The Effects of Unions on Firm Behavior: An Empirical Analysis Using Firm-Level Data

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We evaluate the use of firm-level union coverage rates in empirical models of firm behavior and performance. We focus on the potential for measurement error in both firm- and industry-level data, and find that firm-level union coverage rates provide more precise estimates of union effects. Higher union coverage at a firm is associated with slower employment and sales growth, decreased productivity in nonmanufacturing firms, increased productivity in manufacturing firms, lower profitability, and less investment in durable assets, such as research and development.

Introduction

In recent years, there has been renewed interest in the impact of labor unions on the behavior and performance of firms. Higher unionization rates have been associated with reduced investment in plant, equipment, and research and development (R&D), decreased profitability, slower growth, and higher labor productivity.[1] In many empirical studies, industry unioni-

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[1] These studies include estimation of the impact of unions on profitability (Abowd, 1989; Bronars and Deere, 1990; Hirsch 1991a,b; Ruback and Zimmerman, 1984; Salinger 1984), investment and employment decisions (Bronars and Deere, 1993a; Clark, 1984; Connolly, Hirsch, and Hirschey, 1986; Hirsch 1991b, 1992; Hirsch and Link, 1987), and the productivity of union versus nonunion employees (Allen, 1984; Brown and Medoff, 1978; Clark, 1980).

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zation rates have been used as proxies for firms' unionization rates. This practice involves two potential problems. Industry unionization rates measure a firm's union coverage with error, and their use prevents the inclusion of industry-specific effects in econometric models of firm behavior. A growing number of studies have utilized firm-level union data that can, in principle, provide a more accurate measure of a firm's union coverage.[2]

In this paper, we utilize a new source of firm-level union coverage data to evaluate the use of firm-level data in empirical models of unionization and firm behavior and performance. These unionization rates are calculated from Bureau of Labor Statistics (BLS) union contract data and Standard and Poor's Compustat employment data (see Bureau of Labor Statistics). We focus on the potential for measurement error in both firm- and industry-level coverage data, and examine the implications of omitting industry-specific fixed effects from these empirical models.

Our firm-level unionization rates exhibit substantial variation across firms in the same two-digit SIC industry. For a subsample of our data set, we are able to estimate the degree of measurement error in our firm-level unionization rates by comparing them to independent firm-level coverage rates calculated by Hirsch (1991b). Although nearly half of the observed within-industry variation in our firm-level unionization rates is due to measurement error, substantial within-industry variation remains.

A standard source of union coverage rates for empirical studies of unionization effects is the Current Population Survey (CPS). The CPS does not identify employers, and therefore cannot be used to compute firm-level unionization rates. If the aim in an empirical study is to control for firm-level unionization, then by maintaining the "classical" measurement error assumptions, we show that our admittedly noisy firm-level rates are a better proxy than are the CPS industry rates for a firm's true unionization rate. However, if industry average unionization rates are desired for a specific application, then it is unlikely that an alternative data set could rival the depth of coverage of the CPS.[3]

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[2] These studies include Clark (1984), Ruback and Zimmerman (1984), Abowd (1989), Bronars and Deere (1993b), Card (1990), and Hirsch (1991a,b, 1992).

[3] In some applications, the industry unionization rate may provide a better measure of the union environment facing a firm, particularly when "threat effects" are thought to be important. A limitation of the CPS is that it uses U.S. Census industry codes. Thus, reliable estimates of unionization rates by three-digit SIC codes are generally infeasible from the CPS. The vast majority of publicly traded firms—firms for which other data are readily available—operate in more than one four-digit SIC industry, while many operate in multiple three-digit U.S. Census industries, and several even produce output in more than one two-digit SIC industry. Thus, assigning an industry unionization rate to each firm is problematic. The best approach would be to take weighted averages of industry rates where the weights were proportional to a firm's activity in each industry. The necessary line-of-business data are

An important benefit of firm-level unionization rates is that they permit estimation of models with industry fixed effects, and allow tests of the hypothesis that unobserved industry factors are uncorrelated with union coverage and other regressors. We generally fail to reject these specification tests, and proceed to use random-effects estimators in conjunction with our firm-level data. We then find that the precision of these estimated unionization effects falls sharply, however, if we substitute CPS industry-level coverage rates for our firm-level coverage rates.

In our empirical applications we estimate union effects on several aspects of firm behavior: growth, profitability, productivity, and investment in plant and equipment, R&D, and advertising. We find fairly strong evidence that higher unionization at a firm is associated with slower employment and sales growth, decreased productivity in nonmanufacturing firms, and increased productivity in manufacturing firms. We also find weaker evidence that higher unionization at a firm leads to lower profitability and less investment in durable assets, such as R&D and advertising.

Empirical Methodology

Consider the following econometric model of firm behavior and performance:

$$Y_{ij} = X_{ij}\beta + \tau U_{ij} + \epsilon_{ij}, \tag{1}$$

where the subscript ij denotes firm i and industry j, Y_{ij} is an indicator of firm behavior, X_{ij} is a vector of control variables, U_{ij} is the fraction of a firm's workforce covered by union contracts, and ϵ_{ij} is an error term. In many applications it is reasonable to expect that ϵ_{ij} contains an unobservable industry-specific component, δ_j:

$$\epsilon_{ij} = \delta_j + e_{ij}, \tag{2}$$

where e_{ij} is identically, independently distributed. For example, δ_j may represent unobserved industry-specific R&D and investment opportunities, barriers to entry, or growth opportunities in industry j. If these unobservable industry-specific factors are correlated with union coverage, ordinary-least-squares (OLS) regressions yield inconsistent estimates of union effects.[4] Even if these industry-specific factors are uncorrelated with

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generally quite difficult to assemble. See Salinger (1984) and Bronars and Deere (1993a), however, for empirical analyses that use such a weighted average measure of unionization.

