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What Do Unions Do To CEO Compensation?

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Abstract
In this paper we estimate the relation between union presence and CEO compensation, using a unique panel of publicly listed US firms for the period 1992-2001. We find that, on average, union presence: 1) is significantly associated with lower levels of total CEO compensation; 2) affects the mix of CEO compensation by providing higher levels of pay but much lower stock option values; 3) lowers dispersion of CEO compensation across firms; and 4) does not significantly reduce the performance sensitivity of CEO compensation as compared to non-union firms.

JEL Classification: J33, J51
Keywords: Unions, CEO compensation, implicit regulation

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"A closer look at the tense negotiations between American Airlines and its unions shows how the revelation of executive perks almost drove American into bankruptcy. Workers learned that Donald J. Carty, American's chairman, had hidden executive benefits, including a protected executive pension trust fund, while he was negotiating for deep concessions from the unions. While the two sides eventually reached an agreement to keep the American out of bankruptcy, the airline still faces considerable obstacles, which it must now confront without Mr. Carty, who was ousted."

The New York Times, April 27th 2003.

1. Introduction

Jensen and Murphy [1990], in their seminal paper on CEO compensation, posit that political and regulatory constraints truncate the upper tail of executive remuneration, resulting in lower overall levels of CEO pay and reduced pay-performance sensitivities. They refer to this as the effect of *implicit regulation* and identify government legislation and the presence of unions as two obvious examples of such institutional constraints. Several empirical studies have since confirmed Jensen and Murphy's predictions for firms operating in regulated industries, where government oversight and disclosure rules ensure that executive pay remains a highly visible and contentious subject.[1] Similar research, however, on the possible constraining effects of unions is scarce. One obvious reason is the difficulty in identifying plausible causal links between union presence (which is found at an establishment level) and executive remuneration (which is set at the level of the firm). The other more important obstacle, up to now, has been the lack of a suitable dataset capable of matching union information with executive compensation data.[2]

[1] 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 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.

[2] 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".

This paper fills this gap in the literature by using a unique panel of 2,159 publicly listed US firms, observed from 1992 to 2001, in which the presence (or absence) of a union is identified with workplace data (i.e. NLRB elections, FMCS contract expirations) linked to our sample of parent companies. Our results suggest that union presence is indeed significantly associated with CEO compensation. Specifically, we find that union firms display lower levels of total CEO compensation (i.e. the sum of salary, bonus, benefits, restricted grants, long term incentive plans, and stock options) as compared to non-union firms and that this effect is primarily the result of substantially lower stock option values, which outweigh a positive -but smaller- pay effect associated with union presence. The constraining effect of unionisation on total CEO compensation is also found to be stronger the higher-up in the distribution ladder one moves. That is, the upper tail of CEO compensation is significantly reduced within unionised firms, just as Jensen and Murphy predicted. Somewhat surprisingly, the elasticity of CEO compensation with respect to firm performance does not differ significantly between union and non-union firms. These results distinguish between union and regulatory pressures, since unions appear to re-arrange the mix of CEO compensation without inhibiting incentives.

In addition to the findings above, the paper is unique in several respects. First, as compared to earlier studies by DiNardo et al. [2000] and Singh and Agarwal [2002], we provide the first systematic look at unions and CEO pay since the explosion of stock-based compensation in the early 1990s. Second, we employ a larger and richer dataset, with compensation data taken from SEC fillings[3] and unionisation variables taken from the Bureau of National Affairs' (BNA) extensive electronic archive of NLRB elections and FMCS contract expirations data. Third, rather than remaining agnostic about the source(s) of difference in CEO compensation between union and non-union firms, we provide two

[3] In response to criticisms of excessive executive compensation, the Securities and Exchange Commission (SEC) adopted significantly modified proxy disclosure requirements for executive compensation in October 1992 (SEC Release No. 33-6940).

possible mechanisms – one direct and the other indirect – that can account for our empirical results. Finally, by "breaking down" CEO compensation into its major components (i.e. salary, bonus, and stock-options), we are able to distinguish amongst the various indirect and direct channels of union influence identified in our analytical section.

The rest of the paper is organized as follows: Section 2 provides a brief literature review. Section 3 outlines the possible channels of union influence upon firm corporate governance, financial markets and ultimately on CEO 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. Previous literature

As noted in the introduction, only two studies have estimated the effect of union presence within a firm on the level of CEO compensation. DiNardo et al. [2000] examined the effects of unionisation in US firms for the periods 1971-74, 1975-78 and 1979-82 on levels of CEO pay but not for total compensation (i.e. they exclude the value of stock options).[4] The unionisation data is for 1977 and 1987. They limited their study to unionised firms only and evaluated how the strength of unionisation, measured in terms of union density at the firm level, affects CEO pay. DiNardo et. al., [2000] initially found that greater levels of unionisation were negatively and significantly associated with CEO 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,

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

although they admit that they did not have a satisfactory explanation as to why this period should differ from the rest of their sample.

More recently, 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 greater CEO pay, but that for other compensation components (e.g. stock options) and total compensation (e.g., stock options, bonus and salary) no union effect is found. Nevertheless, the study's cross-sectional nature and narrow industrial focus do not offer a definitive view of the issue.

Given the lack of any clear empirical findings, a closer look at possible channels of union influence with respect to CEO remuneration is in order.

3. The Link(s) Between Union Presence and CEO Pay

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 bear 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. 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 labour 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 actualising worker preferences for a more compressed compensation structure within the firm.

