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# Labor Unions, Operating Flexibility, and the Cost of Equity

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## Abstract

We study whether the constraints on firms' operations imposed by labor unions affect firms' costs of equity. The cost of equity is significantly higher for firms in more unionized industries. This effect holds after controlling for several industry and firm characteristics, is robust to endogeneity concerns, and is not driven by omitted variables. Moreover, the unionization premium is stronger when unions face a more favorable bargaining environment and is highly countercyclical. Unionization is also positively related to various measures of operating leverage. Our findings suggest that labor unions increase firms' costs of equity by decreasing firms' operating flexibility.

## I. Introduction

What determines firms' costs of equity? We study this question from the perspective of firms' operating flexibility. A number of recent studies have shown theoretically that the degree of firms' operating flexibility could have a nontrivial impact on their costs of equity,[1] yet direct empirical evidence in this regard is relatively scant. One of the main problems in testing these theories is that operating

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[1]Cochrane (1991), (1996), Jermann (1998), Gomes, Kogan, and Zhang (2003), and Zhang (2005) study the return implications of flexibility in capital inputs; Danthine and Donaldson (2002) focus on labor inputs, while Merz and Yashiv (2007) also examine the interactions between capital and labor.

flexibility is generally difficult to measure. In this paper, we consider a novel setting in which one can empirically evaluate the importance of operating flexibility, namely labor unions. The advantage of using this setting is that we can identify the source of cross-sectional variation in operating flexibility and study its effect on firms' costs of equity.

Although in a neoclassical framework labor inputs are fully flexible and thus have little effect on firm risk, in reality labor unions introduce an important friction that can have real effects on firms' operations. In particular, by reducing firms' operating flexibility, powerful unions can induce an increase in firms' systematic risk and thus in their costs of equity. This can happen in at least 2 ways. First, unions make wages sticky and layoffs costly, which increases firms' operating leverage due to labor and thus makes the adjustment of firms' labor stock more costly. Second, unions often intervene in firms' restructurings, for example by blocking plant closures, which can also make firms' adjustment of physical capital stock more costly. Hence, if unions reduce firms' operating flexibility, then investors should require higher returns on the capital they provide.

Whereas the scale of unions' actions has been generally diminishing over time, unions still play a significant role in firms' operations. As an example, through their negative impact on firms' operating flexibility, unions are considered largely responsible for the deterioration of the operating performance of firms in the auto industry. Among many actions, the United Auto Workers negotiated a contract with the "Big Three" auto producers that enforces a fixed wage irrespective of the performed work. Moreover, it effectively blocked the closure of some of General Motors' (GM) existing plants even though the closure decision was beneficial to GM's shareholders.[2]

Following the literature in labor economics, we use industry unionization rates as a proxy for unions' ability to affect firms' operations. Consistent with our prediction, we first show that firms in more unionized industries have higher costs of equity, measured by the implied cost of equity of Gebhardt, Lee, and Swaminathan (2001). Our results hold after we control for several firm-level characteristics, including revenue cyclicality, financial leverage, asset tangibility, age, sales growth, size, and return volatility, as well as for the industry capital intensity and its concentration. The economic magnitude of the results is considerable: A 1-standard-deviation increase in the unionization rate is associated with an increase in the implied cost of equity of about 1.23 percentage points per year.

Next, we use various additional identification strategies to account for alternative explanations. First, both unionization and the cost of equity may be correlated with a general observable industry characteristic such as the stage of the industry life cycle and thus this correlation may induce bias in the estimate of the effect of unionization on the cost of equity. We address this concern in that we explicitly control for a host of industry-level characteristics including age, "old-economy" and "new-economy" status, research and development (R&D)

[2]For a vivid discussion of the effect of unions on firms' operating flexibility, see Logan Robinson, "Why Detroit Has an Especially Bad Union Problem," *The Wall Street Journal* (Dec. 30, 2008).

expenses, advertising expenses, asset growth rate, and profitability. Our evidence indicates that such an industry effect is unlikely to explain our results.

Second, we address the concern that the unionization premium in the cost of equity might be driven by the omission of unobservable firm or industry characteristics correlated with unionization. To this end, we study the cross-sectional variation in the unionization premium. We find that the effect of unionization on the cost of equity is stronger when unions face a more favorable bargaining environment: in firms located in states with no right-to-work (RTW) laws, in firms located in states that are strongly influenced by the Democratic Party, and also in firms with more concentrated business operations. Hence, we conclude that omitted variables are unlikely to drive our results, as it is not obvious why these omitted variables would have a stronger effect on the cost of equity when unions face a more favorable bargaining environment.

A potential concern with our analysis is that endogeneity issues may affect the sign, magnitude, or statistical significance of our results. For example, the coefficient on unionization could be biased upwards if unions self-select into industries with high costs of equity. Conversely, the coefficient could be biased downwards if unions form in industries with low costs of equity. We use instrumental-variables techniques to address the endogeneity issue. We identify 2 industry-level instrumental variables, the percentage of female workers and the average age of workers, which are strongly correlated with unionization rates and satisfy the exclusion condition in that they do not directly affect the cost of equity. The instrumental-variables results are highly similar to the ordinary least squares (OLS) results and indicate that our findings are robust to endogeneity concerns.

Another way in which we can identify the effect of unionization on the cost of equity is to look at the time-series variation in the unionization premium. If labor unions increase a firm's cost of equity because they reduce the firm's operating flexibility, then the unionization premium should be countercyclical. The reason is that a unionized firm will be unable to adjust its operations during bad times, and it will remain overloaded with unproductive capital or labor, or committed expenses. Such an effect leads to an increase in the ex ante covariance between a firm's cash flows and market returns in economic downturns, which then translates into an increase in the firm's cost of equity. To this end, we study the time-series properties of the spread between the costs of equity of high- and low-unionization firms. Consistent with our story, we find that the unionization premium is higher when economic activity is low, that is, when the gross domestic product (GDP) growth, inflation rate, T-bill rate, and stock market return are low.

Last, we examine the direct effect of unionization on operating flexibility. Our argument is that unions reduce operating flexibility. Although operating flexibility is inherently difficult to measure, we rely on the elasticity of firms' total profits with respect to their sales, a traditional measure of operating leverage. Higher elasticity implies lower flexibility. We find that unionization is positively related to such an elasticity measure, and thus it is negatively related to operating flexibility. In addition, we separately study the effect of labor unions on

2 components of operating flexibility: that driven by labor and that driven by nonlabor factors. We find that unionization is associated with lower operating flexibility in both labor and nonlabor production inputs, which further supports our premise that unions have a pervasive impact on firms' ability to adjust both capital and labor inputs, and that the effect we document is not just a general labor effect.

In addition to our main analysis, we conduct a number of other tests. First, although we measure the cost of equity and other controls at the firm level, the lack of large-sample firm-level data makes it necessary to measure unionization rates at the industry level. As a result, our estimates may suffer from measurement error. Apart from the instrumental-variables approach that generally addresses this problem, we further expand our analysis by conducting our tests at the industry level, and using a small sample of firm-level unionization data. The industry-level results are similar in terms of magnitude and statistical significance to the baseline results. Also, the small-sample results based on firm-level unionization rates provide similar economic magnitudes to those using industry-level rates.

Second, we study the effect of unionization using an alternative measure of the cost of equity: the Fama-French (1993) cost of equity. We further decompose this measure into its 3 components: market beta, size beta, and value beta. We find that unionization is positively associated with the Fama-French cost of equity, and that the effect of unions works primarily through the book-to-market channel. Given that previous studies have linked book-to-market ratios to either operating leverage or costly adjustment of production factors, this evidence is consistent with the view that labor unions induce an increase in the cost of equity because they reduce firms' operating flexibility.

Third, we show that the positive relation between unionization and the cost of equity holds after we control for 2 measures of operating leverage: the elasticity of a firm's earnings with respect to its sales and a firm's labor stock. Hence, unionization captures aspects of operating flexibility that go beyond those captured by these measures of operating leverage. Fourth, given that other studies have shown that unionization is positively related to financial leverage, it is possible that financial leverage may explain part of the effect of unionization on the cost of equity. To address this issue, in all our regressions we explicitly control for differences in financial leverage. In addition, we repeat our tests using the unlevered cost of equity as the dependent variable and find that firms in more unionized industries also face a higher unlevered cost of equity. Hence, the effect of unions on the cost of equity we document is not driven by financial leverage. Finally, we find that unions' shareholder activism is unlikely to be the main channel through which unions affect firms' costs of equity.

The rest of the paper is structured as follows. Section II develops our main hypothesis and reviews related studies. Section III describes the data sources, defines the main variables, and examines the source of variation in the unionization data. Section IV relates unionization to the cross section of the cost of equity. Section V considers various identification issues. Section VI reports additional tests. Section VII concludes.

# II. Hypothesis Development and Related Studies

## A. Conceptual Framework

Previous research in labor economics shows that unionized labor has an impact on firms' operations that is quite distinct from the effect of nonunionized labor. For example, labor unions introduce substantial markups in the salaries of unionized workers (e.g., Lewis (1986)), they reduce profitability and equity values (e.g., Hirsch (1991b), Abowd (1989)), and they also affect corporate investment decisions (e.g., Connolly, Hirsch, and Hirschey (1986)).

We argue that the presence of powerful unions substantially reduces firms' operating flexibility.[3] One important way in which this can happen is that unions make wages more sticky and layoffs more costly. Specifically, employment contracts established through collective bargaining agreements exacerbate wage inflexibility, since multiyear contracts are the norm and wage adjustments are set out in advance or are allowed to vary only with inflation via escalator clauses (e.g., Mitchell (1985), Wunnava and Okunade (1996)). Moreover, job security and severance payments are central issues in unions' bargaining agendas (e.g., McLaughlin and Fraser (1984)), and collective bargaining agreements often include constraints on firms' ability to lay off workers. For example, senior workers at unionized firms enjoy more layoff protection than those at nonunionized firms (Abraham and Medoff (1984)). In addition, due to union opposition, flexible staffing arrangements that allow the firm to use and pay for workers' services only at the time the services are required are far less common in unionized sectors than in nonunionized sectors (e.g., Gramm and Schnell (2001)).

This discussion suggests that unions increase firms' share of fixed labor costs in total labor costs, that is, their operating leverage due to labor. Earlier work by Rubinstein (1973), Lev (1974), and Booth (1991) demonstrates that total operating leverage affects expected returns in the capital asset pricing model (CAPM). The notion of operating leverage in these studies is based on the importance of a firm's total fixed costs, without explicit reference to the sources of these costs. More recently, Danthine and Donaldson (2002) recognize that fixed labor costs may be an important source of operating leverage. In a representative agent model with fixed labor costs that arise from optimal risk-sharing contracts, these authors are able to generate operating leverage and better match the observed equity premium. This evidence suggests a link between labor operating leverage and systematic risk. As a result, shareholders of more unionized firms should require higher returns for bearing the extra risk associated with lower operating flexibility.

Unions can also increase the costs of adjusting the labor stock and physical capital over and above the costs firms face in the absence of unions. In our paper, we define costly adjustment as a firm's costly ability or inability to unwind its

[3]Firms typically face some costs of adjusting their operations in a changing economic environment. Some of these costs arise due to the nature of firms' technologies, production processes, and ability to adjust their inputs. Some examples include the costs of fine-tuning new equipment, the adaptation of workers to new machinery, the settling-in period for new workers, and the severance payments for workers who are fired. As a result of these costs, even in the absence of unions, firms' operations are not fully flexible.

labor stock and capital stock. The adjustment of the labor stock is more costly because unions impose frictions on the layoffs of workers (e.g., McLaughlin and Fraser (1984), Abraham and Medoff (1984)). In addition, unions often take an activist role by blocking restructurings and plant closings. Thus, unions may also generate costly adjustment in physical capital. Since the theory (e.g., Gomes et al. (2003), Zhang (2005), Cooper (2006), and Merz and Yashiv (2007)) shows that costly adjustment and irreversibility should be positively related to required stock returns in the cross section, we predict a positive relation between unionization and the cost of equity.

Ex ante, it is difficult to argue which aspects of operating flexibility are the most important to generate cross-sectional differences in the cost of equity. Both operating leverage and costly adjustment have a solid grounding in the theory. In addition, labor unions may reduce a firm's operating flexibility and thus increase risk in various other ways. For example, unions may oppose the adoption of new technologies (Dowrick and Spencer (1994)).[4] Overall, the important point is that the variation in the strength of labor unions across firms is associated with important differences in firms' operating flexibility. This leads us to the following general hypothesis:

*Hypothesis.* By reducing firms' operating flexibility, labor unions increase firms' costs of equity.

