Managerial Incentives in the Presence of Envious Workers

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Abstract

In large corporate hierarchies, senior management is typically motivated through a combination of promotions and incentives directly related to the market value of the firm. Using a large panel of executives in publicly listed US firms, we provide new evidence that promotion-based and firm-value-based incentives are partially traded-off in the presence of unions. In particular, we find that senior managers in unionized firms are offered significantly lower levels of stock option compensation and a more compressed cash pay structure, which is offset by a larger promotion prize to the CEO position while keeping constant the probability of insider succession. These findings confirm theoretical predictions for incentive adjustments in firms where fairness considerations are present and offer an explanation for how unionized firms can remain competitive in the labor market for managers.

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"*Boards have been placed under enormous pressure by the left-wing, anti-business press and the envious leaders of unions and other so-called 'CEO Comp Watchers', and therefore Boards are being forced to protect themselves irrespective of the potential negative long-term impact on public companies.*" [1]

Angelo Mozilo, Countrywide's former CEO & Chairman

1. INTRODUCTION

The degree of pay inequality within a firm is an important determinant of organizational performance. As Lazear (1989) illustrates in his model of industrial politics, pay equality can be desirable on efficiency grounds when workers worry about reference groups and cooperation is necessary in production. More recently, the growing literature on incentives and fairness suggests that worker envy could alter the optimal incentive structure within a firm, especially when a worker's choice of reference for compensation comparison resides within his/her own organization rather than across firms (Bartling and Siemens, 2006; Dur and Glazer, 2008; Englmaier and Wambach, 2005; Itoh, 2004).

In the labor economics literature, trade unions are often characterized as a grouping of envious workers with strong preferences for raising employee pay levels, resisting high-powered incentive schemes, and reducing overall pay dispersion.[2] Lack of pay fairness (actual or perceived) within unionized environments can therefore produce adverse effects, such as the production of defective goods (Krueger and Mas, 2004; Mas, 2008), reduction in job performance (Mas, 2006) and interference with strategic plans (Kole and Lehn, 2000). Could the presence of unions and associated fairness concerns also extend throughout an organization's corporate structure, including to managers in the upper echelon of the firm?

While the effect of unions on workers' wage structure has been well documented (Card et al., 2003; Freeman, 1980) and anecdotal accounts of union opposition to excessive executive compensation abound, there is a lack of systematic empirical work documenting union influence on managerial incentives at the firm level. In this paper, however, we provide such evidence by examining the empirical connection between union presence and

[1] Extract from an email to John England, a Towers Perrin consultant, in October 2006. The email was released by the U.S. House of Representatives' Committee on Oversight and Government Reform in the context of an examination of Mr. Mozilo's compensation given that his company was implicated in the mortgage and banking crisis.

[2] For instance, see Bewley (2004), Burks et al. (2007), Carruth and Oswald (1987), Fehr and Kirchsteiger (1994), Fehr et al. (2007), and Skott (2005).

(non-CEO) senior manager incentives in publicly listed US firms. Based on the theoretical predictions of incentive models in the presence of envious agents in firms (Bartling and Siemens, 2006; Dur and Glazer, 2008), as well as Jensen and Murphy's (1990) well noted assertion that unions can act as implicit regulators of executive compensation, we predict that managers receive less firm-value-based incentives and more compressed cash pay in firms with union presence than in non-unionized firms. This, in turn, implies that unionized firms need to offer more promotion-based incentives in order to maintain managerial effort and to compete in the market for managerial talent (Core et al., 2003; Gibbs, 1995; Main, O'Reilly and Wade, 1993). Otherwise, in equilibrium, lowering of high-powered incentives, with no offsetting benefits, would result in unionized firms attracting only poor quality managers and, ultimately, failing in the product market.

This paper addresses the aforementioned conjectures and provides important new insights. Consistent with the expected tradeoff between firm-value-based and promotion-based incentives, we find that unionized firms offer greater promotion-based incentives; in the sense of combining a substantially larger promotional prize to the CEO position while keeping the probability of insider succession constant (if not slightly higher). We also illustrate that unionized firms substantially reduce the variable part of managerial compensation (i.e., stock options) but offer a more compressed cash pay distribution. Furthermore, we find that the elasticity of managerial compensation with respect to firm performance is not significantly different between unionized and non-unionized firms; implying that union presence, on average, does not appear to inhibit managerial incentives. This is a somewhat surprising finding but one that is consistent with the offsetting changes to incentive structures of managers in unionized versus non-unionized firms.

Overall, this study is part of a growing research trend that examines the linkages between labor and finance at the firm level.[3] In this sense, it contributes in several important ways to the empirical literature on incentives. *First*, our results provide an answer to the puzzle of how unionized firms can attract and motivate managerial talent, given that unions channel strong membership preferences for pay fairness and appear to operate as implicit regulators of executive remuneration. Indirectly, these results illustrate the importance of promotion-based incentives in managerial compensation, especially in an era of heightened stock-market volatility, which is diluting the power of stock-based

[3] For instance, see Atanassov and Kim (2007), Chen et al. (2007), Klasa et al. (2009), Pagano and Volpin (2008), and Philippon and Reshef (2008).

incentives. *Second*, the paper empirically confirms the noted union aversion to stock-option incentive systems and indicates that unions extend their influence throughout the corporate hierarchy, beyond those workers covered by union-negotiated wage scales. This is consistent with union preferences for seniority-based pay and compensation compression within and across positions (Hirsch, 2008). *Third*, we provide evidence that despite overall union membership declines during the last several decades, union presence appears to be effective in its objectives.[4] This could be attributed to unions' increased use of proxy issue proposals as a result of the 1992 SEC Proxy Reforms (Choi, 2000), and the increased pressure through union-controlled pension funds, which actively urge corporate governance reforms (Gillan, and Starks, 2000). *Fourth*, our paper appears to offer an alternative causal direction to that of Cronqvist et al. (2009), who used Swedish data to argue that both entrenched managers and unionized workers are extracting higher remuneration in poorly governed firms. In particular, they find that workers are paid more when CEOs are entrenched and less when CEOs have financial incentives through cash flow rights.[5] The higher worker wage creates private benefits for the CEO in the form of improved social relations with employees and reduced conflict during collective bargaining. Unlike Cronqvist et al. (2009), in the context of managerial entrenchment and rent extraction, our paper implies that unions and managers are competing for rent extraction, rather than tacitly collaborating.

The rest of the paper is organized as follows: Section 2 presents the empirical methodology. Section 3 describes the data and sample design construction, while Sections 4 presents the results. Section 5 concludes.

2. Methodology

2.1. Empirical strategy

Our empirical approach rests on the premise that unions offer an independent institutionalized voice for workers' interests, such as labor concerns for pay fairness and subsequent demands for a more compressed compensation structure in the firm. Trade unions can enforce these preferences by threatening industrial peace at the workplace,

[4] In a similar fashion, Hirsch (2008) demonstrates that the union wage premium for workers has little fluctuation during the 1970-2006 period

[5] Cronqvist et al. (2009) focus on union aggressiveness, rather than union presence, since 70% of the Swedish private sector workforce is unionized. In contrast, roughly 10% of the US private sector employees are unionized.

applying pressure through worker-controlled pension funds, influencing committee decisions through employee-nominated representatives on the boards of directors, and by using high executive compensation as a justification for increased wage demands in labor negotiations (Jensen and Murphy, 1990). Moreover, from an information perspective, unions conduct and publicize salary surveys of workers and managers, all while going to great lengths to keep themselves informed about firms' financial conditions (Bewley, 1999, 2003). These aspects of union activity might constrain the efficient design of senior management incentives and thereby misalign executive compensation with shareholder interests.

The empirical approach used to test our hypotheses about union presence and senior management incentives proceeds in three stages as follows. First, we examine the alignment of managerial incentives to shareholder interests in unionized and non-unionized firms by estimating the pay-for-performance relation. To the extent that incentives are misaligned in the presence of worker envy, unionized firms should display lower pay-performance elasticity. However, consistent with Rosen's (1986) suggestion that a promotion essentially has an option value, it is just as likely that unionized firms adjust by changing their optimal compensation strategies in response to the constraints placed on variable-based incentives (e.g., stock options). As a result, we test for the presence of stronger promotion-based incentives within unionized firms, in the form of a higher likelihood of an insider becoming CEO, and a higher compensation gap between the CEO and top managers. Finally, in an effort to establish compression effects in the presence of envious workers, we examine compensation structures across matched samples of firms in order to establish whether unionized firms display a more compressed distribution of managerial compensation than those in the non-union sector.

2.2 Testing for potential endogeneity

In the empirical literature of worker wage differentials between unionized and non-unionized workers, one has to differentiate between the pure union effect and the influence of other observable and unobservable factors correlated with union presence, such as increased tenure and worker skills (Card et al., 2003; Freeman, 1980).

In the context of managerial compensation, however, this endogeneity problem is severely ameliorated since the distribution of managerial skills between unionized and non-unionized firms is expected to be similar. More specifically, firm size and industry classification are widely used during compensation-setting as a proxy for managerial skill requirements, while personal characteristics such as age, experience and education are sidelined (Murphy, 1999).[6] At the same time, union presence is unlikely to be correlated with individual characteristics of senior managers because, unlike lower-level workers, senior managers are not union members nor are covered by collective agreements. Moreover, the top four (non-CEO) managers, which are the focus of this study, hold similar positions across firms (e.g., President, Vice-President, Chief Financial Officer, Chief Operating Officer, Treasurer) thus revealing comparable duties. Also, as a matter of fact, in our sample, there is evidence of managerial mobility between the two types of firms (i.e., with and without union presence) in both directions.[7] Lastly, unlike the private-public sector choice whose endogeneity has been well-established in the literature (Gregory and Borland, 1999), all firms in the sample belong to the private sector thus being uniformly exposed to product market competition. Based on the aforementioned, in contrast to lower-level employees, managers are unlikely to self-select into a publicly listed firm based on the criterion of union presence.

Nevertheless, it could be that unions are more likely to organize firms with substantial rents and that this in turn could also affect the choice of senior manager incentives. For instance, recent work by Klasa et al., (2009) suggests that unionized firms maintain lower cash holdings in order to shelter corporate income from organized labor demands. Thus, for robustness purposes, we account for possible endogeneity of union presence by utilizing firm-specific instrumental variables concerning the *location of company headquarters* and the *age of the firm*.

The location of company headquarters has been found to influence union presence given the variation in state-level labor regulation (e.g., right-to-work laws) and union approval (Herod, 1998; Holmes, 2006). To capture these effects we utilize three dummy variables: BIBLE-BELT and SUN-BELT in order to identify companies that are headquartered in two areas traditionally seen as having an anti-union bias, and RUST-BELT in order to identify companies that are headquartered in an area traditionally friendly to unions (Lopez,

[6] Unlike blue-collar and lower-level professional jobs, for which unionized firms attract more skilled persons, general human capital variables (e.g., schooling years and standardized test scores) have little effect on managers, whose labor skill is typically assessed from aggregate firm performance rather than human capital requirements and piece-rate output.

[7] There are 148 executives in our sample who are observed to have moved from a non-union company to a union-company, or vice-versa. Although this number may appear small (about 1% of the total number of executives in this study), one has to bear in mind that it underestimates actual managerial mobility between companies due to the nature of available data in Execucomp. Notably, Execucomp typically has information for the top-five paid executives, meaning that a substantial part of managerial mobility is unobserved since managers that were in the top-five tier in one company, need not be in the top-five tier in the next company (and vice-versa).

2004).[8] We focus on headquarter location because a firm's headquarters are typically in close proximity to its establishments for transportation/communication considerations, and because headquarter location could be indicative of the firm's organizational culture, especially towards trade unions. For instance, during our sample period (1992-2001), the private-sector union density in the bible-belt, sun-belt and rust-belt was 5.1%, 10.6% and 14.6%, respectively (Hirsch and Macpherson, 2003).[9]

Firm age is important since firm-wide union decertification events are rare (Nilsson, 1997), meaning that firms established during the more union-friendly period before the 1970s are more likely to be observed as having union presence during our sample period (Fiorito, 2007; Freeman, 1988; Palley and LaJeunesse, 2007).[10] In calculating firm age, we use as the date of origin the date that each firm appears either in the Compustat dataset or the CRSP files; when there is substantial difference between Compustat and CRPS, we choose the earlier date (Pastor and Veronesi, 2003).

