Labor Unions and Expected Stock Returns*

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Abstract

We examine the effect of labor unions on the cross-section of expected stock returns. We find that expected returns are higher for firms in more unionized industries. 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, and are robust to potential endogeneity concerns. Our findings are consistent with the view that labor unions increase expected stock returns because they decrease firms' operating flexibility.

Keywords: labor unions, operating flexibility, expected stock returns

JEL: G12, J51

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1. Introduction

Little is known about how features of the market for labor inputs affect equilibrium risk and expected stock returns. Given that labor expenses constitute a large fraction of firms' total costs, the labor market might have a nontrivial impact on stock prices and returns. In this paper, we empirically study the effect of one feature of the labor market, labor unions, on the cross-section of expected stock returns.

Although in a neoclassical framework labor inputs are fully flexible and 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 increase firms' systematic risk and thus expected stock returns. This can happen in several ways. For example, labor unions make wages sticky and layoffs costly, thus increasing a firm's operating leverage due to labor and making the adjustment of a firm's labor stock more costly. In addition, unions often intervene in firms' restructurings, for example by blocking plant closures, which may also make firms' adjustment of physical capital stock more costly. Regardless of the precise mechanism, if unions reduce firms' operating flexibility, then investors should require higher returns on the capital they provide.[1]

Consistent with our prediction, we show that firms in more unionized industries are associated with higher ex-ante expected stock returns, measured either by the cost of equity arising from the Fama-French (1993) three-factor model or from 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, fixed assets, age, growth, size, and return volatility, as well as for the industry capital

[1] Throughout this study, we use the terms expected stock returns, cost of equity, and required returns interchangeably. These three quantities are identical in a standard equilibrium framework.

intensity and industry concentration. In the sample for which both measures of expected returns are available, a one-standard-deviation increase in the unionization rate increases the Fama-French cost of equity by about one percentage point per year and the implied cost of equity by about 1.5 percentage points per year.

To ensure that our tests truly identify the unionization premium, we use additional identification strategies. First, it is possible that both unionization and expected returns may be correlated with a general observable industry characteristic, such as the stage of the industry life cycle. To address this concern, we explicitly control for a host of industry-level characteristics including age, "old-economy" and "new-economy" status, R&D expenses, one- and five-year assets growth rates, and profitability. Our results indicate that such industry effect is unlikely to explain our results.

Second, we address the argument that the unionization premium in expected returns might be driven by the omission of unobservable firm or industry characteristics correlated with unionization. One way we address this issue is to study the cross-sectional variation in the unionization premium. We find that the effect of unionization on expected returns is stronger when unions face a more favorable bargaining environment: in industries with low unemployment rates and that are strongly influenced by the Democratic Party, and also in firms with more concentrated business operations. Thus, omitted variables are unlikely to drive our results, as it is not obvious why these omitted variables should have a stronger effect on expected returns when unions face a more favorable bargaining environment. In addition, our findings suggest that the unionization premium is more likely to be explained by differences in risk rather than in relative mispricing.

Last, we use instrumental-variables techniques to address the concern that endogeneity problems – omitted variables, reverse causality, and measurement error – may bias our inference. Motivated by research in labor economics, we identify two industry-level variables, the percentage of female workers and the average age of workers, that are strongly correlated with unionization rates. We show that these variables do not directly affect expected returns and thus are valid instruments. Our results show that our OLS estimates do not suffer from the omitted-variable bias or the measurement-error problem, and that unions cause an increase in expected returns.

In addition to our main analysis, we conduct a number of supplementary tests. First, we conjecture that labor unions increase expected returns by reducing firms' operating flexibility. We provide some evidence consistent with this conjecture. In particular, we decompose the Fama-French cost of equity into its three components – market beta, SMB beta, and HML beta – and find that the effect of unions on expected returns 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, this evidence is consistent with the premise that labor unions increase expected returns because they reduce firms' operating flexibility.

Second, one consequence of the lack of operating flexibility is higher operating leverage. We show that the positive relation between unionization and expected returns holds after controlling for two measures of operating leverage: the sensitivity of a firm's earnings to its sales and a firm's labor stock. Hence, unionization may capture aspects of operating flexibility that go beyond those captured by these measures of operating leverage.

Third, we establish the effect of unions on expected returns over and above their effect through financial leverage. Given that other studies have shown that unionization is positively related to financial leverage, it is possible that financial leverage may explain some of the effect of unionization on expected returns. To this end, we explicitly control for differences in financial leverage in all our regressions. 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 higher unlevered costs of equity.

Fourth, we show that the positive relation between unionization and expected returns holds, and is both statistically and economically significant, for most of the individual years in our sample. Fifth, while our analysis uses standard errors clustered by industry, the statistical significance of our findings is similar or higher if we use several alternative methods to cluster standard errors. Finally, our results hold, with improved statistical significance, if we use the Fama-MacBeth (1973) methodology with Newey-West standard errors.

The paper is structured as follows. Section 2 develops our main hypothesis and reviews related studies. Section 3 describes the data sources, defines the main variables, and examines the source of variation in the unionization data. Section 4 relates unionization to the cross-section of expected stock returns. Section 5 takes a closer look at various identification concerns. Section 6 reports additional tests. Section 7 concludes.

2. Hypothesis Development and Related Studies

We argue that labor unions might affect expected returns by reducing firms' operating flexibility. One important way in which unions reduce operating flexibility is that they make wages sticky and layoffs costly. As a result, unions increase firms' share of fixed labor costs in total labor costs, that is, 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 fixed 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 match the observed equity premium. This evidence suggests a link between labor operating leverage and systematic risk. As a result, shareholders of firms that are more unionized should require higher returns for bearing extra risk.

Unions can also generate costly adjustment in the labor stock and in physical capital. In our paper, we define "costly adjustment" as a firm's costly ability or inability to unwind its labor stock and capital stock. Adjustment of the labor stock is costly, given that unions impose frictions on the layoffs of workers. In addition, unions often tend to take an activist role by blocking restructurings. For example, the United Auto Workers have recently blocked the closure of some of General Motor's existing plants even though the closure decision was beneficial to GM shareholders. Thus, unions may also lead to costly adjustment in physical capital. Since theory (e.g., Kogan (2001), Zhang (2005), Merz and Yashiv (2006)) predicts that costly adjustment and irreversibility are positively related to expected stock returns in the cross-section, we argue that higher unionization should lead to higher expected returns.

Ex ante, it is difficult to argue which aspects of operating flexibility are important for generating the premium in expected returns. Both operating leverage and costly adjustment have a solid grounding in the theory. Moreover, 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)).[2] Regardless of the precise mechanism, our hypothesis is as follows:

<u>Hypothesis:</u>

By reducing firms' operating flexibility, labor unions increase expected stock returns.

The precise empirical 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 sensitivity coefficients (e.g., Mandelker and Rhee (1984)). Second, by design the measures are backward looking, but operating leverage in and of its nature should be forward looking (see Novy-Marx (2007)). Last, the measures generally capture the total effect rather than only the labor effect.

Although empirically little is known about how operating flexibility affects expected returns, there are empirical studies that link a specific aspect of operating flexibility – operating leverage – to systematic risk (e.g., Mandelker and Rhee (1984), and Xing and Zhang (2004)). In the context of labor markets, Rosset (2001) uses a firm's fixed labor commitments as a proxy for operating leverage due to labor, and finds that such a measure is positively related to a firm's market beta and stock return volatility. While he limits his attention to firms with the highest unionization rate and studies the

[2] Though probably the most important, operating flexibility may not be the only mechanism through which unions drive expected returns. In particular, some studies directly associate higher unionization with higher financial leverage and thus implicitly with higher expected returns. We address this issue in Section 6.3. Other candidates are less obvious. For example, anecdotal evidence suggests that unions increase strike risk, which could affect expected returns. Whether such risk should be priced is unclear, largely because such risk could be potentially diversified away.

risk implications caused by the variation in the labor stock, we study the expected returns implications for the entire cross-section of firms with different unionization rates.[3]

3. Data Sources and Variable Definitions

3.1. Data Sources and Sample Selection

We obtain unionization data for the period 1983-2004 from the Union Membership and Coverage Database at www.unionstats.com, which is maintained by Barry Hirsch and David Macpherson. The data are compiled from the Current Population Survey (CPS), based on the method used by the Bureau of Labor Statistics (BLS). 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 (IBES). We compute industry unemployment and workforce demographics using data from the Current Population Survey Labor Extracts. The presidential election data come from Dave Leip's Atlas of U.S. Presidential Elections, available at www.uselectionatlas.org.

To construct our sample, 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 (SIC codes 6000 to 6999) and utilities (SIC codes 4900 to 4999) industries. Our final sample contains information about all nonfinancial, nonutility firms in the CRSP-Compustat merged database for which we can compute expected returns, and have data on unionization rates and control variables.

[3] Our study differs from his also in other aspects. Apart from operating leverage, our unionization measure may capture other aspects of operating flexibility. Also, we study the impact of unions on expected returns that goes beyond market beta.

We measure our data at an annual frequency. Because our analysis uses lagged explanatory variables, unionization rates and our control variables span the period 1983-2004, while our dependent variables span the period 1984-2005. The exception is the implied cost of equity, which we can only compute for the period 1984-2004 due to the limited availability of analyst forecast data for the year 2005.

3.2. Labor Unions: Measures and Source of Variation in the Data

We 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 the collective bargaining with the employers.[4] Our data set comprises 199 CIC industries, which correspond roughly to 3-digit SIC industries.

Although we can measure expected returns at the firm level, a typical problem in studies on labor unions is the lack of firm-level unionization data. Consequently, we rely on labor unionization data measured at the industry level. We believe that this limitation does not pose a serious problem for our study. To the extent that industry-level unionization is a noisy proxy for firm-level unionization, this measurement-error problem could bias our tests against finding any relation between unionization and expected returns.

Because our study relates the variation in industry unionization to the variation in expected returns, we document the source and amount of the variation in *UNION*. Table 1 reports summary statistics on unionization rates for the ten most (least) unionized industries. The table illustrates both the cross-sectional and time-series variation in the unionization rates.

