00)	(0.01)	(0.01)
1991	0.32	0.35	–0.02	0.04	0.03	0.01
	(0.01)	(0.02)	(0.02)	(0.00)	(0.01)	(0.01)
1992	0.31	0.34	–0.03	0.04	0.03	0.01
	(0.01)	(0.02)	(0.02)	(0.00)	(0.01)	(0.01)

Notes: Standard errors in parentheses. For description of counterfactual weights see text.

early part of our sample, union workers were more likely than nonunion workers to desire more work hours. However, over the entire sample frame the proportion constrained in this regard rose much more for nonunion workers than for union workers so that beginning in the mid-1980s, the gap between the two groups was numerically and statistically insignificant. Similarly, the final columns of the table show that union workers are slightly less likely to desire fewer work hours; however, the small percentage of workers expressing this view means that the difference between union and nonunion workers is not significantly different from zero at conventional levels.

7.3 Health Insurance, Pensions, and Other Benefits

As suggested by the quote by Sidney and Beatrice Webb (1897) in our introduction, labor unions have long played an important role in securing social insurance benefits for workers. Indeed, according to Munts (1967), many early union organizations were established for the provision of such benefits and only later became engaged in bargaining over wages. Later on, pressure from organized labor provided significant impetus for the growth of pension plans in the 1940s and 1950s (Allen and Clark 1986).

There are several reasons to expect a positive effect of union membership on health insurance, retirement plans, and related benefits.

First, in nonunion workplaces, where entry and exit are the primary adjustment by which worker preferences are expressed, compensation policies will be tailored to suit the "marginal" workers, who tend to be young and mobile and therefore likely to have a low demand for health and retirement benefits. In contrast, the political processes of unionized workplaces often result in greater weight being placed on the preferences of older, less-mobile workers, who are likely to have a greater demand for such benefits.[12]

Second, to the extent that unions play a role in administering benefit programs, they may also lower the actual cost of such programs to employers. In many cases, the union's role involves providing information to employees about the value and tax advantages of nonwage benefits, and thus influencing employee demand. This notion of unions as providing information to employees is consistent with the evidence that union workers are more likely to take up publicly funded benefits, such as workers' compensation and unemployment insurance (Hirsch, Macpherson, and Dumond 1997; Budd and McCall 1997). Also worthy of note is the possibility that in repeated bargaining, established benefits such as health insurance are difficult for the firm to eliminate—they become part of a nonnegotiable "base" as compared to a negotiable "increment."

12. See Goldstein and Pauly (1976) for a more formal treatment.

A number of studies from the 1970s and early 1980s show large union effects on the receipt and quantity of fringe benefits, particularly health insurance and pensions (Freeman 1981; Alpert 1982; Freeman and Medoff 1984; Feldman and Scheffler 1992; Allen and Clark 1986; Belman and Heywood 1990). Pierce (2001) exploited the Bureau of Labor Statistics (BLS) data used to compute the Employment Cost Index (for the years 1981–1997) and found significant union effects on paid leave, pensions, and health insurance. In this section we update this research by estimating the relationship between union membership and health insurance and pensions from the 1980s to the mid-1990s, using data from several special supplements to the CPS. Both benefits are very important from a policy perspective. There is great concern among health economists and other policy analysts over the decline in health insurance coverage over the past several decades (see, for example, Kronick and Gilmer 1999). Likewise, the aging of the baby boom cohort combined with questions concerning the long-run viability of Social Security heighten the policy importance of employer-sponsored retirement programs.

In the last part of this section, we also present union/nonunion differences in the receipt of several other nonwage benefits, using additional data sources besides the CPS supplements. Largely for reasons of data availability, we focus on the offer or receipt of the various benefits rather than the level or generosity of coverage. Because several studies indicate significantly positive union effects on the latter (Freeman 1981; Freeman and Medoff 1984; Allen and Clark 1986; Buchmueller, DiNardo, and Valletta 2002), the estimates we report here can be thought of as lower bounds of sorts.