The precise identification of the operating flexibility channel is generally difficult, mostly because some aspects of operating flexibility are only measurable with noise (operating leverage) or cannot be measured directly (costly adjustment). In measuring operating leverage, at least a few points are noteworthy. First, current measures of operating leverage require the estimation of elasticity coefficients relating profits or costs to various demand proxies (e.g., Mandelker and Rhee (1984)). Second, by design, the measures are backward looking, but the concept of operating leverage in its nature is forward looking. Last, existing measures generally do not identify the precise sources of inflexibility, and thus they capture the total effect rather than only the labor effect. The advantage of our setting is that we can identify the cross-sectional variation in operating flexibility that originates in labor unions and study its effect on firms' costs of equity.

## B. Related Empirical Studies

Although empirically little is known about how operating flexibility affects the cost of equity, some empirical studies link a specific aspect of operating flexibility—operating leverage—to systematic risk. In particular, Mandelker and Rhee (1984) measure operating leverage as the elasticity of a firm's profits with respect to its sales and find a positive relation between this measure and market beta.

[4]Though probably the most important, operating flexibility may not be the only mechanism through which unions increase cost of equity. In particular, some studies associate higher unionization with higher financial leverage and thus implicitly with higher cost of equity. Other candidates are less obvious. For example, anecdotal evidence suggests that unions increase strike risk, which could affect cost of equity. Whether this risk is priced is unclear, largely because this risk could be potentially diversified away.

Novy-Marx (2011) derives new empirical predictions related to operating leverage and empirically tests them by focusing on differences between the within- and across-industry variations in operating leverage. In a recent study, García-Feijóo and Jorgensen (2010) find a positive association between operating leverage and the book-to-market equity ratio as well as between operating leverage and stock returns in the cross section.

In the specific context of labor markets, Rosett (2001) uses a firm's committed labor expenses as a proxy for operating leverage due to labor and finds that such a measure is positively related to a firm's equity risk measured by market beta and stock return volatility. Our paper differs from his in several important dimensions. First, he studies the effect of fixed labor costs on equity risk in highly unionized firms, but we study the effect of unionization on equity risk for all firms. Second, we measure equity risk using the implied cost of equity, which Pástor, Sinha, and Swaminathan (2008) show is a more accurate measure of the required return on equity. Third, we provide causal evidence that unions increase firms' costs of equity by reducing firms' operating flexibility. Fourth, we provide evidence in support of the theoretical prediction that the effect of operating flexibility on equity risk varies with the business cycle.

More broadly, our paper is related to studies that examine how unions affect firms' profitability by looking at their effect on firms' equity values (e.g., Ruback and Zimmerman (1984), Abowd (1989)). We extend this literature by showing that unions reduce the equity values because they increase equity risk and thus firms' costs of equity, independent of any effect they may have through their impact on profitability. Last, our paper is related to studies that incorporate labor into the asset pricing models (e.g., Jagannathan and Wang (1996), Santos and Veronesi (2006)). None of these papers, however, directly studies the implications of unionization for firms' discount rates.

## III. Data Sources and Variable Definitions

### A. Data Sources and Sample Selection

We obtain unionization data for the period 1984–2006 from the Union Membership and Coverage Database at www.unionstats.com, maintained by Barry Hirsch and David Macpherson. The data are compiled from the Current Population Survey based on the method of the Bureau of Labor Statistics. Hirsch and Macpherson (2003) provide details on the construction of this unique and comprehensive data set. Most of our additional data come from the Center for Research in Security Prices (CRSP), Compustat, and the Institutional Broker Estimates System. We compute industry workforce demographics using data from the Current Population Survey Labor Extracts. The data on political environment come from Dave Leip's Atlas of U.S. Presidential Elections, available at www.uselectionatlas.org. The macroeconomic data come from the St. Louis Federal Reserve Economic Database (GDP growth, inflation, and T-bill rate) and the University of Michigan's Survey of Consumers (expected inflation).

We begin with the set of firms in the CRSP-Compustat merged database. We include only companies with ordinary shares (CRSP share codes 10 or 11) and

exclude companies in the financial (Standard Industrial Classification (SIC) codes 6000–6999) and utilities (SIC codes 4900–4999) industries. Our final sample contains firms in the CRSP-Compustat merged database for which we can compute the cost of equity and have data on unionization rates and control variables. More precisely, the data we use in our main tests contain 31,307 observations corresponding to 5,580 firms during the period 1984–2006.

## B. Labor Unionization: Measures and Source of Variation in the Data

Although we can measure the cost of equity at the firm level, a typical problem in studies on labor unions is the lack of large-sample, firm-level unionization data. We follow previous work (e.g., Connolly et al. (1986)) and measure labor force unionization (UNION) as the percentage of employed workers in a firm's primary Census Industry Classification (CIC) industry covered by unions in collective bargaining with employers. Our data set comprises 188 CIC industries, which correspond roughly to 3-digit SIC industries.

Since our study relates the variation in industry unionization to the variation in the cost of equity, we document the source and amount of the variation in UNION. Table 1 reports summary statistics on the cross-sectional and time-series variation in unionization rates for the 10 most and the 10 least unionized industries.

TABLE 1

Unionization Rates for Selected Industries (1984–2006)

Table 1 reports the percentage of an industry's workers covered by unions in the collective bargaining with the firm. Based on the average industry unionization rates over the period 1984–2006, we identify the 10 most and 10 least unionized industries from the group of 188 Census Industry Classification (CIC) industries included in our data.

| Industry | Avg. 1984–2006 | 1984 | 1995 | 2006 |
|---|---|---|---|---|
| *Panel A. Highest Unionization Industries* | | | | |
| Railroads | 74.72% | 85.00% | 75.20% | 65.03% |
| Pulp, paper, and paperboard mills | 48.20% | 60.60% | 45.00% | 31.05% |
| Blast furnaces, steelworks, rolling and finishing mills | 47.44% | 60.20% | 50.80% | 25.06% |
| Motor vehicles and motor vehicle equipment | 44.26% | 61.60% | 44.00% | 26.32% |
| Air transportation | 41.74% | 38.90% | 40.80% | 45.86% |
| Primary aluminum industries | 39.05% | 48.60% | 34.80% | 26.17% |
| Railroad locomotives and equipment | 38.99% | 62.00% | 35.80% | 32.64% |
| Telephone (wire and radio) | 37.93% | 58.60% | 36.40% | 21.64% |
| Engines and turbines | 37.14% | 49.30% | 36.00% | 19.13% |
| Coal mining | 36.72% | 60.40% | 34.70% | 23.10% |
| *Panel B. Lowest Unionization Industries* | | | | |
| Apparel and accessory stores, exc. shoe | 2.61% | 4.10% | 2.80% | 1.41% |
| Sporting goods, toys, and hobby goods | 2.51% | 1.70% | 1.40% | 2.63% |
| Agricultural production, crops | 2.48% | 3.10% | 2.60% | 3.02% |
| Computer and data processing services | 2.32% | 4.30% | 2.40% | 2.03% |
| Book and stationery stores | 2.25% | 4.80% | 0.70% | 1.77% |
| Eating and drinking places | 2.19% | 4.00% | 2.40% | 1.58% |
| Sporting goods, bicycles, and hobby stores | 2.13% | 3.30% | 1.80% | 0.64% |
| Household appliances, TV, and radio stores | 1.83% | 2.50% | 1.20% | 0.90% |
| Shoe stores | 1.72% | 4.00% | 4.70% | 2.20% |
| Accounting, auditing, and bookkeeping services | 1.65% | 2.00% | 2.20% | 1.70% |

In the cross section, we observe a significant variation in unionization rates across different industries. Railroads, pulp and paper, steelworks, and motor vehicles are among the most unionized industries, with average rates above 44%

during the period 1984–2006. In contrast, small retail industries have an average unionization rate of about 2%. In the time series, we see a decreasing trend in unionization rates over time, with aggregate union coverage decreasing from 22.7% in 1984 to 10.4% in 2006. We further inspect the variation in unionization rates by estimating a regression of unionization rates on CIC industry fixed effects. We find that 89% of the total variation in unionization rates in our panel data set can be attributed to the between-industry variation and 11% accounts for the within-industry variation. Hence, we conclude that the primary source of variation in our data originates in the cross-sectional differences in the industry unionization rates. We use this variation to identify our results in the subsequent sections.

## C. Variables for the Cost of Equity Regressions

The dependent variable in our empirical model is the implied cost of equity (ICOE). It is obtained by assuming a valuation model and inferring the cost of equity using equity prices and other variables in the model, such as future cash flows. In our study, we follow Gebhardt et al. (2001) and calculate the ICOE using the residual income model. While several methods are currently available, we choose their method because it has received the most acclaim in academic research. For example, Guay, Kothari, and Shu (2006) argue that the ICOE is the best measure among those they survey. Similarly, this measure is used in other studies of the cost of equity (e.g., Lee, Ng, and Swaminathan (2009), Pástor et al. (2008)). All of these studies also provide the detailed methodology to construct this variable.

In addition to our test variable (UNION), we include several control variables. Sales beta (SALESBETA) is the cyclicality of revenues in a firm's CIC industry, which we compute, using quarterly data, as the slope from the full-sample time-series regression of changes in log industry net sales over the 1-year period on log GDP growth. Financial leverage (FINLEV) is the ratio of total liabilities (Item 181) to book assets (Item 6). The ratio of fixed assets to total assets (FA/TA) is net property, plant, and equipment (Item 8) over book assets. The industry capital to labor ratio (INDKL) is the total amount of net fixed assets in the firm's CIC industry divided by the total number of employees in that industry (based on Item 29). We divide the ratio by 1,000 to express it in millions of dollars of physical capital per employee. FIRMAGE is the natural logarithm of the number of years since a firm first appeared in CRSP. LOGSALGR is the change in the natural logarithm of a firm's net sales (Item 12). LOGASSETS is the natural logarithm of a firm's book assets. VOLAT is the standard deviation of daily stock returns during the calendar year. INDCONC is a Herfindahl index of sales concentration in a firm's CIC industry, which we average over the past 3 years to minimize the influence of potential data errors. To mitigate the impact of outliers on our results, we Winsorize SALESBETA at the 5% level, and both FINLEV and VOLAT at the 1% level.

Table 2 reports summary statistics. The mean and median ICOE in our sample of 5,580 firms are 10.1% and 11.0%, respectively, with a standard deviation of 5.7%. The distribution of unionization rates is slightly skewed to the right, as

the average unionization rate of 12.5% is higher than the median value of 8.8%. Also, the distribution is quite dispersed, with a standard deviation of 12.2%. All other variables have properties that are similar to those reported in other studies.

TABLE 2

Summary Statistics for Main Variables

Table 2 reports summary statistics for the variables we use in the regressions of the implied cost of equity (ICOE) on unionization rates and control variables. The sample covers the period 1984–2006 and contains 5,580 firms and a total of 31,307 firm-year observations (financial institutions and utilities are excluded from the sample). ICOE is the implied cost of equity of Gebhardt et al. (2001); UNION is union coverage at the Census Industry Classification (CIC) industry level; SALESBETA is the cyclicality of revenues in a firm's CIC industry, computed using quarterly data as the slope from a full-sample time-series regression of changes in log industry net sales over the 1-year period on log GDP growth; FINLEV is book leverage defined as total liabilities divided by total assets; FA/TA is net fixed assets divided by total assets; INDKL is average net fixed assets per employee in $000s within a CIC industry; FIRMAGE is the natural logarithm of the number of years a firm has been listed in CRSP; LOGSALGR is the growth in the natural logarithm of firm sales; LOGASSETS is the natural logarithm of total assets; VOLAT is the standard deviation of daily stock returns during the calendar year; INDCONC is the Herfindahl index measuring the concentration of sales within a CIC industry.

| Variable | Mean | Std. Dev. | Median | 5th Pctile | 95th Pctile |
|---|---|---|---|---|---|
| ICOE | 0.101 | 0.057 | 0.110 | 0.001 | 0.179 |
| UNION | 0.125 | 0.122 | 0.088 | 0.016 | 0.392 |
| SALESBETA | 1.862 | 1.280 | 1.514 | 0.529 | 4.430 |
| FINLEV | 0.187 | 0.185 | 0.153 | 0.000 | 0.529 |
| FA/TA | 0.293 | 0.220 | 0.239 | 0.036 | 0.761 |
| INDKL | 0.073 | 0.107 | 0.040 | 0.015 | 0.263 |
| FIRMAGE | 2.403 | 0.977 | 2.398 | 0.693 | 4.078 |
| LOGSALGR | 0.155 | 0.253 | 0.115 | −0.171 | 0.623 |
| LOGASSETS | 6.234 | 1.695 | 6.114 | 3.698 | 9.267 |
| VOLAT | 0.481 | 0.228 | 0.430 | 0.213 | 0.924 |
| INDCONC | 0.227 | 0.184 | 0.179 | 0.048 | 0.619 |

## IV. Unionization and the Implied Cost of Equity

We examine the univariate association between unionization rates and the ICOE. In each year, we sort companies into quintile portfolios based on their unionization rates. Next, we calculate the equal- and value-weighted cost of equity for each of the 5 portfolios. Last, we average the cost of equity in each portfolio across years. Quintile 1 includes firms with the lowest unionization rates, and Quintile 5 includes firms with the highest. Panel A of Table 3 presents the results.

