Union Effects on Health Insurance Provision and Coverage in the United States

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Abstract

Since Freeman and Medoff's (1984) comprehensive review of what unions do, union density in the U.S. has fallen substantially. During the same period, employer provision of health insurance has undergone substantial changes in extent and form. Using individual data from various supplements to the Current Population Survey and establishment data from the 1993 Robert Wood Johnson Foundation survey, we investigate the effects of unionization on employer provision of health benefits. We find that in addition to increasing coverage by employer-provided health benefits, unions reduce employee cost sharing and substantially increase the probability that employer-provided health plans extend to retirees. The union effects on coverage for current employees and for retirees have risen over time, and our estimates suggest that declining unionization explains about 17-20 percent of the decrease in employer-provided health insurance between 1983 and 1997.

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I. Introduction

Because health insurance coverage in the United States largely is employment-based, there is substantial interest among labor and health economists in the factors that determine the extent, quality, and types of health coverage provided in the workplace. Past research has highlighted the important role of labor unions in determining benefit outcomes. In particular, through the preference revelation and enforcement mechanism inherent in the collective bargaining process, unions raise the level of benefits received by employees and the share of benefits in total compensation (Freeman 1981, Freeman and Medoff 1984). Recent data from the U.S. Department of Labor (1998) suggests that these effects potentially are large: as a share of total compensation, employer expenditures on health insurance in unionized work places are nearly double the level in nonunion workplaces. Understanding the factors that generate these differences will provide insight into the changing nature and extent of health insurance coverage in the U.S. labor market and the role of unions in the contemporary U.S. economy.

The focus on unions is timely for several reasons. First, most existing analyses of union/nonunion differences in fringe benefits used data from the 1970s and early 1980s. Since then, union density and influence have declined along with health insurance coverage for lesser-skilled workers (Farber and Levy 1999, Currie and Yelowitz 1999); the union impact on benefits may have changed as well. Second, in response to rapidly rising health care costs, many employers have required employees to pay a larger share of premiums and have replaced traditional indemnity insurance with less costly but more restrictive managed care plans. Whereas previous studies of union effects focused on health coverage per se, union efforts now may be increasingly oriented towards influencing plan quality and resisting higher employee contributions.

To examine the role of unions in the provision of employer-based health insurance in the United States, we use individual survey data from several supplements to the Current Population Survey (CPS) and establishment data from a survey conducted in 1993 by the Robert Wood Johnson Foundation (RWJF). The individual data enable us to decompose employment-based insurance coverage and changes therein into portions attributable to insurance offers by employers, individual employee eligibility, and employee acceptance of offered insurance (takeup). We find that union workers are more likely than nonunion workers to receive health benefits, and the difference mainly is explained by higher probabilities of insurance offers and higher takeup rates for union workers. Although the union effect on offers is limited to workers in small establishments, union effects on takeup operate within small and large establishments alike.

Two likely explanations for higher takeup rates among union workers are smaller required contributions to health plan financing and higher plan quality. Results from the RWJF data indicate that establishments employing union workers are more generous than nonunion establishments with respect to monthly premium contributions. Differences in plan benefits are somewhat less pronounced. Among indemnity and preferred provider organization (PPO) plans, those offered by union establishments tend to have lower deductibles. However, we find no difference in patient cost-sharing between HMOs offered by union and nonunion establishments.

The CPS and RWJF data sets also provide information on the prevalence and financing of retiree health coverage. While less research has been conducted on this benefit than on standard employee health benefits, the growing size of retirement cohorts, rising incidence of job loss among older, senior workers (Neumark, Polsky, and Hansen 1999, Valletta 1999), and declining incidence of retiree coverage (Loprest 1998; U.S. GAO 1997b, 1998) makes retiree health insurance an increasingly important policy issue. The collective-voice view of union behavior suggests that unions have an especially important role to play in the provision of such benefits. In particular, the union voting and bargaining process is likely to produce outcomes that reflect the preferences and needs of older workers who constitute the union's core constituency, rather than younger marginal workers on whom the firm focuses its recruiting efforts. Consistent with this reasoning, we find larger effects of unionization on the provision of retiree health benefits than on the provision of standard health benefits.

In the next section, we discuss union effects on fringe benefit outcomes, changes in the market for health insurance provision, and their implications for our empirical work. Section III describes our CPS and RWJ data and presents basic tabulations. In Section IV, we present results from regression analyses of health insurance outcomes; these regressions control for worker and establishment characteristics that are likely to differ between union and nonunion workplaces. Section V summarizes the results and discusses their implications.

II. Union Effects and the Market for Health Insurance

Union Effects on Health Insurance Provision

The role of US trade unions in obtaining health and welfare benefits for their members dates to the 18th century. Indeed, according to Munts (1967), many early union organizations were established for the provision of such benefits and only later became engaged in bargaining with employers over wages. Roughly a century ago, Beatrice and Sidney Webb wrote of British unions that "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). Around that time, Samuel Gompers and others in the US advocated the expansion of union benefit plans because they benefited workers directly and helped unions maintain membership during economic downturns (Munts 1967). Despite this tradition, however, it was not until the late 1940s, with the Taft-Hartley Act and the Inland Steel Case, that health and welfare benefits became the subject of collective bargaining.

Early economic studies of fringe benefits (Rice 1964; Lester 1967) noted the likely importance of unions in increasing benefits, but due to data limitations did not investigate union effects in detail. The most comprehensive analysis of the effect of unionization on non-wage benefits is the work of Richard Freeman and James Medoff (summarized in their 1984 book). Freeman and Medoff noted that in nonunion workplaces, where entry and exit are the primary adjustment mechanisms, employment and compensation outcomes are determined primarily by the preferences of "marginal" workers, who tend to be young, mobile, and have little invested in the firm. This adjustment mechanism ignores the preferences of less mobile workers, who are older and often have substantial firm-specific investments. By contrast, in a unionized environment the preferences of such inframarginal workers are explicitly taken into account, through union voting and political processes that give voice to a wider set of workers than those at the margin. The resulting bargained outcomes reflect the preferences of workers who are more representative of the complete bargaining unit—along dimensions such as age and seniority—than are the marginal workers.[1]

[1] The simplest statement of this view posits union bargaining based on the preferences of the median union member. As discussed by Farber (1986), however, the conditions necessary for union objective functions to represent the preferences of the median member are unlikely to hold in typical bargaining situations. Budd (1992) argues that the wage compression and standardization policies pursued by unions in the U.S. and other countries are consistent with a "Rawlsian" objective function, which focuses on the utility of the least well-off member.

Freeman and Medoff termed this process of preference revelation and expression "collective voice." In addition to effectively replacing expression of opinion by individual workers, which may be perilous to their employment status, collective voice is important due to the public goods nature of some employment conditions. Features of the employment relationship such as health insurance, occupational safety, and procedures for layoffs and work sharing involve public goods elements and therefore will not be adequately provided for if left to standard market adjustment mechanisms.[2,3]

Due to collective voice and other factors, union effects on the provision of health insurance are likely to be large. First, one key effect of unions is to raise the level of compensation; unless union members place no value on improvements in their health plan, the union bargaining effect on total compensation will raise the coverage and quality of employer health plans. Second, the inframarginal workers empowered by the collective voice mechanism are more likely than marginal workers to be older, to have dependents, and to face salient health and retirement issues. They will likely have stronger preferences concerning the availability and quality of their health insurance coverage while employed. In addition, older workers are likely to place especially strong weight on employer provision of retiree health benefits.

Unions may increase expenditures on health benefits and alter their form through other channels as well. Attractive health benefit packages are a highly visible and readily understood benefit, and as such may be especially attractive to union leaders, who need the approval of

[2] In one of the few direct tests of the collective voice and median voter hypotheses, Kahn (1990) examined occupational safety outcomes for unionized industries and concluded that the preferences of the most senior and the most junior workers mattered most.