7.3.1 Health Insurance

Table 7.6 presents estimates of union/nonunion differences in health insurance benefits for four different periods: 1983, 1988, 1993, and 1997. The data are from various supplements to the CPS. The 1983 data is the most limited, providing information only on whether the responding worker received health insurance coverage through his employer. For later years, it is possible to distinguish between whether a worker's employer offers health benefits to any workers and whether the worker is covered.[13] The third column in the table reports unadjusted union/nonunion differences for these outcomes—that is, the difference between the figures in the first two columns. The adjusted differences in the fourth column are the coefficients on an indicator variable for union membership from linear probability models that also include individual characteristics and industry

13. Conditional on firm offers, coverage depends on whether a worker is eligible for benefits and take-up conditional on eligibility. See Buchmueller, DiNardo, and Valletta (2002) for more detail on union/nonunion differences in these intermediate outcomes.

Table 7.6 **Union/nonunion differences in health insurance availability and coverage (CPS benefits supplement data)**

			Difference (union-nonunion)		
	Union	Nonunion	Unadjusted	Adjusted	Adjusted (size)
			A. 1983 (N = 15,637)		
Covered	.929	.655	.274	.211	.151
			(.009)	(.009)	(.008)
			B. 1988 (N = 15,254)		
Available	.938	.816	.122	.095	.039
			(.008)	(.009)	(.009)
Covered	.890	.668	.222	.152	.097
			(.010)	(.010)	(.010)
			C. 1993 (N = 15,179)		
Available	.946	.792	.154	.141	.078
			(.010)	(.009)	(.009)
Covered	.870	.624	.246	.194	.132
			(.012)	(.011)	(.011)
			D. 1997 (N = 8,144)		
Available	.928	.816	.112	.100	n.a.
			(.013)	(.013)	
Covered	.835	.620	.215	.175	n.a.
			(.016)	(.016)	

Notes: All estimates were obtained using the survey supplement weights. Standard errors are in parentheses. The estimates in the fourth column are the union coefficients from linear probability models that also control for education (four category dummies), age, age squared, female, whether married, female by married, race/ethnicity (dummy variables for black and hispanic), a dummy variable for metropolitan statistical area (MSA) residency, three region dummies, and eight major industry dummies. The adjusted differences in the final column are based on a specification that also includes five establishment size dummies (10–24, 25–49, 50–99, and 100–249, 250+; <10 is the omitted category; four dummies in 1983); n.a. = not available.

dummy variables.[14] The fifth column adds controls for firm size. This is our preferred specification as both health insurance coverage and union membership are positively related to firm size. The column (4) specification is reported because we would like to trace union effects over the entire period and information on firm size was not available in the 1997 data.

In each year, the percentage of union workers with health insurance is substantially higher than the corresponding figure for nonunion workers, and for both groups the percent covered fell over time. Since the decline between 1983 and 1997 was greater for union workers (10.3 percentage points versus 5.1 percentage points), the union/nonunion difference also fell. A

14. Using some of the same data we use here, Even and Macpherson (1993) show that estimates of union/nonunion differences in health insurance and pension receipt are fairly robust to alternative econometric specification. We use linear probability models for their ease of interpretation.

comparison of the estimates for 1988 and 1997 suggest that for both sectors the decline in coverage was not driven by a cutback in employer offers but, rather, in the percentage of workers in insurance-providing firms that are covered, a result that has been documented in other studies (Cooper and Schone 1997; Farber and Levy 2000).

Differences in worker and firm characteristics account for roughly half of the union/nonunion difference in health insurance coverage. When the regression controls for firm size, the adjusted difference falls from 15 percentage points in 1983 to 9.7 percentage points in 1988 before increasing to 13.2 percentage points in 1993.[15]

One interesting pattern in table 7.6 is that, for all years, the union effect on coverage is greater than the effect on offers. Unreported regressions reveal that this is due to the fact that within firms where insurance is available there are positive union effects on both the probability of being eligible for coverage and take-up among eligible workers. These two effects have opposing time trends. The adjusted union/nonunion difference in the probability a worker is eligible for coverage, conditional on employer provision, fell from 6 percentage points in 1988 to 2 percentage points in 1997 (without controlling for firm size). The comparable difference in take-up rates (conditional on being eligible for coverage) increases from 3 to 10 percentage points over this same period.

The higher take-up rate for union workers is partly explained by their lower required employee contributions and more generous benefits available in offered plans. Table 7.7 lists the evidence for this, which comes from a 1993 survey of employers conducted by the Robert Wood Johnson Foundation.[16] Adjusting for establishment characteristics (size, industry, state, years in business, and employee demographics), the mean employer premium contribution (expressed as a percentage of the premium) is between 9 and 10 percent higher in unionized establishments. Part of this difference arises from the fact that establishments with union workers are between 15 and 20 percent (depending on coverage tier—i.e., single or family) more likely to pay the full cost of coverage (results not listed).