[4] While we follow the literature in labor economics and use union coverage as our test variable, our data also contain union membership, defined as the fraction of the industry's workers that are union members. Our results are similar if we use union membership instead.

In the cross-section, we find 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 47% during the period 1983-2004. In contrast, small retail industries have an average unionization rate of about 1%. In the time series, we see a decreasing trend in unionization rates over time, with an aggregate union coverage decreasing from 20.43% in 1983 to 11.20% in 2004.[5] We further inspect the variation in unionization rates by estimating two separate regression models of unionization rates on CIC industry and year dummies. We find that the industry variation explains about 88% of the total variation in unionization rates, but the time variation accounts for only 4%. Hence, we conclude that the primary source of variation in our data originates in the cross-sectional differences in the unionization rates.

3.3. Variables for Expected Returns Regressions

Our empirical model uses expected returns as the dependent variable. We measure this variable using either the Fama-French (1993) cost of equity (*FFCOE*) or the implied cost of equity (*ICOE*) of Gebhardt, Lee, and Swaminathan (2001). The importance of using ex-ante measures of expected returns has been forcefully emphasized by Elton (1999) in his Presidential address. Other studies that argue for ex-ante measures of expected returns include Friend, Westerfield, and Granito (1978), Kaplan and Ruback (1995), Claus and Thomas (2001), Fama and French (2002), Brav, Lehavy, and Michaely (2005), and Pástor, Sinha, and Swaminathan (2007), among others. Appendix A provides further details on the construction of these variables.

[5] Although union membership as a percentage of the total workforce has declined slowly but steadily from its peak in the mid-1950s, unions still exert an important impact on firms' decisions. For example, in recent years, unions have opposed General Motors' labor restructuring plans, the merger between U.S. Airways Group and America West, the Sprint Nextel Corp.'s plan to spin off its local phone business into an independent company, and the level of pay of Verizon's CEO Ivan Seidenberg.

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 by using quarterly data as the slope from the full-sample time-series regression of changes in log industry net sales over the one-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 net 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 three 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.

Panel A of Table 2 reports summary statistics for the sample used in the Fama-French cost of equity regressions. Panel B reports statistics for the sample used in the implied cost of equity regressions. The mean and median *FFCOE*, in our sample of 8682 firms, equal 14.3% and 14.0%, respectively, with a standard deviation of 8.3%. Likewise, the mean and median *ICOE*, in our sample of 4835 firms, are 10.4% and 11.4%, respectively, with a standard deviation of 5.6%.

4. Unionization and Expected Returns

We examine the univariate association between unionization and expected returns measured as the Fama-French cost of equity (*FFCOE*) or the implied cost of equity (*ICOE*). In each year, we sort companies into quintile portfolios based on their previous year's unionization rates. Next, we calculate the equal- and value-weighted expected returns for each of the five portfolios. Last, we average the portfolio returns across years. Quintile 1 includes firms with the lowest, and Quintile 5 includes firms with the highest unionization rates. Panel A of Table 3 presents the results.

The results indicate an increasing pattern in expected returns as we move from the lowest- to the highest-unionization portfolio for both the equal- and value-weighted portfolios. For both the *FFCOE* and *ICOE* samples, the average unionization rate varies from about 2.5% for Quintile 1 to about 33% for Quintile 5. The difference in the Fama-French cost of equity between Quintile 5 and Quintile 1 equals 0.94 percentage points per year for the equal-weighted portfolio and 2.31 percentage points per year for the value-weighted portfolio. Both numbers are statistically significant. Likewise, for the implied cost of equity the respective differences are approximately 2.73 and 2.1 percentage points per year.

Although our objective is to test whether unionization is related to expected returns, 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 pooled cross-sectional OLS model:

$$ER_{ijt+1} = a_0 + a_1\, UNION_{jt} + a_2\, Controls_{ijt} + \varepsilon_{ijt+1}, \qquad (1)$$

where *i* indexes firms, *j* indexes a firm's CIC industry, and *t* indexes year. Our dependent variable is the expected return (*ER*), for which either *FFCOE* or *ICOE* serves as a proxy.

Our test variable (*UNION*) is the percentage of workers in the industry covered by unions. Current theories predict that expected stock returns should be affected by firms' revenue cyclicality (*SALESBETA*) and financial leverage (*FINLEV*). We also include two variables that are related to the production technology: the nature of firms' assets (*FA/TA*) and the industry-level capital intensity (*INDKL*). We add in other firm characteristics that earlier research has found to be correlated with expected returns, including *FIRMAGE*, *LOGSALGR, LOGASSETS, VOLAT,* and *INDCONC*. To control for time and industry variation in expected returns, in each regression we include the one-digit SIC industry dummies and year dummies. The coefficient of interest is a_1, which measures whether companies in more unionized industries have different expected returns than do other companies.

To assess the statistical significance of our estimates, throughout our analysis we use a conservative method of clustering standard errors at the CIC industry level. The problem is 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 (e.g., Moulton (1986)). 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, the clustering method accounts for heteroskedasticity in the residuals. Panel B of Table 3 reports the results from estimating equation (1).

We continue to find a positive and statistically significant relation between unionization rates and expected returns. In column (1), we regress *FFCOE* on *UNION* and industry and year dummies. The coefficient on *UNION* equals 0.046 and is statistically significant. In column (2), we add *SALESBETA, FINLEV, FA/TA,* and *INDKL,* and column (3) includes *FIRMAGE, LOGSALGR, LOGASSETS, VOLAT,* and *INDCONC.* We find that although including all these controls reduces the magnitude of the coefficient on *UNION* to 0.037, the coefficient remains statistically significant.

In terms of economic significance, a one-standard-deviation increase in the unionization rate increases expected returns by approximately 0.47 percentage points per year. The coefficients on *SALESBETA*, *FINLEV*, *LOGASSETS*, and *VOLAT* are positive, and the coefficient on *FIRMAGE* is negative. All coefficients are statistically significant. In contrast, *FA/TA, INDKL, LOGSALGR,* and *INDCONC* are not correlated with the Fama-French cost of equity.

In columns (4)-(6), we repeat the above regressions, but now we use the implied cost of equity (*ICOE)* as the dependent variable. In column (4), we regress *ICOE* on *UNION*, while including industry dummies and year dummies. The coefficient on *UNION* is 0.116 and is statistically significant. Once we add in all the controls, the coefficient on *UNION* remains similar in terms of its magnitude (0.112) and is statistically significant at the 1% level. A one-standard-deviation increase in the unionization rate is associated with an economically significant increase in the implied cost of equity of approximately 1.46 percentage points per year.[6]

We note that the effect of unionization on the expected returns is greater if we measure the expected returns with *ICOE* rather than with *FFCOE*. One reason for such a discrepancy is the qualitative difference in the samples we use for *ICOE* and *FFCOE*.

[6] As an alternative approach to studying the economic impact of unions on the expected returns, we investigate the direct effect of unions on a firm's market value of equity, controlling for cash flows. The main advantage of this approach is that it does not rely on any specific model for estimating required returns on equity. The untabulated results indicate that firms in more unionized industries have lower equity values, even after controlling for differences in the levels and volatility of their cash flows.

The sample for *ICOE* contains only firms with available earnings forecasts, and these are mostly large firms. In contrast, the sample for *FFCOE* covers all firms in the CRSP-Compustat universe, which includes a significant fraction of smaller firms that normally are not covered by analysts.

To check how much any discrepancy in the sample composition can explain the differences in the coefficients, we repeat the *FFCOE* regressions on the subset of firms for which we can compute *ICOE*. For the specification that includes all the control variables, we find that the coefficient on *UNION* increases to 0.077 and is statistically significant. A one-standard-deviation increase in the unionization rate increases the Fama-French cost of equity by one percentage point per year, which is closer to, but still less than, the magnitude of the effect that we find for *ICOE*.[7] Another reason for the discrepancy might be that by construction, *FFCOE* does not capture time variation in factor premiums, which may cause *FFCOE* to differ from the true expected returns.

5. Identification Tests

5.1. Omitted Industry Characteristics

It is possible that both unions and expected returns are 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 one-digit industry fixed effects. However, our concern is that we might still be missing some other

[7] Another manifestation of the differences between the two samples is evident from the coefficients on *LOGASSETS* and *VOLAT*. In particular, the coefficients change their signs when we move from the *FFCOE* to the *ICOE* specification. This difference is largely driven by the differences in the composition of firms in each sample. Fixing the sample composition at the *ICOE* sample makes the coefficients consistent across the two dependent variables.

important observables. For example, 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 have higher expected returns.

We examine this concern in three ways. First, in our baseline specification we have already controlled for firm age, firm sales growth, and one-digit SIC industry fixed effects, which should at least partially capture such an effect. 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. Nevertheless, since this evidence is somewhat circumstantial and some mature industries are highly unionized, we further turn to more formal regression evidence.

To this end, in Table 4 we control for a long list of variables that are our proxies 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*), one- and five-year assets growth rates (*1YR-INDGRWTH*, and 5*YR-INDGRWTH*), and industry profitability (*INDPROFIT*).[8] *INDLOGAGE* is the natural logarithm of age of the oldest company in the firm's CIC industry. *OLDECON* is an indicator variable equal to one if the firm operates in an old-economy industry, and zero 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 one if the firm operates in a new-economy industry, and

[8] One could argue that what we really capture in our analysis is a pure "value effect", that is, firms in more unionized industries may simply be value stocks or value industries. Our untabulated results show that the value effect, measured by the book-to-market ratio, does not explain our results. However, given that our expected return measures are by construction related to the book-to-market ratio we do not include this control in our main tests.

zero 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 assets in a firm's CIC industry. *1YR-INDGRWTH* is the median one-year growth in assets in a firm's CIC industry. *5YR-INDGRWTH* is the median five-year growth in assets in a firm's CIC industry. *INDPROFITS* is the median ROE in a firm's CIC industry.