Union voice is normally observed during formal collective bargaining negotiations, but it can also be channelled informally through local stewards, public awareness campaigns and even through the sponsorship of journalism and public policy research critical of alleged managerial excesses.[5] 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 influence is gained through representation on the boards of directors, or other senior level committees.[6] A further direct effect, put forward by Jensen and Murphy [1990], maintains that unions perceive high executive pay as a signal for the firm's financial health and employ it as a justification for increased wage demands in labour negotiations. This would naturally make senior executives and boards more cautious when negotiating CEO pay packages. This is also why union presence, much like the public oversight faced by CEOs of firms operating within regulated industries, is predicted to lower CEO compensation and make it less sensitive to firm performance.

The second channel by which unions and executive compensation are related is *indirect* and refers to the adverse consequences that union presence can have on financial markets, which, in turn, determine the value of the stock-related part of executive compensation. This particular channel is quite important given that stock options comprise a substantial (and growing) part of top executive remuneration packages within publicly listed US companies; a fact confirmed in our own dataset and represented in figure 1.

[Figure 1]

Stock-option compensation – unlike salary and bonus, which are fixed or accounting-based – depends on financial performance measures. This means that if financial markets respond

[5] In the US, during the mid 1990s, at the height of the Republican "Contract with America", the union movement financed several reports on CEO pay, corporate downsizing and the links with union-busting behaviour.

[6] 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.

negatively towards union presence, then the incentive power of stock-related compensation for senior executives is reduced within unionised firms.

Studies in the United States almost always 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 labour cost changes and concludes that increases in labour costs due to collective bargaining reduce a firm's market value. Ruback and Zimmerman [1984] find that formal union 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 analyses undertaken during the period when "unions really did matter" (i.e. the 1930s) also confirms the large and significantly negative effect of strikes and workplace disputes on stock values [DiNardo and Hallock, 2002]. In a related fashion, Becker and Olson [1989] find that shareholder returns, though less sensitive to economic fluctuations, are generally lower in unionised firms.[7]

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,[8] but rather from the expectation that union influence over a firm's governance leads to inflexibility, greater

[7] 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 unionisation 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.

[8] 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], Hirsch [1991, 1992], and Bronars and Deere [1993] 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 (2006) for a detailed review].

redistribution of rents toward workers, and ultimately lower profitability.[9] Unionised 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.[10]

In sum, the indirect effect(s) of union presence on CEO compensation work in the following way: Since stock options, generally, have an asymmetric payoff function and since financial markets disapprove of unions, it is more difficult for the share price to increase in union firms. In order to counteract this problem, unionised firms can essentially select between two alternatives.[11] The first alternative is to give more stock options. That is, for every additional dollar of share price change, the CEO of a unionised firm would have more options to exercise as compared to a CEO in a non-unionised firm. The problem is that unions tend to frown upon any pay-for-performance scheme that could substantially accentuate the upper tail of compensation, let alone one for executives. The second alternative for a unionised firm is to offer less stock option compensation, but increase the fixed (i.e. salary) and accounting-based (i.e. bonus, which is always 'capped') part of CEO compensation.[12] In this way, unionised firms can still compete in the labour market for CEOs and attract capable executives.

Which alternative is most often chosen is a question that, along with the overall effect of unions on executive compensation, requires empirical investigation. We address this need by examining two specific hypotheses drawn from the direct and indirect channels of the union effect discussed above:

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[9] Freeman and Medoff [1984: p.248] were clearly aware of this: "American unionism is…a minus on the corporate balance sheet…"

[10] Empirical evidence on firms' cost of fighting labour organization is provided in Abowd and Farber [1990], Freeman [1986] and Freeman and Kleiner [1990].

[11] A third alternative would be to lower the exercise price, but -for accounting purposes- US firms already (in almost all cases) equate exercise price with share price at the day of issue.

[12] This possibility may explain why DiNardo et al. [2000] found that pay was higher in firms with strong union presence after controlling for other factors.

Hypothesis 1: Union presence in firms is associated with reduced overall levels of CEO compensation, particularly because the upper tail of the CEO compensation distribution is truncated and the compensation mix is altered (i.e. union presence may be associated with increased pay but more than likely with a lower value of stock options).

Hypothesis 2: As a result of unions 'implicitly regulating' executive compensation, the distribution of CEO compensation will be more compressed within unionised firms and incentives are expected to be less, in terms of lower pay-performance relations in unionised firms.

4. Data

The unique dataset employed in this paper is the result of merging firm-level Standard & Poor's Executive Compensation observations (Execucomp) with establishment-level Bureau of National Affairs' (BNA) NLRB data. From Execucomp we obtained information on CEO compensation, ownership, tenure, and firm characteristics for 14,013 firm-years covering S&P 1500 firms from 1992 to 2001.[13] From the BNA's electronic union dataset, we obtained information on a multitude of union activity variables – contract expirations, NLRB elections with 'win' outcomes, as well as work stoppages and unfair labour practice charges – that could 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 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. As a result we then were able to create firm-year observations with unionisation variables.

[13] For 1992-1993 Execucomp included S&P 500 firms and only a portion of medium and small capitalization firms.

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. As a result of this detailed firm-specific and year-specific matching process, we were able to produce unionisation variables for 4,912 observations in Execucomp (out of 14,013), which refer to 717 firms (out of 2,502) that had some union presence.[14] Given that establishment unionisation data from BNA is extremely exhaustive, the remaining observations without any union presence were treated as non-union firms.

The final dataset -- after dropping observations that correspond to CEO compensation that is not reported on annual basis and dropping observations with missing dependent variables-- contains 10,249 firm-year observations in the 1992-2001 period, covering 2,922 CEOs in 2,159 firms. This sample has 3,689 firm-year union observations drawn from 652 unionised firms. Though this union total as a fraction of all firms may seem high (30 percent) -- given that levels of union density in the United States are now close to 13 percent -- one must remember that these are the largest firms in the country, which is where most union activity is found. Moreover, we purposely used a fairly broad-based definition of union presence so as to be sure that we were not leaving any unionised firm out of our sample.[15]

[14] Since Execucomp is an unbalanced panel, we dropped many union activity observations for firms that are included in Execucomp but not in that particular year of the union activity (e.g. union strike in a year that the firm is not included in Execucomp). Further details about the dataset construction process are available upon request.

