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What Do Unions Do to Executive Compensation?

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Abstract

We estimate the relation between union presence and executive compensation using a unique panel of executives in publicly listed US firms during the period 1992-2001. We find evidence that union presence is associated with lower levels of total executive compensation. We find this union effect to be primarily the result of substantially lower stock option awards, and to a lesser extent due to lower cash pay. Moreover, the negative relation between unionization and executive remuneration becomes larger at the higher end of the conditional distribution of executive remuneration. We also find that the elasticity of cash pay to financial performance is similar across unionized and non-unionized firms, and that union presence is associated with a more compressed intra-firm and inter-firm executive compensation structure.

JEL Classifications: G30, J33, J51, M52
Keywords: Unions, executive compensation, Implicit regulation.

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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

In their seminal paper on CEO compensation, Jensen and Murphy (1990) posited that political and regulatory constraints in publicly listed firms would work to truncate the upper tail of executive remuneration, resulting in lower overall levels of remuneration and reduced pay-performance sensitivities. They referred to this as the *implicit regulation* effect and identified government legislation and the presence of unions as obvious examples of such institutional constraints. Several empirical studies have since confirmed Jensen and Murphy's predictions for public firms operating in regulated industries, where government oversight and disclosure rules ensure that executive remuneration remains a highly visible and contentious subject.[2] Similar research, however, on the possible constraining effects of unions is scarce. This is surprising since anecdotal accounts of union opposition to executive compensation abound (e.g., against Donald J. Carty, the former CEO at American Airlines), while at the same time there could be a relation between the dramatic growth in executive compensation and the substantial decline in private sector union density (Fortin and Lemieux, 1997). Moreover, one would have assumed

[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] An early test of the implicit regulation hypothesis can be found in Joskow et al. (1993), where CEOs in regulated industries in the United States receive substantially lower levels of pay. Joskow et al. (1993, 1996) also find that even within the already regulated electric utility industry, the level of CEO cash pay is negatively related to the degree of regulation. In terms of the pay-performance effects, Crawford et al. (1995) and Hubbard and Palia (1995) find increased CEO pay-for-performance sensitivity after deregulation in the banking industry, thus providing empirical support for the limitations on CEO compensation imposed by regulatory constraints.

that there would be a parallel body of work examining the impact of unions on executive compensation on a non-anecdotal manner, given organized labor's interest in monitoring executive pay and in highlighting alleged executive excesses.

One obvious reason for the scarcity of published research in this area is the difficulty in identifying clear causal links between union presence, which is found at an establishment level, and executive remuneration, which is set at the firm level. The other more important obstacle, up to now, has been the lack of a sizeable dataset matching union information with executive compensation data.[3] This paper fills this gap in the literature by using a unique sample of matched pairs of unionized/non-unionized publicly listed US firms, during the 1992-2001 period, in which the presence (or absence) of a union at the firm level is identified with appropriate workplace data (i.e., NLRB elections, FMCS contract expirations). Moreover, we offer evidence for both the CEO and other top managers (non-CEOs).

Our results suggest that union presence is significantly associated with executive compensation. Specifically, we find that union firms display lower levels of total compensation for their executives (i.e., *total compensation* being the sum of salary, bonus, benefits, restricted grants, long term incentive plans, and stock options) as compared to non-union firms. We find this union effect to be primarily the result of substantially lower stock option awards, and to a lesser extent due to lower cash pay. The constraining effect of unionization on remuneration is also found to be stronger the higher-up in the distribution ladder one moves. However, we find that the elasticity of executive cash pay, with respect to firm performance, does not differ significantly between union and non-union firms. These results provide mixed support for the

[3] The difficulty of matching (for the purposes of this paper) publicly listed firms with establishment-level data on unions is supported by DiNardo et al.'s (2000: p.18) statement that "matching the universe of [union workplace] contracts to the universe of firms is virtually impossible".

implicit regulation hypothesis, therefore distinguishing union presence from regulatory oversight.

Overall, this study is part of a growing research trend that examines the linkages between labor and finance at the firm level. In this sense, it contributes in several important ways to the empirical literature on incentives. *First*, our results empirically confirm the noted union aversion to stock-option incentive systems and indicate that unions possibly extend their influence throughout the corporate hierarchy, beyond those workers covered by union-negotiated wage scales. This is consistent with union preferences for remuneration compression within and across positions (Hirsch, 2008). *Second*, 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). *Third*, 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 executives are competing for rent extraction.

[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 were unionized during our sample period.

The rest of the paper is organized as follows: Section 2 reviews earlier studies on the topic, while Section 3 outlines the possible channels of union influence upon firm corporate governance, financial markets and ultimately on executive compensation. Section 4 describes how the dataset was constructed and provides an overview of the data. Section 5 analyses the results and discusses their relation to the analytical framework presented in section 3. Section 6 concludes.

2. Literature Review

Although there is a substantial body of research examining the effect of unions on the workers' wages (Blanchflower and Bryon, 2004), as noted in the introduction, only two published studies have estimated the effect of union presence within US firms on the level of CEO compensation. DiNardo et al. (2000) examined the effects of unionization in US firms for the periods 1971-74, 1975-78 and 1979-82 on levels of CEO cash pay, but not for total compensation (i.e., they exclude the value of stock options).[6] The unionization data is for 1977 and 1987. They limited their study to unionized firms only and evaluated how the strength of unionization, measured in terms of union density at the firm level, affects CEO cash pay. DiNardo et al., (2000) initially found that greater levels of unionization were negatively and significantly associated with CEO cash pay. However, after controlling for industry and firm effects in alternative specifications, this relation became insignificant and in some cases positive. They characterized these particular findings as 'extreme' and attributed them to the inclusion of a later period (i.e., 1979-1982) in their data, although they admit that they did not have a satisfactory explanation as to why this period should differ from the rest of their sample.

[6] A shorter version of DiNardo et al. (2000) appeared as a working paper at NBER in 1997 (DiNardo et al., 1997).

More recently, Banning and Chiles (2007) examine the relationship between union presence and CEO compensation by exploiting a cross-sectional sample of 170 firms from the Fortune 500 list in 1996. After collecting information on compensation and unionization from the companies' SEC fillings, they find that union presence, as well as the unionization rate at the firm-level, are negatively related to both the level of total CEO compensation and the proportion of CEO compensation that is contingent on firm performance. However, the study's cross-sectional nature and empirical methodology do not provide a definitive view on the issue. In particular, SEC fillings do not necessarily reveal unionization status.[7] Also, the unionization rate at the firm-level is difficult to measure since employment figures in SEC fillings are typically based on full-time employees thus not representing the actual number of workers at the firm (e.g., temporary/seasonal workers). Furthermore, the authors utilize an OLS standardized regression framework even though their key dependent variable, the proportion of performance-contingent CEO compensation, is fractional with a sizeable cluster of zeros. Moreover, by adjusting the dependent variables at the overly narrow 4-digit SIC level, the authors bias the adjusted values for 4-digit industries with very few observations, and overlook the fact that large firms are active in several other 4-digit SIC industries, which are also reported as secondary industry classifications in the firms' financial statements.

From a Canadian perspective, Singh and Agarwal (2002) examined the effect of union presence on CEO compensation with a cross-sectional sample of 86 Canadian mining and manufacturing firms listed in the Toronto Stock Exchange in 1996. After controlling for firm performance and size, they found that union presence is associated with higher CEO cash pay, but not associated with other compensation components (e.g., stock options) and total

[7] This is probably the reason why they find 17.6% of the firms to have union presence, compare to the 22.6% in our sample.

compensation. Nevertheless, the study's narrow industrial focus and the US-Canada differences in attitudes towards unions restrict any definitive inferences on the topic.