The results indicate that we cannot attribute the unionization premium to the stage of the industry's life cycle. Including all the industry controls in the *FFCOE* specification changes the coefficient on *UNION* merely from 0.038 to 0.035, and the *t*-statistic drops from 2.74 to 2.56. The coefficient on *UNION* in the *ICOE* regression drops slightly more from 0.131 to 0.115, but it is still highly significant. Among the industry characteristics we consider, industry age seems to be most closely related to expected returns. Also, *NEWECON* is significant in the *FFCOE* regressions, while *1YR-INDGRWTH* and *INDPROFIT* are important in the *ICOE* regressions. All other characteristics do not seem to add additional explanatory power to our regressions.[9]

5.2. Cross-Sectional Variation in the Unionization Premium

Another way to identify the effect of unions on expected returns 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 affect firms' operating risk, and thus, the expected stock returns. To test this prediction, we estimate the following general regression model:

[9] As an alternative approach, we use a two-stage procedure, in which we first construct a residual unionization variable that is orthogonal to all other industry-level controls included in Table 4, and then we study the effect of this variable on expected returns. The results remain unchanged.

$$ER_{ijt+1} = a_0 + a_1\ UNION_{jt} + a_2\ ENV_{ijt} + a_3\ UNION_{jt} \times ENV_{ijt} + b\ Controls_{ijt} + \varepsilon_{ijt+1}. \qquad (2)$$

Our controls include *SALESBETA*, *FINLEV*, *FA/TA*, *INDKL*, *FIRMAGE*, *LOGSALGR*, *LOGASSETS*, *VOLAT*, *INDCONC*, *INDLOGAGE*, *OLDECON*, *NEWECON*, *INDRDEXP*, *1YR-INDGRWTH*, 5*YR-INDGRWTH*, and *INDPROFIT*. In each regression we also include the one-digit SIC industry and year dummies, and cluster standard errors at the CIC industry level.

We also alternately include three other 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 proxies include the CIC industry-level unemployment rate (*UNEMPL*) and political environment (*DEMOCRAT*), as well as the concentration of a firm's sales across its business segments (*BUSCONC*). *UNEMPL* is the unemployment rate within a firm's CIC industry. *DEMOCRAT* is the fraction of workers within a CIC industry that is located in a democratic state. 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. *BUSCONC* is the Herfindahl Index measuring the concentration of a firm's sales across its business segments.

Our intuition for using these variables is as follows. First, existing empirical studies document that higher industry unemployment undermines unions' bargaining power (e.g., Svejnar (1986)). For example, companies in such industries are more able to substitute their unsubordinated workers with those who are unemployed.[10] Also, the general idea that unemployment disciplines the workforce has been studied theoretically in Shapiro and Stiglitz (1984) and others. Hence, we predict that higher industry

[10] This process has been recently observed in the airline industry, in which the strike of the mechanic workers in Northwest Airlines led management to substitute its workforce with the available free outside workers. As a result, the boycott orchestrated by unionized workers was not successful.

unemployment should weaken unions' bargaining power, which then implies a negative coefficient on *UNION×UNEMPL*. 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))[11]. 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×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×BUSCONC*.

Table 5 reports the results for *FFCOE* and *ICOE.* We demean *UNION* and *ENV* 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 expected returns, conditional on the bargaining environment. We find results consistent with our predictions. The coefficient on *UNION×UNEMPL* is negative, and the coefficients on *UNION×DEMOCRAT* and *UNION×BUSCONC* are both positive. All three coefficients are statistically significant for *FFCOE* and *ICOE*. Thus, the effect of labor unions on expected returns is weaker when the unemployment rate in a firm's industry is higher. Likewise, the effect is stronger when a larger fraction of the industry

[11] For example, in February 2007, the Democrats-controlled House passed the Employee Free Choice Act despite the Republican opposition. This bill aims to make it easier for workers to join unions. Under the current law, unions are required to hold the National Labor Relations Board-supervised elections. If the new law is implemented, employers will be required to recognize a union as long as the majority of workers signs union cards.

workers is located in the Democratic-Party states or a firm's operations are more concentrated across business segments.

Our results suggest that the effect of unionization on expected returns is unlikely to be driven by the omission of unobservable firm or industry characteristics that are correlated with unionization rates and affect expected returns. In fact, if one believes that the unionization premium in expected returns is driven by omitted unobservable variables correlated with unionization, then one would have to explain why these omitted variables have a stronger effect on expected returns when unions face a more favorable bargaining environment. The results also show that differences in risk rather than in relative mispricing are more likely to explain the unionization premium in expected returns.

5.3. Endogeneity

Our analysis suggests that labor unions increase firms' operating risk and thus expected stock returns. However, it is possible that our inference may be biased due to endogeneity – correlation between *UNION* and the error term in equation (1) coming from three different sources: reverse causality, omitted unobserved characteristics, and measurement error. First, greater unionization might arise in firms operating in riskier industries with higher expected returns. Second, our results might spuriously arise if important unobservable firm or industry characteristics are correlated with unionization and affect expected returns. Third, our industry-level unionization variable is an imperfect measure of the unobservable firm-level unionization. This classical errors-in-variables problem could lead to attenuation bias, and we would underestimate the true effect of unionization on expected returns.

We use a two-stage least-squares (2SLS) test to show that our results are robust to these endogeneity concerns. This test is an alternative way of dealing with the omitted-variables bias, and also addresses our concerns regarding reverse causality and measurement error. In our context, good instruments are exogenous variables that are economically related to unionization but are uncorrelated with the error term of the second-stage regression relating expected returns to unionization. To find such instruments, we consider factors that the labor economics literature has shown to affect the demand for union services (e.g., Hirsch (1980, 1982)). We use two demographic characteristics of the industry labor force that we construct from the Census Population Survey: the fraction of female workers (*FEMALE*) and the average age of workers (*WORKERAGE*). We argue that both *FEMALE* and *WORKERAGE* are economically related to *UNION*, but are uncorrelated with the error term of the second-stage regression that relates expected returns to unionization.

More 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 non-wage 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. Because 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 relatively low. Thus, we expect a positive relation between *WORKERAGE* and *UNION*. At the same time, we have no reason to believe that our instruments would have a direct economic impact on expected stock returns, and they are thus unlikely to be correlated with the error term in the second-stage regression.

Table 6 reports the two-stage least-squares (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 first-stage results relating *UNION* to *FEMALE*, *WORKERAGE*, and to the exogenous variables of the model. Panel B reports coefficients from the second-stage regression of expected returns on the value of *UNION* predicted in the first-stage regression and the corresponding exogenous control variables.

Consistent with the hypothesized economic relation between our instruments and unionization, the first-stage results show that *UNION* is negatively affected by *FEMALE* and is positively affected by *WORKERAGE*. Both coefficients are statistically significant at the 1% level. Moreover, our instruments have strong predictive power. The partial R^2 of the first-stage regression indicates that the instruments explain around 14% 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 zero. 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 second-stage regression. We also find (unreported) that our instruments are not statistically significant when included in the OLS test of expected returns on unionization, which indicates that our instruments have no additional effect on expected returns other than through unionization.

The second-stage results continue to provide strong evidence of a statistically significant and positive relation between *UNION* and both the Fama-French cost of equity (*FFCOE*) and the implied cost of equity (*ICOE*). The magnitudes of the estimated 2SLS coefficients are similar to those obtained from the OLS tests reported in Table 3. In addition, using the Hausman (1978) test, we cannot reject the null hypothesis that the 2SLS and OLS coefficients on *UNION* are the same. Hence, we conclude that our OLS results are robust to the endogeneity concerns; that is, greater unionization causes higher expected stock returns and our OLS estimates are not biased due to omitted variables or measurement error.

6. Additional Tests

6.1. Unionization and Loadings on the Fama-French Factors

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, Fisher, and Giammarino (2004), Cooper (2006), and Gourio (2006)) and costly adjustment (e.g., Zhang (2005)). Thus, to the extent that the book-to-market ratio is positively related to the loading on the HML factor and that book-to-market is indeed risk, if unionization increases expected returns through the channel we propose, then unionization should have a stronger positive effect on HML beta than on the other factor loadings in the Fama-French three-factor model.

To identify the specific channel through which unions affect the Fama-French cost of equity, we investigate the relation between unionization rates and each of the loadings on the Fama-French factors: market beta, SMB beta, and HML beta. In each year, we sort firms into quintile portfolios based on their previous year's industry unionization rates and compute equal-weighted averages of each loading. We then average them across years. The results are presented in Panel A of Table 7.

We find that the market beta for Quintile 5 is lower than that for Quintile 1, but we do not find any statistically significant difference in market betas across Quintiles 2-4. In addition, we observe a large negative effect of unionization on SMB beta, which we attribute to the fact that more-unionized firms tend to be larger. Column (5) of the table shows that higher unionization rates are associated with substantially higher HML betas. The difference in the betas between Quintile 5 and Quintile 1 equals 0.42, and it is statistically significant.

Panel B examines the effect of unionization on each loading separately, using a multivariate regression framework that includes all the control variables used in Table 3. In contrast to the univariate tests, we do not find any systematic relation between unionization rates and market betas. In addition, the negative effect of *UNION* on *SMBBETA* vanishes once we include the controls. The only statistically significant result is the strong positive effect of *UNION* on *HMLBETA*. Clearly, unionization increases the Fama-French cost of equity primarily through its effect on *HML* beta. This finding is consistent with the view that labor unions increase expected stock returns because they reduce firms' operating flexibility.

6.2. Operating Leverage

Our results on the HML beta suggest that labor unions increase expected stock returns through the operating flexibility channel. Although some aspects of operating flexibility are inherently unobservable, one aspect that we can measure is operating leverage. Following Mandelker and Rhee (1984) and Xing and Zhang (2004), we 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.[12] We refer to this measure as the Mandelker and Rhee degree of operating leverage (*MRDOL*), which we further winsorize at the 5% level. We also use the labor stock (*LS*), which we define as the number of a firm's employees divided by total assets, as another measure of operating leverage. As Rosett (2003) argues, this variable is a proxy for a firm's committed labor expenses and thus may be related to equity risk.