[15] Our definition of union presence was one in which a firm with even a single union-proven establishment linked to the parent company was considered unionised and given a dummy value of one. We experimented

The key explanatory variable employed for the purpose of testing our three propositions is union presence (UNION$_i$), which is an indicator variable taking on the value 1 if the firm *i* had any establishments that were unionised, and 0 otherwise.[16] We identify union presence from contract expirations and NLRB elections with a 'win' outcome in an establishment that belongs to the parent of the S&P listed firm. On a few occasions, union presence was indirectly identified through "union activity" such as strikes and unfair labour practice petitions.[17] This data was provided to us in an electronic form from the Bureau of National Affairs.

Our key dependent variable, CEO compensation, was calculated from its major components: pay (i.e. the sum of salary and bonus); Black-Scholes value of stock options; and finally total compensation (i.e. the sum of salary, bonus, benefits, LTIPs, restricted grants and stock options). Control variables for CEO and firm characteristics include CEO ownership, tenure, firm size, firm accounting and financial performance, and industry/year effects. Firm size is proxied by the natural logarithm of firm assets, whereas firm performance is measured by return on assets and return to shareholders. A full listing of the variables, along with means and standard deviations, can be found in Table 1. Finally, in Table 2 we observe that union presence is more evident in particular industries. Very little unionisation is found amongst financial firms, but above average presence is found within the manufacturing sector; patterns consistent with labour market union density data.

[Tables 1 and 2]

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with other criteria, which were more restrictive and which narrowed our union sample, but our results (available upon request) did not vary from those presented here, either in significance or qualitative direction.

[16] In theory the union dummy can be time-variant, as union presence within a firm can disappear through decertification at the establishment level or can appear through a successful organizing drive in a non-union firm. In practice, however, this is very rare. For example, for all the establishments in the US during the 1990-1993 period the average annual number of union de-certifications was approximately 400 [Nilsson, 1997: Figure 1]. Even more rare would be all establishments within a firm to become non-unionised in our sample period. Thus, it is not surprising that in our own firm-level dataset the union presence variable is essentially time-invariant.

[17] We would like to thank Katherine Windley from the Bureau of National Affairs (BNA) for alerting us to this potential source of union presence information.

5. Results

5.1 Levels of CEO compensation

We test whether union presence has a significant effect on levels of pay *W*, stock option value *S,* and total compensation *Y* using the following specifications:

[1a] $$\ln W_{it} = \alpha_1 + \delta_1 \cdot UNION_i + \gamma \cdot Z_{it} + \varepsilon_{1it}$$

[1b] $$\ln S_{it} = \alpha_2 + \delta_2 \cdot UNION_i + \kappa \cdot Z_{it} + \varepsilon_{2it}$$

[1c] $$\ln Y_{it} = \alpha_3 + \delta_3 \cdot UNION_i + \lambda \cdot Z_{it} + \varepsilon_{3it}$$

where α is the intercept term, $UNION_i$ is the indicator variable indicating union presence, Z_{it} is a vector of control variables on CEO and firm characteristics (including industry and year effects in 1b an 1c), and ε specifies the error terms. Since stock-related compensation was increasingly employed in the 1990s with a varying degree across industries, year and 2-digit SIC industry controls are employed in the panel data estimations for stock options and total compensation.

In [1a] and [1c] we employ random effects panel data estimation. [18] A Hausman [1978] specification test justifies the use of random effect over fixed-effect (or OLS) estimates for both CEO pay and total compensation estimations. In the regression for stock-option compensation [1b], we follow Yermack [1995] and use Tobit estimation, in order to tackle the problem of clustering around zero values, arising from the fact that our dependent variable includes a number of observations with no stock option pay component.[19] The results from the Tobit are robust to alternative estimations that control for selection of firms

[18] We drop the time subscript in the UNION dummy given that our key explanatory variable is essentially time invariant in our sample, another reason making fixed effects estimates unfeasible. Although in theory union presence could of course vary with time –i.e. firms could systematically close down union workplaces and workers could decertify all workplaces where a union is present– in practice firms remain unionised in our sample. See footnote 16.

[19] Random-effects tobit is not applied because its estimates are based on the Gauss-Hermite quadrature approximation which becomes highly unstable in the presence of constant within-group variables, such as the union presence variable in this study.

choosing to offer stock options, such as the two-part estimation (Jones, 2000; Manning et al, 1987) and Heckman estimation with the inverse Mills ratio (Heckman, 1979).

Table 3 presents the findings for levels of total CEO compensation and its major components (i.e. pay and stock option value) 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 10 percent lower total CEO compensation. As anticipated, the negative association with total compensation occurs because of the large negative impact of union presence on stock-option compensation, an effect which dominates the (smaller) positive union effect for pay.

[Table 3]

As an added check of our results and as a direct test of whether unions have a greater effect on the upper tail of the CEO compensation distribution, we employ a quantile regression on our total compensation specification used in [1c]. [20] The results are graphed in figure 2. The figure is fairly clear: the negative association between unions and CEO compensation is much stronger as we move up the compensation ladder. That is, unions appear to significantly limit the upper tail of the CEO pay distribution much more than the lower tail, as Jensen and Murphy [1990] predict. In particular, union presence is associated with 25 percent lower levels of total CEO compensation for the top 90th percentile as compared to only 5 percent for the bottom 10th percentile. The average impact (straight line) picked up in our initial random-effects regression, was 10 percent.