3. The Link between Union Presence and Executive Remuneration

There are two basic channels – one direct and the other indirect – by which unions are linked to executive compensation. The *direct channel* is based on the pressure that unions can bring to bare within the firm and its various establishments. This pressure, as noted by Jensen and Murphy (1990) is exercised primarily through the voicing of fairness concerns and the threat of industrial disruption at the workplace, either of which would be expected to occur if union members perceive executive salaries as excessive.[8] Union pressure is normally observed during formal collective bargaining negotiations, but it can also be channeled informally through local stewards, public awareness campaigns and even through the sponsorship of journalism and public policy research critical of alleged managerial excesses. 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). Union pressure for more constrained executive compensation can also be exercised through union-controlled pension funds, which are increasingly taking part in firms' annual shareholder meetings and actively urging corporate governance reforms. Less frequently, union

[8] The analytical justification for the direct channel is rooted in the work of Freeman and Medoff (1984), who in their seminal study of the role of organized labor in the US economy, argued that in addition to being a bargaining institution that raises wages for its members, unions can also act as an institutional counterbalance to managerial power, one that provides a structured means (i.e., a 'voice') for addressing employee concerns and securing fairness at work; in this case actualizing worker preferences for a more compressed compensation structure within the firm.

influence is gained through representation on the boards of directors, or other senior level committees.[9]

A further direct effect, put forward by Jensen and Murphy (1990), maintains that unions perceive high executive remuneration as a signal for the firm's financial health and employ it as a justification for increased wage demands in labor negotiations. This would naturally make senior executives and boards more cautious when negotiating remuneration packages. This is also why union presence, much like the public oversight faced by executives of firms operating within regulated industries, is predicted by Jensen and Murphy to truncate executive remuneration and make it less sensitive to firm performance.

The second channel by which unions and executive compensation are linked is *indirect* and relates to the adverse consequences that union presence can have on share price fluctuation, which, in turn, determines the value of the stock-related part of executive compensation. This particular channel is quite important given that stock options comprise a substantial part of top executive remuneration packages within publicly listed US companies. Unlike salary and bonus, which are fixed or accounting-based, stock-option compensation depends on financial performance measures. This means that if financial markets respond negatively towards union presence, then the incentive power of stock-related compensation for senior executives is reduced within unionized firms.

Studies in the United States uniformly find that there is a negative relationship between union activity in a firm and its market value. Abowd (1989) examines share price movements in response to labor cost changes and concludes that increases in labor costs due to collective bargaining reduce a firm's market value. Ruback and Zimmerman (1984) find that formal union

[9] The practice is common in Japan and Germany, and even present in Canada and the UK. However, it is still rare in the US and found mostly when unions participate in broad-based ESOP plans.

organizing within establishments owned by a parent firm significantly lowers that firm's equity value. They also find that the reduction in firm equity value due to petitions that lead to union wins in NLRB elections is almost three times bigger compared to petitions that lead to union losses in the year after the certification date. Historical research focusing on the interwar period also confirms the large and significantly negative effect of strikes and workplace disputes on stock values (DiNardo and Hallock, 2002).[10]

By and large, it appears that financial market antipathy toward union presence stems not so much from adverse productivity effects or lower R&D spending,[11] but rather from the expectation that union influence over a firm's governance leads to inflexibility and greater redistribution of rents toward workers. Unionized firms (or even firms that have been the targets of organizing drives) are perceived as having to devote considerable resources to dealing with unions, or in trying to counteract their spread within the organization.[12] These labor constraints result in unionized firms' reduced share price fluctuations, which could potentially necessitate a different dividend policy to signal future profitability to investors. This point is reflected in our sample of matched pairs of unionized and non-unionized firms (see Section 4.2), in which we find that they have an average (median) volatility of their stock returns over the previous sixty months of 0.32 (0.29) and 0.35 (0.32), respectively. Also, the median dividend yield (3-year

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[10] Union activity in one firm also seems to have negative spillover effects on the equity price of other firms in the same industry, presumably because union activity in one firm may increase the threat of unionization in others (a result also found in DiNardo and Hallock, 2002). Findings by Bronars and Deere (1994) reinforce this pattern and show that the petition for union representation in one firm has a substantial negative impact on the share price of other firms in the same 4-digit SIC industry group.

[11] The well-established negative impact of union presence on a firm's share price described above has not been met with equally straightforward results for other firm activities and outcomes such as R&D spending or productivity. For instance, Connolly et al. (1986) and Hirsch (1992) suggest that union power has a negative effect on R&D spending, thus impeding innovation and the long-term profitability of the firm. However, Bronars et al. (1994) find that such a relation becomes insignificant after controlling for industry effects or estimating with first-differences. In a similar fashion, the large body of work concerning union effects on firm level productivity has produced mixed results [see Hirsch (2007) for a detailed review].

[12] Empirical evidence on firms' cost of fighting labor organization is provided in Abowd and Farber (1990) and Freeman and Kleiner (1990).

average) is 1.72% for unionized firms and 1.13% for non-unionized firms.[13] Given the similarity of the remaining characteristics of executive stock options (i.e., non-tradable, 10-year maturity period, vesting period between 1-3 years, exercise price equal to the share price at the time of the issue), the indirect effect of union presence on executive remuneration materializes in the form of lower stock option value, due to the lower share price volatility and the higher dividend rate of unionized firms.

Overall, we expect to find that union presence in firms is associated with reduced levels of executive remuneration, particularly at the upper tail of the remuneration distribution. As suggested by the implicit regulation hypothesis, we also expect unionized firms to display lower pay-performance elasticity for their executives' cash pay compared to non-unionized firms. Finally, we hypothesize that union aversion to inequality will be associated with less dispersed executive compensation both within-firms and across-firms.

4. Sample Design

This section describes how we merge a firm-level dataset on executive compensation and an establishment-level dataset on union activity, in order to create a merged dataset containing compensation information for the CEO and the four highest-paid managers. It also explains our focus on matched pairs of union-nonunion firms, in order to mitigate endogeneity concerns.

4.1. Combining datasets

[13] S&P Execucomp offers these figures on volatility (*bs_volat*) and dividend rate (*bs_yield*), which are used in calculating Black-Scholes values for executive stock options. The difference in the mean/median for these two variables across unionized and non-unionized firms in our sample is statistically significant at the 1% level. Also, we observe zero dividend yield (3-year average) in 21% (31%) of the firm-year observations for unionized (non-unionized) firms in our sample of matched pairs. Moreover, the vast majority of executive stock options in our sample have 10-year maturity period and an exercise price which is equal to the share price at the time of the issue.

Data on executive compensation and firm characteristics are obtained from the Standard and Poors (S&P) Executive Compensation database (Execucomp).[14] 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 sales, return on assets and return to shareholders), we drop observations that correspond to less-than-annual compensation.[15] In the case of CEOs, we identify these observations using the information provided in Execucomp on CEO tenure. For non-CEO executives, 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.[16] 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-paid non-CEO executives (hereafter referred to as *Managers*).[17] 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.

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[14] 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.

[15] 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).

[16] 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.

[17] 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.

Executive remuneration is measured in terms of *cash pay* (i.e., the sum of salary and bonus), Black-Scholes value of *stock options* awarded each year, and finally *total compensation* (i.e., the sum of salary, bonus, benefits, LTIPS, restricted grants and stock options). Examining Figure 1, drawn from the final dataset, provides a revealing picture of executive compensation in the United States during the 1990s. The most striking features are the growth of total compensation, the constancy of cash pay levels, and the remuneration gap between managers and CEOs. The growth in total compensation can be attributed to the dramatic increase in stock options, both in terms of usage and magnitude. In addition, the sizeable difference between the mean and median value of total compensation provides evidence that some executives enjoyed exceptionally large stock option contracts, thus skewing the compensation distribution and attracting considerable public attention.

[Insert Figure 1 about here]

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.[18] 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. Then, by determining a discrete time dimension for each activity

[18] 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.

entry and sorting the establishments that corresponded to firms included in Execucomp, we were able to consolidate information on establishments belonging to the same firm. We match the firm-level data from Execucomp with the establishment-level data from BNA.