The results indicate an increasing pattern in the cost of equity as we move from the lowest- to the highest-unionization portfolio for both the equal- and value-weighted portfolios. The average unionization rate varies from about 2.1% for Quintile 1 to about 31% for Quintile 5. The difference in the ICOE between Quintiles 5 and 1 is economically large and equals 3.11 percentage points per year for the equal-weighted portfolio and 2.93 percentage points per year for the value-weighted portfolio. Both numbers are highly statistically significant.

Although our objective is to test whether unionization is related to firms' costs of equity, our task is complicated by the fact that highly unionized firms may differ from other firms across several dimensions. Thus, we turn to a multivariate analysis in which we estimate the following regression model:

$$\text{(1)} \qquad \text{ICOE}_{ijt} = a_0 + a_1\,\text{UNION}_{jt} + a_2\,\text{CONTROLS}_{ijt} + \varepsilon_{ijt},$$

where $i$ indexes firms, $j$ indexes a firm's CIC industry, and $t$ indexes year. Our dependent variable is the ICOE. Our test variable (UNION) is the percentage of

TABLE 3

Unionization and the Implied Cost of Equity

Panel A of Table 3 reports portfolio sorts. For each year, we sort firms into quintile portfolios based on their unionization rate. We then compute the equal- and value-weighted implied cost of equity (ICOE) for each quintile portfolio, and subsequently take the average for each quintile across years. The last 2 columns report the differences in means between Quintile 5 and Quintile 1 and their corresponding *t*-statistics. Panel B reports the regression results. The dependent variable is ICOE. UNION is the Census Industry Classification (CIC)-industry unionization rate; SALESBETA is the cyclicality of revenues in a firm's CIC industry, computed using quarterly data as the slope from a full-sample time-series regression of changes in log industry net sales over the 1-year period on log GDP growth; FINLEV is book leverage, which we define as total liabilities divided by total assets; FA/TA is net fixed assets divided by total assets; INDKL is average net fixed assets per employee in $000s within a CIC industry; FIRMAGE is the natural logarithm of the number of years a firm has been listed in CRSP; LOGSALGR is the growth in the natural logarithm of firm sales; LOGASSETS is the natural logarithm of total assets; VOLAT is the standard deviation of daily stock returns during the calendar year; and INDCONC is the Herfindahl index measuring the concentration of sales within a CIC industry. All regressions include year and major SIC-division fixed effects. The absolute values of *t*-statistics (in parentheses) are based on robust standard errors clustered at the CIC industry level. *, **, and *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

*Panel A. Portfolio Sorts*

| | Unionization Quintile | | | | | | |
|---|---|---|---|---|---|---|---|
| Statistic | Q1 | Q2 | Q3 | Q4 | Q5 | Q5 – Q1 | *t*-Stat. |
| Equal-weighted ICOE (%) | 9.378 | 9.673 | 10.640 | 11.355 | 12.488 | 3.110 | 5.81 |
| Value-weighted ICOE (%) | 7.621 | 8.024 | 8.196 | 10.025 | 10.550 | 2.929 | 5.23 |
| Unionization (%) | 2.095 | 4.966 | 9.622 | 14.553 | 31.028 | | |

*Panel B. Regressions*

| Explanatory Variable | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| UNION | 0.126*** | 0.127*** | 0.120*** | 0.119*** | 0.101*** |
| | (4.15) | (3.97) | (3.87) | (3.93) | (3.60) |
| SALESBETA | | 0.004* | 0.004* | 0.004* | 0.003* |
| | | (1.67) | (1.85) | (1.85) | (1.74) |
| FINLEV | | 0.028*** | 0.024** | 0.025** | 0.025*** |
| | | (2.66) | (2.38) | (2.56) | (2.79) |
| FA/TA | | | 0.022** | 0.020** | 0.016* |
| | | | (2.17) | (2.08) | (1.82) |
| INDKL | | | −0.022 | −0.017 | −0.012 |
| | | | (0.44) | (0.35) | (0.26) |
| FIRMAGE | | | | 0.001 | −0.001 |
| | | | | (0.58) | (1.44) |
| LOGSALGR | | | | −0.014*** | −0.012*** |
| | | | | (3.03) | (2.79) |
| LOGASSETS | | | | −0.001 | −0.002*** |
| | | | | (1.28) | (3.11) |
| VOLAT | | | | | −0.031*** |
| | | | | | (5.01) |
| INDCONC | | | | | 0.047*** |
| | | | | | (3.97) |
| Intercept | 0.087*** | 0.066*** | 0.059*** | 0.065*** | 0.057** |
| | (6.69) | (4.17) | (3.46) | (3.77) | (2.55) |
| No. of observations | 31,307 | 31,307 | 31,307 | 31,307 | 31,307 |
| Adjusted $R^2$ | 0.28 | 0.29 | 0.29 | 0.30 | 0.33 |

workers in the industry covered by unions. Current theories predict that firms' costs of equity should be affected by firms' revenue cyclicality (SALESBETA) and financial leverage (FINLEV). We also include 2 variables that are related to the production technology: asset tangibility (FA/TA) and the industry-level capital intensity (INDKL). We add in other firm characteristics that earlier research has found to be correlated with the cost of equity, including FIRMAGE, LOGSALGR, LOGASSETS, VOLAT, and INDCONC. We control for differences in the cost of

equity across broadly defined industries by including major SIC-division industry fixed effects. Specifically, we follow Kahle and Walkling (1996) and use 2-digit SIC codes to aggregate industries into 10 major SIC divisions A–J.[5] Moreover, we control for differences in the cost of equity across time by including year fixed effects.

Throughout the paper, we estimate our empirical models using the conservative method of running pooled (panel) OLS regressions and calculating the standard errors by clustering at the CIC industry level (using our 188 CIC industry groupings). The clustering of standard errors at the CIC industry level responds to the concern that, conditional on the independent variables, the errors may be correlated within industry groupings, especially given that UNION is an industry-level variable. Clustered errors assume that observations are independent across industries but not necessarily independent within industries. Clustering at the industry level addresses the concern that residuals may be serially correlated within a firm, and also the concern that the residuals may be correlated across firms within the industry in the same or different periods of time. In addition, clustering accounts for heteroskedasticity in the residuals.

The results in Panel B of Table 3 show a positive and statistically significant relation between unionization and the cost of equity. In column (1), we regress ICOE on UNION and major SIC-division and year fixed effects. The coefficient on UNION equals 0.126 and is statistically significant at the 1% level. In columns (2)–(5), we sequentially add SALESBETA, FINLEV, FA/TA, INDKL, FIRMAGE, LOGSALGR, LOGASSETS, VOLAT, and INDCONC. Although including these controls reduces the magnitude of the coefficient on UNION to 0.101, the coefficient remains statistically significant. In terms of economic significance, a 1-standard-deviation increase in the unionization rate increases the cost of equity by approximately 1.23 percentage points per year. The coefficients on SALESBETA, FINLEV, FA/TA, and INDCONC are positive, and the coefficients on LOGSALGR, LOGASSETS, and VOLAT are negative. They are all statistically significant. In contrast, INDKL and FIRMAGE are not correlated with the cost of equity. Thus, consistent with unions reducing firms' operating flexibility, we find a positive relation between unionization rates and the cost of equity.

## V. Identification Tests

### A. Omitted Industry Characteristics

It is possible that both unionization rates and the cost of equity may be correlated with some general industry characteristic, and that what we really capture is an effect largely due to an omitted industry characteristic rather than unions per se. To preserve the parsimony of the model, our main specification in equation (1) considers only some characteristics of that type, such as industry capital-to-labor ratio, industry concentration, and major SIC-division fixed effects. However, our

[5]All our results remain similar if we replace these effects with 1-digit SIC industry fixed effects.

concern is that we might still be missing some other important observables. In particular, the unionization premium could be no more than an industry "life cycle" effect, in that mature industries tend to be more unionized, they might be more risky, and they might have higher costs of equity.

We address this concern in 3 ways. First, in our baseline specification we control for firm age, firm sales growth, and major SIC-division fixed effects. Second, we look in our data for evidence that mature industries might be more unionized. This is not always the case. For example, in our sample, department stores, oil and gas, drugs, and plastics are all mature, established industries, yet they have very low-unionization rates. Last, since this evidence is somewhat circumstantial, as some mature industries are highly unionized, we further turn to more formal regression evidence.

To this end, in Table 4 we extend our baseline specification and further control for a number of variables that proxy for the stage of the industry's life cycle. They include industry age (INDLOGAGE), old-economy (OLDECON) and new-economy (NEWECON) status, R&D expenses (INDRDEXP), advertising expenses (INDADVEXP), 1-year asset growth rates (INDGRWTH), and industry profitability (INDPROFIT).[6] INDLOGAGE is the natural logarithm of the age of the oldest company in the firm's CIC industry. OLDECON is an indicator variable equal to 1 if the firm operates in an old-economy industry, and 0 otherwise. Following Ittner, Lambert, and Larcker (2003), old-economy industries are defined as industries with SIC codes less than 4000 that are not in the computer, software, Internet, telecommunications, or networking industries. Likewise, NEWECON is an indicator variable equal to 1 if the firm operates in a new-economy industry, and 0 otherwise. New-economy industries include firms in the computer, software, Internet, telecommunications, or networking industries. INDRDEXP is the median ratio of R&D expenses to sales in a firm's CIC industry. INDADVEXP is the median ratio of advertising expenses to sales in a CIC industry. INDGRWTH is the median 1-year growth in assets in a firm's CIC industry. INDPROFITS is the median return on assets (ROA) in a firm's CIC industry.

The results in Table 4 indicate that the unionization premium cannot be attributed to the stage of the industry's life cycle. In fact, including all the industry controls in the ICOE specification slightly increases the coefficient on UNION from 0.101 (in column (5) of Panel B in Table 3) to 0.109, and increases the $t$-statistic from 3.60 to 3.92. Including the additional industry variables increases the $R^2$ by 2 percentage points, which suggests that these variables do capture some of the relevant variation in the cost of equity. Among the industry characteristics we consider, INDLOGAGE seems to be most closely related to the cost of equity. Also, NEWECON, INDADVEXP, and INDPROFIT are statistically significant. All other characteristics do not seem to add much explanatory power to our regressions.

[6]One could argue that what we capture in our analysis is a pure value effect, that is, firms in more unionized industries may simply be value stocks or value industries. In unreported results, we find that controlling for the book-to-market ratio does not affect our results. However, given that our cost of equity measure is by construction related to the book-to-market ratio, we do not include this control in our main tests.