[3] Goldstein and Pauly (1976) also pointed out the public good nature of fringe benefits, focusing specifically on health insurance. Like Freeman and Medoff, they assumed that unions choose compensation packages based on the preferences of the median worker within a firm, whereas the decisions of nonunion employers are made to satisfy marginal workers.

union members in order to stay in power.[4] Moreover, if the union helps to administer a health insurance program across multiple work sites, the resulting economies of scale in plan provision may provide the basis for expansion of coverage and improvements in plan quality or choice characteristics.

Existing empirical results broadly support this view of union effects on employer-provided health insurance. Using data from the 1970s, Freeman and Medoff (1984) found that unionization substantially raised the probability workers were covered by employer-provided health plans. Woodbury and Bettinger (1991) used CPS data from 1979 and 1988 to investigate changes in fringe benefit coverage during the 1980s. According to their decomposition analysis, declining union membership was the most important measured factor explaining the decline in employer-provided health insurance over that period. Their results also suggest that as unions were declining in membership, their impact on benefit outcomes also was waning: the marginal effect of union membership on health insurance coverage fell by 28 percent between 1979 and 1988. Even and McPherson (1991) also found that the impact of unionization on insurance coverage fell during the 1980s. Since neither of these studies distinguished between employer offers and employee takeup of coverage, the exact reasons for this decline are unclear.

Freeman and Medoff also analyzed data on employer expenditures for life health, and accident insurance combined, and found that the unionization effect on expenditures is larger than the unionization effect on incidence.[5] Assuming it does not simply reflect greater

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[4] Mabry (1973, p. 98) summarized the view that fringe benefits serve the interests of union leaders as follows: "The administration of [fringe benefit] programs requires a bureaucracy which tends to strengthen the rationale of union existence, membership dependency, and, hence, organizational survival."

[5] The results discussed in Freeman and Medoff (1984) are based on simple linear regressions. However, Freeman (1981) subjected his data to additional tests that accounted for establishment-specific effects and spillovers to nonunion employees. His results largely are consistent with those from the less complex analyses discussed in their 1984 book.

inefficiency in fringe provision, the larger union effect on expenditures than on incidence suggests that improvements in plan quality are an important feature of union effects. Despite the potential effects of unionization on specific features of employer-provided health plans, little analysis of these effects has been conducted, perhaps due to data scarcity. Based on unadjusted comparisons from the 1981 National Medical Care Expenditure Survey, Freeman and Medoff found that health plans in union establishments provide more flexibility in regard to obtaining a second opinion, and that the proportion of health insurance premiums paid by employers was 14 percent higher in union settings. Similarly, using establishment data from the year 1971, Goldstein and Pauly (1976) found that conditional on a set of establishment characteristics, unionization significantly raises the probability that employees offer noncontributory health plans; this was one of the key predictions of their model of union effects on benefit provision.

Changes in Health Insurance Markets

An updating of union effects on health insurance is especially important given the changes in health insurance markets—in the workplace and more generally—that have occurred over the past decade or so. Employers have responded to the rising cost of health care in several ways affecting both the number of workers with any insurance and the nature of the coverage held by insured workers. One response has been to increase the amount that employees are required to contribute directly for insurance in addition to implicit payments in the form of foregone wages (GAO 1997a; Gabel 1999). Higher employee contributions have been shown to reduce the percentage of workers who accept health insurance offered by their employers (Chernew et al. 1997; Shore-Sheppard et al. 1999). Several studies indicate that a decline in takeup among workers who are offered health benefits is the primary reason private insurance coverage has fallen over the past decade or so (Cooper and Schone 1997; Rice et al. 1997; Farber and Levy 1999).

Another health care development that is pertinent to understanding the role of unions in financing health care is a dramatic decline in retiree health benefits over the past 10 to 15 years.[6] Loprest (1998) reports tabulations from BLS surveys indicating that the percentage of workers in medium and large firms that could continue their health insurance into retirement declined from 75 percent in 1985 to 46 percent a decade later. Other survey data also show a large decline in retiree health benefits (GAO 1997b, 1998). Early retirees who are not yet eligible for Medicare are especially affected by this development. Individually-purchased insurance can be quite expensive for this group, and despite recent insurance market reforms persons with serious health conditions may be unable to obtain such coverage at all.[7]

III. Data

To examine the role of unions in the provision of employer-based health insurance in the United States, we use individual and establishment survey data. Our individual data come from several special supplements to the Current Population Survey (CPS): the Benefits Supplements conducted in May 1983, May 1988, and April 1993, a supplement regarding retiree health benefits conducted in August 1988, and the Contingent Work Supplements conducted in February 1995 and February 1997. In the 1983 Benefits Supplement survey, respondents were asked about receipt of employer-provided insurance. Beginning with the 1988 Benefits

[6] In addition to rising health care costs, a major factor influencing this trend is a 1990 ruling by the Financial Accounting Standards Board (FAS 106) requiring employers to report anticipated retiree health care costs as liabilities on their balance sheets.

[7] The Health Insurance Portability and Accountability Act (HIPAA) of 1997 prohibits insurers from denying individually-purchased coverage to persons leaving employer-sponsored group plans.

Supplement, respondents also were asked about employer insurance offers and individual eligibility.[8] The 1993 Benefits Supplement included the widest range of health insurance questions, including ones about retiree health benefits and a limited set of health plan characteristics. This additional information is not available in the other Benefits Supplements or in the Contingent Work Supplements. One additional drawback of the Contingent Work Supplement data is the absence of information on establishment or firm size, which is an important determinant of both union status and health coverage. For all analyses discussed below, we restricted our CPS samples to employed individuals aged 20-64 at the time of the survey, and we excluded self-employed individuals and government workers.[9]

Table 1 presents tabulations that indicate the distribution of unionization and employer-provided health insurance by establishment size (where available) in each CPS sample. The figures show that coverage by employer-provided health insurance plans declined by about 8 percentage points between 1983 and 1997; most of the decline occurred between 1988 and 1993, and it leveled off after 1993.[10] Union membership density declined by about 9 percentage points between 1983 and 1997. Table 1 also documents the well-known positive relationship between

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However, since HIPAA does not regulate the premiums, it does not ensure the availability of affordable coverage for early retirees with costly health conditions.

[8] Currie and Yelowitz (1999) noted that the ordering and wording of the health insurance questions differs between the Benefits Supplements and the Contingent Work Supplements. Although this may affect comparisons over time, Currie and Yelowitz also noted that the trends evident in these data sets are similar to those evident for the same time period in the Survey of Income and Program Participation, in which the insurance questions did not change.

[9] The Contingent Work Supplement samples used in our analyses are smaller than the full sample from the monthly CPS because the questions regarding union status and earnings are posed only to those respondents that will be rotating out of the sample at the end of that month (one quarter of the sample). In the Benefits Supplement data, the BLS matched information on earnings and union status from the May CPS survey, so we are not constrained to use only the outgoing rotation group observations; however, the Benefits Supplements were administered to only one-half of the monthly CPS sample.

[10] Our figures are largely consistent with those presented by Farber and Levy (1999, Table 2), except our coverage rates in 1995 and 1997 are about 2 percentage points lower.

establishment size and union membership, which we account for in analyses using the Benefits Supplements.[11]

We also use establishment data from a telephone survey conducted in 1993 by the Robert Wood Johnson Foundation (RWJF). These data provide a means for validating and reinforcing results from the CPS data, and also provide substantial independent detail on health plan characteristics. The RWJF sample was drawn from ten states: Colorado, Florida, Minnesota, New Mexico, New York, North Dakota, Oklahoma, Oregon, Vermont and Washington. Although the sample is designed to be representative of employers in these states rather than the nation as a whole, aggregate economic and health insurance statistics for this group are fairly comparable to those for the nation as a whole (Cantor et al. 1995).

