THE JOURNAL OF FINANCE • VOL. LXI, NO. 4 • AUGUST 2006

CEOs' Outside Employment Opportunities and the Lack of Relative Performance Evaluation in Compensation Contracts

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ABSTRACT

Although agency theory suggests that firms should index executive compensation to remove market-wide effects (i.e., RPE), there is little evidence to support this theory. Oyer (2004, *Journal of Finance* 59, 1619–1649) posits that an absence of RPE is optimal if the CEO's reservation wages from outside employment opportunities vary with the economy's fortunes. We directly test and find support for Oyer's (2004) theory. We argue that the CEO's outside opportunities depend on his talent, as proxied by the CEO's financial press visibility and his firm's industry-adjusted ROA. Our results are robust to alternate explanations such as managerial skimming, oligopoly, and asymmetric benchmarking.

AGENCY THEORY PREDICTS THAT THE MARKET-WIDE COMPONENT of a firm's returns should be removed from the compensation package since executives cannot affect the overall market by their actions and it is costly for an executive to bear the relative risks. Such market indexing of compensation is also referred to as relative performance evaluation (RPE). However, there is little empirical evidence in the literature of RPE (e.g., Antle and Smith (1986), Janakiraman, Lambert, and Larcker (1992)). In fact, a widespread feature of CEO pay packages, especially those which include stock option plans, is that they reward managers for stock price increases due to general market trends. Proponents of the optimal contracting view argue that rewarding chief executive officers (CEOs) for riding a bull market is optimal if the CEO's reservation wage stemming from outside employment opportunities varies with the economy's fortunes (Oyer (2004), Himmelberg and Hubbard (2000)). Assuming that CEO talent is scarce, demand for talented CEOs rises as the economy booms, in which case firms must pay CEOs more to retain them. In other words, allowing pay to increase with rising market levels during boom periods potentially enables firms to retain talented executives. A direct empirical test of the outside

*Rajgopal and Shevlin are from the University of Washington; Zamora is from Boston College. Rajgopal and Shevlin acknowledge support from the Accounting Development Fund at the University of Washington. Zamora acknowledges support from Boston College. We appreciate written comments from an anonymous referee, Luzi Hail, Frank Hodge, and Dawn Matsumoto, and verbal comments from James Jiambalvo, Eric Noreen, Jane Kennedy, and other workshop participants at the University of Washington. We thank Jennifer Francis, Allen Huang, and Amy Zang for graciously agreeing to share their CEO press citations data with us.

opportunities-based explanation for the lack of RPE is missing from the literature largely because CEO talent is typically difficult to observe.[1] We attempt to address this empirical void in the current paper.

We focus on a sample of S&P 500 firms for the period 1993 to 2001, and we collect data on two proxies for CEO talent in these firms: (i) the number of articles in the major U.S. and global business newspapers and wire services in which the CEO's name appears, as identified by searches of the Lexis/Nexis database; and (ii) past industry-adjusted return on assets (ROA) performance of the firm. We argue that more talented CEOs are likely to be cited and recognized by the business press more often than less talented CEOs. Moreover, talented CEOs are more likely than less talented CEOs to report superior industry-adjusted ROA performance. To ensure that the number of articles is not merely a reflection of CEO infamy as opposed to talent, we conduct three validation checks. First, we code the tone (i.e., favorable, unfavorable, or neutral) of coverage for 50 articles picked at random every year over 1993 to 2001 and find that the tone is favorable toward the CEO 95% of the time. Second, we find that the number of articles is correlated with CEOs appointed from outside the firm, a proxy for reputation used by Milbourn (2003). Third, the number of articles is highly correlated with explicit recognition of the CEO as determined by the inclusion in "Top CEO" lists compiled by various sources. Under the outside opportunities explanation, the propensity of firms to let CEO pay "ride the bull market" should increase with these proxies for CEO talent. That is, more talented CEOs should face less RPE.

However, finding results consistent with such a prediction does not necessarily single out the optimal contracting model as a satisfactory explanation for the absence of RPE. An important alternate hypothesis, put forward by several governance activists and economists, is that rent-seeking managers exploit captured governance mechanisms to set their own pay with little shareholder oversight (e.g., Crystal (1991), Bertrand and Mullainathan (2001), Bebchuk, Fried, and Walker (2002), Bebchuk and Fried (2003)). These critics argue that paying managers for market-wide increases in stock prices is inconsistent with the optimal contracting framework because such an arrangement only makes the compensation contract riskier (Holmstrom (1979)). To empirically distinguish the outside opportunities view from the rent-seeking view, we investigate whether the absence of RPE is increasing in proxies for poor governance and whether such an effect dominates the outside opportunities explanation for the lack of RPE.