3. Sample design and Descriptive Statistics

Data on executive compensation and firm characteristics are obtained from the Standard and Poors (S&P) Executive Compensation database (*Execucomp*). *Execucomp* delivers data on executive compensation provided by the various US Securities and Exchange Commission (SEC) filings, as a result of an SEC regulation change in 1992 for firms in the S&P 1500 index and other supplemental S&P indices. Typically, it contains official compensation data for up to five of the highest paid executives in each firm.

[8] In particular, we construct three dummy variables for firm headquarters being located in a state that is either in the bible-belt, the sun-belt, or the rust-belt. Firm headquarters located in the remaining states serve as the omitted group. BIBLE-BELT is a dummy variable taking value 1 if the firm has its headquarters in a bible-belt state (namely AL, AR, FL, GA, KS, KY, LA, MO, MS, NC, OK, SC, TN, TX and VA), and 0 otherwise. SUN-BELT is a dummy variable taking value 1 if the firm has its headquarters in a sun-belt state (namely CA, OR, and WA), and 0 otherwise. RUST-BELT is a dummy variable taking value 1 if the firm has its headquarters in a rust-belt state (namely IL, IN, MI, OH, PA, and WV), and 0 otherwise. We opted to utilize these three dummies rather than state-specific dummies because observations would drop from the sample due to perfect prediction (e.g., if all firms headquartered in a specific state are non-union, their respective observations would be dropped from the sample). Also, the IV estimator (2SLS) has large biases when numerous instruments are used. Information on firms' headquarters location is included in Execucomp. The validity of these instruments is reflected in our data. The median number of unionized establishments per state in our sample is 77 for the rust-belt, 36 for the sun-belt and 26 for the bible-belt.

[9] Union density is averaged by state and weighted by employment.

[10] For instance, during the 1965-1980 period, union density averaged 26%, while in the period 1981-2001 it averaged 15%.

Our initial sampling pool began with 119,518 observations on executive compensation (15,069 CEO observations and 104,449 non-CEO observations) found in *Execucomp* during the 1992-2001 period. After dropping observations from non-US firms and observations containing missing variables for key independent variables utilized in this paper (namely, firm assets, return on assets and return to shareholders), we drop observations that correspond to less-than-annual compensation.[11] In the case of CEOs, we identify these observations using the information provided in *Execucomp* on CEO tenure. For managers, we identify these observations utilizing a number of sources, namely successive editions of the *S&P Register of Corporations, Directors and Executives*, corporate websites, proxy statements and Hoovers.com.[12] Also, given that firms report a variable number of executives each year, for uniformity purposes in estimations, we keep only observations for the CEO and the four highest non-CEO executives (hereafter referred to as *Managers*).[13] In this way, we create a dataset that uniformly contains all annual compensation information for the CEO and the four highest-paid managers in each firm-year. This final dataset contains 46,465 executive-year observations (9,293 CEO-years and 37,172 Manager-years), which come from 2,070 firms and 16,416 executives.

The key explanatory variable in the dataset is union presence (UNION_i=0/1), which is an indicator variable taking on the value 1 if the firm *i* had any establishments that were unionized, and 0 otherwise. In other words, a firm is considered unionized if it has at least one unionized establishment linked to the parent company.[14] In addition, in order to add variation to our unionization measure, we interact the 'Union_Presence' dummy with another dummy variable (High_Union_Presence) that indicates whether there are three or more unionized establishments in the firm.[15]

[11] It is important to drop observations with non-annual compensation not only because it is pro-rated but also because it typically includes transition perquisites (e.g., severance pay, signing bonus).

[12] We also verify continuing tenure of managers by whether they re-appear in SEC fillings of that firm in some subsequent year. For this purpose we explore *Execucomp* observations up to fiscal year 2006. However, we constrain the sample in the 1992-2001 period because the union data available from Bureau of National Affairs (BNA) extends to 2001.

[13] In particular, 35% of the firms in *Execucomp* (during the 1992-2001 period) report less than five executives' compensation and 10% of the firms report more than seven executives' compensation.

[14] Despite the seemingly low threshold for union presence in a firm, it is the most appropriate since establishments can be quite large in size, and since in the majority of unionized firms in the sample (i.e., 62%) we identify two or more unionized establishments that belong to the firm. Unfortunately, we cannot create a firm-level union-density variable because the Execucomp/Compustat variable on full-time employment is not representative of the actual number of workers at the firm and because BNA has a substantial portion of missing values for number of unionized workers at the establishment level.

[15] We choose 'three unionized establishments' as the cutoff point for the union strength dummy given that the median number of unionized establishments for firms with union presence in our sample is two. Notably, the results are robust to the selection of an alternative cutoff point (e.g., four or five unionized establishments).

We identify unionized establishments using the Bureau of National Affairs' (BNA) union activity data archive, from which we obtain information on a multitude of union activity variables, namely contract listings and NLRB elections with 'win' outcomes (and in some cases work stoppages and unfair labor practice petitions) that demonstrate evidence of union presence in 220,380 establishments in the United States between 1990 and 2002. We match the firm-level data from *Execucomp* with the establishment-level data from BNA using ownership information that is already included in BNA, as well as additional information from Hoovers Online, Dun & Bradstreet's Online, Harris Info-Source on firm establishments, firms' annual statements and firms' official websites.

Notably, even though we determine a discrete time dimension for each union activity entry before the BNA-*Execucomp* matching, the output firm-level dummy on union presence is essentially time-invariant ($UNION_i$ rather than $UNION_{it}$). This can be attributed to two factors: continuously updated union contract listings and multiple unionized establishments in firms with union presence. This pattern is also consistent with studies emphasizing the rarity of union decertification at the establishment level in the United States during the 1990s (Nilsson, 1997).

Using the constructed dummy on union presence, Table 1 offers an overview of union presence by industry in our sample. Consistent with US labor market union density data by industry, we observe that union presence at the firm-level is more evident in particular industries. For instance, substantial presence is found within the manufacturing sector, while only traces of union presence are found in the financial industry.

[Table 1 about here]

Table 2 presents descriptive statistics for the remaining variables used in this paper. The key dependent variables are *cash pay* (i.e. the sum of salary and bonus), *stock options* (i.e., Black-Scholes value at the time of the award), and *total compensation* (i.e. the sum of salary, bonus, benefits, long-term incentive plans, restricted grants and stock options). Independent variables include firm and executive characteristics and as instruments for union presence, we utilize three binary variables that denote headquarters location (i.e., *rust-belt*, *bible-belt*, *sun-belt*) and a continuous variable for *firm age*.

[Table 2 about here]

Figure 1 offers the unconditional distribution of the two main compensation components (*cash pay* and *stock options*) in executive compensation across unionized and non-unionized firms. More specifically, it shows that executives (both *Managers* and *CEOs*) in unionized firms have higher cash pay and lower stock option compensation compared to those in non-unionized firms.[16] At the same time, the compensation distributions seem to be more compressed (leptokurtic) for executives in unionized firms, as manifested by the higher peak and the thinner tails in all unconditional distributions presented in Figure 1. This initial indication of the different distributions of managerial compensation in unionized and non-unionized firms offers a fertile starting point for our analysis.

[Figure 1 about here]

4. Results

4.1 Pay for performance

This section estimates the relation between managerial compensation and market performance. If union presence within firms distorts managerial incentives, as Jensen and Murphy (1990) predict, we would expect to find lower pay-performance relation for managerial compensation in unionized firms.

The first-difference OLS regression specification for estimating the pay-performance elasticity for managers in the pooled sample of firms is as follows:

$$\begin{aligned}\Delta \ln(\text{Cash Pay})_{jit} = \beta_0 + \beta_1 \cdot \Delta \ln SV_{it} + \beta_2 \cdot \left(\Delta \ln SV_{it} \times UNION_i\right) + \beta_3 \cdot UNION_i \\ + \beta_4 \cdot \left(\Delta \ln SV_{it} \times cdfRisk_{it}\right) + \beta_5 \cdot cdfRisk_{it} + \varepsilon_{jit}\end{aligned} \quad [1]$$

where the independent variable is the percentage change in the cash pay compensation for manager j in firm i at year t, while β_1 indicates the pay-for-performance estimate associated with changes in shareholder value $(\Delta \ln SV)$, and β_2 denotes the slope-effect on pay-for-performance from union presence, after controlling for firm risk. We opt for pay-performance elasticity -rather than sensitivity- because it reduces the impact of firm size bias and better illustrates linearity in agency contracts.[17] The change in shareholder value

[16] Notably, in Figure 1, there is no difference in the binary incidence of stock option awards between non-unionized and unionized firms (circa 74%, respectively, offer stock options to their top-5 executives in the sample). In this way, the omission of zero stock option awards due to logarithmic transformation does not affect the results.

[17] Pay sensitivity measures the sharing rate between the CEO and the firm, by regressing changes in CEO compensation to changes in shareholder wealth. In larger firms this sharing rate has to decline because executives, being risk-averse and having limited personal liability, can only bear a much smaller fraction of the firm's fluctuation in total value. As Rosen (1992: p. 200) first argued, when the pay-performance specification involves arithmetic (rather than logarithmic) differences, then larger firms in the sample dominate the estimates. As a result, the pay-performance link seems far weaker. In contrast, pay elasticity involves logarithmic differences (i.e. growth in a percentile form) that are largely independent of size and, thus, reflect managerial incentives more accurately.

equals $\Delta \ln(1 + \text{Return to Shareholders})$, while firm risk is defined as the cumulative distribution of three-year monthly variation in firms' share prices (Murphy, 1999; Aggarwal and Samwick, 1999).[18] Following Conyon and Murphy (2000: p. 661), the effect of the binary variable is captured by interacting the union presence dummy with the firm performance measure. Compensation elasticity to financial performance is estimated only for cash pay, since managers' wealth is much less related to the firm's share price compared to the CEO wealth. In particular, managers receive substantially less stock-related compensation compared to their CEO and they have much lower stock ownership than the CEO.[19]

Table 3 demonstrates that performance elasticity of cash pay with respect to market performance is positive and significant across unionized and non-unionized firms. The lack of significance for the interaction term $(\Delta \ln SV \times UNION)$ in Table 3 is consistent with union presence having no significant negative effect on CEO incentives. The results do not change even after performing IV regression, thus allowing for endogeneity of union presence, or including firm risk in the specification. Notably, the findings in Table 3 are consistent with principal-agent theory predictions, according to which increased risk has a negative effect on the alignment of incentives (Prendergast, 2002). Finally, the sign and statistical significance for both firm risk and the interaction between firm risk and firm performance are consistent with the results in Aggarwal and Samwick (1999).

[Table 3 about here]

4.2 Promotion-based incentives

Given that the pay-performance elasticity is positive and not significantly different across unionized and non-unionized firms, managers in unionized firms appear to be as motivated as their counterparts in non-unionized firms. Our hypothesis is that managers are given greater promotion-based incentives to offset the decrease in stock-based compensation. This

[18] For a small number of firms in our sample that were newly listed (and, consequently, had less than three years of trading history), we calculate their risk as the average risk of firms in the same 4-digit SIC industry (or alternatively 3-digit SIC industry).

[19] For instance, the average percentage of the company's shares owned by senior managers and CEOs in our sample is 0.36% and 2.85%, respectively. In absolute dollar amount, the median stock-option award for senior managers and CEOs is $141,570 and $426,236 respectively. Also, the median ratio of "stock option awards" to "total compensation" for senior managers is 24%, while the respective ratio for the CEOs is 29%. Notably, utilizing Schaefer's (1998) definition of total compensation in the context of pay-performance elasticity, we confirm the findings from our cash-pay regressions. However, due to the design of his independent variable the sample is smaller (Results available from the authors upon request).

would be the case if (*a*) the relative compensation gap between managerial compensation and the compensation of the CEO is significantly larger than that of managers in non-unionized firms, and/or if (*b*) the possibility of the CEO's successor being drawn from the internal pool of managers –rather than being an external hire– would be higher than that of managers in non-unionized firms. We test for these associations below.

