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Union Coverage and Profitability Among U.S. Firms

by

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Abstract

This paper utilizes unique survey data on labor union coverage at the firm level to examine union effects on the profitability of 705 U.S. companies during the 1970s. Market value and earnings are estimated to be about 10-15 percent lower in an average unionized company than in a nonunion company, following extensive control for firm and industry characteristics. Deleterious union effects on firm profitability are sizable throughout the 1972-80 period, but vary considerably across industries. The relatively poor profit performance of unionized companies may help explain the recent decline in U.S. union membership.

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Introduction

Although the determination of firm and industry profitability has been the subject of much study, the impact of labor unions on company earnings and market value only recently has received careful attention. This neglect is surprising, given the importance of labor in production cost, clear-cut evidence of union-nonunion compensation differentials, and the marked decrease in U.S. union membership since the late 1970s. Recent studies examining the impact of unions on profitability have found unionism to be associated with significantly lower profitability, regardless of the profit measure, the unit of observation, time period, specification, or estimation method. A serious limitation of most previous studies, however, has been the absence of firm-level measures of the extent of union coverage.[1]

This study examines the earnings and market valuation of 705 publicly-traded U.S. companies over the 1972-80 period. Firm-specific union coverage data, based on a survey conducted by the author, are employed to measure carefully union-nonunion differences in profitability. The analysis includes use of alternative profit measures, examines the stability of union-nonunion differentials over time, identifies differences in union profit effects across broad industry categories, distinguishes between the effects of firm- and industry-level union coverage, and includes a particularly detailed set of firm and industry controls.

Union Impact on Earnings and Market Value

Union bargaining power allows rank-and-file to appropriate (tax) some share of a firm's returns from market power and special advantages, and the quasi-rents that make up the normal returns to fixed tangible and intangible capital. Unions have some degree of monopoly power owing to the costly

[1] Among the earliest and most frequently cited papers are Clark (1984), Ruback and Zimmerman (1984), Salinger (1984), Karier (1985), and Connolly et al. (1986). U.S. studies are evaluated by Becker and Olson (1987) and Addison and Hirsch (1989), while Metcalf (1988) provides a survey of British studies. Clark (1984) analyzes lines-of-business (LB) from the PIMS database using LB-specific union information, but is unable to examine the impact of unionization on firm market value. Ruback and Zimmerman (1984) and Bronars and Deere (1990) employ data on union representation elections to analyze their effects on stock prices, but elections generally cover a small portion of a firm's total workforce and a sample of companies with new union elections is unlikely to be representative of the population of existing union and nonunion companies. Becker and Olson (1989) construct a company-specific union measure based on union coverage among workers covered by pensions in 1977, and estimate union effects on investor returns and risk (beta) during the 1970s. Abowd (1989) and Bronars, Deere, and Tracy (1989) utilize a firm coverage measure calculated by matching data on union contract settlements to Compustat data on the parent companies. It is difficult to estimate accurately the extent of coverage with contract data, however, since the data on contracts are incomplete (e.g., they include only large contracts), matching is tedious (there are no common identifier codes), and coverage from multi-employer contracts cannot be easily allocated among firms.

substitution of a nonunion for union workforce (due, for example, to firm-specific skills) and the requirement of good faith bargaining mandated by U.S. labor law. Union bargaining power is constrained by the level and elasticity of labor demand, although bargaining outcomes need not be on the demand curve, and by the legal rules and enforcement surrounding the NLRB union representation process. Bargaining outcomes may be either noncooperative, as in the case of sequential wage and employment determination on the demand curve, or cooperative, wherein the joint value of the enterprise (the sum of market value and union rents) is maximized (Abowd, 1989).

Regardless of the nature of union settlements or the sources from which unions appropriate rents, union bargaining power implies that profitability will be lower among union than among similar nonunion firms.[2] Following all long run adjustments, “marginal” union and nonunion companies must have equivalent risk-adjusted profit rates. But given some level of new union organizing, unionized companies will have lower average profit rates at any point in time than their nonunion counterparts. Firms differ as well in the cost of deterring union organizing (or of decertifying existing unions). Therefore, “inframarginal” firms will have different profit rates; those with low deterrence costs will be nonunion and have higher profitability, while those with high deterrence costs will be unionized and have lower profitability (Lazear, 1983). Moreover, unions can appropriate rents for long periods if capital is long-lived and nontransferable, if returns (prior to the union tax) emanating from market power are sustainable, and if substitution of production from union to nonunion plants is costly.

