12-1997

Unions and Managerial Pay

John DiNardo
University of California-Irvine

Kevin F. Hallock
Cornell University, kfh7@cornell.edu

Jorn-Steffen Pischke
Massachusetts Institute of Technology

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Abstract

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Keywords

unions, management, wage distributions, managerial pay, compensation

Comments

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Working Paper **6318**

NBER WORKING PAPER SERIES

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Working Paper 6318
http://www.nber.org/papers/w6318

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
December 1997

We thank John Abowd, Danny Blanchflower, Barry Hirsch, Kevin Murphy, Joe Tracy and Michael Wallerstein for data, Ben Gordon for excellent research assistance, and Daron Acemoglu, Danny Blanchflower, Donald Deere, Sue Dynarski, Wally Hendricks, and Barry Hirsch for comments. The American Compensation Association provided generous support. Any opinions expressed are those of the authors and not those of the National Bureau of Economic Research.

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NBER Working Paper No. 6318
December 1997
JEL Nos. J31, J44, J51

ABSTRACT

Unions compress the wage distribution among workers covered by union contracts. We ask whether unions also have an effect on the managers of unionized firms. To this end we collected and assembled data on unionization and managerial pay within firms and industries in the U.S. and across countries. Generally, we find a negative correlation between executive compensation and unionization in our cross-section data, but no relationship of changes in unionization on the growth of compensation of executives over time. Using NLRB elections data, we find that a loss of union members due to decertification elections is associated with higher CEO pay, although our estimates are imprecise. With CPS data we consistently find that where unions are stronger, fewer managers are employed.

John DiNardo
Department of Economics
University of California-Irvine
Irvine, CA 92717
and NBER
jdinardo@uci.edu

Kevin Hallock
Department of Economics
University of Illinois at Urbana-Champaign
Champaign, IL 61820
hallock@uiuc.edu

Jörn-Steffen Pischke
Department of Economics
Massachusetts Institute of Technology
50 Memorial Drive
Cambridge, MA 02139
and NBER
jsp@mit.edu

1 Introduction

Union representation limits management's discretion to set the wages of covered workers. Management is required to bargain with the workers' union. The effect of this process on the wages of union workers is clear: unions raise the wages of their members. Moreover, union wage setting is more egalitarian, equalizing wages *within* establishments (Freeman 1982) and *across* establishments (Slichter, Healy and Livernash 1960). "Single rate" and related policies have long been associated with unionism in the U.S. and abroad, *inter alia* out of a concern for equity (Burda 1995), worker solidarity and organizational unity. In this paper we look elsewhere for "union wage effects." Specifically we ask whether unions also impact the pay of non–union workers in the firm, specifically a firm's managers and executives.

According to rules of the National Labor Relations Board, managers cannot form unions themselves. Nevertheless, there are several mechanisms by which the unions of workers in an establishment might affect the pay of managers. At bargaining time and more visibly during periods of industrial disputes, workers often voice concern over the pay and compensation of executives. Anecdotal evidence suggests that when workers are organized, executives or directors feel more pressure to limit the pay of top managers. DeAngelo and DeAngelo (1991), for example, find that corporate executives and other white collar workers accepted major pay cuts in years where steel firms sought to obtain pay concessions from blue collar workers in union negotiations. To the best of our knowledge, however, the link between unionization and executive pay has not been previously formally investigated.

Our work is related to three different literatures. First, our paper is linked to the literature on the impact of unions on economic outcomes. While a tremendous amount of evidence exists that unions raise wages and benefits for their members, the effect of unions on the wages of non-union workers in the same or competing firms is less clear (Freeman and Medoff 1984, Neumark and Wachter 1995, Rosen 1969).[1] Abowd (1989) finds, for instance, that collective bargains are "efficient" (in the sense of maximizing the sum of shareholders' and union members' wealth). Given the evidence that profits are lower in unionized firms (Ruback and Zimmerman 1984) this suggests that unions may redistribute rents from the owners of firms to the workers. If this is the case, it would be interesting to know whether other constituents in the firm gain or lose from the presence of unions. Freeman and Medoff (1984) briefly discuss the impact of unions on white collar workers within the same firm (who seem to gain from unionization), but they do not discuss the most highly paid workers: managers and top executives.

Second, our work contributes to the growing literature on CEO compensation. The compensation of top executives varies widely across firms. Although the focus of much of the literature has been on the role of firm performance in CEO compensation (Jensen and Murphy 1990), it appears to account for little of the variation in the level of CEO pay across firms (abstracting from variability introduced by revaluations of stock and option grants as stressed by Hall and Liebman (1997)). A related literature has attempted to explain cross-sectional differences in CEO pay. Joskow, Rose and Shepard (1993), for example, ask whether political pressures may limit the level of CEO pay. They find that CEO compensation is lower in regulated industries and that pay is less sensitive to firm performance in these industries. Joskow, Rose and Wolfram (1994) link CEO pay in the electric utility industry even more directly to specific regulatory practices. Akin to regulation, which gives a particular stakeholder group (customers) more voice, unions may be able to affect the pay of CEOs for their own benefit.

Third, there is a link between our investigation and the recent literature on changes in

[1]Indeed, Pencavel (1994) notes that "while there may be a good deal of evidence of a qualitative kind to support the notion that the presence of trade unions affects the wages paid to nonunion workers, good, solid evidence on the direction, size, and variations of these effects does not exist."

wage inequality. Freeman (1993), Card (1992), and DiNardo, Fortin and Lemieux (1996) have found that the decline in unionization can explain roughly 30 percent of the recent increase in wage inequality in the U.S., while Bell and Pitt (1995) put this number at about 25 percent for the U.K. Moreover, DiNardo and Lemieux (1997) and Freeman (1996) also find an association between unionization and the differences and changes in wage inequality across countries. These studies, however, focus only on the unions' wage impacts on covered workers. Since unionized workers typically occupy the lower and middle part of the wage distribution, these papers do not attempt to account for changes in wage inequality that have taken place among non-unionized workers in the top of the wage distribution. If the presence of unions has an impact on the pay of the highly paid as well, however, the decline of unionization in countries like the U.S. and the U.K. might explain an even bigger part of the spreading wage distributions.

Data on unionization and managerial pay are not readily available. We have therefore collected data from a variety of sources and we perform a number of different comparisons in our investigation. We start in section 2 by investigating cross-national data on unionization and CEO pay, which differs substantially across countries. Abowd and Bognanno (1995) have assembled information on CEO compensation for 12 countries, which we use in conjunction with unionization rates obtained for these countries.

In section 3, we use several sources of data on the compensation of CEOs (originally collected by Kevin Murphy) and unionization of firms. Since the pay of the top five executives of publicly held companies has to be disclosed annually, we focus on CEO compensation in this paper as does most of the literature on executive pay. Hirsch (1991) and Bronars, Deere and Tracy (1994) have assembled data on the unionization of particular firms. We use their data together with the pay data for CEOs to investigate the link between unionization and executive pay.

In reporting correlations between unionization and managerial pay, we are not necessarily able to establish whether unions actually cause differences in compensation for the groups we study. For example, it might merely be the case that unions are to be found in industries or firms which have higher rents. If these rents are shared among various groups, including workers and managers, we would find a positive correlation between unionization and managerial pay. We try a variety of strategies to understand these issues, however each is imperfect. For example, we use control variables to capture the level of rents in different firms, but these may well be inadequate. We also look at changes in unionization over time, but these changes themselves may have a particular correlation with changes in the well-being of firms.

In another attempt to address the potential endogeneity of unionization we match our CEO data to data from the NLRB election files on union certification and decertification elections in section 4. This allows us to investigate the impact of changes in unionization on executive pay growth conditional on an election. These changes are arguably more likely to be random than the choice of establishments which are targetted for organization. The impact of representation elections on numerous outcomes has been analyzed in the literature (Ruback and Zimmerman 1984, Bronars and Deere 1993). Unfortunately, our election sample is small, and other explanations are also consistent with the results.

The analyses discussed above only refer to top executives. In section 5, we use a panel of grouped data by region and industry, which we constructed from the 1983 to 1993 Current Population Surveys to gain some idea about the impact of unions on the pay of managers other than CEOs. Related results using a similar strategy are also reported by Neumark and Wachter (1995).