The time-consuming sorting and matching process between establishment-level BNA data and the firm-level Execucomp dataset was constructed and verified both mechanically (i.e., software assisted) and manually. This dual approach was undertaken in order to tackle matching problems arising from such things as variations in company names and abbreviations, and in order to identify establishments that had similar names with listed firms, but belonged to unrelated private companies. These functions were carried out with 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 indicator 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 indicator 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.

[Insert Table 1 about here]

4.2. Selecting matched pairs

Sectors with union presence might have different production technologies that require different type of executive talent or organizational structures. Hence, one has to differentiate between the pure union effect and the influence of other observable and unobservable factors correlated with union presence. We focus our analysis on matched pairs of unionized and non-unionized firms, in order to match the distribution of managerial skills between unionized and non-unionized firms. 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).[19] At the same time, union presence is unlikely to be correlated with individual characteristics of executives because, unlike lower-level workers, executives are not union members and are not covered by collective agreements. Moreover, the top five executives, which are the focus of this study, hold similar positions across firms (e.g., CEO, 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.[20] Lastly, unlike the private-public sector choice whose endogeneity has been well-established in the labor economics literature, all firms in the sample belong to the private sector thus being uniformly exposed to product market competition. Based on the

[19] 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.

[20] 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).

aforementioned, in contrast to lower-level employees, executives are unlikely to self-select into a publicly listed firm based on the criterion of union presence.

We construct the matched pairs of firms by utilizing the nearest-neighbor matching algorithm from Abadie et al. (2004). 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 quite similar 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 that have been widely found to influence compensation patterns; whether the CEO is also Chairman of the Board, the founder of the firm, or a relative of the founder. Two samples of executives in unionized and non-unionized firms are created from the pairs of matched firms.[21] Each of the two samples contains 773 firm-year observations (i.e., 773 CEOs and 3796 Managers). In sum, after matching on observables, we control for a large part of observed managerial quality and skill (Gabaix and Landier, 2008; Murphy, 1999), thus creating two comparable populations of executives that differ only in that they work either in a firm with or without union presence. A full listing of the aforementioned variables, along with means and standard deviations, can be found in Table 2. Notably, the matched pairs sample generally reflects the wider sample, with the exception of industry composition (see Tables 1 and 2). This is anticipated since it is difficult to find matched-pairs in industries with scarce union presence. We also exclude regulated utilities (SIC 49) from the matching process, in order to distinguish between union presence and regulatory oversight.

[Insert Table 2 about here]

[21] 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. Notably, we also remove utility firms (SIC 49) from the matching process, given that the executive compensation in those firms is subject to regulatory pressures.

5. Empirical Strategy and Results

We evaluate the association between union presence and senior management incentives in four stages as follows. First, we check for the relation between union presence and levels of executive compensation. We then evaluate whether that relation is affected by endogeneity and heterogeneity concerns. Based on our analysis in Section 3, it is expected that unions would be associated with lower levels of stock-option rewards, balanced by higher levels of cash pay compensation. We also examine the alignment of executive 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 unions, unionized firms should display lower pay-performance elasticity. Finally, we examine whether union presence is associated with a more compressed cross-firm distribution of executive compensation both within-firms and across-firms. This effectively tests whether the well-documented spillover properties of the union compression on employees' salaries extend to the top of the hierarchy.

5.1. Unions and level of executive compensation

We test whether union presence is significantly related with the levels of executive cash pay and total compensation for firm *i* at time *t* using the following specifications, separately for CEOs and Managers:

[1] $\ln(\text{Total Compensation})_{it} = \alpha_1 + \beta_1 UNION_i + \gamma_1' Z_{it} + \delta_1' Industry_i + \varepsilon_{2it}$

[2] $\ln(\text{Cash Pay})_{it} = \alpha_2 + \beta_2 UNION_i + \gamma_2' Z_{it} + \delta_2' Industry_i + \varepsilon_{2it}$

where α is the intercept term, $UNION_i$ is the indicator variable indicating union presence, Z_{it} is a vector of control variables on executive and firm characteristics, $Industry_i$ is a vector of industry controls (2-digit SIC), and ε specifies the error terms. In particular, the control variables involve stock ownership, board presence (Chairman or member of the board), firm size (natural logarithm of total sales), as well as accounting and financial performance.

We also perform estimations with stock-option compensation as the dependent variable. In terms of econometric methods, while in [1] and [2] we employ OLS regression, for stock-option compensation we follow Cragg's (1971) two-tier estimation, given that 22.5% of the CEOs and 22.9% of the Managers in our sample involve firm-year observations with no stock option component.[22] The specification for the first-tier reflects parameters influencing the adoption of stock option awards, namely stock ownership, firm riskiness (i.e., three-year standard deviation of monthly returns), industry effects and year effects. The specification for the second-tier contains the Z_{it} and $Industry_i$ vectors, discussed in [1] and [2].

Table 3 presents the findings for levels of total compensation and its two major components (i.e., cash pay and stock options) across union and non-union firms. We find strong evidence that union presence within a firm is significantly associated with all three compensation (dependent) variables. Specifically, union presence is significantly associated with 12 percent lower total compensation for both CEOs and Managers.[23] As hypothesized, the negative association with total compensation occurs because of the large negative impact of union presence on stock-option compensation, an impact which dominates the much smaller negative union impact on cash pay especially given that stock option awards generally constitute a substantial part of total compensation. Given that there is negligible difference in the binary incidence of stock option awards across unionized and non-unionized firms in our sample, we

[22] Cragg's (1971) two-tier model addresses the main shortcoming of the Tobit model that the factors affecting the two stages of the choice are identical. Moreover, the results are robust to alternative estimations that control for selection of firms choosing to offer stock options, such as the two-part estimation, Heckman estimation with the inverse Mills ratio, and Tobit.

[23] The effects are found with the following formula: $e^{\beta} - 1$, where β is the respective union presence estimate.

can attribute this significant difference in total compensation across firm types to differences in the size of the stock option awards.[24] Finally, the coefficients for the control variables have signs consistent with expectations when significant.

[Insert Table 3 about here]

5.2. Robustness checks

5.2.1 Heterogeneity of union estimate

As an added check of our results and as a straightforward test of whether unions have a greater effect on the upper tail of the executive compensation distribution, we employ a quantile regression on specifications [1] and [2]. By using this approach one can detect any heterogeneity in the union estimate over executive remuneration since regression parameters are allowed to vary across different points in the conditional distribution. As Koenker and Xiao (2002: p. 1583) suggest "[b]y supplementing least squares estimation of conditional mean functions with techniques for estimating a full family of conditional quantile functions, quantile regression is capable of providing a much more complete statistical analysis of the stochastic relationships among random variables."

Table 4 presents the estimates from the median regression on total compensation. The median effects, for both the CEOs and the Managers, are similar to the average impact reflected in our initial OLS estimation. However, the heterogeneity is pronounced when we illustrate the coefficient estimate for union presence across the conditional distribution. The results, graphed in Figures 2 and 3, suggest that the negative association between unions and executive remuneration is much stronger as we move up the compensation ladder. That is, unions appear to

[24] More specifically, 76% and 78% of the Managers in the matched unionized and non-unionized firms, respectively, are awarded stock options. The portion of CEOs receiving stock options in unionized and non-unionized firms is 77% and 78%, respectively.

significantly limit the upper tail of the compensation distribution much more than the lower tail, as Jensen and Murphy (1990) predict. For instance, union presence is associated with 15% (15%) lower level of total compensation for CEOs (Managers) at the 0.7 conditional quantile, as compared to only 9% (10%) at the 0.3 conditional quantile (see Figure 2). Both coefficient estimates are statistically significant at the 5% level. In Figure 3, the quantile regressions for cash pay reveal even more heterogeneity, given that the union presence's coefficient estimate for CEOs (Managers) is statistically insignificant at the 5% level in the first four (two) conditional quantiles.

Overall, the negative union effect with respect to executive remuneration strengthens as we move to the upper tail of the conditional remuneration distribution, thus providing evidence in support of Jensen and Murphy's (1990) assertion that unions are particularly concerned with constraining high-end remuneration packages that 'super-star' executives often can command.

