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Shareholder Risk and Returns in Union and Nonunion Firms

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Abstract

This study examines shareholder risk and rates of return in union and nonunion companies in 1973-87. Shareholder risk declined with the extent of union coverage in the 1970s, and returns were lower among highly unionized companies than among other companies during the late 1970s and early 1980s. Union-nonunion differences in risk were small and insignificant by the mid-1980s, however, and there was no systematic relationship between union coverage and shareholder returns in the mid-1970s or mid-1980s. Finally, firms with relatively low rates of return to investors in 1977-82 tended to experience larger than average declines in firm-level union coverage between 1977 and 1987. As a partial explanation for these findings, the authors posit a relationship between shareholder risk and the differential use of COLAs across industries and time.

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Recent research has focused considerable attention on union-nonunion differences in profitability, market value, and other dimensions of firm performance or behavior. Becker and Olson (1989) provided a unique analysis of differences in shareholder risk and returns in union and nonunion companies. They found that during most of the 1970s, shareholder risk was systematically lower in union than in nonunion companies, and shareholders in more highly unionized companies realized lower rates of return than did owners of nonunion companies. Becker and Olson argued that lower shareholder risk in highly unionized companies is the result of risk shifting from owners to labor, and they suggested that lower investor returns during the 1970s led to increased pressure on management to limit the extent of union representation in the 1980s.

This paper provides new evidence on union-nonunion differences in shareholder risk and returns using a sample of companies and a measure of firm unionization different from those used by Becker and Olson. Among the questions addressed are the following: 1) Was shareholder risk lower in union than in nonunion companies during the 1970s, and, if so, did that relationship continue into the 1980s? 2) Did lower rates of return to investment in union companies continue during the 1980s? 3) Do differences in shareholder risk result because of risk shifting from shareholders to labor, or are there other reasons for this relationship? 4) How does stock market performance affect subsequent unionization?

Below, we briefly review the Becker and Olson study. We then present evidence on the relationship of firm union coverage with shareholder risk and rates of return for some 400 publicly traded manufacturing sector companies during the 1973-87 period.

Background

Becker and Olson (1989) examined data on about 1,200 publicly traded U.S. firms, approximately two-thirds of which were in the manufacturing sector. An important contribution of their paper is the use of a firm-level union coverage measure, constructed for the year 1977 based on pension coverage information collected from Department of Labor surveys. Union coverage is approximated by

the number of workers enrolled in union-negotiated pension plans, divided by enrollment in all firm pension plans.¹

Becker and Olson estimated shareholder risk or “beta” by the “market model” relating firm to market rates of return, based on weekly shareholder and return data for the years 1970-81, with controls for industry and with and without controls for firm size and leverage (the ratio of debt to equity). They found no significant union-nonunion differences in beta during 1970-72, but found a significant negative relationship between union coverage and beta in eight of the nine years during the 1973-81 period. The value of their union coverage coefficient for the entire period was $-.102$, indicating a beta about $.04$ lower for a company with, say, 40% coverage than for a nonunion company in the same industry (mean beta is close to one).

Why might there be lower shareholder risk for owners of union firms than for owners of nonunion firms? Becker and Olson posited that there is risk shifting from shareholders to labor in union firms. They argued that monitoring and rewarding (penalizing) worker effort (shirking) is more difficult in unionized than in nonunion workplaces. An alternative way to sustain or increase workers’ productivity is to link their wages to firm profitability, thus making their interests more compatible with those of shareholders. Union workers bear greater risk in the form of wage or employment fluctuations, but are compensated through receipt of a higher wage. An implication of Becker and Olson’s work is that some portion of the union wage premium represents a compensating differential for bearing risk. Implicit in their model is that there is no systematic difference in total enterprise (shareholder plus worker) risk in union and nonunion companies; rather, a larger proportion of total risk is shifted from shareholders to labor in union companies.

Ex ante rates of return to investors should not differ between union and nonunion companies, other than from premiums associated with higher systematic risk (higher betas).² Becker and Olson,

¹ Becker and Olson obtain an estimate of 32.4 percent coverage among manufacturing companies (derived from Table 1 “simple averages” weighted by number of companies), as compared to 33.4 percent coverage in 1977 reported by Hirsch (1991b) from a firm survey, and 36.8 percent in Kokkelenberg and Sockell (1985) based on individual data from the Current Population Survey (the CPS figure is for membership rather than contract coverage, and for eligible workers rather than all workers; these differences approximately offset each other).

however, found substantially lower ex post rates of return for unionized companies than for nonunion companies (controlling only for beta) during much of the 1970s. Following the inclusion of industry dummies in the rate of return equation, estimated union-nonunion differences were small and variable in sign, the exceptions being significant negative union coefficients in 1971, 1972, and 1980. But the cumulative 1971-81 union-nonunion differential was nontrivial, indicating a deterioration of about 10% in the relative value of equity for companies that were 40% organized. It was this poor performance, Becker and Olson reasoned, that strengthened pressure on management to combat union bargaining and organizing activities.

Data and Analysis

The data set employed here consists of an unbalanced panel of approximately 400 U.S. companies for the years 1973-87. Data are obtained from several sources. Firm data on rates of return, employment, capital stocks, depreciation, debt, equity, and principal industry are obtained from the Manufacturing Sector Master File (Hall, 1990), a panel of firm-year observations on publicly traded manufacturing sector companies, comprising data compiled from Compustat and other sources. Firm-level union coverage data for the years 1977 and 1987 were collected in a 1987 phone and mail survey of firms conducted by Hirsch (1991a), which covered a subset of the firms contained in the Manufacturing Sector File. Stock market betas are published in the *Value Line Investment Survey*. Company age is based on the year of incorporation, as listed in *Ward's Business Directory* or *Moody's Industrial Manual*. Industry concentration ratios, adjusted for trade and the regional distribution of sales, are constructed based on calculations for 1972 and 1977 by Weiss and Pascoe (1986) and on 1982 Census of Manufactures figures. Concentration figures are matched to firms' most detailed primary SIC code listed in Compustat (at the 2-, 3-, or 4-digit level). The final data set consists of firm-year observations for which all necessary information is available.

² Equivalent stock market rates of return are not inconsistent with evidence (e.g., Hirsch, 1991b; Becker and Olson, 1992) that union firms have lower profitability (current earnings divided by sales or assets) and stock market valuation of assets (measured by Tobin's q or "excess value" divided by sales).