None of these data sources are ideal so we would like to interpret our results with caution. Together, however, the various data and approaches provide an interesting and informative portrait of the association of managerial pay and unionization. To preview our findings, we generally find that unions are associated with lower CEO pay in cross-sectional regressions. Changes in managerial pay, on the other hand, bear less of a relationship with changes in unionization rates, and sometimes we find a positive relationship in changes. Finally, there is consistent evidence that where unions are stronger fewer managers are employed.

2 International Comparisons

Many observers have noted that American CEOs tend to be paid much more than CEOs in other countries, whether in terms of direct comparisons or relative to other workers in the economy (see for example Crystal (1992), and Bok (1993)). Abowd and Bognanno (1995) have assembled comparable data on CEO pay from a variety of compensation consulting firms for 12 OECD countries for the most comprehensive international comparison. They document clearly that CEO pay is highest in the U.S. In addition, they try to explain international compensation patterns with the structure of taxation. While they find no impact of corporate taxes, personal marginal tax rates seem to have a large impact on the level of CEO compensation. Nevertheless, a significant U.S. residual remains unexplained.

If it is true that unions lower executive compensation, then the relatively low U.S. unionization rate may help explain the higher levels of pay. In order to assess this possibility, we have obtained the compensation data used by Abowd and Bognanno (1995). We will briefly describe the main features of the dataset, the details of which are documented in the appendix to their paper. The data are for the years 1984 and 1988 to 1992, and cover the countries Belgium, Canada, France, Germany, Italy, Japan, the Netherlands, Spain, Sweden, Switzerland, the U.K., and the U.S. They are based on reports by various compensation consulting firms; and in some instances data from more than one source are available for a single year. Abowd and Bognanno collected data for a variety of high level managers. Like them, we report results for CEOs and top human resource managers only. These are the groups for whom the largest samples are available.

The compensation amounts refer to total compensation, including cash compensation, benefits and perquisites, and long term compensation in the form of stock or option grants. Abowd and Bognanno refer to this as the total compensation cost, the amount a firm has to spend on an executive. An important question is how to make these costs comparable across countries. We use three approaches. The first is a real measure of compensation at actual exchange rates by converting compensation into U.S. dollars for each year and deflating by the U.S. consumer price index. Comparisons using the real measure will reflect the strong depreciation of the dollar over the sample period. This is avoided by using OECD purchasing power parity exchange rates and the OECD index of consumer prices. However, there is necessarily a degree of arbitrariness in PPP exchange rates. The third method is to look at executives relative to a comparison group within the same country. We compare executives to non-supervisory manufacturing workers. This comparison is of course affected by the impact of unions on the pay of manufacturing workers and by trends in their compensation. Nevertheless, in practice we obtain relatively similar results with all three methods. The data on unionization are obtained from the Trade Union Membership Database by Visser (1997).

Tables I to III contain our regression results for the various compensation measures. For each measure, we start by presenting results including the country's unionization rate, time dummies, and dummies for source of the compensation data. Unionization is associated with 50 log points lower CEO pay in the first column of Table I. Column (2) includes a dummy variable for the U.S. and a U.S. specific time trend. It reveals that about half of the previous union effect is due to the presence of the U.S. in the dataset. In the next column, we run a specification similar to Abowd and Bognanno (1995). These regressions include a time trend (instead of time dummies), a U.S. dummy, a U.S. trend, the marginal employee tax rate including both income and payroll taxes, the marginal corporate tax rate, and dummy variables for the data source. Unlike Abowd and Bognanno, we do not control for the average size of a company in our regressions, so that our exact results differ slightly from theirs. Changing the specification in this way does relatively little to the union coefficient although the marginal employee tax rate and the unionization rates have a correlation of about 0.5 in this sample. The fourth column presents results controlling f. r country fixed effects. Unfortunately, there is not enough within country variation in unionization to estimate the effect of unions anymore.

The right hand panel of each of the Tables I to III presents results using the PPP measure instead of the real measure of compensation. The effects on the unionization rate become stronger in the cross-sectional regressions now, but the basic picture is unchanged. In Table II, which presents analogous regressions for top HR managers we also find similar results. The union coefficient in the HR manager regressions is small and insignificant using the real compensation measure, except when the marginal tax rate variables are included. However, the coefficient on the employee tax rate is positive rather than negative in this case. The results are much more similar to the CEO results for the PPP measure. The PPP measure is probably a more sensible measure of cross country differences in pay over this period of large exchange rate fluctuations.

Table III contains results for compensation relative to a manufacturing operative. These results are little different than those using the PPP measure. While unionization rates are associated with a large negative effect on the executive/worker pay gap across countries, this effect becomes rather unstable once we control for country effects.

Overall, we find a large negative association between unionization rates and executive pay in a cross section of countries. Going from no unionization to 100 percent unionization typically implies about 40 to 80 percent lower compensation for CEOs and top HR managers. H wever, these effects are due to between country differences and are not mirrored by a similar association between pay growth and changes in unionization rates within countries.

One question in assessing these results is whether union membership rates are the right variables in this exercise. What matters for the wages of most workers in many countries are coverage rates, not membership. France, for example, had a unionization rate of around 15 percent in the mid-1980s, but 90 percent of the workforce where covered by collective bargaining agreements. Freeman (1996) finds that coverage rates across countries are more strongly associated with wage inequality than unionization rates.

Nevertheless, we think there is a good case to be made that it is union membership we should be interested in here. One channel we are interested in is the role unions play in putting a lid on executive compensation. Obviously, CEOs are not subject to collective bargaining agreements or mandatory extensions. Instead, we believe that such a link would work through more informal political pressures. It is likely that this pressure is larger if more workers care about pay equity, something they voice by being organized in unions.

Another reason for preferring membership rates is more pragmatic: these are available more widely. The OECD (1997) has published data for union coverage for the years 1980, 1990, and 1994 for the countries in our sample. Using these, and interpolating the intervening years, we generally found similar correlations (not reported in the tables) with executive pay to those shown for the membership data. However, the coefficients for coverage rates tend to be somewhat closer to zero than those for membership rates. This lends support to our conjecture that membership might be the more sensible measure in this context.

3 The Unionization of U.S. Firms and CEO Pay

Data on unionization rates of individual firms are not readily available. Nevertheless, it is possible to gather these data through a variety of channels. Bronars et al. (1994) constructed unionization measures for a sample of large U.S. firms from the Bureau of Labor Statistics Bargaining Calendar. The Bargaining Calendar contains information on the workers covered by collective bargaining agreements in particular establishments. By dividing this information on the number of covered workers by employment from Compustat. the authors were able to construct coverage rates for individual firms. One shortcoming of this exercise is that the sample is limited to firms which had any collective bargaining agreements, i.e. a unionization rate greater than zero. This is simply because matching the universe of contracts to the universe of firms is virtually impossible. The sample is also limited to firms with contracts covering at least 1000 workers. Because the union employment data have to be interpolated between contract dates, Bronars, Deere, and Tracy constructed four year averages of their unionization measures to lessen the impact of this smoothing. Their data cover the periods 1971-1974, 1975-1978, and 1979-1982.

A different approach was pursued by Hirsch (1991) who conducted a survey of large U.S. employers in manufacturing about their unionization rates in 1987. The survey contained questions on the current unionization level and retrospective questions about unionization in 1977. The Hirsch data refer to the firm's entire workforce in the U.S. and Canada.

We obtained both the Bronars, Deere, and Tracy data and the Hirsch data for 1977. We merged both datasets with information about the firms from Compustat and with CEO pay data which are published annually in Forbes magazine. We obtained the latter data for the relevant sample years from Kevin Murphy.

In Table IV we run OLS regressions of CEO pay on the Bronars, Deere, and Tracy measure of unionization and a variety of other covariates. Our measure of CEO pay is the logarithm of salary, bonuses, and the value of other compensation but excluding the value of stock or option grants. The latter part of compensation is potentially large for many CEOs but the Securities and Exchange Commission has only recently required firms to uniformly disclose options in a given year. Forbes does report the value of exercised options in a given year which some authors add to salary, bonus, and other compensation to arrive at total compensation. However, this is not really a good measure of total compensation because the value of options should be included in compensation when they are received. not when they are exercised. If we additionally include the value of exercised options in compensation in this section we get very similar results (which are not reported in the tables).