In terms of plan benefits, indemnity and PPO plans offered to union employees have significantly lower deductibles. For preferred provider organization (PPO) plans, the mean in-network and out-of-network coinsurance rates are 2 percentage points lower for union plans. Among health maintenance organization (HMO) plans, there is no significant difference in office visit co-payments between union and nonunion plans. The likely explanation is that in the case of HMOs, provider panel size and "quality" are more

15. Because insurance provision is essentially universal among firms with 100 or more employees, the union effect on offers reported here represent the average of large effects for small and medium-sized firms and zero effects for larger ones. See Buchmueller, DiNardo, and Valletta (2002).

16. See Buchmueller, DiNardo, and Valletta (2002) for more details on this data set.

Table 7.7 Union effects on employer premium contributions and health plan cost sharing

	Mean (standard deviation)		Difference: union-nonunion (standard error)	
	Union	Nonunion	Unadjusted	Adjusted
Employer premium contribution (as a percent of premium)				
Single coverage (N = 19,450)	88.3	81.8	6.5	9.1
Family coverage (N = 19,102)	76.3	64.9	11.4	10.1
Plan cost sharing: Indemnity plans (N = 8,891)				
Deductible ($)	200.05	300.70	–100.65	–54.32
	(164.10)	(164.10)	(16.71)	(18.69)
Coinsurance (%)	17.22	17.41	–0.19	–0.37
	(9.16)	(9.61)	(1.25)	(1.18)
PPOs (N = 6,543)				
In-network				
Deductible ($)	163.64	206.46	–42.82	–14.17
	(214.64)	(257.48)	(23.32)	(21.05)
Coinsurance (%)	14.55	16.85	–2.31	–2.15
	(10.18)	(8.60)	(1.12)	(1.06)
Out-of-network				
Deductible ($)	275.36	343.90	–68.64	–55.06
	(286.14)	(365.19)	(27.05)	(25.45)
Coinsurance (%)	25.88	27.88	–2.00	–2.23
	(10.26)	(11.25)	(1.19)	(1.08)
HMOs (N = 4,753)				
Office visit copayment ($)	6.62	7.25	–0.64	–0.46
	(4.50)	(4.48)	(0.49)	(0.52)

Notes: All figures are weighted by plan enrollment. Adjusted differences are based on linear probability model regressions. The regression specification includes indicator variables for establishment size (six categories), industry (ten categories), state, and whether or not the firm has another location. The model also includes the percentage of workers in five demographic categories (males under age twenty-five, females under age twenty-five, females aged twenty-five to fifty-four, males aged fifty-five and older, females aged fifty-five and older) and the natural log of the ratio of annual payroll to the number of employees. The regression standard errors have been adjusted to account for establishments that offer multiple plans.

important variables than co-payment rates for distinguishing between more- or less-attractive plans. We have no data on these other plan attributes.

An aspect of employer-provided health insurance that has received somewhat less attention from researchers, but is of increasing policy importance, is retiree health benefits. Because nongroup health insurance is typically quite expensive and in some cases not available at any price for older workers, retiree health benefits make it possible for many workers to retire before the age of sixty-five without suffering a loss of insurance coverage. Also, there are substantial gaps in the coverage offered by Medi-

Table 7.8 **Union-nonunion differences in retiree health benefits (CPS benefits supplement data)**

	Union	Nonunion	Difference (union-nonunion) Unadjusted	Adjusted	Adjusted (size)
A. 1988 retiree health insurance supplement (N = 1,098)					
Retiree coverage	.740	.639	.101	.045	n.a.
			(.031)	(.034)	
B. 1993 benefits supplement (N = 1,806)					
Retiree coverage	.766	.598	.167	.146	.128
			(.026)	(.029)	(.028)

Notes: All estimates were obtained using the survey supplement weights. Standard errors are in parentheses. The estimates in the fourth column are the union coefficients from linear probability models that also control for education (four category dummies), age, age squared, female, whether married, female by married, race/ethnicity (dummy variables for black and hispanic), a dummy variable for msa residency, three region dummies, and eight major industry dummies. The adjusted differences in the final column are based on a specification that also includes five establishment size dummies (10–24, 25–49, 50–99, and 100–249, 250+; <10 is the omitted category; four dummies in 1983); n.a. = not available.

care—that is, no coverage for prescription drugs or health insurance from a former employer that supplements Medicare provides significant benefits for retirees over the age of sixty-five.