[4] If δ_j is uncorrelated with union coverage but is correlated with some other regressor in X, then the OLS coefficient on union coverage will still be biased unless X_k and unionization are uncorrelated.

a firm's observable characteristics, the OLS-reported standard errors for β and τ will be biased toward zero.

Industry fixed-effects models consistently estimate β and τ, regardless of this correlation.[5] A fixed-effects specification of (1) amounts to estimating

$$DY_{ij} = DX_{ij}\beta + \tau DU_{ij} + e_{ij}, \tag{1'}$$

where the operator D denotes deviation from industry means; hence $DU_{ij} = U_j - \overline{U}_j$, where $\overline{U}_j$ is the mean unionization rate in industry j. Estimates based on within-industry variation are not contaminated by unobservable industry factors, even if these factors are correlated with unionization or other regressors.[6] Using (1′), it is possible to estimate τ consistently even if δ_j is correlated with U_{ij}.

Alternatively, industry random-effects (GLS) models yield consistent estimates of β and τ only if δ_j is uncorrelated with X_{ij} and the unionization rate. In this case OLS, fixed-effects, and GLS estimates are all consistent, but the GLS estimates have the lowest variances. Given firm-level data, it is possible to test for correlation between unobserved industry effects and regressors by comparing fixed-effects estimates of (1′) that treat the δ_j as parameters to GLS estimates of (1) that treat the δ_j as random (see Hausman, 1978). These specification tests provide valuable information on the importance of omitted industry factors and the appropriate econometric model for estimating unions' effects on firms.

Because of data limitations, many studies use the CPS industry unionization rates, denoted by $UCPS_j$, as a proxy for U_{ij}. This approach presents two potential problems. Estimates of τ using industry-level union data are biased toward zero because of measurement error. This bias occurs because a firm's behavior depends on U_{ij}, and $UCPS_j$ measures U_{ij} with error. If $UCPS_j$ accurately measures $\overline{U}_j$, then DU_{ij} is the measurement error component of the proxy variable $UCPS_j$. If $UCPS_j$ measures $\overline{U}_j$ with error, then the within-industry variance, $Var(DU_{ij})$, gives a lower bound for the variance of the measurement error component of $UCPS_j$.[7] The magnitude of this downward bias will be larger, the greater the true within-industry variation in union coverage. However, measurement error is not unique to industry-level data; firm-level union coverage rates are also likely to be measured with error. The smaller the measurement error in firm-level union rates, and the greater the within-industry varia-

[5] Note, however, that β_k is identified only if X_{ijk} varies within industries.

[6] It is also possible that the inclusion of industry effects reduces measurement error that may be present in our firm-level union variable.

[7] See Mellow and Sider (1983) for a discussion of possible errors in CPS industry classifications that could introduce measurement error problems.

tion in union coverage, the greater the benefit from using firm-level unionization data.

In addition, the use of $UCPS_j$ precludes industry fixed-effects estimation because $UCPS_j$ has no within-industry variation. If δ_j is correlated with X_{ij} or $UCPS_j$, then as noted above GLS and OLS estimates using UCPS are inconsistent. Without prior information (see Hausman and Taylor, 1981), it is impossible to test whether such correlation between industry factors and unionization is present when using industry-level union data such as UCPS.

Last, with panel data equation (1) can be first-differenced to eliminate δ_j. Letting the subscript t denote the time period, this yields

$$Y_{ijt} - Y_{ijt-1} = (X_{ijt} - X_{ijt-1})\beta + \tau(U_{ijt} - U_{ijt-1}) + \epsilon_{ijt} - \epsilon_{ijt-1}, \quad (1'')$$

where β, τ, and δ_j are assumed to be time invariant. A key problem with this approach is that measurement error in U_{ijt} may be exacerbated by differencing (Freeman, 1984). In the next section, we show that both our firm-level unionization rates and UCPS are rather noisy measures of the true U_{ij}. We then provide estimates of the relative variance of the measurement error component of UBLS using an independent source of firm-level unionization rates.

Evaluating the Reliability of Firm-Level Unionization Rates

We construct firm-specific unionization rates from all "major" U.S. collective bargaining agreements reported to the BLS, i.e., those involving at least 1,000 workers. Our BLS contract sample contains over 10,000 contracts in manufacturing and nonmanufacturing industries (excluding public sector and construction contracts) from 1971 to 1989.[8] We match these union employment data to the total employment of publicly traded firms in the Compustat database as described in the appendix. In our matched sample, we compute the ratio of union to total employment, denoted as $UBLS_{ij}$ (for firm i in industry j), in each year from 1971 to 1982 for 560 firms. In our empirical applications, we use average levels of $UBLS_{ij}$ computed over three different time periods: 1971–74, 1975–78, and 1979–82.[9]

Table 1 provides for each time period the decomposition of the total variance in $UBLS_{ij}$, into the variance within and between two-digit SIC industry classifications. In each time period at least two-thirds of the total

[8] These data were collected from various issues of the annual BLS publication, *Bargaining Calendar,* which provides information on major contracts scheduled to expire in that year.

[9] We chose these time periods because the number of union contracts declines dramatically in the 1980s and substantially reduces our potential sample size.

TABLE 1

WITHIN AND ACROSS INDUSTRY VARIANCES OF FIRM-LEVEL UNIONIZATION RATES (UBLS)
(percentage of total variation in parentheses)

Year	Within Two-Digit SIC Industry Variance	Across Two-Digit SIC Industry Variance	Total Variance
1971–74	.04010	.01772	.05782
	(69.4)	(30.6)	(100.0)
1975–78	.04582	.02066	.06648
	(68.9)	(31.1)	(100.0)
1979–82	.04005	.01993	.05998
	(66.8)	(33.2)	(100.0)

variation in $UBLS_{ij}$ is within a two-digit industry. These results suggest that there may be substantial differences in true union coverage rates across firms within the same two-digit industry, which implies that the CPS industry average may be a poor proxy for a firm's union coverage. Whether $UBLS_{ij}$ is a superior proxy for U_{ij} depends on how much of the observed variation in $UBLS_{ij}$ is due to measurement error compared to true variation in U_{ij}.