The union tax on company earnings implies that accounting profit rates will be lower among union than among similar nonunion firms. Lower present and future earnings in turn lowers the market value of the firm until investors' expected (risk-adjusted) return on investment is equivalent across union and nonunion companies. Market value measures of profitability such as Tobin's q (the ratio of firm market

[2] An exception is the case where unions increase productivity sufficiently to offset cost increases (Clark, 1984). Extant evidence suggests that union productivity effects on average are close to zero, positive effects that exist may be in response to decreased profitability, and positive productivity effects are not generally sufficient to offset union wage increases (Addison and Hirsch, 1989).

value to the replacement cost of tangible assets) thus should be lower in union than in nonunion companies.[3] Competitive forces eventually must narrow the profitability gap between union and nonunion companies or there will be a decline in the size of the union sector. The decline results because of lower employment and investment among unionized companies, and because the cost of deterring union organizing in new plants (firms) is low relative to the cost of decertifying unions in existing organized plants (firms).

Data and Descriptive Evidence

Studies using industry union data fail to account for the considerable intraindustry variation in coverage and entangle to an unknown degree union and industry effects on profitability. In order to obtain information on company coverage, a survey of U.S. firms listed on the R&D Master File was conducted by the author during late 1987 and 1988. The R&D Master File is comprised of all publicly-traded manufacturing sector companies operating in 1976 that were included on Compustat tapes during 1976-78. In the union survey, contacted firms were asked to answer the following question for 1977 and 1987:

> *To the best of your knowledge, approximately what percentage of your corporation's total North American workforce is covered by collective bargaining agreements?*

Beginning with a list of 1904 firms, those for which addresses were available were mailed a questionnaire. The largest (approximately 1000) of these firms also were contacted by phone and/or mail and received a follow-up questionnaire if they did not initially respond. Responses were received from about one-third of the firms located. Efforts were made to ensure that responses regarding 1977 coverage corresponded to firms' extant corporate structure in 1977.

This paper utilizes data from a sample of 705 firms for which a measure of 1977 union coverage is available. A direct measure of 1977 union coverage was obtained from 578 firms in the survey. Union coverage was estimated for 20 additional firms reporting 1987 but not 1977 coverage by multiplying the 1987 figures by 1.21, based on the ratio of 1977-to-1987 coverage data among the 567 firms for which both

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[3] The rate of return on equity, measured by current earnings divided by the equity value of the firm, does not differ greatly between union and nonunion companies since union coverage lowers both the numerator and denominator. This ratio can differ with respect to union status, however, owing to differences in systematic risk (Becker and Olson, 1989), debt financing (Bronars and Deere, forthcoming), and life-cycle earnings patterns among companies with equivalent present values.

years of data were available (the simple correlation between 1977 and 1987 coverage is 0.87). Coverage figures for 125 additional firms are estimated based on data from an independent 1972 Conference Board Survey. Of the 723 companies for which union coverage data were available, 705 had sufficiently complete information on other variables to be included in subsequent empirical work.[4] Company union coverage in 1977 (UN) averaged 33 percent among this sample of firms (means and standard deviations by industry category are provided in Table 3). Substantial interindustry and intraindustry variation in coverage are found, supporting the proposition that measurement of unionization at the firm level is important. Apart from primary metals, all industries had at least one nonunion firm, and all industries except office, computer, and accounting equipment had at least one firm with coverage above 60 percent.

Our data set is constructed by matching the firm union coverage survey information with firm and industry data. Most firm-level information is from the R&D Master File (Cummins et al., 1985), comprised of all publicly-traded manufacturing sector companies operating in 1976, constructed from information provided in Compustat (which is based primarily on company 10-K reports), along with supplemental data on patents and R&D. The data set includes information on firm market value, accounting rates of return, sales, employment, gross and net plant, R&D investment and stock, and advertising expenditures. Data on company age are from Ward's Business Directory or Moody's Industrial Manual. Four-digit industry data on shipments are obtained from the Bureau of Industrial Economics tape consolidating data from the Annual Survey of Manufactures, while data on industry concentration (adjusted for imports and regional concentration) and import penetration are available for 1972 and 1977 in data assembled by Weiss and Pascoe (1986). Industry union coverage data for 1976-78 at an approximate 3-digit level is obtained from Kokkelenberg and Sockell (1985). All industry data are matched to the firm at the 2-, 3-, or 4-digit levels, based on Compustat's SIC-code variable designating the firm's principal industry in 1976. Complete data