Column (1) includes regressors which are typically found in cross-sectional regressions for the level of CEO compensation: the years of age and tenure of the CEO and the square of these variables, the logarithm of sales and employment to control for firm size effects, and the percent change in the value of the firm (return) during the four year period and its lag. Except for the percent change in the value of the firm (return), these regressors are constructed as the average over the four year period (before taking logs or squares). Since CEOs may turn over during any four year interval we also control for whether the CEO changes during the period. Finally, we control for period effects to capture aggregate growth in CEO pay, and the impact of changing market returns for these firms, etc.

Our key regressor is the fraction of employees covered by union contracts in the firm. We find that CEO pay is lower by about 3 percent for a 10 percentage point increase in unionization. In order to probe whether this relationship is indeed roughly linear in the data we also broke the unionization rates into five categories for 0 to 20 percent unionized, 21 to 40 percent, etc. Dummy variable results are shown in the second column of coefficients in Table IV and they reveal indeed a roughly linear relationship. The only category which does not follow this pattern is firms with more than 80 percent union workers. In those firms CEO pay is as high as in firms where only half the workforce is covered by collective bargaining agreements. However, this cell is very small, so that we do not want to put too much weight on this result.

Hirsch (1991), Bronars et al. (1994), and others have found that unionized firms have lower investment, more capital, lower growth rates, and lower profits. Since such firms may be deemed as less successful by the directors or shareholders, CEO pay could suffer indirectly from the consequences unions have on the firm. In addition, unions are more likely to organize firms with higher capital-labor ratios or with higher rents. The results discussed above reflect both direct effects of unions on CEO pay as well as indirect effects working throught the effect of unions on investment, growth, and profits, or selective organization of higher rent plants. In order to estimate only the direct association of unions and CEO pay we include a number of additional variables which we constructed from Compustat data: sales growth and employment growth from the previous four year period, the log of property, plant, and equipment per worker, a crude measure of Tobin's Q (total assets divided by the sum of the book value of debt plus the market value of equity), and the ratio of investment in property, plant, and equipment to sales. Including these regressors lowers the value of the unionization coefficient by about 40 percent. However, CEO pay is still lower by almost 2 percent for a 10 percentage point change in unionization.

For comparability with the Hirsch data we also ran these regressions separately for manufacturing and non-manufacturing industries. Results are displayed in Table V. The relationship between CEO pay and unionization is weaker within sectors, a 10 percentage point increase in unionization means only about 1 percent lower executive pay. In addition, the relationship is no longer significant. In Table VI we go one step further and exploit the panel nature of the dataset. By including 44 two digit industry effects in the regression we again find an insignificant effect of -0.1 but this effect goes to zero once 103 three digit industry effects are included. Thus, the association between unionization and CEO pay seems to reflect mostly industry differences. A curious result appears when we control for firm effects: in this case we find a positive and significant association between unionization rates and CEO pay. A 10 percentage point increase in unionization means 2.6 percent higher executive pay.[2]

Table VII presents our results using the Hirsch data. The regressions are specified anal-

[2]The positive effect is entirely due to the changes from the 1975-1978 to the 1979-1982 period. However, we have not been able to discern what is different during that period from before.

ogously to the previous tables but the sample is a simple cross section. For comparability across data sets, and because there is large variability in CEO pay from year to year, we used the same four year aggregates of the variables for these regressions as well. Thus, all variables refer to the period 1975-1978, except for the unionization measure, which is supposed to represent unionization in 1977. However, recall that this is a retrospective survey measure, so that the respondent may well have reported something closer to an average for the years roughly ten years ago. When we run these regressions with data only for 1977 we find basically the same results.

A 10 percentage point increase in the unionization rate in the Hirsch data is associated with 1.5 percent lower executive pay. The effect is slightly larger when we limit the sample to firms with some unionization, the sample which is most comparable with the Bronars, Deere, and Tracy data. In fact, these coefficients are of the same magnitude as what we found in the Bronars, Deere, and Tracy data for manufacturing only. The effects are not very precisely estimated because this dataset is rather small.

A curious result in this dataset is that CEOs in firms with positive unionization rates up to about 60 percent actually get paid *more* than CEOs in firms without unions. However, firms where more than 60 percent of the workers are unionized have lower paid CEOs than any of the other groups. It is essentially this difference which shows up in the regressions with just a linear unionization rate. We are unable to discern whether this is actually a meaningful result or whether it is just due to the small sample size. Remember that the Bronars, Deere, and Tracy sample only contained firms with some unionized workers, so that we cannot do an analogous comparison of firms with and without unions.

Including covariates like growth rates and measures which capture rents does not change the union coefficient much in the Hirsch dataset. As before, including industry effects attenuates the union effect but the attenuation is also not as strong as in the Bronars, Deere, and Tracy data. Regressions including two and three digit industry dummies are displayed in Table VIII. In neither case is the coefficient on fraction unionized significantly different from zero. We obviously cannot include firm effects as in Table VI since this is a simple cross section.

One issue in all of these regressions is that the measures of unionization are rather imperfect, so that our coefficients are likely to be biased because of measurement error. Because the Hirsch measure and the Bronars, Deere, and Tracy data are constructed independently, it seems sensible to assume that the errors in these measures might be independent. Bronars, Deere, and Tracy analyze the correlation between the two measures with a subsample of the data with the same firms. Under the assumption of classical measurement error, they conclude that the relevant attenuation factors for both their data and the Hirsch data is about 0.55. Furthermore, the same applies when they isolate within industry variation in the data. This means that the coefficients should be roughly multiplied by 2 to eliminate the attenuation bias from measurement error. We found roughly similar changes in the coefficients using IV estimates on the overlapping sample, indicating that these conclusions are unchanged partialling out the effects of our covariates.

Attenuation due to measurement error is unlikely a good explanation of our findings when we control for industry or firm effects. Attenuation does not seem to be much greater within industries, while we find that the union coefficient is substantially closer to zero. We would expect that the signal to total variance ratio for the unionization rate declines more strongly in the within firm data, but this is extremely unlikely to lead to a positive coefficient with a t-statistic of 2.5. Thus, changes in unionization rates seem to have a very different relationship with CEO pay than the levels of unionization.

4 Representation Elections and CEO Pay

Since both the presence of unions and CEO pay may be determined by some variables like rents or other aspects of the well-being of the firm which are difficult to capture, (12 percent) are decertification elections. All but 3 of the decertification elections are initiated by employees covered by the bargaining agreement. 70 of the elections in our sample (25 percent) are won by the union. This is much lower than the aggreagate union win rate during the sample period (about 45 percent). Unions tend to be less successful ir larger establishments like the ones in our sample. The success rate in our sample is not unrepresentative for establishments with more than 250 eligible voters. The union is more likely to win decertification elections (the win rate in the sample is 57 percent) than certification elections (where the win rate is only 20 percent).

Our dependent variable is the change in the logarithm of salary, bonus and other compensation of the CEO between the year after the election and year before the election. Since the matched sample spans a period from 1975 to 1991, CEO pay is deflated by the consumer price index. The key independent variables are dummies for whether the union gains or looses members. The dummy for a membership gain will be one if the union wins a certification election, and the dummy for a membership loss will be one if the union looses a decertification election. We also experimented with a measure capturing the change in the number of workers unionized but that measure turned out to be very imprecise. CEO tenure and age refer to the year before the election, all other variables are changes between the period after the election and the year before.

Table IX presents regression results. The first column only includes the dummies for the change in union membership and a control for whether the CEO turned over either in the election year or the year after (which is used in calculating the change in CEO salary). Changes in unionization induced by elections seem to have little effect on CEO salaries. Unfortunately this effect is also estimated very imprecisely because there are only few elections which induce a change in union membership. Adding other covariates, like CEO tenure and age, stock returns and growth rates of the firm, and changes in the capital/labor ratio, Tobin's Q, or the investment rate does not change the estimates. This might indicate this could easily lead to the observed correlations between unionization and CEO pay. Even in changes, it is unlikely that the factors which move CEO pay and unionization are unrelated. Thus, it would be interesting to look at changes in unionization which are more likely to be random. Therefore, we analyze the changes in union membership induced by the outcomes of union representation elections conditional on the fact that an election took place. The election outcome will be random if the lobbying effects of unions and management are not themselves dependent on the same factors which lead a union or firm to seek an election in the first place. Bronars and Deere (1993) report that union wins and losses are uncorrelated with previous changes in sales, employment, or investment at the firm. Bronfenbrenner (1997), on the other hand, shows that higher company profitability is systematically associated with *lower* union win rates for a sample of certification elections in the late 1980s. However, she also concludes that there seems to be a large idiosyncratic component to election outcomes. Thus, election outcomes are probably not quite perfect for our purpose, but might give valuable additional insights into the effect of unions on CEO pay.