[Figure 2]

[20] In contrast to typical linear regression, in which the dependent variable is the mean, quantile regression uses a percentile (e.g., 10th, 90th) as the dependent variable. This method can be useful to explore whether the effect of union presence varies depending on the quantile chosen. Quantile regression is based on minimizing asymmetrically weighted absolute residuals, and estimates models for the full range of conditional quantile functions. By using this approach one can detect any heterogeneity in the union effect over CEO compensation 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."

To address the possibility that union presence may be linked to CEO compensation in some endogenous manner (i.e. CEOs with higher compensation may actively oppose the establishment of unions in their workplaces) we provide a set of endogeneity robustness checks below.

5.2. Union presence as an endogenous variable

In this subsection we treat $UNION_i$ in equations [1a] and [1c] as an endogenous variable. In particular, we employ robustness checks using IV estimation. As instrumental variables for union presence we use labour intensity and indicators for the region where the firm's headquarters are located, variables that are likely to affect unionisation but not our compensation measures.

Labour-intensive companies are prime targets of union organization. The instrumental variable for labour intensity is the natural logarithm of the ratio of sales to employment for each firm (*ln* INTENSITY). Also, it is well established in the literature that firm location influences union presence (Herod, 1998; Holmes, 2006). In particular, firms located in the bible-belt and the sun-belt have substantially less union activity – due in part to local anti-union labour regulation and attitudes.[21] Firms located in the remaining states have a higher likelihood of being unionised. To capture this positive effect, we construct two dummy variables for firm headquarters being located in a state that is either in the rust-belt (RUST) or the remaining states (REST). RUST is a dummy variable taking value 1 if the firm has its headquarters in a rust-belt state (namely IL, IN, MI, MN, NY, OH, PA, WI and WV), and 0 otherwise. REST is a dummy variable taking value 1 if the firm has its headquarters in a state that does not belong to the bible-belt, sun-belt or rust-belt (namely AK, CO, CT, DC, DE, HI,

[21] The bible-belt is considered to consist of the following fifteen states: AL, AR, FL, GA, KS, KY, LA, MO, MS, NC, OK, SC, TN, TX and VA. In a related fashion, the sun-belt consists of the following five states: AZ, CA, NV, OR and WA

IA, ID, MA, MD, ME, MT, ND, NE, NH, NJ, NM, RI, SD, UT and VT, as well as companies with headquarters outside the United States).

All three variables are highly correlated with union presence but uncorrelated to CEO pay and total compensation. In addition, the F-stat for the first-stage regression in our union prediction – that tests whether the instruments are jointly equal to zero – is 62.19, well above the benchmark of 10 as suggested by Staiger and Stock (1997), further supporting our choice of instruments.[22]

Table 4 presents the findings for levels of CEO pay and total compensation from the IV estimations. We find strong evidence that union presence within a firm is significantly related to lower total compensation. There is also weak evidence (at 15% percent significance level) that union presence is positively associated with CEO pay.[23] Both IV regressions pass the Sargan's test for over-identifying restrictions. We perform several endogeneity and misspecification tests, which basically confirm the lack of endogeneity in $UNION_i$ dummy.

[Table 4]

5.3 Within-group inequality of CEO compensation

Inequality of CEO compensation between non-union and union firms is proxied by the standard deviation of the natural logarithm of CEO compensation, a suitable method when earnings are log normally distributed, as in the case of CEO compensation. Both actual and predicted (from estimations 1*a*, 1*b* and 1*c*) values of CEO compensation are employed in the comparison. The effect of union presence on the dispersion of total CEO compensation across-firms and its components is initially tested with a standard equality of variance test.[24]

[22] Results available upon request

[23] The fact that the 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, 1985a/b).

[24] The standard deviations F-test compares the variances of the CEO compensation components in non-unionised and unionised firms. Even though these tests for heterogeneity of variance are sensitive to non-normality because the variance converges slowly to a normal distribution, the quite large sample employed in

Comparison of the variation across CEO remuneration in our sample of union firms – i.e. pay, stock option value and total compensation –is made against the variation in CEO remuneration within non-union firms:

[2] $$H_0 : Var(Y_u) = Var(Y_n) \quad , \quad H_1 : Var(Y_u) \neq Var(Y_n),$$

where subscripts *u* and *n* represent unionised and non-unionised firms respectively. Based on our second proposition, $Var(Y_u)$ should be significantly different (i.e. lower) than $Var(Y_n)$. The results from *F*-tests for a variance comparison, at the bottom of Table 3, uniformly show that across union firms, CEO pay dispersion is lower than across non-union firms at any conventional level of significance. Similar results are found for total compensation and the predicted value of stock options.

Table 5 indicates that our results are also robust across several alternative measures of dispersion as well as across different firm sizes and degrees of strength of the union presence within the industry in which the firm belongs.[25] For the standard deviation of the natural logarithm of total compensation, dispersion of CEO compensation is smaller across unionised firms than across non-unionised firms (top panel) in situations when the union presence in the industry is low as well as high. This union compression effect is greater in small firms (second panel) than in large firms (bottom panel).

[Table 5]

The expected truncation effect associated with union presence at the upper tails of the CEO pay distribution is also apparent in Table 5. For small firms (bottom panel) the 90th/10th and 90th/50th gap are always smaller across the union sample, whether the union presence in the industry is high or low. For large firms (middle panel) the 90th/10th and 90th/50th gap are

this research enhances the normal approximation suggested by the central limit theorem. The results from the Levene's [1960] statistic for equality of variances, which is less dependent on the assumption of normality, are also reported.

[25] Based on Table 2, industries are distinguished into 'high' and 'low' union presence. High union presence industries include Mining, Construction, Manufacturing, Transportation, Communications and Utilities (i.e. SIC 10-49). Low union presence industries include (Wholesale & Retail) Trade, Finance, Insurance and Real Estate and Services (i.e. SIC 50-89).

smaller only when the union presence in the industry is high; when it is low the gaps are very similar across union and non-union firms.