The most direct approach to assessing the measurement error component of $UBLS_{ij}$ is to obtain an independent measure of firms' union coverage ratios for our sample.[10] Hirsch (1991b) has constructed firm-level unionization rates for a sample of large publicly traded firms in 1977, which we denote by UH_{ij}.[11] We can obtain estimates of measurement error variances by examining the covariation between these independently calculated measures of union coverage for the 194 firms that are common to both samples.

In our empirical applications, there is considerable evidence that industry-specific effects explain a substantial portion of the observed variation in behavior and performance across firms. Throughout the empirical analysis that follows, we reject the simple OLS specifications in favor of industry-fixed or random-effects models. Therefore our discussion here emphasizes obtaining estimates of the within-industry measurement error in our firm-level data, i.e., measurement error in $DUBLS_{ij}$.

Suppose that both the BLS contract data and Hirsch's data provide

[10] Greene (1990, pp. 293–303) provides a useful discussion of the issues involved.

[11] See Hirsch (1991b) for details. The means and standard deviations of UH and UBLS in the matched sample are .468 (.225) and .327 (.225), respectively. The comparable statistics for the full samples are .380 (.300) and .329 (.281) for UBLS and UH, respectively.

unbiased but noisy measures of unionization at firm i in industry j. We can then write $UBLS_{ij}$ and UH_{ij} as

$$UBLS_{ij} = U_{ij} + EBLS_{ij} \tag{3a}$$

and

$$UH_{ij} = U_{ij} + EH_{ij}, \tag{3b}$$

where $EBLS_{ij}$ and EH_{ij} are zero mean i.i.d. measurement error components. Assuming that $EBLS_{ij}$ and EH_{ij} are orthogonal, and each is orthogonal to U_{ij}, it follows that

$$\mathrm{Var}(U) = \mathrm{Cov}(UBLS, UH) \tag{4a}$$

and

$$\begin{aligned}\mathrm{Var}(EBLS) &= \mathrm{Var}(UBLS) - \mathrm{Var}(U)\\ &= \mathrm{Var}(UBLS) - \mathrm{Cov}(UBLS, UH).\end{aligned} \tag{4b}$$

The inclusion of industry dummy variables in regression analyses of union effects implies that the resulting estimates depend on only within-industry variation. The variation in true unionization rates can be decomposed into within-industry and between-industry components. If the within-industry deviation in UBLS and UH, denoted DUBLS and DUH, are independent, then the true within-industry variance is given by[12]

$$\mathrm{Var}(DU) = \mathrm{Cov}(DUBLS, DUH), \tag{5a}$$

while

$$\mathrm{Var}(DEBLS) = \mathrm{Var}(DUBLS) - \mathrm{Cov}(DUBLS, DUH). \tag{5b}$$

In a similar fashion the between-industry variance, $\mathrm{Var}(\overline{U})$, equals

$$\mathrm{Var}(\overline{U}) = \mathrm{Cov}(\overline{UBLS}, \overline{UH}), \tag{6a}$$

while

$$\mathrm{Var}(\overline{EBLS}) = \mathrm{Var}(\overline{UBLS}) - \mathrm{Cov}(\overline{UBLS}, \overline{UH}) \tag{6b}$$

where the "$\overline{\ \ }$" indicates industry means. Equations (3)–(6) can be used to estimate the fraction of the variance in UBLS (total variance and its within- and between-industry components) that is due to measurement error.

[12] If we had prior information that measurement error was more of a problem in some industries, we would weight observations differently on the basis of a firm's industry. Our maintained hypothesis throughout the analysis is that measurement error variances are equal for each industry.

Finally, information on the measurement error component of $UCPS_j$ can also be obtained. The mean squared error of $UCPS_j$ around the true unionization rate, U_{ij}, can be written as

$$MSE(UCPS) = Var(DU) + \Delta, \tag{7}$$

where $\Delta \geq 0$ is the weighted (by number of firms) average across industries of the squared difference between $UCPS_j$ and $\overline{U}_j$. Thus the true within-industry variance, which equals Cov(DUBLS,DUH) from (5a), provides a lower bound on the error from using $UCPS_j$ as a proxy for U_{ij}. Also, the inclusion of $UCPS_j$ precludes the use of industry dummies.

In Table 2 we use estimates of equations (3)–(6) for the 1975–78 time period to decompose the total observed variation in UBLS into within- and between-industry components and to estimate the percentages of the overall variation and of these components that are due to measurement error. We find that the covariance between UBLS and UH, which is our estimate of the true variance in union coverage (i.e., Var(U)), is 56.1 percent of the total observed variance of UBLS. For comparison we find that 56.2 percent

TABLE 2
ESTIMATED VARIANCE DECOMPOSITIONS OF FIRM-LEVEL UNIONIZATION RATES (UBLS) 1975–78

	True within-Industry Variance	Within-Industry Variance due to Measurement Error	True between-Industry Variance	Between-Industry Variance due to Measurement Error	True Total Variance	Total Measurement Error Variance	Total Observed Variance
Variation in UBLS	0.02470	0.02112	0.01260	0.00806	0.03730	0.02918	0.06648
Percent of Observed Total Variation	37.1	31.8	19.0	12.1	56.1	43.9	100.0
Percent of Observed within-Industry Variation	53.9	46.1	—	—	—	—	—
Percent of Observed between-Industry Variation	—	—	60.0	40.0	—	—	—
Percent of True Total Variation	66.2	—	33.8	—	100.0	—	—

of the total observed variation in Hirsch's measure of union coverage reflects true variation in unionization. Using the covariance of DUBLS and DUH to estimate the within-industry variance in true unionization (i.e., Var(DU)), we find that true within-industry variation in unionization accounts for 53.9 percent of the observed within-industry variation in UBLS. The analogous estimate of Var($\overline{\text{U}}$) implies that true between-industry variation accounts for 61.0 percent of the observed between-industry variation in UBLS. The comparison figures for Hirsch's data are that 46.7 percent of the observed within-industry and 84.7 percent of the observed between-industry variation in UH reflect true variation in unionization.