The results, based on 2,343 CEO-firm-year observations drawn from 1993 to 2001, suggest that compensation committees do not practice RPE, that is, they do not completely filter out industry-wide or market-wide performance from the CEO's total compensation, which consists of stock- and non-stock-based

[1] An alternative test of the CEO talent retention hypothesis is to examine CEO turnover in up and down markets. However, such a test has low power because observed turnover occurs after the retention effects of allowing CEO pay to reflect market returns. That is, if the CEO talent retention hypothesis for the lack of RPE is valid, we expect to observe little ex post CEO turnover.

components.[2] Consistent with the outside opportunities explanation, we find that the sensitivity of CEO compensation to industry-wide and market-wide performance is systematically higher for CEOs who enjoy greater press visibility and superior industry-adjusted ROA during the prior 3 years. These results are consistent with the view that industry-wide and market-wide shocks increase demand for CEO talent outside the firm, which in turn, forces some firms to increase compensation levels to retain their more talented CEOs.

Further analyses test the robustness of this result to several alternate explanations. First, we find that a portion of the sensitivity of CEO compensation to market-wide shocks that is attributable to CEO talent persists after controlling for the governance quality of the firm, and as such, is robust to the skimming hypothesis mentioned earlier. Second, we show that the CEO talent explanation is incremental to Aggarwal and Samwick's (1999b) argument that lower RPE might be optimal in oligopolistic industries, to prevent rival managers from indulging in excess competition. Third, when we allow for Garvey and Milbourn's (2003) hedging explanation that RPE is lower for older CEOs and for CEOs with greater firm-specific wealth (as a proxy for the CEO's unobservable total wealth) because such CEOs can hedge the exposure of their compensation to market-wide shocks, our CEO talent explanation for the lack of RPE still generally holds.

In the last set of analyses, we evaluate first whether CEO pay is allowed to float with the market index only when the market is up but not when the market is down, and second whether such asymmetric benchmarking of CEO pay with the market index, which is interpreted as rent extraction by Bertrand and Mullainathan (2001) and Garvey and Milbourn (2004), is systematically higher or lower for talented managers. We find that CEO pay is shielded from a market downturn in most specifications although the results related to talented CEOs are mixed depending on the proxy for talent and the specific index chosen (industry index or market index). Thus, our results reveal no systematic pattern of asymmetric benchmarking for either talented or less talented executives, and thus suggest that rent extraction is not the dominant explanation for asymmetric benchmarking.

We contribute to the literature on executive compensation in several ways. Recent reviews of the compensation literature (Murphy (1999), Abowd and Kaplan (1999), Prendergast (1999)) identify the lack of RPE as an important unresolved puzzle. We provide evidence on an economically plausible explanation for the lack of RPE, namely compensation for the CEO's outside opportunities (Oyer (2004)). In an unpublished working paper, Himmelberg and Hubbard (2000) exploit Rosen's (1982) insight that talented managers will be employed by larger, more complex firms and empirically document that the sensitivity of CEO pay to systematic market-wide factors is greater for larger firms. Himmelberg and Hubbard (2000), in effect, proxy for CEO talent with

[2] A priori, we are not sure whether compensation committees employ either the industry return or the overall market return as the benchmark while considering RPE decisions. Hence, we report results related to both the industry and the market index.

firm size—a firm characteristic. Our proxies for CEO talent are more direct and are tied to the executive (e.g., media cites of a CEO's name and superior industry-adjusted performance under the CEO's tenure). Thus, our paper is among the first to show that a firm's use of RPE varies with the *executive*'s characteristics. Garvey and Milbourn (2003) also relate an executive's characteristics to the lack of RPE but they focus on the role of age and CEO wealth. Unlike Garvey and Milbourn (2003), we directly control for the rent extraction explanation and test whether the CEO talent explanation still holds.