We constructed a sample of 284 elections in large U.S. firms, using the NLRB election files from 1977 to 1992, which we can match to Compustat and to CEO pay data. The data contain both certification elections and decertification elections (either initiated by workers or by management). First, we selected elections with 250 or more eligible voters from the election files. For these elections, we tried to match the company names, which appear on the election files, to the names as they appear in Compustat. In this process, we made sure to capture subsidiaries of larger conglomerates. This resulted in 1,191 matches. However, we can only match 487 of these to both the Compustat and the CEO pay data for the relevant years. Excluding those elections with missing values on some of the covariates we use, we are left with 284 observations.

Of the 284 sample cases, 249 elections (88 percent) are certification elections and 35 that the elections provide a better experiment than the cross section regressions in affecting the union status of workers in a fashion unrelated to changes in firm characterisitics.

In Table X we reran these regressions on the subsample of elections which are not associated with the turnover of the CEO. Other conditions might have changed at the firm as the CEO turns over, and changes in pay may be less meaningful when they are observed for different individuals. Table X shows that there is a larger effect of successful decertification elections on the pay of incumbent CEOs. Winning such an election boosts CEO pay by around 10 to 15 percent. This result could indicate that CEOs get rewarded when they manage to end the representation of a union for a bargaining union. This effect is still not significant in the small sample at hand.

5 Unionization and the Pay of Other Managers

CEOs, no doubt, are a very special group of employees. Therefore, it seems useful to also have an assessment of the influence of unions on other executives and managers. Unfortunately, no comparable data for other managers are available as there are for CEOs. This makes a firm level analysis of this question impossible. Instead, we follow the strategy of Neumark and Wachter (1995), and analyze the pay of managers at the level of industry and regional cells, which we construct from the Current Population Survey (CPS). Neumark and Wachter use industry level data from the CPS for the period 1973 to 1989 and find that a 10 point higher fraction of blue collar workers unionized in an industry is associated with 1.8 percent lower wages for managers and professionals.

We use the 1983 to 1993 Merged Outgoing Rotation Groups from the CPS to construct cells at the industry/region/year level. Our sample starts in 1983 because this is the first year where union status is available for the outgoing rotation groups. In order to have cells of sufficient size, we aggregated two digit industries into 26 broader industries, and states into 16 regions. A complete list of the industry and state cells used is given in the appendix.

While we are mostly interested in the effects of unions on managers, we also look at spillover effects from unionization on three other groups. To this end, we break occupations into five groups. The first is executives and managers (1980 3-digit SIC occupation codes 3-22). These are primarily occupations who supervise others. In principle, workers in this group cannot form unions under NLRB rules. The second is management related occupations (occupation codes 23-37), which includes business specialists like accountants, analysts, inspectors, etc. In practice, the line between these two groups is probably not quite as clear. The third group is professionals (occupation codes 43-199), which includes a wide range of professions like engineers, scientists, physicians, nurses, and other health professions, teachers, librarians, social scientists, social workers, lawyers, artists, etc. This group may be unionized itself, but professionals tend to be highly educated and often resemble managers more than they resemble other workers. While we do not want to lump them together with the group of managers, they seem like an interesting group to study as well. We refer to all remaining occupations (codes 203-889) as workers. We split this group into union workers, those who are covered by collective bargaining themselves, and non-union workers. Our samples are restricted to those working five or more hours, who are in the private sector and not self-employed, and with valid wage and occupation information in the CPS.

Table XI provides means of the characteristics of these five groups for the years 1983 and 1993. Managers, management related occupations, and professionals have comparable wages but professionals tend to have more education. Managers and union workers tend to be older than the other groups.

For each of these groups we report two measures of unionization. The fraction unionized in the occupation refers to the fraction of observations in this occupation who respond that they are covered by a union contract. Between 5 and 6 percent of managers and related occupations in 1983 report that they are covered by collective bargaining. This is surprising since we do not expect managers to be organizeed in unions. This result may be due to measurement error because either the union or the occupation question are answered or coded incorrectly. However, 5 percent seems a little too high compared to typical misclassification rates (Card 1996) so that some of this may reflect actual unionization of this group. Some managers may be in relatively low level positions, who do not directly supervise others, and who are therefore not subject to NLRB rules. When we run standard cross-sectional wage regressions, we find that unionized managers have roughly 8 percent higher wages than other managers, about half the union wage differential typically found for all workers. This also seems to indicate that at least some of this reflects actual unionization of the group. Not surprisingly, unionization is higher among professionals since this is the broadest of our upper income categories, but not as high as among other workers.

The penultimate row in each panel in Table XI reports the fraction of unionized *workers* (those who are neither managers, related, or professionals) in the industry/region cells to which a manager, professional, etc. belongs. If all occupations were equally distributed across these cells, this number would be constant across the different columns (except for sampling variation) and would reflect the unionization level of workers in the economy. But different occupations tend to be more concentrated in certain industries. For example, there are more managers in banking and other finance than there are in the construction industry. Since the unionization rates of workers in these industries differ, and since the reported means for managers, say, are weighted by the number of managers in a cell, the fraction of union workers reported for each occupation differs. Thus, the fraction of union workers refers to the exposure of the occupation to unionized workers in the particular cell. The most notable, but unsurprising result about this fraction is that union workers tend to be concentrated in industry/region cells which are more highly unionized. Exposure to unionized workers has declined for all occupations over time.

The final row in each panel reports the fraction of managers in the industry/region cell an occupation belongs to. This fraction does not differ much for different occupations. For each occupation there is about a 2 percentage point increase in the fraction of managers over the sample period.

In order to assess the effect of unions on the pay of managers, we estimated a regression of the form

$$\ln w^s_{ijt} = X^s_{ijt}\beta + \gamma u^s_{jt} + \mu_j + \eta^s + \delta_t + \varepsilon^s_{ijt}.$$

w^s_{ijt} is the wage of individual i in industry j and state s at time t. X^s_{ijt} is a set of covariates which includes years of schooling, potential experience and its square, and dummies for female, black, whether the individual lives in an SMSA, and part-time status (less than 35 hours a week). u^s_{jt} is the fraction of workers who are unionized in each state/industry/time cell. Notice that this variable does not vary at the individual level. In order to implement the estimation of this model and to obtain standard errors robust to a group structure in ε_{ijt} we first regressed $\ln w^s_{ijt}$ on X^s_{ijt} and a full set of industry/state/year interactions. In a second stage, we regressed the coefficients from these dummy variables on the fraction of workers unionized in the industry. The second stage regression is weighted by the number of observations in the cell. The results from the second stage regression are shown in Tables XII and XIII.

Column (1) in Table XII displays results which control for separate industry, region, and year effects. The results indicate a positive association between unionization of workers and the wages of professionals, but not managers or related occupations. In addition, unionization in the industry and region is associated with higher pay for both covered and uncovered workers. The differential between the two groups is about 16 percent, roughly corresponding to the union wage differential.

We know from the literature on industry wage differentials that unions tend to be more concentrated in high wage industries. Thus, if these industry wage differentials change over time, and this affects trends in unionization, we may be more likely picking up changes in the industry wage structure than the effects of unions on other occupations. Therefore, we also control for the interaction of industry and year effects in column (2). The results are now much more uniform. Higher unionization is associated with higher pay for all occupations now, except professionals, and these results are mostly very significant. The pay of managers is about 0.8 percent higher for any 10 point change in worker unionization. For non-union workers the effect is lower at around 1.8 percent, still twice as big as the effect on managers.

We also added a control for the fraction of the occupation covered by a union contract themselves because we were concerned that the results may simply reflect correlations in unionization rates in an industry across different groups. These results are displayed in column (3). They are hardly changed from the previous specification.