At this point, it is worth considering further the possible sources of measurement error in UBLS and UH. UBLS is derived from contemporaneous information on collective bargaining contracts covering at least 1,000 workers. As noted in the appendix, we adjust for workers in smaller contracts using data from 1974. Our coverage rates exclude workers covered in multiemployer contracts because we cannot identify the employers. Hirsch's data for 1977 are derived from three sources: (1) a retrospective survey of 1977 coverage conducted in 1987; (2) adjusted figures from responses on 1987 coverage from the same survey; and (3) adjusted figures on 1972 coverage from a survey conducted around 1972. The survey questions for 1977 and 1987 coverages referred to a firm's entire North American workforce. There are several potential sources of measurement error in Hirsch's 1977 coverage rates: (1) retrospective recall bias; (2) inclusion of Canadian plants where unionization is higher than in the United States; and (3) inaccuracies in the data adjustments.[13] Our priors about the greater potential for between-industry errors in UBLS was somewhat confirmed by the fact that UH is a less noisy measure of between-industry variation. The relatively better performance of UBLS in measuring within-industry variation may largely reflect the effect of using contemporaneous information.

Despite the sizable within-industry measurement error in both UBLS and UH, the results in Table 2 confirm that UCPS_j is an even noisier measure of U_{ij}. Recall that true within-industry variation in U_{ij} is a lower bound for the measurement error component of UCPS_j. Table 2 shows that, in the absence of measurement error in union coverage, within-industry variation accounts for 66.2 percent (.0247/.0373) of the total varia-

[13] The data adjustments are made for about 27 percent of Hirsch's entire sample but an unknown fraction of the sample in common with UBLS. Inaccuracies in these adjustments are exacerbated by the inclusion of Canadian plants because of differences in unionization trends between the two countries. Note also that the inclusion of Canadian plants helps explain the higher mean of UH as reported in note 11.

tion in U_{ij}. Thus $UCPS_j$ can account for at most 33.8 percent of the total variation in union coverage across firms. Hence $UCPS_j$ appears to be a noisier measure of U_{ij} than does either of the firm-level unionization rates, UH or UBLS. When coupled with the fact that the use of UCPS precludes the estimation of industry fixed-effects, it follows that firm-level union coverage rates can provide more accurate estimates of union effects on firm behavior.[14]

Empirical Applications

In the next four subsections, we examine the relationship between a firm's union coverage and its behavior and performance. We estimate two cross-sectional regressions using variable means calculated over the time periods 1975–78 and 1979–82, respectively. Annual observations on each firm's union coverage are combined with annual data on firms from Compustat before averaging. We estimate separate regressions for the manufacturing and nonmanufacturing sectors, and we estimate the impact of union coverage on firm behavior with and without a fairly extensive set of control variables in the regression.[15] The coefficient on unionization in the specifications with the control variables estimates only the direct effect of unionization, holding several firm characteristics constant. On the other hand, as stressed by Hirsch (1991b), the union coefficient in the regressions without other controls estimates both the direct effect of unionization and its indirect effect through changes in the control variables. Each row of our tables presents a different empirical specification of the model using either UBLS or UCPS as the measure of union coverage, while each column represents a different dependent variable/time period combination. Because of space limitations, our tables present only the estimated union coefficient and associated standard errors.[16]

For regressions using UBLS, we estimated industry fixed-effects, random-effects, and first-differenced regression models.[17] For each

[14] Industry fixed-effects and industry average unionization rates can be included in the same regression if the industry level of aggregation differs. For example, Connolly, Hirsch, and Hirschey (1986) assign unionization rates by three-digit U.S. Census industry, and include one- and two-digit SIC industry dummy variables in their regressions. A drawback to this approach is that it requires a large firm to be assigned to a rather narrow industry classification, even though its employment and sales may cut across a variety of three-digit U.S. Census industries.

[15] Sample sizes vary across dependent variables and time periods because of missing data. If the regressions with larger samples are reestimated with the smaller samples of other regressions the results are not materially changed.

[16] More details on the regression results are available from the authors.

[17] The first-differenced specifications entail regressing $\bar{Y}_{79-82} - \bar{Y}_{75-78}$ on $\bar{X}_{79-82} - \bar{X}_{75-78}$. First differ-

random-effects regression, we test the hypothesis that δ_j is uncorrelated with X_{ij} and $UBLS_{ij}$. When we fail to reject this hypothesis we report the more efficient random-effects estimates as the "Industry Effects" specifications. When we reject the hypothesis that δ_j is uncorrelated with X_{ij} or $UBLS_{ij}$, we report the consistent fixed-effects estimates. For purposes of comparison we also present estimated union effects from OLS, random-effects, and first-differenced regressions using UCPS.

It is important to note that the estimated union effects we present based on UBLS have not been adjusted for measurement error bias. If 46.1 percent of within-industry variation in UBLS is due to measurement error, consistent estimates of unionization effects are obtained by multiplying the reported results by $1.855 = 1/(1-.461)$. This adjustment factor applies to our estimated standard errors as well, and hence does not affect the significance level of the reported results. An alternative approach to unbiased estimation is to utilize an instrumental variables (IV) technique with UH employed as an instrument for UBLS. We did not take this approach because it would have greatly reduced our sample sizes, especially in nonmanufacturing. As a check on the accuracy of this adjustment, however, we calculated IV estimates for manufacturing for the 1975–78 time period and compared them to GLS estimates derived from the same (smaller) samples. The results of this comparison were quite encouraging. For the profitability and investment regressions, the IV estimates of the union coefficient were about twice as large as the GLS estimates and the significance levels were very similar. For the growth measures, the IV estimates of the union coefficient were about 25 percent larger than the GLS estimates and the significance levels were again similar. The IV estimates of union productivity effects were, on the other hand, insignificant. Thus the suggested average upward coefficient adjustment of about 80 percent seems consistent with the results of the IV procedure.