Note further that Milbourn (2003) also relies on the number of CEO citations in the press as a proxy for CEO talent and shows that such a proxy is related to the observed pay-for-performance sensitivity (PPS). Our research question, whether CEO talent explains the absence of RPE, is different from Milbourn's question of whether PPS is linked to CEO talent. Nevertheless, we show that the empirical relation between CEO citations in the press and the absence of RPE is incremental to the link between citations in the press and PPS found by Milbourn (2003). Finally, we point out that asymmetric benchmarking by itself need not indicate rent extraction by managers (Bertrand and Mullainathan (2001), Garvey and Milbourn (2004)). Shielding CEO pay from industry or market downturns has been shown to be optimal in other contexts such as option repricing (e.g., Saly (1994), Acharya, John, and Sundaram (2000), Carter and Lynch (2001, 2004), Chidambaram and Prabhala (2003)).

The remainder of the paper is organized as follows. Section I discusses the background literature and lays out the CEO's outside opportunities hypothesis. Section II describes the sample, the empirical specifications, and the results of testing the CEO's outside opportunities hypothesis. Section III reports the robustness of the CEO's outside opportunities hypothesis to alternative hypotheses. Section IV concludes.

I. Background and Hypothesis

A. Prior Research

Several authors present strong economic arguments for assessing an agent's performance relative to a peer group if the performances of all agents are affected by a common factor such as industry-wide or market-wide returns (Baiman and Demski (1980), Diamond and Verrechhia (1982), Holmstrom (1979, 1982)). Holmstrom (1982) shows that as the number of peer agents becomes sufficiently large to form a precise estimate of the common factor and if the actions of each agent have no effect on the performance of any other agent, then the common uncertainty should be completely filtered out of the agent's performance. Thus, reducing the *systematic risk* attached to a common factor would allow for a better assessment of the agent's actions and in turn provide a stronger incentive. The prediction, also known as the strong form version of the RPE hypothesis, is that the common factor will be completely filtered out and performance evaluation will be a function of *only* the unsystematic component of the performance measure. A weaker form of the RPE hypothesis

argues that if the common factor is measured with error or if an agent could influence the performance of other agents, performance evaluation may be a function of both the systematic and unsystematic components of the performance measure.

For a more concrete understanding of the issue, consider a typical empirical specification wherein compensation is regressed on firm returns and a common factor proxied as returns to the industry:

$$\begin{aligned} \textit{Compensation} = \beta_0 + \beta_1 \textit{Firm returns} + \beta_2 \textit{Industry returns} \\ + \beta_3 \textit{Controls} + \textit{error}. \end{aligned} \tag{1}$$

In equation (1), β_1 is traditionally referred to as the pay-for-performance sensitivity (PPS) measure. Although equation (1) is written as though the common factor is industry returns, we can also think of the specification in terms of the common factor being market returns. Under the strong form of RPE, we expect β_2 to be negative and equal in magnitude to β_1 (or $\beta_1 + \beta_2 = 0$). Under the weak form of RPE, we expect β_2 to be negative but less than β_1 in absolute magnitude. These predictions are consistent with the idea that under the strong (weak) form of RPE, the compensation committee filters industry-wide or market-wide performance completely (partially) out of executive compensation.

The extant empirical evidence on both the weak and the strong form of RPE is mixed at best. Gibbons and Murphy (1990) find modest support for the weak form of RPE when they regress the change in the natural log of CEO compensation on the firm's stock returns and measures of market and industry return as proxies for peer performance. However, market and one-digit industry returns contribute a substantial portion of the explanatory power in the regression, suggesting that systematic risk is imposed on executives, or, in other words, executives are rewarded for market and industry increases. Several studies are consistent with a lack of evidence for the strong form of the RPE hypothesis and with relatively mild support for the weak form of RPE (Murphy (1985), Antle and Smith (1986), Barro and Barro (1990), Sloan (1993), Janakiraman et al. (1992)).

Evidence supporting RPE is somewhat stronger with respect to predicting CEO turnover. Several papers report an association between CEO turnover and net-of-market stock returns (e.g., Gibbons and Murphy (1990), Murphy and Zimmerman (1993), Pourciau (1993), Weisbach (1988)). Recently, Defond and Park (1999) show that RPE is more closely associated with CEO turnover in high-competition industries.

Corporate governance activists such as Crystal (1991), Rappaport (1999), and Bebchuk et al. (2002) interpret the limited evidence in favor of the RPE hypothesis as support for the rent extraction perspective, according to which managers exploit captured boards of directors to reward themselves for industry or economy-wide increases in stock prices. Bertrand and Mullainathan (2001) document that pay is as sensitive to exogenous luck as it is to firm-specific performance and that the association with exogenous luck is stronger when shareholders are diffused and arguably passive.