4.2.1 Intra-Firm Compensation differentials

In this sub-section we examine whether there are differences between unionized and non-unionized firms in terms of compensation differentials between the CEO and the top managers. This differential can be measured by the distance in remuneration within the top echelon inside the firm, effectively approximating the promotional prize between senior managers and the CEO. Distance is defined as the difference between CEO remuneration and the median remuneration for the top 4 managers in firm i at time t [$Dist_{it} = (W_{CEO} - W_{Executives})_{it}$].[20] Distance is measured for both cash pay and total compensation. The transformed value of the distance measure is then regressed against union presence and a number of control variables.[21] Put formally:

$$\text{Distance}_{it} = \alpha_0 + \alpha_1 \cdot UNION_i + \gamma \cdot z_{it} + \varepsilon_{it} \qquad [2]$$

, where α_0 is the intercept term, $UNION_i$ is the indicator variable indicating union presence, z_{it} is a vector of control variables on CEO characteristics[22], firm characteristics, median compensation of top-4 managers (to control for initial conditions) as well as year effects (to address any differential effect in the growth of stock option compensation), and ε specifies the error term.

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[20] As explained in Section 2, the top four managers consist of the four highest paid non-CEO managers in the firm.

[21] Due to the presence of negative outliers, for the transformation of distance values we utilize the inverse hyperbolic sine function (Burbidge et al., 1988) that not only yields very similar results to those of the logarithmic function for positive numbers, but also transforms negative numbers. Negative distance values occur in cases of CEOs who have substantial wealth, mostly due to their founder status in the firm, and are employed for substantially less compensation than their counterparts. For instance, Apple's CEO Steve Jobs symbolically gets $1 as annual salary. The values of the dependent variables are windsorized at the 5% level in order to mitigate influence by extreme outliers, in particular observations with negative distance values (i.e., CEO compensation is less than the median compensation of the top four managers). See Appendix 1.

[22] CEO characteristics that can influence the distance measure are *CEO/Chairman duality* (a CEO/Chairman is expected to be paid more because she has dual responsibilities and faces more complex tasks, compared to a CEO who is not the chairman of the board), *CEO tenure* (the marginal returns from CEO tenure may be different than those for managerial tenure), and *CEO share ownership* (incentives are better aligned for CEOs with higher ownership).

Due to the presence of negative outliers (see Appendix 1) and our expectation of a positive, but diminishing, union effect due to union aversion to inequality, we utilize quantile regression estimation with the union presence dummy being exogenous. The benefit of quantile regression is that it examines the impact of union presence on the entire conditional distribution of compensation, rather than as a single central tendency measure. For robustness purposes, we also perform IV (2SLS) regression where union presence is endogenous.[23]

Table 4 presents the findings for both cash pay and total compensation. Consistent with our hypothesis, it is found that union presence substantially increases intra-firm cash pay differentials between the CEO and the top managers. The same holds for the union presence relation with respect to total compensation differentials. In particular, union presence is found to increase the median cash pay differential by 10% for unionized firms and by 13% for firms with high union presence (Columns 1-2). The respective median effects for total compensation differentials are 9% for all unionized firms, and 12% for firms with high union presence (Columns 4-5).

Estimations that account for the endogeneity of union presence yield similar results (Columns 3 and 6). In particular, utilizing IV regressions, it is found that union presence increases the differentials for cash pay and total compensation by 48% and 45% respectively.[24] Not surprisingly, IV regression estimates are much higher than the quantile regression ones since they are influenced by outliers and they do not account for the binary nature of the endogenous variable. Nevertheless, the fact that both estimations produce positive and significant coefficient estimates for union presence offers robust evidence for a higher promotion prize in unionized firms.

Figures 2a and 2b capture the heterogeneity in the union effect across different levels of compensation differentials, utilizing the quantile regression specifications in Table 4 (Columns 1 and 4, respectively). The union effects at lower and higher quantiles, respectively, are significantly different from each other, thus verifying that firms with union

[23] Recently, there have been developments in the estimation methods for IV quantile regression (e.g., Abadie, et al., 2002; Froelich and Melly, 2008). However, we do not utilize these methods since they can accommodate only a single binary instrument.

[24] The aforementioned results are robust to alternative definitions of distance (e.g., if distance is estimated as the difference between the CEO remuneration and the mean, instead of the median, compensation of the top-4 managers) or alternative measures of distance [e.g., taking the ratio between the CEO's and Managers' compensation rather than the natural logarithm of the arithmetic difference in their compensation, or, creating the cumulative distribution function of distance thus indicating the transition from more equivalent differences ($cdfDist_{it} \to 0$) to more unequal differences ($cdfDist_{it} \to 1$)].

presence favor an increased promotion prize, but not at the detriment of generating dramatic compensation inequality between the CEO and the lower-level employees. For example, union presence increases the "*Distance in Cash-Pay*" by 21% at the 0.2 quantile, but only by 5% at the 0.8 quantile. In the case of "*Distance in Total Compensation*", union presence causes an increase of 38% at the 0.1 quantile, but has an insignificant effect at the 0.9 quantile.

[Table 4 & Figures 2a/b about here]

As an additional robustness check we estimate the following specifications for the pooled sample, in order to test the relation between union presence and its differential effect on cash pay and total compensation, between managers and their CEO along the conditional compensation distribution:

$$\begin{aligned}\ln(\text{Cash Pay})_{jit} = {} & \alpha_0 + \alpha_1 \cdot (UNION \times CEO)_{jit} + \alpha_2 \cdot (UNION \times Manager)_{jit} \\ & + \alpha_3 \cdot CEO_{jit} + \alpha_4 \cdot Chairman_{jit} + \alpha_5 \cdot \ln Assets_{it} \\ & + \alpha_6 \cdot ROA_{it} + \alpha_7 \cdot RET_{it} + \gamma \cdot Industry_i + \delta \cdot Year_t + \varepsilon_{jit}\end{aligned} \quad [3]$$

$$\begin{aligned}\ln(\text{Total Comp})_{jit} = {} & \beta_0 + \beta_1 \cdot (UNION \times CEO)_{jit} + \beta_2 \cdot (UNION \times Manager)_{jit} \\ & + \beta_3 \cdot CEO_{jit} + \beta_4 \cdot Chairman_{jit} + \beta_5 \cdot \ln Assets_{it} \\ & + \beta_6 \cdot ROA_{it} + \beta_7 \cdot RET_{it} + \gamma \cdot Industry_i + \delta \cdot Year_t + u_{jit}\end{aligned} \quad [4]$$

, where the independent variables are the natural logarithms of cash pay and total compensation for executive *j*, in firm *i* at year *t*. The two interactions of union presence with the *CEO* and *Manager* (i.e., non-CEO senior manager) dummies provide the slope parameters at different conditional quantiles of the compensation distribution. This is estimated after controlling for positional and firm characteristics. In particular, *CEO* and *Chairman* are dummies controlling for the compensation premium of the CEO and Chairman positions, respectively, while firm characteristics refer to firm size (natural logarithm of assets) and performance (return on assets, return on shareholders).

Figure 3 illustrates the estimated coefficients for the two interaction terms, which confirm the stronger negative relation found for union presence on managers' total compensation as compared to CEOs' total compensation, throughout the conditional compensation distribution. Post estimation hypothesis testing for the equality of the estimates of the interaction terms (i.e., Wald test with H_0: $\alpha_1 = \alpha_2$, or, H_0: $\beta_1 = \beta_2$) rejects these at the 10 percent significance level for all quantiles in specification [4], and for all -but two- quantiles in specification [5].

The negative union effect with respect to total compensation strengthens as we move to the upper tail of the conditional compensation distribution, thus providing evidence in support of Jensen and Murphy's (1990) assertion that unions are particularly concerned with constraining high-end-compensation packages that 'super-star' executives often can command. Notably, Figure 3 also indicates the disproportionally negative and significant effect of union presence on stock options for both managers and CEOs.

[Figure 3 about here]

4.2.2 Patterns of internal CEO succession

Using Execucomp, we identify 781 turnovers during our sample period (1992-2001). After excluding 63 turnovers that involve founders taking the helm, we are left with 718 turnovers.[25] Table 5 provides descriptive statistics about the nature of succeeding CEOs. Although most of the CEO turnovers come from non-unionized firms, there is no substantial difference in terms of turnover rate (i.e., 33% vs. 37%), after accounting for the population of each firm type in the sample. Also, the two types of firms have very similar rates of *outsider* CEO succession, 43% and 41% for non-unionized and unionized firms, respectively. However, they dramatically differ in terms of prior firm tenure for insider CEOs: the median tenure of insider CEOs in unionized firms is ten years longer than that of insider CEOs in non-unionized firms.

[Table 5 about here]

These patterns of CEO succession suggest substantial variation across the prior firm-specific tenure of insider CEOs. Further illustrating this point, Figure 4 presents the distribution of firm-specific tenure prior to promotion in the sample of 442 insider CEOs in both unionized and non-unionized firms. Notably, there is a wide variation that spans well into 10- or 20-year careers in the firm; only 29% of the insider CEOs spent between 1-5 years in the firm prior to their promotion, while 41% spent more than 15 years in the firm. Firm-specific experience for insider CEOs are even longer in unionized firms. Using the sample of 442 insider CEOs, Figure 5 illustrates the distribution of prior firm-specific experience across three types of firms: non-unionized firms, firms with low union presence (<3 unionized establishments), and firms with high union presence (≥3 unionized

[25] When an executive is the firm's founder or co-founder then she is a de-facto insider, thus perfectly predicting the outcome in the context of equation [5].

establishments). In non-unionized firms, 50% of insider CEOs have at most ten years of prior tenure in their firms when they were promoted to the CEO position; this sharply contrasts the 28% of insider CEOs in firms with high union presence. On the other hand, only 21% of insider CEOs in non-unionized firms have more than twenty years of prior tenure, compared to 56% of insider CEOs in firms with high union presence. Another interesting pattern illustrated in Figure 5 is that the firm-specific experience pattern of companies with low union presence falls roughly between those of non-unionized and highly-unionized firms, thus indicating a potential effect of union strength on promotion patterns within a firm.

[Figures 4 and 5 about here]

The aforementioned descriptive statistics call for a dual examination of the union presence effect on CEO promotion patterns, specifically on the outsider/insider choice of CEO and the length of his firm-specific tenure. In analyzing these two effects, we utilize binary choice models to explain whether union presence influences the binary outcome of the CEO being an insider, and we employ a Tobit model to investigate whether union presence influences the expected level of firm-specific experience in becoming a CEO.[26]

In order to examine whether unionized firms differ in terms of insider succession to the CEO position, we estimate the following likelihood function for the probability that the CEO is an insider, instead of an external hire:

$$CEOi\text{sInsider}_{ji} = \begin{cases} 1 & \text{if } \delta_0 + \delta_1 \cdot UNION_i + \mu \cdot x_{ji} + \varepsilon_{ji} > 0 \\ 0 & \text{otherwise} \end{cases} \qquad [5]$$

where *CEOisInsider* is a dummy variable for each unique CEO-firm observation indicating that the CEO j worked in firm i for more than one year before his appointment as the firm's CEO; *UNION* is our dummy variable for union presence, and x is a vector of control variables influencing the origins of CEO succession.[27] More specifically, in terms of control variables we employ firm size (assets) at the time of succession, industry homogeneity, and a dummy

[26] Although the number of years is a continuous variable, we do not utilize OLS regression because the dependent variable contains a large cluster of zeros (i.e., 43% of the observations have zero years, reflecting outside CEO succession), and thus OLS results would be give biased and inconsistent.

[27] We do not need to control for performance since we do not estimate the reason for CEO succession, but the likelihood of the CEO being an insider rather than an outsider. The financial literature on CEO turnover provides evidence that long-term firm performance influences the likelihood of forced CEO turnover but not the decision on an internal or external succession.

variable indicating whether the CEO is a relative of the firm's (co-)founder.[28] *Firm size* positively influences the pool of potential internal candidates, while when the founding family has a strong grip on the firm's control, we would expect an executive who is *a relative of the founder* to have a better chance of being promoted to the CEO position. Also, Parrino's (1997) measure of industry homogeneity is included since he finds that it influences the nature of CEO succession (i.e. internal/external).[29] In order to account for potential endogeneity in union presence, besides the standard probit estimation, we also perform a bivariate probit estimation, which is widely utilized when both the outcome variable and the endogenous variable are binary choice variables (Greene, 1998; Neal, 1997).