The full RWJF sample consists of 22,347 private establishments. We exclude from our analysis 493 observations (2.2 percent of the full sample) for which information on the union status of the firm's employees is missing.[12] The RWJF survey also provides detailed information on all the health plans offered by each employer, and some of our analysis is done at the plan level. Our health plan sample contains observations on a total of 20,218 plans offered by 14,737 private establishments for which union status could be determined.[13]

Tables 2 presents sample sizes and summary statistics on union status for the establishment portion of the RWJF data. Survey respondents were asked what percent of the

[11] Prior studies (Bramley, Wunnava and Robinson 1989; Wunnava and Ewing 1999) have found that the effect of unions on benefits is strongest for employees of smaller firms.

[12] Firms with missing data on the union question tend to be larger than average (the mean number of employees is 90.5, compared to 54.4 for firms with valid union information) and, conditional on size, more likely to offer insurance.

[13] We lose 1003 health plan observations (4.7 percent of the total sample) due to missing data on union status.

firm's employees were union members.[14] In much of our analysis we compare establishments with any union employees (hereafter union establishments) with those employing no union workers (nonunion establishments). As shown in the first row of the table, union establishments constitute 6.5 percent of the unweighted sample and 20.9 percent of the employee-weighted sample.[15] In some analyses we divide the union establishments into two groups based on the percentage of employees who are union members, using 50 percent as the cut-off point. The figures in the table show that union establishments are split fairly evenly between these two categories. Similar to the individual data, the figures show that union membership is quite uncommon among employees of small establishments—fewer than 3 percent of firms with less than 10 workers employ any union members—and increases steadily with establishment size. Roughly one third of the establishments in the largest size category (250 or more employees) have some union employees.

Ideally, in estimating the effect of unions on plan characteristics, we would like to distinguish between effects operating within as well as across establishments. Unfortunately, this is not possible since there is ambiguity in the data as to which types of workers are eligible for which plans.[16] Therefore, our plan-level analysis represents a comparison of plans offered by

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[14] The survey was administered to the person at each establishment most knowledgeable about health benefits and firm and worker characteristics.

[15] The unionization rates in Table 2 are not directly comparable to the rates calculated using the CPS data. The employee-weighted mean for the "percent union" variable is, however. For the full sample it equals 10.2 percent, which is slightly less than the rate of 12.5 percent in the April 1993 CPS Benefit Supplement.

[16] Two survey questions elicit broad information on within-establishment differences. Establishments that provide insurance and employ some union workers were asked if nonunion workers were eligible for health benefits. A valid response is available for 65 percent of the relevant (unweighted) observations. Of this group, 94 percent answered that nonunion workers were eligible for benefits. These respondents were then asked whether there were differences in the health benefits offered to union and nonunion employees. Slightly less than half (48 percent) reported that there were differences. The survey provides no information on the exact nature of these differences.

union and nonunion establishments, controlling for employee and firm characteristics that vary at the establishment level.[17]

IV. Results

Health Insurance for Active Employees

Table 3 lists union/nonunion differences in health insurance offers and receipt, estimated using our CPS data.[18] We provide the same decomposition as used by Farber and Levy (1999). For years besides 1983, we are able to identify whether an individual's employer offers health insurance to any of its employees ("employer offers"), whether that employee is eligible for coverage ("eligible"), and whether the employee chooses to accept coverage ("takeup"); eligibility is defined conditional on employer offers, and takeup is defined conditional on offers and eligibility. The coverage rate is the product of these three components:

$$\Pr(\text{covered})=\Pr(\text{employer offer})\bullet\Pr(\text{eligible} \mid \text{offered})\bullet\Pr(\text{takeup} \mid \text{offered, eligible}).$$

In the table, we list the union and nonunion means for each outcome (e.g., the percentage of individuals whose employer offers insurance), the unadjusted difference between the union and nonunion means, and several adjusted estimates of the union/nonunion difference (the "union effect"). The adjusted differences in the fourth column are the coefficients on a union membership dummy variable from linear probability models that also include various individual characteristics and industry dummies, as listed at the bottom of the table.[19] The adjusted union/nonunion differential from these regressions combines the effect of unionization on total

[17] The data set provides information on whether each plan covers union employees. Since over 90 percent of establishments in the survey have no union employees, this variable is highly correlated with an indicator for whether the firm employs any union workers; the two variables yield similar results.

[18] The results for 1995 are similar to those for 1997 and therefore are omitted.

compensation with an effect that relates to the union effect on the share of compensation received in the form of health insurance. If we observed each worker's total compensation, we could separate these two effects. Data on total compensation is not available, but we observe each worker's cash wage. For the regression results reported in the fifth column, we added ln(hourly wage).[20] Regressions reported in the final column include 5 establishment size dummies as explanatory variables (firm size information is unavailable in the 1997 data).

In the third column of Table 3, the unadjusted union/nonunion differences in health insurance receipt range from 22 percentage points in 1988 and 1997 to 27 percentage points in 1983. In years for which we are able to perform our decomposition, differences in the probability that employers offer insurance make a consistently large contribution to the union/nonunion difference in coverage. When we control for individual characteristics and industry in the fourth column, the union effects on all components of the decomposition are reduced somewhat.

The inclusion of the hourly wage in the fifth column reduces the union effect, suggesting that the effect estimated in the fourth column is due in part to the greater total compensation received by union workers. Controlling for firm size (column six) further reduces the union/nonunion gap.[21] Although we are unable to control for establishment size in the 1997 data, the pattern over time in the union effect on outcomes is similar in columns five and six.

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[19] For all regressions with dichotomous dependent variables reported in this paper, we verified that estimation of probits produces results that are similar to those from the linear probability model; we use the latter for ease of interpretation.

[20] The coefficients on the wage variable reflect two opposing forces: a compensating differential effect that produces a negative relationship between wages and fringe benefits, and an unobserved heterogeneity or productivity component that produces a positive relationship. The significant positive coefficient on the wage that we obtain in most specifications suggests that the heterogeneity effect dominates. The important role of heterogeneity in this context is consistent with the simulation results presented by Hwang, Reed, and Hubbard (1992).

The union effect on offers and coverage rose between 1988 and 1993 and then remained approximately constant or fell a bit. Most interesting is the sharply rising union effect on takeup between 1988 and 1997. By 1997, the union effect on takeup was a bit larger than the union effect on employer offers. This is consistent with the view that in addition to bargaining for employer provision of health plans, unions bargain over various aspects of health plan quality, and that the attractiveness of union plans relative to nonunion plans increased between 1988 and 1997.

Recall from Table 1 that between 1983 and 1997 health insurance coverage and union membership among private sector workers fell by 8 and 9 percentage points, respectively. Based on the results presented in Table 3, we can estimate what fraction of the decline in insurance coverage is explained by the decline in unionization. Holding constant employee characteristics other than wages, the decline in union membership explains 20 percent of the decline in insurance coverage.[22] When we also control for wages, the change in the union variable accounts for 17 percent of the change in health insurance coverage between 1983 and 1997.

Table 4 presents additional regressions for the 1988 and 1993 CPS samples in which the union effect is allowed to vary by establishment size.[23] These results show that pooling workers from all establishment sizes obscures large union effects for employees of smaller firms and large changes over time in several of the outcomes. In both 1988 and 1993, the union effect on

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[21] When we reversed the order of inclusion for the wage variable and establishment size dummies, we found in general that the inclusion of establishment size makes a larger marginal contribution to the reduction in the union effect than does inclusion of the wage.

[22] This intertemporal decomposition is done by comparing the actual change in insurance coverage with the change that would have occurred if union membership had remained at 1983 levels. The counterfactual coverage rate is calculated using the regression coefficients and control variable means from the 1997 analysis and the 1983 sample mean for the union variable.

[23] Results for each year are based on a single regression in which we fully interact the union and establishment size dummies. Standard errors for the union effects by size category are obtained through a transformation of the relevant F-test statistic.

insurance offers generally is restricted to establishments with fewer than 25 employees (with the exception of a small but statistically significant effect for establishments with 50 to 99 employees in 1993). Among workers in the smallest size category, the effect on offers increased considerably, from 15.8 percentage points in 1988 to 24.9 percentage points in 1993. As a result of this change and smaller increases in eligibility and takeup, the union/nonunion difference in insurance coverage among workers in the smallest establishment size category more than doubled between 1988 and 1993.