An alternative explanation for the limited evidence supporting RPE is the optimal contracting perspective. For example, Janikiraman et al. (1992) suggest that the lack of RPE to compensation arises because managers at higher levels might have incentives to forecast and respond to luck. They also suggest that in oligopolistic settings, strategic interactions among firms prevent a simple incorporation of the industry index into performance evaluation. Aggarwal and Samwick (1999b) argue that shareholders would be worse off if firms screened out industry-wide effects because such screening would provide managers an incentive to lower industry-wide returns by engaging in excessive competition, which in turn would lower profits.

B. RPE and CEO Outside Opportunities

Recently, Oyer (2004) and Himmelberg and Hubbard (2000) posit that we do not observe RPE in wage contracts because the value of the executives' outside opportunities is market sensitive. When we observe industry or market-wide increases in stock prices, the reservation wage of talented CEOs increases and firms pay their CEOs more simply to match their increased outside opportunities. This is because aggregate industry or market returns reflect, among other things, shocks to demand and productivity, and hence demand for managerial talent. If the supply of managerial talent is scarce and hence inelastic, then increases in aggregate industry or market returns should increase the equilibrium wage paid to CEOs. Moreover, if increases in aggregate industry or market returns raise the demand for CEO effort, and if such shocks are contractible, then the principal will demand more effort (or talent) and compensation levels have to rise to meet the agent's reservation utility. The models proposed by Oyer (2004) and Himmelberg and Hubbard (2000), from which the above arguments are drawn, critically depend on the assumption that executive turnover is costly and adjusting the parameters of the pay scheme after observing the state of the economy is onerous. In particular, the firm may initially pay a fixed wage, and upon being faced with an outside offer, the firm can adjust its wage to the market wage. In such a scenario, there is no need to make pay contingent on market-wide returns. However, revising a flat wage in response to an outside offer, as opposed to linking the wage to market returns, may lead to negotiation, information gathering, and turnover costs to the firm.

We provide a direct test of the CEO outside opportunities-based explanation in the current paper. In particular, we argue that talented CEOs should have greater outside opportunities, and we construct proxies for CEO talent using the frequency of media cites and recent prior industry-adjusted ROA firm performance. Our empirical prediction is as follows:

Hypothesis 1: The sensitivity of CEO pay to systematic market-wide factors is an increasing function of CEO talent.

We also account for the rent-extraction explanation, the strategic influence among oligopolistic industries explanation, hedging explanation, and the

asymmetric benchmarking explanation in our empirical tests as alternate explanations. These alternative explanations are developed below.

II. Sample and Data

A. CEO Talent Proxies

Our sample consists of officers named as CEOs of all S&P 500 companies over the 9-year period 1993 to 2001 in the ExecuComp database that are matched with firm-level data from COMPUSTAT and monthly stock return data from Center for Research in Security Prices (CRSP). We focus on S&P 500 firms for two reasons. First, talented mangers are likely to gravitate to this group of larger firms, wherein we expect to find the greatest support for the outside opportunities hypothesis. Second, we gather our CEO talent proxies through extensive hand collection of data. Hence, cost and time considerations require us to concentrate on only a subset of firms in the population. We exclude CEOs of subsidiaries and divisions.

Table I describes the data filters we use to compile the sample. Our initial sample of CEO-years with data on press cites of CEOs (*Articles*) consists of 464 firms and 3,487 CEO-firm-years after eliminating firms with missing data on either COMPUSTAT or CRSP. Computing industry-adjusted average ROA over the last 3 years (one of our two talent proxies) entails a loss of 1,144 CEO-firm-years. Of these 1,144 CEO-firm-years, we lose the most observations (855 CEO-firm-years) on account of our requirement that we compute ROAs over

Table I
Sample Selection

The initial sample of 464 firms and 3,487 CEO-firm years with data on CEO citations in the press (*Articles*) consists of officers named as CEO of all S&P 500 companies during the period 1993 to 2001 in the ExecuComp database after eliminating firms with missing firm-level data from COMPUSTAT and monthly stock return data from CRSP. The final sample of 403 firms and 2,343 CEO-firm-years with data on both *Articles* and industry-adjusted average ROA over the last 3 years (*IndadjROA*) consists of those observations wherein return on assets can be computed over the same CEO's tenure after eliminating CEO-firm-years with missing CEO-level data from ExecuComp and return data from CRSP.