TABLE 4

Unionization and the Implied Cost of Equity: Additional Industry Controls

In Table 4, the dependent variable is the implied cost of equity (ICOE). UNION is the Census Industry Classification (CIC)-industry unionization rate; SALESBETA is the cyclicality of revenues in a firm's CIC industry, computed using quarterly data as the slope from a full-sample time-series regression of changes in log industry net sales over the 1-year period on log GDP growth; FINLEV is book leverage, which we define as total liabilities divided by total assets; FA/TA is net fixed assets divided by total assets; INDKL is average net fixed assets per employee in $000s within a CIC industry; FIRMAGE is the natural logarithm of the number of years a firm has been listed in CRSP; LOGSALGR is the growth in the natural logarithm of firm sales; LOGASSETS is the natural logarithm of total assets; VOLAT is the standard deviation of daily stock returns during the calendar year; and INDCONC is the Herfindahl index measuring the concentration of sales within a CIC industry. The new industry-level control variables are as follows. INDLOGAGE is the logarithm of the age of the oldest firm in a CIC industry. OLDECON is an indicator variable equal to 1 if a firm operates in an "old-economy" industry, and 0 otherwise. We define old-economy industries as industries with SIC codes less than 4000 that are not the computer, software, Internet, telecommunications, or networking industries. NEWECON is an indicator variable equal to 1 if a firm operates in a "new-economy" industry, and 0 otherwise. We define new-economy industries to comprise the computer, software, Internet, telecommunications, and networking industries. INDRDEXP is the median ratio of R&D expenses to sales in a CIC industry. INDADVEXP is the median ratio of advertising expenses to sales in a CIC industry. INDGRWTH is the median 1-year growth in the logarithm of firm assets in a CIC industry. INDPROFIT is the median ROA in a CIC industry. All regressions include year and major SIC-division fixed effects. The full set of control variables in our benchmark specification (column (5) of Panel B in Table 3) is omitted for brevity. The absolute values of *t*-statistics (in parentheses) are based on robust standard errors clustered at the CIC industry level. *, **, and *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

| Explanatory Variable | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| UNION | 0.107*** | 0.103*** | 0.105*** | 0.109*** |
| | (3.88) | (3.53) | (3.69) | (3.92) |
| All columns include the full vector of control variables used in column (5) of Panel B, Table 3: SALESBETA, FINLEV, FA/TA, INDKL, FIRMAGE, LOGSALGR, LOGASSETS, VOLAT, and INDCONC. | | | | |
| *Additional Industry Controls* | | | | |
| INDLOGAGE | −0.015*** | −0.013*** | −0.012*** | −0.011*** |
| | (3.29) | (3.12) | (2.93) | (2.83) |
| OLDECON | | −0.000 | −0.000 | 0.000 |
| | | (0.04) | (0.03) | (0.01) |
| NEWECON | | −0.014** | −0.013** | −0.012** |
| | | (2.53) | (2.34) | (2.14) |
| INDRDEXP | | | −0.002 | −0.001 |
| | | | (0.74) | (0.66) |
| INDADVEXP | | | 0.335** | 0.252** |
| | | | (2.44) | (2.16) |
| INDGRWTH | | | 0.045** | 0.022 |
| | | | (2.35) | (1.28) |
| INDPROFIT | | | | 0.144*** |
| | | | | (3.15) |
| No. of observations | 31,307 | 31,307 | 31,307 | 31,307 |
| Adjusted $R^2$ | 0.34 | 0.34 | 0.34 | 0.35 |

## B. Cross-Sectional Variation in the Unionization Premium

Another way to identify the effect of unions on the cost of equity is to study the cross-sectional variation in the unionization premium that arises from differences in the bargaining environment. Such differences could affect the strength of unions. We posit that in more favorable bargaining situations, labor unions should be more able to constrain firms' decisions and thus increase operating risk, resulting in a higher ICOE. To test this prediction, we estimate the following regression model:

$$\text{(2)} \quad \text{ICOE}_{ijt} = a_0 + a_1\,\text{UNION}_{jt} + a_2\,\text{ENV}_{ijt} + a_3\,\text{UNION}_{jt} \times \text{ENV}_{ijt} + b\,\text{CONTROLS}_{ijt} + \varepsilon_{ijt}.$$

We alternately include 3 variables (ENV) and their interactions with UNION. These variables are proxies for the environment in which the bargaining between the firm and the union takes place. Our first proxy is RTW, an indicator variable equal to 1 if the firm's operations are principally located in a state with right-to-work laws. Our second proxy is DEMOCRAT, an indicator variable equal to 1 if the firm's operations are located in a state under the influence of the Democratic Party. We define a Democratic state as one in which the Democratic Party has won the majority of electoral votes in the most recent presidential election. Our last proxy is BUSCONC, the Herfindahl index measuring the concentration of a firm's sales across its business segments.

Our reasoning for using these variables is as follows: First, we argue that unions are less powerful in states with RTW laws. This is because these laws prohibit agreements between unions and employers that make the membership or payment of union fees a condition of employment, which then introduces a free-rider problem in unions' actions (e.g., Ellwood and Fine (1987)). Hence, we predict that unions representing workers in states with RTW laws should have less bargaining power, which implies a negative coefficient on UNION $\times$ RTW. Second, in the U.S. the Democratic Party tends to favor unions and thus to enhance the effectiveness of unions' actions. Historically, unions became identified with the Democratic Party during the Great Depression and remain so today (e.g., Lipset and Katchanovski (2001)). Therefore, unions representing workers in states under the influence of the Democratic Party are more likely to enjoy support from the Party and thus to have stronger bargaining power. Hence, we predict a positive coefficient on UNION $\times$ DEMOCRAT. Last, Rose (1991) shows, both theoretically and empirically, that firms with more diversified business operations possess a bargaining advantage over unions because they can use their "deep pockets" to cross-subsidize strikes or any other costs related to unions' activity. Thus, we predict a positive coefficient on UNION $\times$ BUSCONC.

Our controls include SALESBETA, FINLEV, FA/TA, INDKL, FIRMAGE, LOGSALGR, LOGASSETS, VOLAT, INDCONC, INDLOGAGE, OLDECON, NEWECON, INDRDEXP, INDADVEXP, INDGRWTH, and INDPROFIT. In each regression, we also include the major SIC-division and year fixed effects, and cluster standard errors at the CIC industry level.

Panel A of Table 5 reports the results. To facilitate interpretation, we demean UNION and the ENV variables before forming interaction terms in equation (2). The coefficients on the control variables are omitted for brevity. The coefficient of interest is $a_3$, which represents the effect of unionization on the cost of equity, conditional on the bargaining environment. We find results consistent with our predictions. The coefficient on UNION $\times$ RTW is negative, and the coefficients on UNION $\times$ DEMOCRAT and UNION $\times$ BUSCONC are both positive. All 3 coefficients are statistically significant. Thus, the effect of labor unions on the cost of equity is weaker in states with RTW laws. Likewise, the effect is stronger when unions operate in states under the influence of the Democratic Party, or when a firm's operations are more concentrated across business segments.

Our results in Panel A of Table 5 suggest that the effect of unionization on the cost of equity is unlikely to be driven by the omission of unobservable firm or industry characteristics that are correlated with unionization rates and affect the

cost of equity. In fact, if one believes that the unionization premium in the cost of equity is driven by omitted unobservable variables correlated with unionization, then one would have to explain why these omitted variables have a stronger effect on the cost of equity when unions face a more favorable bargaining environment.

In addition, specification (2) offers a suitable setting to address the concern that unobserved time-invariant industry characteristics may drive our results. An ideal approach to deal with such a problem would be to introduce CIC

TABLE 5

Cross-Sectional Variation in the Unionization Effect

In Table 5, the dependent variable is the implied cost of equity (ICOE), which is regressed on unionization (UNION), interaction terms between UNION and proxies for the nature of the bargaining environment, and all the control variables included in Table 4, which include all those in our benchmark specification in column (5) of Panel B in Table 3 plus all the additional industry-level control variables from Table 4. The interacting variables are RTW, a dummy variable equal to 1 if the firm's operations are principally located in a state with right-to-work laws, and 0 otherwise; DEMOCRAT, a dummy variable equal to 1 if the firm's operations are principally located in a Democratic Party state, and 0 otherwise, where a state is considered to be a Democratic Party State for the 4 consecutive years following a presidential election if the majority of electoral votes in the state were cast for Democrats; BUSCONC is the Herfindahl index measuring the concentration of a firm's sales across its business segments. All interacted variables (UNION, RTW, DEMOCRAT, and BUSCONC) are first demeaned. The control variables are SALESBETA, the cyclicality of revenues in a firm's Census Industry Classification (CIC) industry, computed using quarterly data as the slope from a full-sample time-series regression of changes in log industry net sales over the 1-year period on log GDP growth; FINLEV, book leverage, which we define as total liabilities divided by total assets; FA/TA, net fixed assets divided by total assets; INDKL, average net fixed assets per employee in $000s within a CIC industry; FIRMAGE, the natural logarithm of the number of years a firm has been listed in CRSP; LOGSALGR, the growth in the natural logarithm of firm sales; LOGASSETS, the natural logarithm of total assets; VOLAT, the standard deviation of daily stock returns during the calendar year; INDCONC, the Herfindahl index measuring the concentration of sales within a CIC industry; and INDLOGAGE, the logarithm of the age of the oldest firm in a CIC industry. OLDECON is an indicator variable equal to 1 if a firm operates in an "old-economy" industry, and 0 otherwise. We define old-economy industries as industries with SIC codes less than 4000 that are not the computer, software, Internet, telecommunications, or networking industries. NEWECON is an indicator variable equal to 1 if a firm operates in a "new-economy" industry, and 0 otherwise. We define new-economy industries to comprise the computer, software, Internet, telecommunications, and networking industries. INDRDEXP is the median ratio of R&D expenses to sales in a CIC industry; INDADVEXP is the median ratio of advertising expenses to sales in a CIC industry; INDGRWTH is the median 1-year growth in the logarithm of firm assets in a CIC industry; and INDPROFIT is the median ROA in a CIC industry. All regressions include year fixed effects. The coefficients on the control variables are omitted for brevity. The absolute values of *t*-statistics (in parentheses) are based on robust standard errors clustered at the CIC industry level. *, **, and *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively. In Panel A the regressions include major SIC-division fixed effects. In Panel B the regressions include CIC industry fixed effects corresponding to each of the 188 CIC industries in our data.

*Panel A. OLS Regressions with Year and Major SIC-Division Fixed Effects*

| Explanatory Variable | (1) | (2) | (3) |
|---|---|---|---|
| UNION | 0.113*** | 0.123*** | 0.121*** |
| | (4.09) | (4.43) | (4.17) |
| RTW | 0.006*** | | |
| | (3.54) | | |
| UNION × RTW | −0.042** | | |
| | (2.37) | | |
| DEMOCRAT | | −0.006*** | |
| | | (4.28) | |
| UNION × DEMOCRAT | | 0.083*** | |
| | | (4.84) | |
| BUSCONC | | | −0.013* |
| | | | (1.93) |
| UNION × BUSCONC | | | 0.145*** |
| | | | (3.89) |
| All columns include the full set of control variables used in column (4) of Table 4: SALESBETA, FINLEV, FA/TA, INDKL, FIRMAGE, LOGSALGR, LOGASSETS, VOLAT, INDCONC, INDLOGAGE, OLDECON, NEWECON, INDRDEXP, INDADVEXP, INDGRWTH, and INDPROFIT. | | | |
| No. of observations | 30,398 | 30,398 | 30,449 |
| Adjusted $R^2$ | 0.35 | 0.35 | 0.35 |

*(continued on next page)*

TABLE 5 (continued)

Cross-Sectional Variation in the Unionization Effect

*Panel B. OLS Regressions with Year and CIC Industry Fixed Effects*

| Explanatory Variable | (1) | (2) | (3) |
|---|---|---|---|
| UNION | −0.038 | −0.024 | −0.027 |
| | (0.91) | (0.56) | (0.66) |
| RTW | 0.001 | | |
| | (1.05) | | |
| UNION × RTW | −0.015 | | |
| | (1.27) | | |
| DEMOCRAT | | −0.003** | |
| | | (2.36) | |
| UNION × DEMOCRAT | | 0.034*** | |
| | | (3.37) | |
| BUSCONC | | | −0.006 |
| | | | (1.35) |
| UNION × BUSCONC | | | 0.103*** |
| | | | (3.46) |
| All columns include the full set of control variables used in column (4) of Table 4: SALESBETA, FINLEV, FA/TA, INDKL, FIRMAGE, LOGSALGR, LOGASSETS, VOLAT, INDCONC, INDLOGAGE, OLDECON, NEWECON, INDRDEXP, INDADVEXP, INDGRWTH, and INDPROFIT. | | | |
| No. of observations | 30,398 | 30,398 | 30,449 |
| Adjusted $R^2$ | 0.51 | 0.51 | 0.51 |

industry-level fixed effects. Unfortunately, we are not able to do this in our main specification (1), since unionization rates are highly persistent, and thus we do not have enough time-series variation in unionization rates to identify the coefficient on UNION. The advantage of introducing CIC industry-level fixed effects in specification (2) is that the interaction terms UNION × ENV have much more time-series variation due to the varying nature of the interacting variables. The only exception is UNION × RTW, since there is virtually no change in RTW legislation during our sample period. Consequently, although with this approach we cannot precisely identify the coefficient $a_1$, we have enough statistical power to identify the coefficient $a_3$. Thus, we can evaluate whether introducing CIC-industry fixed effects in specification (2) impacts our inference based on the interaction terms UNION × ENV. Panel B of Table 5 presents the results.

As expected, when we introduce CIC industry-level fixed effects we are unable to precisely estimate the coefficient on UNION, which is no longer statistically significant. Importantly, the qualitative aspects of the coefficients on the interaction terms are similar to those reported in Panel A of Table 5. As before, the coefficients on UNION × DEMOCRAT and UNION × BUSCONC are positive and highly statistically significant, and the coefficient on UNION × RTW continues to be negative, as predicted, but it is no longer statistically significant. In sum, we conclude that our main results are unlikely to be driven by time-invariant unobserved industry characteristics.