In contrast to the results for offers, differences between union and nonunion workers in takeup are not limited to the smallest firms. In the 1988 sample, the union effect on takeup is between 4.5 and 5.0 percentage points for the first four size categories (up to 99 employees) and a smaller but significant effect of 2.6 percentage points for workers in establishments with 100 to 249 employees. With one exception (the 50 to 99 category), the union effect on takeup increased between 1988 and 1993. In 1993, the union effect on takeup was significant for all size categories. As a result of these effects on takeup, we find that as of 1993 union workers in establishments of all sizes were more likely to have employer-provided health insurance than were nonunion workers.

Analysis of the RWJF establishment data provides further information regarding the effect of unions on employer provision of health insurance. Table 5 compares health insurance offer rates for union and nonunion establishments for the full RWJF sample and the sample broken down by establishment size, using the same size categories used for the CPS samples in Table 4. The results from the two data sets are quite similar. As in the individual data, the establishment-level results indicate that the effect of unions on health insurance offers is most pronounced for small establishments and essentially zero for large ones. Among establishments with fewer than 10 workers, those with union employees are 21.4 percentage points more likely to offer insurance than nonunion establishments with similar observed characteristics. This matches fairly closely with the 24.9 percentage point effect on offers in the 1993 CPS data. Although this effect is large, it is important to keep in mind that fewer than 3 percent of establishments in this size grouping employ any union workers. The regression-adjusted union/nonunion difference falls, both in magnitude and as a proportion of the unadjusted difference, in each of the next two size categories, though it remains statistically significant at conventional levels. At establishment with 50 employees or more, the adjusted union effect is not significant.

In unreported regressions we split union establishments into two groups: those in which the percent unionized was more than 50 percent and those in which the percent unionized was positive but below 50 percent. Point estimates from these regressions suggest that the probability a firm offers health insurance increases with the percentage of employees who are union members, though differences between these two union categories are very small and not statistically significant. Similarly, when we include separate variables indicating the presence of union employees and the percent of the establishment's employees who are union members, the coefficient on the latter is statistically insignificant.

Employee Premium Contributions

We now turn to an examination of health plan characteristics, using the plan-level data from the RWJF employer survey, and beginning with the employer's premium contribution. Specifically, we investigate union effects on the percentage share of single and family premiums paid by employers. We use this share variable rather than a dollar-denominated measure because variation in the latter is likely to reflect cost considerations that are unrelated to the influence of unions, whereas the share variable is more likely to reflect the direct impact of union bargaining power.

Our analysis is complicated by the distribution of the employer contribution variable. The employer's percentage share, S, is distributed as a continuous variable on the percentage point interval [0,100], but a large fraction of the observations take on the maximum value of 100 (and a small fraction take on the minimum value of 0). The large density mass at the maximum makes it difficult to choose an appropriate functional form for regression analysis and raises concern that the results will be sensitive to specification.[24] We therefore apply a semi-parametric estimation approach that controls for establishment characteristics without imposing parametric restrictions on the distribution of the dependent variable or the union effect.

This approach is an application of the technique developed by DiNardo, Fortin, and Lemieux (1996) and applied by DiNardo and Lemieux (1997) to a problem similar to ours. We want to compare the observed distribution of S in union establishments with the distribution that would prevail in nonunion establishments if they had the same characteristics as union establishments—i.e., the union effect conditional on the distribution of control variables. This is achieved by reweighting the nonunion observations by $p(U{=}1 \mid X)/(1-p(U{=}1 \mid X))$, where $p(U{=}1 \mid X)$ is the probability that an establishment is unionized, conditional on characteristics X. This technique works through assigning greater weight to nonunion observations that are similar to union observations in terms of characteristics and lesser weight to nonunion observations that are less similar to union observations in terms of characteristics. Whereas differences between the unadjusted union and nonunion distributions are due to unionization and establishment

[24] This setting may seem like a natural application for a Tobit model. However, unlike the classic Tobit case, in which excess density mass arises from censoring, in our case S=100 is a meaningful outcome and limit that does not reflect censored measurement. Moreover, Tobit models may be biased and inefficient in the presence of heteroscedasticity (Johnston and DiNardo 1997).

characteristics, differences between the unadjusted union distribution and the adjusted nonunion distribution are due to unionization only. This approach is analogous to a Oaxaca (1973) decomposition, except that our approach imposes no parametric restrictions on the relationship between unionization and the outcome variable *S*. The conditional probabilities p(U=1 | X) can be estimated through various means. We use a logit specification to obtain the appropriate fitted probabilities. The vector of control variables *X* is the same set of control variables that was used for analysis of union effects on employer offers (Table 5).

The results of this analysis are reported in Table 6 and Figure 1. The top panel of Table 6 presents results for union effects on employers' share of single coverage premiums. The table lists results for the unadjusted union and nonunion distributions of *S* and the nonunion distribution adjusted for differences in establishment characteristics. The results reported include the mean and median of *S* along with the percentage of employers that pay full cost (*S*=100). Consistent with previous studies using data from the 1970s (Goldstein and Pauly 1976; Freeman and Medoff 1984), the results indicate strong effects of unionization on the generosity of employer premium contributions. On an unadjusted basis, plans offered by union establishments are 12 percentage points more likely to be fully financed by employers (49.4 percent vs. 37.4 percent), and the mean and median employer share both are noticeably larger in unionized establishments.

Controlling for establishment characteristics increases the size of the union/nonunion differential in employer contributions for single coverage. Conditional on establishment characteristics, plans offered by union establishments are about 20 percentage points more likely to be fully financed by employers (49.4 percent vs. 29.6 percent).[25] The difference in the median value of *S* between plans offered by union and nonunion establishments is 13 points (98 vs. 85). Because of the way *S* is truncated, the mean difference is somewhat smaller (8.6 percentage points). To put these differences in perspective, the median and mean premiums for single coverage in the RWJF data set are $148 and $157 per month, respectively. Thus, the 13 percentage point difference in the median values of *S* implies that union workers pay roughly $20 less per month for single coverage than nonunion employees; the difference of 8.6 percentage points in the means of *S* implies a difference of about $13. These differences are visually displayed in Figure 1A, where we plot the actual distribution of the employer's share of premiums for single coverage for plans offered by union establishments and the distribution for nonunion plans adjusted for establishment characteristics. The biggest difference between the two densities is near employer contributions of 75-80 percent, where the mass is greater for the adjusted nonunion density, and 100 percent, where the mass is greater for the union density; union/nonunion differences at lower values of *S* are less pronounced.

The distributions of employer contributions for family coverage (lower panel of Table 6 and Figure 1B) are different from those for single coverage. Most notably, employers are less likely to pay the entire family coverage premium. However, the contrast between union and nonunion establishments is similar to that for single coverage contributions. Conditional on establishment characteristics, union establishments are 15.3 percentage points more likely to pay the full premium for family coverage than are nonunion ones (27.6 percent vs. 12.3 percent). The average union effect is roughly 6 percentage points when the distributions are compared at either the mean or the median. Applied to the median family premium in the RWJF data set

[25] We used a bootstrap technique to estimate the sampling distribution of the union/nonunion differences reported in Table 6. All of the differences are statistically significant at the .01 level, except

($381), this translates to a difference of roughly $23 in the amount that union and nonunion workers are required to contribute each month for family coverage. Figure 1B shows that differences in the densities are most pronounced near 60 percent and 100 percent, where the mass is greater for the unadjusted union density, and near 75 percent, where the mass is greater for the adjusted nonunion density.[26]

Plan Benefit Design

Plans offered to union and nonunion workers may differ in terms of their comprehensiveness. The best information on this aspect of plan quality in the RWJF data comes from questions on the cost-sharing provisions of each plan. Since the relevant cost-sharing variables differ by plan type, we estimate separate models for three plan types: indemnity plans, preferred provider organizations (PPOs) and Health Maintenance Organizations (HMOs). Forty-four percent of the plans in the data set are indemnity plans, 32 percent are PPOs, and 24 percent are HMOs.[27]

With traditional indemnity insurance, the comprehensiveness of a policy is captured by two parameters: the plan deductible and the co-insurance rate. The main difference between traditional indemnity insurance and PPO plans is that PPOs are designed to control costs by giving patients a financial incentive to receive care from a panel of providers who have agreed to accept the insurer's (discounted) fee schedule and oversight. For example, a common PPO

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for the difference in the median family contribution, which is significant at the .05 level.