One hypothesis why unions affect the pay of other groups including managers is that different groups of workers split the rents accruing to a firm. For a given amount of rents to go around, we would expect that managers and other white collar employees get fewer of the rents as workers get more. Thus, apart from the fraction of workers unionized it is also the size of the union wage premium which matters. Columns (4) through (6) display similar specifications with the union wage differential in the industry/state/year cell included as a separate covariate. We obtained this regressor by running a regression on the pooled micro data for union and non-union workers including the vector X^{s}_{ijt} and a full set of industry/state/year/union interactions. The union wage differential is then given by the difference between the union and non-union effect for a particular industry, state, and year. The results on the fraction of workers unionized are again hardly changed. The union wage differential has a negative effect on the wages of all occupations. The effect is close to zero for managers and strongest for non-union workers.

Table XIII makes another attempt to address the issue that the fraction of workers unionized (or the union wage differential) could reflect the level of rents in an industry and state. One way to capture these rents may be to include the industry wage differential for non-union workers in the regression. This was again created from the same first stage regression which we ran to compute the cell level union wage differential. The industry wage differential has a large and strongly significant effect on the wages in each occupation. After controlling for industry*year interactions the elasticity of an occupation's wage with respect to the wage of non-union workers in the cell is about 0.45. At the same time, the impact of both the fraction of workers unionized and the union wage differential is diminished. For professionals the effect of the fraction of workers unionized now turns negative and the union wage differential positive. For management related occupations neither variable matters. For managers only the union wage differential seems to have a small positive effect. These results seem to indicate that the fraction of workers unionized in a cell is likely to capture some aspects of rents since including the state-industry wage effect matters for the results. Overall, there is relatively little evidence that unions have an effect on the pay of managers.

Our results differ from Neumark and Wachter (1995) which is similar in spirit to our study although they consider a different time period. Using industry level data, they find that higher unionization of workers is associated with lower pay for non-union workers and managers, controlling for main industry and year effects. On the other hand, when they use city level data, they find the opposite effect on non-union workers. In part, our results suggest that either it might be important to control for industry and time interactions, or to employ other strategies to capture the level of rents in an industry and its change over time. However, without these controls we generally find a positive association between the wages of different occupations and the fraction of workers unionized rather than a negative one. We feel that this positive association is likely to reflect industry rents.

In order to probe the influence of unions on other occupations further we also pooled the observations for all occupations in a cell, and regressed the fraction of managers in the cell on the fraction of unionized workers. This is the quantity analogue to the pay regressions above. Table XIV displays the results. We find consistently significant negative effects of both the level of unionization on the fraction of managers in the cell and of the union wage differential. If unions are ind d causing these differences, the pattern suggests that unions depress both the supply of managers to a cell and the demand. A supply side effect might be that there is a compensating differential for managing a unionized firm: it may be more difficult or less enjoyable to manage a unionized firm or establishment. A demand side effect might be that firms higher fewer managers when unions are present, because unions and managers are substitutes. This could be because the higher wages in unionized establishments prevent worker shirking at a lower level of supervision. However, it seems to be mostly the union wage differential that matters for managers' wages and the fraction of workers unionized that matters for quantities, which seems to suggest that such a supply and demand driven explanation is unlikely.

6 Discussion and Conclusion

In general, we find a negative cross-sectional association between CEO pay and unionization. This is both true across countries and across firms in the U.S. The effects in this case are relatively large, CEO pay is roughly 2 percent lower for each 10 point rise in unionization rates in U.S. firms. While these results may not be individually significant, they all point in the same direction. The one caveat to these results is that we found in the Hirsch sample that firms with low unionization rates actually pay their CEOs more than those without any unions. We have no check on this result in any of the other samples. Since the res ılts are not strongly significant, we are not sure whether we should put too much weight on this finding.

Our second basic finding is that changes in unionization are not associated with changes in CEO pay in the opposite direction. Again, we find this both in the international comparisons and in the comparisons across firms. In fact, in the Bronars, Deere, and Tracy data we find a relatively sizeable *positive* association between unions and CEO pay. CEO pay is roughly 2.5 percent higher for each 10 point rise in unionization rates. A similar result is echoed in the CPS data for other managers. Their pay seems to move in unison with the degree of unionization of their industry and region unless we introduce the industry wage differential as a control. If the industry wage effect really captures rents, this may indicate that the finding in the Bronars, Deere, and Tracy data may also reflect the role of changing rents over time.

We also did not find any evidence of a positive effect of unions on CEO pay in the unionization election data. Our strongest results there indicate that a loss of union members due to decertification elections is possibly associated with higher CEO pay, but the results are noisy. Since we regard election outcomes as relatively random changes in unionization, this corroborates the findings that any positive association between unionization and managerial pay may be due to rents.

While the various datasets we have analyzed seem to reveal relatively little evidence that unions depress the pay for managers, unions may have an impact on managers nevertheless. In the CPS we found a very consistent pattern that fewer managers are employed where unions are stronger. Thus, unions may redistribute rents towards workers not by lowering the pay of managers but by reducing their number and therefore their wage bill.

Appendix: States and Industries Used in the CPS Analysis

For the analysis in section 5 we have grouped states and two digit industries into larger aggregates based on size and proximity, in order to assure that more of the cells we are using actually contain any observations.

The 16 regional aggregates are:

1. Maine, New Hampshire, Vermont, Massachusetts, Rhode Island
2. Connecticut, New York
3. New Jersey, Pennsylvania
4. Michigan, Ohio
5. Indiana, Illinois
6. Wisconsin, Minnesota, Iowa
7. Missouri, North Dakota, South Dakota, Nebraska, Kansas
8. Delaware, Maryland, Virginia, DC
9. West Virginia, North Carolina, South Carolina, Kentucky, Tennessee
10. Georgia, Florida
11. Alabama, Mississippi, Arkansas, Louisiana
12. Oklahoma, Texas
13. Idaho, Wyoming, Montana, Washington, Oregon
14. Colorado, New Mexico, Utah, Arizona, Nevada
15. California
16. Alaska, Hawaii

The 26 industry aggregates are (the CPS Detailed Industry Recodes are in parentheses):

1. Agriculture service, other agriculture, forestry and fisheries (01–02, 46)
2. Mining (03)
3. Construction (04)
4. Lumber, wood, furniture, fixtures, stone, clay, glass, and concrete (05–07)
5. Primary and fabricated metals, metals industries not specified (08–10)

6. Machinery, including electric, professional and photographic equipment, watches (11–12, 16)
7. Auto vehicles, aircraft, parts, and other transportation equipment (13–15)
8. Toys, amusement, and sporting goods, and miscellaneous manufacturing (17–18)
9. Food, tobacco, and kindred products (19–20)
10. Textiles, apparel, and leather products (21–22, 28)
11. Paper, printing, publishing, and allied industries (23–24)
12. Chemicals, petroleum and coal products, rubber, miscellaneous plastic, and allied products (25–27)
13. Transportation (29)
14. Communications (30)
15. Utilities and sanitary services (31)
16. Wholesale trade (32)
17. Retail trade (33)
18. Banking and other finance (34)
19. Insurance and real estate (35)
20. Private household and personal services (36, 39)
21. Business and other professional services (37, 45)
22. Repair services (38)
23. Entertainment and recreation services (40)
24. Hospitals and health services (41–42)
25. Education services (43)
26. Social services (44)

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Table I:

Cross Country Panel Regressions
CEOs

	Dependent Variable:							
	Ln(Real Total Compensation)				Ln(PPP Total Compensation)			
Unionization Rate	-.530	-.293	-.229	.385	-1.019	-.729	-.451	.185
	(.114)	(.085)	(.092)	(1.052)	(.152)	(.106)	(.161)	(.808)
Time Trend	—	—	.101	—	—	—	.494	—
			(.013)				(.066)	
US Dummy	—	.819	.785	—	—	.580	.020	—
		(.070)	(.063)			(.124)	(.011)	
US Incremental Trend	—	-.045	-.044	—	—	.022	.054	—
		(.013)	(.013)			(.016)	(.023)	
Marginal Employee	—	—	-.147	—	—	—	-.602	—
Tax Rate			(.108)				(.206)	
Marginal Corporate	—	—	-.245	—	—	—	.220	—
Tax Rate			(.198)				(.310)	
Time Dummies	Yes	Yes	No	Yes	Yes	Yes	No	Yes
Country Dummies	No	No	No	Yes	No	No	No	Yes

Data sources: Data on CEO compensation are from Abowd and Bognanno (1995). Unionization rates are from Jelle Visser with assistance from Danny Blanchflower and Michael Wallerstein. Data refer to 12 countries and the periods 1984, 1988-1992. Number of observations is 81. All regressions also include dummies for the data source. Multiple observations for a particular year and country may be available from different sources. Standard errors in parentheses are adjusted for this grouping.