5.4 CEO pay-performance elasticity

Finally, the specifications employed for estimating the pay-performance elasticity of unionised and non-unionised firms, respectively, are as follows:

[3a] $$\Delta \ln(CEO\, compensation)_{it}\big|_{Union=1} = \alpha_1 + \beta_1 \cdot \Delta \ln SV_{it} + u_{it}$$

[3b] $$\Delta \ln(CEO\, compensation)_{jt}\big|_{Union=0} = \alpha_2 + \beta_2 \cdot \Delta \ln SV_{jt} + v_{it}$$

where α_1 and α_2 indicate intercept terms, β_1 and β_2 are the pay-for-performance estimates, and u_{it} / v_{it} specify the residuals. The equations are applied for changes in both CEO pay and total compensation. The change in shareholder value is denoted by $\Delta \ln SV$, and equals $\ln(1 + R\text{eturn to Shareholders}_{it})$ which yields the continuously accrued rate of return (Murphy, 1999). Parametric t-tests are then performed in order to investigate whether there are significant differences between pay-for-performance estimates in the independent samples of unionised and non-unionised firms. As an added robustness check, we also use the entire sample to estimate the following model that includes an interaction between changes in firm market value *V*, and union presence:

[3c] $$\Delta \ln(CEO\, compensation) = \alpha_3 + \beta_3 \cdot \Delta \ln SV + \beta_4 \cdot (\Delta \ln SV \cdot Union) + \varepsilon$$

Table 6, Panel A, demonstrates that the performance elasticity of pay and total compensation is positive and significant across union and non-union firms. In terms of differences between the two samples, the elasticity of total compensation with respect to changes in market value is lower across unionised firms. However, the coefficient does not differ significantly from that found in non-union firms. For the CEO pay component, the results actually show a stronger pay-performance effect across unionised firms, but again when compared to the non-union sample, the difference in elasticity is not significant. The lack of significance for the interaction term $(\Delta \ln V \cdot UNION)$ in Table 6, Panel B, is also consistent with union presence having no significantly negative effect on CEO incentives within firms.

[Table 6]

The results presented in Table 6 have several possible interpretations. First, they appear to cast some doubt on Jensen and Murphy's [1990] assertion that unions inhibit efficient managerial incentive design due to their demands for wage fairness and a re-allocation of rents. This would mean that unlike regulatory pressures – which have been found to reduce CEO compensation *and* its performance sensitivity – union presence appears to reduce average CEO compensation and lower its variance across firms, but without any apparent sacrifice in overall performance sensitivity. The findings in Table 6 could also be linked, more generally, to 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 associated with awarding the payouts and the excessive levels of risk that executives have to bear. 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 CEO stock option payouts– keep option wealth closer to an optimal threshold, thereby maintaining the incentive power of stock option design.

5.5 Implications

These results reveal a new aspect of executive compensation in listed firms but also confirm an older picture of what unions do. Unions appear to have similar effects for top executives as those found for other employees, in that they increase the fixed-part of compensation (i.e.

pay) but decrease performance-based premiums and pay dispersion, especially at the upper end of the compensation structure. These findings verify Jensen and Murphy's [1990] implicit regulation hypothesis, and are consistent with Freeman and Medoff's [1984] view of unions as institutional channels of employee voice.

The results above 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 options in the United States, a phenomenon that is not replicated in other OECD countries.[26] 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 Korea, Japan 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 unions in a number of OECD countries could therefore be one important reason keeping CEO compensation lower than that in the United States.

Overall, the findings suggest that unions may indeed operate as a "fairness factor and/or implicit regulator", translating union members' desire for reduced intra-firm wage dispersion into reality. It is well established in the literature that union presence is associated with a wage-premium for workers and a reduction in wage dispersion.[27] This effect,

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

[27] See Blanchflower and Bryson [2004], Hirsch and Schumacher [2001] and Lewis [1986] for review of the union wage premium literature. Concerning wage dispersion, Freeman [1980,1982] finds that unions reduce workers wage dispersion both at the aggregate and intra-firm level. Card [2001] partly attributes the increase in male wage inequality to declining unionisation rates.

combined with the finding that unionised firms also award lower CEO compensation, further confirms the union role in lowering intra-firm wage dispersion and acting as a voice mechanism for worker demands (legitimate or otherwise) for a more compressed distribution of income inside the workplace. Notably, the findings in this paper draw a distinction between pressures from unions and regulation, since unions re-arrange the mix of CEO compensation without (seemingly) inhibiting incentives, while regulated industries have been found to reduce all aspects of CEO compensation as well as CEO incentives.

6. Concluding remarks

In this paper we identify that union presence is associated with lower levels of total CEO compensation. We show that this effect is the result of lower stock option values for CEOs in unionised firms, which outweighs the positive pay-effect associated with union presence. The negative association between union presence and total compensation is also found to be stronger the higher-up in the CEO compensation distribution ladder one moves. Additionally, the variance of CEO compensation components across unionised firms is significantly smaller. These findings on altered compensation mix and more compressed CEO compensation distributions in unionised firms provide empirical support for Jensen and Murphy's (1990: p. 262) prediction that "truncating the upper tail of the payoff distribution requires that the lower tail of the distribution also be truncated in order to maintain levels of compensation consistent with equilibrium in the managerial labour market". More detailed work, however, is needed to see whether the estimates presented here are moderated by factors like union strength and membership militancy on the shop floor.