For consistency purposes, we utilize the same control variables in the Tobit model examining whether unionized firms differ in terms of the expected level of firm-specific experience in becoming CEO. The only difference is that the dependent variable is the number of firm-specific tenure (in years) prior to becoming a CEO. More specifically, the dependent variable is observable only when the latent variable is positive:

$$\text{Firm - specific Tenure}_{ji} = \max(0, \alpha_0 + \alpha_1 \cdot UNION_i + \mu \cdot x_{ji} + u_{ji}) \qquad [6]$$

Table 6 offers mixed results for the union presence effect on the outsider/insider CEO choice. The probit estimation (Columns 1 and 3) shows that firms with union presence are not significantly different in terms of choosing an insider over an outsider during CEO succession. However, once we account for potential endogeneity with the bivariate probit estimation we find that union presence increases by 5.2% the likelihood of the CEO being an insider (Column 2). This finding holds for all unionized firms and is not driven by firms with high union presence (Column 4).

[Table 6 about here]

Unlike the binary outcome estimates, estimates from the Tobit regressions (Table 6, Columns 5-6) provide strong evidence for a union effect on the expected firm-specific

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[28] Regarding CEO's status as a founder or a relative (by blood of marriage) of a founder, detailed information was manually collected on executive biography and corporate history from firms' annual reports, Dun & Bradstreet's *Million Dollar Database*, Hoovers *Online*, the Standard and Poor's *Register of Corporations, Directors, and Executives*, company press releases and official websites, and the business press.

[29] Parrino (1997) calculates intra-industry homogeneity using the partial correlation between 2-digit SIC industry returns and common stock returns of the firms in that industry, hence reflecting the degree to which firms belonging to the same industry tend to react similarly to news. Notably, industry homogeneity deals with firm reaction to industry-wide shocks, and not with intra-industry performance. This means that firms in a homogeneous industry can have greatly different performance. We follow Parrino's (1997) methodology in calculating the homogeneity index across the 1992-2001 period.

experience. In particular, these estimates suggest an increase in the expected firm-specific tenure by three years across unionized firms. When we include a measure of high union presence, we observe that the positive effect is absorbed by the interaction, indicating that only in highly unionized firms the succeeding CEO has 5.6 years more firm-specific experience compared to incoming CEOs in the other two firm types, i.e., non-unionized firms and firms with low union presence. Indirectly, these findings reveal an oversight of the traditional definition of insider CEO utilized by the economic and financial literature (i.e., having spent at least one year with the firm prior to her appointment to the CEO position); by assuming a power-law distribution of insider tenure the traditional definition does not consider variations in insider CEOs.

The results above clarify two issues. First, there is no tradeoff between promotion prize and probability of insider succession; unionized firms have a larger promotion prize and similar probability of insider succession. Second, the prior firm-specific tenure for insider CEOs varies substantially across firm type; insider CEOs in unionized firms have substantially longer tenures as managers with the firm. An explanation for this finding is that unionized firms utilize successive promotions –all the way up to the CEO position– to induce managers to acquire nonverifiable firm-specific human capital (Prendergast, 1993).

Notably, the control variables have the expected signs in all the aforementioned estimations. Both firm size and being a relative of the firm's founder are found to be positive and statistically significant. Industry heterogeneity is found to be positive but weakly significant in the case of the likelihood of insider CEO succession; this may be explained by the reduction in industry heterogeneity during the 1990s.[30] Finally, the sign and the significance of the results in Table 6 do not change when we include additional control variables on executive and firm characteristics at the time of CEO turnover, such as CEO age, and firm's idiosyncratic risk and sales growth.

Overall, the combination of larger promotion prize to the CEO position and higher firm-specific human capital acquisition suggest that during our sample period, unionized firms induce managerial effort through promotion-based incentive schemes as compared to non-unionized firms which opt for wider use of performance-based incentives like stock options.

[30] In Parrino's (1997) sample period of 1970-1988, the average score for forty 2-digit SIC industries is 0.2974, while in our sample the average score for the same 2-digit SIC industries is 0.2688. This difference is statistically significant at the 1% level.

4.3 Do unions compress the distribution of managerial compensation?

The previous sub-sections focused on the average/median effect of union presence. In this sub-section we focus on the effect of union presence on the entire distribution of managerial compensation. This is because one of the most pervasive union effects is often not confined to wage-premium, but to the compression of the overall distribution of wages across firms (Card, 2001; Card et al. 2003; DiNardo et al., 1996).

In assessing variation in the distribution of managerial compensation, we measure the inequality in the remuneration awarded to managers in unionized and non-unionized firms. As discussed previously, by matching firm size, firm performance and industry classification we also control for a large part of observed managerial quality and skill (Gabaix and Landier, 2008; Murphy, 1999; Tervio, 2008), thus creating two comparable populations of managers that differ only in that they work either in a firm with or without union presence. More specifically, we seek to match each unionized firm with a unique non-unionized firm that operates in the same year, in the same industry (2-digit SIC), and that has almost identical firm size (in terms of market value, sales and assets) and firm performance (in terms of return on assets and return to shareholders). The pairs are also matched in terms of three CEO characteristics; whether the CEO is also Chairman of the Board, the founder of the firm, or a relative of the founder.[31]

We construct the matched-pairs by utilizing the nearest-neighbor matching algorithm from Abadie et al. (2004). Two samples of unionized and non-unionized firms are created from the pairs of matched firms.[32] Each of the two samples contains 943 firm-year observations. Thus, we compare the distributions of the two populations each having 3796 managers (i.e., 943 firms times 4 non-CEO top managers per firm), and the key differentiating factor is union presence in the firm. After matching on observables, we assume that the two populations of executives have similar underlying abilities and skills, but they differ in terms of the binary treatment of union presence.

[31] Matching for CEO's founder status is important since CEOs that are either founders or part of the founding family often receive different compensation from professional managers (Anderson and Bizjak, 2003; Gomez-Mejia et al., 2004).

[32] In order to avoid poor matching, we drop 10% of the matched pairs with the poorest matching score, as measured by the *distance* metric. Post-matching mean and median tests for the two groups of unionized and non-unionized firms indicate no statistically significant difference for all the variables utilized as matching criteria at any conventional significant level.

Inequality of remuneration within each group of managers is measured with three indices, namely the Gini coefficient, the Theil entropy measure and the standard deviation of logs. The motivation for using multiple inequality measures is that they have different degrees of decomposability among population subgroups (Shorrocks, 1984; Toyoda, 1980), different emphases of distribution transfers (Blackorby and Donaldson, 1978), and possibly intersecting Lorenz curves (Champernowne, 1974). Also, the three aforementioned indices were utilized by DiNardo et al. (1996) to examine the effect of unions on US wage inequality and provide a set of comparable findings. The large sample properties of these indices allow us not only to consider their point estimates, but also to apply statistical inference procedures in order to test for differences in the inequality indices between the matched samples (Rongve and Beach, 1997; Zheng and Cushing, 2001).

[Table 7 about here]

The results in Table 7 (Panel A) indicate that union presence does compress the cash pay distribution of managers. For instance, the Gini coefficient for unionized firms is 0.323, while for non-unionized it is 0.347. However, we find no such compression for the distribution of stock option compensation and total compensation. Although, there are differences in each inequality index, these are not collectively significant at the 5% level. Notably, the aforementioned results do not change if we isolate the comparison between matching pairs of highly unionized firms and non-unionized firms.[33]

Besides the indices of distribution inequality, we also perform mean and median comparisons for managerial compensation in the two matched samples (Table 7, Panel B).[34] In terms of statistically significant differences, we find that unionized firms offer -on average- 24% lower stock option compensation and 7% lower total compensation. Given that there is negligible difference in the binary incidence of stock option awards (i.e., 73.94% and 72.64% of the managers in the matched unionized and non-unionized firms, respectively, are awarded stock options), we can attribute this significant difference in total compensation across firm types to differences in the size of the stock option awards. These results further support Freeman's (1980, 1982) long-held assertion that a primary objective of trade unions is to reduce wage dispersion through the use of standard rates (i.e., salary and capped accounting-based bonuses in the case of managers).

[33] For brevity purposes these results are not included, but they are available from the authors upon request.

[34] Basically, the difference in means between the matched samples is the Average Treatment Effect (Abadie et al., 2004) for the binary treatment of union presence.

Overall, two main findings emerge regarding the distribution of managerial compensation across the matched samples of unionized and non-unionized firms. Cash pay distribution in unionized firms is more compressed but with a higher mean. In contrast, stock option compensation and total compensation are similarly dispersed across unionized and non-unionized firms, but with a lower mean/median for unionized firms.

5. Conclusions

The concept of fairness has been used in economics to explain involuntary unemployment (Akerlof and Yellen, 1990), wage rigidity (Bewley, 2004) and salary dispersion in cohesive groups (Lazear, 1989; Levine, 1991). In this paper, fairness is viewed from the perspective of workers comparing themselves with those in dissimilar occupations, in particular with top-executives, within the same firm. Fairness concerns, in this context, are related to the concept of *envy*, whereby worker utility is reduced with increased inequality. Union presence offers one way of capturing envious workers within a firm, since unions not only seek overall equity at the workplace but they can also disrupt firm performance when their fairness expectations are not met.

Our results indicate that firms adjust managerial compensation in the presence of unions. In particular, we find that union presence is associated with greater promotion-based incentives by offering a larger promotion prize to the CEO position coupled with a slightly higher insider succession probability. It is also found that union presence increases fixed compensation (i.e., cash pay) and compresses the corresponding distribution, while at the same time substantially reducing the variable compensation (i.e., stock options) component without compressing its respective distribution. This means that by boosting promotion-based incentives and offsetting lower variable-compensation with higher fixed-compensation, unionized firms appear to prevent a flight of managerial talent to the non-unionized firms.

These findings also suggest that a trade-off exists between stock-option compensation and promotion-based incentives, which is consistent with incentive design when individual effort is unverifiable, as is the typical case for senior managers. Notably, our results regarding union presence differ from those examining other cases of implicit regulation, namely government oversight in utility firms.[35] One of the reasons, perhaps, is that unlike utilities that typically operate in local monopolies or concentrated markets, publicly listed US firms (unionized or otherwise) operate in competitive environments subject to product market pressures, thus having to compete in the labor market for executives.

This paper also ultimately serves as an impetus for further work on the interaction between labor institutions and managerial remuneration. Future research could focus on middle-level managers whose compensation information is not typically disclosed, or managerial compensation changes across the private sector as a result of rising industrial action.[36] These topics also have a timely appeal in light of the US private sector union density increases since 2007, reversing a three-decade-long declining trend, as well as the potentially sizeable gains emanating from the proposed Employee Free Choice Act (EFCA), which is currently pending Congressional approval and plans to amend the National Labor Relations Act making it significantly easier for unions to organize employees.

[35] See Joskow et al. (1993, 1996).

[36] Notably, during our sample period, labor relations at the macro-level in the United States remained stable without any disputes spikes. This is based on indicators on work stoppages, industrial disputes and quality of labor relations from the Bureau of Labor Statistics and the IMD World Competitiveness Yearbook.

References

Abadie, A., Angrist, J., Imbens, G.W. 2002. Instrumental variable estimates of the effect of subsidized training on the quantiles of trainee earnings. *Econometrica*, 70(1), pp. 91-117.

Abadie, A., Drukker, D., Leber-Herr, J., Imbens, G.W. 2004. Implementing matching estimators for average treatment effects in STATA. *Stata Journal*, 4(3), pp. 290-311.

Aggarwal, R.K., Samwick, A.A. 1999. The other side of the trade-off: The impact on executive compensation. *Journal of Political Economy*, 107, pp. 65-105.

Akerlof, G.A., Yellen, J.L. 1990. The fair wage-effort hypothesis and unemployment. *Quarterly Journal of Economics*, 105(2), pp. 255-283.

Anderson, R.C., Bizjak, J.M. 2003. An empirical examination of the role of the CEO and the compensation committee in structuring executive pay. *Journal of Banking and Finance*, 27, pp. 1323-1348.

Atanassov, J., Kim, E.H. 2007. Labor and corporate governance: International evidence from restructuring, *Journal of Finance*, forthcoming.

Burbidge, J.B., Magee, L., Robb, A.L. 1988. Alternative transformations to handle extreme values of the dependent variable. *Journal of the American Statistical Association*, 83, pp. 123-127

Bartling, B., Siemens, F. 2006. The intensity of incentives in firms and markets: Moral hazard of envious agents. GESY Discussion Paper No. 115, University of Munich.