Shareholder Risk Results

Shareholder risk is calculated from the “market model” relating firm rates of return to the market rate of return. *Value Line* presents information on b_i based on the following regression(s) using five years of weekly data:

$$(1) R_{it} = a_i + b_i R_{mt} + e_{it}$$

R_{it} is the rate of return to firm i in period t , measured by dividends plus the change in the stock price between period t and $t-1$, all divided by the price in period $t-1$; R_{mt} measures the market rate of return for a broad-based market portfolio (the New York Stock Exchange Composite Index) during the same time period; a_i are firm-specific intercepts; and e_{it} are error terms with zero means and constant variances (the residuals measure non-systematic firm risk or the variability that is not correlated with market movements). The parameters b_i represent firm-specific “betas” measuring the degree of “systematic” risk to shareholders. A $b_i=1.0$ implies that firm rates of return, on average, exhibit changes equivalent to changes in the market rate; $b_i<1.0$ implies less than equal changes in firm relative to market rates (lower shareholder risk); and $b_i>1.0$ implies changes in firm rates larger than market rate changes (higher shareholder risk). All betas in our sample are positive, although $b_i<0$ is theoretically possible.

Firm betas then become the dependent variable in the following regression, estimated annually for the period 1973-87:

$$(2) b_{iy} = \alpha_y + \Psi_y \text{COV}_{iy} + \sum \beta_{yk} X_{iyk} + e_{iy},$$

where b_{iy} are betas for firm i in year y , COV_{iy} measure firm union coverage in year y , $\sum \beta_{yk} X_{iyk}$ represents control variables and their coefficients indexed by k , α_y are annual intercepts, Ψ_y are estimates of annual union coverage effects on b , and e_{iy} are random error terms with zero means and constant variances by year.³ Variables included in X are firm size as measured by the log of employment (SIZE), the firm’s

³ *Value Line* betas are calculated over a five-year period and do not measure precisely year-to-year changes in beta. In statistical work shown in the paper, we use a measure of beta *centered* on each year, rather than ending in the year. For example, for the year 1987 we use an end of year 1989 *Value Line* measure, based on the period 1985-89. We also estimate the models with two alternative measures of beta. The first alternative is to use a measure of beta ending in each year; for example, the year 1987 based on data for 1983-87. Not surprisingly, we obtain similar results to those shown in the paper, but intertemporal changes show up with a two-year lag. The second alternative is to estimate a *marginal* or annual beta, based on changes in the five-year beta. Letting b_0 and b_1 be the five-year

capital-labor ratio (K/L), the estimated age of the capital stock as measured by the log of accumulated depreciation divided by annual depreciation (CAPAGE), company age measured by years since incorporation (AGE), leverage as measured by the debt-equity ratio (LEV), industry concentration adjusted for regional and international market shares (CR), and 19 two-digit industry dummies (IND). A data appendix provides fuller detail on the variables.

Regression results from the beta equations are reported in Table 1, where coefficients Ψ_y , on the union coverage variable are presented from equations with and without firm/industry controls. Standard errors have been corrected for heteroskedasticity using the method proposed by White (1980); results are highly similar to those obtained from OLS without the correction.⁴

The regression results strongly confirm the previous findings by Becker and Olson. The negative union coefficients are of substantial magnitude and highly significant during the years 1973-81 (1981 was the last year in the Becker-Olson study). For example, a union coverage coefficient of -0.160, the unweighted average of coefficients from regression (2) (i.e., with controls) for 1973-81, implies that a company with union coverage of 40% (the approximate mean coverage among unionized companies in our sample) would have a beta .064 lower than an otherwise similar nonunion company. Although statistically significant and nontrivial in magnitude, the union-nonunion difference in beta is only about one-third the size of the standard deviation of beta (equal to .210). Our coefficients for the 1973-81 period, however, are larger (in absolute value) than those in the Becker-Olson study for the same period, and our standard errors are smaller. These differences may result in part because of differences in the sample of firms. They also may result from less measurement error in our union coverage variable than in Becker and Olson's, since random measurement error will bias toward zero the union coefficient while increasing the standard errors.

betas ending in years 0 and 1, respectively, we calculate the marginal beta by $b_m = [b_1 - (4/5)b_0]5$, based on the approximation $b_1 \cong 4/5b_0 + 1/5b_m$. Use of b_m produces a pattern of results highly similar to that shown in the paper, but coefficients and significance levels exhibit greater year-to-year variability.

⁴ The use of a regression parameter (e.g., beta) as a dependent variable can produce heteroskedastic errors. A common approach is the use of weighted least squares, with the inverse of the coefficient error variance as a weight (Saxonhouse, 1976). Because Value Line does not publish the standard errors associated with its estimates of beta, we did not pursue this approach.

Although our results confirm the finding of lower shareholder risk among union companies than among nonunion companies through 1981 (the final year of the Becker-Olson study), this relationship does not continue into the late 1980s. The magnitude of the union coefficient decreases during the early 1980s, and is statistically insignificant and close to zero by 1985. In short, differences in shareholder risk between union and nonunion companies diminished during the 1980s, and remaining differences now appear to be of minor importance. What explains this change over time in union-nonunion differences, and what changes led to a decreased shift of risk from shareholders to union labor? We will return to these questions in a subsequent section.⁵

Below, we present the coefficients on the control variables from a pooled 1973-87 equation in which intercepts and union coefficients are allowed to vary by year (these coefficients and standard errors are highly similar to those presented in Table 1), while other coefficients are fixed (standard errors are biased downward in the pooled model). The following results are obtained (|t| in parentheses):

$$b_{iy} = .021\text{SIZE} + .174\text{K/L} - .224\text{CAPAGE} + .002\text{LEV} - .052\text{CR} - .111\text{AGE} + \Sigma\alpha_y + \Sigma\Psi_y\text{COV}_{iy} + \Sigma\Phi_k\text{IND}_k,$$

(10.25) (4.87) (1.80) (0.72) (2.39) (12.04) n = 4,851, R² = .257

Although there is no consensus in the theoretical or empirical literature on the determinants of systematic risk, results are largely consistent with expectations. Large, capital-intensive firms display greater shareholder risk, while older companies, those in less competitive markets, and those with older capital stocks provide lower risk. No significant relationship is found between systematic risk and leverage.⁶

⁵ Although the Master File ends in 1987, substitution of 1988 and 1989 beta values (measured for 1986-90 and 1987-91, respectively) in equations with a full set of 1987 control variables yields union coverage coefficients (|t| ratios based on heteroskedasticity adjusted standard errors) of 0.004 (0.08) and 0.019 (0.39), respectively, indicating that union-nonunion differences in shareholder risk remained small and insignificant beyond 1987. Using a measure of b_m (see footnote 4), the union coefficients from similar equations are 0.049 (0.61), -0.131 (0.92), 0.138 (0.91), and 0.057 (0.62) during 1988-91.