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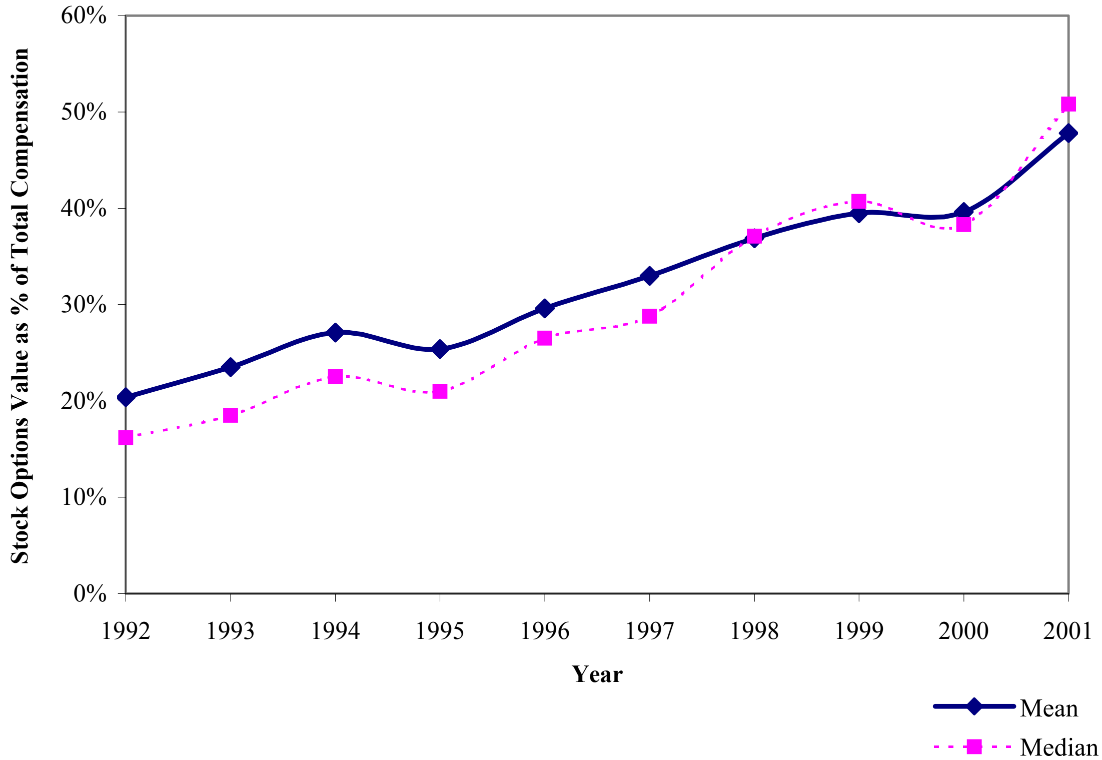
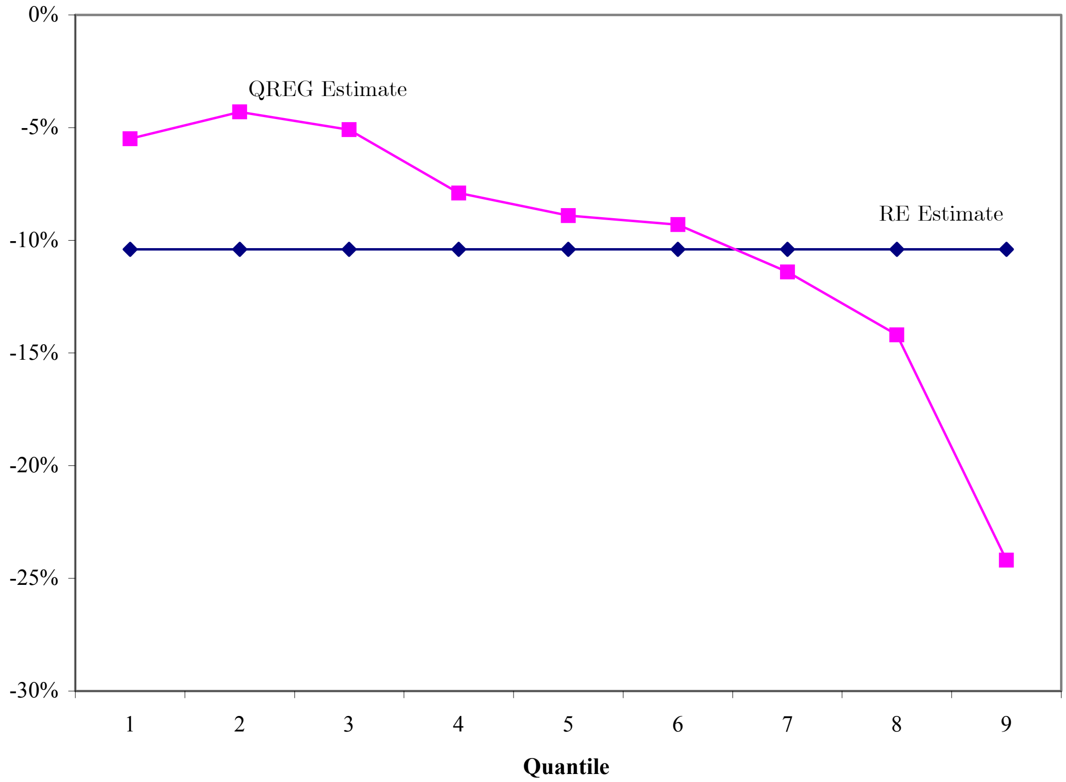
Figure 1: Stock-option Value as a Percentage of Total CEO Compensation, 1992-2001



Notes: The sample consists of an unbalanced panel with 10,249 observations for 2,159 firms in the 1992-2001 period. For each year we take the mean (thick line) and median (dotted line) of the ratio of the Black-Scholes value of stock options awarded to the CEO to the total CEO compensation. Total compensation is the sum of salary, bonus, benefits, restricted shares, LTIPs, and stock options.