## C. Are the Results Robust to Endogeneity Concerns?

Our OLS analysis reveals a positive relationship between unionization rates and firms' cost of equity that is both statistically and economically significant.

However, it is possible that our OLS estimates may be biased due to endogeneity, a correlation between UNION and the error term in equation (1). Three different factors may potentially contribute to this effect: reverse causality, omitted unobserved characteristics, and measurement error. In particular, the magnitude of the effect would be overstated if workers unionize more in industries that are more risky. Conversely, if unions tend to become stronger in industries that are less risky, the effect would be understated. Another reason for a downward bias is measurement error, since we use industry-level unionization rates to proxy for firm-level unionization rates.

Since endogeneity could affect the sign, magnitude, or statistical significance of our results, we estimate a 2-stage least squares (2SLS) regression model to determine whether our OLS results are robust to these endogeneity concerns. In our context, good instruments are exogenous variables that are economically related to unionization but are uncorrelated with the error term of the 2nd-stage regression relating the cost of equity to unionization. To find such instruments, we consider factors that the labor economics literature has shown to empirically affect the demand for union services (e.g., Hirsch (1980), (1982)). We use 2 demographic characteristics of the industry labor force that we construct from the Census Population Survey: the logarithm of the fraction of female workers (FEMALE) and the logarithm of the average age of workers (WORKERAGE).

Specifically, female workers are less likely to unionize because women, on average, have less permanent attachment to the labor market and to specific internal job ladders than do men. In addition, the expected benefits (particularly nonwage benefits) from being a union member may be smaller for female workers and their costs of organizing may be greater. Thus, we expect a negative relation between FEMALE and UNION. Similarly, the industry workforce age structure may be related to the unionization level. Since senior workers have a relatively strong job attachment and low mobility, their expected benefits from unionization are likely to be high (in the form of institutionalized work rules, strict seniority systems, grievance procedures, and health and pension benefits), while their organizing costs may be low. Thus, we expect a positive relation between WORKERAGE and UNION. Simultaneously, we have no reason to believe that our instruments have a direct economic impact on firms' costs of equity, and thus they are unlikely to be correlated with the error term in the 2nd-stage regression.

Table 6 reports the 2SLS estimates of equation (1), in which we treat UNION as an endogenous variable that we instrument with FEMALE and WORKERAGE. Panel A reports the 1st-stage results relating UNION to FEMALE, WORKERAGE, and to the exogenous variables of the model. Panel B reports coefficients from the 2nd-stage regression of the cost of equity on the value of UNION predicted in the 1st-stage regression and the corresponding exogenous control variables. In both panels, column (1) includes the control variables in our benchmark specification, while column (2) further includes the additional industry-level variables we used in Section V.A.

Consistent with the hypothesized economic relation between our instruments and unionization, the 1st-stage results show that UNION is negatively related to FEMALE and positively related to WORKERAGE. Both coefficients are statistically significant at the 1% level. Also, our instruments have strong predictive

power. The partial $R^2$ of the 1st-stage regression indicates that the instruments explain around 24% of the variation in unionization rates, net of any effect they may have through other explanatory variables. In addition, the $F$-test rejects the null hypothesis that the coefficients on both instruments are jointly 0. In further support of our instruments, the test of overidentifying restrictions cannot reject the joint null hypothesis that our instruments are uncorrelated with the error term and are correctly excluded from the 2nd-stage regression.

The 2nd-stage results continue to provide strong evidence of a statistically significant and positive relation between UNION and ICOE. For the most comprehensive specification, a 1-standard-deviation increase in the unionization rate is associated with an increase in the cost of equity of about 1.4 percentage points.

TABLE 6

Unionization and the ICOE: Instrumental-Variables Estimation

Panel A of Table 6 reports the results from the 1st-stage regressions of unionization (UNION) on the instrumental variables (FEMALE and WORKERAGE) and the exogenous control variables included in the 2nd-stage regression. FEMALE is the logarithm of the percentage of female workers in a firm's Census Industry Classification (CIC) industry, and WORKERAGE is the logarithm of the average age of workers in a firm's CIC industry. SALESBETA is the cyclicality of revenues in a firm's CIC industry, computed using quarterly data as the slope from a full-sample time-series regression of changes in log industry net sales over the 1-year period on log GDP growth; FINLEV is book leverage, which we define as total liabilities divided by total assets; FA/TA is net fixed assets divided by total assets; INDKL is average net fixed assets per employee in $000s within a CIC industry; FIRMAGE is the natural logarithm of the number of years a firm has been listed in CRSP; LOGSALGR is the growth in the natural logarithm of firm sales; LOGASSETS is the natural logarithm of total assets; VOLAT is the standard deviation of daily stock returns during the calendar year; and INDCONC is the Herfindahl index measuring the concentration of sales within a CIC industry. Column (2) further includes the additional industry-level control variables used in Table 4. INDLOGAGE is the logarithm of the age of the oldest firm in a CIC industry. OLDECON is an indicator variable equal to 1 if a firm operates in an "old-economy" industry, and 0 otherwise. We define old-economy industries as industries with SIC codes less than 4000 that are not the computer, software, Internet, telecommunications, or networking industries. NEWECON is an indicator variable equal to 1 if a firm operates in a "new-economy" industry, and 0 otherwise. We define new-economy industries to comprise the computer, software, Internet, telecommunications, and networking industries. INDRDEXP is the median ratio of R&D expenses to sales in a CIC industry. INDADVEXP is the median ratio of advertising expenses to sales in a CIC industry. INDGRWTH is the median 1-year growth in the logarithm of firm assets in a CIC industry. INDPROFIT is the median ROA in a CIC industry. The partial $R^2$ is the fraction of the variation in UNION explained by the instruments, net of their effect through the predetermined variables. The cluster-robust $F$-statistic tests the joint statistical significance of the instruments. The test of overidentifying restrictions tests the joint null hypothesis that the instruments are uncorrelated with the error term and are correctly excluded from the 2nd-stage equation. Panel B reports the results from the 2nd-stage regressions of the implied cost of equity (ICOE) on unionization (UNION) and control variables, in which we treat unionization as the endogenous variable. The Hausman test examines whether the OLS and 2SLS coefficients on UNION are statistically different from each other. All regressions include year and major SIC-division fixed effects. The absolute values of $t$-statistics (in parentheses) are based on robust standard errors clustered at the CIC industry level. *, **, and *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

*Panel A. First-Stage Regressions of UNION and Validity of Instruments*

| Explanatory Variable | (1) | (2) |
|---|---|---|
| *Instruments* | | |
| FEMALE | −0.087*** | −0.085*** |
| | (4.43) | (4.21) |
| WORKERAGE | 0.580*** | 0.581*** |
| | (4.52) | (4.65) |
| *Unreported Control Variables Included in Regression* | | |
| All predetermined variables in main specification | Yes | Yes |
| All additional industry-level predetermined variables | No | Yes |
| Major SIC-division fixed effects | Yes | Yes |
| Year fixed effects | Yes | Yes |
| No. of observations | 31,307 | 31,307 |
| Adjusted $R^2$ | 0.58 | 0.59 |
| *Predictive Power of Instruments* | | |
| Partial $R^2$ | 0.244 | 0.232 |
| Robust $F$-statistic | 15.26 | 15.43 |
| $p$-value | 0.000 | 0.000 |
| *Test of Overidentifying Restrictions* | | |
| Hansen's $J$-statistic | 2.125 | 0.954 |
| $p$-value | 0.273 | 0.458 |

*(continued on next page)*

TABLE 6 (continued)

Unionization and the ICOE: Instrumental-Variables Estimation

*Panel B. Second-Stage Regressions of the ICOE on Unionization*

| Explanatory Variable | (1) | (2) |
|---|---|---|
| *Potentially Endogenous Instrumented Variable* | | |
| UNION | 0.114** | 0.160*** |
| | (1.99) | (2.94) |
| *Unreported Control Variables Included in Regression* | | |
| Predetermined variables in main specification | Yes | Yes |
| Additional industry-level predetermined variables | No | Yes |
| Major SIC-division fixed effects | Yes | Yes |
| Year fixed effects | Yes | Yes |
| No. of observations | 31,307 | 31,307 |
| *Hausman Test for the Effect of Unionization (Coefficient 2SLS = Coefficient OLS)* | | |
| Cluster-robust *t*-statistic | 0.260 | 1.330 |
| *p*-value | 0.799 | 0.186 |

This estimated effect is similar in magnitude to that obtained using the OLS coefficients. Moreover, although the magnitude of the estimated 2SLS coefficient is slightly higher than that obtained from the OLS tests reported in Table 3, using the Hausman test we cannot reject the null hypothesis that the 2SLS and OLS coefficients on UNION are the same. Hence, we conclude that our OLS estimates of the effect of unionization rates on the cost of equity are robust to endogeneity concerns.

Our analysis of the 1st-stage regression indicates that the correlation between FEMALE and UNION exhibits a decreasing pattern over time. Hence, it is important to examine whether this pattern affects the quality of our instrument and thus the reliability of our instrumental-variables results. To this end, we first repeat our analysis for 4 subperiods: 1984–1988, 1989–1994, 1995–2000, and 2001–2006. We find that the effect of unionization on the cost of equity is positive and statistically significant in all of the subperiods. Second, we repeat the analysis for each year and find that the coefficient on UNION is positive for all years except 1985, when it is negative but statistically insignificant. The positive coefficients are significant in 20 of the remaining 22 years.

## D. The Time-Series Variation in the Unionization Premium

Another way to identify the effect of unionization on the cost of equity is to look at the time-series variation in the unionization premium, that is, the difference in the ICOE of portfolios composed of high- and low-unionization firms. A growing theoretical literature directly links a firm's operating inflexibility to its cost of capital and provides a rationale for the countercyclical time-series variation in equity risk (e.g., Carlson, Fisher, and Giammarino (2004), Kogan (2004), Zhang (2005), and Cooper (2006)). The argument these studies put forward is that firms with inflexible operations are unable to adjust their operations during times of low demand for their products, when they are under more pressure to restructure their operations, and thus they remain loaded with unproductive

capital or labor, or committed expenses. Such an effect then leads to an increase in the ex ante covariance between a firm's cash flows and business conditions in economic downturns, which then translates into an increase in the firm's cost of equity. Given that labor unions reduce firms' operating flexibility these theories suggest that the unionization premium should be larger in periods of low economic activity.

Guided by this theory, we study the time-series properties of the unionization premium. Our prediction is that the unionization premium should be countercyclical. To this end, we estimate the following univariate time-series regression model using the 23 annual observations in our sample period:

$$\text{(3)} \quad \text{UNIONIZATION\_PREMIUM}_t = a_0 + a_1 \text{ BUSINESS\_CYCLE}_t + \varepsilon_t,$$

where UNIONIZATION_PREMIUM is the spread between the cost of equity of the portfolio of high-unionization firms (top quintile) and the cost of equity of the portfolio of low-unionization firms (bottom quintile). We consider both the equal- and value-weighted unionization premium. BUSINESS_CYCLE is a generic variable indicating market conditions. Following the asset pricing literature, our proxies for market conditions include contemporaneous and future values of GDP growth, inflation rate, and T-bill rate, as well as the contemporaneous value-weighted stock market return. GDPGROWTH is the year-over-year growth in the 4th quarter's GDP; INFLATION is the year-to-year change in the December Consumer Price Index; EXPINFLATION is the inflation rate expected by consumers as of December of the current year; TBILL is the 3-month T-bill rate as of December of the corresponding year; and MKTRET is the return on CRSP value-weighted index. Since we use 7 different business-cycle predictors, we estimate a total of 14 different regression models (1 set for equal- and 1 set for value-weighted premium). In each regression, the coefficient of interest is $a_1$, which measures whether the unionization premium varies with market conditions. Our standard errors are corrected for autocorrelation using the Newey-West (1987) adjustment.

Since our tests include only 23 annual observations, our tests may lack statistical power. Nevertheless, the results, reported in Table 7, show that the unionization premium is higher when current and future market conditions are weaker, that is, when GDP growth, inflation rate, T-bill rate, and market return are lower. The coefficient is statistically significant in each of the 14 models we consider. The magnitudes of the estimates also indicate that the business-cycle variation in the unionization premium is economically significant. Depending on the specification we consider, a 1-standard-deviation movement of our empirical proxies toward the market downturn is associated with a 0.7%–1.7% increase in the equal-weighted and a 0.5%–1.6% increase in the value-weighted unionization premiums. Given that the finding of a strongly countercyclical unionization premium is consistent with the view that the constraints on firms' operations imposed by labor unions are more likely to bind when economic activity is low, this time-series test provides further evidence that labor unions increase the cost of equity because they reduce firms' operating flexibility.