[26] An alternative estimate of union effects could be obtained by reversing our approach and applying the nonunion distribution of characteristics to the union sample. In general, this produces approximately the same results as those reported in Table 6, except we obtain a larger estimate of the union effect on the median employer contribution for family coverage using the alternative approach.

[27] Interpreting union effects from these stratified regressions would be difficult if union and nonunion establishments tended to offer different types of plans. However, for the most part, this is not the case. Controlling for observable firm characteristics, there is no significant difference between union and nonunion establishments in the probability of offering employees at least one HMO, at least one PPO, or at least one non-HMO plan (PPO or indemnity). The only significant difference in plan offerings is

design might require patients who receive their care in the network and who have met their deductible to pay 10 percent of the cost, whereas those seeing "out-of-network" providers will have to pay 30 percent (after the deductible). Thus, for PPOs we examine union/nonunion differences in deductibles and coinsurance rates for both in- and out-of-network care. HMOs require less cost-sharing by patients than PPOs or traditional indemnity plans but place greater restrictions on which providers they can see. In the typical HMO, patients face no deductible and are charged a fixed dollar amount (usually between $5 and $25) per physician visit. We use the office visit copayment as the cost-sharing outcome for HMO plans.

The results for the plan cost-sharing outcomes are reported in Table 7. The layout is similar to that of previous tables. In the second and third columns we report the (enrollment-weighted) means for plans offered by union and nonunion establishments, respectively. The fourth column contains the unadjusted differences between the means, and the fifth column reported the differences that remain after adjusting for establishment characteristics using a linear regression.[28] The last column reports the number of observations used in each regression.

For all outcomes, the union-nonunion difference is negative, which implies that union establishments offer more comprehensive coverage; for several of the comparisons, however, the difference is not statistically significant. In the case of indemnity plans, the mean deductible for nonunion plans is 50 percent larger than that for union plans ($300 vs. $200). When we control for establishment characteristics the differential is cut roughly in half ($54), but remains statistically significant at the .01 level. The indemnity plan coinsurance rates are nearly identical

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that, controlling for other factors, union establishments are 7 percentage points more likely to offer their employees at least one indemnity plan.

[28] The control variables are the same as those used in the RWJF offer regressions. Unlike the offer regressions, the estimated union effect does not vary materially with establishment size, so we do not report models in which the union and establishment size variables are interacted. In estimating the

for union and nonunion establishments. This is not surprising given how little variation there is in this variable: 72 percent of the indemnity plans in the sample have coinsurance rates of 20 percent.

Union-nonunion differences in deductibles are less pronounced for PPO plans. The mean in-network deductibles are $163 for union plans and $206 for nonunion plans, respectively. This difference is significant at the .10 level, though when we control for establishment characteristics the difference is substantially smaller ($14) and has a t-statistic less than one. For the out-of-network deductible, the regression results imply a statistically significant difference of $55. For both PPO coinsurance rates (in- and out-of-network), there is a significant union effect of roughly two percentage points. Among HMO plans, the union/nonunion difference in mean copayments for office visits is very small and statistically insignificant. This result may arise because in HMO plans it is the breadth and quality of the provider network, rather than cost-sharing parameters, that differentiates higher and lower quality plans.[29]

Retirement Coverage

In the final part of our analysis, we examine union/nonunion differences in retiree health benefits. In the August 1988 and April 1993 CPS files, respondents were asked whether their current employer will provide health insurance at a group rate through their retirement years. Results for this outcome are reported in Table 8. The table layout and regression specifications are similar to those from Table 3, which presented results for current health insurance coverage. In the 1988 survey, the questions regarding retiree coverage were asked only of workers who had

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standard errors we account for the fact that some establishments offer several plans and therefore contribute multiple observations to the estimation sample.

[29] Information on the size and quality of HMO provider panels is not available in the RWJF data.

current employer-provided coverage. We apply this restriction to both samples.[30] Therefore, our estimates indicate the effect of unions on retiree coverage only, not a combination of active and retiree coverage.[31]

The results show that the union effect on employer provision of retiree benefits increased substantially between 1988 and 1993. The unadjusted union effect rose from 10.1 percentage points (Panel A, third column) to 16.7 percentage points (Panel B, third column). Controlling for individual characteristics and cash wages (fifth column), the adjusted differential rose about 10 percentage points, from a statistically insignificant effect of 3.4 percentage points in 1988 to a significant effect of 13.2 percentage points in 1993.[32] The increase in the union effect on retiree benefits between these two years is consistent with the rising union effect on coverage for active employees and the implied improvement in the quality of union plans associated with the rising union effect on takeup (Table 3).

There is no information on establishment size in the 1988 data, so we can not compare results over time for our most complete specification. However, in the 1993 data adding establishment size dummies reduces the estimated union effect only slightly, from 13.2 to 11.7 percentage points.[33] The latter figure is slightly larger in absolute terms than the corresponding union effect for active employee coverage (10.1 percentage points; Table 3, Panel C), as is the

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[30] In the 1993 survey, respondents were asked the retiree coverage questions if their employers offered current insurance to any employees. The results for the 1993 sample are virtually identical to those reported below when we use employer offers rather than employee coverage to define the analysis sample.

[31] The questions on retiree insurance were asked of workers 40 and older in the August 1988 survey and 45 and older in April 1993. For the sake of comparability, we use the latter cut off for both years.

[32] Compared to our results, MacPherson (1992) found a slightly larger and statistically significant effect of unionization on retiree coverage in the August 1988 supplement data. The difference in our results is explained by his inclusion of public sector employees and small differences in regression specification.

[33] Unreported regressions using the 1993 CPS data indicate no significant differences in the union effect across the establishment size categories.

implied effect in adjusted terms. Relative to the average incidence of retirement coverage for nonunion employees, unions raise the incidence of retirement coverage by 20 percent (.117/.513), compared to a 16 percent (.101/.624) effect on coverage for active employees.

The 1993 CPS Benefits Supplement provides additional information on retiree health benefits beyond whether employers provide the benefit throughout respondents' retirement years. Respondents were asked whether their employer will provide health insurance at a group rate at least until the respondent becomes eligible for Medicare at age 65. We combined this variable with the variable indicating coverage throughout their retirement years to form a "pre-Medicare coverage" variable, which takes the value one if retiree coverage will be provided at least until Medicare eligibility and zero otherwise. Respondents also were asked whether they expect that their employer will pay the full cost of retiree coverage; the analysis sample for this variable is restricted to individuals whose employers offer retiree coverage.[34] Panel B of Table 8 lists estimates of the union effect on these outcomes in the 1993 CPS. The unadjusted union/nonunion difference in pre-Medicare coverage is about 17 percentage points, and the adjusted difference (controlling for wages and establishment size) is a bit over half that, though still statistically significant. In addition, on an unadjusted basis union employees are about twice as likely as nonunion employees to be eligible for a retirement health plan for which their employer pays the full cost; the adjusted differences are nearly as large.