Figure 2: Quantile Regression Estimates of the Effect of Union Presence on Natural Log of Total CEO Compensation



Notes: Quantile regression coefficients for UNION dummy variable from ten quantile regressions (at 0.10, 0.20, 0.30, 0.40, 0.50, 0.60, 0.70, 0.80, 0.90) are plotted along with the co-efficient from random-effects panel data estimation (horizontal line). The figure depicts the negative effect of union presence on total compensation. Each union effect is the coefficient from the respective quantile regression in each decile. All effects are significant at the 1 percent level, except for the first three deciles which are significant at the 5 percent level. Finally, the straight (horizontal) line represents the average union effect as estimated in the random-effects panel data regression.

Table 1: Variables Used in Our Analysis of the Union Effect on CEO Compensation

Variables	Definition	Mean	Std.Dev.
CEO Total Compensation	The sum of salary, bonus, benefits, B-S value of stock options, restricted grants and LTIPs – *in $million*	3.53	11.10
CEO Pay	Pay (salary & bonus) – *in $million*	1.08	1.59
CEO Stock Options	Black-Scholes value of CEO's stock options awarded – *in $million*	1.89	8.73
Union Presence	Dummy variable taking value 1, if the firm has at least one unionised establishment, and 0 otherwise.	0.36	0.48
CEO Ownership	CEO's stock ownership of the firm, as a percentage of the total outstanding shares (excluding options) – *in percentage form*	2.92	6.74
CEO Tenure	Tenure as CEO in the firm at the end of the fiscal year – *in years*	8.27	7.43
Assets	Firm size, in terms of total assets – *in $billion*	8.36	31.22
Return on Assets	ROA is defined as net income before extraordinary items and discontinued operations divided by total assets – *in percentage form*	3.87	14.70
Return to Shareholders	RET is defined as total return to shareholders, including the monthly reinvestment of dividends – *in percentage form*	28.80	617.81

Notes: Definitions and descriptive statistics for variables used in our analysis of CEO compensation. The sample contains an unbalanced panel of 10,249 firm-year observations in the 1992-2001 period, corresponding to 2,922 CEOs in 2,159 companies. Data was obtained from Bureau of National Affairs' labour databases and Standard & Poors' Execucomp. The number of observations in dependent variables (i.e. compensation variables) may differ due to logarithmic transformation. All level variables have been adjusted for inflation and are stated in 1992 dollars.

Table 2: Percentage Union Presence in our Sample by Industry, 1992-2001 (%)

	Firms			Firm-year observations		
Selected Industries	Union	Non-Union	Diff	Union	Non-Union	Diff
1. Mining & Construction	28.8	71.2	-42.4	35.2	64.7	-29.5
2. Manufacturing	39.4	60.6	-21.2	45.5	54.5	-9.0
3. Transportation, Communications & Utilities	40.5	59.5	-19.0	44.5	55.5	-11.0
4. Wholesale & Retail Trade	27.8	72.2	-44.4	31.8	68.2	-36.4
5. Finance, Insurance & Real Estate	7.3	92.7	-85.4	10.1	89.9	-79.8
6. Services	17.6	82.4	-64.8	21.3	78.7	-57.4
Observations	647	1504		3661	6541	
(%)	(30.1)	(69.9)	(-39.8)	(35.9)	(64.1)	(-28.2)

Notes: The cell values represent percentage of firms in our sample with and without union presence by industry. The third column for each period indicates differences between union and non-union percentages for each industry. Industries 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 sample contains an unbalanced panel with 10,249 observations from 2,159 firms. Forty-seven observations from eight firms in SIC 99 are excluded from this table, since they do not belong in any of the previously described industries.

Table 3: The Effect of Union Presence on CEO Pay, Stock Options and Total Compensation

	Ln(CEO Pay)	Ln(CEO Stock Option Value)	Ln(CEO Total Compensation)
Independent Variables	RE Panel data estimates	Tobit estimates	RE Panel data estimates
Intercept	8.308*** (49.61)	3.972*** (15.68)	5.748*** (22.33)
Union Presence	0.071** (2.14)	-0.281*** (-8.43)	-0.104*** (-2.65)
CEO Ownership	-0.018*** (-9.76)	0.003 (1.02)	-0.022*** (-11.00)
CEO Tenure	0. 010*** (7.21)	0.003 (1.58)	0.005*** (3.54)
Ln Total Assets	0.244*** (30.45)	0.428*** (43.45)	0.389*** (37.17)
Return on Assets	0.004*** (8.48)	0.0005 (0.59)	0.003*** (6.59)
Return to Shareholders	0.00002 (0.23)	0.00003* (1.92)	0.00003*** (2.75)
Industry effects (2-digit SIC)	No[28]	Yes	Yes
Year effects	No	Yes	Yes
Observations	10218	7467	10249
R^2 (or pseudo-R^2)	0.271	(0.127)	0.400
Hausman test (p-value)	6.47 (0.26)	__	16.95 (0.25)

Tests for equality of compensation variance between non-unionised and unionised firms

Ho: $\sigma(Y_{non}) = \sigma(Y_{union})$			
F-test	1.126***	1.028	1.135***
Levene's statistic	7.006***	0.132	6.722***
Ho: $\sigma(\hat{Y}_{non}) = \sigma(\hat{Y}_{union})$			
F-test	1.349***	0.835***	0.935**
Levene's statistic	60.692***	25.813***	13.527***

Notes: The sample consists of an unbalanced panel with 10,249 observations for 2,159 firms in the 1992-2001 period. All three dependent variables are in natural logarithm form. The values of *z*-statistics (*t*-statistics for Tobit estimation) appear in parentheses below each coefficient estimate. Number of observations differs due to logarithmic transformation. The *p*-value of Hausman specification tests is reported wherever applicable. The absolute values of the dependent variables are employed in the *F*-tests and Levene's tests for variance comparison between non-unionised and unionised firms. Asterisks denote significance at 1 percent (***), 5 percent (**), and 10 percent (*) levels. Finally, tests for the equality of compensation variance between non-unionised and unionised firms are performed using both *actual* and *predicted* compensation (in natural logarithm form).