TABLE 7

Business Cycle Variation in the Unionization Premium

Table 7 reports the results of OLS time-series univariate regressions of the annual unionization premium (both equal- and value-weighted) on business cycle indicators. The dependent variable is the unionization premium, calculated as the difference between the equal- or value-weighted average implied cost of equity for firms in the highest unionization quintile and that for firms in the lowest unionization quintile. GDPGROWTH is the year-over-year growth in the 4th quarter's GDP; INFLATION is the year-over-year change in the December Consumer Price Index; EXPINFLATION is the inflation rate expected by consumers as of December of the current year; TBILL is the 3-month T-bill rate as of December of the corresponding year; and MKTRET is the annual return on the value-weighted CRSP index. The sample covers 23 years of data during the period 1984–2006. The absolute values of *t*-statistics (in parentheses) are calculated using standard errors obtained from a Newey-West (1987) procedure that accounts for any significant autocorrelation. *, **, and *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

| Explanatory Variable | Equal-Weighted Unionization Premium (1) | Value-Weighted Unionization Premium (2) |
|---|---|---|
| (1) $\text{GDPGROWTH}_t$ | −0.006*** | −0.004** |
| | (3.61) | (2.33) |
| (2) $\text{GDPGROWTH}_{t+1}$ | −0.005*** | −0.004** |
| | (3.45) | (2.67) |
| (3) $\text{INFLATION}_t$ | −0.008*** | −0.007*** |
| | (3.55) | (3.66) |
| (4) EXPINFLATION | −0.025*** | −0.021*** |
| | (5.05) | (6.04) |
| (5) $\text{TBILL}_t$ | −0.008*** | −0.007*** |
| | (6.40) | (4.89) |
| (6) $\text{TBILL}_{t+1}$ | −0.008*** | −0.007*** |
| | (5.93) | (3.57) |
| (7) $\text{MKTRET}_t$ | −0.042*** | −0.044*** |
| | (2.98) | (3.44) |

## E. Labor Unions and Operating Flexibility

Our hypothesis implies a negative relation between unionization and operating flexibility. Since an important aspect of operating flexibility is operating leverage, one would expect that unionization should increase operating leverage. To explore this issue, we follow Mandelker and Rhee (1984) and estimate total operating leverage as the elasticity of a firm's operating income after depreciation with respect to its sales, using the 15 most recent quarterly observations.[7] We refer to this measure as the Mandelker and Rhee degree of operating leverage (MRDOL), which we further Winsorize at the 5% level. Conceptually, a higher MRDOL is associated with higher operating leverage and thus with lower operating flexibility. We estimate the following multivariate regression model:

$$\text{MRDOL}_{ijt} = a_0 + a_1\ \text{UNION}_{jt} + a_2\ \text{CONTROLS}_{ijt} + \varepsilon_{ijt}, \tag{4}$$

where $i$ indexes firms, $j$ indexes a firm's CIC industry, and $t$ indexes year. Our vector of controls includes LOGASSETS, TOBQ, FA/TA, FINLEV, INDKL, INDLOGAGE, and INDCONC. TOBQ is Tobin's Q, defined as the market value of assets divided by the book value of assets. We also include a proxy for the firm's labor stock (LS), defined as the number of a firm's employees divided by total

[7]For firms with positive earnings before interest and taxes (EBIT), we calculate this elasticity by estimating the time-series quarterly regression of log EBIT on log SALES. For firms with at least one negative value of EBIT within a given series, we approximate the elasticity by estimating a similar regression of EBIT on SALES and then multiplying the coefficient on SALES by the ratio of average sales to average operating income calculated over the estimation period.

assets, to control for the importance of labor in firms' operations. The other variables are defined as before. Finally, we include major SIC-division and year fixed effects. All standard errors are clustered at the CIC industry groupings. Our prediction is that UNION should be positively related to MRDOL.

Consistent with the hypothesis that unionization reduces operating flexibility, our results in column (1) of Table 8 indicate that unionization is positively associated with MRDOL. The effect is statistically significant at the 1% level. Column (2) further indicates that including additional controls in our regression model reduces the magnitude of the coefficient on UNION, but the effect remains statistically significant at the 5% level.

TABLE 8

Unionization and Additional Measures of Operating Leverage

Table 8 reports the results from OLS regressions of 3 alternative measures of the degree of operating leverage (MRDOL, LDOL, and NLDOL) on unionization (UNION) and a set of control variables. MRDOL is the Mandelker and Rhee (1984) measure of operating leverage, calculated as the elasticity of a firm's operating income after depreciation with respect to its sales; LDOL is a measure of operating leverage due solely to labor costs (estimated as the elasticity of a firm's labor costs with respect to its sales); NLDOL is a measure of operating leverage due solely to nonlabor costs (estimated as the elasticity of a firm's nonlabor costs with respect to its sales). A higher MRDOL and a lower LDOL or NLDOL are associated with higher operating leverage. Our controls include the natural logarithm of total assets (LOGASSETS), Tobin's Q (TOBQ), the ratio of net fixed assets to total assets (FA/TA), book leverage (FINLEV), average net fixed assets per employee in $000s within a Census Industry Classification (CIC) industry (INDKL), the logarithm of the age of the oldest firm in a CIC industry (INDLOGAGE), the Herfindahl index measuring the concentration of sales within a CIC industry (INDCONC); and the labor stock (LS) defined as the number of a firm's employees divided by total assets. All regressions include year and major SIC-division fixed effects. The absolute values of *t*-statistics (in parentheses) are based on robust standard errors clustered at the CIC industry level. *, **, and *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

| | MRDOL | | LDOL | | NLDOL | |
|---|---|---|---|---|---|---|
| Explanatory Variable | (1) | (2) | (3) | (4) | (5) | (6) |
| UNION | 0.029*** | 0.016** | −0.527*** | −0.267*** | 0.036 | −0.231*** |
| | (3.54) | (2.27) | (5.24) | (3.08) | (0.44) | (3.10) |
| LOGASSETS | | 0.002*** | | −0.004 | | 0.023*** |
| | | (5.89) | | (0.47) | | (3.83) |
| TOBQ | | −0.001*** | | 0.008 | | −0.041*** |
| | | (4.08) | | (0.46) | | (3.77) |
| FA/TA | | 0.004** | | −0.030 | | 0.039 |
| | | (2.17) | | (0.43) | | (0.88) |
| FINLEV | | −0.003* | | 0.048 | | −0.056 |
| | | (1.69) | | (0.58) | | (0.97) |
| INDKL | | 0.001 | | −0.934*** | | 0.487*** |
| | | (0.10) | | (4.42) | | (2.69) |
| INDLOGAGE | | −0.002* | | 0.069* | | 0.021 |
| | | (1.87) | | (1.78) | | (0.86) |
| INDCONC | | 0.006** | | −0.041 | | 0.079 |
| | | (2.31) | | (0.59) | | (1.64) |
| LS | | 0.102** | | 1.240* | | 2.381*** |
| | | (2.07) | | (1.89) | | (4.01) |
| Intercept | 0.007* | 0.002 | 1.072*** | 0.913*** | 0.988*** | 0.768*** |
| | (1.69) | (0.31) | (10.41) | (6.26) | (34.92) | (6.92) |
| No. of observations | 92,338 | 92,338 | 2,876 | 2,876 | 2,876 | 2,876 |
| Adjusted $R^2$ | 0.01 | 0.02 | 0.14 | 0.19 | 0.04 | 0.15 |

To better understand the mechanism through which labor unions reduce operating flexibility, we further decompose operating leverage into its 2 components: operating leverage due to labor and operating leverage due to capital. Our premise is that labor unions should affect flexibility of both labor and capital inputs.

We define 2 variables: Operating leverage due to labor (LDOL) is the elasticity of a firm's labor and related expenses (Item 42) with respect to its sales. We calculate this elasticity by estimating the time-series regression model of log labor costs on log SALES using the 10 most recent annual observations. Similarly, operating leverage due to nonlabor costs (NLDOL) is the elasticity of a firm's nonlabor costs (cost of the goods sold plus selling, general, and administration expense minus labor and related expenses) to its sales. We Winsorize both LDOL and NLDOL at the 1% level. Since we look at the elasticity of costs rather than profits, conceptually, both higher LDOL and NLDOL are associated with lower operating leverage and thus with higher operating flexibility.

The specification of our model mirrors the regression model in (4), except that now our dependent variable is either LDOL or NLDOL. Our prediction is that the coefficient on UNION should be negative. Columns (3)–(6) of Table 8 present the results. We observe that unionization is negatively related to both LDOL and NLDOL. The coefficients obtained from multivariate regressions are significant at the 1% level in both cases. We also observe that unionization alone explains a more significant fraction of the variation in LDOL ($R^2$ equal to 5.4%) and significantly less so in NLDOL ($R^2$ equal to 1.1%), which confirms labor unions' primary role in driving operating flexibility due to labor. These results also provide more direct evidence that labor unions rather than just a general labor effect drive our main findings. Finally, they illustrate that labor unions have a pervasive impact on firms' ability to adjust both capital and labor inputs.

# VI. Additional Tests

## A. Measurement Error in the Unionization Rates

A potential limitation of our study is that unionization is measured at the industry level and not at the firm level. This limitation is not specific to our paper but is a generic problem of any large-sample study on labor unions. The consequence of such a problem is that our results may suffer from measurement error. In Section V.C, we use the instrumental-variables approach and show that measurement error does not significantly bias our coefficients. In this section, we further strengthen our analysis in 2 ways. First, we collapse all the data into CIC industry equal- and value-weighted averages and then conduct our tests at the industry level. Second, we provide small-sample evidence using unionization data at the firm level.

### 1. Industry-Level Analysis

Table 9 provides the estimates of the industry-level multivariate regression models, similar to those in equation (1), on a sample of 188 CIC industries over the period 1984–2006. Columns (1)–(3) are based on the equal-weighted industry variables and columns (4)–(6) use the value-weighted industry variables. In columns (1) and (4), we report the results from the OLS regressions, which give equal importance to each industry-level observation. Since this estimation approach ignores the relative importance of each industry in the entire population of firms in our sample, we further report the results from the weighted least squares

(WLS) regressions. Specifically, in columns (2) and (5), each industry-level observation is weighted by the number of firms in the CIC industry; in columns (3) and (6), each industry-level observation is weighted by the market capitalization of the firms in the CIC industry.

TABLE 9

Unionization and the ICOE: Industry-Level Regressions

In Table 9, the dependent variable is the Census Industry Classification (CIC) industry-level implied cost of equity (ICOE). UNION is the industry-level unionization rate. All firm-level variables are aggregated into industry-level variables using 2 different approaches. In columns (1)–(3) all firm-level variables are converted into CIC industry equal-weighted averages. In columns (4)–(6) all variables are converted into CIC industry value-weighted averages. Columns (1) and (4) report the results of OLS regressions. Columns (2) and (5) report the results of weighted least squares (WLS) where the weight on each industry-year observation is the number of firms in the industry in that year. Columns (3) and (6) report the results of WLS where the weights on each industry-year observations are the market capitalizations of firms in the industry in that year. SALESBETA is the cyclicality of revenues in a firm's CIC industry, computed using quarterly data as the slope from a full-sample time-series regression of changes in log industry net sales over the 1-year period on log GDP growth; FINLEV is book leverage, which we define as total liabilities divided by total assets; FA/TA is net fixed assets divided by total assets; INDKL is average net fixed assets per employee in $000s within a CIC industry; FIRMAGE is the natural logarithm of the number of years a firm has been listed in CRSP; LOGSALGR is the growth in the natural logarithm of firm sales; LOGASSETS is the natural logarithm of total assets; VOLAT is the standard deviation of daily stock returns during the calendar year; and INDCONC is the Herfindahl index measuring the concentration of sales within a CIC industry. All regressions include year and major SIC-division fixed effects. The absolute values of *t*-statistics (in parentheses) are based on robust standard errors clustered at the CIC industry level. *, **, and *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

