Market Structure, Union Rent Seeking, and Firm Profitability

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Abstract

Unique survey data on firm union coverage, market share, and concentration are used to examine the hypothesis that profits deriving from market power provide a primary source for union gains. Based on evidence from 247 companies during 1972-1980, no support is found for this hypothesis.

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Introduction

Considerable attention has been given to studies concluding that monopoly profits associated with industry concentration provide the primary source from which labor unions capture rents (Salinger, 1984; Karier, 1985; Freeman, 1983). This conclusion has obvious appeal. It implies that the impact of concentration on profitability has been systematically understated in empirical studies owing to omitted variable bias (concentration and unionization are positively correlated), and that deleterious union effects on firm survival are mitigated since union rent seeking is restricted to firms realizing supra-competitive profits.

The conclusion that concentration-related profits provide the principal source for union gains has been attacked by Hirsch and Connolly (1987). They first note that evidence on this issue is mixed. Domowitz, Hubbard, and Petersen (1986) find little evidence that union effects on price-cost margins are more negative in highly concentrated industries (they find this relationship to vary over the business cycle), while Clark (1984) obtains the surprising result that union profit effects are most deleterious among lines-of-business with low market shares. Evidence from the labor market, moreover, reveals union wage premiums that are typically smaller in highly concentrated industries (Lewis, 1986, pp. 154-55), just the opposite of what is implied by the profitability studies cited at the outset. Hirsch and Connolly also provide firm-level evidence indicating that results from Salinger, in particular, are fragile.

Previous studies, however, have been hobbled by serious data limitations. This note presents empirical evidence on the relationship between profitability, market structure, and union coverage using a unique data set intended to overcome deficiencies inherent in past studies. Specifically, we employ firm-level measures of both market value and accounting profitability, utilize company-specific measures of union coverage collected by the author, and use sales-weighted measures of industry concentration and firm market shares.

Data and specification

Data are from an unbalanced panel of 247 large manufacturing-sector companies during 1972-1980. These firms represent the union of three data sets: (1) the R&D Master File, containing financial and other information on a large panel of U.S. manufacturing firms publicly-traded in 1976 and listed in Compustat;
(2) a file of 723 companies for which firm-specific union coverage data for 1977 has been collected (Hirsch, forthcoming); and (3) 347 Fortune 500 companies for which 1977 data on market share and industry concentration, weighted to reflect the distribution of firm sales across 4-digit industries, are available.  

Profitability is measured by \( \ln(q) \), the log of Tobin’s \( q \) (market value of the firm divided by the replacement cost of assets) and by \( \delta_k \), the rate of return on capital (gross earnings divided by the inflation-adjusted gross capital stock). A time-series/cross-section profit equation of the following general form can be specified for the 1972-80 period:

\[
D_{it} = \hat{\delta}_j X_{jit} + \hat{\delta}_k \text{YEAR}_{it} + \hat{\delta}_m \text{IND}_{li} + \hat{\delta}_n \text{MKST}_{mi} + \hat{\delta}_m \text{UN}_i \cdot \text{MKST}_{mi} + e_{it}
\]

where \( D_{it} \) is the profitability of firm \( i \) in year \( t \), measured alternatively by \( \ln(q) \) and \( \delta_k \), \( X \) is a vector including a constant and \( j-1 \) firm- and industry-specific variables that are \textit{time-variable} and \( \hat{\delta}_j \) is the attaching coefficient vector, \( \text{YEAR} \) represents \( k \) year dummies and \( \hat{\delta}_k \) its coefficient vector, \( \text{IND}_i \) represents 19 industry dummies and \( \hat{\delta}_m \) its coefficient vector, \( \text{UN}_i \) measures union coverage in firm \( i \) and \( \hat{\delta}_n \) is its coefficient, \( \text{MKST} \) represents the two market structure variables and \( \hat{\delta}_n \) is its coefficient vector, the coefficients \( \hat{\delta}_m \) attach to the union-market structure interactions, and \( e_{it} \) is an error term. The fixed or time-invariant variables/vectors \( \text{UN}, \text{IND}, \) and \( \text{MKST} \) are measured for 1977.