The RWJF establishment survey also asked about employer-provided retiree health benefits; regressions from that data set offer additional evidence on the topic and a check on the CPS results. Table 9 displays the effect of unions on the provision of retiree health benefits in

[34] The sample for the employer cost variable is reduced further by "don't know" and missing responses. The 1993 CPS file also indicates whether the respondent expects her employer to pay part of the cost of retiree coverage. We found no significant union effect on this outcome, perhaps because it is so broad as to be largely non-informative with respect to the generosity of employer fringe programs.

the RWJF data. The layout is similar to Table 5, with one exception: because the sample size is reduced by restriction of the retiree coverage sample to establishments that offer health insurance to active employees, we report the sample sizes in the first column. The figures in the first row show that 56 percent of union establishments and 31 percent of nonunion establishments that offer health insurance to active employees also offer retiree health benefits, implying an unadjusted union effect of 25 percentage points. Controlling for observable firm and worker characteristics reduces the union effect to 7.7 percentage points. This is much larger than the adjusted union effect on coverage for active employees in the RWJF data (1.8 percentage points; Table 5). Moreover, compared to union effects on coverage for active employees that were small and insignificant in establishments with 50 or more employees, the union effect on retiree benefits is fairly large and statistically significant for all but the very largest establishment size category. Overall, we find relatively large and consistent union effects on employer provision of retiree health benefits in our 1993 individual and establishment data.[35]

V. Conclusions

The collective-voice role of unions suggests the likelihood of large union effects on fringe benefits such as health insurance. Prior studies found positive union effects on the provision of employer-provided health insurance and expenditures on it. Most of this research, however, was based on data from the 1970s and early 1980s. Since then, there has been a gradual decline in union membership and significant changes in the U.S. health care delivery system. These changes suggest the need for updated analysis of union impacts on the extent and

[35] We also estimated models that account for differences among unionized establishments by replacing the single union dummy with two indicator variables denoting establishments in which fewer or greater than one-half of the employees are union members. Although we do not report these results in a

form of employer-provided health insurance, using data that are recent and also enable analysis of outcomes beyond whether or not workers have insurance coverage.

In this paper we used individual and establishment data to estimate union effects on employer-provided insurance for active employees and retirees. Using the individual data we decomposed the effect of union membership on health insurance coverage into effects on intermediate outcomes that determine coverage: employer offers, individual employee eligibility, and employee takeup. We found that union effects on these outcomes vary by establishment size. In very small firms unions appear to focus on getting employers to provide any insurance at all. For workers in firms with fewer than 25 employees, we found large and significant union effects on employer offers of insurance. This effect, which increased between 1988 and 1993, is the most important factor explaining the large difference in insurance coverage between union and nonunion workers in very small firms.

In contrast, provision of health insurance is essentially universal among firms with more than 100 employees, and offer rates are quite high among firms with 50 to 99 workers. Thus, in larger firms union efforts with respect to health insurance will focus on how much employees are required to pay for their coverage. Using the RWJF establishment data, we found that unionization substantially reduces workers' required premium payments in all plan types and workers' required cost-sharing payments in indemnity and PPO plans. These results on employee cost sharing are consistent with our finding from the CPS of a significant positive effect of union membership on the probability that workers accept (takeup) insurance offered by their employers. The union takeup effect holds for all establishment size categories, and it

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table, we found that the union effect on retiree health coverage is significantly larger in majority-union establishments than it is in minority-union establishments.

became more pronounced between 1988 and 1997, at the same time that the market for health insurance was being transformed.

Both the CPS and RWJF data also enable examination of the determinants of post-retirement health insurance coverage. Consistent with the "collective-voice" hypothesis, which predicts that unions will respond more to the preferences of older workers, we find larger union effects on retiree coverage than on coverage for current employees, especially in our RWJF establishment data. In the CPS data, the union effect on the probability that employers pay the full cost of retiree coverage is particularly large. The union effect on retiree coverage grew substantially between 1988 and 1993, at the same time that the union effect on coverage for current employees was growing. Moreover, while union effects on active employee coverage are limited to small firms (since nearly all firms with more than 50 employees offer insurance), our RWJF results indicate that unions raise access to retiree health benefits in firms of all sizes.

These results are quantitatively important and have implications for the changing provision of health insurance for workers and retirees. Our estimates suggest that declining unionization explains about 17-20 percent of the decrease in employer-provided health insurance among private sector employees during the period 1983-97. This is comparable to the contribution of declining unionization to the rise in male earnings inequality during the 1980s (Fortin and Lemieux 1997). The union effect on retiree coverage is even larger than the effect on current coverage, and declining unionization is likely to explain an even larger share of declining retiree benefits. The associated decline in health insurance for the elderly suggests that public resources for elderly care may become increasingly strained as current and future generations of workers retire, unless expansion of collective bargaining or other means are used to encourage private provision of retiree health benefits.

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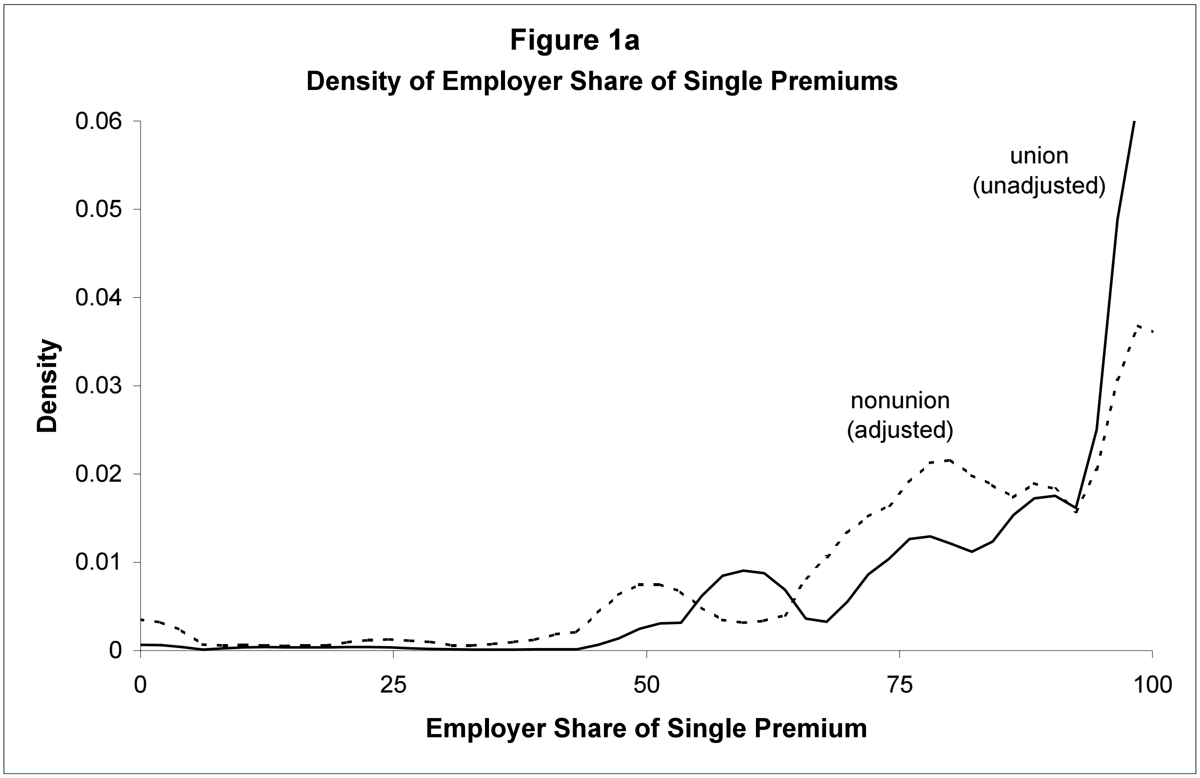

Figure 1a
Density of Employer Share of Single Premiums
0.06
0.05
0.04
0.03
0.02
0.01
0
Density
union
(unadjusted)
nonunion
(adjusted)
0
25
50
75
100
Employer Share of Single Premium