Unionization and profitability. A number of authors (including Abowd, 1989; Bronars and Deere, 1990; Hirsch, 1991a,b; Ruback and Zimmerman, 1984; and Salinger, 1984) have examined the impact of unionization on the profitability of firms. In this section, we examine the impact of higher union coverage on three different measures of a firm's profitability:

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ences in UBLS and UCPS may be even noisier measures of the true change in U_{ij} (see Bound and Krueger, 1991). Although we present estimates of union effects based on first-differenced data, these estimated effects may contain a substantial bias and should be viewed with some skepticism. See Freeman (1984) for a related discussion pertaining to estimates of union wage effects in panel data based on changes in union status. In these studies, measurement error due to misclassification or coding errors can result in sizable bias in estimated union wage effects.

(1) Tobin's q; (2) Excess Market Value/Sales, the ratio of excess market value of the firm to sales; and (3) Net Operating Income/Sales, the ratio of a firm's annual net operating income to its sales.[18]

We model a firm's profitability as a function of several lagged variables: log(Capital Stock/Employment), the logarithm of the firm's capital/labor ratio; Sales Growth, the annual rate of growth in sales; Advertising/Sales, the ratio of advertising expenditures to sales; and R&D/Sales, the ratio of R&D expenditures to sales. Each variable is calculated using Compustat data averaged over the preceding four-year period. For specifications in the manufacturing sector, we include Four-Firm Concentration Ratio, the four-firm concentration ratio in the firm's primary industry, adjusted for imports.

Table 3 presents the estimated profitability effects of a change in union coverage. These effects are negative and statistically significant in manufacturing in the random-effects regressions using UBLS and no other control variables. After controlling for a firm's observable characteristics, the profitability effects are reduced by about one-half and are statistically significant only for Net Operating Income/Sales. In nonmanufacturing, there are significant negative effects of unionization on Tobin's q and Excess Market Value/Sales in the random-effects regressions with no other control variables. When other regressors are included, only the effect on Tobin's q remains statistically significant. The first-differenced regressions generally yield insignificant results, with only marginally significant effects on Excess Market Value/Sales in nonmanufacturing.

In general, we find fairly strong and significant evidence that the total effect of higher union coverage rates is to reduce profitability. There is rather weak evidence that unions directly decrease profits, conditional on lagged values of a firm's capital/labor ratio, sales growth, and investment in R&D and advertising. After adjusting the random-effects estimates for measurement error bias, a 10 percent (between 3.5 and 4 percentage points) increase in union coverage reduces Tobin's q by 1.5 to 2.5 percent in nonmanufacturing, and reduces Net Operating Income/Sales by 1.5 to 4 percent in manufacturing. These results are consistent with Connolly, Hirsch, and Hirschey's (1986) finding that unions indirectly reduce profitability through changes in investment behavior and growth.

Estimated profitability effects using UCPS vary substantially, depending on the empirical specification. Both OLS and random-effects regressions yield sizable negative effects in manufacturing. The OLS standard

[18] Tobin's q is the ratio of a firm's market value of equity plus book value of debt to the book value of its tangible assets. Excess Market Value/Sales is a firm's market value of equity plus book value of debt minus the book value of its assets, divided by its annual sales.

TABLE 3

IMPACT OF A CHANGE IN UNION COVERAGE ON FIRM PROFITABILITY
(standard errors in parentheses)

	Measure of Profitability and Time Period					
	Tobin's q		Excess Market Value/Sales		Net Operating Income/Sales	
	1975–78	1979–82	1975–78	1979–82	1975–78	1979–82
Manufacturing[a]						
Mean (S.D.)	1.080	1.109	.500	.445	.120	.106
	(.760)	(.680)	(.368)	(.420)	(.057)	(.060)
Union Coverage Variable: UBLS						
Industry Effects	−.425**	−.379**	−.229**	−.232**	−.039**	−.061**
	(.172)	(.163)	(.083)	(.094)	(.013)	(.015)
Industry Effects with Controls	−.078	−.238	−.138	−.129	−.028**	−.042**
	(.148)	(.173)	(.085)	(.097)	(.012)	(.015)
First Difference		−.212		−.032		.015
		(.198)		(.110)		(.016)
First Difference with Controls		−.320		−.089		.006
		(.208)		(.115)		(.016)
Union Coverage Variable: UCPS						
OLS with Controls	−.837**	−1.075**	−.423**	−.557**	−.076**	−.121**
	(.291)	(.317)	(.161)	(.172)	(.023)	(.025)
Industry Effects with Controls	−.462	−1.177	−.312	−1.109	−.053	−.110
	(1.284)	(1.861)	(.327)	(1.142)	(.061)	(.070)
First Difference with Controls		.566		−.884		.001
		(1.129)		(.787)		(.091)
Nonmanufacturing[b]						
Mean (S.D.)	.978	.904	.028	−.086	.194	.184
	(.456)	(.361)	(.770)	(1.035)	(.142)	(.143)
Union Coverage Variable: UBLS						
Industry Effects	−.303**	−.315**	−.052	−.333**	.046*	.064**
	(.097)	(.109)	(.194)	(.162)	(.025)	(.031)
Industry Effects with Controls	−.178*	−.229**	.037	−.169	.020	.034
	(.091)	(.112)	(.182)	(.160)	(.026)	(.029)
First Difference		−.144		−.302*		.005
		(.177)		(.175)		(.033)
First Difference with Controls		−.084		−.329*		.009
		(.192)		(.197)		(.038)
Union Coverage Variable: UCPS						
OLS with Controls	.352*	.524**	.074	.390	.044	.069
	(.199)	(.239)	(.299)	(.365)	(.052)	(.065)
Industry Effects with Controls	.456	.618	−.812	.652	.014	−.017
	(.577)	(.670)	(1.347)	(.871)	(.102)	(.094)
First Difference with Controls		−2.158**		−1.066		−.563**
		(.716)		(.891)		(.136)

* Significant at .10 level; **significant at .05 level.