	Firms	CEO-years
1993 to 2001 ExecuComp firm-years with non-missing data for S&P 500 firms and top 5 executives	513	28,250
Less: non-CEO executives	13	24,356
Less: missing COMPUSTAT data	25	280
Less: missing other data	11	127
Initial sample with data on *Articles*	464	3,487
Less: Firm-years in industry-years with <10 firms		30
Less: non-incumbent CEO-firm years		855
Less: missing ExecuComp data		186
Less: missing CRSP data		73
Final sample with data on both *Articles* and *IndadjROA*	403	2,343

the *same* CEO's tenure.[3] Thus, we obtain a smaller sample of 403 firms and 2,343 CEO firm-years for which both *Articles* and *IndadjROA* are available. To simplify the exposition, we only tabulate results run on these 2,343 CEO-firm-years because both proxies for talent (*Articles* and *IndajROA*) are available for this particular sample. However, we verify in untabulated results that all the inferences reported in the paper hold when we consider the larger sample of 3,487 CEO-firm-years for which only the *Articles* proxy for CEO talent is available.

Our observations are evenly divided among our sample period of 1993 to 2001; in particular, 1993 contributes 9.26% of the sample and 2000 contributes 12.73% of the sample. Furthermore, our sample covers a wide range of industries with the three largest industry groups in the sample coming from Electric, Gas, and Sanitary Services (SIC code 4900, 8.40% of the sample), Chemicals and Allied Products (SIC code 2800, 7.89% of the sample), and Depository Institutions (SIC code 6000, 7.51% of the sample).

For each CEO-year, we collect data on several variables related to CEO talent. One proxy is intended to capture whether external parties view the CEO as talented. The first of these external measures is the number of articles containing a CEO's name and company affiliation that appear in the major U.S. and global business newspapers and newswires in a calendar year.[4] The newspapers we consider are the *Wall Street Journal*, *New York Times*, *Washington Post*, *USA Today*, *Financial Times*, *Asian Wall Street Journal*, *Wall Street Journal Europe*, and *International Herald Tribune*. Our text search uses both the CEO's last name and company name (e.g., Akers and International Business Machines or IBM). We include an article only once, irrespective of how many times the CEO's name appears in the article. Based on Milbourn (2003) and Francis et al. (2004), we classify CEOs with larger values of press coverage (*Articles*) as more talented than CEOs with smaller values for this variable.[5] Descriptive data on *Articles* appears in Panel A of Table II. The median CEO receives 11 press citations in 1 year. This count is consistent with Milbourn (2003), who finds a median of 55 references to the CEO's name in the Dow Jones News Service (DJNS) over a 5-year period (see his table 2A). We perform three validation tests to ensure that the press visibility measures capture constructs such as CEO talent:

[3] Computing industry-adjusted ROA over three prior years restricts the final sample to firms in which the CEO has at least three prior years of seniority. To address the seniority issue and yet provide sufficient time for the CEO to stay in the job such that the industry-adjusted ROA can serve as a credible proxy for talent, we compute industry-adjusted ROA over the prior 2 years of the same CEO's tenure. As expected, the number of usable CEO-firm-years increases from 2,343 for the 3-year ROA sample to 2,754 for the 2-year ROA sample. In untabulated results, we find that the reported inferences hold even when the 2-year ROA sample is used.

[4] As our starting point, we use the CEO's name as reported in ExecuComp. To avoid understatements, we also search for shortened names (e.g., Bill for William), and to avoid overstatements (potentially associated with common names, such as Smith), we include the company name in the search.

[5] We thank Francis et al. (2004) for sharing their data with us.

Table II
Descriptive Statistics

The table presents descriptive statistics for the final sample of 403 firms and 2,343 CEO-firm-years with data on both CEO citations in the press (*Articles*) and industry-adjusted average ROA over the last 3 years (*IndadjROA)*, and consists of officers named as CEOs of all S&P 500 companies during the period 1993 to 2001 in the ExecuComp database after eliminating CEO-firm-years with missing data from COMPUSTAT and CRSP. The table presents the mean, standard deviation, and median for all variables including the talent proxies in Panel A, compensation data in thousands of dollars in Panel B, wealth data in millions of dollars in Panel C, and control and other variables in Panel D. Panel E presents OLS regression results to validate the CEO talent proxies. The dependent variable is the indicator variable *Top*, which is coded as one if the CEO is recognized as a top manager in lists published by the financial press. The independent variables are *IndadjROA* and the empirical cumulative distribution function (CDF) for *Articles* and *Size* (market value of equity). The first column of Panel E uses 3,487 CEO-firm-year observations with data on *Articles*. The second and third columns of Panel E use 2,343 CEO-firm-year observations with data on both *Articles* and *IndadjROA*. Sources and computations for all variables are described at the end of the table.