[4] No significant response bias is found in the survey. The 1972 data were provided by David C. Hershfield, who developed the figures from data collected in a 1972 survey by the Conference Board (details available on request). Results are highly similar when the analysis is restricted to those companies for which a direct 1977 union measure is available. Coverage across all classes of workers, rather than just among production workers, was deemed the appropriate measure since it accounts for interfirm differences in the number of and coverage among nonproduction workers. Whereas the union coverage variable is defined for North American operations, all other firm data reflect worldwide operations. Subsequent conclusions should be qualified in light of this limitation.

are available for 705 firms in 1976, while missing observations increase as one moves back to 1972 and forward to 1980.

Table 1 presents descriptive evidence on firms' rate of return and market valuation of assets over the 1972-80 period, cross-tabulated by union status. Total sample sizes are less than 9 times 705 owing to missing data. “Nonunion” is defined as firms with no union coverage, “low” union as covered firms with less than 30% coverage, “medium” union as firms with coverage of at least 30 but less than 60%, and “high” union as firms with coverage 60% or higher. Tobin’s *q*, measuring the ratio of firm market value to the replacement cost of company assets, decreases substantially as one moves from the nonunion to high union categories. A similar although less marked pattern is found for the rate of return to capital, measured by gross net earnings divided by the gross capital stock. Average *q* among unionized companies during the 1972-80 period is 43% lower than among the nonunion companies; the corresponding figure for ð$_k$ is 27%.

Union Effects on Profitability: Specification and Results

Profitability equations using the natural logarithm of Tobin’s *q* (ln(*q*)) and the rate of return on capital (ð$_k$) as dependent variables are estimated.[5] The variable ð$_k$ measures realized annual earnings relative to an asset base, whereas the market value measure q is forward-looking, reflects expected performance over time rather than accounting performance for a single period, measures risk-adjusted returns, and is less likely to be affected by differences in accounting procedures across firms. Company earnings (present and future) reflect the difference between revenues and costs; thus, ln(*q*) and ð$_k$ are determined by factors affecting prices, production, and costs. A simple form of a profit equation is:

(1) Ð$_{it}$ = á + Óâ$_j X_{jit}$ + Óã$_k Z_{kit}$ + Óè$_d IND_{di}$ + Óä$_m YEAR_{mt}$ + ø$UN_i + e_{it}$,

where Ð$_{it}$ is the profitability of firm *i* in year *t*, measured alternatively by ln(*q*) and ð$_k$, á is an intercept, *X* represents *j* firm-specific variables affecting revenues and costs and the â$_j$ are the attaching coefficients, *Z* represents *k* industry variables and the ã$_k$ are the attaching coefficients, *IND* represents *d* industry dummies

[5] Use of the Box-Cox transformation allows us to reject soundly the linear specification with *q* as the dependent variable, while providing support for the semi-logarithmic specification (Hirsch and Seaks, 1990).

and $è_d$ are the corresponding coefficients, *YEAR* represents *m* year dummies with $ä_m$ the coefficients, UN_i measures 1977 union coverage in firm *i* and ø is its coefficient, and e_{it} is the error term.[6]

Estimation of (1) results in an error term that is serially correlated across years within firms, since unmeasured firm-specific profit determinants in year *t* are likely to be positively correlated with such determinants in year *t* - 1. Although coefficient estimates from (1) need not be biased, standard errors will be biased downward. In order to purge the model of positively correlated errors, a two-step model is estimated. In the first step, profitability is regressed on all *time-varying* profit determinants, plus 705 firm dummy variables (the intercept is suppressed). In the second-step, the coefficients on the firm dummies, measuring firm fixed effects "averaged" over the 1972-80 period, are used as the dependent variable in a second-step weighted least squares regression, with all time-invariant variables, including firm means of variables in *X* and *Z*, as regressors. Specifically, the first-step OLS regression is:

(2) $Ð_{it} = Óâ'_j X'_{jit} + Óã'_k Z'_{kit} + Óä'_m YEAR_{mt} + Óö_i FIRM_i + e'_{it}$,

where $FIRM_i$ represents 705 firm dummies with coefficients $ö_i$, *X'* is equivalent to *X* except for exclusion of company age (*AGE* is a linear combination of the firm and year dummies), and *Z'* is equivalent to *Z* except for exclusion of the time-invariant industry union variable, *I-UN*.