[a] Control variables include Four-Firm Concentration Ratio and lagged values of log(Capital Stock/Employment), Sales Growth, Advertising/Sales, and R&D/Sales. Sample sizes ranged from 280 to 300 for 1979–82 and from 320 to 360 for 1975–78.

[b] Control variables include lagged values of log(Capital Stock/Employment), Sales Growth, Advertising/Sales, and R&D/Sales. Sample sizes ranged from 120 to 130 for 1979–82 and from 130 to 150 for 1975–78.

errors appear to contain a substantial downward bias, due to the within-industry correlation in the disturbance term. GLS models generally lead to imprecisely estimated and insignificant profitability effects. In non-manufacturing, estimates based on first-differenced regressions using UCPS yield large negative and statistically significant profitability effects for Tobin's q and Net Operating Income/Sales. In general, however, the first-differenced estimates differ in both direction and magnitude from the OLS and GLS estimates.

Unionization and investment. A number of researchers have examined the relationship between a firm's investment in durable assets and its unionization rate (see Bronars and Deere, 1993a; Connolly, Hirsch, and Hirschey, 1986; Hirsch, 1991b, 1992; and Hirsch and Link, 1987). The maintained hypothesis is that unions are able to extract some of the returns (quasi rents) to such intangible sunk assets, hence reducing the firm's incentive to invest in these assets.[19] We test this hypothesis by estimating the impact of union coverage on a firm's investment in plant and equipment, research and development, and advertising. We also estimate the impact of union coverage on the firm's capital/labor ratio.

The dependent variables in our empirical analysis are averages over 1975–78 and 1979–82, respectively, of (1) Capital Expenditures/Sales, the ratio of expenditures on plant and equipment to sales, (2) R&D/Sales, (3) Advertising/Sales, and (4) log(Capital Stock/Employment). We include Tobin's q and Four-Firm Concentration Ratio (for firms in the manufacturing sector) as controls for a firm's potential profitability. The additional control variables are lagged values (averaged over the previous four-year period) of Sales Growth; log(Capital Stock/Employment); R&D/Sales; and Capital Expenditures/Capital Stock, the ratio of current investment to the stock of plant and equipment.[20] Capital Expenditures/Capital Stock is a measure of the portion of a firm's capital stock that has been put in place in the recent past, and hence should control for some unobserved differences in growth potential across firms.

For two random-effects regressions in nonmanufacturing, Capital Expenditures/Sales in 1979–82 and R&D/Sales in 1975–78, we reject the hypothesis that δ_j is uncorrelated with X_{ij} and $UBLS_{ij}$. Therefore, Table 4 presents fixed-effects estimates in these two cases and the more efficient random-effects estimates for the remainder of the Industry Effects specifications.

[19] Grout (1984) and Baldwin (1983) develop theoretical models that imply that firms will underinvest in capital as a response to unionization.

[20] The controls that are included vary some across dependent variables so that own lags are never included. See the notes to the tables for the exact content of the controls in each regression.

TABLE 4

IMPACT OF A CHANGE IN UNION COVERAGE ON FIRM INVESTMENT
(standard errors in parentheses)

	Measure of Investment and Time Period							
	Capital Expenditures/ Sales		R&D/Sales		Advertising/Sales		log(Capital Stock/ Employment)	
	1975–78	1979–82	1975–78	1979–82	1975–78	1979–82	1975–78	1979–82
Manufacturing[a]								
Mean (S.D.)	5.779	6.929	1.188	1.402	1.159	1.323	3.010	3.145
	(4.173)	(4.684)	(1.353)	(1.573)	(2.053)	(2.371)	(.864)	(.895)
Union Coverage Variable: UBLS								
Industry Effects	−1.766**	−.911	−.623**	−.997**	−.915**	−1.148**	.196	−.015
	(.867)	(1.176)	(.265)	(.364)	(.400)	(.532)	(.135)	(.156)
Industry Effects with Controls	−1.155	−.162	−.281	−.813**	−1.073**	−.865	.360**	.192
	(.885)	(1.233)	(.274)	(.365)	(.404)	(.543)	(.111)	(.130)
First Difference		−1.859		.260		.000		.179
		(1.643)		(.291)		(.397)		(.110)
First Difference with Controls		−1.521		.204		.207		.222*
		(1.743)		(.310)		(.412)		(.113)
Union Coverage Variable: UCPS								
OLS with Controls	4.360**	3.916*	−1.284**	−1.238*	−2.825**	−3.992**	1.394**	1.014**
	(1.835)	(2.330)	(.609)	(.722)	(.924)	(1.152)	(.269)	(.293)
Industry Effects with Controls	7.152	2.334	−1.049	−.979	−3.738	−6.232	2.306*	1.568
	(6.367)	(8.568)	(3.094)	(3.117)	(4.649)	(5.867)	(1.347)	(1.510)
First Difference with Controls		3.557		−2.174		−1.748		1.460**
		(9.332)		(1.697)		(2.251)		(.610)