Table II:

Cross Country Panel Regressions
Human Resource Managers

	Dependent Variable: Ln(Real Total Compensation)				Ln(PPP Total Compensation)			
Unionization Rate	-.104	-.096	-.266	1.505	-.623	-.567	-.514	.477
	(.122)	(.126)	(.109)	(1.704)	(.074)	(.086)	(.119)	(.878)
Time Trend	—	—	.065	—	—	—	-.007	—
			(.055)				(.061)	
US Dummy	—	.360	.386	—	—	.058	.031	—
		(.088)	(.086)			(.045)	(.042)	
US Incremental Trend	—	-.060	-.055	—	—	.013	.015	—
		(.011)	(.009)			(.007)	(.008)	
Marginal Employee Tax Rate	—	—	.425	—	—	—	-.113	—
			(.198)				(.155)	
Marginal Corporate Tax Rate	—	—	.067	—	—	—	.258	—
			(.412)				(.340)	
Time Dummies	Yes	Yes	No	Yes	Yes	Yes	No	Yes
Country Dummies	No	No	No	Yes	No	No	No	Yes

Data sources: See Table I. Number of observations is 45.

Table III:

Cross Country Panel Regressions
Relative Compensation:

	Dependent Variable: Ln(CEO/Manufacturing Operative)				Ln(HR Manager/Manuf. Operative)			
Unionization Rate	-.840	-.686	-.662	-1.277	-.493	-.563	-.529	.011
	(.125)	(.110)	(.129)	(.627)	(.148)	(.148)	(.183)	(1.092)
Time Trend	—	—	.024	—	—	—	.016	—
			(.011)				(.013)	
US Dummy	—	.164	.161	—	—	-.353	-.348	—
		(.095)	(.099)			(.041)	(.046)	
US Incremental Trend	—	.042	.041	—	—	.035	.035	—
		(.015)	(.015)			(.010)	(.010)	
Marginal Employee Tax Rate	—	—	-.082	—	—	—	-.149	—
			(.215)				(.307)	
Time Dummies	Yes	Yes	No	Yes	Yes	Yes	No	Yes
Country Dummies	No	No	No	Yes	No	No	No	Yes

Data sources: See Table I. Observations only refer to the years 1984, 1988, and 1992. Number of observations is 45 for the CEO results and 33 for the HR manager results.

Table IV:

Firm Level Regressions
Dependent Variable:
Ln(Salary, Bonus and Other CEO Compensation)
Bronars, Deere, and Tracy Data

	Means	Coefficients			
Fraction Union	.319	-.303 (.054)	—	-.182 (.057)	—
Union 21-40%	.295	—	-.055 (.030)	—	-.041 (.029)
Union 41-60%	.194	—	-.143 (.034)	—	-.111 (.034)
Union 61-80%	.088	—	-.313 (.048)	—	-.231 (.050)
Union 81-100%	.036	—	-.144 (.069)	—	-.056 (.069)
CEO Age	57.6	.019 (.023)	.021 (.023)	.030 (.023)	.031 (.023)
CEO $Age^2/100$	33.5	-.015 (.021)	-.016 (.020)	-.024 (.020)	-.025 (.020)
CEO Tenure	7.66	.018 (.006)	.017 (.006)	.015 (.006)	.014 (.006)
CEO $Tenure^2/100$	.943	-.032 (.020)	-.032 (.020)	-.023 (.020)	-.023 (.020)
New CEO	.245	-.064 (.030)	-.073 (.030)	-.067 (.029)	-.074 (.029)
ln(Sales)	7.89	.148 (.026)	.152 (.026)	.145 (.039)	.140 (.039)
ln(Employment)	3.19	.176 (.025)	.166 (.025)	.162 (.041)	.163 (.040)
Stock Return	.127	.070 (.016)	.075 (.016)	.070 (.017)	.075 (.017)
Stock Return last period	.270	.017 (.009)	.019 (.009)	.011 (.010)	.013 (.010)
Sales Growth	.183	—	—	.105 (.090)	.113 (.089)
Employment Growth	.086	—	—	-.068 (.083)	-.085 (.083)
ln(Capital per Worker)	3.68	—	—	.017 (.025)	.023 (.025)
Tobin's Q	.962	—	—	-.177 (.043)	-.174 (.043)
Investment/Sales	.103	—	—	-.813 (.203)	-.813 (.201)
Period 1974-78	.332	.322 (.035)	.318 (.034)	.323 (.034)	.319 (.034)
Period 1979-82	.299	.738 (.033)	.735 (.032)	.771 (.035)	.768 (.035)

Data sources: Unionization rates are from Bronars, Deere, and Tracy (1994), CEO pay data are from Kevin Murphy, all other variables are from Compustat. Data refer to four year aggregates for 1971-74, 1975-78, and 1979-82. Standard errors are in parentheses. Number of Observations is 782.

Table V:

Firm Level Regressions
Dependent Variable:
Ln(Salary, Bonus and Other CEO Compensation)
Bronars, Deere, and Tracy Data

	Manufacturing		Non-Manufacturing	
Fraction Union	-.124	—	-.099	—
	(.063)		(.104)	
Union 21-40%	—	-.039	—	-.042
		(.029)		(.072)
Union 41-60%	—	-.122	—	-.057
		(.034)		(.076)
Union 61-80%	—	-.076	—	-.172
		(.065)		(.087)
Union 81-100%	—	-.019	—	.001
		(.083)		(.117)
CEO Age	.040	.040	-.018	-.018
	(.022)	(.022)	(.064)	(.064)
CEO $Age^2/100$	-.034	-.035	.024	.024
	(.020)	(.019)	(.056)	(.056)
CEO Tenure	.019	.019	.020	.020
	(.007)	(.007)	(.010)	(.010)
CEO $Tenure^2/100$	-.060	-.061	-.003	-.005
	(.027)	(.027)	(.032)	(.032)
New CEO	-.052	-.056	-.087	-.097
	(.030)	(.030)	(.062)	(.062)
ln(Sales)	.203	.204	.091	.085
	(.048)	(.048)	(.080)	(.080)
ln(Employment)	.067	.066	.221	.227
	(.050)	(.050)	(.076)	(.077)
Stock Return	.067	.067	.101	.104
	(.017)	(.017)	(.038)	(.038)
Stock Return	.010	.011	-.002	-.001
last period	(.012)	(.012)	(.016)	(.016)
Sales Growth	.001	.010	.255	.251
	(.100)	(.100)	(.160)	(.160)
Employment Growth	.082	.076	-.229	-.235
	(.091)	(.091)	(.156)	(.157)
ln(Capital per Worker)	.007	.006	.003	.015
	(.037)	(.037)	(.037)	(.038)
Tobin's Q	-.127	-.122	.064	.072
	(.043)	(.043)	(.125)	(.125)
Investment/Sales	.699	.656	-.443	-.468
	(.452)	(.450)	(.269)	(.268)
Period 1974-78	.364	.363	.174	.169
	(.035)	(.035)	(.077)	(.077)
Period 1979-81	.798	.798	.645	.641
	(.036)	(.036)	(.078)	(.078)
Number of Observations	577	577	205	205

Data sources: Unionization rates are from Bronars, Deere, and Tracy (1994), CEO pay data are from Kevin Murphy, all other variables are from Compustat. Data refer to four year aggregates for 1971-74, 1975-78, and 1979-82. Standard errors are in parentheses.