In Table 8, we sequentially add the measures of operating leverage to our base specification, regressing expected stock returns on unionization and control variables. Our objective is to examine how the addition of these variables affects the coefficient on *UNION*. Columns (1) and (4) repeat the results for the benchmark specifications in Table 3 for *FFCOE* and *ICOE*, respectively. In columns (2) and (5) we control for *LS*. In columns (3) and (6) we further control for *MRDOL*. 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. Thus, labor unionization has strong predictive power for expected stock returns over and above the effects captured by the labor stock and total operating leverage.

We note that although controlling for operating leverage does not seem to significantly affect our unionization effect, 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 may not fully capture the

[12] For firms with positive *EBIT*, we calculate this elasticity by estimating the time-series quarterly regression model 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 model 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.

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.

6.3. Financial Leverage

Our results are consistent with the view that labor unions increase expected returns by decreasing operating flexibility. A possible concern with this interpretation is that the unionization effect may arise due to differences in financial leverage and not in operating flexibility. For example, Bronars and Deere (1991) and Matsa (2005) find that unionization is positively associated with financial leverage and argue that firms rely more on debt financing to shelter their cash flows from union demands. Since financial leverage is positively related to expected returns, the effect of unionization on expected returns 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 expected returns 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 support for this finding, we consider the unlevered cost of equity calculated from the Modigliani-Miller formula with taxes. Working with the unlevered cost of equity eliminates any concerns that financial leverage might alter our results. The formula we use to unlever cost of equity is given by:

$$UNLCOE = \frac{ER + (D/S)(1-t)r_D}{1 + (D/S)(1-t)}, \tag{3}$$

where *UNLCOE* denotes the unlevered cost of equity; *ER* is the levered cost of equity, which we define as either *FFCOE* or *ICOE*; r_D denotes the cost of debt; *D* and *S* are the levels of debt and market equity, respectively; and *t* is the maximum corporate tax rate in a given year.

We estimate the cost of debt for each firm-year in our sample by mapping a firm'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 to 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, revenue cyclicality, profitability, interest coverage, the natural logarithm of a firm's age, excess returns, 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 obtained from SDC Platinum. The estimated bond yields are then entered into equation (3) to obtain the unlevered Fama-French cost of equity (*UFFCOE*) and the unlevered implied cost of equity (*UICOE*).

Next, we estimate regression models similar to those in Table 3. We use the unlevered cost of equity (*UNLCOE*) as the dependent variable and exclude financial leverage from our set of controls. Table 9 reports the results. We find a positive relation between unionization rates and the unlevered cost of equity. The coefficients on *UNION* are statistically and economically significant for both *UFFCOE* and *UICOE*. A one-standard-deviation increase in the unionization rate increases the unlevered Fama-French cost of equity by 0.33 percentage points and the unlevered implied cost of equity by 1.15 percentage points per year. We conclude that financial leverage does not explain the impact of unions on expected stock returns.

6.4. Year-by-Year Analysis

Although our primary source of variation in the unionization rates is cross-sectional, we also observe a downward trend in the unionization rates over time. Thus, it is possible that the interpretation of our results might not precisely reflect the true cross-sectional impact of unions on expected returns, or that our results might be driven by a few particular periods in the data. To address these concerns, we redo our analysis by considering the cross-sectional regressions in specification (1) for each individual year. Table 10 presents the estimated coefficients on *UNION*, along with their respective *t*-statistics. Columns (1)-(2) and (4)-(5) report the results for *FFCOE* and *ICOE*, respectively.

Our findings indicate that the relation for the entire sample also holds true for most of the individual years. The coefficient on *UNION* is positive and statistically significant for 14 out of 22 years for *FFCOE,* and for 20 out of 21 years for *ICOE*. Among those coefficients that are not significant, the effect is positive in all but one case for *FFCOE* and in the only case for *ICOE*.

To provide further evidence on the time-series impact of labor unions, we evaluate the contribution of unions to total expected returns. We estimate this contribution by calculating the product of the coefficient on *UNION* and the average unionization rate in a given year. Columns (3) and (6) of Table 10 present the results. We observe that the contribution of unions to expected stock returns is relatively stable over time for both *FFCOE* and *ICOE*. These results suggest that despite the downward time trend in the unionization rates, the effect of unions on discount rates does not diminish over time.

6.5. Alternative Econometric Methods

In our analysis, we estimate pooled OLS test models and assess the statistical significance of our estimates by using the conservative method of clustering standard errors at the CIC industry groupings. For robustness, we estimate our main specification (1) by using various alternative methods to calculate standard errors. Table 11 reports the coefficients on *UNION* and their statistical significance.

Row 1 repeats the results for the base-case specification with standard errors clustered at the CIC industry level. In rows 2 and 3, we cluster standard errors at the firm and year levels, respectively. Rows 4 and 5 utilize two-dimensional clustering. In row 4, we follow Thompson (2006), clustering the standard errors by firm and year. This approach allows for correlations among different years in the same firm and different firms in the same year. In row 5, we cluster standard errors by industry and year, using the method of Cameron, Gelbach, and Miller (2006). This approach, in turn, allows for correlation not only among firms in the same industry, but also among firms in different industries in the same year. Row 6 reports the results from the cross-sectional regressions using the Fama-MacBeth (1973) method with the Newey-West (1987) standard errors. To alleviate any concern regarding the possible autocorrelation in the estimated coefficients, we set the lag length at six, which is longer than the length of the estimation window we use for estimating the Fama-French three-factor loadings.

Compared to the base-case specification, in both the *FFCOE* and in the *ICOE* regressions the *t*-statistics associated with the coefficients on *UNION* more than double when we cluster standard errors by firm, by year, or by both firm and year. But when errors are simultaneously clustered by industry and year, the *t*-statistics are similar to those reported in the base-case specification. This finding suggests that there is little correlation among different industries in the same year. The estimated effect of *UNION* arising from the Fama-MacBeth estimation is higher in magnitude than is that arising from the pooled OLS tests, particularly when we measure expected returns by *FFCOE*. For both *FFCOE* and *ICOE*, the *t*-statistics based on the Newey-West standard errors are considerably larger.

6.6. Additional Tests

In our analysis we control for differences across industries by using the one-digit SIC industry dummies. Doing so ensures that it is not the variation in the unionization rates across one-digit SIC industries that identifies our results. We further explore the impact of the industry effects by replacing the one-digit SIC dummies with the two-digit SIC industry dummies. Although this finer industry partition might provide better control for the unobserved time-invariant industry characteristics that might be correlated with unionization, it would ignore the large variation in the unionization rates across two-digit industries and identify the estimated coefficients by using only the much smaller variation within the two-digit industries.[13] In our sample, the results remain qualitatively similar for both cases.

In our study, we assign to each firm the unionization rate corresponding to its primary CIC industry, which we identify based on the firm's primary SIC code. Although

[13] We measure unionization at the CIC industry level, which roughly corresponds to three-digit SIC codes, and, as we argue in Section 3, its main source of variation is cross-sectional rather than time series. Therefore, while including CIC or three-digit SIC industry dummies in our estimation would appear potentially appealing, it would completely eliminate the variation in the unionization rates across industries and would force us to identify the estimated coefficients only from the small time-series variation. Likewise, we would face a similar problem if we wanted to identify the union effect from the model using the first-difference transformation.

this procedure is error free for single-segment firms, it may not be accurate for capturing the unionization rate of multi-segment firms whose operations span several industries with potentially different unionization rates. To address this concern, we repeat our analysis using the employee-weighted unionization rate averaged across the industries corresponding to each firm's business segments. Both the coefficient and its standard error remain almost identical to those reported before.

We estimate expected stock returns by using the four-factor model, which additionally includes the momentum factor. The momentum beta is unrelated to unionization and the results for expected returns remain qualitatively similar. We also include Nasdaq and S&P 500 dummies to control for any systematic differences across firms that could arise from correlation of unionization with firms' exchange listing or index membership. Our results remain unaffected.

Our empirical analysis relies on ex-ante measures of expected returns. Alternatively, we could gauge expected stock returns by using realized returns, under the assumption that, in the long run, average realized returns converge to expected returns. We show that unionization rates are not significantly related to realized returns. We argue that realized returns are a poor proxy for expected returns because the unexpected component of returns is negatively related to unionization. The details of the analysis are omitted for brevity.

7. Conclusion

We examine what determines expected returns from a new perspective, namely, that imperfections in the market for labor inputs may have a significant impact on expected stock returns. 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 equilibrium expected returns.

Consistent with our hypothesis, we find that firms in more-unionized industries exhibit statistically and economically higher expected returns. The results cannot be attributed to an industry life-cycle effect. Furthermore, the effect of unions on expected returns is stronger when unions face a more favorable bargaining environment, and thus are more able to affect firms' operating strategies. Moreover, we show that our results do not suffer from endogeneity concerns. We also find that unions increase expected returns primarily through the book-to-market channel, which previous research has related to operating flexibility.

Our findings support the hypothesis that labor unions increase expected stock returns by decreasing operating flexibility. More generally, our results point to the importance of the market for labor inputs in explaining the cross-sectional variation in expected stock returns.

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Appendix A. Expected Returns Estimation

FFCOE derives from the Fama and French (1993) three-factor model. First, we estimate annual three-factor model loadings. Specifically, for each stock *j* in year *t* (between 1984 and 2005), we estimate the following time-series regression model using monthly data from January of year *t-4* to December of year *t*:

$$r^j - r^f = \alpha_j + \beta^j_{m,t}(r^M - r^f) + \beta^j_{h,t} HML + \beta^j_{s,t} SMB + \varepsilon^j , \qquad \text{(A.1)}$$

where the dependent variable (*r*) is the monthly return on stock *j* minus the risk-free rate, and the independent variables are given by the returns of the following three zero-investment factor portfolios (month subscripts are omitted for brevity). The term $r^M - r^f$ denotes 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 difference between small and large capitalization stocks. We then construct the Fama-French cost of equity capital (*FFCOE*) in year *t* as follows:

$$FFCOE_t = \overline{r^f} + \hat{\beta}^j_{m,t}(\overline{r^M - r^f}) + \hat{\beta}^j_{h,t}\overline{HML} + \hat{\beta}^j_{s,t}\overline{SMB} , \qquad \text{(A.2)}$$

where $\overline{r^f}$, $(\overline{r^M - r^f})$, $\overline{HML}$, and $\overline{SMB}$ are the average annualized returns of the risk-free asset and three Fama-French factors calculated over the period 1925-2005.[14] For simplicity, throughout the paper we refer to the beta loadings on each of these factors as *MKTBETA*, *HMLBETA*, and *SMBBETA*, respectively.