Bewley, T.F. 2004. Fairness, reciprocity, and wage rigidity. IZA Discussion Paper No. 1137.

Bewley, T.F. 2003. Labor market behavior. Yale University Working Paper.

Bewley, T.F. 1999. Why wages don't fall during a recession. Cambridge, MA: Harvard University Press.

Blackorby, C., Donaldson, D. 1978. Measures of relative inequality and their meaning in terms of social welfare. *Journal of Economic Theory*, 18, pp. 59-80.

Burks, S.V., Carpenter, J.P., Verhoogen, E. 2007. Fairness and freight-handlers: Local labor market conditions and wage fairness perceptions in a trucking firm. *Industrial and Labor Relations Review*, 60(4), pp. 477-498.

Card, D., Lemieux, T., Riddell, W.C. 2003. Unions and wage inequality. In J.T. Addison and C. Schnabel (Eds.), *The international handbook of trade unions*, Cheltenham: Edward Elgar, pp. 246-92

Card, D. 2001. The effect of unions on wage inequality in the US labor market. *Industrial and Labor Relations Review*, 54(2), pp. 296-315.

Carruth, A., Oswald, A.J. 1987. On the union preferences and labour market models: Insiders and outsiders. *Economic Journal*, 97, pp. 31-45.

Champernowne, D. 1974. A comparison of measures of inequality of income distribution. *Economic Journal*, 84, pp. 787-816.

Chen, H., Kacperczyk M., Ortiz-Molina H. 2007. Do labor unions affect agency costs of debt?, Working Paper, University of British Columbia.

Choi, S. 2000. Proxy issue proposals: Impact of the 1992 SEC proxy reforms. *Journal of Law, Economics and Organization*, 16(1), pp. 233-268.

Conyon, M.J., Murphy, K.J. 2000. The prince and the pauper? CEO pay in the United States and United Kingdom. *Economic Journal*, 110, pp. F640-F671.

Core, J.E., Guay, W.R., Larcker, D.F. 2003. Executive equity compensation and incentives: A survey. *FRBNY Economic Policy Review*, April 2003.

Cronqvist, H., Heyman, F., Nilsson, M., Svaleryd, H., Vlachos, J. 2009. Do entrenched managers pay their workers more?, *Journal of Finance*, 64(1), pp. 309- 339.

DiNardo, J., Fortin, N.M., Lemieux, T. 1996. Labor market institutions and the distribution of wages, 1973-1992: A semiparametric approach. *Econometrica*, 64(5), pp. 1001-1044.

Dur, R., Glazer, A. 2008. Optimal contracts when a worker envies his boss. *Journal of Law, Economics and Organization*, 24(1), pp. 120-137.

Englmaier, F., Wambach, A. 2005. Optimal incentive contracts under inequity aversion. IZA Discussion Paper No. 1643.

Fehr, E., Goette L., Zehnder, C. 2007. The behavioral economics of the labor market: Central findings and their policy implications. University of Zurich Working Paper.

Fehr, E., Kirchsteiger, G. 1994. Insider power, wage discrimination and fairness. *Economic Journal*, 104, pp. 571-583.

Fiorito, J. 2007. The state of unions in the United States, *Journal of Labor Research*, 28(1), pp. 43-68.

Freeman, R.B. 1980. Unionism and the dispersion of wages. *Industrial and Labor Relations Review*, 34(1), pp. 3-23.

Freeman, R.B. 1982. Union wage practices and wage dispersion within establishments. *Industrial and Labor Relations Review*, 36(1), pp. 3-21.

Freeman, R.B. 1988. Contraction and expansion: The divergence of private sector and public sector unionism in the United States, *Journal of Economic Perspectives*, 2(2), pp. 63-88.

Froelich, M., Melly B., 2008. Quantile treatment effects in the regression discontinuity design. IZA Discussion Paper No. 3638.

Gabaix, X., Landier A. 2008. Why has CEO pay increased so much?, *Quarterly Journal of Economics*, 123(1), pp. 49-100.

Gibbs, M., 1995. Incentive compensation in a corporate hierarchy. *Journal of Accounting and Economics*, 19, pp. 247-277.

Gillan, S.L., Starks, L.T. 2000. Corporate governance proposals and shareholder activism: the role of institutional investors. *Journal of Financial Economics*, 57(2), pp. 275-305.

Gomez-Mejia, L.R., Larraza-Kintana, M., Makri, M. 2004. The determinants of executive compensation in family-owned firms. *Academy of Management Journal*, 46, pp. 226-238.

Greene, W.H. 1998. Gender economics courses in liberal arts colleges: Further results. *Journal of Economic Education*, 29(4), pp. 291-300.

Gregory, R. G., Borland, J. 1999. Public sector labor markets. In: O. Ashenfelter and D. Card (Eds.), *Handbook of Labor Economics*, Volume 3c, Chapter 53, Elsevier, Amsterdam, pp. 3573-3630.

Herod, A., 1998. *Organizing the landscape: Geographical perspectives on labor unionism*. Minneapolis, MN: University of Minnesota Press.

Hirsch, B.T., 2008. Sluggish institutions in a dynamic world: Can unions and industrial competition coexist? *Journal of Economic Perspectives*, 22(1), 153-176.

Hirsch, B.T., Macpherson, D.A. 2003. Union membership and coverage database from the current population survey: Note. *Industrial and Labor Relations Review*, 56(2), pp. 349-354.

Holmes, T.J., 2006. *Geographic spillover of unionism*. Federal Reserve Bank of Minneapolis Staff Report No. 368.

Hubbard, R.G., Palia, D. 1995. Executive pay and performance: Evidence from the US banking industry. *Journal of Financial Economics*, 39(1), pp. 105-130.

Itoh, H. 2004. Moral hazard and other-regarding preferences. *Japanese Economic Review*, 55(1), pp. 18-45.

Jensen, M.C., Murphy, K.J. 1990. Performance pay and top-management incentives. *Journal of Political Economy*, 98(2), pp. 225-264.

Joskow, P.L., Rose, N.L., Shepard, A. 1993. Regulatory constraints on CEO compensation, *Brooking Papers: Microeconomics*, 1, pp. 1-72

Joskow, P.L., Rose, N.L., Wolfram, C.D. 1996. Political constraints on executive compensation: Evidence from the electric utility industry, *RAND Journal of Economics*, 27(1), pp. 165-182.

Kahn, L.M., Curme, M. 1987. Unions and nonunion wage dispersion. *Review of Economics and Statistics*, 69(4), pp. 600-607.

Klasa, K., Maxwell, W.F., Ortiz-Molina H. 2009. The strategic use of corporate cash holdings in collective bargaining with labor unions. *Journal of Financial Economics*, forthcoming.

Kole, S.R., Lehn, K. 2000. Workforce integration and the dissipation of value in mergers: The case of USAir's acquisition of Piedmont Aviation. In S. Kaplan (Ed.) *Mergers and productivity*, University of Chicago Press, pp. 239-279.

Krueger, A., Mas, A. 2004. Strikes, scabs and tread separations: Labor strife and the production of defective Bridgestone/Firestone tires. *Journal of Political Economy*, 112(2), pp. 253-289.

Lazear, E.P. 1989. Pay equality and industrial politics. *Journal of Political Economy*, 97(3), pp. 561-580.

Levine, D. 1991. Cohesiveness, productivity, and wage dispersion. *Journal of Economic Behavior and Organizations*, 15, pp. 237-255.

Lopez, S.H. 2004. *Reorganizing the rust belt: An inside study of the American Labor Movement*, Berkeley: University of California Press.

Main, B., O'Reilly, C. , Wade J. 1993. Top executive pay: Tournament or teamwork. *Journal of Labor Economics*, 11, pp.606-628.

Mas, A. 2008. Labor unrest and the quality of production: Evidence from the construction equipment resale market. *Review of Economic Studies*, 75(1), pp. 229-258.

Mas, A. 2006. Pay, reference points, and police performance. *Quarterly Journal of Economics*, 121(3), pp. 783-821.

Murphy, K.J. 1999. Executive compensation. In O. Ashenfelter and D. Card (Eds.), *Handbook of labor economics*, vol. 3, Amsterdam: North Holland, pp. 2486-2563.

Neal, D. 1997. The effects of catholic secondary schooling on educational achievement. *Journal of Labor Economics*, 15(1), pp. 98-123.

Nilsson, E. 1997. The growth of union decertification: A test of two nonnested theories. *Industrial Relations*, 36(3), pp.324-43.

Palley, T.I., LaJeunesse R.M. 2007. Social attitudes, labor law, and union organizing: Toward a new economics of union density, *Journal of Economic Behavior and Organization*, 62(2), pp. 237-254

Parrino, R. 1997. CEO turnover and outside succession: A cross-sectional analysis, *Journal of Financial Economics*, 46(2), pp. 165-197.

Pagano, M., Volpin, P. 2008. Labor and Finance. London Business School Working Paper.

Pastor, L, Veronesi, P. 2003. Stock valuation and learning about profitability. *Journal of Finance*, 58(5), pp. 1749-1790.

Philippon, T., Resef, A. 2008. Wages and human capital in the US financial industry: 1909-2006. NYU Stern Working Paper.

Prendergast, C. 2002. The tenuous trade-off between risk and incentives. *Journal of Political Economy*, 110(5), pp. 1071-1102.

Prendergast, C. 1993. The role of promotion in inducing specific human capital acquisition. *Quarterly Journal of Economics*, 1993, 108(2), pp. 523-534.

Rongve, I., Beach, C. 1997. Estimation and inference for normative inequality indices. *International Economic Review*, 38(1), pp. 83-96.

Rosen, S., 1992. Contracts and the market for executives. In L. Werin and H. Wijkander (Eds.) *Contract economics*, Blackwell, Cambridge (MA), pp. 181-217.

Rosen, S. 1986, Prizes and incentives in elimination tournaments, *American Economic Review*, 76, 701-715.

Schaefer, S. 1998. The dependence of pay-performance sensitivity on the size of the firm. *Review of Economics and Statistics*, 80, pp. 436-443.

Shorrocks, A.F. 1984. Inequality decomposition by population subgroups. *Econometrica*, 52(6), pp. 1369-1386.

Skott, P. 2005. Fairness as a source of hysteresis in employment and relative wages, *Journal of Economic Behavior & Organization*, 57, pp. 305-331.

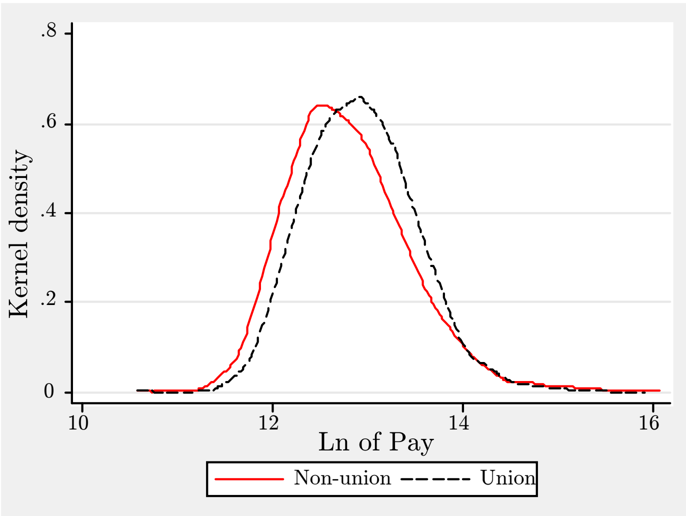
Tervio, M. 2008. The difference that CEOs make: An assignment model approach. *American Economic Review*, 98(3), pp. 642-668.

Toyoda, T. 1980. Decomposability of inequality measures. *Economic Studies Quarterly*, 31, pp. 207-216.

Zheng, B., Cushing, B.J. 2001. Statistical inference for testing inequality indices with dependent samples. *Journal of Econometrics*, 101, pp. 315-335.