Estimation of (1) results in an error term that is serially correlated across years within firms, so that standard errors in (1) are biased downward (Hsiao, 1986). In order to purge the model of positively correlated errors, a two-step model is estimated. In the first step, firm profitability is regressed on all time-varying profit determinants and 246 firm dummy variables. In a second-step, firm fixed effects averaged over the 1972-80 period, measured by the coefficients on the firm dummies (plus a zero for the omitted first-step firm) are used as the dependent variable in a second-step GLS regression, with the time-invariant variables as regressors. The second-step regression is weighted by the inverse of the standard errors from the firm dummy coefficients. That is, the first-step OLS regression is:

\[^1\text{The R&D Master File is described in Cummins et al. (1985). The union variable measures the proportion of a company's workforce covered by a collective bargaining agreement in 1977. These data were collected on a retrospective basis in a recent survey by Hirsch and supplemented by data collected in a 1972 survey by the Conference Board (Hirsch, forthcoming). The market share and industry concentration data were provided by Mark Hirschey (Hirschey, 1982), and have been previously used by Hirsch and Connolly (1987).}\]
\( (1') \quad D_{it} = \hat{\theta} x_{jit} + \hat{\alpha} \text{YEAR}_{it} + \hat{\phi}_{n-1} FIRM_{n-1, i} + \hat{\delta}_{it}, \)

where \( FIRM \) is a vector of 246 firm dummies with coefficients \( \phi \). A second-step GLS regression is then estimated \( (n=247) \); the value of \( \phi \) for the omitted firm is zero:

\[ \varphi_i = \hat{\alpha} + \hat{\alpha} \text{IND}_{i} + \hat{\theta} \text{UN}_i + \hat{\phi} \text{MKST}_{mi} + \hat{\theta} \text{UN}_i \cdot \text{MKST}_{mi} + \hat{\delta}_{it} \]

which purges the error term of serial correlation, while using all available information for the 1972-80 period. Interpretation of the second-step regression coefficients is identical to that in (1).

Included in vector \( X \) are R&D intensity, measured by the R&D stock divided by sales and a dummy for firms with no reported R&D; advertising intensity, measured by the ratio of advertising expenditures to sales and a dummy for firms with no reported advertising; firm size, measured by the log of employment; capital intensity, measured by the log of the capital-labor ratio; firm-specific sales growth, measured by the annualized logarithmic two-year growth rate; industry sales growth, measured by the annualized four-year growth rate in the firm’s principal industry; and foreign competition, measured by the share of imports in domestic sales in the firm's principal industry. The time-invariant variables, measured for 1977, are \( \text{UN} \), representing firm-specific union coverage; \( \text{CR} \), measuring the sales-weighted 4-firm concentration ratio across industries; \( \text{MS} \), measuring the firm’s sales-weighted market share across industries; and \( \text{I-UN} \), measuring industry union coverage in the firm’s principal 2- or 3- digit industry (a detailed description of data and full regression results are available on request).

**Results**

Table 1 presents GLS regression results for the second-stage profit regressions (eq. 2), where the dependent variables are firm fixed effects on \( \text{ln}(q) \) and \( \delta_k \). The estimated union impact on profitability is sizable and consistent with previous estimates (Addison and Hirsch, 1989, Table 1). Following inclusion of industry dummies and union density in specifications 2 and 2 , a change from nonunion \( \text{UN} = 0.0 \) to 50 percent coverage \( \text{UN} = 0.5 \) implies a decrease in Tobin’s \( q \) by 9.0 percent and in \( \delta_k \) by 6.9 percent.\(^2\)

Estimates are substantially larger when industry dummies are excluded. Both industry concentration (\( \text{CR} \))