Our second measure is the implied cost of equity capital (*ICOE)*, which 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 a recent study, Pástor, Sinha, and Swaminathan (2007) show that if both dividend growths and conditional expected returns follow AR(1) processes, then *ICOE* is a perfect proxy for expected returns. In our study, we follow Gebhardt, Lee, and Swaminathan (2001), hereafter GLS, and calculate the implied cost of equity using the residual income model. While several methods are currently available, we choose the GLS method, because it has received the most acclaim in academic research. For example, Guay, Kothari, and Shu (2005) argue that the implied cost of equity is the best measure among those they survey. Similarly, the GLS measure is used in other studies of cost of equity, e.g., Lee, Ng, and Swaminathan (2003), Pástor, Sinha, and Swaminathan (2007).

To obtain *ICOE*, we start with the valuation formula implied by the dividend discount model.

[14] Our results are similar if we calculate factor premiums using data that are observed by an investor.

$$P_t = \sum_{i=1}^{\infty} \frac{E_t(D_{t+i})}{(1+r_e)^i} \tag{A.3}$$

In the formula, P is the stock price, D is dividends, r_e is the discount rate, and $E(.)$ is the expectation operator. Assuming clean surplus accounting (Change in Book Equity = Net Income – Dividends), we obtain the residual income equity valuation model:

$$P_t = B_t + \sum_{i=1}^{\infty} \frac{E_t(ROE_{t+i} - r_e)B_{t+i-1}}{(1+r_e)^i}, \tag{A.4}$$

where ROE is the return on equity and B is the book value of equity.

We subsequently solve for the implied cost of equity (r_e) from the equation above using the current stock price, current book value of equity, and forecasts of future ROE and book value of equity. Following GLS, we obtain the forecasts in the following way. First, we estimate earnings forecast in two steps: 1) We forecast earnings explicitly for the following three years using IBES earnings per share (*EPS*) and *EPS* growth forecast, and 2) We forecast earnings beyond year 3 implicitly, by assuming that the ROE at period $t+3$ mean reverts by period T to the industry median ROE (described below). In our specification, we set T at 12 years. We obtain the forecasts using a simple linear interpolation between ROE at period $t+3$ and the industry median ROE at time t. The industry median ROE is a moving median of the previous ten-year ROEs from all firms in the same industry from the 48-industry Fama and French (1997) classification. Second, by assuming a clean-surplus accounting system and assuming a constant dividend payout ratio, we forecast future book value of equity using the forecasted future earnings.

With future earnings and book value of equity, we can calculate future ROE ($FROE$) and so the stock price at year t as:

$$\begin{aligned} P_t &= B_t + \frac{FROE_{t+1} - r_e}{1+r_e} B_t + \frac{FROE_{t+2} - r_e}{(1+r_e)^2} B_{t+1} + TV \\ &= B_t + \frac{FROE_{t+1} - r_e}{1+r_e} B_t + \frac{FROE_{t+2} - r_e}{(1+r_e)^2} B_{t+1} + \sum_{i=3}^{T-1} \frac{FROE_{t+i} - r_e}{(1+r_e)^i} B_{t+i-1} + \frac{FROE_{t+T} - r_e}{r_e(1+r_e)^{T-1}} B_{t+T-1} \end{aligned} \tag{A.5}$$

where $FROE_{t+1}$ and $FROE_{t+2}$ are ROE forecasts for the following two years, and TV is the terminal value estimate. Thus, the implied cost of equity (*ICOE*) is the numerical solution (r_e) of the above equation. Despite its intuitive appeal, one limitation of this measure is its reliance on the availability of analyst forecast data. This problem biases our sample towards the larger firms, for which IBES reports analyst data. Also, the limited availability of data in IBES constrains our *ICOE* sample to the period 1984-2004.

Table 1
Unionization Rates for Selected Industries: 1983-2004

The table reports the percentage of the industry workers covered by unions in their collective bargaining with the firm. Based on the average industry unionization rates over the period 1983-2004, we identify the ten most and ten least unionized industries from the group of 199 Census Industry Classification (CIC) industries included in our data.

Panel A: Highest-Unionization Industries

Industry	1983-2004	1983	1993	2004
Railroads	76.2%	85.4%	78.0%	68.0%
Pulp, paper, and paperboard mills	50.3%	62.7%	59.4%	38.7%
Blast furnaces, steelworks, rolling and finishing mills	50.2%	67.3%	56.9%	31.0%
Motor vehicles and motor vehicle equipment	46.6%	61.7%	47.9%	29.9%
Bus service and urban transit	42.1%	52.3%	41.5%	37.1%
Air transportation	41.6%	45.7%	42.1%	44.9%
Railroad locomotives and equipment	41.5%	62.7%	29.4%	15.9%
Primary aluminum industries	40.8%	52.6%	42.6%	38.7%
Telephone communications	40.5%	63.1%	44.65	24.0%
Leather tanning and finishing	40.5%	48.3%	58.5%	12.5%

Panel B: Lowest-Unionization Industries

Industry	1983-2004	1983	1993	2004
Radio, TV, and computer stores	1.8%	1.4%	1.4%	1.5%
Restaurants and other food services	1.7%	2.5%	4.0%	N/A
Lodging places, except hotels and motels	1.7%	0.5%	1.1%	2.7%
Accounting, auditing, and bookkeeping services	1.6%	1.4%	1.1%	1.2%
Gasoline service stations	1.6%	2.0%	1.8%	0.8%
Agricultural production, livestock	1.5%	1.2%	1.8%	2.4%
Beauty shops	1.4%	2.4%	1.8%	0.8%
Sewing, needlework, and piece goods stores	1.3%	0.0%	7.1%	0.0%
Retail nurseries and garden stores	0.9%	0.0%	0.0%	1.0%
Mobile home dealers	0.1%	0.0%	0.0%	N/A

Table 2
Summary Statistics for Main Variables

Panel A: Fama-French cost of equity regressions (1984-2005). *FFCOE* is the cost of equity calculated using the Fama-French (1993) three-factor model; *UNION* is union coverage at the Census Industry Classification (CIC) industry level; *SALESBETA* is a firm's revenue cyclicality; *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. **Panel B: Implied cost of equity regressions (1984-2004).** *ICOE* is the implied cost of equity. Other variables are as previously defined.

Panel A: Variables for Fama-French Cost of Equity Regressions (8682 firms)

Variable	N	Mean	Std. Dev.	Median	5^{th} Pctile	95^{th} Pctile
FFCOE	63,505	0.143	0.083	0.140	0.013	0.284
UNION	63,505	0.139	0.127	0.102	0.018	0.405
SALESBETA	63,505	1.859	1.244	1.514	0.507	4.326
FINLEV	63,505	0.240	0.212	0.211	0.000	0.627
FA/TA	63,505	0.304	0.221	0.251	0.038	0.763
INDKL	63,505	0.064	0.074	0.039	0.014	0.219
FIRMAGE	63,505	2.568	0.744	2.565	1.386	3.970
LOGSALGR	63,505	0.084	0.420	0.077	-0.365	0.555
LOGASSETS	63,505	4.963	2.067	4.853	1.768	8.542
VOLAT	63,505	0.602	0.343	0.520	0.222	1.283
INDCONC	63,505	0.225	0.180	0.176	0.044	0.601

Panel B: Variables for Implied Cost of Equity Regressions (4835 firms)

Variable	N	Mean	Std. Dev.	Median	5^{th} Pctile	95^{th} Pctile
ICOE	25,835	0.104	0.056	0.114	0.001	0.180
UNION	25,835	0.137	0.130	0.099	0.017	0.413
SALESBETA	25,835	1.867	1.289	1.514	0.507	4.430
FINLEV	25,835	0.214	0.187	0.192	0.000	0.551
FA/TA	25,835	0.308	0.219	0.257	0.043	0.763
INDKL	25,835	0.066	0.074	0.041	0.014	0.237
FIRMAGE	25,835	2.401	0.986	2.398	0.693	4.094
LOGSALGR	25,835	0.158	0.311	0.114	-0.181	0.633
LOGASSETS	25,835	6.129	1.721	5.999	3.520	9.201
VOLAT	25,835	0.503	0.242	0.449	0.218	0.987
INDCONC	25,835	0.211	0.172	0.161	0.039	0.567

Table 3
Unionization and Expected Returns: Portfolio Sorts and Regression Analysis

Panel A reports portfolio sorts. For each year we sort firms into quintile portfolios based on their previous year's unionization rate. We then compute the equal- and value-weighted Fama-French cost of equity (*FFCOE*) and implied cost of equity (*ICOE*) for each quintile portfolio, and subsequently take the average for each quintile across years. The last row reports p-values corresponding to the t-test of the differences in means between Quintile 5 and Quintile 1. **Panel B** reports the results from OLS tests of expected returns (both *FFCOE* and *ICOE*) on lagged unionization (*UNION*) and a set of control variables. *SALESBETA* is a firm's revenue cyclicality; *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 Census Industry Classification (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 one-digit SIC dummies (not reported). 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 10%, 5%, and 1% levels, respectively.