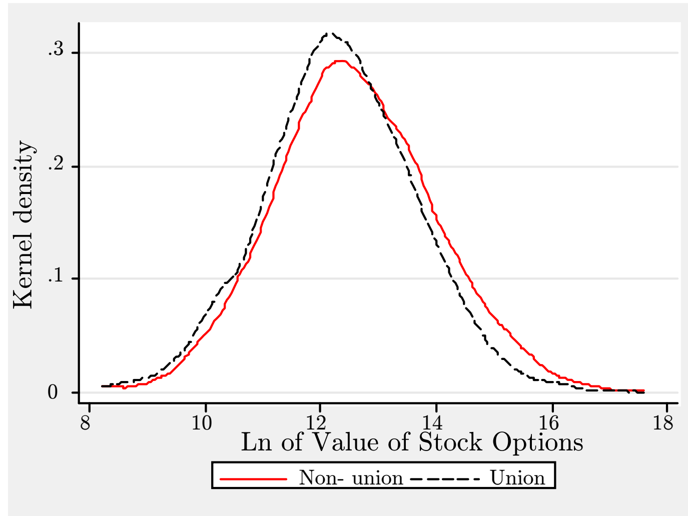
Figure 1. Distribution of cash pay and stock options for Managers and CEOs across firms with and without union presence.

(i) Cash pay (Managers)



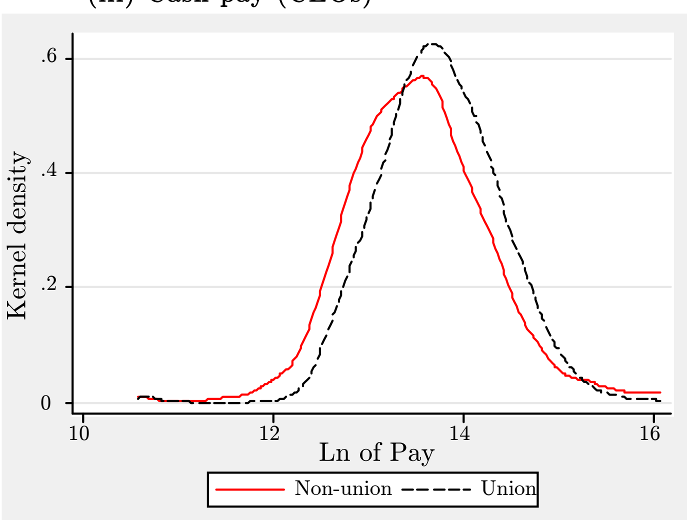
*Two-sample mean-comparison t-test (t-statistic): -16.30****

(ii) Stock options (Managers)



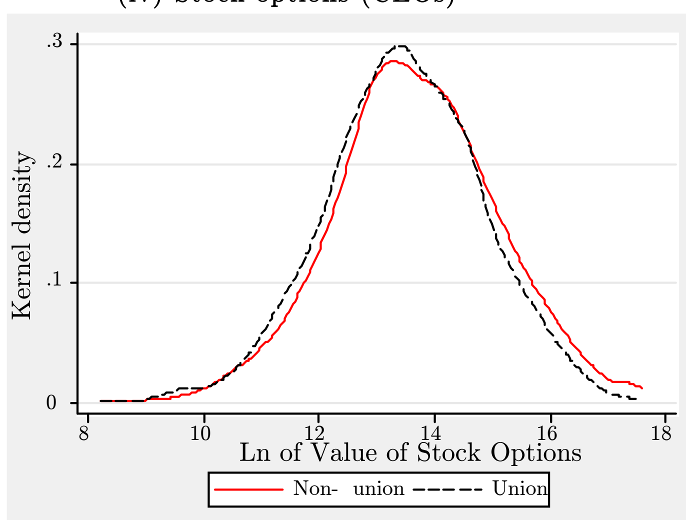
*Two-sample mean-comparison t-test (t-statistic): 13.92****

(iii) Cash pay (CEOs)



*Two-sample mean-comparison t-test (t-statistic): -12.04****

(iv) Stock options (CEOs)

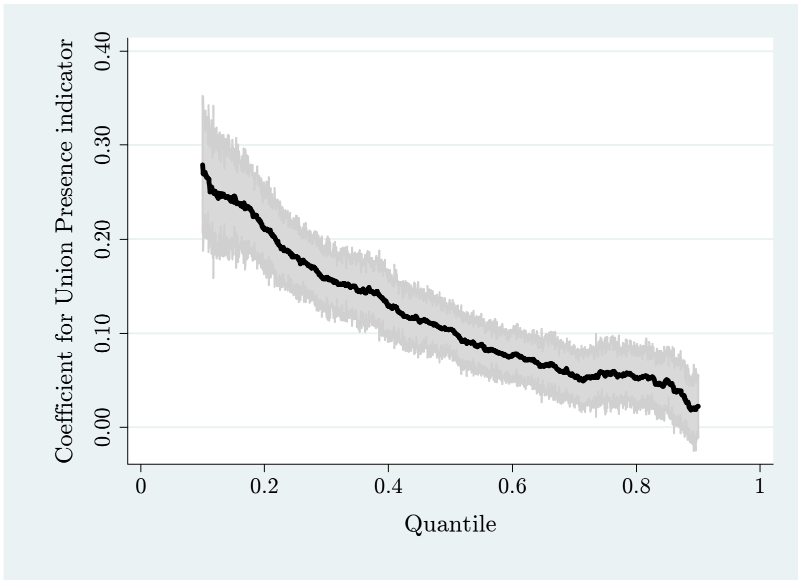


*Two-sample mean-comparison t-test (t-statistic): 5.16****

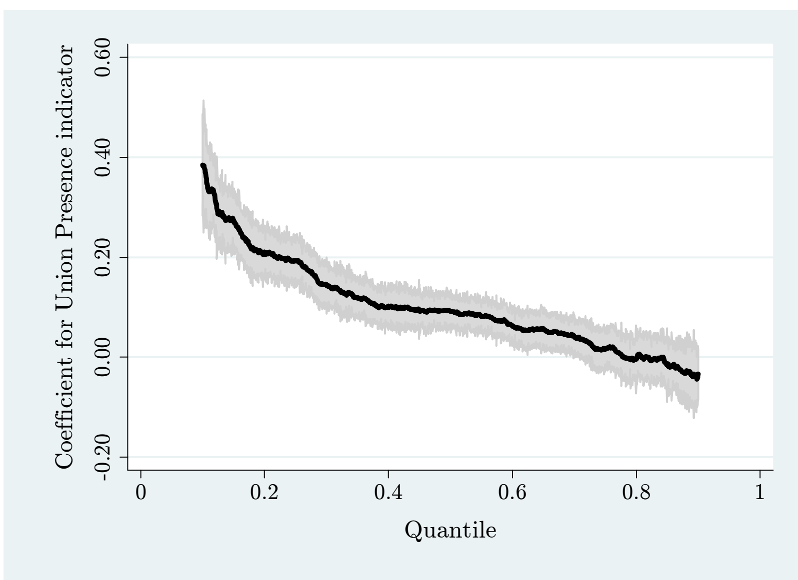
Notes: The figure depicts the unconditional distributions for the two main compensation components, cash pay (i.e., the sum of salary and bonus) and stock options, for Managers and CEOs across firms with and without union presence. The density distributions are estimated with the Epanechnikov kernel estimator. The horizontal axes present the natural logarithm of cash pay (i & iii) and total compensation (ii & iv), while the vertical axes present the kernel density.

FIGURE 2. DIFFERENTIAL EFFECT OF UNION PRESENCE ACROSS THE CONDITIONAL DISTRIBUTION OF DISTANCE MEASURES.

A. Effect of union presence across the conditional distribution of '*Distance in Cash-Pay*' between the CEO and the top-4 Managers

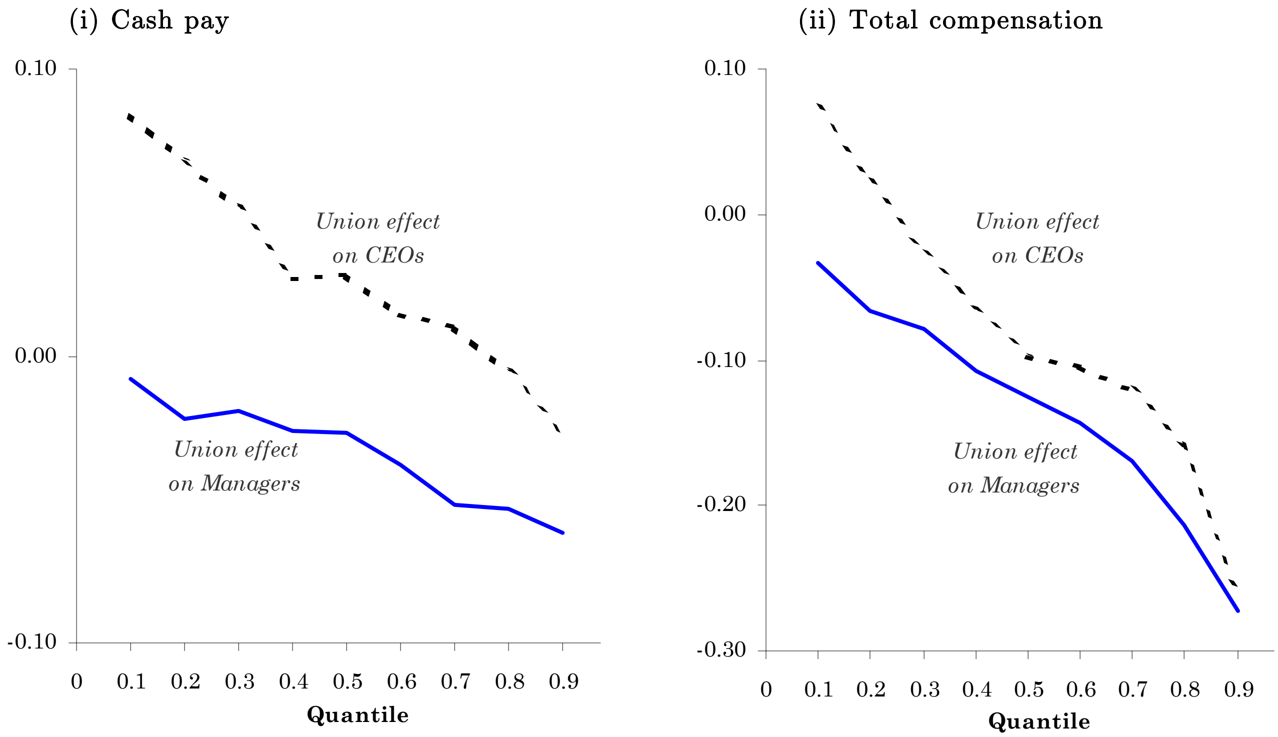


B. Effect of union presence across the conditional distribution of '*Distance in Total Compensation*' between the CEO and the top-4 Managers



NOTES – The figures report the effect of union presence on the '*Distance in Cash Pay*' and '*Distance in Total Compensation*', respectively, utilizing the quantile regression specifications in Table 4 (Columns 1 and 4). The shaded region is the 95% confidence band using bootstrapped standard errors (1000 replications).

FIGURE 3. DIFFERENTIAL EFFECT OF UNION PRESENCE ON CEOS' AND MANAGERS' COMPENSATION



F-statistic from Wald tests for H_0: $\alpha_1 = \alpha_2$

Quantile	.1	.2	.3	.4	.5	.6	.7	.8	.9
F-statistic	*24.09****	*36.18****	*26.93****	*16.06****	*17.14****	*16.95****	*18.85****	*8.44****	*3.40**

F-statistic from Wald tests for H_0: $\beta_1 = \beta_2$

Quantile	.1	.2	.3	.4	.5	.6	.7	.8	.9
F-statistic	*18.62****	*16.62****	*5.95***	*5.94***	*1.93*	*3.13**	*4.11***	*4.73***	*0.05*

NOTES: The figure depicts the differential effect of union presence on the compensation of CEOs and senior managers. We estimate a quantile regression for the pooled sample of 46,465 executives (i.e., CEO and senior managers) for the natural logarithm of cash pay (or, alternatively, total compensation) against the two interactions of union presence with the *CEO* and *Manager* dummies, and a number of control variables, namely dummy variables for *CEO* and *Chairman* positions, the natural logarithm of firm asset value, return on assets, return on shareholders, as well as 2-digit SIC industry and year effects. The standard errors are bootstrapped using 1000 replications. Quantile regression coefficients for the two interactions dummies ('Union×CEO' and 'Union×Manager') from nine quantile regressions (starting at quantile value 0.10 with step 0.10) are plotted for both CEOs (dotted line) and senior managers (solid line). Wald tests are performed in each quantile testing the equality of the coefficients for the two interactions dummies ('Union×CEO' and 'Union×Manager'). Finally, in the Wald tests, asterisks denote significance at 1 percent (***), 5 percent (**), and 10 percent (*) levels.