\(^2\) Letting \( \hat{\theta} \) represent the estimated coefficient on union coverage, the average percentage effect of coverage on profitability is calculated by \( [\exp(0.50 \hat{\theta}) - 1]100 \) for \( q \) and by \( (0.50 \hat{\theta} / \hat{\delta})100 \) for \( \delta_k \). Mean \( \hat{\delta} \) among nonunion companies is 0.109. Calculation of the effects of \( \text{CR} \) and \( \text{MS} \) proceeds in similar fashion.
and firm market share ($MS$) have positive and significant impacts on profitability. An increase in industry concentration of 10 percentage points is estimated to increase $q$ by 6.0 percent and $\delta_k$ by 3.2 percent. Each 1 percentage point increase in market share is associated with a 2.3 increase in $q$ and a 1.1 increase in $\delta_k$.

Specifications 3 and 3’ in Table 1 introduce interactions of firm union coverage with the market structure variables $CR$ and $MS$. The hypothesis that unions capture profits generated by market power implies positive coefficients on $CR$ and $MS$, and negative coefficients on $UN \cdot CR$ and $UN \cdot MS$. We find no empirical support for the hypothesis that $CR$ and $MS$ provide sources for union rents. The $UN \cdot CR$ coefficient is very close to zero, while the coefficient on $UN \cdot MS$ is positive and significant. This latter result, suggesting more deleterious union effects in companies with low market shares, is consistent with the finding by Clark (1984) in his study of union effects on the rate of return to capital among lines-of-businesses.

Also estimated is the pooled OLS time-series/cross-section model (eq. (1). Again, no evidence is found for the hypothesis that monopoly profits provide the primary source for union rents. The following results were obtained ($t$ in parentheses; $I-UN, IND$, and all time-varying regressors included; $n = 2,175$):

$$
\ln(q) = \hat{\alpha}X - 0.557UN + 0.134CR + 0.341UN \cdot CR - 0.142MS + 2.154UN \cdot MS, \quad R^2 = 0.582,
$$

\begin{align*}
(4.52) & & (0.90) & & (1.14) & & (0.51) & & (3.71) \\
\delta_k = \hat{\alpha}X - 0.043UN + 0.001CR + 0.042UN \cdot CR + 0.018MS + 0.109UN \cdot MS, \quad R^2 = 0.418.
\end{align*}

\begin{align*}
(5.60) & & (0.10) & & (2.26) & & (1.05) & & (2.99)
\end{align*}

Neither the $UN \cdot CR$ nor $UN \cdot MS$ coefficients are negative. If anything, the suggestion is that union profit effects are most deleterious among competitive companies with low industry concentration and market shares.

The failure to find support for the hypothesis that profits from market structure provide the primary source of union gains is not unique to the regressions results presented above. In work not shown, concentration and market share are entered as categorical dummies, separately and in interaction with union coverage, to explore possible nonlinearities in these relationships. We arrive at similar conclusions. In addition, we substituted an alternative industry concentration measure (not sales-weighted) for $CR$ and
obtained highly similar results. Other specifications, including various union interaction terms for which there exist reasonable theoretical arguments, provide a rather diverse set of results. But in no case do we obtain significant negative interaction terms of union coverage with concentration or market share.

**Interpretation and conclusion**

Unlike previous studies, the data set developed here contains a firm-level measure of union coverage, both market value and accounting measures of profitability, sales-weighted company measures of industry concentration and market share, and a detailed set of control variables. No evidence is found to support the contention that union rent seeking is more effective among firms with large market shares or in highly concentrated industries. The latter finding casts serious doubt on the frequently cited conclusions by Salinger (1984) and Karier (1985), but is consistent with results reported by Hirsch and Connolly (1987) and Domowitz, Hubbard, and Petersen (1986), as well as labor market evidence indicating smaller union-nonunion wage differentials in highly concentrated industries. The finding that market share enhances profitability by larger amounts in union than in nonunion companies is consistent with Clark’s (1984) evidence on lines-of-business using the PIMS data. Clark speculates that similar union wage premiums across firms will have more detrimental effects on low share firms unable to pass through cost increases.