Variables	Mean	*SD*	$N = 2{,}343$ Median
Panel A: Talent Proxies			
Articles	17.81	28.40	11.00
Top	0.22	0.41	0.00
IndadjROA	0.68	0.20	0.70
Panel B: Compensation Data ($ thousands)			
Salary	816.75	370.94	795.80
Bonus	1,098.74	1,713.40	682.44
Other annual payments	62.01	229.36	0.00
Option grants (Black–Scholes)	5,617.47	20,460.92	1,628.22
Restricted stock grants	880.43	13,667.83	0.00
Long-term incentive payments	413.64	1,709.91	0.00
Other long-term compensation	298.33	2,887.85	51.96
Total flow compensation	9,187.38	25,580.40	4,283.36
ΔMarket value of stock	65,411.13	1,429,771.55	904.91
ΔMarket value of options	4,353.42	87,167.45	609.23
ΔMarket value of stock and options	69,764.54	1,474,862.21	2,530.54
ΔTotal compensation	78,951.92	1,477,389.64	7,609.44
Panel C: Wealth Data ($ millions)			
CEO firm specific wealth at the start of the year	237.47	1,863.32	28.65
ΔSW−firm	2,309.50	16,556.61	591.02
ΔSW−industry	1,902.89	11,305.53	626.56
ΔSW−market	333.17	2,857.21	232.47
Panel D: Other Variables			
Size ($ millions)	13,779.19	30,252.50	5,386.66
CEO age	57.82	6.25	58.00
Variance of SW ($ millions)	104.23	301.35	34.38
Tenure	9.39	6.86	7.43
Governance variables			
D_{gscore}	0.90	0.31	1.00
$D_{gscore} * g$ score	8.54	3.90	9.00
CEO-Chair	0.86	0.35	1.00
Onboard	0.34	0.19	0.29
Interlock	0.01	0.05	0.00
Meetings	7.33	2.82	7.00
Herfindahl–Hierschman Index	0.20	0.15	0.15

(*continued*)

Table II—*Continued*

Panel E: Validation of Talent Proxies			
Variables	Top	Top	Top
Intercept	144.93 × 10e-3*	86.19 × 10e-3†	61.85 × 10e-3†
F(Articles)	3.22 × 10e-3*		2.07 × 10e-3*
IndadjROA		190.21 × 10e-3†	89.22 × 10e-3†
F(Size)			0.00 × 10e-3*
N	3,487	2,343	2,343
Adjusted-R^2	0.066	0.008	0.151

Variable definitions are as follows

Articles	The number of articles containing the CEO's name and company affiliation that appear in the major U.S. and global business newspapers and newswires in the year preceding the sample year. An article is included only once irrespective of how many times the CEO's name appears in the article.
Top	Equals 1 if the CEO is recognized as a top manager in any one of the following lists: *Business Week*'s "The Best Managers" (available for 1992–2001), *Worth*'s "The Best CEOs" (available for 1998–2001), the *Financial Times'* "World's Most Respected Business Leaders" (available for 1997–2001), *Time*'s "The Time/CNN 25 Most Influential" (available for 2001), and *Fortune*'s "The 50 Most Powerful Women in Business" (available for 1998–2001) and "Most Powerful Black Executives in America" (available for 2001, and zero otherwise).
IndadjROA	Average of the empirical CDF of within industry-year rankings for the prior 3 years.
Salary, bonus, and other annual payments	ExecuComp data SALARY, BONUS, and OTHANN, respectively.
Option grants	Estimated grant-date value of options granted during the year using a modified version of the Black–Scholes (1973) model. The exercise price, stock price at the grant date, number of securities granted, and time to maturity are obtained from ExecuComp. Risk-free rate of interest is the approximate average yield in the data year from a 7-year U.S. Treasury bond. Expected stock return volatility is estimated as the annualized standard deviation of daily stock returns over the 120 trading days preceding the end of the fiscal year of the option grant.
Restricted stock grants, long-term incentive payments, and other long-term compensation	ExecuComp data RSTKGRNT, LTIP, and ALLOTHTOT, respectively.
Total flow compensation	Sum of salary, bonus, other annual payments, option grants, restricted stock grants, long-term incentive payments, and other long-term compensation.
ΔMarket value of stock	Number of shares held by the CEO at the start of the year (including restricted stock) multiplied by the stock return (including dividends) during the year.
ΔMarket value of options	Number of options held by the CEO at the start of the year, multiplied by an estimate of the change in the average option value during the year. The average option value is estimated based on the method presented by Core and Guay (2002), except that the time to exercise is equal to 70% of the Core and Guay assumed time to maturity, and the value at the start (end) of the year is based on the start (end) of the year stock price, as adjusted for stock splits during the data year. All other parameters used are as described for new option grants above.