The 1993 CPS Benefit Supplement and a 1988 supplement focusing on retiree health benefits provide information on whether current workers will be eligible for employer-sponsored health insurance when they retire. Union/nonunion differences estimated using these data are reported in table 7.8. For each year, the sample is limited to workers between the ages of forty-five and sixty-four whose employer offers health insurance to active employees.[17]

The figures in the first panel of table 7.8 show that in 1988 there was a 10 percentage point difference in eligibility for retiree health coverage between union and nonunion workers (74 percent versus 64 percent). Controlling for worker characteristics and employer's industry produces an adjusted difference of roughly half this magnitude. The union/nonunion difference increased between 1988 and 1993 due to an increase in the percentage of union workers with retiree coverage and a decline among nonunion workers. Adjusting for all observables (including firm size) results in a 12.8 percentage point difference in 1993.

To summarize, the data from several CPS supplements show significant, though declining, union effects on health insurance coverage. This trend is

17. It is important to keep the second sample selection criterion (employer health insurance offers) in mind when interpreting the reported union effects. The union effect for all workers combines the retiree coverage effect with the effect on employer offers listed in table 7.6.

consistent with the notion of declining union bargaining power. However, other results suggest a slightly more complicated story in which union efforts are increasingly focused not merely on getting employers to offer health insurance but, rather, on the scope and quality of the benefits that are offered (including the availability of retiree benefits).

7.3.2 Pensions

Table 7.9 presents our results on union/nonunion differences in pension plans. The data are from the same CPS supplements used in the health insurance analysis, and the layout of table 7.9 is the same as table 7.6. As with the health insurance analysis, we examine two outcomes: whether a pension is offered by the worker's firm and, if so, whether he or she is covered.

Table 7.9 Union/nonunion differences in pension availability and coverage (CPS benefits supplement data)

			Difference (union-nonunion)		
	Union	Nonunion	Unadjusted	Adjusted	Adjusted (size)
			A. 1983 (N = 15,637)		
Available	.843	.494	.349	.310	.227
			(.009)	(.009)	(.009)
Covered	.766	.388	.378	.303	.236
			(.009)	(.009)	(.009)
			B. 1988 (N = 15,254)		
Available	.868	.536	.333	.302	.212
			(.011)	(.011)	(.010)
Covered	.772	.415	.358	.281	.212
			(.011)	(.011)	(.011)
			C. 1993 (N = 15,179)		
Available	.862	.586	.275	.268	.172
			(.012)	(.012)	(.011)
Covered	.750	.432	.318	.269	.196
			(.012)	(.012)	(.011)
			D. 1997 (N = 8,371)		
Available	.820	.578	.242	.227	n.a.
			(.017)	(.016)	
Covered	.719	.438	.281	.236	n.a.
			(.017)	(.016)	

Notes: All estimates were obtained using the survey supplement weights. Standard errors are in parentheses. The estimates in the fourth column are the union coefficients from linear probability models that also control for education (four category dummies), age, age squared, female, whether married, female by married, race/ethnicity (dummy variables for black and hispanic), a dummy variable for metropolitan statistical area (MSA) residency, three region dummies, and eight major industry dummies. The adjusted differences in the final column are based on a specification that also includes five establishment size dummies (10–24, 25–49, 50–99, and 100–249, 250+; <10 is the omitted category; four dummies in 1983); n.a. = not available.

For all years, the union/nonunion differences for pensions are larger than those for health insurance. In 1983, union workers were nearly twice as likely to participate in an employer-sponsored pension plan: 76.6 percent versus 38.8 percent. This result is due mainly to a large difference in whether the worker's firm offered a pension plan at all (84.3 percent versus 49.4 percent), though there also is a union edge in coverage conditional on employer offers (0.909 versus 0.784, figures not shown). Holding constant worker and firm characteristics (including establishment size), in 1983 union workers were 23.4 percentage points more likely than nonunion workers to participate in a pension plan.

Over time, the coverage rates for the two groups move in opposite directions, generally falling for union workers and increasing steadily for nonunion workers. In 1993, the adjusted union/nonunion difference was 19.6 percentage points according to our preferred specification and 26.9 percentage points when we do not control for firm size. The more restricted specification implies a slightly smaller union effect, though still large, of 23.5 percentage points in 1997.