FIGURE 4. DISTRIBUTION OF PRIOR FIRM-SPECIFIC TENURE FOR INSIDER CEOS.

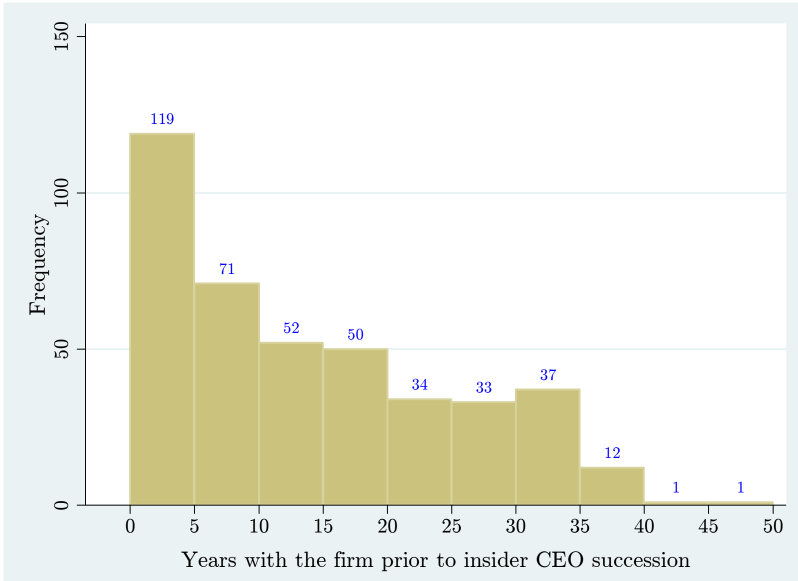


FIGURE 5. DISTRIBUTION OF PRIOR FIRM-SPECIFIC TENURE FOR INSIDER CEOS.

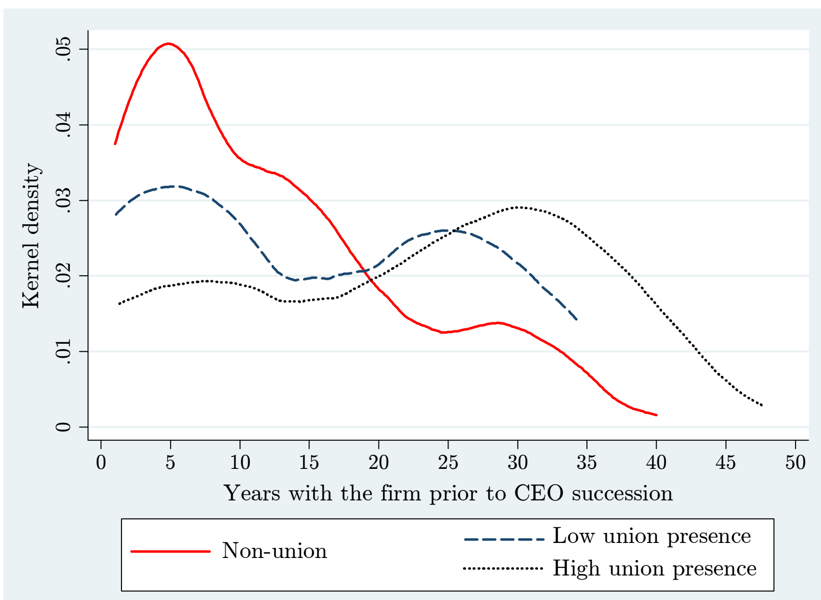


Table 1. Union Presence in the Sample by Major Sector (%), 1992-2001

Major sectors	Firms: *Union presence*	Firms: *No Union presence*	Firm-year observations: *Union presence*	Firm-year observations: *No Union presence*
Mining & Construction	16.7	83.3	16.4	83.6
Manufacturing	34.5	65.5	39.2	60.8
Transportation, Com. & Utilities	25.9	74.1	28.0	72.0
Wholesale & Retail Trade	19.1	80.9	22.0	78.0
Finance, Insurance & Real Estate	1.4	98.6	2.9	97.1
Services	10.5	89.5	13.6	86.4
Other	16.7	83.3	7.7	92.3
Number of observations	468	1602	12400	34065
(%)	(22.6)	(77.4)	(26.7)	(73.3)

Notes: The cell values represent percentage of firms in the sample with and without union presence by sector. Sector categorization includes Mining & Construction (SIC 1-19), Manufacturing (SIC 20-39), Transportation, Communications and Utilities (SIC 40-49), Trade (SIC 50-59), Finance, Insurance and Real Estate (SIC 60-69) and Services (i.e. SIC 70-89). The few remaining industries are consolidated under the 'Other' category. The sample contains an unbalanced panel with 46,465 observations from 2070 firms. Moreover, the frequency of firms in the sample period 1992-2001 is similar between unionized and non-unionized firms (3.6 versus 3.3 observations, on average, for each firm). The same holds for the executives (both Managers and CEOs) in the unionized and non- unionized firms, averaging 3.1 and 2.7 observations per executive, respectively, during the sample period.

Table 2. Descriptive Statistics and Definitions of Variables

Variables	Definition	Mean	Std.Dev.
CEO Cash pay	Cash pay (salary & bonus) – *in $million*	1.07	1.61
CEO Stock Options	Black-Scholes value of CEO's stock options awarded – *in $million*	1.82	8.65
CEO Total Compensation	The sum of salary, bonus, benefits, B-S value of stock options, restricted grants and LTIPs – *in $million*	3.44	11.10
Managerial Cash pay	Cash pay (salary & bonus) – *in $million*	0.48	0.58
Managerial Stock Options	Black-Scholes value of CEO's stock options awarded – *in $million*	0.56	1.90
Managerial Total Compensation	The sum of salary, bonus, benefits, B-S value of stock options, restricted grants and LTIPs – *in $million*	1.23	2.91
Union Presence	Dummy variable taking value 1, if the firm has at least one unionized establishment, and 0 otherwise.	0.26	0.44
High_Union	Dummy variable taking value 1, if the firm has three or more unionized establishments, and 0 otherwise.	0.12	0.32
Founder	Dummy variable taking value 1 if the CEO is the firm's sole founder or co-founder, and 0 otherwise.	0.17	0.37
Relative of Founder	Dummy variable taking value 1 if the CEO is a direct or distant relative or any other family member of the firm's sole founder or co-founder (e.g., son, son-in-law, nephew, grandson, great-grandson), and 0 otherwise.	0.06	0.23
Firm size (Assets)	Firm size, in terms of total assets – *in $billion*	7.89	31.40
Return on Assets	ROA is defined as net income before extraordinary items and discontinued operations divided by total assets – *in percentage form*	3.78	13.49
Return to Shareholders	RET is defined as total return to shareholders, including the monthly reinvestment of dividends – *in percentage form*	22.86	72.67
Industry Homogeneity	Partial correlation between industry returns (2-digit SIC) and common stock returns of the firms in that industry.	0.26	0.06
Firm Age	Number of years that the firm has been publicly listed.	26.11	18.92
Rust-belt	Dummy variable taking value 1 if the firm has its headquarters in a rust-belt state (namely IL, IN, MI, OH, PA, and WV), and 0 otherwise	0.20	0.40
Bible-belt	Dummy variable taking value 1 if the firm has its headquarters in a bible-belt state (namely AL, AR, FL, GA, KS, KY, LA, MO, MS, NC, OK, SC, TN, TX and VA), and 0 otherwise	0.28	0.45
Sun-belt	Dummy variable taking value 1 if the firm has its headquarters in a sun-belt state (namely CA, OR, and WA), and 0 otherwise.	0.17	0.38

Notes: Definitions and descriptive statistics for variables used in the analysis of executive compensation. The sample contains an unbalanced panel of 46465 executive-year observations in 2070 companies in the 1992-2001 period. For each firm-year observation in the sample we have compensation data for the CEO and the four-highest paid managers, resulting in 9293 CEO-year observations and 37172 manager-year observations. Summary statistics for *Founder* and *Relative of Founder* refer to a sub-sample of 8914 CEO-year observations for which we were able to collect reliable information. Data was obtained from merging the Bureau of National Affairs' Labor database and Standard & Poors' Execucomp, using additional information from Hoovers Online, Dun & Bradstreet's Online, Harris Info-Source on firm establishments, firms' annual statements, firms' official websites and S&P Register of Corporations, Directors and Executives. Regarding founder status, detailed information was manually collected on executive biography and corporate history from firms' annual reports (especially 10-K forms), Dun & Bradstreet's Million Dollar Database, Hoovers Online, the Standard and Poor's Register of Corporations, Directors, and Executives, company press releases, and other official sources (such as litigation documents). In some cases, additional information was retrieved from the companies' internet sites and the business press. All level variables have been adjusted for inflation and are stated in 1992 dollars.

TABLE 3. UNION PRESENCE EFFECT ON MANAGERIAL INCENTIVES

	Δ ln(Cash pay)			
	OLS regression estimates	IV regression estimates	OLS regression estimates	IV regression estimates
Independent Variables	[1]	[2]	[3]	[4]
Intercept	0.045***	0.046***	0.030***	0.024**
	(15.21)	(9.15)	(5.68)	(2.15)
Δ ln(Shareholder Value)	0.140***	0.146***	0.207***	0.245***
	(16.26)	(10.19)	(9.49)	(4.86)
Δ ln(Shareholder Value) × Union pres.	0.027	-0.006	0.007	-0.064
	(1.28)	(0.09)	(0.36)	(0.71)
Union presence	-0.005	-0.008	-0.001	0.011
	(1.08)	(0.57)	(0.22)	(0.54)
Δ ln(Shareholder Value) × cdf(Risk)	—	—	-0.090***	-0.126***
			(3.15)	(2.42)
cdf (Risk)	—	—	0.022**	0.027**
			(2.52)	(2.28)
Observations	20208	20208	20208	20208
R^2 (Centered R^2)	0.064	(0.063)	0.066	(0.023)
Hansen J-statistic (p-value)	—	2.391 (0.792)	—	3.625 (0.604)

NOTES: The absolute values of robust *t*-statistics and z-statistics with firm clustering appear in parentheses below each coefficient estimate from OLS estimation and IV (2SLS) estimation, respectively. Cash pay is the sum of salary and bonus. Notably, the dependent variable is windsorized at the 1% level in order to reduce the influence of extreme outliers or erroneous data entries in Execucomp. Both IV estimations (columns 2 and 4) pass the tests for overidentification and for weak instruments at any conventional level of confidence. Also, both IV estimations allow for two endogenous variables: *union presence* and the interaction of *union presence* with *Δln(Shareholder Value)*; as a result, we increase the number of valid instruments interacting the three HQ location dummies with the Δln(Shareholder Value). Asterisks denote significance at 1 percent (***), 5 percent (**), and 10 percent (*) levels.