⁶ See, for example, Lev (1974), Thompson (1976), Subrahmanyam and Thomadakis (1980), and Mandelkar and Rhee (1984). Lev correctly predicts higher betas for more capital-intensive firms (i.e., those with higher operating leverage or ratio of fixed to variable costs); Subrahmanyam and Thomadakis argue the opposite. Binder (1992) reviews the literature relating systematic risk to concentration and size. He shows that negative coefficients on CR and firm size (a common finding) are consistent with a simple competitive model, as well as with the usual market power arguments. Size is measured here by employment, rather than sales or market value. The positive coefficient on size is reduced when other variables are excluded.

The robustness and generality of our results are examined in several additional ways. (The results, though not shown here, are available on request). We first examine the linearity of the beta-coverage relationship. Inclusion of both the union coverage and a coverage-squared variable produces in most but not all years a slightly more negative coefficient on COV and a positive (but generally insignificant) coefficient on its square. We also estimate a specification which includes both COV, the continuous union coverage variable, and a dummy variable equal to one if coverage is positive. Coefficients on the coverage dummies are negative, relatively small, and statistically insignificant in nearly all years, while coefficients on COV are affected relatively little.

As a final check on linearity, we include categorical variables for low, medium, and high coverage levels, with nonunion companies being the omitted base group. The breakpoint between low and medium coverage is .25, while that between medium and high coverage is .50. The breakpoints 0, .25, and .50 divide the sample of firms into approximate quartiles. During the 1973-81 period, annual coefficients on the coverage dummies average approximately -.06, -.08, and -.12 for low, medium, and high coverage companies, respectively. These results indicate a somewhat larger union-nonunion difference in beta than implied by the linear profile previously estimated. The corresponding averages for 1982-87 are -.02, -.05, and -.03, respectively. During the years 1985-87, coefficients on the high coverage dummy are effectively zero. The step function results confirm the basic conclusions reached previously, while also indicating deviations from linearity during some of the years, and a more varied relationship than is evident in Table 1.

One concern is that since firm union coverage is measured at two points in time – 1977 and 1987 – with estimated values used for other years, union coefficients for years furthest away from 1977 and 1987 may be the least reliable. Becker and Olson (1989), who use a single 1977 coverage measure for the entire 1970-1981 period, face a similar problem.

Although we do not have measures of firm-level coverage for all years, we are confident that the results shown are fairly reliable. Union coverage changes slowly, and values are highly correlated over time (Hirsch, 1991a, p. 26, reports a correlation of .87 between the 1977 and 1987 measures). And

because product and financial market variables and relationships exhibit considerable change over time, the advantages gained by looking at all years over a long time period are likely to exceed disadvantages stemming from an imperfect measure of union coverage.

An additional concern is that because the 1977 union coverage data were collected retrospectively in 1987, there is a potential for recall error in data for the 1970s and early 1980s. Measurement error from this source should bias the union coverage coefficients toward zero in the earlier years of our sample relative to the later years. Yet, the pattern we observe is exactly the reverse, indicating that recall bias is not responsible for our conclusions.

An additional issue to be examined is the effect of a changing panel composition over time. To this point, all estimates have been based on an unbalanced panel. The advantages of using an unbalanced panel are that it provides increased efficiency owing to larger sample sizes, and the sample in any year is more representative of the larger population of publicly traded companies than is the sample in a balanced panel. The major disadvantage of an unbalanced panel is that we may mistake changes resulting from differences in the sample of companies for true changes in parameter values over time.⁷

Although estimation with a balanced panel considerably reduces our sample size and causes moderate changes in the coefficient estimates, the pattern and conclusions reached previously continue to hold. For example, the coefficients (and t-ratios adjusted for heteroskedasticity) on COV from the model with full controls and a balanced panel of 202 firms are -.136 (2.25) and -.027 (0.46) for 1977 and 1987, respectively. These results are similar to those shown in Table 1, model (2), and indicate that the pattern of change shown in the coefficients is not due primarily to changes in the composition of firms. For the reasons stated above, we have more confidence in the results based on use of the larger unbalanced panel.

Construction of a balanced panel also allows estimation of a longitudinal model that accounts for omitted firm-specific determinants of beta fixed over the two time periods. For convenience, we drop subscripts i for firm, designate 1977 as year 1 and 1987 as year 2, and let the change operator Δ represent

⁷ Of course, changes in the risk-coverage relationship owing to changes in the sample of firms over time are accounted for in part by the inclusion of detailed control variables. And even in a balanced panel, the structure of included companies can change considerably over time.

values in year 2 minus those in year 1. Beginning with a model in levels similar to (2) but with a firm-specific fixed effect and with different coverage coefficients in the two years, we obtain the following difference equation:

$$(3) \quad \Delta b = \Delta \alpha + \Psi_1 \Delta \text{COV} + \Delta \Psi \text{COV}_2 + \sum \beta_k \Delta X_k + \Delta e.$$

The coefficient on ΔCOV provides an estimate of the 1977 Ψ , while the coefficient on the 1987 coverage level (COV_2) measures the change in w between 1977 and 1987. Unmeasured firm specific fixed effects on beta drop out in this model. As widely recognized, a disadvantage of longitudinal models is that measurement error in a variable in levels will be magnified in difference form. Thus, the noise-to-signal ratio of ΔCOV will be quite high, biasing its coefficient toward zero. When the following model is estimated, the following results are obtained:

$$(4) \quad \Delta b = \Delta \alpha - 0.046 \Delta \text{COV} + 0.139 \text{COV}_2 + \sum \beta_k \Delta X_k, \quad n = 202, \quad R^2 = 0.080$$

(0.35) (2.15)

The change in the union coefficient over the 1977-87 period is estimated to be .139, similar to the .109 change observed in the levels estimates of Ψ from the balanced panel. The estimate of Ψ for 1977, based on the coefficient of ΔCOV , is close to zero with a large standard error. We place relatively little weight on this result, however, owing to the substantial measurement error in the *change* in coverage variable.

Rate of Return Results

We turn next to an examination of rate of return differences among shareholders of union and nonunion companies. Although there should not exist systematic ex ante differences in risk-adjusted rates of return, ex post differences will result from unanticipated events that differentially affect union and nonunion companies. One would not expect the effects of unanticipated events to be large, sustained for long time periods, or systematically correlated with firm union coverage, since investor expectations adjust relatively rapidly to new information. Becker and Olson, however, did find a fairly sustained loss in market value among union companies during the late 1970s, supporting the hypothesis that unionized companies were subjected to continuing negative effects during this period (their results, however, are rarely statistically significant when industry dummies are included).