[Insert Table 4 and Figures 2/3 about here]

5.2.2. Union presence as an endogenous variable

In this subsection we treat $UNION_i$ in equation [1] as an endogenous variable. One could argue that unions are more likely to organize firms with substantial rents, which in turn could also affect the choice of executive incentives. 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 and union approval (Herod, 1998; Holmes, 2006). To capture these effects we utilize two indicator variables: RIGHT-TO-WORK in order to identify companies that are headquartered in states with right-to-work laws, effectively proxying for anti-union bias (Davis and Huston, 1995; Moore, 1998), and RUST-BELT in order to identify companies that are headquartered in an area traditionally friendly to unions (Lopez, 2004).[25] 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. Exemplifying this point, during our sample period, the nonagricultural private-sector union density in the 'right-to-work' states and 'rust-belt' states was 5.1% and 14.7%, respectively, while in the rest of the United States it was 11.7%.[26]

Firm age is important since firm-wide union decertification events are rare (Nilsson, 1997), meaning that firms established before the union-busting era of the 1980s are more likely to be observed as having union presence during our sample period (Fiorito, 2007; Palley and LaJeunesse, 2007). Reflecting this change in unionization, during the 1965-1980 period, union density averaged 26%, while in the period 1981-2001 it averaged 15%. 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 CRSP, we choose the earlier date (Pastor and Veronesi, 2003).

[25] RIGHT-TO-WORK is a indicator variable taking value 1 if the firm has its headquarters in a state with right-to-work laws (namely AL, AZ, AR, FL, GA, ID, IA, KS, LA, MS, NE, NV, NC, ND, SC, SD, TN, TX, UT, VA and WY), and 0 otherwise. RUST-BELT is a indicator 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. Firm headquarters located in the remaining states serve as the omitted group. 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.

[26] Union density is averaged by state and weighted by employment. The source of union density data is Hirsch and Macpherson (2003).

Table 4, Columns 3 and 4, present the IV (2SLS) estimates for the levels of executives' total compensation. The fact that an endogenous variable (i.e., union presence) is binary is not influenced by the first-stage linear regression, which still produces consistent IV estimators (Heckman and Robb, 1985). We find strong evidence that union presence within a firm is negatively related to total compensation for both the CEO and the Managers. Notably, after controlling for potential endogeneity, the estimated coefficients are much larger than those from the OLS/Median regressions. This is typical of instrumented variables and could be attributed to unobserved differences between unionized and non-unionized firms. Nevertheless, the fact that all three estimation methods yield uniform results in terms of sign and statistical significance offers strong evidence for a different structure of executive remuneration in unionized firms.

5.3. Pay-performance elasticity

In order to assess whether unionized firms have distorted incentives alignment in their executives' remuneration, we estimate the relation between executive cash pay and firm's 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 executive cash pay is as follows:

[3] $$\Delta \ln(\text{Cash Pay})_{jit} = \beta_0 + \beta_1 \Delta \ln SV_{it} + \beta_2 \left(\Delta \ln SV_{it} \times UNION_i\right) + \beta_3 UNION_i + \varepsilon_{jit}$$

where the independent variable is the percentage change in the cash pay compensation for executive *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. 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. Following Murphy (1999), the change in shareholder value equals the continuously accrued rate of return, $\Delta \ln(1 + \text{Return to Shareholders})$. Also, 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. Remuneration 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.[27] Moreover, estimating the performance sensitivity of stock options is not straightforward since it requires assumptions on the executives' risk aversion, wealth and diversification.

[Insert Table 5 about here]

Table 5 demonstrates that performance elasticity of cash pay with respect to market performance is positive and significant, while β_2 and β_3 are found to be statistically insignificant at the 10% level. Notably, the results do not change even if we would control for firm risk in the specification. These estimates suggest that union presence has no significant effect on the alignment of executive incentives, thus casting doubt on one aspect of Jensen and Murphy's (1990) assertion that unions inhibit efficient executive incentive design due to their demands for wage fairness and a re-allocation of rents. Unlike regulatory pressures, which have been found to dramatically reduce executive compensation and its performance sensitivity, union presence

[27] Cash pay is a major component of executive remuneration, representing 44% and 35% of the total compensation for Managers and CEOs, respectively. The respective proportion of stock option compensation is 39% and 48%, while 23% of the executive-year observations in our sample reflect zero stock option awards. In absolute dollar amount, the median stock-option award for senior managers and CEOs is $150,257 and $479,107 respectively (in 1992 dollars). Furthermore, the median percentage of the company's shares owned by Managers and CEOs in our sample is 0.04% and 0.29%, respectively.

appears to moderately reduce executive remuneration, but without any apparent sacrifice in the alignment of executive incentives to firm performance.

5.4. Intra-firm and inter-firm dispersion of executive compensation

The degree of remuneration inequality within a firm is an important determinant of organizational performance. As Lazear (1989) illustrates in his model of industrial politics, remuneration equality can be desirable on efficiency grounds when workers worry about reference groups and cooperation is necessary in production. Among labor market institutions, organized labor has been found to reduce remuneration dispersion not only for its members, but also for employees in non-unionized firms through a spillover mechanism. 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), reduction in job performance (Mas, 2006), and interference with strategic plans (Kole and Lehn, 2000). Could the presence of unions and associated remuneration compression preferences also extend throughout an organization's corporate structure, including to managers in the upper echelon of the firm? In the following two sub-sections we examine whether union presence is associated with reduced remuneration dispersion both within-firms and across-firms.

5.4.1. Unions and dispersion of executive compensation within firms

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. Distance is defined as the difference between CEO remuneration and the mean remuneration for the four highest-paid managers in firm *i* at time *t*, $Dist_{it} = \left(W_{CEO} - \overline{W}_{Managers}\right)_{it}$.

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 relevant control variables.[28] Put formally:

[4] $\text{Distance}_{it} = \alpha_0 + \alpha_1 \cdot UNION_i + \gamma \cdot Z_{it} + \delta \cdot Industry_i + \varepsilon_{it}$

, 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 and firm characteristics, $Industry_i$ corresponds to a vector of 2-digit SIC industry effects, and ε specifies the error term. Given unions' aversion to inequality, we expect that α_1 will be negative. CEO characteristics that could positively influence the distance measure are CEO/Chairman duality (due to dual responsibilities and task complexity), founder status (reflecting possible entrenchment), CEO tenure and age (the marginal returns for the CEO's human capital may be different than those for Managers), and CEO stock ownership (incentives are better aligned for CEOs with higher ownership). Also, the participation of top managers in the board of directors would improve their compensation, thus reducing the differential with the CEO. In terms of firm characteristics, we include firm size and performance which are expected to have a positive sign given the norms of proportionality between successive levels in organizational hierarchies (Mahoney, 1979; Simon, 1957). Importantly, our estimations deal with the monetary value of the distance measure, rather than on the CEO remuneration 'slice' (i.e., the CEO's remuneration divided by the total remuneration for

[28] Negative outliers reflect 1.6% of the sample for cash pay differentials and 4.1% of the sample for total compensation differentials. 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 receives $1 as annual salary.

the top five executives of the firm) because the latter measure is fractional, highly leptokurtic with fat tails, and it does not distinguish cases in which all top executives in a firm are overpaid.

Due to the presence of negative outliers skewing the distribution, OLS estimation would produce inefficient estimates. As a result we utilize robust estimators, namely quantile regression, and robust regression with M-estimation and Huber weighting. The benefit of quantile regression is that it examines the impact of union presence on the entire conditional distribution of executive remuneration, rather than as a single central tendency measure. Also, robust regression downweights influential outliers, while closely approximating OLS efficiency.

Table 6 presents the findings for both cash pay and total compensation. Consistent with our hypothesis, it is found that union presence substantially decreases intra-firm remuneration differentials between the CEO and the top managers, in terms of both cash pay and total compensation. In particular, union presence is found to decrease the median (average) cash pay differential by 8% (6%) for unionized firms. The respective median (average) estimate for total compensation differential is 10% (14%) for unionized firms.