7.3.3 Other Fringe Benefits

The National Organizational Survey Data

In this section, we examine union effects on a variety of other benefits using data from the 1993 CPS Employee Benefit Supplement and two other surveys conducted in the 1990s–the 1996 Medical Expenditure Panel Survey (MEPS) and the 1991 National Organization Survey (NOS). Of the three data sets we use in this section, only the NOS is not well known to labor and health economists, so a few words are in order.

The 1991 NOS was designed for research on organizational behavior.[18] The sample was drawn in the following way. Employed respondents to the 1991 General Social Survey (GSS), a nationally representative survey of U.S. adults, were asked to provide the names of their employers. These organizations were then contacted, and their representatives were asked about a number of organizational characteristics and policies, including whether the organization provided a number of nonwage employee benefits. While previous researchers using the NOS have treated the organization as the unit of observation,[19] the public-use file contains both organization-level variables and the individual-level variables from the GSS, including whether each respondent is a union member. We treat the combined GSS/NOS as an individual-level data set and estimate the relationship between an individual's union status and the probability that his or her

18. See Kalleberg et al. (1994) for an overview of the objectives of the NOS and Spaeth and O'Rourke (1994) for details on its design and implementation.
19. For example, see Knoke (1994), Knoke and Ishio (1994), and Huffman (1999).

employer offers various benefits. The drawback of these data is our inability to identify whether the employee is eligible for offered benefits.

Results

Table 7.10 presents regression-adjusted union effects on the receipt of several fringe benefits. The econometric specification varies slightly across these data sets due to differences in data availability, though in all cases it is fairly comparable to our preferred specification for the health insurance and pension regressions discussed previously.

Each of the outcomes in this table is dichotomous (provided or not). The top panel presents results for three types of insurance: dental, life, and long-term disability. The results here are similar to what we find for health insurance and pension benefits. For all three types of insurance, the adjusted union/nonunion difference is statistically significant. The difference is largest for dental insurance (16.5 percentage points) and smallest for life insurance (4.5 percentage points). In the case of long-term disability, our two data sources provide fairly similar estimates—9 percentage points for the 1991 NOS and 6 percentage points for the 1996 MEPS.[20]

The estimated union effects on the various types of paid leave are generally smaller. All three of the surveys have information on paid sick leave. The largest union effect, and the only one that is statistically significant at conventional levels, is estimated using the 1993 CPS Employee Benefit Supplement. The estimate from the NOS is of a comparable magnitude but, because of a much smaller sample size, is not statistically significant.

In the NOS, the employers of union members are 6 percentage points more likely to offer some type of maternity leave. Unfortunately, the data set provides no further details on the nature of that coverage—that is, paid versus unpaid or length of time. There is a similarly sized union effect on vacation coverage from the 1996 MEPS. This is consistent with our previous evidence from the PSID. Union workers are more likely than nonunion workers to have *any* paid vacation.

The last panel presents union/nonunion differences in dependent care benefits from the NOS. Those results indicate no significant difference in the percentage of union and nonunion workers receiving employer-sponsored child care or elder care benefits. Given the unfortunately vague way in which these last two outcomes are defined—child care benefits could include anything from subsidized on-site day care to Section 125 benefit programs that allow employees to pay child care expenses with their money on a pretax basis—we are reluctant to read too much into these results.

Finally, because the NOS and MEPS also contain information on pen-

20. Given differences in survey design, we are reluctant to interpret differences across these sources as reflecting time trends.

Table 7.10 Union effects of fringe benefits other than health insurance in the 1990s

Outcome	Data set	Adjusted union-nonunion
Other insurance		
Dental	1991 NOS	0.165
		(.045)
Life	1991 NOS	0.047
		(0.037)
Long-term disability	1991 NOS	0.092
		(0.045)
Pension/retirement plan		
Offered	1991 NOS	0.187
		(0.041)
	1993 CPS Benefit Supplement	0.172
		(0.011)
	1996 MEPS	0.197
		(0.017)
Leave		
Paid sick leave	1991 NOS	0.052
		(0.042)
	1996 MEPS	0.014
		(0.018)
Maternity leave	1991 NOS	0.065
		(0.037)
Paid vacation	1996 MEPS	0.062
		(0.017)
Dependent care		
Child care benefits	1991 NOS	–0.021
		(0.040)
Elder care benefits	1991 NOS	–0.023
		(0.041)