Union-nonunion differences in shareholder rates of return are estimated by the following annual regressions for each year y :

$$(5) \quad R_{iy} = \varphi_y + \tau_y \text{COV}_{iy} + \zeta_y b_{iy} + \sum \Psi_{yk} \text{IND}_{ik} + e_{iy},$$

where R_{iy} are inflation adjusted rates of return for firm i in year y , COV_{iy} measure firm coverage in year y , $\zeta_y b_{iy}$ measures firm betas in year y and their coefficients by year (i.e., the market return on risk), $\sum \Psi_{yk} \text{IND}_{ik}$ represents two-digit industry dummies ($k=1, \dots, 19$) fixed over time and their annual coefficients, φ_y are annual intercepts (rates of return for nonunion companies in the omitted industry), τ_y are estimates of annual union coverage effects on R , and e_{iy} are random error terms with zero means and constant variances by year.

Table 2 presents the coefficients τ_y for specifications with and without industry dummies. Both models include beta to control for risk differences across firms. Differences in returns between union and nonunion companies during the 1973-76 period, although not statistically significant, indicate a tendency toward greater ex post returns among more highly unionized companies. Beginning in 1977, however, unionized companies sustained lower returns than did nonunion companies for six straight years, although differences are not always large, and standard errors are sizable. The cumulative sum of the 1977-82 coefficients, however, is -0.608, with a standard error of .188 (based on the specification with industry controls), indicating a 24% decrease in market value over the six year period for unionized companies with mean coverage of 40%, as compared to nonunion companies in the same industry group and with similar risk. The cumulative sum absent industry controls is -0.689, with a standard error of .177. The cumulative mean return over the 1977-82 period for the entire panel of firms was .586 (an 8.0% annualized real rate of return). The finding of a sustained difference in union and nonunion returns to shareholders during the 1977-82 period supports the previous findings of Becker and Olson.

Not surprisingly, the disadvantage to investors in highly unionized companies did not continue indefinitely. Indeed, the 1983-87 period indicates a substantial advantage to investors in union companies. The cumulative sum of τ_y (with industry controls) over the 15 year period from 1973-87 is -.032 (with a standard error of .334), indicating no difference between union and nonunion companies over

the very long run. The mean 15 year cumulative return for the entire sample of firms was 1.531; \$100 at the beginning of 1972 would be worth an average of \$253 at the end of 1987 (in constant dollars). Or equivalently, the average annualized real rate of return over the 15 year period was 6.4%. The similarity in returns over long periods is to be expected. The market permits ex post differences in risk-adjusted rates of return, but should not permit sustained or predictable ex ante differences in shareholder returns. Less obvious are causes and consequences of the sustained poor performance of union companies during the late 1970s and early 1980s, followed by a relatively strong performance after 1982.⁸

Interpretation

Risk Shifting and Worker Monitoring

Although our results strongly confirm the Becker and Olson's findings for the 1970s, it is important to ask whether their thesis of greater risk shifting from shareholders to labor in union companies is a reasonable one, and whether it is consistent with the 1980s evidence. Becker and Olson explain the existence of lower betas in union companies by positing a shift of risk from shareholders to workers. But why would workers in general, and union workers in particular, agree to explicit or implicit contracts in which they bear a substantial degree of risk? Workers are risk-averse and far less able than shareholders to adjust their portfolios to acquire a preferred degree of market risk. Tying workers' income and employment too closely to the ups and downs of their company's performance also means they must bear unsystematic risk, which shareholders can eliminate through diversification. And if union bargaining power permits workers to acquire wage premiums, should not risk-averse workers want to realize some of their "rent" in the form of *lower* wage and employment risk?

Becker and Olson suggest that unions and unionized workplaces severely limit employer monitoring of worker effort, so that unionized firms must shift risk in order to make worker goals more fully incentive-compatible with those of shareholders. This monitoring cost explanation for greater risk shifting in union companies than in nonunion companies is reasonable, and consistent with lower betas in

⁸ The magnitude and significance of the union coefficients in the rate of return equations display moderate sensitivity (in both directions) to the inclusion of company-years with extreme rate of return values and the omission of beta. Company-year observations with nominal $R > 3.0$ or $R < -0.75$ are excluded. Coefficients on beta vary considerably from year to year.

union than in nonunion firms during the 1970s. But can we reconcile it with the narrowing of union-nonunion beta differences during the 1980s and the approximate equality in betas by 1987? To do so, we must accept the argument that shareholders in union companies during the 1980s accepted an increasing share of total enterprise risk in return for gaining the ability to implement more detailed monitoring of worker effort. In fact, although we have no direct evidence on this point, the argument has some plausibility. Weakened bargaining power during the 1980s, coupled with the much-discussed transformation in industrial relations (Kochan, Katz, and McKersie, 1986), may have produced a union workplace that allowed substantially greater employer monitoring than before. On the other hand, weakened union bargaining power might have permitted, *ceteris paribus*, increased monitoring to be accompanied by increased rather than decreased shifting of risk to workers.

What are the consequences of the risk-shifting hypothesis? A shift of risk from shareholders to labor in union firms should be evinced not only by a compensating wage premium, but also by greater employment or wage variability (or both).⁹ Empirical evidence indicates that union workers face both greater wage and employment variation than do nonunion workers. For example, Raisian (1983) analyzed the variability in weekly wages, weekly hours, and annual weeks worked among a sample of male heads of households in the Michigan PSID for 1967-79. He regressed individual deviations from means in each of these variables on a measure of the business cycle (proxied by annual industry unemployment rate in workers' industry of employment, minus the average industry unemployment rate for the entire period), plus a number of control variables. Raisian found negative coefficients on the unemployment measure in each of the equations, implying procyclical wage, hours, and employment. Similarly, he found that blue-collar workers bore greater risk than did white-collar workers. Of interest for this paper is his finding that union workers exhibited greater cyclicity (more negative coefficients) in wages, hours, and employment than did nonunion workers. Thus, the labor market evidence presented by Raisian is consistent with the finding of lower betas for shareholders in union companies, since union workers face larger income and employment risk than do nonunion workers.

⁹ Li (1986) provides a thorough analysis relating worker and shareholder risk associated with employment fluctuations.

Freeman and Katz (1991) utilized very different data and estimating equations, yet arrived at conclusions similar to Raisian's. Using annual manufacturing industry data for 428 industries for the period 1958-84 (with additional analysis employing individual data from the CPS), Freeman and Katz found that more highly unionized industries displayed greater variability in both wages and hours in response to changes in industry sales than did less unionized industries. Wage variability was particularly large among unionized industries in which workers realized large union wage premiums (estimated separately from the CPS).