A second-step weighted least squares (*WLS*) regression is then estimated (*n* = 705) with the firm fixed effects $ö_i$ as the dependent variable, using the inverse of the standard errors from the firm dummy coefficients in equation (2) as weights, and right-hand variables being the time-invariant variables: firm union coverage, industry union density, industry dummies (*IND*), company age in 1977, and the firm-specific means of variables in *X'* and *Z'*. That is, we estimate:

(3) $ö_i = á' + øUN_i + ôI\text{-}UN_i + æAGE_i + Óè'_d IND_{di} + ÓÃ_j \overline{X}'_{ji} + ÓÈ_k \overline{Z}'_{ki} + v_{it}$.

[6] Regression results from specifications including interactions between *UN* and variables in *X* and *Z* are provided in Hirsch (1989; 1990). Such specifications allow inferences to be made about the *sources* from which unions appropriate rents. Analysis in Hirsch (1990) provides no support for the hypothesis that monopoly rents associated with industry concentration and firm market share provide sources for union gains (see Clark, 1984; Salinger, 1984; Karier, 1985; and Hirsch and Connolly, 1987). Evidence provided in Hirsch (1989) suggests that unions appropriate quasi-rents associated with fixed long-lived tangible and intangible capital and engage in rent sharing of profits (and losses) associated with disequilibria and changes in demand.

Coefficients from the first-step measure within-firm effects, while between-firm effects are captured by the firm dummies. In the second step, between-firm effects are explained by differences in firm and industry unionization, age, industry, and firm means of variables in *X*' and *Z*'. Interpretation of the second-step union coefficient ø' in (3) is analogous to interpretation of ø from (1).

Firm-level variables included in *X*' are firm size (*L*), capital intensity (*K/L*), R&D intensity (*R&D/S*), advertising intensity (*ADV/S*), and logarithmic sales growth (*GROWTH*) (firm-years for which *GROWTH* is less than -1.0 or greater than 1.0 are excluded from the sample). Industry-level variables in *Z*' are industry sales growth (*I-GROWTH*), concentration (*I-CR*), and import penetration (*I-IMPORT*). Included in the second-step regression are firm means over time of the variables in *X*' and *Z*', firm union coverage (*UN*), industry union density (*I-UN*), company age (*AGE*), and industry dummies (*IND*). A data appendix provides definitions for all variables.

It is not possible to present complete regression results from OLS estimation of equations (1) or (2), based on the 1972-80 sample of firm-years, nor discuss the relationship of firm profitability with firm and industry variables other than unionization (full results are available on request). It is worth noting, however, that union profit effect estimates ø from OLS estimation of (1) are larger (in absolute value) than are estimates from (3), and estimates are substantially larger when we do not account for firm-specific growth and industry control variables.

Table 2 provides the second-step WLS regression results from equation (3), where firm dummy coefficients from the first-step ln(*q*) and $ð_k$ regressions are the dependent variables. As a measure of firm-level union coverage, we use not only *UN*, measuring the proportion of a firm's workforce covered by a collective bargaining agreement, but also dummies representing low, medium, and high union coverage, with nonunion the excluded category. The results indicate clearly that unionized companies have significantly lower market valuation of assets and profit rates than similar nonunion companies. Coefficient estimates from specifications (1) and (1') indicate that *q* and $ð_k$ are lower by an average 12.4% and 9.2%, respectively, in an average unionized firm (with *UN* = 0.43) than in a similar nonunion firm.[7] The coefficients on the union coverage dummies indicate that union profitability effects vary continuously with the extent of firm coverage. We find union-nonunion differentials in q of -4.5%, -14.7%, and -19.5% for companies with low, medium, and high union coverage, respectively, and differentials in $ð_k$ of -8.7%, -12.6%, and -18.5%.