[28] Random effect results with industry and time dummies do not change appreciably (results available upon request). We omit the specification with these dummies, however, as it fails the Hausman test

Table 4: IV estimates for Union Presence Effect on CEO Pay and Total Compensation

	Ln(CEO Pay)	Ln(CEO Total Compensation)
Independent Variables	IV estimates	IV estimates
Intercept	7.483***	5.962***
	(34.47)	(24.50)
Union Presence (from IV)	0.208+	-0.395**
	(1.45)	(-2.46)
CEO Ownership, CEO Tenure, Ln Total Assets, ROA, ROS	Yes	Yes
Industry effects (2-digit SIC)	Yes	Yes
Year effects	Yes	Yes
Observations	10218	10249
Centered R^2	0.324	0.390
Sargan test (p-value)	3.317	1.557
	(0.19)	(0.45)

Notes: As instrumental variables for union presence we employ three variables: RUST, REST and the natural logarithm of labour intensity. RUST is a dummy variable taking value 1 if the firm has its headquarters in a rust-belt state (namely IL, IN, MI, MN, NY, OH, PA, WI and WV), and 0 otherwise. RUST is a dummy variable taking value 1 if the firm has its headquarters in a state that does not belong to the bible-belt, sun-belt or rust-belt (namely AK, CO, CT, DC, DE, HI, IA, ID, MA, MD, ME, MT, ND, NE, NH, NJ, NM, RI, SD, UT and VT, as well as companies with headquarters outside the United States). Labour intensity is defined as the ratio of sales to employment. The sample consists of an unbalanced panel with 10,249 observations for 2,159 firms in the 1992-2001 period. The dependent variable is in natural logarithm form. The values of *z*-statistics appear in parentheses below each coefficient estimate. The *p*-value of the Sargan test for overidentification is reported in the last row. Asterisks denote significance at 1 percent (***), 5 percent (**), 10 percent (*), and 15 percent (+) levels.

Table 5: Summary Measures of Dispersion in Ln(Total CEO Compensation) Between Union and Non-Union Firms, by Firm Size and Strength of Union Presence Within Industry

	Union Presence Within Industry					
	Low			High		
A. All Firms	Union	Non-Union	Gap	Union	Non-Union	Gap
Std Deviation	1.12	1.17	-0.05	1.04	1.11	-0.07
90-10 gap	2.89	2.82	0.07	2.41	2.58	-0.17
90-50 gap	1.62	1.52	0.10	1.28	1.44	-0.16
Observations	763	2919		2897	3623	

	Union Presence Within Industry					
	Low			High		
B. Small Firms	Union	Non-Union	Gap	Union	Non-Union	Gap
Std Deviation	0.83	1.11	-0.28	0.75	0.91	-0.16
90-10 gap	2.09	2.54	-0.45	1.80	2.25	-0.45
90-50 gap	1.25	1.36	-0.11	0.89	1.23	-0.34
Observations	289	1516		1028	2263	

	Union Presence Within Industry					
	Low			High		
C. Large Firms	Union	Non-Union	Gap	Union	Non-Union	Gap
Std Deviation	1.04	1.05	-0.01	1.04	1.20	-0.16
90-10 gap	2.53	2.50	0.03	2.27	2.61	-0.34
90-50 gap	1.40	1.37	0.03	1.12	1.52	-0.40
Observations	474	1403		1869	1360	

Notes: Small firms are defined as having fewer than the median in total assets and large firms defined as having greater than the median in total assets. High union presence industries include Mining, Construction, Manufacturing, Transportation, Communications and Utilities (i.e. SIC 10-49). Low union presence industries include (Wholesale & Retail) Trade, Finance, Insurance and Real Estate and Services (i.e. SIC 50-89). These categorizations are taken from Table 2.

Table 6: CEO Pay-for-performance Elasticity in Union and Non-union Firms

Panel A: Separate OLS estimations of independent samples of union and non-union firms

	Δ ln(CEO Pay)		Δ ln(Total CEO Compensation)	
Independent Variables	Union	Non-union	Union	Non-union
Intercept	0.029*** (2.93)	0.042*** (4.65)	0.072*** (5.44)	0.096*** (7.26)
$\Delta \ln SV$	0.248*** (9.28)	0.210*** (10.50)	0.265*** (4.78)	0.324*** (11.52)
Observations	2697	4546	2702	4562
R^2	0.032	0.031	0.021	0.037
T-test that β differs between the two groups (p-value in parentheses)	1.150 (0.25)		-0.947 (0.34)	

Panel B: OLS estimations for the entire sample of firms with $(\Delta \ln SV_{it} \times UNION_i)$ interaction term

Independent Variables	Δ ln(CEO Pay)	Δ ln(Total CEO Compensation)
Intercept	0.037*** (5.49)	0.087*** (9.02)
$\Delta \ln SV$	0.211*** (10.58)	0.326*** (11.72)
$\Delta \ln SV \times UNION$	0.033 (1.04)	-0.067 (-1.13)
Observations	7243	7264
R^2	0.031	0.033

Notes: The sample consists of an unbalanced panel with 7,264 observations for 1,890 firms belonging in the S&P 1500 during the 1993-2001 period. The values of *t*-statistics from the OLS estimations (using the Huber/White robust estimator of variance) appear in parentheses below each coefficient estimate. Asterisks denote significance at 1 percent (***), 5 percent (**), and 10 percent (*) levels.

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