| | Equal-Weighted Industry ICOE | | | Value-Weighted Industry ICOE | | |
|---|---|---|---|---|---|---|
| Explanatory Variable | (1) | (2) | (3) | (4) | (5) | (6) |
| UNION | 0.046*** | 0.068*** | 0.101*** | 0.059*** | 0.119*** | 0.166*** |
| | (3.25) | (2.79) | (3.66) | (3.54) | (3.12) | (3.04) |
| SALESBETA | 0.001 | 0.004** | 0.001 | −0.001 | 0.003 | 0.002 |
| | (0.42) | (2.19) | (0.35) | (0.49) | (1.36) | (0.53) |
| FINLEV | 0.040*** | 0.048** | 0.062* | 0.044*** | 0.082*** | 0.120*** |
| | (3.21) | (2.28) | (1.94) | (3.04) | (3.73) | (3.55) |
| FA/TA | 0.002 | 0.030 | 0.019 | 0.004 | 0.006 | 0.001 |
| | (0.20) | (1.39) | (0.52) | (0.31) | (0.35) | (0.03) |
| INDKL | −0.066** | −0.017 | −0.014 | −0.041 | 0.025 | 0.025 |
| | (2.30) | (0.34) | (0.88) | (1.38) | (0.56) | (1.17) |
| FIRMAGE | −0.000 | −0.003 | −0.004 | 0.002 | 0.003 | 0.008 |
| | (0.09) | (0.92) | (0.63) | (0.85) | (1.01) | (1.29) |
| LOGSALGR | −0.015** | −0.026 | −0.040* | −0.017*** | −0.020* | −0.029 |
| | (2.26) | (1.37) | (1.65) | (2.89) | (1.76) | (1.58) |
| LOGASSETS | −0.002 | −0.002 | 0.003 | −0.006*** | −0.011*** | −0.017*** |
| | (0.95) | (0.63) | (0.74) | (3.03) | (4.14) | (4.07) |
| VOLAT | −0.010 | −0.079*** | −0.068** | 0.018 | −0.001 | 0.037 |
| | (0.77) | (3.84) | (2.16) | (1.35) | (0.03) | (1.01) |
| INDCONC | 0.016* | 0.049*** | 0.052* | 0.017* | 0.044*** | 0.056* |
| | (1.70) | (3.65) | (1.84) | (1.66) | (3.06) | (1.74) |
| Intercept | 0.091*** | 0.065* | 0.027 | 0.103*** | 0.086*** | 0.079* |
| | (4.30) | (1.79) | (0.55) | (5.07) | (2.61) | (1.68) |
| Estimation Method | OLS | WLS (# firms) | WLS (Market Cap.) | OLS | WLS (# firms) | WLS (Market Cap.) |
| No. of observations | 3,400 | 3,400 | 3,400 | 3,400 | 3,400 | 3,400 |
| Adjusted $R^2$ | 0.36 | 0.58 | 0.58 | 0.36 | 0.59 | 0.55 |

We find that the industry-level results are similar to our baseline results: The coefficient on UNION is positive and statistically significant at the 1% level for all 6 specifications. In addition, the WLS specifications produce significantly larger point estimates. We believe these differences in estimates reflect the importance of using WLS in this context. Comparing the economic magnitude of the WLS

results from this table to that from our specification in Table 3, we can see that the magnitudes are very close to each other. The effect of unionization on the cost of equity is still economically large.

### 2. Firm-Level Unionization

A more direct way of measuring the relation between unionization and cost of equity would be to use firm-level data. However, since firms are not required to disclose the unionization rates of their labor force, no large data set is available. The only firm-level data have been collected through surveys and correspond to the 1970s and mid-1980s.

To this end, we use firm-level unionization rates obtained by Barry Hirsch (see Hirsch (1991a)) from survey data for a small sample of companies for 1972, 1977, and 1987. The 1977 and 1987 data were derived from his 1987 survey of manufacturing firms, and the 1972 data were collected in an independent 1972 Conference Board Survey. Since the cost of equity is available from 1984, our sample period is limited to 1987 only. Also, since the survey has been conducted for a small number of manufacturing firms randomly chosen by Hirsch, our sample includes only 212 firms.

Subsequently, we estimate the regression model in equation (1), with the exception that the unionization variable (FLUNION) is now measured at the firm level. Table 10 presents the results. The coefficient on FLUNION in the most comprehensive specification equals 0.047 and is statistically significant at the 1% level. In terms of economic magnitude, a 1-standard-deviation increase in the firm-level unionization rate is associated with an increase in the ICOE of about 1.13 percentage points per year. This magnitude is very close to the 1.3 percentage points obtained using industry-level unionization rates for the same year. This result is reassuring despite the fact that it has been obtained using a much smaller sample. Overall, the firm-level analysis bolsters our argument that the effect of unionization on the cost of equity we identify in other sections of the paper is distinct from an industry effect, and that the possible measurement error in industry-level unionization rates does not affect our inferences.

## B. Unionization, the FFCOE, and Loadings on the Factors

In our analysis, we use the ICOE. In this section, we consider an alternative measure: the cost of equity derived from the Fama-French (1993) 3-factor model (FFCOE). This model assumes that the cost of equity is a linear projection of returns on a set of 3 factors: market, size, and value. To obtain factor loadings, for each stock $j$ in year $t$ (between 1984 and 2004), we estimate the following time-series regression model using monthly data from year $t+1$ to year $t+3$:

$$r_j - r_f = \alpha_{jt} + \beta_{jt}^{m}\text{MKT} + \beta_{jt}^{h}\text{HML} + \beta_{jt}^{s}\text{SMB} + \varepsilon_j, \tag{5}$$

where $r_j - r_f$ is the monthly return on stock $j$ minus the risk-free rate, MKT is the excess return of the market portfolio over the risk-free rate, HML is the return difference between high and low book-to-market stocks, and SMB is the return

TABLE 10

Firm-Level Unionization and the ICOE

Table 10 reports the results from OLS regressions of the implied cost of equity (ICOE) on firm-level unionization (FLUNION) and a set of control variables. The sample consists of 212 firms in the year 1987 for which firm-level unionization rates are available. SALESBETA is the cyclicality of revenues in a firm's Census Industry Classification (CIC) industry, computed using quarterly data as the slope from a full-sample time-series regression of changes in log industry net sales over the 1-year period on log GDP growth; FINLEV is book leverage, which we define as total liabilities divided by total assets; FA/TA is net fixed assets divided by total assets; INDKL is average net fixed assets per employee in $000s within a CIC industry; FIRMAGE is the natural logarithm of the number of years a firm has been listed in CRSP; LOGSALGR is the growth in the natural logarithm of firm sales; LOGASSETS is the natural logarithm of total assets; VOLAT is the standard deviation of daily stock returns during the calendar year; and INDCONC is the Herfindahl index measuring the concentration of sales within a CIC industry. The absolute values of *t*-statistics (in parentheses) are based on heteroskedasticity-robust standard errors. *, **, and *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

| Explanatory Variable | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| FLUNION | 0.042*** | 0.040*** | 0.047*** | 0.047*** | 0.047*** |
| | (4.95) | (4.25) | (4.78) | (4.97) | (5.20) |
| SALESBETA | | 0.001 | −0.001 | −0.000 | −0.001 |
| | | (0.58) | (0.30) | (0.17) | (0.40) |
| FINLEV | | 0.045** | 0.058*** | 0.059*** | 0.069*** |
| | | (2.02) | (3.04) | (3.11) | (3.66) |
| FA/TA | | | 0.006 | 0.008 | 0.006 |
| | | | (0.28) | (0.43) | (0.35) |
| INDKL | | | −0.271*** | −0.268*** | −0.279*** |
| | | | (8.32) | (8.29) | (10.10) |
| FIRMAGE | | | | 0.002 | 0.002 |
| | | | | (0.50) | (0.51) |
| LOGSALGR | | | | 0.026* | 0.030* |
| | | | | (1.69) | (1.92) |
| LOGASSETS | | | | −0.001 | −0.001 |
| | | | | (0.64) | (0.78) |
| VOLAT | | | | | −0.048* |
| | | | | | (1.68) |
| INDCONC | | | | | 0.018 |
| | | | | | (1.38) |
| Intercept | 0.127*** | 0.116*** | 0.125*** | 0.122*** | 0.140*** |
| | (39.28) | (17.62) | (13.29) | (8.11) | (8.49) |
| No. of observations | 212 | 212 | 212 | 212 | 212 |
| Adjusted $R^2$ | 0.09 | 0.10 | 0.28 | 0.29 | 0.31 |

difference between small and large capitalization stocks (month subscripts are omitted for brevity). We then construct the FFCOE in year $t$ as follows:

$$\text{FFCOE}_t = \overline{r_f} + \hat{\beta}^m_{jt}\overline{\text{MKT}} + \hat{\beta}^h_{jt}\overline{\text{HML}} + \hat{\beta}^s_{jt}\overline{\text{SMB}}, \tag{6}$$

where $\overline{r_f}$, $\overline{\text{MKT}}$, $\overline{\text{HML}}$, and $\overline{\text{SMB}}$ are the average annualized returns of the risk-free asset and 3 Fama-French factors over the period 1926–2007. For simplicity, we refer to the loadings on these factors as MKTBETA, HMLBETA, and SMB-BETA, respectively.

We subsequently relate FFCOE to unionization rates using a multivariate regression framework. The specification of our empirical model mirrors that in equation (1). Our findings, presented in column (1) of Table 11, are consistent with the results for ICOE. The coefficient on UNION is positive and statistically significant at the 5% level. This result gives us some comfort that the findings we document are not driven by a particular choice of our cost of equity measure.

TABLE 11

Unionization and the Fama-French Cost of Equity

In Table 11, the dependent variables are the Fama-French (1993) cost of equity (FFCOE), MKTBETA, SMBBETA, and HMLBETA. Factor loadings are calculated using 5 years of monthly data. FFCOE is the annualized cost of equity projected from the Fama-French 3-factor model calculated using the average of each factor's premium from July 1926 to December 2007; MKTBETA is the loading on the market factor; SMBBETA is the loading on the SMB factor; HMLBETA is the loading on the HML factor in the same model. All loadings are estimated using 3 years of future monthly return data. UNION is the Census Industry Classification (CIC)-industry unionization rate; SALESBETA is the cyclicality of revenues in a firm's CIC industry, computed using quarterly data as the slope from a full-sample time-series regression of changes in log industry net sales over the 1-year period on log GDP growth; FINLEV is book leverage, which we define as total liabilities divided by total assets; FA/TA is net fixed assets divided by total assets; INDKL is average net fixed assets per employee in $000s within a CIC industry; FIRMAGE is the natural logarithm of the number of years a firm has been listed in CRSP; LOGSALGR is the growth in the natural logarithm of firm sales; LOGASSETS is the natural logarithm of total assets; VOLAT is the standard deviation of daily stock returns during the calendar year; and INDCONC is the Herfindahl index measuring the concentration of sales within a CIC industry. All regressions include year and major SIC-division fixed effects. The absolute values of *t*-statistics (in parentheses) are based on robust standard errors clustered at the CIC industry level. *, **, and *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

| Explanatory Variable | FFCOE (1) | MKTBETA (2) | SMBBETA (3) | HMLBETA (4) |
|---|---|---|---|---|
| UNION | 0.032** | −0.139 | 0.071 | 0.832*** |
| | (2.20) | (1.53) | (0.51) | (4.14) |
| SALESBETA | 0.001 | 0.004 | 0.001 | 0.018 |
| | (1.34) | (0.61) | (0.10) | (1.33) |
| FINLEV | 0.029*** | −0.060* | 0.181*** | 0.596*** |
| | (9.62) | (1.67) | (4.27) | (7.20) |
| FA/TA | −0.002 | −0.177*** | −0.197*** | 0.344*** |
| | (0.57) | (6.18) | (4.21) | (4.75) |
| INDKL | 0.048* | 0.353* | 0.235 | 0.302 |
| | (1.85) | (1.75) | (1.01) | (0.88) |
| FIRMAGE | −0.003*** | −0.061*** | −0.084*** | 0.077*** |
| | (4.21) | (9.25) | (10.74) | (6.61) |
| LOGSALGR | 0.003* | 0.102*** | 0.017 | −0.116*** |
| | (1.71) | (5.60) | (1.17) | (5.64) |
| LOGASSETS | 0.005*** | 0.111*** | −0.057*** | −0.048*** |
| | (4.69) | (15.54) | (6.29) | (3.54) |
| VOLAT | 0.011*** | 0.325*** | 0.164*** | −0.387*** |
| | (4.63) | (11.00) | (5.95) | (6.65) |
| INDCONC | 0.000 | −0.182*** | −0.105** | 0.350*** |
| | (0.05) | (3.59) | (2.43) | (4.10) |
| Intercept | 0.087*** | 0.287*** | 1.231*** | −0.117 |
| | (10.05) | (3.77) | (13.48) | (0.98) |
| No. of observations | 64,251 | 64,251 | 64,251 | 64,251 |
| Adjusted $R^2$ | 0.03 | 0.07 | 0.06 | 0.08 |

Looking at the specification based on the FFCOE has additional merit for identification of the unionization effect. In particular, recent theoretical work on asset pricing in the production economy provides some guidance as to how we could potentially identify the effects on equity risk associated with operating leverage and costly adjustment. This literature suggests that book-to-market equity explains the cross section of stock returns because it is correlated with both firms' operating leverage (e.g., Carlson et al. (2004), Cooper (2006), and Gourio (2007)) and costly adjustment (e.g., Zhang (2005)). Thus, to the extent that the book-to-market ratio is positively related to the loading of returns on the HML factor and that book-to-market is indeed risk, if unionization increases the cost of equity through the channel we propose, then unionization should have a strong positive effect on HML beta.