Panel A: Portfolio Sorts

	FFCOE			*ICOE*		
Unionization Quintile	Unionization (%)	Equal-weighted (%)	Value-weighted (%)	Unionization (%)	Equal-weighted (%)	Value-weighted (%)
Q1 (Lowest)	2.548	13.841	9.571	2.21	9.811	8.277
Q2	5.877	14.002	10.228	5.32	9.895	8.392
Q3	11.269	14.017	10.451	10.53	11.040	8.532
Q4	16.494	14.469	11.515	15.77	11.792	10.574
Q5 (Highest)	33.261	14.777	11.878	33.23	12.545	10.374
Q5 – Q1		0.936***	2.307***		2.734***	2.098**
p-value		0.002	0.000		0.001	0.036

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Panel B: Regressions

	FFCOE			*ICOE*		
	(1)	(2)	(3)	(4)	(5)	(6)
UNION	0.046***	0.042***	0.037**	0.116***	0.122***	0.112***
	(3.38)	(3.30)	(2.56)	(3.91)	(4.14)	(4.10)
SALESBETA		0.002**	0.002*		0.004**	0.004**
		(2.17)	(1.78)		(2.45)	(2.53)
FINLEV		0.039***	0.034***		0.034***	0.036***
		(11.43)	(11.10)		(4.08)	(4.68)
FA/TA		-0.002	-0.003		0.017*	0.011
		(0.43)	(0.65)		(1.74)	(1.43)
INDKL		0.017	0.010		-0.119**	-0.111**
		(0.84)	(0.50)		(2.45)	(2.43)
FIRMAGE			-0.004***			-0.001
			(3.69)			(1.25)
LOGSALGR			-0.001			-0.006**
			(1.48)			(2.42)
LOGASSETS			0.004***			-0.003***
			(3.59)			(3.81)
VOLAT			0.012***			-0.036***
			(5.12)			(5.31)
INDCONC			0.004			0.042***
			(0.90)			(3.53)
Intercept	0.129***	0.110***	0.097***	0.089***	0.064***	0.068***
	(22.67)	(12.79)	(9.72)	(5.76)	(3.65)	(2.86)
Industry Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	63,505	63,505	63,505	25,835	25,835	25,835
R-squared	0.02	0.03	0.03	0.26	0.30	0.33

Table 4
Unionization and Expected Returns: Additional Industry Controls

The table reports the results from OLS tests of expected returns (both *FFCOE* and *ICOE*) on lagged unionization (*UNION*) and a set of control variables. The firm and industry level control variables included in previous specifications are *SALESBETA*, a firm's revenue cyclicality; *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 Census Industry Classification (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 *INDLOGAGE*, which is the logarithm of the age of the oldest firm in a CIC industry; *OLDECON*, which is an indicator variable equal to one if a firm operates in an "old-economy" industry, and zero otherwise. We define old-economy industries as industries with SIC codes less than 4000, comprising firms that are not in the computer, software, internet, telecommunications, or networking industries. *NEWECON* is an indicator variable equal to one if a firm operates in a "new-economy" industry, and zero otherwise. We define new-economy industries as those that comprise firms that are in the computer, software, internet, telecommunications, or networking industries. *INDRDEXP* is the median ratio of R&D expenses to assets in a CIC industry; *1YR-INDGRWTH* is the median one-year growth in the logarithm of firm assets in a CIC industry; *5YR-INDGRWTH* is the median five-year growth in the logarithm of firm assets in a CIC industry; and *INDPROFIT* is the median *ROE* in a CIC industry. All regressions include year and one-digit SIC dummies (not reported). 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 10%, 5%, and 1% levels, respectively.

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	FCOE				*ICOE*			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Variables in Main Specification								
UNION	0.038***	0.037***	0.037***	0.035**	0.131***	0.129***	0.127***	0.115***
	(2.74)	(2.64)	(2.66)	(2.56)	(4.25)	(3.99)	(4.03)	(3.91)
SALESBETA	0.002*	0.002**	0.002**	0.002*	0.004**	0.004**	0.004**	0.004**
	(1.75)	(2.13)	(2.03)	(1.97)	(2.31)	(2.58)	(2.42)	(2.23)
FINLEV	0.034***	0.033***	0.033***	0.033***	0.034***	0.034***	0.033***	0.033***
	(11.16)	(10.99)	(10.91)	(10.86)	(4.29)	(4.20)	(4.15)	(4.46)
FA/TA	-0.003	-0.004	-0.004	-0.004	0.014*	0.013*	0.012	0.008
	(0.60)	(0.86)	(0.94)	(1.04)	(1.77)	(1.67)	(1.57)	(1.22)
INDKL	0.019	0.021	0.021	0.022	-0.118**	-0.116**	-0.114**	-0.085
	(0.93)	(1.08)	(1.09)	(1.18)	(2.20)	(2.15)	(2.10)	(1.57)
FIRMAGE	-0.004***	-0.004***	-0.004***	-0.004***	-0.001	-0.001	-0.001	-0.001
	(3.59)	(3.64)	(3.64)	(3.69)	(1.04)	(1.04)	(1.16)	(1.24)
LOGSALGR	-0.001	-0.001*	-0.001	-0.001	-0.006**	-0.007***	-0.006**	-0.006***
	(1.62)	(1.65)	(1.53)	(1.54)	(2.55)	(2.63)	(2.59)	(3.12)
LOGASSETS	0.004***	0.004***	0.004***	0.004***	-0.003***	-0.003***	-0.003***	-0.002***
	(3.63)	(3.68)	(3.70)	(3.69)	(3.68)	(3.76)	(3.78)	(3.22)
VOLAT	0.013***	0.013***	0.013***	0.013***	-0.035***	-0.035***	-0.035***	-0.024***
	(5.48)	(5.48)	(5.43)	(5.73)	(5.04)	(5.11)	(5.07)	(4.62)
INDCONC	0.002	0.000	-0.000	-0.001	0.047***	0.046***	0.045***	0.039***
	(0.41)	(0.01)	(0.01)	(0.11)	(3.60)	(3.37)	(3.43)	(3.45)
Additional Industry Control Variables								
INDLOGAGE	-0.008**	-0.008**	-0.008**	-0.008**	-0.008**	-0.008**	-0.008**	-0.010**
	(2.54)	(2.38)	(2.47)	(2.41)	(2.04)	(2.01)	(2.08)	(2.54)
OLDECON		-0.005	-0.005	-0.005		-0.004	-0.003	-0.000
		(0.67)	(0.66)	(0.64)		(0.41)	(0.38)	(0.01)
NEWECON		-0.012**	-0.011**	-0.011**		-0.008	-0.007	-0.004
		(2.04)	(2.05)	(2.00)		(1.36)	(1.24)	(0.84)
INDRDEXP			0.044	0.037			0.131	0.035
			(0.67)	(0.57)			(1.24)	(0.39)
1YR-INDGRWTH			-0.012	-0.017			-0.023	-0.041**
			(1.06)	(1.37)			(1.25)	(2.52)
5YR-INDGRWTH			-0.001	-0.001			0.000	0.001
			(0.82)	(0.83)			(0.41)	(0.99)
INDPROFIT				0.011				0.145***
				(1.38)				(4.49)
Intercept	0.151***	0.152***	0.154***	0.153***	0.147***	0.149***	0.154***	0.138***
	(9.67)	(10.08)	(10.20)	(10.19)	(6.63)	(7.93)	(7.97)	(7.74)
Industry Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	63,131	63,131	63,131	63,131	23,182	23,182	23,182	23,182
R-squared	0.03	0.03	0.03	0.03	0.32	0.32	0.32	0.34

Table 5
Cross-Sectional Variation in the Unionization Effect

The table reports the results from OLS tests of expected returns (both Fama-French cost of equity (*FFCOE*) and implied cost of equity (*ICOE*)) on lagged unionization (*UNION*), interaction terms, *and all the control variables we list below.* The interacting variables are *UNEMPL*, the unemployment rate within a Census Industry Classification (CIC) industry; *DEMOCRAT*, the fraction of the workers within a CIC industry that are located in a Democratic-Party state, where we identify a Democratic-Party state by an indicator variable that is equal to one for four consecutive years following a presidential election if the majority of electoral votes in the state where a firm is located were cast for Democrats, and zero otherwise; *BUSCONC* is the Herfindahl index measuring the concentration of a firm's sales across its business segments. All interacted variables (*UNION*, *UNEMPL*, *DEMOCRAT*, and *BUSCONC*) are first demeaned. The control variables are *SALESBETA*, a firm's revenue cyclicality; *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; *INDLOGAGE*, the logarithm of the age of the oldest firm in a CIC industry. *OLDECON* is an indicator variable equal to one if a firm operates in an "old-economy" industry, and zero otherwise. We define old-economy industries as industries with SIC codes less than 4000, comprising firms that are not in the computer, software, internet, telecommunications, or networking industries. *NEWECON* is an indicator variable equal to one if a firm operates in a "new-economy" industry, and zero otherwise. We define new-economy industries as those that comprise firms in the computer, software, internet, telecommunications, or networking industries. *INDRDEXP* is the median ratio of *R&D* expenses to assets in a CIC industry. *1YR-INDGRWTH* is the median one-year growth in the logarithm of firm assets in a CIC industry; *5YR-INDGRWTH* is the median five-year growth in the logarithm of firm assets in a CIC industry; and *INDPROFIT* is the median *ROE* in a CIC industry. All regressions include a constant term and year and one-digit SIC dummies (not reported). 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 10%, 5%, and 1% levels, respectively.