Figure 4 captures the notable heterogeneity in the union effect across different levels of compensation differentials, utilizing the quantile regression for specification [4]. The union estimates are considerably different across lower and higher quantiles, thus verifying that firms with union presence avoid a large remuneration differential, effectively reducing the remuneration dispersion between the CEO and the lower-level employees. For example, the union presence estimate for the '*Distance in Cash-Pay*' is statistically insignificant at the 0.3 quantile, but negative (-12%) and highly statistically at the 0.7 quantile. Similarly, in the case of '*Distance in Total Compensation*', the union presence relation is statistically insignificant at the 0.3 quantile, but negative (-20%) and highly statistically at the 0.7 quantile. Finally, besides union presence, three control variables are found to have a statistically significant and sizeable estimate across regressions, namely CEO/Chairman duality, Managers' board membership, and firm size. These results for the control variables underline the monetary premium of board responsibilities and the increasing complexity of larger firms.

[Insert Table 6 and Figure 4 here]

5.4.2. Unions and compression of executive compensation across firms

In this sub-section we examine the relation between union presence and the entire distribution of executive 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 inter-firm variation in the distribution of executive compensation, we measure the dispersion in the remuneration awarded to executives (CEO and Managers) in unionized and non-unionized firms. As discussed in Section 4.2, we effectively compare compensation distributions using two comparable populations of executives where the key differentiating factor is union presence in the firm. Dispersion of remuneration within each group of executives is measured with three indices, namely the Gini coefficient, the Theil entropy measure and the standard deviation of logs. The motivation for using multiple dispersion measures is that they have different degrees of decomposability among population subgroups, different emphases of distribution transfers, and possibly intersecting Lorenz curves. Also, the three aforementioned indices were utilized by DiNardo et al. (1996) to examine the effect of unions on US wage dispersion 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 dispersion indices between the matched samples (Mills and Zandvakili, 1997).

[Insert Table 7 about here]

The results in Table 7 indicate that union presence is associated with a compressed cash pay distribution for both CEOs (Panel A) and Managers (Panel B). For instance, in the case of CEOs (Managers) the Gini coefficient for cash pay in unionized firms is 0.317 (0.311), while in non-unionized firms it is 0.381 (0.337). Despite some differences in each dispersion index, these differences are statistically significant at the 5% level. Regarding total compensation, we also observe lower dispersion measures for unionized firms, but they are not different in a statistically significant manner compared to those in non-unionized firms. Overall, these results further support Freeman's (1980, 1982) long-held assertion that a primary objective of trade unions is to reduce wage dispersion in terms of standard rates, such as base salary and 'capped' accounting-based bonuses in the case of executives.

6. Discussion

In this paper we demonstrate that union presence is associated with moderately lower levels of total compensation for executives, though with quite different magnitudes for cash pay and stock options. The negative association between union presence and executive remuneration is also found to be stronger the higher-up in the remuneration distribution ladder one moves. In addition, we find no evidence of lower elasticity of executive cash pay to firm performance, a finding that runs against one of the predictions of the implicit regulation hypothesis.

These results confirm a traditional picture of what unions do. Unionism at the firm-level appears to have similar effects for top executives as those found for workers, in that it decreases performance-based awards and reduces fixed-salary dispersion. These findings are consistent with Freeman and Medoff's (1984) view of unions as institutional channels of worker demands, legitimate or otherwise, for a more compressed distribution of income inside the workplace.

Our findings could also offer a new perspective on international differences in CEO compensation. Abowd and Bognanno (1995) studied executive compensation in twelve OECD countries for the period 1984-92 and found that CEO compensation in the United States was substantially larger than that of comparable companies in other advanced OECD countries. Moreover, they showed that this difference was due to the greater value of stock-related compensation in the United States, a phenomenon that is not replicated in other OECD countries.[29] Since it appears that the mere presence of a union within a firm restrains the value of stock options awarded to the CEO, it seems likely that stronger (and often more militant) unions – like those found in and most European countries – can have an important restraining effect on executive compensation, and in particular over the growth of CEO stock option compensation. The presence of stronger union movements in most OECD countries could therefore be one important reason keeping average CEO compensation lower than that in the United States.

Notably, the findings in this paper draw a clear distinction between pressures from unions and regulation since regulatory oversight in specific industries (such as utilities) has been found to dramatically reduce all aspects of executive compensation, as well as inhibit the alignment of executive incentives to firm performance. The difference between these two 'implicit regulators' could be attributed to the fact that utility firms typically operate in local monopoly markets, while many unionized firms operate in competitive environments subject to product market pressures thus having to compete in the labor market for executives with other non-unionized

[29] Weak union presence is not discussed explicitly as a cause of greater stock option value for CEOs in the United States. The dominant explanation is that stock option awards were influenced by more favorable accounting and tax treatment (Abowd and Bognanno, 1995; Conyon and Murphy, 2002) and weak corporate governance arrangements (Bertrand and Mullainathan, 2001; Core et al., 1999).

firms in their respective industry. It appears that by restricting only the right tail of the executive remuneration distribution, unionized firms can effectively compete in the executive labor market.

In a similar fashion, the disproportionate union estimate across the conditional distribution of total compensation may not be detrimental given the decreasing incentive power of stock options. Tian (2004) argues that at high levels of stock option exposure, stock based compensation can become wasteful and unproductive due to rising agency costs and the excessive levels of risk that executives have to bear since executives cannot sell or hedge their stock options. He further suggests that stock options generate incentives to increase shareholders' wealth only if executive's option wealth, as a fraction of total wealth, does not exceed an optimal threshold. Thus, union presence within a firm can perhaps inadvertently, by curbing extremely large stock option payouts, keep option wealth closer to an optimal threshold, thereby maintaining the incentive power of stock option design.

Overall, our findings serve as an impetus for further work on the interaction between labor institutions and executive remuneration, especially that of middle-level managers whose compensation information is not typically disclosed. Future work should focus on whether the estimates presented here are moderated by factors like union strength and membership militancy.[30] Another important topic for future work is testing whether unionized firms offer executives more promotion-based incentives or a lower stock-option forfeiture probability in order to counterbalance the lower stock-option compensation.

[30] 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 from the Bureau of Labor Statistics and the IMD World Competitiveness Yearbook on work stoppages, industrial disputes and quality of labor relations.

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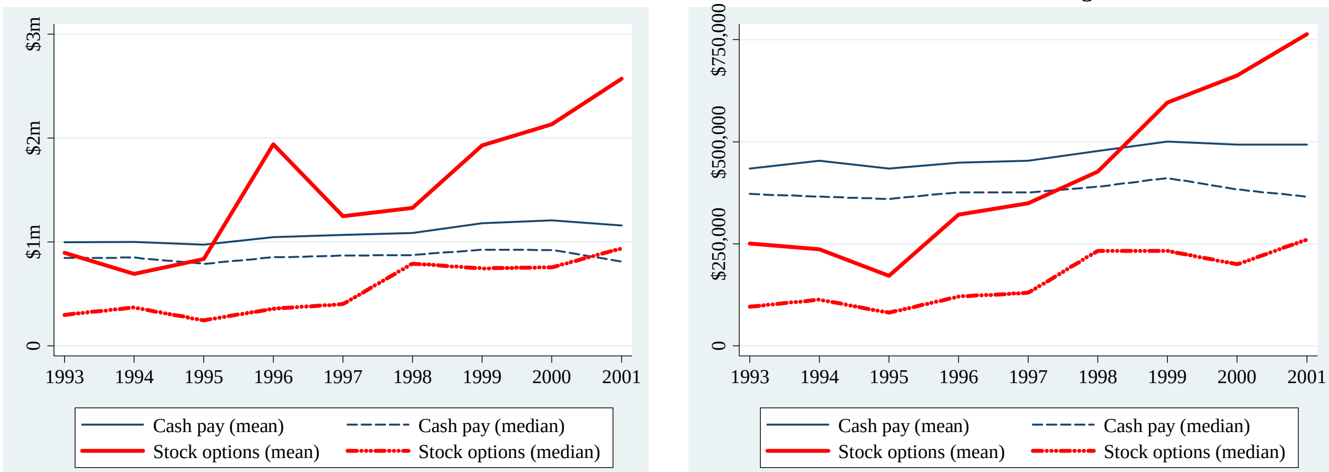
FIGURE 1