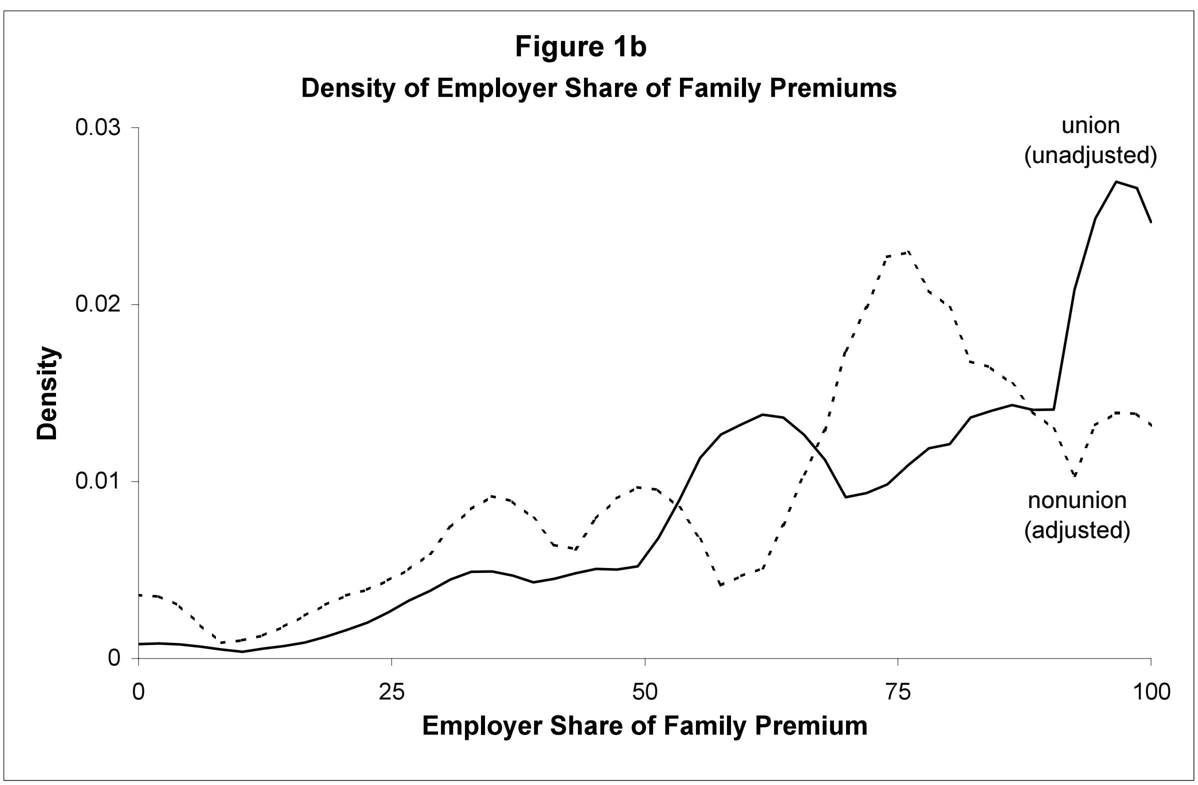

Figure 1b
Density of Employer Share of Family Premiums
0.03
0.02
0.01
0
Density
union
(unadjusted)
nonunion
(adjusted)
0
25
50
75
100
Employer Share of Family Premium

Table 1: Unionization and Health Coverage by Establishment Size, CPS Benefits Supplement Data

	May 1983 (N=15,634)		
Establishment Size	sample share	% union	% covered
Full sample	--	.209	.712
<25	.388	.088	.493
25–99	.224	.221	.746
100–499	.210	.310	.875
500–999	.067	.274	.898
1000+	.110	.354	.946
	May 1988 (N=15,253)		
Establishment Size	sample share	% union	% covered
full sample	--	.149	.701
<10	.212	.039	.430
10–24	.145	.071	.586
25–49	.126	.134	.687
50–99	.107	.158	.754
100–249	.136	.190	.803
250+	.274	.251	.967
	April 1993 (N=15,179)		
Establishment Size	sample share	% union	% covered
full sample	--	.125	.655
<10	.207	.037	.379
10–24	.146	.061	.531
25–49	.121	.099	.620
50–99	.114	.144	.707
100–249	.145	.156	.766
250+	.268	.211	.860
	February 1995 (N=8,979)		
	Sample share	% union	% covered
full sample	--	.114	.617
	February 1997 (N=8,149)		
	Sample share	% union	% covered
full sample	--	.115	.629

Note: All tabulations were weighted using the supplement weights. The samples are restricted to private-sector employees aged 20-64 at the time of the survey.

Table 2: Unionization by Establishment Size, RWJF Data

	Sample Size	Unweighted % of Employees in a Union			Employee Weighted % of Employees in a Union		
		>0	0 to 50	>50	>0	0 to 50	>50
All Firms	21,854	6.53%	2.83%	3.61%	20.85%	10.33%	10.30%
By Establishment Size							
< 10 employees	10,426	2.34	1.10	1.23	2.62	1.14	1.47
10 to 24 employees	5,532	4.66	2.19	2.40	5.73	2.99	2.70
25 to 49 employees	2,360	8.81	3.43	5.34	10.81	4.36	6.44
50 to 99 employees	1,483	14.11	6.00	7.82	19.07	6.83	12.10
100 to 249 employees	1,249	20.78	7.13	12.89	23.53	7.34	15.92
250 + employees	779	31.50	15.53	15.79	38.25	23.23	14.97

Note: There are 25 establishments for which it is possible to determine the presence of a union but the percent of workers who are members is missing. Because of this (and rounding) the second and third column of each panel may not sum to equal the first.

Table 3: Union/Nonunion Differences in Health Insurance Offers and Receipt, CPS Benefits Supplement Data

Panel A: 1983 (N=15,634)

			Difference (union-nonunion)			
	Union	Nonunion	Unadjusted	Adjusted	Adjusted (wages)	Adjusted (wages & size)
Covered	.929	.655	.274	.211	.171	.128
			(.007)	(.009)	(.008)	(.008)

Panel B: 1988 (N=15,253)

			Difference (union-nonunion)			
	Union	Nonunion	Unadjusted	Adjusted	Adjusted (wages)	Adjusted (wages & size)
Employer Offers	.938	.816	.122	.095	.070	.025
			(.007)	(.009)	(.008)	(.008)
Eligible	.962	.881	.081	.056	.041	.038
			(.006)	(.008)	(.007)	(.008)
Takeup	.987	.929	.057	.033	.029	.024
			(.004)	(.029)	(.006)	(.006)
Covered	.890	.668	.222	.152	.112	.072
			(.009)	(.010)	(.010)	(.010)

Panel C: 1993 (N=15,179)

			Difference (union-nonunion)			
	Union	Nonunion	Unadjusted	Adjusted	Adjusted (wages)	Adjusted (wages & size)
Employer Offers	.946	.792	.154	.141	.107	.056
			(.007)	(.009)	(.009)	(.009)
Eligible	.961	.908	.053	.032	.020	.018
			(.007)	(.007)	(.007)	(.007)
Takeup	.957	.867	.090	.068	.059	.050
			(.007)	(.009)	(.009)	(.009)
Covered	.870	.624	.246	.194	.147	.101
			(.010)	(.011)	(.011)	(.011)

Panel D: 1997 (N=8,149)

			Difference (union-nonunion)			
	Union	Nonunion	Unadjusted	Adjusted	Adjusted (wages)	Adjusted (wages & size)
Employer Offers	.931	.820	.112	.099	.084	N/A
			(.009)	(.013)	(.013)	
Eligible	.946	.909	.037	.021	.016	N/A
			(.010)	(.011)	(.010)	
Takeup	.951	.835	.117	.098	.093	N/A
			(.010)	(.014)	(.014)	
Covered	.826	.604	.222	.176	.152	N/A
			(.015)	(.016)	(.016)	

Note: 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 (4 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, 3 region dummies, and 8 major industry dummies. The adjusted differences in the fifth column also include a control for ln(hourly wage), and the adjusted difference in the final column adds 5 establishment size dummies (10-24, 25-49, 50-99, and 100-249, 250+; <10 is the omitted category; 4 dummies in 1983).