		FFCOE			*ICOE*	
	(1)	(2)	(3)	(4)	(5)	(6)
UNION	0.030**	0.036***	0.036**	0.106***	0.126***	0.119***
	(2.15)	(2.61)	(2.57)	(3.66)	(4.36)	(3.89)
UNEMPL	0.001***			0.001**		
	(2.84)			(2.02)		
UNION×UNEMPL	-0.004***			-0.006*		
	(2.65)			(1.72)		
DEMOCRAT		-0.013***			-0.022***	
		(2.81)			(2.68)	
UNION×DEMOCRAT		0.040**			0.143***	
		(2.21)			(5.37)	
BUSCONC			-0.008***			-0.003
			(2.73)			(0.70)
UNION×BUSCONC			0.049**			0.064**
			(2.33)			(1.99)
All Controls From Table 4	Yes	Yes	Yes	Yes	Yes	Yes
Industry Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	63,131	63,131	60,618	23,182	23,182	22,615
R-squared	0.04	0.03	0.03	0.34	0.36	0.34

Table 6
Unionization and Expected Returns: Instrumental-Variables Estimation

Panel A reports the results from the first-stage regressions of unionization (*UNION*) on the instrumental variables (*FEMALE* and *WORKERAGE*) and the exogenous control variables included in the second-stage regression. *FEMALE* is the percentage of female workers in a firm's CIC industry, and *WORKERAGE* is the average age of workers in a firm's CIC industry. The control variables are *SALESBETA*, a firm's revenue cyclicality; *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; and *INDLOGAGE* is the logarithm of the age of the oldest firm in a CIC industry. *OLDECON* is an indicator variable equal to one if a firm operates in an "old-economy" industry, and zero otherwise. We define old-economy industries as industries with SIC codes less than 4000, comprising firms are not in the computer, software, internet, telecommunications, or networking industries. *NEWECON* is an indicator variable equal to one if a firm operates in a "new-economy" industry, and zero otherwise. We define new-economy industries as those that comprise firms in the computer, software, internet, telecommunications, or networking industries. *INDRDEXP* is the median ratio of R&D expenses to assets in a CIC industry. *1YR-INDGRWTH* is the median one-year growth in the logarithm of firm assets in a CIC industry; *5YR-INDGRWTH* is the median five-year growth in the logarithm of firm assets in a CIC industry; and *INDPROFIT* is the median ROE 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 second-stage equation. **Panel B** reports the results from the second-stage regressions of expected returns (both *FFCOE* and *ICOE*) on lagged unionization (*UNION*) and control variables, in which we treat unionization as the endogenous variable. The Hausman (1978) test examines whether the OLS and 2SLS coefficients on *UNION* are statistically different from each other. All regressions include a constant term and year and one-digit SIC dummies (not reported). 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 10%, 5%, and 1% levels, respectively.

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Panel A: First-Stage Regressions of UNION and Validity of Instruments

	FFCOE Sample		*ICOE Sample*	
	(1)	(2)	(3)	(4)
Instruments				
FEMALE	-0.002***	-0.002***	-0.002***	-0.002***
	(3.24)	(3.20)	(3.15)	(3.18)
WORKERAGE	0.012***	0.012***	0.012***	0.012***
	(2.86)	(2.84)	(3.00)	(2.97)
Predetermined Variables in Main Specification				
SALESBETA	-0.004	-0.003	-0.001	-0.001
	(0.56)	(0.55)	(0.19)	(0.08)
FINLEV	0.011	0.006	0.032*	0.025*
	(1.05)	(0.68)	(1.97)	(1.81)
FA/TA	0.084***	0.069**	0.099***	0.083**
	(2.80)	(2.38)	(2.84)	(2.42)
INDKL	0.027	0.046	0.059	0.093
	(0.20)	(0.35)	(0.40)	(0.63)
FIRMAGE	0.009***	0.006***	0.002	-0.000
	(3.77)	(2.93)	(1.05)	(0.26)
LOGSALGR	-0.003*	-0.003*	-0.005*	-0.002
	(1.80)	(1.91)	(1.91)	(0.80)
LOGASSETS	0.006***	0.006***	0.006***	0.007***
	(3.37)	(3.76)	(2.67)	(3.37)
VOLAT	-0.005	0.002	-0.024*	-0.006
	(0.86)	(0.59)	(1.74)	(0.69)
INDCONC	0.045	0.034	0.037	0.025
	(1.52)	(1.25)	(1.10)	(0.81)
Additional Industry-Level Predetermined Variables				
INDLOGAGE		0.020		0.018
		(1.28)		(1.33)
OLDECON		0.044		0.061**
		(1.44)		(2.04)
NEWECON		-0.022		-0.012
		(1.22)		(1.01)
INDRDEXP		0.123		0.335
		(0.36)		(0.98)
1YR-INDGRWTH		-0.112***		-0.119***
		(3.32)		(3.99)
5YR-INDGRWTH		0.002		0.001
		(1.45)		(0.87)
INDPROFIT		0.208***		0.175***
		(3.38)		(2.80)
Industry Fixed Effects	Yes	Yes	Yes	Yes
Year Fixed Effects	Yes	Yes	Yes	Yes
Observations	63,131	63,131	23,180	23,180
R-squared	0.50	0.52	0.52	0.54
Predictive Power of Excluded Instruments				
Partial R-Squared	0.143	0.131	0.152	0.145
Robust F-Statistic	9.310	8.570	9.220	9.080
p-value	0.000	0.000	0.000	0.000
Test of Overidentifying Restrictions				
Hansen's J-Statistic	1.186	1.528	0.541	0.559
p-value	0.276	0.216	0.462	0.455

Panel B: Second-Stage Regressions of Expected Returns on Unionization

	FFCOE		*ICOE*	
	(1)	(2)	(3)	(4)
Potentially Endogenous Instrumented Variable				
UNION	0.059*	0.064*	0.160**	0.161**
	(1.83)	(1.89)	(2.13)	(2.17)
Predetermined Variables in Main Specification				
SALESBETA	0.002**	0.002**	0.005**	0.004**
	(1.97)	(2.15)	(2.15)	(2.13)
FINLEV	0.034***	0.033***	0.034***	0.031***
	(11.29)	(10.90)	(3.93)	(4.14)
FA/TA	-0.004	-0.006	0.011	0.006
	(0.81)	(1.33)	(1.59)	(0.94)
INDKL	0.009	0.017	-0.133**	-0.096
	(0.41)	(0.79)	(2.30)	(1.64)
FIRMAGE	-0.005***	-0.004***	-0.002	-0.001
	(4.05)	(4.01)	(1.61)	(1.35)
LOGSALGR	-0.001	-0.001	-0.006***	-0.006***
	(1.48)	(1.41)	(2.75)	(3.23)
LOGASSETS	0.004***	0.004***	-0.003***	-0.003***
	(3.27)	(3.37)	(4.27)	(3.91)
VOLAT	0.013***	0.013***	-0.035***	-0.024***
	(5.26)	(5.68)	(4.41)	(4.25)
INDCONC	0.005	-0.002	0.049***	0.037***
	(0.98)	(0.38)	(3.88)	(3.61)
Additional Industry-Level Predetermined Variables				
INDLOGAGE		-0.008**		-0.011***
		(2.38)		(2.63)
OLDECON		-0.007		-0.004
		(0.90)		(0.36)
NEWECON		-0.011**		-0.004
		(2.07)		(0.86)
INDRDEXP		0.053		0.042
		(0.79)		(0.42)
1YR-INDGRWTH		-0.012		-0.033**
		(0.89)		(2.23)
5YR-INDGRWTH		-0.001		0.001
		(0.97)		(0.82)
INDPROFIT		0.004		0.136***
		(0.31)		(3.98)
Industry Fixed Effects	Yes	Yes	Yes	Yes
Year Fixed Effects	Yes	Yes	Yes	Yes
Observations	63,131	63,131	23,180	23,180

Hausman Test for the Effect of Unionization

Cluster-Robust t-Statistic	0.750	0.940	0.600	0.860
p-value	0.455	0.347	0.550	0.393

Table 7
Unionization and Loadings on the Fama-French Factors

Panel A reports portfolio sorts. For each year, we sort firms into quintile portfolios based on their previous year's unionization rate. We then compute the equal-weighted averages of each loading on the Fama-French factors: Market beta (*MKTBETA*), SMB beta (*SMBBETA*), and HML beta (*HMLBETA*). Finally, we average the loadings across years. The last row reports p-values corresponding to the t-test of the differences in means between Quintile 5 and Quintile 1. **Panel B** reports the results from OLS tests of *MKTBETA*, *SMBBETA*, and *HMLBETA* on lagged unionization (*UNION*) and a set of control variables. *SALESBETA* is a firm's revenue cyclicality; *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 Census Industry Classification (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 one-digit SIC dummies (not reported). 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 10%, 5%, and 1% levels, respectively.

Panel A: Loadings on the Fama-French Factors

Unionization Quintile	*UNION* (%)	*MKTBETA*	*SMBBETA*	*HMLBETA*
Q1 (Lowest)	2.548	1.020	0.894	-0.062
Q2	5.877	0.967	0.842	0.081
Q3	11.269	0.970	0.915	0.040
Q4	16.494	1.009	0.843	0.115
Q5 (Highest)	33.261	0.960	0.667	0.358
Q5 – Q1		-0.060**	-0.227***	0.420***
p-value		0.039	0.000	0.000

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Panel B: Regressions of MKTBETA, SMBBETA, and HMLBETA on UNION

	MKTBETA		*SMBBETA*		*HMLBETA*	
	(1)	(2)	(3)	(4)	(5)	(6)
UNION	-0.018	-0.135	-0.607***	-0.047	1.321***	0.916***
	(0.16)	(1.40)	(3.17)	(0.31)	(5.60)	(4.86)
SALESBETA		0.007		0.003		0.017
		(0.88)		(0.34)		(1.26)
FINLEV		-0.028		0.129***		0.601***
		(0.91)		(3.40)		(8.96)
FA/TA		-0.181***		-0.303***		0.388***
		(5.09)		(5.28)		(6.02)
INDKL		0.251*		-0.184		-0.170
		(1.72)		(0.85)		(0.49)
FIRMAGE		-0.100***		-0.169***		0.143***
		(11.53)		(10.06)		(7.81)
LOGSALGR		0.009		-0.002		-0.050***
		(1.20)		(0.16)		(3.61)
LOGASSETS		0.107***				-0.050***
		(16.72)				(3.67)
VOLAT		0.292***		0.386***		-0.323***
		(12.07)		(9.52)		(5.42)
INDCONC		-0.128***		-0.079		0.366***
		(2.84)		(1.64)		(4.21)
Intercept	0.699***	0.436***	0.875***	1.226***	0.274***	-0.337***
	(28.22)	(5.45)	(10.44)	(11.12)	(3.67)	(2.79)
Industry Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	63,505	63,505	63,505	63,505	63,505	63,505
R-squared	0.02	0.09	0.04	0.10	0.06	0.11

Table 8
Unionization and Expected Returns: Controlling for Operating Leverage

The table reports the results from OLS tests of expected returns (both *FFCOE* and *ICOE*) on lagged unionization (*UNION*) and a set of control variables. *SALESBETA* is a firm's revenue cyclicality; *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 Census Industry Classification (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 one-digit SIC dummies (not reported). In this table, we extend the analysis in Table 3 to include two additional controls: *MRDOL* is the Mandelker and Ree (1984) measure of operating leverage, which we calculate as a sensitivity of a firm's operating income after depreciation to its sales; and *LS* is labor stock, which we define as the number of a firm's employees divided by total assets. 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 10%, 5%, and 1% levels, respectively.