Notes: Estimated union-nonunion differences are based on linear probability models. All regressions control for the following: age (and age squared), years of education (and education squared), and indicator variables for gender, marital status (married/not married), gender × marital status, firm size (five categories), geographic region (four categories) and race/ethnicity. The NOS and CPS regressions also control for industry. The CPS and MEPS regression controls for whether the individual lives in an MSA. The NOS sample sizes range from 636 to 650, depending on the outcome. Standard errors are in parentheses.

sions and retirement plans, we also compute estimates similar to those in section 7.3.2. The similarity of the estimates from all three data sets is reassuring.

7.4 Conclusions and Discussion

Consistent with other research, and with the comprehensive review in Freeman and Medoff (1984), we find generally significant union effects on the large variety of nonwage aspects of work that we can measure. For a worker with the attributes of a typical union member, a union job offers more vacation, fewer hours per week, the greater likelihood of dental, health, maternity, retirement, and pension benefits, inter alia. We also uncover evidence that there has been a decline in the magnitude of the various differentials over time. This is consistent with the decline in the extent of unionization and possibly a decline in their influence within unionized workplaces, although the inconsistency in measurement across our different data sets makes this conclusion tentative in some cases.

One open question is the extent to which our findings reflect a direct causal impact of unionization on nonwage outcomes. For example, the results may instead reflect systematic worker sorting between union and nonunion workplaces (endogenous unionization) so that omitted worker characteristics and choices rather that unionization per se are the cause of differences in benefit outcomes.

Although our estimates did not account for endogenous unionization, estimates of the union wage effect using data on individuals that attempt to account for endogeneity indicate a substantial positive, causal impact of union status (see, for example, Robinson 1989). Moreover, Blau and Kahn (2000) observe that the impacts of unionization are much the same whether measured within countries over time or across countries subject to similar global influences but with differences in unionization.

A related issue is the mechanism by which unions provide these "premiums." In most variants of the monopoly union model, union gains come exclusively at the expense of nonunion workers and the level of economic activity in the economy. The mechanism is straightforward—unions raise the cost of inputs to the firms. Profit maximizing firms respond by lowering output and substituting away from the higher-priced input. Although this view is widespread among economists, direct supporting evidence is quite scarce. Indeed, the presence of a "deadweight welfare loss" to unionization is a staple of textbook treatments of unionization. Even Freeman and Medoff (1984) who, inter alia, highlight the potential for allocation improvements under unionization, essentially stipulate the existence of such a welfare loss, although they note that their estimate of this loss is small.[21] More recent evidence from DiNardo and Lee (2001) suggests that the employment losses resulting from union impacts on establishment closure are quite small.

Overall, our results suggest that unions have had substantive allocative

21. The "Harberger triangle" in this setting equals one-half the product of the union wage effect, the associated employment decline in the union sector, the fraction of the workforce unionized, and the fraction of total costs associated with labor. Assuming an elasticity of demand for labor of –2/3, an upper bound to the union wage effect of 25 percent, a union share of the workforce of 25 percent in 1981, and a labor share of gross national product (GNP) of 3/4, Freeman and Medoff (1989) estimate the efficiency loss as a share of GNP is $1/2 \times 0.20 \times 0.13 \times 0.25 \times 0.75 = 0.0040$. They note that this estimate is very close to the calculations in Rees (1962).

impact on important nonwage aspects of jobs. These include key benefits, such as health insurance and pensions, for which market conditions for their provision have changed substantially in recent decades. The apparent continuing union impact on these key benefits may both provide a link to the historical origins of the labor movement and a means by which unions may continue to have important impacts on the terms and conditions of employment in the future despite the decline in unionization.

Appendix A
A Simple Model

In this appendix we discuss a parameterization of a simple theoretical model, due to Johnson (1990), of hours determination in a unionized setting. The point of departure is the work of Pencavel and MaCurdy (1986) and Brown and Ashenfelter (1986) and the class of models sometimes referred to as "contract curve" models (hereafter CC models), which constitute an alternative to standard "labor demand" models of union behavior. In the labor demand model, the effect of unions is similar to that of the minimum wage: the union raises the wage above the competitive level, leading to a decrease in employment (assuming hours per worker are fixed). In the basic CC model, however, the union is aware that its wage demands affect employment and takes that effect into account when formulating its bargaining stance. In Brown and Ashenfelter (1986), for example, the union's optimal choice for the level of employment maximizes firm profits, and any resulting rents are divided between the union and the owners of the firm. As in implicit contract models, this process implies that the wage no longer plays the allocative role it does in the simple textbook model and that it is not possible to consider hours worked as the outcome solely of individual preferences interacting with a fixed wage rate.