Nonmanufacturing[b]								
Mean (S.D.)	16.68	18.47	.097	.075	.570	.646	4.252	4.518
	(16.34)	(15.69)	(.377)	(.295)	(1.024)	(1.160)	(1.742)	(1.539)
Union Coverage Variable: UBLS								
Industry Effects	6.583	8.975*	−.163*	−.088	−.492**	−.493*	.371*	.339*
	(4.020)	(4.676)	(.085)	(.108)	(.240)	(.288)	(.202)	(.192)
Industry Effects with Controls	7.284	2.655[c]	−.213**[c]	−.033	−.375	−.323	.374**	.051
	(4.717)	(6.103)	(.086)	(.106)	(.279)	(.290)	(.145)	(.157)
First Difference		−8.166		.015		.190		.148
		(5.754)		(.087)		(.345)		(.177)
First Difference with Controls		−8.845		.026		.076		.211
		(6.418)		(.087)		(.372)		(.158)
Union Coverage Variable: UCPS								
OLS with Controls	41.554**	28.390**	−.229	−.392*	−1.420**	−1.716**	2.289**	2.201**
	(8.189)	(8.812)	(.200)	(.235)	(.569)	(.677)	(.506)	(.576)
Industry effects with Controls	20.482	17.770	−.774	−.168	−.480	−1.698	2.063*	1.709
	(15.490)	(14.729)	(.582)	(.584)	(1.556)	(2.061)	(1.065)	(1.338)
First Difference with Controls		−48.806**		.544		−2.987**		−1.822*
		(23.046)		(.572)		(1.494)		(.954)

* Significant at .10 level; **significant at .05 level.

[a] Controls include Four-Firm Concentration Ratio and lagged values of Sales Growth and R&D/Sales for the Capital Expenditures/Sales regression; Four-Firm Concentration Ratio and lagged values of log(Capital Stock/Employment), Sales Growth, and Capital Expenditures/Capital Stock for the R&D/Sales and Advertising/Sales regressions; and Four-Firm Concentration Ratio and lagged values of Capital Expenditures/Sales and Sales Growth for the log(Capital Stock/Employment) regression. Sample sizes ranged from 280 to 300 for 1979–82 and from 320 to 360 for 1975–78.

[b] Controls include lagged values of Sales Growth and R&D/Sales for the Capital Expenditures/Sales regression; lagged values of log(Capital Stock/Employment), Sales Growth, and Capital Expenditures/Capital Stock for the R&D/Sales and Advertising/Sales regressions; and lagged values of Capital Expenditures/Sales and Sales Growth for the log(Capital Stock/Employment) regression. Sample sizes ranged from 120 to 130 for 1979–82 and from 130 to 150 for 1975–78.

[c] Based on fixed-effects rather than random effects.

Table 4 shows that there are generally strong negative effects of UBLS on investment in plant and equipment, R&D, and advertising in manufacturing, when no other control variables are included in random-effects regressions. Controlling for a firm's observable characteristics, the estimated direct effects of union coverage are roughly equal in magnitude but less precisely measured. The capital/labor ratio is higher in manufacturing firms with higher values of UBLS in all the random-effects specifications. After adjusting our estimates for measurement error bias, a 10 percent increase in UBLS decreases R&D/Sales by 3.5–5 percent in manufacturing and decreases Advertising/Sales by 5–7 percent. The first-differenced estimates using UBLS are imprecisely measured, and yield no significant results except for log(Capital Stock/Employment).

In nonmanufacturing there is some evidence that an increase in UBLS is associated with reduced investment in R&D and advertising and a higher capital/labor ratio, but these effects often become statistically insignificant after controlling for a firm's observable characteristics. After adjusting for measurement error bias, a 10 percent increase in UBLS decreases Advertising/Sales by 3.5–6.5 percent. The first-differenced regressions using UBLS in nonmanufacturing yield statistically insignificant results for all of our investment variables.

The regression models using UCPS again yield widely different point estimates, depending on the empirical specification. Random-effects estimates are generally statistically insignificant with large standard errors; hence, the OLS reported standard errors are again misleading. In nonmanufacturing the negative and significant effects in the first-differenced regressions are generally of the opposite sign of either the random-effects or the OLS estimates.

Unions and firm growth. Bronars and Deere (1993a,b) and Hirsch (1991b) have examined the relationship between union coverage and firm

TABLE 5

IMPACT OF A CHANGE IN UNION COVERAGE ON SALES AND EMPLOYMENT GROWTH
(t-statistics in parentheses)

	Measure of Growth and Time Period			
	Sales Growth		Employment Growth	
	1975–78	1979–82	1975–78	1979–82
Manufacturing[a]				
Mean (S.D.)	.0216	−.0145	.0119	−.0272
	(.0879)	(.1168)	(.0969)	(.1263)
Union Coverage Variable: UBLS				
Industry Effects	−.0668**	−.0298	−.0788**	−.0431
	(.0204)	(.0275)	(.0225)	(.0308)
Industry Effects with Controls	−.0325	−.0342	−.0526**	−.0563*
	(.0200)	(.0298)	(.0214)	(.0341)
First Difference		−.0699		−.1643**
		(.0499)		(.0582)
First Difference with Controls		−.0591		−.1327**
		(.0513)		(.0533)
Union Coverage Variable: UCPS				
OLS with Controls	.0009	−.0569	.0212	−.0400
	(.0338)	(.0464)	(.0362)	(.0544)
Industry Effects with Controls	−.0416	.1081	.0344	−.0850
	(.0828)	(.2946)	(.0944)	(.2305)
First Difference with Controls		.7208**		.0154
		(.3271)		(.3409)
Nonmanufacturing[b]				
Mean (S.D.)	.0467	.0382	.0156	.0168
	(.0933)	(.0942)	(.0819)	(.1107)
Union Coverage Variable: UBLS				
Industry Effects	−.0217	−.0639**	−.0360	−.0922**
	(.0286)	(.0269)	(.0278)	(.0263)
Industry Effects with Controls	−.0523*	−.1052**	−.0456	−.1178**
	(.0295)	(.0299)	(.0279)	(.0301)
First Difference		−.1595*		−.1365**
		(.0908)		(.0689)
First Difference with Controls		−.1553**		−.1491**
		(.0766)		(.0603)
Union Coverage Variable: UCPS				
OLS with Controls	−.0225	.1030*	−.0085	.1136*
	(.0497)	(.0546)	(.0409)	(.0627)
Industry Effects with Controls	.0329	.1089	−.0156	.0650
	(.1353)	(.1303)	(.0925)	(.1883)
First Difference with Controls		−.5058*		.5353**
		(.2862)		(.2691)

* Significant at .10 level; ** significant at .05 level.