It is worth noting that *industry* union coverage, *I-UN*, has no significant independent effect on profitability, after accounting for firm-level coverage. The effect of industry coverage is indeterminate, a priori. On the one hand, industry coverage might reduce industry output and raise price, thus increasing company earnings (holding constant firm coverage). On the other hand, industry coverage may proxy union power and the threat of union organizing, thus raising current and future wage costs for union *and* nonunion firms.[8]

A potential problem in the preceding analysis is that of simultaneity between union coverage and profitability, wherein the level of firm coverage is influenced by firm profitability (Voos and Mishel, 1986). On the one hand, union organizing is likely to be more successful among firms with larger monopoly profits and quasi-rents from which unions can appropriate gains. If higher profits lead to greater union coverage, the negative effect of unionization on profitability may be *understated* using OLS estimation. On the other hand, workers may be more inclined to vote for union representation in low-profit firms that treat workers opportunistically. Because we are measuring existing levels of firm coverage rather than current union

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[7] As noted previously, unadjusted profit differentials in *q* and $ð_k$ based on data from Table 1 are 43% and 27%, respectively. Letting ø' represent the estimated coefficient on union coverage, the average percentage effect of union coverage on profitability is calculated by [exp(.43ø')-1]100 for *q* and by (.43ø'/.103)100 for $ð_k$, where .103 is mean $ð_k$ among *nonunion* companies. For comparison with previous estimates in the literature, see Addison and Hirsch (1989, Table 1). Reported WLS estimates are highly similar to unweighted second-step OLS estimates. Union coefficient estimates also are similar when 105 detailed industry dummies are substituted for the 18 2-digit dummies in specification 1/1', the *UN* coefficients ($|t|$) changing to -0.301 (4.13) and -0.019 (3.44), respectively. Because industry variables and dummies are matched imperfectly to companies operating across multiple industry categories, industry effects may be captured in part by firm-level variables.

[8] *I-UN* coefficients are highly sensitive to specification. When the variables *UN*, *I-UN*, and the interaction of *UN* and *I-UN* are included, the coefficient on *I-UN* is positive and that on *UN* · I-*UN* is negative. Neither, however, is statistically significant. Inclusion of *I-UN* but not *UN* results in union coefficients highly sensitive to specification and estimation method, supporting the proposition that industry union density is a poor proxy for firm level coverage. The simple correlation between *UN* and *I-UN* is 0.467.

organizing or changes in coverage, however, simultaneity may be less serious a problem than commonly thought. That is, existing levels of coverage have been determined well in the past and may be only weakly affected by current firm profitability. Attempts here to deal with simultaneity, while not completely satisfactory, suggest that the union profit effect is more deleterious than indicated by our reported estimates.

A related concern is that despite the inclusion of detailed control variables in our regressions, there may be omitted determinants of profitability correlated with the union coverage variable. A potential method to account for omitted variable bias is the use of a fixed-effects or difference model, wherein changes in profitability are estimated as a function of changes in union coverage. But such estimation is not possible here, since union coverage is measured only for 1977.[9] A final potential bias is that of selectivity. Companies realizing lower rates of profit may contract in size and be less likely to survive over time. Therefore, our sample of firms may not include companies for which the union impact on earnings has been most deleterious. Estimates of the union tax on profits thus may provide biased estimates of unionism's true impact on a representative firm.[10]

Union Profitability Effects by Year and Industry

Neither intertemporal nor interindustry differences in union profit effects have been examined previously. Union-nonunion differences in profitability across broad industry categories are estimated by:

(4) $$\varphi_i = \alpha'' + \Sigma\psi_{d+1} UN \cdot IND_{d+1,i} + \tau' I\text{-}UN + \Sigma\theta''_d IND_{di} + \zeta AGE_i + \Sigma\Gamma'_j \overline{X}'_{ji} + \Sigma\Theta'_k \overline{Z}'_{ki} + v'_i$$

where, as in Table 2, the φ_i are the coefficients on the firm dummies in the first-step regression, while ψ_{d+1} are the coefficient estimates on the interactions between *UN* and the 19 industry dummies from a second-step WLS regression. Table 3 presents these union coefficients by industry category for both the ln(q) and π_k regressions.

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[9] The author is currently conducting a study relating changes in firm profitability to changes in coverage between 1977 and 1987. The estimation of fixed effects models, of course, introduces its own set of econometric problems.