Table VI:

Firm Level Regressions
Dependent Variable:
Ln(Salary, Bonus and Other CEO Compensation)
Bronars, Deere, and Tracy Data
Industry and Firm Effects

Fraction Union	-.107	.011	.255
	(.056)	(.058)	(.105)
CEO Age	-.008	-.012	-.001
	(.023)	(.024)	(.030)
CEO $Age^2/100$	.009	.011	.006
	(.020)	(.021)	(.027)
CEO Tenure	.016	.021	.019
	(.005)	(.005)	(.006)
CEO $Tenure^2/100$	-.035	-.052	-.039
	(.019)	(.018)	(.019)
New CEO	-.060	-.064	-.058
	(.026)	(.024)	(.020)
ln(Sales)	.253	.132	.013
	(.052)	(.058)	(.096)
ln(Employment)	.027	.149	.291
	(.051)	(.058)	(.092)
Stock Return	.067	.068	.054
	(.015)	(.014)	(.014)
Stock Return	.014	.019	.000
last period	(.009)	(.009)	(.012)
Sales Growth	-.021	.012	.062
	(.087)	(.086)	(.081)
Employment Growth	.041	.046	.015
	(.080)	(.080)	(.078)
ln(Capital per Worker)	-.014	.100	-.031
	(.033)	(.038)	(.069)
Tobin's Q	-.076	-.033	-.233
	(.044)	(.047)	(.063)
Investment/Sales	.043	.223	.593
	(.202)	(.210)	(.218)
Period 1974-78	.323	.341	.370
	(.031)	(.029)	(.028)
Period 1979-81	.752	.768	.828
	(.032)	(.031)	(.037)
Industry Effects	44	103	No
Firm Effects	No	No	327

Data sources: Unionization rates are from Bronars, Deere, and Tracy (1994), CEO pay data are from Kevin Murphy, all other variables are from Compustat. Data refer to four year aggregates for 1971-74, 1975-78, and 1979-82. Standard errors are in parentheses. Number of Observations is 782.

Table VII:

Firm Level Regressions
Dependent Variable:
Ln(Salary, Bonus and Other CEO Compensation)
Hirsch Data

	Means	Full Sample				Firms With Any Union
	Means	Coefficients				
Fraction Union	.367	-.145	—	-.179	—	-.209
		(.078)		(.087)		(.090)
Union 1-20%	.181	—	-.001	—	.039	—
			(.071)		(.072)	
Union 21-40%	.189	—	.021	—	.042	—
			(.073)		(.075)	
Union 41-60%	.256	—	.044	—	.059	—
			(.069)		(.073)	
Union 61-80%	.211	—	-.131	—	-.109	—
			(.070)		(.076)	
Union 81-100%	.026	—	-.044	—	-.084	—
			(.133)		(.140)	
CEO Age	57.0	.028	.016	.023	.014	-.069
		(.054)	(.054)	(.055)	(.055)	(.060)
CEO $Age^2/100$	32.7	-.014	-.003	-.010	-.002	.065
		(.049)	(.049)	(.049)	(.049)	(.052)
CEO Tenure	7.54	-.000	.001	.000	.002	.018
		(.010)	(.010)	(.010)	(.010)	(.011)
CEO $Tenure^2/100$	.92	-.016	-.017	-.019	-.021	-.070
		(.031)	(.031)	(.031)	(.031)	(.035)
New CEO	.198	-.123	-.117	-.113	-.107	-.124
		(.052)	(.052)	(.051)	(.052)	(.049)
ln(Sales)	7.53	.172	.158	.077	.055	.064
		(.040)	(.041)	(.075)	(.076)	(.071)
ln(Employment)	3.13	.155	.168	.249	.270	.292
		(.040)	(.041)	(.075)	(.076)	(.072)
Stock Return	.048	.132	.122	.179	.163	.154
		(.035)	(.035)	(.038)	(.038)	(.036)
Stock Return	.070	.048	.043	.042	.035	.054
last period		(.027)	(.028)	(.031)	(.032)	(.030)
Sales Growth	.252	—	—	-.220	-.159	-.127
				(.191)	(.193)	(.185)
Employment Growth	.123	—	—	.111	.090	.216
				(.164)	(.165)	(.159)
ln(Capital per Worker)	3.28	—	—	.119	.116	.084
				(.066)	(.065)	(.063)
Tobin's Q	.930	—	—	-.118	-.131	-.092
				(.068)	(.069)	(.067)
Investment/Sales	.069	—	—	-.089	-.119	.402
				(.798)	(.801)	(.766)
Number of Observations	227	227	227	227	227	196

Data sources: Unionization rates are from Hirsch (1991), CEO pay data are from Kevin Murphy, all other variables are from Compustat. All variables except unionization rates are aggregates for the four years 1975-1978. Standard errors are in parentheses. Unionization rates refer to 1977.

Table VIII:

Firm Level Regressions
Dependent Variable:
Ln(Salary, Bonus and Other CEO Compensation)
Hirsch Data
Industry Effects

Fraction Union	-.100	-.118
	(.101)	(.132)
CEO Age	.020	.002
	(.061)	(.073)
CEO $Age^2/100$	-.008	.007
	(.054)	(.064)
CEO Tenure	-.004	-.004
	(.011)	(.014)
CEO $Tenure^2/100$	.004	-.005
	(.035)	(.042)
New CEO	-.097	-.071
	(.053)	(.065)
ln(Sales)	-.070	-.131
	(.099)	(.153)
ln(Employment)	.390	.439
	(.096)	(.149)
Stock Return	.173	.201
	(.041)	(.056)
Stock Return	.035	.026
last period	(.033)	(.039)
Sales Growth	-.103	-.067
	(.220)	(.274)
Employment Growth	.063	.004
	(.181)	(.220)
ln(Capital per Worker)	.116	.106
	(.083)	(.118)
Tobin's Q	-.074	-.086
	(.079)	(.096)
Investment/Sales	.428	.028
	(.906)	(1.117)
Industry Effects	30	85

Data sources: Unionization rates are from Hirsch (1991), CEO pay data are from Kevin Murphy, all other variables are from Compustat. All variables except unionization rates are aggregates for the four years 1975-1978. Unionization rates refer to 1977. Standard errors are in parentheses. Number of observations is 227.

Table IX:

Regressions for Union Election Data
Dependent Variable:
Change in Ln(Salary, Bonus and Other CEO Compensation)

Union Gains Members	-.032	-.033	-.038	-.050
	(.060)	(.060)	(.057)	(.058)
Union Loses Members	.036	.037	.021	.038
	(.101)	(.103)	(.098)	(.098)
CEO Tenure	—	.000	-.002	-.001
		(.004)	(.004)	(.004)
CEO Age	—	-.002	.001	.001
		(.004)	(.004)	(.004)
New CEO	-.140	-.138	-.126	-.123
	(.054)	(.057)	(.054)	(.054)
Stock Return	—	—	.105	.076
			(.051)	(.083)
Market Return	—	—	-.279	-.278
			(.125)	(.125)
Sales Growth	—	—	.695	.806
			(.169)	(.187)
Employment Growth	—	—	-.455	-.577
			(.162)	(.189)
Change in ln(Capital per Worker)	—	—	—	-.277
				(.127)
Change in Tobin's Q	—	—	—	-.115
				(.233)
Change in Investment/ Sales	—	—	—	.461
				(.518)

Sources: Election outcomes are from the NLRB Election Files. CEO pay data are from Kevin Murphy, all other data are from Compustat. Dependent variable is the change in CEO pay from the year before the election to the year after. Independent variables refer to the year of the election; growth rates refer to the change from the year before to the year after the election. Standard errors are in parentheses. Number of observations is 284.

Table X:

Regressions for Union Election Data
Dependent Variable:
Change in Ln(Salary, Bonus and Other CEO Compensation)
Sample Restricted to Elections Without CEO Turnover

Union Gains Members	.002	.002	-.000	-.003
	(.070)	(.070)	(.065)	(.066)
Union Loses Members	.151	.140	.107	.134
	(.113)	(.115)	(.107)	(.107)
CEO Tenure	—	.003	.004	.004
		(.005)	(.004)	(.004)
CEO Age	—	-.002	.001	.000
		(.005)	(.005)	(.005)
Stock Return	—	—	.038	.042
			(.053)	(.083)
Market Return	—	—	-.301	-.272
			(.134)	(.135)
Sales Growth	—	—	.885	.940
			(.176)	(.197)
Employment Growth	—	—	-.610	-.711
			(.170)	(.200)
Change in ln(Capital per Worker)	—	—	—	-.306
				(.133)
Change in Tobin's Q	—	—	—	.004
				(.242)
Change in Investment/ Sales	—	—	—	.111
				(.673)

Sources: See Table IX. Standard errors are in parentheses. Number of observations is 220.