	FFCOE			*ICOE*		
	(1)	(2)	(3)	(4)	(5)	(6)
UNION	0.037**	0.037**	0.036**	0.112***	0.113***	0.111***
	(2.56)	(2.55)	(2.54)	(4.10)	(4.20)	(4.20)
SALESBETA	0.002*	0.002*	0.002*	0.004**	0.004**	0.004**
	(1.78)	(1.75)	(1.71)	(2.53)	(2.35)	(2.28)
FINLEV	0.034***	0.034***	0.034***	0.036***	0.037***	0.037***
	(11.10)	(11.21)	(11.21)	(4.68)	(4.79)	(4.79)
FA/TA	-0.003	-0.004	-0.005	0.011	0.005	0.005
	(0.65)	(0.93)	(0.98)	(1.43)	(0.67)	(0.57)
INDKL	0.010	0.015	0.015	-0.111**	-0.095**	-0.095**
	(0.50)	(0.74)	(0.75)	(2.43)	(2.05)	(2.08)
FIRMAGE	-0.004***	-0.004***	-0.004***	-0.001	-0.002*	-0.002**
	(3.69)	(3.87)	(3.90)	(1.25)	(1.84)	(2.07)
LOGSALGR	-0.001	-0.001	-0.001	-0.006**	-0.006**	-0.005**
	(1.48)	(1.54)	(1.52)	(2.42)	(2.30)	(2.17)
LOGASSETS	0.004***	0.004***	0.004***	-0.003***	-0.002***	-0.002***
	(3.59)	(3.63)	(3.62)	(3.81)	(2.68)	(2.70)
VOLAT	0.012***	0.013***	0.013***	-0.036***	-0.033***	-0.034***
	(5.12)	(5.21)	(5.32)	(5.31)	(5.21)	(5.29)
INDCONC	0.004	0.004	0.004	0.042***	0.041***	0.041***
	(0.90)	(0.86)	(0.84)	(3.53)	(3.44)	(3.46)
LS		0.125*	0.122*		0.476***	0.475***
		(1.81)	(1.78)		(2.79)	(2.83)
MRDOL / 100			0.034***			0.137***
			(2.77)			(3.15)
Intercept	0.097***	0.096***	0.096***	0.068***	0.062***	0.061***
	(9.72)	(9.73)	(9.77)	(2.86)	(2.64)	(2.60)
Industry Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	63,505	63,505	63,505	25,835	25,835	25,835
R-squared	0.03	0.03	0.03	0.33	0.34	0.34

Table 9
Unionization and Unlevered Cost of Equity

The table reports the results from OLS tests of the unlevered cost of equity on lagged unionization (*UNION*) and a set of control variables. We unlever cost of equity using the Modigliani-Miller formula. Columns (1) and (2) correspond to the unlevered Fama-French cost of equity (*UFFCOE*) and columns (3) and (4) correspond to the unlevered implied cost of equity (*UICOE*). *SALESBETA* is a firm's revenue cyclicality; *FA/TA* is net fixed assets divided by total assets; *INDKL* is average net fixed assets per employee in $000s within a Census Industry Classification (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 one-digit SIC dummies (not reported). 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 10%, 5%, and 1% levels, respectively.

	UFFCOE		*UICOE*	
	(1)	(2)	(3)	(4)
UNION	0.023**	0.026**	0.092***	0.086***
	(2.43)	(2.45)	(3.59)	(3.61)
SALESBETA	0.001	0.001	0.004**	0.004**
	(1.48)	(1.58)	(2.44)	(2.60)
FA/TA	-0.004	-0.003	0.017**	0.015**
	(1.34)	(0.83)	(2.06)	(2.19)
INDKL	0.013	0.014	-0.105**	-0.097**
	(0.88)	(0.93)	(2.49)	(2.47)
FIRMAGE	-0.005***	-0.003***	-0.002*	-0.002*
	(4.42)	(3.07)	(1.90)	(1.75)
LOGSALGR		0.001		-0.005**
		(1.03)		(2.19)
LOGASSETS /100		-0.044		-0.328***
		(0.50)		(4.87)
VOLAT		0.007***		-0.025***
		(3.17)		(3.85)
INDCONC		0.001		0.037***
		(0.22)		(3.40)
Intercept	0.136***	0.129***	0.079***	0.082***
	(15.65)	(14.03)	(4.80)	(3.88)
Industry Fixed Effects	Yes	Yes	Yes	Yes
Year Fixed Effects	Yes	Yes	Yes	Yes
Observations	43,867	43,867	20,355	20,355
R-squared	0.02	0.02	0.29	0.33

Table 10
Year-by-Year Regressions

Columns (1) and (4) of the table report the coefficient on *UNION* from the year-by-year regressions of expected returns (Fama-French cost of equity (*FFCOE*) and the implied cost of equity (*ICOE*)) on unionization rates and a set of control variables. *UNION* and all control variables are lagged one period. *SALESBETA* is a firm's revenue cyclicality; *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 Census Industry Classification (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 one-digit SIC dummies (not reported). The absolute values of *t*-statistics (in parentheses) are based on robust standard errors clustered at the CIC industry level. Columns (3) and (6) show the contribution of unions to total expected returns. We estimate this contribution by taking the product of the estimated coefficient on *UNION* and the average unionization rate in the corresponding year. We express this number in percentage points. *, **, and *** indicate statistical significance at 10%, 5%, and 1% levels, respectively.

		FFCOE			*ICOE*	
	Coef. on *UNION*	*t*-stat	Coef × Avg. Unionization	Coef. on *UNION*	*t*-stat	Coef × Avg. Unionization
	(1)	(2)	(3)	(4)	(5)	(6)
1984	0.035***	(3.39)	0.820	0.007	(0.41)	0.176
1985	0.034**	(1.98)	0.712	0.023*	(1.72)	0.498
1986	0.026	(1.16)	0.511	0.059***	(3.41)	1.168
1987	0.034**	(2.56)	0.621	0.063***	(2.87)	1.194
1988	0.034*	(1.85)	0.597	0.066**	(2.60)	1.238
1989	0.048***	(3.40)	0.790	0.074**	(2.47)	1.336
1990	0.073***	(4.44)	1.168	0.102***	(2.79)	1.797
1991	0.059***	(3.21)	0.900	0.121***	(3.91)	2.032
1992	0.045*	(1.76)	0.678	0.103***	(3.43)	1.673
1993	0.04	(1.45)	0.591	0.119***	(4.06)	1.802
1994	0.058**	(2.25)	0.832	0.141***	(4.61)	2.075
1995	0.027	(1.14)	0.372	0.125***	(3.81)	1.751
1996	0.027	(0.88)	0.338	0.134***	(4.49)	1.788
1997	0.045*	(1.71)	0.549	0.123***	(4.55)	1.566
1998	0.023	(0.85)	0.274	0.158***	(4.29)	1.889
1999	0.012	(0.48)	0.138	0.211***	(5.60)	2.399
2000	-0.011	(0.42)	-0.121	0.176***	(4.15)	1.990
2001	0.02	(0.88)	0.198	0.167***	(4.98)	1.655
2002	0.046*	(1.82)	0.458	0.190***	(5.55)	1.855
2003	0.078***	(2.79)	0.721	0.202***	(5.12)	1.910
2004	0.097***	(3.61)	0.877	0.227***	(4.84)	2.086
2005	0.104***	(3.50)	0.907	-	-	-

Table 11
Alternative Econometric Methods

This table reports the coefficient on *UNION* from estimating regressions of expected returns on unionization rates and a set of control variables using different econometric methodologies. Column (1) reports coefficients for the Fama-French cost of equity (*FFCOE*). Column (2) reports coefficients for the implied cost of equity (*ICOE*). *UNION* and all control variables are lagged one period. *SALESBETA* is a firm's revenue cyclicality; *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 Census Industry Classification (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 one-digit SIC dummies. The coefficients on the control variables are omitted for brevity. As a benchmark, in row 1 we repeat the results from OLS estimation with standard errors clustered by CIC industry. In rows 2 and 3, we cluster standard errors by firm and by year, respectively. Using two-dimensional clustering, in row 4 we cluster standard errors by firm and year, while in row 5 we cluster standard errors by CIC industry and year. Row 6 reports the coefficient from the Fama-MacBeth regressions with Newey-West standard errors adjusted for autocorrelation using six lags. The absolute values of *t*-statistics are reported in parentheses. *, **, and *** indicate statistical significance at 10%, 5%, and 1% levels, respectively.

	Coefficient on *UNION*	
Econometric Method	*FFCOE*	*ICOE*
	(1)	(2)
OLS clustering by CIC	0.037**	0.112***
	(2.56)	(4.10)
OLS clustering by firm	0.037***	0.112***
	(5.91)	(13.63)
OLS clustering by year	0.037***	0.112***
	(7.42)	(9.65)
OLS clustering by firm and year	0.037***	0.112***
	(5.02)	(8.08)
OLS clustering by CIC and year	0.037***	0.112***
	(2.47)	(3.79)
Fama-MacBeth with Newey-West standard errors	0.043***	0.123***
	(5.83)	(4.55)