EVOLUTION OF EXECUTIVE REMUNERATION IN S&P 1500 FIRMS (1993-2001)

PANEL A: CEOs PANEL B: Managers

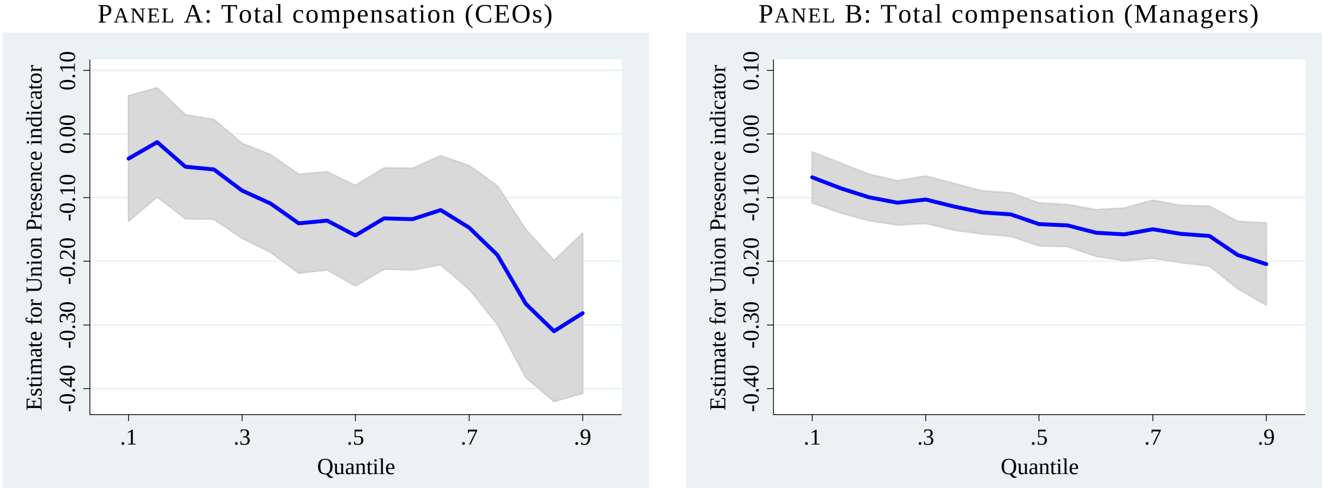


NOTES – The two panels illustrate the evolution of mean/median cash pay and stock option awards during the 1993-2001 period (separately for CEOs and Managers). The figures are based on the overall sample of 46465 executive-year observations and 9293 firm-year observations, and omit year 1992 since Execucomp does not cover the entire S&P1500.

FIGURE 2

DIFFERENTIAL EFFECT OF UNION PRESENCE ACROSS THE CONDITIONAL DISTRIBUTION OF TOTAL COMPENSATION

PANEL A: Total compensation (CEOs) PANEL B: Total compensation (Managers)

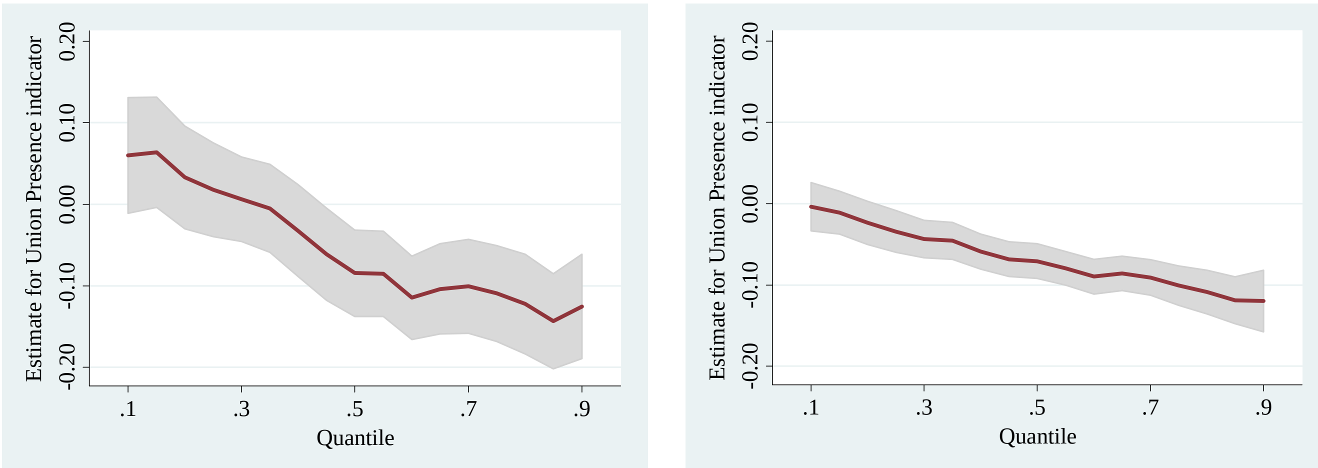


NOTES – The two panels report the effect of union presence on executives' total compensation (separately for CEOs and Managers), utilizing the quantile regression for specification [1]. The shaded region is the 95% confidence band using bootstrapped standard errors (500 replications).

FIGURE 3

DIFFERENTIAL EFFECT OF UNION PRESENCE ACROSS THE CONDITIONAL DISTRIBUTION OF CASH PAY

PANEL A: Cash pay (CEOs) PANEL B: Cash pay (Managers)

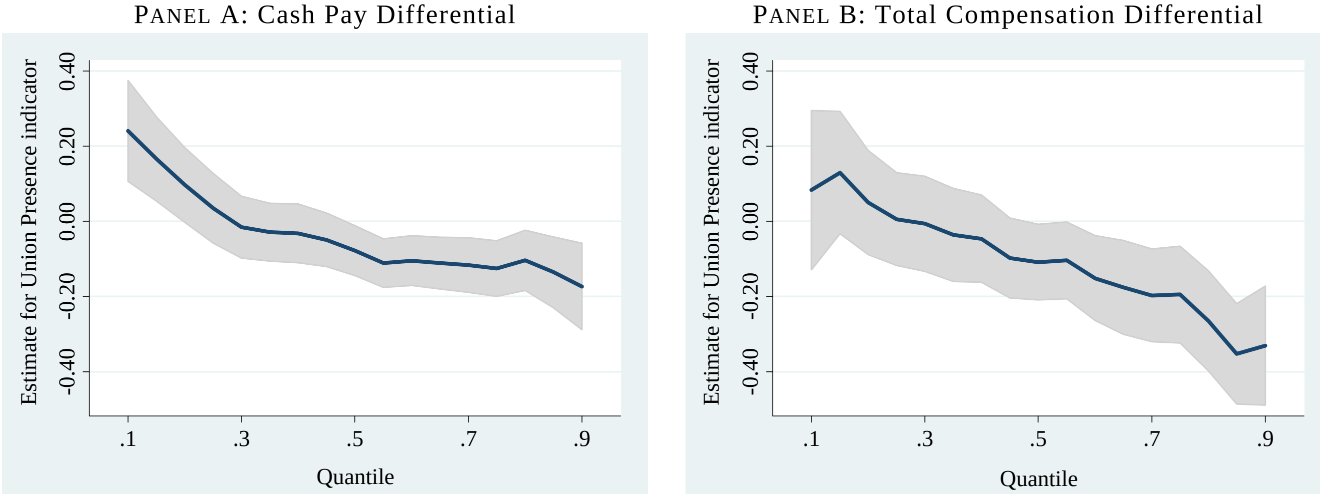


NOTES – The two panels report the effect of union presence on executives' cash pay (separately for CEOs and Managers), utilizing the quantile regression for specification [2]. The shaded region is the 95% confidence band using bootstrapped standard errors (500 replications).