The results in the Raisian and Freeman-Katz studies not only lend support to the finding that unionized companies realize lower shareholder risk by shifting risk to workers, but also indicate that union workers sustain risk through both employment and wage variability. The finding of greater employment risk among union than among nonunion workers is well established (e.g., Medoff, 1979; much weaker evidence, however, is found in Montgomery, 1991). The finding of larger wage variability is more surprising, given the prevalence of explicit contractual wages in the union sector. Freeman and Katz contend, however, that unionized industries with large wage premiums can exercise far more wage discretion in the face of output changes than can industries paying competitive wages.

These studies leave unanswered whether the narrowing of union-nonunion differences in shareholder risk (beta) during the 1980s coincided with a lessening of wage and employment risk for union workers relative to nonunion workers. We are unable to provide a definitive answer to this question, since no comprehensive study of which we are aware compares union-nonunion differences in wage and employment variability into the late 1980s. The narrowing of the beta differential is, however, consistent with much anecdotal evidence in the business press indicating that wage and employment risk for nonunion white-collar workers increased during the 1980s, suggesting a shift of risk from shareholders to workers in nonunion companies.

Bell and Neumark (1991) found a sharp increase in the use of lump-sum payments and profit sharing plans in union contracts during the 1980s. They found some evidence that profit sharing plans were associated with greater labor cost flexibility (no such evidence was found for lump-sum payments).

Such plans should shift greater wage (but not employment) risk to workers, the opposite of what is implied by the risk-shifting hypothesis and our evidence of narrowing union-nonunion differences in shareholder risk during the 1980s. Kruse (1991), however, provided evidence that over the 1971-85 period, manufacturing companies with profit sharing plans exhibited significantly greater employment stability. Therefore, a possible explanation for our results is that the increased use of profit-sharing plans by union companies in the 1980s decreased relative employment risk by an amount sufficient to lessen differences in shareholder risk between union and nonunion companies.

Enterprise Risk and the Use of COLAs

Other explanations for union-nonunion differences in risk are worth exploring. Can the use of cost-of-living agreements (COLAs) in union contracts help explain union-nonunion differences in systematic risk, as well as offer an alternative to the thesis of risk shifting between shareholders and labor?¹⁰ If there exists certainty about future product prices and the price level, contractual agreements on nominal and real wages are identical and there is no need for COLAs. Price uncertainty creates uncertainty about real wages. In the absence of COLAs, nominal wage increases will be contracted to reflect both the expectations of the parties and a premium for risk. Both parties may benefit from a COLA. If workers expect higher inflation than does the firm, both parties will prefer, ex ante, indexed wages over contractual nominal wage increases. Even if the parties' expectations (mean and variance) of price increases are identical, both may prefer to lower risk by indexing (fully or partially) the labor contract. In particular, if workers are risk-averse and unable to diversify, the ex ante costs to the firm of an indexed contract should be lower than one specified entirely in nominal wages. That is, workers will accept lower expected wage increases over the life of a three-year contract if they are partially protected from large unanticipated price increases.

COLAs therefore may lower total enterprise or joint risks to shareholders and labor (ex ante) during periods of high price uncertainty. But only union companies are likely to adopt explicit COLAs,

¹⁰ There is a large literature on the use of indexing, much of it dealing with contract length. Frequently cited studies on indexing include Gray (1978), Hendricks and Kahn (1983), Ehrenberg, Danziger, and San (1983), and Card (1986). Woglom (1990) explicitly examines the relationship between systematic risk and wage indexation.

because they face lower contractual and enforcement costs than do nonunion firms. Therefore, union companies may have exhibited lower systematic risk to shareholders *and* workers than did nonunion companies during the 1970s and early 1980s when inflation was high and the use of COLAs was most widespread. As the benefits of COLAs declined in the 1980s, their use declined and differences in shareholder risk based on union status narrowed. Similarly, firms in industries with relatively low contractual and monitoring costs, and relatively high ex ante benefits from COLAs, are most likely to have adopted COLAs.

Unlike Becker and Olson's explanation, the COLA explanation does not require that there be a greater shifting of risk to labor in union companies than in nonunion companies; rather, lower contracting costs in union companies lowers joint enterprise risk during periods of price uncertainty. Of course, both explanations for union-nonunion differences in beta may be operable. We cannot easily distinguish between the risk-shifting and lower enterprise risk views unless we can measure labor as well as shareholder risk.

Although the use of COLAs can lower shareholder as well as employee risk, the prediction of a negative beta-COLA relationship need not always follow. Assume initially that there exist unanticipated demand shocks during the length of a contract, and that the firm's product price is positively correlated with the economywide price index used in the contractual COLA (*anticipated* real and monetary shocks are reflected in the contract). In this case, a union firm in which there is partial indexing would have a lower beta than a firm with wages fixed over the period.¹¹ That is, positive demand shocks will lead to higher market and firm returns, but firm returns will increase less than proportionately owing to upward wage adjustments resulting from the COLA. Similarly, a negative demand shock will have a smaller impact on a firm with a COLA provision than on a firm with constant wages, all else equal.

Although a COLA with partial adjustment to price changes lower shareholder risk relative to a policy of fixed wages, it produces exactly the opposite result compared to a practice of full wage

¹¹ Typically, U.S. contracts with COLAs have partial rather than full indexing, combined with frequent use of restrictive "minimums" and "caps". For detailed evidence, see Kaufman and Woglom (1986). Card (1983) provides Canadian evidence indicating less than complete indexing in manufacturing.

adjustment (i.e., invariant real wages). The relevant question is then the speed of adjustment in nonunion wages to unanticipated demand shocks. If nonunion wages adjust rapidly to changes in demand, then shareholder risk (beta) will be higher rather than lower in the union sector. Evidence in Raisian and Freeman-Katz, however, suggests that wages and employment adjusted to output changes more slowly in the nonunion sector than the union sector. The declining use of COLAs over time does imply, *ceteris paribus*, that the degree of shareholder risk in unionized firms should grow more and more similar to that in nonunion companies.

If unanticipated price changes are the result of supply rather than demand shocks (e.g., higher prices of imported oil or unanticipated environmental regulations), the conclusions stated above are reversed. A company with partial indexing will have higher betas than will firms with fixed wages, since higher wages will exacerbate supply-side cost increases resulting from negative supply shocks. Similarly, supply-side cost decreases (such as unanticipated oil price decreases) allow firms with indexed wages to lower their labor costs more than firms with wages fixed at the beginning of an explicit or implicit contract period. Although unanticipated supply shocks to the economy are typically less important than are demand shocks, their existence complicates clear-cut theoretical predictions about a relationship between indexing and systematic risk.