(*continued*)

Table II—*Continued*

Variable definitions are as follows	
ΔTotal compensation	Sum of total flow compensation, ΔMarket value of stock and ΔMarket value of options.
CEO-firm-specific wealth at the start of the year	Sum of beginning-of-year stock and options, where the number of shares (options) held by the CEO at the start of the year is valued at the share price (option values) as of the beginning of the year.
ΔSW − firm, industry, market	Change in shareholder wealth (i.e., beginning-of-year market value) multiplied by the firm's return, industry index return, and market index return, respectively.
Size (MVE)	Number of common stock shares outstanding times the stock price at the beginning of the year.
CEO age	CEO's age in the sample year.
Variance of SW	Beginning-of-the-year market value times the variance of monthly stock returns, expressed in decimal form (not percentage), over the 60 months preceding the sample year.
Tenure	Number of years the CEO has held the top-ranking position in the firm
D_{gscore}	Equals 1 if a firm's g score exists in the list from Gompers et al. (2001), and zero otherwise.
$D_{gscore} * g$ score	D_{gscore} multiplied by the g score. G score is a measure of shareholder power.
CEO-Chair	Equals 1 if the CEO is also the chairman of the board, and zero otherwise.
Onboard	Proportion of the top five officers that serves on the board calculated as the average of cases in which the dummy variable for EXEC_DIR as per ExecuComp is "TRUE" across officers of a firm for each firm-year.
Interlock	Proportion of the top five officers that is subject to interlock relations calculated as the average of cases in which the dummy variable for PINTRLOC flag, as per ExecuComp is "TRUE" across officers of a firm for each firm-year.
Meetings	Number of board meetings during the year.
Herfindahl-Hirschman Index	Sum of an industry's (two-digit SIC) squared market shares (in percentages) as in Defond and Park (1999).

* and † represent statistical significance at p-value < 0.1 and 0.05, respectively, and is two tailed.

1. *Tone of coverage*: We randomly select 50 articles every year over the 9-year period of 1993 to 2001, and we classify the tone of the articles with regard to the CEO as favorable, unfavorable, or neutral. More than 95% of the articles are neutral or favorable.
2. *Prior position*: Milbourn (2003) argues that CEOs appointed from outside the firm are better-reputed CEOs because the hurdle for hiring an outside CEO is higher than for hiring an inside CEO since insiders have the advantage of possessing firm-specific knowledge. We find that the (untabulated) correlation between *Prior*, a dummy that is set equal to one if the CEO is hired from outside, and *Articles*, is 0.06 (p-value < 0.01). Thus, *Articles* captures some of the outside-hire aspect of CEO talent.
3. *Explicit recognition*: Francis et al. (2004) show that the press visibility measure is positively correlated with the explicit recognition of the CEO

as a top manager by business publications. In particular, the authors code a *Top* dummy as equal to one (zero) for a fiscal year if the CEO is recognized (not recognized) as a top manager in any one of the following lists: *Business Week*'s list of "The Best Managers" (available 1992–2001), *Worth*'s list of the "The Best CEOs" (available 1999–2001), the *Financial Times*' list of the "World's Most Respected Business Leaders" (available 1998–2001), *Time*'s list of "The Time/CNN 25 Most Influential" (available for 2001), and *Fortune*'s lists of the "The 50 Most Powerful Women in Business" (available 1998–2001) and the "Most Powerful Black Executives in America" (available 1998–2001).[6] Descriptive data in Panel A of Table II show that 2% of the CEOs in the sample are rated as "Top CEOs" by newspapers and magazines. As reported in Panel E, Table II of our paper, in a regression of *Top* on *Articles*, the coefficient on *Articles* is positive and reliably significant (coefficient = 0.00322, p-value = 0.01).