FIGURE 4

DIFFERENTIAL EFFECT OF UNION PRESENCE ACROSS THE CONDITIONAL DISTRIBUTION OF REMUNERATION GAP

PANEL A: Cash Pay Differential PANEL B: Total Compensation Differential



NOTES – The two panels report the coefficient estimate of union presence on the 'Distance in Cash Pay' and 'Distance in Total Compensation', respectively, utilizing the quantile regression for specification [4]. The shaded region is the 95% confidence band using bootstrapped standard errors (500 replications).

TABLE 1

UNION PRESENCE IN THE SAMPLE BY MAJOR SECTOR (%), 1992-2001

	Union presence (Overall sample)		*Sector composition*	
Major sectors	Firm-year observations	Unique firms	Overall sample	Matched pairs sample
Mining & Construction	16.4 %	16.7 %	5.1 %	3.2 %
Manufacturing	39.2 %	34.5 %	44.7 %	69.7 %
Transportation & Com. & Utilities	28.0 %	25.9 %	13.2 %	5.7 %
Wholesale & Retail Trade	22.0 %	19.1 %	11.5 %	12.6 %
Finance, Insurance & Real Estate	2.9 %	1.4 %	12.7 %	–
Services	13.6 %	10.5 %	12.2 %	8.8 %
Other	7.7 %	16.7 %	0.6 %	–
Number of observations (Percentage with union presence)	9293 (26.7%)	2070 (22.6%)		

NOTES – The first column identifies the major sectors, while the second and third columns present the percentage of firms in the overall sample with 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 fourth and fifth columns indicate the sector composition for both the overall sample and the matched pairs sample of unionized and non-unionized firms. Finally, the overall sample contains an unbalanced panel with 46465 executive-year observations and 9293 firm-year observations from 2070 unique firms, while the matched sample corresponds to 7730 executive-year observations and 1546 firm-year observations (773 unionized and 773 non-unionized). Notably, utilities are excluded from the matched pairs sample.

TABLE 2

DESCRIPTIVE STATISTICS AND DEFINITIONS OF VARIABLES

Variables	*Definition*	*Overall sample*		*Matched sample*	
		Mean	Std. Dev.	Mean	Std. Dev.
CEO Cash Pay	Cash pay (salary & bonus) – *in $million*	1.07	1.62	1.08	0.91
CEO Stock Options	Black-Scholes value of CEO's stock options awarded – *in $million*	1.82	8.65	1.50	5.37
CEO Total Compensation	The sum of salary, bonus, benefits, Black-Scholes value of stock options, restricted grants and LTIPs – *in $million*	3.45	11.13	3.10	6.24
Managerial Cash Pay	Cash pay (salary & bonus) – *in $million*	0.48	0.58	0.47	0.39
Managerial Stock Options	Black-Scholes value of CEO's stock options awarded – *in $million*	0.56	1.91	0.42	1.07
Managerial Total Compensation	The sum of salary, bonus, benefits, Black-Scholes value of stock options, restricted grants and LTIPs – *in $million*	1.23	2.92	1.07	1.58
Union Presence	Indicator variable taking value 1, if the firm has at least one unionized establishment, and 0 otherwise.	0.26	0.44	0.50	0.50
Executive is Chairman	Indicator variable taking value 1, if the executive is the Chairman of the Board, and 0 otherwise.	0.17	0.37	0.18	0.39
Executive is Board Member	Indicator variable taking value 1, if the executive is a member of the Board of Directors (other than Chairman), and 0 otherwise.	0.22	0.42	0.20	0.40
Ownership	The executive's stock ownership of the firm, as a percentage of the total outstanding shares (excluding options) – in percentage form	0.82	3.56	0.59	2.68
Firm Size (Sales)	Firm size, in terms of total sales – *in $billion*	3.44	9.22	3.76	6.53
Return on Assets	ROA is defined as net income before extraordinary items and discontinued operations divided by total assets – *in percentage form*	3.79	13.50	5.79	4.64
Return to Shareholders	RET is defined as total return to shareholders, including the monthly reinvestment of dividends – *in percentage form*	21.56	59.29	13.08	35.53
Firm Age	Number of years that the firm has been publicly listed.	26.11	18.92	31.42	19.85
Rust Belt	Indicator 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	0.27	0.44
Right-to-Work	Indicator variable taking value 1 if the firm has its headquarters in a state with right-to-work laws (namely AL, AZ, AR, FL, GA, ID, IA, KS, LA, MS, NE, NV, NC, ND, SC, SD, TN, TX, UT, VA and WY), and 0 otherwise.	0.28	0.45	0.29	0.45

NOTES – For each firm-year observation in the sample we have compensation data for the CEO and the four-highest paid managers. The overall sample contains an unbalanced panel of 46465 executive-year observations from 2070 unique companies in the 1992-2001 period. The matched sample corresponds to 7730 executive-year observations and 1546 firm-year observations (773 unionized and 773 non-unionized). 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. All level variables have been adjusted for inflation and are stated in 1992 dollars.

TABLE 3

UNION PRESENCE AND EXECUTIVE COMPENSATION

Independent Variables	*CEO TotalComp.* [1]	*CEO Cash Pay* [2]	*CEO Stock Option Value* [3]	*Manager TotalComp.* [4]	*Manager Cash Pay* [5]	*Manager Stock Option Value* [6]
Union Presence	-0.127***	-0.055*	-0.198**	-0.133***	-0.066***	-0.257***
	(2.78)	(1.79)	(2.42)	(3.91)	(3.00)	(3.37)
Executive is Chairman	0.545***	0.518***	0.496***	0.321***	0.386***	0.243*
	(5.04)	(7.87)	(2.61)	(4.61)	(5.76)	(1.93)
Executive is Board Member	0.348***	0.406***	0.463**	0.302***	0.290***	0.236***
	(2.89)	(5.89)	(2.25)	(10.77)	(15.28)	(4.23)
Stock Ownership	-0.011*	0.000	-0.006	-0.017***	0.001	-0.050*
	(1.77)	(0.00)	(0.40)	(2.49)	(0.23)	(1.94)
Firm Size (*lnSales*)	0.433***	0.328***	0.444***	0.402***	0.297***	0.420***
	(18.64)	(20.82)	(11.16)	(25.18)	(30.27)	(11.81)
Return on Assets	0.026***	0.020***	0.018**	0.018***	0.011***	0.018**
	(4.85)	(5.42)	(2.04)	(4.55)	(5.36)	(2.41)
Return to Shareholders	0.001*	0.001***	-0.001	0.000	0.001***	-0.002**
	(1.73)	(3.74)	(0.72)	(0.91)	(4.27)	(2.30)
Industry effects (2-digit SIC)	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1546	1546	1546	6184	6184	6184
R^2	0.470	0.543	—	0.485	0.585	—

NOTES – Each of the columns 1, 2, 4 and 5 present an OLS regression, while columns 3 and 6 present a two-tier estimation based on Cragg (1971). The specification for the Cragg (1971) first-tier includes stock ownership, firm riskiness (i.e., three-year standard deviation of monthly returns), industry effects and year effects. All dependent variables are are windsorized at the 0.5% (both tails), in order to reduce the effect of severe outliers. Total compensation and cash pay are in natural logarithm form. Due to the presence of a cluster of zero values for stock option compensation, 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 keeps zero values in the sample. A constant term is included but not reported. The absolute values of robust *t*-statistics [*z*-statistics in columns 3 and 6] with firm-clustering appear in parentheses below each coefficient estimate. Asterisks denote significance at 1 percent (***), 5 percent (**) or 10 percent (*) level.