N/A = not available

Table 4: Union Effects on Health Insurance Outcomes, by Establishment Size, 1988 and 1993 CPS Benefits Supplements

	Union Effects by Establishment Size (number of employees)					
	< 10	**10 - 24**	**25 - 49**	**50 - 99**	**100 - 249**	**250 +**
1988						
Offer	.158	.112	.013	.008	-.003	.011
	(.030)	(.027)	(.021)	(.022)	(.018)	(.012)
Eligible	-.013	.076	.040	.054	.038	.033
	(.032)	(.027)	(.021)	(.020)	(.017)	(.011)
Take-up	.045	.046	.047	.050	.026	.005
	(.026)	(.177)	(.016)	(.016)	(.013)	(.008)
Covered	.112	.177	.074	.097	.053	.049
	(.036)	(.033)	(.026)	(.026)	(.022)	(.014)
1993						
Offer	.249	.151	.038	.052	.026	.029
	(.032)	(.030)	(.026)	(.022)	(.020)	(.013)
Eligible	.032	.019	.047	.018	.009	.014
	(.030)	(.026)	(.022)	(.019)	(.016)	(.011)
Take-up	.057	.106	.116	.050	.045	.028
	(.035)	(.030)	(.026)	(.022)	(.019)	(.012)
Covered	.229	.201	.159	.090	.071	.067
	(.039)	(.036)	(.032)	(.027)	(.024)	(.016)

Note: All estimates were obtained using the survey supplement weights. Standard errors are in parentheses. The estimated union effects are obtained from the union coefficients and size interaction coefficients from linear probability models that include the same variables as used in the final column of Table 3.

Table 5: Union Effects on Employer Offers of Health Insurance, RWJF Data

	Health Insurance Offer Rates		Difference (union – nonunion)	
	Union	Nonunion	Unadjusted	Adjusted
All Establishments	.989	.836	.153	.018
			(.006)	(.005)
By Establishment Size				
< 10 employees	.875	.524	.351	.214
			(.030)	(.028)
10 to 24 employees	.945	.746	.199	.104
			(.025)	(.024)
25 to 49 employees	.985	.859	.126	.061
			(.022)	(.022)
50 to 99 employees	.968	.922	.046	-.013
			(.017)	(.017)
100 to 249 employees	.990	.957	.033	.016
			(.012)	(.013)
250 + employees	.999	.997	.002	.005
			(.003)	(.004)

Note: All figures are employee-weighted. Standard errors are in parentheses. Sample sizes are reported in Table 2. Adjusted differences are based on linear probability model regressions. The regression specification includes indicator variables for establishment size (full sample only; 6 categories) industry (10 categories), state, and whether or not the firm has another location. The model also includes the percentage of workers in four demographic categories (males under age 25, females under age 25, females 25 to 54, males 55 and older, females 55 and older), and the natural log of the ratio of annual payroll to the number of employees.

Table 6: Union Effects on the Employer's Share of Premium Payments

Employer's Percentage Share of:	**Union, Unadjusted**	**Nonunion, Unadjusted**	**Nonunion, Adjusted**	**Difference (union-nonunion)**	
				Unadjusted	Adjusted
Single Coverage Premium					
Mean Percentage	88.3	81.8	79.7	6.5	8.6
Median Percentage	98	89	85	9	13
% of Employers Paying Full	49.4%	37.9%	29.6%	11.5%	19.8%
number of observations	2,635	16,815	16,815	--	--
Family Coverage Premium					
Mean Percentage	76.3	64.9	66.5	11.4	9.8
Median Percentage	81	70	75	11	6
% of Employers Paying Full	27.6%	15.9%	12.3%	11.7%	15.3%
number of observations	2,615	16,487	16,487	--	--

Notes: The employer's share of premiums is expressed in percentage terms. All statistics are weighted by plan enrollment. Adjusted nonunion figures also are weighted by conditioning weights that account for union-nonunion differences in the distribution of establishment characteristics, as described in the text. The list of establishment characteristics is the same as in Table 5.

Table 7: Union Effects on Health Plan Cost Sharing, by Type of Plan

	Mean (Std. Dev.)		Difference: Union-Nonunion (Std. Error)		Sample Size
	Union	Nonunion	Unadjusted	Adjusted	
Indemnity Plans					
Deductible ($)	200.05	300.70	-100.65	-54.32	8891
	(164.10)	(164.10)	(16.71)	(18.69)	
Coinsurance (%)	17.22	17.41	-0.19	-0.37	8891
	(9.16)	(9.61)	(1.25)	(1.18)	
PPOs					
In-Network					
Deductible ($)	163.64	206.46	-42.82	-14.17	6543
	(214.64)	(257.48)	(23.32)	(21.05)	
Coinsurance (%)	14.55	16.85	-2.31	-2.15	6543
	(10.18)	(8.60)	(1.12)	(1.06)	
Out-of-Network					
Deductible ($)	275.36	343.90	-68.64	-55.06	6216
	(286.14)	(365.19)	(27.05)	(25.45)	
Coinsurance (%)	25.88	27.88	-2.00	-2.23	6215
	(10.26)	(11.25)	(1.19)	(1.08)	
HMOs					
Office Visit	6.62	7.25	-0.64	-0.46	4783
Copayment ($)	(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 (6 categories), industry (10 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 25, females under age 25, females 25 to 54, males 55 and older, females 55 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.

Table 8: Union Effects on Retiree Health Benefits, 1988 and 1993 CPS

Panel A: 1988 Retiree Health Insurance Supplement (N=1098)

	Union	Nonunion	Difference (union-nonunion) Unadjusted	Adjusted	Adjusted (wages)	Adjusted (wages & size)
Retiree Coverage	.740	.639	.101 (.031)	.045 (.034)	.034 (.034)	N/A

Panel B: 1993 Benefits Supplement (N=1806)

	Union	Nonunion	Difference (union-nonunion) Unadjusted	Adjusted	Adjusted (wages)	Adjusted (wages & size)
Retiree Coverage	.766	.598	.167 (.026)	.146 (.029)	.132 (.028)	.117 (.028)
Pre-Medicare Coverage	.806	.638	.168 (.026)	.124 (.028)	.110 (.027)	.094 (.027)
Employer Pays Full Cost	.253	.127	.126 (.026)	.099 (.029)	.098 (.029)	.098 (.029)

Note: All estimates were obtained using the survey supplement weights. Standard errors are in parentheses. The adjusted union effects are the union coefficients from linear probability models that include the same variables as listed at the bottom of Table 3 and in the column headings above. Each sample is restricted to private sector employees aged 45-64 who at the time of the survey were receiving employer-provided health insurance in their name. The 1993 employer cost share regression is restricted to the 979 individuals whose employers provide retiree coverage.
N/A = not available

Table 9: Union Effects on Retiree Health Benefits, RWJF Data, by Establishment Size

	Sample Size	Retiree Health Insurance Offer Rates: Union	Nonunion	Difference (union – nonunion): Unadjusted	Adjusted
All Establishments	14,739	.559	.308	.251	.077
				(.009)	(.009)
By Establishment Size					
< 10 employees	5,182	.313	.152	.161	.118
				(.025)	(.025)
10 to 24 employees	4,164	.267	.162	.104	.061
				(.023)	(.023)
25 to 49 employees	2,029	.283	.179	.104	.102
				(.027)	(.027)
50 to 99 employees	1,385	.413	.185	.228	.160
				(.028)	(.029)
100 to 249 employees	1,211	.438	.242	.196	.159
				(.030)	(.032)
250 + employees	777	.668	.581	.087	.049
				(.036)	(.036)

Note: All figures are employee-weighted. The sample is restricted to establishments offering health insurance to active employees. Standard errors are in parentheses. Adjusted differences are based on linear probability model regressions. The regression specification includes indicator variables for establishment size (full sample only; 6 categories) industry (10 categories), state, and whether or not the firm has another location. The model also includes the percentage of workers in four demographic categories (males under age 25, females under age 25, females 25 to 54, males 55 and older, females 55 and older), and the natural log of the ratio of annual payroll to the number of employees.