[a] Controls include Four-Firm Concentration Ratio and lagged values of R&D/Sales, Capital Expenditures/Capital Stock, log(Capital Stock/Employment), Sales, log(Employment), and Tobin's q. Sample sizes ranged from 280 to 300 for 1979–82 and from 320 to 360 for 1975–78.

[b] Controls include lagged values of R&D/Sales, Capital Expenditures/Capital Stock, log(Capital Stock/Employment), Sales, log(Employment), and Tobin's q. Sample sizes ranged from 120 to 130 for 1979–82 and from 130 to 150 for 1975–78.

Table 5 shows that in both GLS and first-differenced regression models, there are significant negative effects of UBLS on employment growth in manufacturing and nonmanufacturing. Higher union coverage is also negatively associated with a firm's sales growth in both manufacturing and nonmanufacturing. In both the GLS and first-differenced models, these effects are generally insignificant in manufacturing and significant in nonmanufacturing. Adjusting for measurement error bias, a 10 percent increase in UBLS reduces sales growth by about .25–.5 percentage points in manufacturing, and .5–.75 percentage points in nonmanufacturing. Similarly, employment growth rates are reduced significantly in all specifications; declining by .5–1.0 percentage points in manufacturing, and .6–1.1 percentage points in nonmanufacturing in response to a 10 percent increase in UBLS.

In contrast, the GLS estimates using UCPS are in all cases insignificantly different from zero. The first-differenced regressions indicate that an increase in UCPS is associated with higher sales growth in manufacturing, and higher sales growth and lower employment growth in nonmanufacturing. However, the first-differenced estimates again typically differ in both magnitude and direction from either the GLS or the OLS estimates.

Unions and labor productivity. In an influential study, Brown and Medoff (1978) estimate the relative productivity of union and nonunion labor in U.S. manufacturing industries. Brown and Medoff hypothesize that union labor is more productive because union workers are likely to receive more training, have more job seniority, and have higher morale than are nonunion workers. Alternatively, unions may have a negative effect on productivity because of the work rules and restrictions they impose on management.

Brown and Medoff used weighted averages of industry/state data to estimate a Cobb-Douglas production function with capital and an aggregate labor input:

$$\begin{aligned}\ln(\mathrm{Y/L}) = {} & \gamma + \alpha\ln(\text{Capital Stock/Employment}) \\ & + (\beta + \alpha - 1)\ln \mathrm{L} + \tau \mathrm{U} + \epsilon,\end{aligned} \tag{8}$$

where Y is output, K is capital, L is employment, U is the ratio of union employment to total employment, α is the elasticity of output with respect to capital, and β is the elasticity of output with respect to labor. Brown and Medoff's estimate of τ is consistent with a 22–24 percent productivity differential between union and nonunion establishments.[21] In this section,

[21] Brown and Medoff estimate equation (4) using cross-state data (there are 29 states or state groups in the 1972 May CPS) for 20 two-digit SIC manufacturing industries. Output data for each state/industry cell are obtained from the value added data reported in the Census of Manufactures (COM).

we estimate (4) using firm-level data on output, employment, capital stock, and union coverage in our sample over the periods 1975–78 and 1979–82.

As an alternative to the restrictive Cobb-Douglas specification, we employ a nonparametric method to evaluate the relative productivity of unionized firms. Specifically, we use the production frontier approach outlined in Färe, Grosskopf, and Lovell (1985) and Burgess and Wilson (1992). First, we establish a nonparametric production frontier for each two-digit industry and time period in our sample. Firms with input/output vectors on the production frontier are assigned a technical efficiency measure of unity. Firms with input/output vectors in the interior of the production set are assigned technical efficiency measures between 0 and 1. The higher the measure of technical efficiency, the smaller the distance between the firm's input/output vector and the production frontier. We then model a firm's technical efficiency as a function of its unionization rate and other observable characteristics.

Table 6 presents the impact of an increase in union coverage on a firm's technical efficiency based on the nonparametric production frontier and the Cobb-Douglas production function in equation (4). We include lagged values of log(Capital Stock/Employment), Capital Expenditures/Capital Stock, R&D/Sales, Sales Growth, the logarithm of employment, and Four-Firm Concentration Ratio (in manufacturing) as additional control variables. We reject the hypothesis that δ_j is uncorrelated with X_{ij} and $UBLS_{ij}$ in one case, and thus present fixed-effects estimates for the Industry Effects specification of the frontier approach in nonmanufacturing from 1975–78.

The results in Table 6 indicate a significant positive impact of UBLS on productivity in manufacturing, using both GLS and first-differenced regressions, for both the nonparametric and Cobb-Douglas production functions. In nonmanufacturing there is evidence of a negative impact of UBLS on productivity when using fixed- or random-effects models. In both the manufacturing and nonmanufacturing sectors, the implied union/nonunion productivity differentials are surprisingly similar whether productivity is measured relative to the nonparametric or the Cobb-Douglas production function.

In contrast, industry average unionization rates from the CPS provide

Brown and Medoff use predicted wages from May CPS data on state-by-industry employment for the 1973–75 period to weight the "raw" data on the number of employee hours reported in the COM. See Brown and Medoff (1978) for a more complete discussion of this weighting procedure. All variables are measured per establishment in each state/industry cell.