[10] The issue of selectivity is rather complex. Relatively small and young establishments are more likely to fail (Dunne, Roberts, and Samuelson, 1989) and *less* likely to be unionized. Evidence in Dunne and Macpherson (1991), examining individual plant data aggregated to the 3-digit industry by size level, suggests that employment expansions are inversely related to industry union density, employment contractions are positively related to density, while business "deaths" are not significantly related to unionization.

Union bargaining power, labor relations, and union effects on wages and productivity differ substantially across industries. It is not surprising, therefore, that union-nonunion profit differentials vary across industries. Most of the estimates cannot be made with precision, given the relatively small number of companies within each industry category. No evidence is found for a sizable and statistically significant *positive* relationship between coverage and profitability within industry categories, although several of the coefficients are positive. The most clear-cut evidence of negative union effects is found in foods, chemicals, rubber and plastics, primary metals, non-electric machinery, and lumber and paper.[11] It is difficult, however, to discern clear-cut patterns in the estimated industry effects. Factors that might explain or be correlated with differences in union power and wage premiums – industry concentration, import penetration, firm capital intensity, firm and industry growth, and the like – are accounted for in the regressions. Providing explanations for interindustry differences in union profit effects thus remains a challenge for future research.

Estimates of union-nonunion profitability differences by year, presented in Table 4, are based on the union coefficients from OLS ln(q) and $\eth_k$ equations, estimated separately by year (identical to equation 1, minus the year dummies). Corresponding estimates from the pooled model (equation 1) with year dummies included are provided on the bottom line. Union coefficients are large (in absolute value) and statistically significant throughout the 1972-80 period. Union effects appear to be particularly detrimental during 1972-73 (and, to a lesser extent, 1979-80), years in which average q and $\eth_k$ are relatively high.[12] This finding supports the thesis that unions engage in rent sharing with respect to demand shifts and disequilibrium returns (Hirsch, 1989) and that some risk is shifted from shareholders to labor (Becker and Olson, 1989).

[11] Standard errors of the industry estimates are much lower when pooled time-series/cross-section regressions are estimated separately by industry category. Industry results display some sensitivity to the measurement of union coverage (e.g., use of dummies rather than a continuous coverage variable) since the level and dispersion of coverage vary substantially among industry categories. Hirsch (1991) reports an inverse relationship across industry categories between estimates of union effects on profitability and productivity.

[12] For related estimates beginning in 1968, see Hirsch (1991). F-tests comparing the annual regressions and the pooled regression with year dummies (bottom row) yield $F(232,5554)=1.694$ for the ln(q) equations and 2.031 for the $\eth_k$ equations. The hypothesis of equal slope coefficients over time is rejected using the classical criterion, but cannot be rejected using an alternative criterion that adjusts the critical value *upward* as sample size increases (Leamer, 1978, p. 114). Measurement error in *UN*, which measures union coverage in 1977, should bias the *UN* coefficient toward zero as we move away from 1977. Regressions using a balanced panel of firms (n = 480) exhibit a similar intertemporal pattern, but lower coefficient estimates.

Conclusions

This paper utilizes a company-specific measure of collective bargaining coverage to examine the relationship between firm profitability and unionization. The results provide broad support for the hypothesis that unions appropriate a share of the returns accruing from profit-enhancing firm and industry characteristics. Union coverage at the firm level exhibits a strong negative relationship with company earnings and market value, even after controlling in detail for firm and industry characteristics. Union effects on profitability remained sizable throughout the 1972-80 period. Differences in the union profit effect across industry categories are substantial, however, and do not lend themselves to simple explanation. The evidence reported here confirms that unionization has an important influence on company earnings and market value and reinforces the conclusion that unionism should not be ignored in empirical studies of profitability (Karier, 1985). Equally important, the results show that there exists substantial intraindustry variability in unionization and that firm-level measures of collective bargaining coverage are highly preferable to more commonly-used industry measures of union density.

The poor profit performance of unionized companies during the 1970s may also provide an important explanation for the recent marked decline in union membership. As shown by Linneman, Wachter, and Carter (forthcoming), employment declines have been concentrated among union workers, while nonunion employment has expanded even in highly unionized industries. Changes in industry demand explain a small proportion of the total decline in private sector unionism. Although the evidence presented here is consistent with the thesis that declines in union membership and coverage have been in response to the continuing poor profit performance of unionized companies throughout the 1970s, future research is needed to examine such a proposition more directly.