To identify the specific channel through which unions affect the FFCOE, we investigate the relation between unionization rates and each of the loadings on the Fama-French (1993) factors: market beta, SMB beta, and HML beta using a multivariate regression framework that includes all the control variables used in Table 3. The results are presented in columns (2)–(4) of Table 10.

The only statistically significant result is the strong positive effect of UNION on HMLBETA. The coefficients on other factor loadings are statistically insignificant. Clearly, unionization increases the FFCOE primarily through its effect on HML beta. Since previous work relates operating flexibility to book-to-market ratios, these findings are consistent with the view that labor unions increase firms' costs of equity because they reduce firms' operating flexibility.

## C. Controlling for Operating Leverage

Although some aspects of operating flexibility are inherently unobservable, one aspect that we can measure, to some extent, is operating leverage. Thus, it is important to determine if UNION captures aspects of operating flexibility that go beyond those captured by existing measures of operating leverage. To the extent that unionization is correlated with these proxies, our results could become weaker. If, in turn, unionization captures other dimensions of operating flexibility, then it should remain statistically significant after we control for the measures of operating leverage. Our primary measure of operating leverage is MRDOL. We also use LS, which Rosett (2003) argues is a proxy for a firm's committed labor expenses and thus may be related to equity risk.

Table 12 presents the results in which we sequentially add the measures of operating leverage to our base specification. In column (1) we control for LS, in column (2) for MRDOL, and column (3) includes both variables. When we control for the labor stock and total operating leverage, we find that it has virtually no effect on the coefficient on UNION, which remains the same in both magnitude and statistical significance. At the same time, both LS and MRDOL have positive and statistically significant effects on the cost of equity. Thus, labor unionization has strong predictive power for the cost of equity over and above the effects captured by the labor stock and total operating leverage.

We note that although controlling for operating leverage does not significantly affect the impact of unionization on the cost of equity, we cannot rule out the possibility that operating leverage is still an important aspect of flexibility in explaining the unionization premium. The problem is that our operating leverage measures might not fully capture the true nature of operating leverage. Labor unions negotiate contracts for future years, but our empirical proxies are based on past accounting data. Thus, unionization may capture additional information about a firm's future labor costs that is not contained in these operating leverage measures.

## D. Controlling for Financial Leverage

A possible concern with the interpretation of our results is that the effect of UNION on ICOE may arise spuriously due to a correlation between UNION and

TABLE 12

Unionization and the ICOE: Controlling for Operating Leverage

In Table 12, the dependent variable is the implied cost of equity (ICOE). UNION is the Census Industry Classification (CIC)-industry unionization rate. SALESBETA is the cyclicality of revenues in a firm's CIC industry, computed using quarterly data as the slope from a full-sample time-series regression of changes in log industry net sales over the 1-year period on log GDP growth; FINLEV is book leverage, which we define as total liabilities divided by total assets; FA/TA is net fixed assets divided by total assets; INDKL is average net fixed assets per employee in $000s within a CIC industry; FIRMAGE is the natural logarithm of the number of years a firm has been listed in CRSP; LOGSALGR is the growth in the natural logarithm of firm sales; LOGASSETS is the natural logarithm of total assets; VOLAT is the standard deviation of daily stock returns during the calendar year; INDCONC is the Herfindahl index measuring the concentration of sales within a CIC industry; MRDOL is the Mandelker and Rhee (1984) measure of operating leverage, which we calculate as the elasticity of a firm's operating income after depreciation to its sales; and LS is labor stock, defined as the number of a firm's employees divided by total assets. All regressions include year and major SIC-division fixed effects. The absolute values of *t*-statistics (in parentheses) are based on robust standard errors clustered at the CIC industry level. *, **, and *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

| Explanatory Variable | (1) | (2) | (3) |
|---|---|---|---|
| UNION | 0.103*** | 0.100*** | 0.103*** |
| | (3.79) | (3.60) | (3.79) |
| SALESBETA | 0.004* | 0.003* | 0.004* |
| | (1.85) | (1.73) | (1.84) |
| FINLEV | 0.025*** | 0.025*** | 0.025*** |
| | (2.90) | (2.79) | (2.90) |
| FA/TA | 0.010 | 0.016* | 0.010 |
| | (1.24) | (1.81) | (1.22) |
| INDKL | −0.004 | −0.012 | −0.004 |
| | (0.09) | (0.26) | (0.09) |
| FIRMAGE | −0.002** | −0.001 | −0.002** |
| | (2.16) | (1.46) | (2.18) |
| LOGSALGR | −0.011*** | −0.012*** | −0.011*** |
| | (2.69) | (2.78) | (2.68) |
| LOGASSETS | −0.001* | −0.002*** | −0.001* |
| | (1.75) | (3.13) | (1.76) |
| VOLAT | −0.028*** | −0.031*** | −0.028*** |
| | (4.66) | (4.98) | (4.64) |
| INDCONC | 0.045*** | 0.046*** | 0.045*** |
| | (3.96) | (3.97) | (3.96) |
| LS | 0.671*** | | 0.671*** |
| | (3.47) | | (3.48) |
| MRDOL | | 0.019** | 0.019** |
| | | (2.53) | (2.47) |
| Intercept | 0.045** | 0.056** | 0.045** |
| | (2.03) | (2.54) | (2.02) |
| No. of observations | 31,307 | 31,307 | 31,307 |
| Adjusted $R^2$ | 0.34 | 0.33 | 0.34 |

financial leverage. For example, Bronars and Deere (1991) and Matsa (2010) find that unionization is positively associated with financial leverage and argue that unionized firms rely more on debt financing to shelter their cash flows from union demands. More generally, Jagannathan and Wang (1996) argue that labor risk may be correlated with financial leverage. Since financial leverage is positively related to the cost of equity, the effect of unionization on the cost of equity could naturally follow. To address this possibility, throughout our analysis we include financial leverage as a control. Thus, we identify the effect of unions on the cost of equity net of any effect they might have through higher financial leverage. We find that financial leverage does not significantly weaken the coefficient on UNION.

To provide additional evidence that financial leverage does not drive our results, we further study the effect of unionization on the *unlevered* cost of equity

calculated from the Modigliani-Miller (1958) formula with taxes. Working with the unlevered cost of equity eliminates any concerns that financial leverage might alter our results.

To use the Modigliani-Miller (1958) formula, we need estimates of firms' costs of debt. We estimate the cost of debt for each firm-year in our sample by mapping a firm's Standard & Poor's (S&P) debt rating to the average bond yield in its rating category. Only a limited number of firms in our sample have credit ratings. We estimate missing credit ratings for other firms as follows: For the subset of companies with credit ratings, we use a set of explanatory variables and estimate an ordered logit model that predicts the S&P debt rating. Our predictors are the natural logarithm of a firm's assets, financial leverage, market beta, profitability, interest coverage, the natural logarithm of a firm's age, and the volatility of excess returns. Next, we use the estimated coefficients from this model to predict the debt rating for all the companies whose ratings are missing, but have the complete set of predictors. For each year, we match a firm's debt rating to the average bond yield in its rating category, based on individual yields on new debt issues obtained from Securities Data Company (SDC) Platinum. The estimated bond yields are then entered into the Modigliani-Miller formula to obtain the unlevered ICOE.

Next, we estimate regression models similar to those in Table 3. We use the unlevered cost of equity as the dependent variable and exclude financial leverage from our set of controls. In the untabulated results, we find a positive relation between unionization rates and the unlevered cost of equity. The coefficient on UNION is statistically and economically significant. A 1-standard-deviation increase in the unionization rate increases the unlevered ICOE by 0.95 percentage points per year. This magnitude is consistent with that obtained in the tests that use the cost of levered equity but control for financial leverage. We conclude that financial leverage does not explain the impact of unions on the cost of equity.

## E. Shareholder Activism by Labor Unions and the Cost of Equity

Recent years have seen increased activity of unions in filing shareholder proposals through their pension funds. Consequently, unions' shareholder activism may be an important channel through which unions affect firms' costs of equity. However, the direction of the effect is a priori unclear. On one hand, unions' activism may increase the cost of equity by imposing frictions in the adjustment of capital and labor inputs; on the other hand, it may decrease the cost of equity by improving a firm's governance. To explore this issue, we obtain data from RiskMetrics on all shareholder proposals sponsored by unions in firms in the S&P 1500 index over the period 1994–2006.

We conduct 2 different tests. In the first test, we examine whether a firm's ex ante probability of being subject to unions' shareholder activism affects its cost of equity. To this end, we first estimate the ex ante probability that labor unions will file at least 1 shareholder proposal that is eventually voted. For this purpose, in each year we estimate a probit model using 5 lagged predictors that Karpoff, Malatesta, and Walkling (1996) identify as determinants of the likelihood of

shareholder proposals,[8] and use the estimated coefficients to calculate the predicted probability (PROB). We then regress ICOE on PROB, all control variables in Table 3, and both firm and year fixed effects. Standard errors are clustered at the firm level. The coefficient on PROB is positive (0.012), but only marginally significant at the 10% level.

In the second test, we use a difference-in-differences (DID) approach to test whether a firm's cost of equity changes when unions file shareholder proposals. The treatment sample includes all firms in which labor unions file at least 1 shareholder proposal that is eventually voted. Similar to Karpoff et al. (1996), each firm in the treatment sample is matched with a firm in a control sample, that is, with a firm that did not receive any union-sponsored shareholder proposal during the sample period, resides in the same 2-digit SIC industry, and is closest in size (measured by assets) to the firm in the treatment group. To implement the DID method, we define an indicator variable (AFTER) equal to 1 for firms in the treatment group in the year the proposal is filed, and 0 otherwise. We then regress ICOE on AFTER, all control variables in Table 3, and both firm and year fixed effects. The standard errors are clustered at the firm level. The coefficient on AFTER is positive (0.001), but statistically insignificant ($t = 1.03$).

In sum, we find only weak evidence that firms that are more likely to attract union shareholder activism have higher costs of equity, and we find no statistically significant evidence that firms' costs of equity increase when the firms are subject to such activism. The evidence from these tests suggests that union-sponsored shareholder activism is unlikely to be the main channel through which unions affect firms' costs of equity.

## VII. Conclusion

We study the importance of operating flexibility for the cost of equity in the novel setting of the market for labor inputs. We focus on an important friction, that generated by labor unions, and hypothesize that by decreasing firms' operating flexibility, labor unions may increase firms' systematic risk and hence their costs of equity.

Consistent with our hypothesis, we find that firms in more-unionized industries exhibit a statistically and economically higher implied cost of equity. The effect holds after we control for a host of industry- and firm-level characteristics and is stronger when unions face a more favorable bargaining environment. The results cannot be attributed to an industry life-cycle effect or omitted industry characteristics, and their direction and magnitude are unlikely to be biased due to potential endogeneity issues. Moreover, the spread in the cost of equity between high- and low-unionization portfolios is highly countercyclical, which is consistent with the lack of operating flexibility imposing more risk during market downturns. Unionization is also positively related to various measures of operating leverage, and it affects both its labor and nonlabor components. Overall,

[8]They include the logarithm of assets, the logarithm of firm age, sales growth, the logarithm of the number of institutional investors, and the annual stock return in excess of the return on the value-weighted index.

our results point to the importance of operating flexibility in explaining the cross-sectional variation in the cost of equity.

More broadly, our analysis raises the question about the net effect of unions on the total surplus generated by the firm. Rents captured by unions are in part a transfer from shareholders and bondholders to workers. However, unions' actions affect corporate decisions and thus may generate additional welfare costs or benefits. Although it is possible that unions may increase productivity (empirical evidence in this regard is largely inconclusive), it has been shown that unions decrease firm value by distorting investment, R&D spending, and capital structure choices. Thus, the collective evidence from previous studies suggests that the overall effect of unions on a firm's surplus is largely negative. Quantifying the magnitude of this effect is a difficult task that we leave for future research; our results suggest that the effect may be economically significant.[9]

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[9]Back-of-the-envelope calculations suggest that a 1-standard-deviation increase in the unionization rate decreases the combined value of the firm to all of its stakeholders by about 7%.

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