Table XI:

Industry/State Level Data from the CPS
Weighted Cell Means

Occupation	Managers	Management Related	Professionals	Non-Union Workers	Union Workers
			1983		
Ln Hourly Wage	2.32	2.27	2.25	1.71	2.15
Years of School	14.4	14.8	15.6	12.3	11.8
Potential Experience	20.8	17.7	16.8	17.2	22.3
Female	.324	.438	.466	.509	.273
Black	.032	.042	.041	.080	.119
In SMSA	.656	.712	.681	.559	.631
Part-time	.047	.048	.178	.258	.086
Unionized	.059	.056	.112	.000	1.000
Fraction Union Workers	.188	.171	.176	.160	.368
Fraction Managers	.081	.081	.063	.064	.061
			1993		
Ln Hourly Wage	2.67	2.63	2.71	2.10	2.47
Years of School	14.4	14.6	15.7	12.4	12.2
Potential Experience	20.8	18.3	18.1	18.5	23.4
Female	.414	.576	.521	.501	.286
Black	.043	.058	.053	.090	.132
In SMSA	.819	.862	.837	.722	.774
Part-time	.054	.065	.172	.243	.086
Unionized	.038	.035	.087	.000	1.000
Fraction Union Workers	.130	.107	.122	.115	.272
Fraction Managers	.097	.107	.093	.078	.078

Data source: Current Population Survey Merged Outgoing Rotation Groups from 1983 and 1993. Means of state/industry cells. Means are weighted by the number of observations in the cells.

Table XII:

Regressions on Industry/State/Year Level Data from the CPS
Dependent Variable: Ln(Hourly Wage Coefficient)

Independent Variable	(1)	(2)	(3)	(4)	(5)	(6)
	Managers					
Fraction of Workers Unionized	.021	.080	.076	.041	.082	.077
	(.024)	(.022)	(.022)	(.024)	(.022)	(.023)
Union Wage Differential	—	—	—	-.092	-.012	-.013
				(.013)	(.013)	(.013)
Fraction Managers Unionized	—	—	.040	—	—	.041
			(.033)			(.033)
	Management Related Occupations					
Fraction of Workers Unionized	-.043	.084	.084	-.022	.087	.087
	(.030)	(.030)	(.030)	(.030)	(.030)	(.030)
Union Wage Differential	—	—	—	-.098	-.034	-.034
				(.015)	(.015)	(.015)
Fraction in the Occupation	—	—	-.001	—	—	.000
Unionized			(.032)			(.027)
	Professionals					
Fraction of Workers Unionized	.112	.010	-.012	.125	.014	-.010
	(.027)	(.023)	(.023)	(.027)	(.023)	(.023)
Union Wage Differential	—	—	—	-.219	-.040	-.044
				(.016)	(.014)	(.014)
Fraction Professionals Unionized	—	—	.072	—	—	.078
			(.022)			(.022)
	Non-union Workers					
Fraction of Workers Unionized	.173	.183	—	.211	.195	—
	(.019)	(.013)		(.019)	(.013)	
Union Wage Differential	—	—	—	- .169	- .107	—
				(.011)	(.007)	
	Union Workers					
Fraction of Workers Unionized	.330	.299	—	—	—	—
	(.016)	(.015)				
Industry, State, Year Effects	Yes	Yes	Yes	Yes	Yes	Yes
Industry*Year Effects	No	Yes	Yes	No	Yes	Yes

Data source: Current Population Survey Merged Outgoing Rotation Groups from 1983 and 1993. The number of observations ranges from 4,123 for management related occupations to 4,502 for non-union workers. The dependent variable is the state/industry/year coefficient of a regression of the ln hourly wage on years of schooling, potential experience and its square, dummies for female, black, in an SMSA, and part-time, and a full set of state/industry/year interactions. The first-stage regression is run separately for each occupation. The union wage differential is obtained from a similar regression for union and non-union workers which also included a full set of state/industry/year/union interactions. It is obtained by subtracting the union coefficient from the non-union coefficient for the cell. Regressions are weighted by the number of observations in the cells. Standard errors are in parentheses.

Table XIII:

Regressions on Industry/State/Year Level Data from the CPS
Dependent Variable: Ln(Hourly Wage Coefficient)

Independent Variable	(1)	(2)	(3)	(4)	(5)	(6)
		Managers				
Fraction of Workers Unionized	-.051	.016	.008	-.054	.007	.000
	(.021)	(.022)	(.022)	(.022)	(.022)	(.022)
Union Wage Differential	—	—	—	.010	.040	.039
				(.013)	(.013)	(.013)
Fraction Managers Unionized	—	—	.063	—	—	.062
			(.032)			(.032)
Industry Wage Differential	.543	.433	.435	.547	.454	.456
	(.017)	(.026)	(.026)	(.018)	(.027)	(.027)
		Management Related Occupations				
Fraction of Workers Unionized	-.072	.023	.022	-.073	.021	.020
	(.028)	(.029)	(.030)	(.028)	(.030)	(.030)
Union Wage Differential	—	—	—	.006	-.006	.006
				(.014)	(.014)	(.014)
Fraction in the Occupation	—	—	.005	—	—	.005
Unionized			(.026)			(.026)
Industry Wage Differential	.601	.444	.444	.604	.447	.447
	(.023)	(.034)	(.034)	(.024)	(.035)	(.035)
		Professionals				
Fraction of Workers Unionized	-.044	-.053	-.077	-.054	-.062	-.083
	(.023)	(.022)	(.023)	(.023)	(.022)	(.023)
Union Wage Differential	—	—	—	.055	.051	.047
				(.015)	(.015)	(.015)
Fraction Professionals Unionized	—	—	.079	—	—	.073
			(.022)			(.022)
Industry Wage Differential	.763	.480	.482	.796	.514	.513
	(.018)	(.027)	(.027)	(.020)	(.028)	(.028)
		Union Workers				
Fraction of Workers Unionized	.255	.237	—	—	—	—
	(.014)	(.014)				
Industry Wage Differential	.547	.413	—	—	—	—
	(.017)	(.019)	—	—	—	—
Industry, State, Year Effects	Yes	Yes	Yes	Yes	Yes	Yes
Industry*Year Effects	No	Yes	Yes	No	Yes	Yes

Data source: Current Population Survey Merged Outgoing Rotation Groups from 1983 and 1993. The number of observations ranges from 4,123 for management related occupations to 4,502 for non-union workers. The dependent variable is the state/industry/year coefficient of a regression of the ln hourly wage on years of schooling, potential experience and its square, dummies for female, black, in an SMSA, and part-time, and a full set of state/industry/year interactions. The first-stage regression is run separately for each occupation. The union wage differential is obtained from a similar regression for union and non-union workers which also included a full set of state/industry/year/union interactions. It is obtained by subtracting the union coefficient from the non-union coefficient for the cell. The industry (state/year) wage differential is the the set of coefficients for non-union workers from the same regression. Regressions are weighted by the number of observations in the cells. Standard errors are in parentheses.

Table XIV:

Regressions on Industry/State/Year Level Data from the CPS

Independent Variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Dependent Variable: Fraction Managers								
Fraction of Workers Unionized	-.082	-.035	-.074	-.033	-.095	-.038	-.088	-.036
	(.006)	(.004)	(.006)	(.004)	(.006)	(.004)	(.006)	(.004)
Union Wage Differential	—	—	-.038	-.011	—	—	-.024	-.009
			(.004)	(.003)			(.004)	(.003)
Industry Wage Differential	—	—	—	—	.007	.018	.068	.013
					(.005)	(.005)	(.005)	(.006)
Dependent Variable: Fraction Managers and Related Occupations								
Fraction of Workers Unionized	-.128	-.043	-.111	-.041	-.163	-.051	-.152	-.049
	(.010)	(.005)	(.010)	(.005)	(.009)	(.005)	(.009)	(.005)
Union Wage Differential	—	—	-.083	-.017	—	—	-.042	-.012
			(.006)	(.003)			(.006)	(.003)
Industry Wage Differential	—	—	—	—	.216	.045	.201	.039
					(.008)	(.006)	(.008)	(.007)
Industry, State, Year Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry*Year Effects	No	Yes	No	Yes	No	Yes	No	Yes

Data source: Current Population Survey Merged Outgoing Rotation Groups from 1983 and 1993. The number of observations is 4,502. The dependent variable is the fraction of managers or managers and related occupations in a state/industry/year cell. The union wage differential is obtained from a similar regression for union and non-union workers which also included a full set of state/industry/year/union interactions. It is obtained by subtracting the union coefficient from the non-union coefficient for the cell. The industry (state/year) wage differential is the the set of coefficients for non-union workers from the same regression. Regressions are weighted by the number of observations in the cells. Standard errors are in parentheses.