As discussed above, theoretical considerations suggest an uncertain relationship between the use of COLAs and risk to shareholders and labor. Is there at least circumstantial evidence suggesting a link between the use of COLAs and union-nonunion differences in beta? We do not have firm-level information on the contractual use of COLAs. Rather, we use information on the proportion of union workers who are covered by a COLA, provided for two-digit SIC industries for years since 1975 (*Monthly Labor Review*, January issues). COLA coverage by industry-year is matched with regression estimates of union-nonunion beta differences, estimated separately by two-digit industry and year from our firm-level data set. Specifically, a pooled regression, with firm betas as the dependent variable, is first estimated for all firm-years during 1975-87 (n=4,186; because of a very small number of tobacco industry firms, firm-years in SIC 21 are deleted). In addition to control variables, 247 separate industry-

by-year union coverage coefficients are estimated (19 two-digit manufacturing industries times 13 years).

That is, we first estimate

$$(6) \quad b_{iy} = \alpha_y + \sum \theta_{jy} \text{COV}_{iy} \text{YEAR}_y \text{IND}_{ij} + \sum \beta_k X_{iyk} + e_{iy},$$

where b_{iy} are betas for firm i in year y , $\sum \text{COV}_{iy} \text{YEAR}_y \text{IND}_{ij}$ represents COV in firm i in year y times 247 dummy variables interacting y years (13) and j two-digit manufacturing industries (19). As in equation (2), the control variables included in X_{iyk} are SIZE, K/L, AGE, LEV, CR, and CAPAGE.

The 247 year-by-industry coefficients θ_{jy} provide estimates of differential union coverage effects on beta. These values constitute the dependent variable in the following weighted least squares (WLS) regressions (n=247).

$$(7a) \quad \theta_{jy} = \alpha + \gamma \text{COLA}_{jy} + e_{jy}$$

$$(7b) \quad \theta_{jy} = \alpha + \gamma \text{COLA}_{jy} + \sum \vartheta_y \text{YEAR}_y + e_{jy}$$

$$(7c) \quad \theta_{jy} = \alpha + \gamma \text{COLA}_{jy} + \sum \Psi_j \text{IND}_j + e_{jy}$$

The variable COLA_{jy} represents the extent of COLA coverage among unionized workers in two-digit industry j in year y , while the coefficient γ is an estimate of the relationship between the union beta differential and COLA coverage. A finding of $\gamma < 0$ would indicate that differential use of COLAs across time, industries, and between union and nonunion companies helps account for union-nonunion differences in systematic risk. The coefficient γ in equation (7a) captures union-nonunion risk differences correlated with COLA coverage differences across industries and across years, (7b) controls for year and thus captures coverage differences across industries, and (7c) controls for industry and captures coverage variation across years.

In addition to estimating (7a)-(7c), we provide estimates of equations in which the COLA coefficients are allowed to differ between 1975-81 and 1982-87. The year 1982 provides a convenient breakpoint, since 1981 was the last year examined in the Becker and Olson study, union organizing activity fell sharply at about that time (Chaison and Dhavale, 1990), and, most important, inflation

dropped from double digits to under 4% per year.¹² All equations are estimated by WLS, where weights are the inverse of the error variances on θ_{jt} from equation (6). Such a procedure is recommended when regression parameters are used as the dependent variable (Saxonhouse, 1976).

Table 3 presents regression estimates for equations (7a)-(7c). A negative and statistically significant relationship between θ and COLA is found. Highly unionized companies have lower shareholder risk than do similar companies with low union coverage in those industries and years where COLA coverage is most extensive. A one standard deviation change in COLA (equal to .35) implies a change in the union-nonunion beta differential of about .07 (i.e., $-.20 \times .35$), over half the sample mean of the differential (the mean of θ is $-.11$). The differential use of COLA clauses across industries and time, therefore, helps explain our earlier finding with respect to union-nonunion differences in shareholder risk. Note that we find a significant θ -COLA relationship in spite of the fact that the COLA variable may measure with considerable error industry-specific COLA coverage among the small number of firms in our sample for many of the two-digit industries (random measurement error biases γ toward zero). Confidence in our qualitative results is reinforced by the finding of similar estimates of γ in (7a) and (7b). The estimate of γ in (7c) is smaller in magnitude, indicating that interindustry rather than intertemporal variation in COLA coverage is the more important determinant of union-nonunion differences in shareholder risk.

When coefficients are allowed to differ between periods, we find an even stronger negative relationship during the 1975-81 period, and a far weaker relationship after 1981. This finding indicates that COLAs were effective in lowering risk during a period of high and uncertain inflation, but had little effect during a period of relative price stability. Lower shareholder risk in union than in nonunion firms during the 1970s and early 1980s, first noted by Becker and Olson and confirmed in this study, may well have been unique to this period and tied closely to a high degree of macroeconomic uncertainty.

¹² December to December changes in CPI-U in 1979, 1980, and 1981 were 13.3, 12.5, and 8.9 percent, respectively, as compared to 3.8, 3.8, 3.9, 3.8, 1.1, and 4.4 percent during the years 1982-87 (Council of Economic Advisors, 1992, p. 365).

The negative θ -COLA relationship shown in Table 3 is consistent with our earlier hypothesis that COLAs lower the ex ante joint risk to shareholders and workers, assuming that most price changes result from demand shocks rather than supply shocks and that COLAs increase nominal wage flexibility (or decrease real wage variability) relative to that observed in union companies without COLAs and nonunion companies without collective bargaining agreements. Absent these assumptions, the prediction of a negative θ -COLA relationship need not follow. The observed negative relationship between θ and COLA coverage then becomes a “fact” in search of an explanation. Although our hypothesis of union-nonunion differences in joint risks owing to differential use of COLAs finds empirical support and offers an alternative to the Becker-Olson risk-shifting hypothesis, the two explanations are not mutually exclusive.¹³ Theory and evidence for each hypothesis are plausible; evidence in support of neither hypothesis is conclusive.

The Implications of Rate of Return Differentials

Our evidence supports the finding by Becker and Olson of risk-adjusted rates of return to investors that decreased substantially with respect to union coverage during the late 1970s. Although ex post returns can certainly differ at any point in time, evidence of sustained poor market performance over the 1977-82 period is not so easily explained. The evidence suggests substantial and sustained environmental shocks that negatively affected the long-run expected earnings of union companies relative to nonunion companies. The long time period over which a difference in returns is observed suggests the existence of numerous unanticipated shocks throughout the period or investor sluggishness in fully capitalizing shocks into firm valuation. The pattern of union-nonunion return differences after 1982 suggests that the relative performance of union companies was far more favorable than anticipated by investors.