Johnson (1990) develops a variant of the basic CC model where hours per worker are not fixed and where the union negotiates jointly over wages, employment levels, and hours worked per worker. As it is hard to rationalize hours of work decisions by union members as labor supply responses to higher wages, this model provides a potentially useful framework for understanding the differences between hours worked in union and nonunion environments.[22]

Profits to a firm are given by a revenue function, $V(\cdot)$, that depends on

22. Estimating a standard labor supply equation, with union status used to instrument for the wage, typically results in labor supply elasticity estimates that are much larger than those found in other studies. See DiNardo (1992).

total man-hours and labor costs, which equal the (average) wage rate times total man-hours employed.[23]

$$\text{(A1)} \qquad \pi = V[f(hN)] - whN,$$

where π is profits, h is hours per worker, N is numbers of workers, and w is average wage rate.

The union has preferences over individual worker utility as well as the total number of employed members:

$$\text{(A2)} \qquad R = [U(wh, -h) - U_a]^{\beta}N$$

It is convenient to specify a specific (Stone-Geary) utility function with utility increasing in earnings and decreasing in the number of hours worked.

$$\text{(A3)} \qquad U(wh, -h) = (wh)^{\theta}(T - c - h)^{1-\theta},$$

where $R(\cdot)$ is the union objective function, T is total hours, c is committed leisure, β is parameter of union preferences, θ is parameter of individual preferences, and U_a is nonunion utility.

Equation (A1) defines profits for employers, equation (A2) defines union utility, and equation (A3) is a simple characterization of individual preferences.

When $\beta = 0$, this model is a standard competitive model, and as $\beta \to \infty$, the union cares more about wages and hours per worker. If unions and firms negotiate over w, N, and h, the observed values of h and N for a given contract should satisfy the following conditions.

$$\text{(A4)} \qquad \frac{\partial \pi}{\partial h}\bigg|_{\pi = \pi_0} = \frac{\partial R}{\partial h}\bigg|_{R = R_0}$$

$$\text{(A5)} \qquad \frac{\partial \pi}{\partial N}\bigg|_{\pi = \pi_0} = \frac{\partial R}{\partial N}\bigg|_{R = R_0}$$

These two equations imply the following:

$$\text{(A6)} \qquad (V'f' - w)N = \beta[U(wh, -h) - U_a]^{\beta-1}N(U_1w - U_2)$$

$$\text{(A7)} \qquad (V'f' - w)h = [U(wh, -h) - U_a]^{\beta}$$

Together, equations (6) and (7) imply that

$$\text{(A8)} \qquad \frac{U_2}{U_1} = w\left\{1 - \left[\frac{U_a\mu}{(U_1\beta h)}\right]\right\}, \text{ where } \mu = \frac{[U(wh, -h) - U_a]}{U_a},$$

23. A fuller analysis might treat labor as a quasi-fixed factor as in Oi (1962) so that firms might prefer fewer workers and greater hours per work. However, our simplifying assumption suffices for characterizing differences between union and nonunion workers.

or, to write it in a more intuitive way,

$$\text{(A9)} \qquad \frac{U_2}{U_1} = w\left\{1 - \frac{\mu}{(1 + \mu)\beta E}\right\},$$

where E is elasticity of utility with respect to income.

Recall that in the absence of a union, a utility-maximizing individual worker would set

$$\text{(A10)} \qquad \frac{U_2}{U_1} = w.$$

The condition that $\mu > 0$ (unions place some emphasis on hours and wages, and union workers are better off than their nonunion worker counterparts) implies that union workers would prefer to work more hours at the actual wage rate. The intuition underlying this result is straightforward. In this model, the firm is indifferent between combinations of N and h that yield the same amount of labor. By contrast, the union faces a trade-off between more members and higher utility per member.