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Table 1
Mean Profitability by Union Category, 1972-80

Variable	Nonunion	Low Union	Medium Union	High Union	All Firms
UN	0.000	0.138	0.451	0.709	0.334
Q	1.493	1.023	0.827	0.681	0.990
$\eth_k$	0.103	0.084	0.073	0.068	0.081
N	1305	1473	1630	1416	5824

Nonunion (*UN* = 0); Low Union (0 < *UN* .30); Medium Union (.30 < *UN* < .60); High Union (*UN* .60).

Table 2
Second-Step WLS Profitability Regression Results

	ln(q)		ð$_k$	
	(1)	(2)	(1')	(2')
UN	-0.309	–	-0.022	–
	(4.45)		(4.32)	
UN-LOW	–	-0.046	–	-0.009
		(0.92)		(2.34)
UN-MED	–	-0.159	–	-0.013
		(3.07)		(3.43)
UN-HIGH	–	-0.217	–	-0.019
		(3.87)		(4.46)
I-UN	-0.211	-0.223	0.016	0.014
	(1.04)	(1.10)	(1.05)	(0.94)
AGE/100	-0.142	-0.140	-0.014	-0.013
	(2.57)	(2.50)	(3.30)	(3.06)
IND	yes	yes	yes	yes
$\bar{R}^2$	0.471	0.470	0.326	0.326

n = 705. $|t|$ in parentheses. Dependent variables are the firm dummy coefficients, ö$_i$, from first-step regressions (equation 2) of ln(q) and ð$_k$. Included in the first-step regressions are the time-varying variables *L*, *K/L*, *R&D/S*, *ADV/S*, *GROWTH*, *I-GROWTH*, *I-CR*, *I-IMPORT*, and year dummies. The second-step GLS regressions are weighted by the inverse of the standard errors of the firm dummy coefficients and, in addition to the variables shown above, include the firm means of the time-varying variables listed above. *UN* is the proportion of a firm's workforce covered by a collective bargaining agreement. *UN-LOW* = 1 if (0 *UN* .30); *UN-MED* = 1 if (.30 < *UN* < .60); and *UN-HIGH* = 1 if (*UN* .60). *I-UN* is the proportion of workers who are union members in the firm's principal industry; *AGE* is years since incorporation; and *IND* is a group of 18 approximate 2-digit industry dummies.

Table 3
Union Coverage and Profitability Effects by Industry Category, Second-Step WLS Estimates

				ln(q) equations		ð$_k$ equations	
Industry Group	n	mean *UN*	s.d. *UN*	*UN*	$\|t\|$	*UN*	$\|t\|$
All Manufacturing	705	0.332	(0.28)	-0.309	(4.45)	-0.022	(4.32)
Food & kindred products	71	0.419	(0.27)	-0.323	(1.71)	-0.032	(2.27)
Textiles & apparel	39	0.233	(0.30)	0.279	(1.37)	-0.018	(1.16)
Chemicals, excluding drugs	43	0.286	(0.20)	-1.157	(3.57)	-0.039	(1.60)
Drugs & medical instruments	38	0.138	(0.18)	-0.001	(0.00)	0.023	(0.74)
Petroleum refining	29	0.281	(0.19)	0.016	(0.03)	-0.055	(1.53)
Rubber & miscellaneous plastics	25	0.381	(0.25)	-0.692	(2.50)	-0.030	(1.42)
Stone, clay, & glass	24	0.444	(0.26)	-0.346	(0.93)	-0.044	(1.59)
Primary metals	41	0.613	(0.22)	-0.760	(2.79)	-0.043	(2.11)
Fabricated metal products	39	0.316	(0.28)	0.133	(0.60)	-0.005	(0.31)
Engines, farm, & const. equip.	25	0.384	(0.23)	0.025	(0.06)	-0.000	(0.01)
Office, computer, & acct. equip.	22	0.043	(0.07)	-0.106	(0.07)	0.207	(1.72)
Other machinery, not electric	48	0.369	(0.28)	-0.588	(3.14)	-0.042	(2.97)
Electrical equipment & supplies	46	0.083	(0.17)	-0.272	(0.61)	0.033	(0.97)
Communication equipment	27	0.456	(0.24)	-0.463	(1.13)	-0.004	(0.14)
Motor vehicle & trans. equip.	43	0.523	(0.26)	-0.243	(0.77)	0.011	(0.45)
Aircraft & aerospace	11	0.305	(0.23)	-0.072	(0.13)	-0.007	(0.16)
Professional & scientific equip.	31	0.116	(0.19)	-0.662	(1.66)	-0.010	(0.32)
Lumber, wood, & paper	55	0.387	(0.30)	-0.484	(2.39)	-0.024	(1.59)
Misc. manuf. & conglomerates	48	0.343	(0.25)	-0.112	(0.49)	-0.004	(0.26)