If above-normal returns associated with market structure do not provide a significant source for union wage gains, what does? Careful analysis of this issue is complex and beyond the scope of this note. But theoretical and empirical analysis 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 changes in demand and other sources of disequilibria. Of course, the hypothesis that unions capture returns associated with market power remains plausible. To date, however, product and labor market evidence that might support this hypothesis is not persuasive.

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3 Hirsch (forthcoming) provides references and detailed evidence on these issues. He uses the complete union survey data base, but not the concentration and market share measures employed here.
References


### Table 1
Second-stage GLS profitability regression results

<table>
<thead>
<tr>
<th>Variable</th>
<th>ln($q$)</th>
<th></th>
<th></th>
<th>$\delta_k$</th>
<th></th>
<th></th>
</tr>
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<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
<td>(3)</td>
<td>(1)</td>
<td>(2)</td>
<td>(3)</td>
</tr>
<tr>
<td>$UN$</td>
<td>-0.850</td>
<td>-0.189</td>
<td>-0.494</td>
<td>-0.037</td>
<td>-0.015</td>
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<td></td>
<td>(5.75)</td>
<td>(1.31)</td>
<td>(1.31)</td>
<td>(4.97)</td>
<td>(1.94)</td>
<td>(1.97)</td>
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<tr>
<td>$CR$</td>
<td>0.885</td>
<td>0.581</td>
<td>0.643</td>
<td>0.048</td>
<td>0.035</td>
<td>0.027</td>
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<tr>
<td></td>
<td>(3.03)</td>
<td>(1.99)</td>
<td>(1.42)</td>
<td>(3.31)</td>
<td>(2.23)</td>
<td>(1.10)</td>
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<tr>
<td>$UN \cdot CR$</td>
<td>–</td>
<td>–</td>
<td>-0.251</td>
<td>–</td>
<td>–</td>
<td>0.017</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>(0.27)</td>
<td></td>
<td></td>
<td>(0.34)</td>
</tr>
<tr>
<td>$MS$</td>
<td>1.516</td>
<td>2.278</td>
<td>0.200</td>
<td>0.058</td>
<td>0.115</td>
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<tr>
<td></td>
<td>(2.57)</td>
<td>(4.44)</td>
<td>(0.23)</td>
<td>(1.97)</td>
<td>(4.23)</td>
<td>(0.53)</td>
</tr>
<tr>
<td>$UN \cdot MS$</td>
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<td>–</td>
<td>5.207</td>
<td>–</td>
<td>–</td>
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<td></td>
<td></td>
<td></td>
<td>(2.89)</td>
<td></td>
<td></td>
<td>(2.33)</td>
</tr>
<tr>
<td>$I-UN$</td>
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<td>-0.247</td>
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<td>–</td>
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<td>0</td>
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<td></td>
<td>(0.71)</td>
<td></td>
<td>(0.18)</td>
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</tr>
<tr>
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<td>yes</td>
<td>no</td>
<td>yes</td>
<td>yes</td>
</tr>
<tr>
<td>$R^2$</td>
<td>0.182</td>
<td>0.507</td>
<td>0.524</td>
<td>0.151</td>
<td>0.418</td>
<td>0.433</td>
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</table>

Sample size is 247. $t$-values in parentheses. Dependent variables are the firm dummy coefficients from first-stage regressions of pooled time-series/cross-section models of ln($q$) and $\delta_k$. Included in the first-stage regressions are year dummies and time-varying variables measuring the R&D stock divided by sales, a dummy for no reported R&D, advertising expenditures divided by sales, a dummy for no reported advertising, log of employment, log of net inflation-adjusted capital stock divided by employment, firm-specific 2-year growth rate in sales, industry-specific 4-year growth rate in sales, and import penetration in firms’ principal industry. The second-stage regression is weighted by the inverse of the standard error of the firm dummy coefficients.