TABLE 4. THE EFFECT OF UNION PRESENCE ON THE COMPENSATION DISTANCE BETWEEN THE CEO AND THE TOP-4 MANAGERS

Independent Variables	*(Distance)_Cash pay* Quantile reg. estimates [1]	*(Distance)_Cash pay* Quantile reg. estimates [2]	*(Distance)_Cash pay* IV regression estimates [3]	*(Distance)_Total compensation* Quantile reg. estimates [4]	*(Distance)_Total compensation* Quantile reg. estimates [5]	*(Distance)_Total compensation* IV regression estimates [6]
Intercept	1.716***	1.783***	3.057***	0.685***	0.712***	2.243***
	(9.99)	(9.75)	(9.22)	(4.28)	(4.39)	(6.05)
Union Presence	0.104***	0.077***	0.485***	0.091***	0.067***	0.452***
	(7.94)	(4.97)	(5.31)	(5.40)	(3.01)	(3.12)
Union presence × High_Union	—	0.059**	—	—	0.057**	—
		(2.94)			(2.10)	
Firm size (*lnAssets*)	0.085***	0.083***	0.081***	0.049***	0.048***	0.095***
	(14.66)	(14.12)	(6.82)	(8.16)	(8.03)	(5.68)
Return on Assets	0.006***	0.006***	0.004***	0.002**	0.002**	0.003*
	(7.46)	(7.38)	(4.52)	(2.35)	(2.29)	(1.80)
Return to Shareholders	0.001***	0.001***	0.001***	0.001**	0.001**	0.001***
	(2.87)	(3.00)	(5.18)	(2.45)	(2.27)	(3.06)
CEO is Chairman	0.202***	0.201***	0.195***	0.205***	0.198***	0.179***
	(12.42)	(12.18)	(6.66)	(10.70)	(10.05)	(3.96)
CEO Tenure	0.006***	0.006***	0.008***	0.002**	0.002*	0.003
	(5.32)	(5.08)	(3.59)	(2.07)	(2.09)	(0.91)
CEO Ownership	-0.016***	-0.016***	-0.020***	-0.028***	-0.027***	-0.046***
	(8.25)	(8.02)	(6.29)	(9.23)	(9.60)	(6.91)
Median Cash Pay of Managers (*ln*)	0.713***	0.711***	0.604***	—	—	—
	(37.51)	(35.40)	(17.26)			
Median Total Comp of Managers (*ln*)	—	—	—	0.872***	0.872***	0.664***
				(65.35)	(62.29)	(20.55)
Year effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	9293	9293	9293	9293	9293	9293
Pseudo-R^2 (Centered R^2)	0.309	0.310	(0.401)	0.301	0.302	(0.304)

NOTES: The sample consists of an unbalanced panel with 9293 firm-year observations for 2744 CEOs in 2070 firms during the 1992-2001 period. The values of the dependent variables are windsorized at the 5% level in order to reduce the influence of extreme outliers. The absolute values of *t*-statistics (in Columns 1, 2, 4, 5) and robust *z*-statistics corrected for clustering at the firm-level (in Columns 3 and 6) appear in parentheses below each coefficient estimate. Also, standard errors in quantile regressions are bootstrapped to address potential heteroskedasticity (1000 bootstrap replications), while standard errors in IV (2SLS) regressions are clustered at the firm-level. The Hansen J-statistic for columns [3] and [6] is 2.409 (p-value: 0.492) and 0.997 (p-value: 0.802), respectively. Asterisks denote significance at 1 percent (***), 5 percent (**), and 10 percent (*) levels.

TABLE 5. CEO TURNOVER - DESCRIPTIVE STATISTICS (1992-2001 PERIOD)

	Non-unionized firms	Unionized firms	Total
CEO turnovers *(#)*	530	188	718
% CEO turnovers *(scaled by firm population in the sample)* [37]	30.1%	37.2%	31.8%
Outsider CEOs *(<1 year prior tenure in the firm)*	230 (43.4%)	78 (41.4%)	308 (42.9%)
Insider CEOs *(≥1 year prior tenure in the firm)*	300 (56.6%)	110 (58.5%)	410 (57.1%)
Mean prior tenure of Insider CEOs at succession time	12.2	18.8	14.0
Median prior tenure of Insider CEOs at succession time	10.0	19.9	11.6

[37] Since this calculation is at the firm-level, we do not double-count any firms that had more than one turnover during the sample period.

Table 6. Union presence effect on the likelihood of insider CEO succession

Independent Variables	$\Pr(\textit{CEO is insider}=1 \mid_{xb})$				Years of firm-specific tenure prior to becoming CEO	
	Probit estimates	Bivatiate Probit estimates	Probit estimates	Bivatiate Probit estimates	Tobit estimates	Tobit estimates
	[1]	[2]	[3]	[4]	[5]	[6]
Union presence	-0.019	0.052**	-0.062	0.043^{+}	2.986**	-0.027
	(0.46)	(2.17)	(1.11)	(1.55)	(2.21)	(0.02)
Union presence × High_Union	—	—	0.080	0.013	—	5.601**
			(1.14)	(0.99)		(2.38)
Relative of the founder	0.340***	0.102***	0.339***	0.101***	9.854***	9.677***
	(2.83)	(3.53)	(2.82)	(3.54)	(4.63)	(4.61)
Firm size (*lnAssets*)	0.044***	0.007***	0.043***	0.007***	2.235***	2.145***
	(4.23)	(2.61)	(4.10)	(2.60)	(7.96)	(7.69)
Industry Homogeneity	0.614*	0.141*	0.640*	0.146*	9.760	11.605
	(1.83)	(1.80)	(1.91)	(1.85)	(1.06)	(1.26)
Observations	718	718	718	718	718	718

Notes: The sample of CEO turnovers consists of 718 unique CEO-firm observations during the 1992-2001 period. In Columns 1-4 the dependent variable is the dummy variable '*CEO is Insider*' that indicates whether the executive spent at least one year in the firm prior to his/her appointment to the CEO position. In Columns 5-6, the dependent variable is the '*Years of firm-specific tenure*' indicating the number of years the executive spent in the firm prior to his/her appointment to the CEO position. Coefficients in Columns 1-4 are marginal effects on the probability of the CEO being an insider (i.e., the firm having an internal succession for the CEO position). Marginal effects for binary independent variables denote the discrete change in probability for the new CEO being an insider as the binary variable changes from 0 to 1. All estimations were run with a constant term. The absolute values of z-statistics (Col. 1-4) and t-statistics (Col. 5-6) appear in parentheses. All standard errors are robust with firm-clustering. In Columns 2 and 4, the first level bivariate probit equation in that *UnionPresence* is a function of *Rust-Belt*, *Bible-Belt*, *Sun-Belt* and *FirmAge*. Asterisks denote significance at 1 percent (***), 5 percent (**), 10 percent (*), and 15 percent ($^{+}$) levels.

TABLE 7. DISTRIBUTION OF MANAGERIAL COMPENSATION WITHIN MATCHED SAMPLES OF UNIONIZED AND NON-UNIONIZED FIRMS

PANEL A: Inequality indices for managers' compensation in unionized and non-unionized firms

	Managers in unionized firms	*Managers in non-unionized firms*	*Difference*	$\lvert t$ - statistic$\rvert$
Inequality of Cash pay (W)				
Gini-coefficient	0.323	0.347	-0.024	2.440 **
Theil entropy measure	0.185	0.235	-0.050	2.011 **
Standard deviation of logs	0.310	0.351	-0.041	3.278 ***
Inequality of Stock option compensation (S)				
Gini-coefficient	0.660	0.631	0.029	1.725 *
Theil entropy measure	0.914	0.819	0.095	1.041
Standard deviation of logs	1.883	1.599	0.284	3.428 ***
Inequality of Total compensation (Y)				
Gini-coefficient	0.490	0.497	-0.007	0.469
Theil entropy measure	0.479	0.495	-0.016	0.332
Standard deviation of logs	0.678	0.725	-0.047	1.899 *
Number of Observations	3772	3772		

PANEL B: Inequality of managers' compensation between matched unionized and non-unionized firms

Cash pay (W)	H_0: $(ln W_{union}) = (ln W_{non})$	
	Mean (Median) % difference:	-2.72% (-1.95%)
	Mean comparison	$\lvert t$ - statistic$\rvert$ = 2.054 **
	Median comparison	$\lvert\chi^2$ - statistic$\rvert$ = 1.783
Stock option compensation (S)	H_0: $(ln S_{union}) = (ln S_{non})$	
	Mean (Median) % difference:	-24.85% (-25.71%)
	Mean comparison	$\lvert t$ - statistic$\rvert$ = 7.003 ***
	Median comparison	$\lvert\chi^2$ - statistic$\rvert$ = 35.493 ***
Total compensation (Y)	H_0: $(ln Y_{union}) = (ln Y_{non})$	
	Mean (Median) % difference:	-7.47% (-9.81%)
	Mean comparison	$\lvert t$ - statistic$\rvert$ = 3.874 ***
	Median comparison	$\lvert\chi^2$ - statistic$\rvert$ = 15.686 ***

NOTES: The tables present and compare the level of inequality in cash pay, stock options and total compensation among the top four (non-CEO) managers for two matched samples. The overall sample contains 7544 manager-year observations from 1886 firms (943 unionized firms and 943 non-unionized firms). The matching process is described in detail in Section 4.3. In *Panel A*, the second and third columns present the three inequality indices for unionized and non-unionized firms, respectively. The fourth column displays the difference in the indices between the two samples, while the fifth column offers the (absolute) t-statistic from an index comparison between the two samples. Bootstrapped standard errors are obtained for each of the three indices. In *Panel B*, the level (rather than the distribution) of compensation is examined by comparing mean and median compensation (in natural logarithm format) of the managers in the matched samples with the use of t-test with unequal variances and nonparametric k-sample test, respectively. Asterisks denote that the difference is significant at 1 percent (***), 5 percent (**), and 10 percent (*), respectively.

Appendix 1

Figure A. Distribution of "*Distance in Cash-Pay*" between the CEO and the top 4-executives (Transformed values)

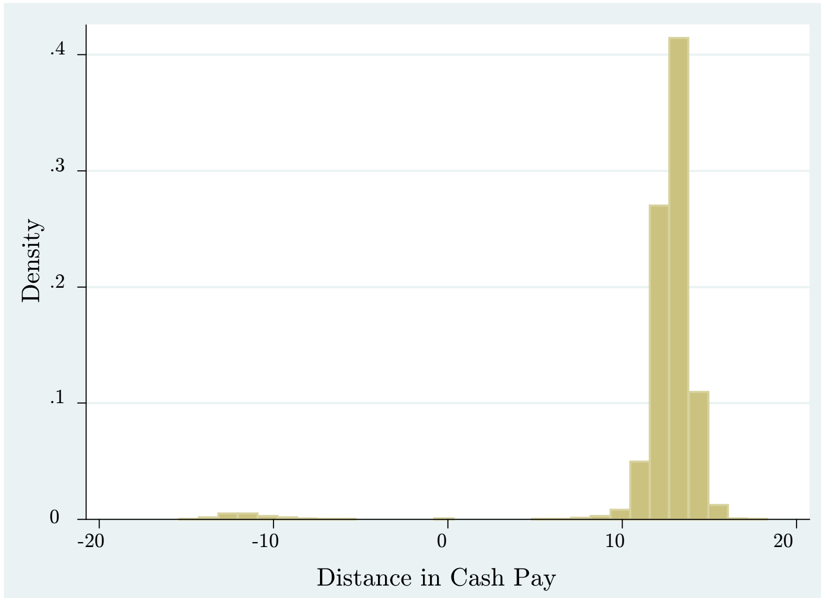
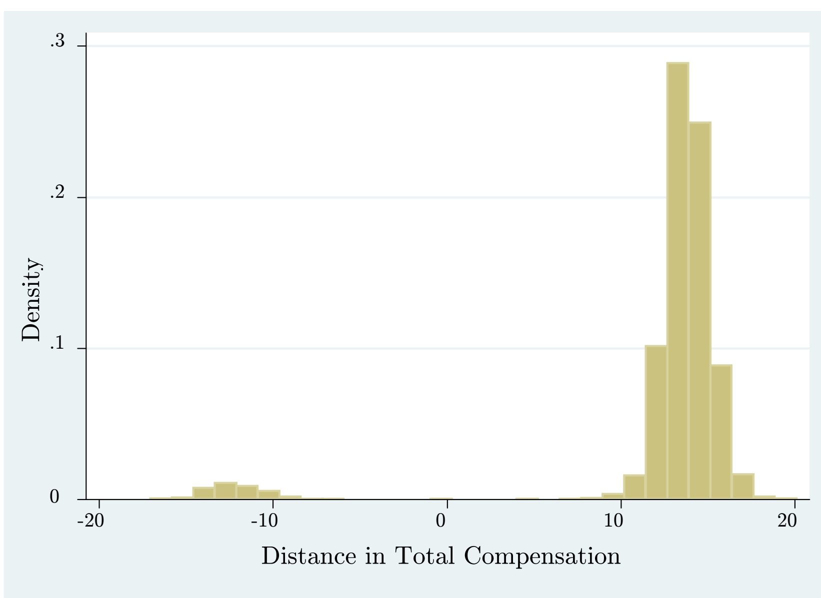


Figure B. Distribution of "*Distance in Total Compensation*" between the CEO and the top 4-executives (Transformed values)



Notes: Due to the presence of negative outliers, for the transformation of distance values we utilize the inverse hyperbolic sine function (Burbidge et al., 1988) that not only yields very similar results to those of the logarithmic function for positive numbers, but also transforms negative numbers.