Several implications that we can compare with the data include:

1. Union work hours generally will be less than nonunion hours.[24]
2. In addition to working less, union workers will be more likely to report being constrained and desiring to work more hours at the union wage rate.
3. The more union wage setting resembles wage setting in the "competitive" sector (the smaller the difference between union member utility and the nonunion alternative), the smaller is the hours differential and the less "constrained" is the union hours outcome.

Appendix B
Statistical Methods for Reweighting

The procedure described here is a straightforward application of "propensity score" weighting (Rosenbaum and Rubin 1983). The exposition follows the discussion in DiNardo, Fortin, and Lemieux (1996) and Johnston and DiNardo (1997). Let the conditional distribution of hours in year i and sector j be given by $f^{i,j}(h \mid X = x_b)$, let x_b denote the characteristics of our base sample and $x_{i,j}$ the characteristics of the sample that we wish to reweight to have the same distribution of characteristics as the base sample. We are interested in computing the counterfactual distribution

24. This is not true for all possible utility functions; the key is that the worker's utility-constant "income effect" of the higher wage should not be too large.

$$\int h \cdot f^{i,j}(h \mid X = x_b)dx,$$

by means of a reweighting of the actul distribution:

$$\int \theta h \cdot f^{i,j}(h \mid X = x_{i,j})dx,$$

where θ is the appropriate weight, and $f^{i,j}(\cdot)$ is the distribution given the structure of hours in year i. The counterfactual distribution yields the distribution of hours that would have obtained in year i for sector j had the distribution of relevant characteristics in the population been x_b instead of $x_{i,j}$. We can use this distribution to characterize the counterfactual mean (or any other moment of the distribution).

Consider calculating θ when we wish to reweight the 1972 distribution of hours of nonunion members so that they have the same distribution of characteristics as union workers in 1992.

We derive the appropriate weight by noting that the distribution of hours worked among union workers in 1992 is given by

$$\int f^{92,U=1}(h \mid X = x_b)dx \equiv \int f^{92}(h \mid X)g(X \mid U = 1, t = 1992)dx,$$

where the term $g(X \mid U = 1, t = 1992)$ denotes the multivariate distribution of X in the union sector in 1992.

The appropriate counterfactual distribution is given by

$$\int f^{72,U=0}(h \mid X = x_b)dx \equiv \int f^{72,U=0}(h \mid X)g(X \mid U = 1, t = 1992)dx,$$

where the term $f^{72,U=0}$ denotes the structure of hours (i.e., the relationship between hours and characteristics) in the nonunion sector in 1972.

The actual distribution of hours in the nonunion sector in 1972 is given by

$$\int f^{72,U=0}(h \mid X)dx \equiv \int f^{72,U=0}(h \mid X)g(X \mid U = 0, t = 1972)dx.$$

We merely need to solve for the value of θ such that

$$\int f^{72,U=0}(h \mid X = x_b)dx \equiv \int \theta f^{72,U=0}(h \mid X)g(X \mid U = 0, t = 1972)dx.$$

A simple application of Bayes law shows that

$$\theta \propto \frac{\Pr[(U = 1, t = 1992) \mid X]}{\Pr[(U = 0, t = 1972) \mid X]} = \frac{\Pr[(U = 1, t = 1992) \mid X]}{1 - \Pr[(U = 1, t = 1992) \mid X]}.$$

We then compute the sequence of these counterfactual means, say, $\int h \cdot f^{72,j}(h \mid X = x_{1992,U=1})dx$, $\int h \cdot f^{73,j}(h \mid X = x_{92,U=1})dx$, $\int h \cdot f^{74,j}(h \mid X = x_{92,U=1})dx$, . . . $\int h \cdot f^{92,j}(h \mid X = x_{92,U=1})dx$, for both union and nonunion sectors.

Holding *X* constant allows us to make a sensible ceteris paribus comparison—comparing the "same" individuals through time we look at changes in the "structure of hours." One advantage relative to a standard regression framework is that if there is significant "treatment-effect heterogeneity," the analysis holds constant the comparison group so that the relative "strength" of the union effect over time can be easily gleaned from the estimates.

As a practical matter, we compute the probabilities in the previous expression using a logit pooling the 1992 union member sample with the 1972 nonunion member sample. We use a complete set of race dummies, school dummies, marital status dummies, SMSA dummies, and a cubic in age as explanatory variables for the PSID data. For the CPS data we use five education categories, five age categories, an SMSA dummy, three regional dummies, marital status, and three race categories.

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