See note to Table 2. Coefficient estimates are from interactions between *UN* and industry dummies, corresponding to vector $ø_{d+1}$ presented in equation (4) in text. The All Manufacturing results are from Table 2, specifications 1 and 1'.

Table 4
Union Profitability Effects by Year, 1972-80

				ln(*q*) equations		ð$_k$ equations	
Year	*n*	$\bar{q}$	$\bar{p}_k$	*UN* coeff.	$\lvert t \rvert$	*UN* coeff	$\lvert t \rvert$
1972	627	1.88	0.089	-0.578	(5.01)	-0.037	(3.95)
1973	646	1.25	0.080	-0.446	(3.93)	-0.032	(4.110
1974	670	0.71	0.061	-0.184	(2.01)	-0.015	(1.95)
1975	687	0.80	0.079	-0.333	(3.60)	-0.026	(4.30)
1976	705	0.90	0.084	-0.273	(3.30)	-0.023	(2.98)
1977	682	0.84	0.087	-0.347	(4.63)	-0.023	(3.31)
1978	652	0.81	0.086	-0.297	(4.04)	-0.022	(2.89)
1979	589	0.84	0.084	-0.470	(5.78)	-0.034	(4.65)
1980	566	0.92	0.080	-0.484	(5.31)	-0.033	(5.35)
1972-80	5824	0.99	0.081	-0.377	(12.21)	-0.028	(10.89)

Annual regressions include *L*, *K/L*, *R&D-STK/S*, *ADV/S*, *GROWTH*, *AGE*, *I-GROWTH*, *I-CR*, *I-IMPORT*, *I-UN*, and *IND* dummies. The pooled 1972-80 regression (equation 1 in text) includes these controls and year dummies.

Data Appendix: Regression Variable Definitions

ln(*q*)	Log of Tobin's *q* – market value of firm (value of preferred and common stock plus short-term debt plus age-adjusted long-term debt minus net short-term assets), divided by replacement cost of tangible assets, proxied by value of net inflation-adjusted capital stock (plant and inventories) (Cummins et al., 1985).
ð$_k$	Gross rate of return on capital; gross cash flows (income plus depreciation plus interest income minus inventory and imputed income adjustments), divided by the gross capital stock adjusted for inflation (Cummins et al., 1985).
UN	Proportion of firm's North American workforce covered by a collective bargaining agreement in 1977.
L	Employment, in thousands.
K/L	Capital stock per employee, in thousands of 1972 dollars. Measured by net inflation-adjusted capital stock (deflated by GNP investment implicit price deflator), divided by employment.
R&D/S	R&D stock in 1972 dollars, divided by sales in 1972 dollars; calculated based on R&D expenditures and assumed 15 percent depreciation rate (Body and Jaffe, no date).
ADV/S	Annual advertising expenses, divided by sales.
GROWTH	Annualized logarithmic growth rate in firm sales between years *t* and *t* - 2; sales deflated by industry-specific price indices.
AGE	Company age in 1977, measured by 1977 minus year of incorporation.
I-GROWTH	Annualized logarithmic growth rate in industry shipments between years *t* and *t* - 4 in firm's primary reported industry; shipments deflated by industry-specific price indices.
I-CR	Four-firm concentration ratio in firm's primary reported industry, adjusted for regional markets and imports, available for 1972 and 1977. Post-1977 data assigned 1977 values; 1973-1976 data assigned values based on linear interpolation.
I-IMPORT	Share of imports in domestic sales in firm's primary reported industry, defined as 100[*IMPORTS*/(*SHIPMENTS* + *IMPORTS* - *EXPORTS*)], available for 1972 and 1977. Post-1977 data assigned 1977 values; 1973-76 data interpolated.
I-UN	Proportion of eligible workers who are union members in firm's primary 2- or 3-digit industry during 1976-78.