Becker and Olson speculate that the poor financial performance of union companies increased investor pressure on management to mitigate union effects, and helped accelerate the decline in union

¹³ Similarly, the incorporation of profit sharing plans may lower total enterprise risk – both workers and shareholders may prefer procyclical variation in compensation in return for lower employment (output) variability. Estimation of (7a)-(7c) using OLS rather than WLS produces less clear-cut results.

coverage. Because we have measures of union coverage in both 1977 and 1987, the thesis that poor financial performance is associated with subsequent reductions in coverage can be tested. Specifically, we regress the change in union coverage in firm i , ΔCOV_i , measured by 1987 minus 1977 coverage, on a measure of past company rates of return, R_i , and k firm/industry controls:

$$(8) \quad \Delta\text{COV}_i = \Phi \bar{R}_i + \sum \beta_k X_{ik} + e_i.$$

\bar{R}_i is measured by the annualized inflation adjusted rate of return over the 1977-82 period, and control variables include 1982 firm size, capital intensity, company age, leverage, industry concentration, capital age, and industry dummies. The panel includes 364 companies for which data are available for 1977-82 and 1987. In order to increase the sample size of the panel, we do not include beta as a control variable (its coefficient is close to zero in a regression using the smaller panel).

Table 4 presents regression results of ΔCOV regressions, with and without firm and industry controls. Evidence is found to support the thesis that poor stock market performance is associated with subsequent declines in firm-level union coverage. A point estimate of $\Phi = 0.12$ implies that union coverage fell by 0.012 in response to, say, a 0.10 decrease in the average rate of return. Estimates in Table 2 indicate that union firms with 40% coverage had annual rates of return approximately 4 percentage points lower (.40 times -0.10) during years of the 1977-82 period. Therefore, union coverage declined an estimated 0.005, or about one-half percentage point, in response to poor stock market performance in the late 1970s and early 1980s. Although far from trivial, the estimated decline in union coverage is quite small relative to the average decline in coverage for this sample of companies during the 1977-87 period.¹⁴

Finally, it is important to note that equivalent market returns to investors do not eliminate shareholder and management incentives to mitigate union bargaining strength, as long as firm coverage continues to have a negative effect on profitability and market value. Union coverage is associated with

¹⁴ The estimate of Φ may be biased toward zero owing to simultaneity between ΔCOV and R . That is, a low R leads to subsequent reductions in coverage, but (unanticipated) declines in coverage are expected to lead to higher rates of return. The use of an annualized R for 1977-82, but 1977-87 changes in coverage is likely to mitigate simultaneity bias.

significantly lower profitability and market value (measured by rates of return on capital and Tobin's q , respectively) during both the late 1970s and 1980s (Hirsch, 1991c). Although market investors may expect to earn equivalent ex ante risk-adjusted rates of return, capital gains or losses still result from unanticipated changes in union coverage and the union "tax" on company earnings.

Conclusion

Our analysis has provided clear-cut evidence corroborating the novel findings of Becker and Olson regarding union-nonunion differences in shareholder risk and returns, despite differences between the two studies in the sample of firms used and the measure of union coverage employed. We find that during the 1970s and early 1980s, more highly unionized companies had substantially lower shareholder risk than otherwise similar nonunion and less unionized companies; and ex post rates of returns to investors were lower among union than among nonunion companies during the 1977-82 period.

On the other hand, we do not find support for either of those conclusions in our analysis of data for the mid-1980s (years that Becker and Olson did not examine). On the contrary, union-nonunion differences in returns after 1982 indicate a pattern more favorable to unionized companies. Differences in shareholder risk between union and nonunion companies weakened considerably during the 1980s, and were small and insignificant by 1985-87.

Finally, we provide direct evidence to support the thesis that firms with relatively low rates of return during the 1977-82 period tended to experience larger than average declines in firm-level union coverage between 1977 and 1987.

We have been less successful in explaining union-nonunion differences in risk and returns during the 1970s, or the change in these patterns in the mid- and late-1980s. Becker and Olson plausibly hypothesized that risk is shifted from shareholders to labor in union firms because union firms have higher monitoring costs than do nonunion firms. That hypothesis would be more compelling, however, if the narrowing of union-nonunion differences in shareholder risk during the 1980s could be explicitly related to changes in monitoring costs. Taking a different tack, we have found some evidence supporting the notion that union-nonunion differences in shareholder risk across industries and time are related to the

prevalence of COLAs. This evidence supports the thesis that union-nonunion differences in shareholder risk may result in part from differences in total enterprise risk, and not just from differences in the allocation of risk between shareholders and employees. A fruitful avenue for future research may be to explore alternative theoretical explanations and more direct empirical tests for differences in shareholder-labor risk between union and nonunion companies.

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Table 1: The Effect of Union Coverage on Shareholder Risk, 1973-87.

Year	n	\overline{COV}	$\bar{\beta}$	$\sigma(\beta)$	(1) No controls		(2) Firm/industry controls	
					Ψ	$ t $	Ψ	$ t $
1973	316	0.349	1.056	0.201	-0.220	(5.12)	-0.167	(3.85)
1974	333	0.350	1.045	0.201	-0.233	(5.61)	-0.166	(4.41)
1975	340	0.350	1.052	0.213	-0.221	(4.99)	-0.180	(4.43)
1976	346	0.343	1.030	0.203	-0.231	(5.59)	-0.170	(4.53)
1977	350	0.343	1.088	0.223	-0.253	(5.77)	-0.172	(4.20)
1978	349	0.340	1.051	0.223	-0.207	(4.78)	-0.161	(3.95)
1979	344	0.330	1.028	0.215	-0.209	(4.87)	-0.157	(3.65)
1980	343	0.315	1.018	0.198	-0.166	(4.16)	-0.148	(3.79)
1981	341	0.305	1.010	0.202	-0.134	(3.31)	-0.121	(3.14)
1982	333	0.297	1.015	0.205	-0.148	(3.43)	-0.128	(3.41)
1983	324	0.290	1.023	0.207	-0.145	(3.18)	-0.103	(2.50)
1984	313	0.280	1.017	0.189	-0.119	(2.73)	-0.091	(2.37)
1985	298	0.267	1.110	0.185	-0.019	(0.39)	-0.024	(0.59)
1986	272	0.253	1.106	0.174	-0.032	(0.66)	-0.008	(0.18)
1987	249	0.254	1.097	0.182	-0.078	(1.30)	-0.044	(0.84)

Sample size is n , \overline{COV} is the mean union coverage rate, $\bar{\beta}$ is the mean beta, and $\sigma(\beta)$ is the standard deviation of β . Presented are annual regression results, with dependent variable b_{iy} , the beta for firm i in year y . Ψ are the coefficients on COV. Column (1) regression includes only COV. Column (2) regression includes SIZE, K/L, CAPAGE, AGE, LEV, CR, and IND (19 dummies). T-ratios based on standard errors adjusted for heteroskedasticity using method by White's (1980) method.