TABLE 4

UNION PRESENCE AND EXECUTIVE COMPENSATION - ROBUSTNESS CHECKS

Independent Variables	*CEO Median reg* [1]	*Non-CEO Median reg* [2]	*CEO IV reg* [3]	*Non-CEO IV reg* [4]
Union Presence	-0.160***	-0.142***	-0.440***	-0.517***
	(3.98)	(8.37)	(2.08)	(3.07)
Executive is Chairman	0.587***	0.280***	0.698***	0.270***
	(2.72)	(5.70)	(4.07)	(3.59)
Executive is Board Member	0.417*	0.297***	—	0.305***
	(1.96)	(12.52)		(9.65)
Stock Ownership	-0.013***	-0.009*	-0.012*	-0.012
	(3.01)	(1.81)	(1.88)	(1.57)
Firm Size (*lnSales*)	0.452***	0.423***	0.439***	0.409***
	(19.44)	(55.14)	(18.45)	(23.15)
Return on Assets	0.026***	0.019***	0.026***	0.019***
	(4.69)	(8.50)	(4.94)	(4.36)
Return to Shareholders	0.001	0.000	0.001*	0.000
	(1.08)	(0.77)	(1.69)	(0.85)
Industry effects (2-digit SIC)	Yes	Yes	Yes	Yes
Year effects	Yes	Yes	Yes	Yes
Observations	1546	6184	1546	6184
Pseudo R^2 (Centered R^2)	0.312	0.308	(0.411)	(0.429)
Hansen *J* statistic	—	—	0.555	4.418
(p-value)			(0.758)	(0.110)

NOTES – The dependent variable is the natural logarithm of total compensation. Columns 1 and 2 present a median regression, while columns 3 and 4 present an IV (2SLS) regression. As instrumental variables for union presence we employ three variables: *RIGHT-TO-WORK*, *RUST-BELT* and *FIRMAGE*. *RIGHT-TO-WORK* is a indicator variable taking value 1 if the firm has its headquarters in a state with right-to-work laws during the 1992-2001 period, and 0 otherwise. *RUST-BELT* is an indicator variable taking value 1 if the firm has its headquarters in a rust-belt state, and 0 otherwise. *FIRMAGE* denotes the number of years that the firm has been publicly listed. Notably, the dependent variables are windsorized at the 0.5% level (both tails), in order to reduce the influence of extreme outliers. The values of *z*-statistics appear in parentheses below each coefficient estimate (bootstrapped standard errors for the median regression and robust standard errors with firm clustering for the IV regression). A constant term is included but not reported. Asterisks denote significance at 1 percent (***), 5 percent (**) or 10 percent (*) level.

TABLE 5

UNION PRESENCE AND CEO PAY-FOR-PERFORMANCE ELASTICITY

	CEO	*Manager*
	Δ ln(Cash pay)	*Δ ln(Cash pay)*
Independent Variables	[1]	[3]
Δ ln(Shareholder Value)	0.375***	0.223***
	(5.02)	(4.42)
Δ ln(Shareholder Value) × Union Presence	-0.062	-0.062
	(0.67)	(0.96)
Union Presence	0.018	-0.021
	(0.72)	(1.16)
Observations	766	2454
R^2	0.124	0.063

NOTES – All columns present OLS regressions. Cash pay is the sum of salary and bonus. The absolute values of robust *t*-statistics with firm clustering appear in parentheses below each coefficient estimate from OLS estimation. A constant term is included but not reported. Notably, the dependent variable is windsorized at the 0.5% level (both tails) in order to reduce the influence of extreme outliers. Given that estimating pay-performance elasticity requires dynamic information (i.e., first differences) that is often missing, we keep only 383 matched pairs of unionized and non-unionized firms that have information for both CEOs. Similarly, we keep only matched pairs of unionized and non-unionized firms with first-differences information for an equal population of Managers. For Managers, there are 142 matched pairs with four Managers in each firm, 179 matched pairs with three Managers in each firm, 58 matched pairs with two Managers in each firm, and 6 matched pairs with only one Manager in each firm. Asterisks denote significance at 1 percent (***), 5 percent (**) or 10 percent (*) level.

TABLE 6

UNION PRESENCE AND REMUNERATION GAP BETWEEN THE CEO AND THE TOP-4 MANAGERS

	(Distance) Cash pay		*(Distance) Total compensation*	
	Median reg.	*Robust reg.*	*Median reg.*	*Robust reg.*
Independent Variables	[1]	[2]	[3]	[4]
Union Presence	-0.078**	-0.061**	-0.109**	-0.149***
	(2.30)	(2.19)	(2.12)	(3.51)
CEO is Chairman	0.222***	0.185***	0.322***	0.325***
	(4.17)	(4.71)	(4.04)	(5.41)
CEO is Founder	-0.145	-0.122*	-0.038	-0.064
	(1.14)	(1.95)	(0.19)	(0.67)
CEO Tenure	0.003	0.002	0.006	0.007*
	(0.95)	(1.01)	(0.93)	(1.82)
CEO Age	0.014***	0.013***	-0.008	-0.006*
	(4.05)	(5.55)	(1.30)	(1.67)
Number of Top-4 Managers in the Board (*ranges between 0-4*)	-0.088***	-0.095***	-0.152***	-0.168***
	(4.02)	(5.72)	(4.98)	(6.63)
Firm Size (*lnSales*)	0.344***	0.354***	0.501***	0.488***
	(17.07)	(25.39)	(18.17)	(22.86)
Return on Assets	0.027***	0.029***	0.036**	0.033**
	(6.05)	(8.65)	(4.90)	(6.44)
Return to Shareholders	0.002***	0.002***	0.001*	0.001*
	(3.22)	(4.36)	(1.81)	(1.91)
Industry effects (2-digit SIC)	Yes	Yes	Yes	Yes
Observations	1546	1546	1546	1546
Pseudo-R^2 (R^2)	0.173	(0.349)	0.122	(0.305)

NOTES – The sample consists of an unbalanced panel with 1546 firm-years. The absolute values of *t*-statistics appear in parentheses below each coefficient estimate. Standard errors are bootstrapped to address potential heteroskedasticity (500 bootstrap replications). A constant term is included but not reported. The goodness-of-fit statistic (R^2) in Columns 2 and 4 is derived from rregfit command in STATA. Asterisks denote significance at 1 percent (***), 5 percent (**), or 10 percent (*) level.

TABLE 7

DISPERSION OF EXECUTIVE COMPENSATION ACROSS FIRMS

PANEL A: Dispersion indices for CEOs' compensation in unionized and non-unionized firms

	Unionized firms	*Non-unionized firms*	*Difference*	|*t*-statistic|
Dispersion of cash pay				
Gini-coefficient	0.317	0.381	-0.064	3.50 ***
Theil entropy measure	0.167	0.283	-0.116	3.34 ***
Standard deviation of logs	0.315	0.441	-0.126	3.69 ***
Dispersion of total compensation				
Gini-coefficient	0.510	0.553	-0.043	1.19
Theil entropy measure	0.499	0.711	-0.212	1.30
Standard deviation of logs	0.785	0.919	-0.134	1.93 *
Number of Observations	773	773		

PANEL B: Dispersion indices for Managers' compensation in unionized and non-unionized firms

	Unionized firms	*Non-unionized firms*	*Difference*	|*t*-statistic|
Dispersion of Cash pay				
Gini-coefficient	0.311	0.337	-0.026	2.51 **
Theil entropy measure	0.166	0.225	-0.059	2.09 **
Standard deviation of logs	0.293	0.333	-0.040	3.08 ***
Dispersion of Total compensation				
Gini-coefficient	0.483	0.484	-0.001	0.07
Theil entropy measure	0.472	0.475	-0.003	0.05
Standard deviation of logs	0.650	0.680	-0.030	1.16
Number of Observations	3092	3092		

NOTES – The table presents and compares the level of dispersion in cash pay and total compensation among the top executives for matched pairs of unionized and non-unionized firms. The overall sample contains 1546 CEO-year observations and 6184 Manager-year observations and from 1546 firms (773 unionized firms and 773 non-unionized firms). The matching process is described in detail in Section 4.2. In each Panel, the second and third columns present the three dispersion 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 (500 replications) are obtained for each of the three indices. Asterisks denote that the difference is significant at 1 percent (***), 5 percent (**) and 10 percent (*), respectively.

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