Table 2: The Effect of Union Coverage on Shareholder Rates of Return, 1973-87.

Year	n	\bar{R}	(1) Industry dummies excluded		(2) Industry dummies included	
			τ	$ t $	τ	$ t $
1973	327	-0.275	0.055	(0.76)	-0.002	(0.02)
1974	345	-0.325	0.080	(1.35)	0.043	(0.69)
1975	346	0.530	0.039	(0.40)	0.039	(0.39)
1976	354	0.354	0.115	(1.48)	0.098	(1.21)
1977	361	-0.047	-0.046	(0.89)	-0.023	(0.42)
1978	360	0.045	-0.134	(2.28)	-0.225	(3.47)
1979	353	0.211	-0.056	(0.71)	-0.029	(0.35)
1980	351	0.192	-0.146	(2.10)	-0.157	(2.05)
1981	355	-0.053	-0.136	(2.12)	-0.108	(1.55)
1982	347	0.238	-0.171	(1.69)	-0.066	(0.62)
1983	339	0.305	0.101	(1.23)	0.096	(1.06)
1984	327	-0.057	-0.001	(0.02)	0.012	(0.19)
1985	313	0.275	0.092	(1.09)	0.179	(1.96)
1986	290	0.116	-0.044	(0.63)	-0.001	(0.01)
1987	264	0.022	0.226	(2.88)	0.112	(1.37)

The dependent variable is R_{iy} , the rate of return in firm i in year y . Means of \bar{R} are listed by year. τ are the coefficients on COV. Column (1) regression includes COV and β_{iy} . Column (2) regression includes β_{iy} and IND (19 dummies). $|t|$ in parentheses.

**Table 3. Union Effects on Risk and COLA Coverage:
Industry-by-Year Observations with WLS Estimation, 1975-87.**

Independent Variable	(1)	(2)	(3)	(4)	(5)	(6)
COLA	-0.199 (6.39)	-0.194 (6.47)	-0.103 (1.37)	-0.257 (7.95)	-0.229 (5.98)	-0.164 (2.31)
COLA x 1982+	-	-	-	0.152 (4.69)	0.089 (1.45)	0.135 (5.87)
YEAR	no	yes	no	no	yes	no
IND	no	no	yes	no	no	yes
R ²	0.143	0.252	0.584	0.214	0.259	0.639

The dependent variable are regression parameters θ_{jy} , representing union coverage effects on beta by two-digit industry (19) and year (13), estimated by equation (5) in text. Sample size in all regressions is 247 (19 x 13). SIC 21 (tobacco) is excluded. COLA measures the proportion of union workers covered by a COLA in the two-digit industry by year. The interaction term COLA x 1982+ is equal to COLA times a dummy equal to one for years beginning in 1982. The t-ratios corresponding to the sum of coefficients for COLA and COLA x 1982+ are 2.91, 2.91, and 0.41 for columns (4)-(6), respectively. Estimation is by weighted least squares, with the inverse of the error variances of θ_{jy} as weights. Columns (1) and (4) include no controls, columns (2) and (5) include 12 year dummies, and columns (3) and (6) include 18 industry dummies. $|t|$ in parentheses.

**Table 4. Shareholder Returns and Changes
in Firm Union Coverage, 1977-87.**

Independent Variable	(1)	(2)
\bar{R}	0.131 (2.34)	0.118 (1.88)
SIZE	–	0.005 (0.94)
K/L	–	-0.063 (0.59)
AGE	–	-0.052 (1.93)
LEV	–	0.045 (0.58)
CR	–	0.039 (0.65)
CAPAGE	–	0.209 (0.71)
IND	no	yes
R ²	0.015	0.087
<i>n</i>	364	364

The dependent variable is ΔCOV_i , equal to firm union coverage in 1987 minus coverage in 1977. \bar{R}_i measures the annualized inflation adjusted rate of return over the 1977-82 period. The mean of ΔCOV is -.058. $|t|$ in parentheses.

Data Appendix: Definition of Variables

COV	Proportion of firm's North American workforce covered by a collective bargaining agreement in 1977 and 1987 (Hirsch, 1991a). For years prior to 1977, COV is set at the 1977 value. For years between 1977 and 1987, values of COV are set by linear interpolation.
<i>b</i>	Shareholder risk, measured by firms' stock market beta, measured with five years of monthly data centered on the year to which beta is matched (<i>Value Line Investment Survey</i> , issues for week closest to the end of each year).
<i>R</i>	Calendar year real rate of return to holding a share of firm's common stock, measured by dividends per share plus capital gain (loss), divided by beginning price, with adjustments made for stock splits (Hall, 1990). Inflation adjustment made using 4th quarter annual changes in Personal Consumption Expenditure (PCE) implicit price deflator.
AGE	Company age, measured by hundreds of years since incorporation (<i>Ward's Business Directory or Moody's Industrial Manual</i>).
SIZE	Log of employment, in thousands (Hall, 1990).
K/L	Capital stock per employee, in millions of 1987 dollars. Measured by net inflation-adjusted capital stock (deflated by GNP investment implicit price deflator), divided by employment (Hall, 1990).
LEV	Leverage, measured by the ratio of debt due over next year to value of common equity (Hall, 1990).
CAPAGE	Age of capital stock, in hundreds of years, approximated by difference in value between gross plant and net plant (accumulated depreciation) divided by current annual depreciation, all divided by 100 (Hall, 1990).
CR	Four-firm industry concentration ratio (range 0-1) adjusted for international trade and regional market concentration. Adjusted and unadjusted concentration ratios are available for 1972 and 1977 in Weiss and Pascoe (1986); unadjusted ratios for 1982 are available in the Census of Manufactures. Adjusted 1982 figures estimated based on ratio of adjusted to unadjusted ratios in 1977. Linear interpolation using adjusted 1972, 1977, and 1982 figures used to determine the annual 1973-82 figures. Concentration after 1982 is assumed unchanged. Matched to firms at 2-, 3-, or 4-digit SIC industry level, determined by firms' principal SIC code in Compustat.
IND	Dummy variables representing firms' two-digit manufacturing industry.
COLA	Proportion of union workers who are covered by a cost of living agreement, for two-digit SIC industries by year, 1975-87 (<i>Monthly Labor Review</i> , January issues).