Beside the press visibility proxy, we use industry-adjusted ROA (*IndadjROA*) for the prior 3 years under the CEO's tenure as a proxy for CEO talent. In particular, we compute a firm's ROA as income before extraordinary items (COMPUSTAT Item 18) scaled by average total assets. We delete firms with missing ROA observations and only retain industries for which we find at least 10 firms in a two-digit SIC code for a given year. We compute the cumulative distribution function (CDF) of ROA for each CEO-firm-year by industry and then calculate the 3-year average of the CDF rank of ROA. If the *same* CEO has not been in office for the last 3 years, we use an average ROA computed over the CEO's tenure.[7] Higher (lower) ranks suggest that the CEO outperformed (underperformed) the industry. Regressions in Panel E of Table II show that *Top* is positively related to *IndadjROA* and the association is statistically significant. The positive associations of the two main talent proxies, *Articles* and *IndadjROA*, with *Top* validate these proxies as good measures of CEO talent. Note that we do not use *Top* as a talent proxy in our empirical tests because *Top* is an indicator variable and 78% of the *Top* observations in the data set are coded as zero (as shown in Panel A of Table II). Thus, *Top* does not have as much power to detect the presence of the CEO talent effect in the data while *Articles* and *IndadjROA* are continuous variables with a reasonable degree of cross-sectional variation and hence have greater power to test the talent hypothesis for the lack of RPE. Moreover, *Top* is likely to be correlated with both

[6] Three of the lists provide explicit guidance as to the criteria examined. *Worth*'s list is based on interviews with Wall Street analysts and fund managers, and identifies the top CEOs in terms of delivering long-term shareholder value and high integrity. The *Financial Times* list is based on survey evidence from CEOs around the world, who are asked to identify the three business leaders they admire and respect most. *Fortune* evaluates women executives on four measures: Revenues and profits she controls, the importance of her business in the global economy, the arc of her career, and her impact on culture and society.

[7] As a sensitivity check, we drop CEOs who have not been in office for the last 3 years from the 3-year ROA sample and find that the reported inferences remain unchanged.

Articles and *IndadjROA* and introducing *Top* as a talent proxy together with *Articles* and *IndadjROA* would deprive these variables of power.[8]

Another observation about the *Articles* proxy also deserves mention. Untabulated results show that *Articles* is not highly correlated with our compensation variable, Δ*Total Compensation,* which is discussed below in Section II.B (correlation = 0.04, p-value < 0.05), or with *IndadjROA* (correlation = 0.08, p-value < 0.01). Thus, the press citations proxy, *Articles*, is not merely a transformed version of the compensation or performance variables, devoid of any economic or statistical signal of its own.

B. Compensation Data

The compensation measure we use in the empirical tests is the change in the CEO's firm-related wealth (Δ*Total Compensation*), following Aggarwal and Samwick (1999a, 1999b) and Himmelberg and Hubbard (2000). The variable Δ*Total Compensation* is measured as the sum of three components: (i) cash compensation, the sum of salary, bonus, and other annual cash payouts over the year; (ii) the Black–Scholes value of options granted and the market value of restricted stock granted during the year and other long-term incentive payouts; and, (iii) the change over the year in the market value of equity and options held by the CEO at the beginning of the year. We use equity and option levels held at the beginning of the year to avoid confounding issues introduced by equity sales and option exercises during the year.

Descriptive data reported in Panel B of Table II reveal that the median CEO receives an annual salary and bonus of $795,800 and $682,440, respectively. As is well known, cash compensation is not the major source of incentives provided to top management. The median CEO receives $1.628 million in option grants and no restricted stock. The change in the market value of equity and options held by the median CEO is $2.531 million. Unreported analyses indicate that when the bull market years starting from 1998 are dropped from the sample, the median change in the market value of equity and options drops to $669,000.

Note that the mean CEO enjoys a staggering average annual appreciation of $69.764 million in his option and stock portfolio. Further analysis reveals considerable right skewness in the data, indicating that most of this appreciation is attributable to the bullish stock market in the year 1999 and to a few CEOs who have large ownership stakes in their firms, such as William Gates III of Microsoft, Larry Ellison of Oracle, Michael Dell of Dell Corporation, Scott McNealy of Sun Microsystems, and Robert Goizueta of Coca Cola Corporation.

[8] In untabulated analyses, we introduce *Top* as a third proxy for CEO talent in model (3) along with *Articles* and *IndadjROA*. Consistent with expectations, we find that the results are somewhat mixed. The interaction of the Shareholder wealth-industry index and *Top* is not significant but the interaction of the Shareholder wealth-market index and *Top* is unexpectedly negative and statistically significant. However, the interaction of Shareholder wealth-index with either *Articles* or *IndadjROA* continues to remain positive and significant, regardless of whether the industry or the market index is considered.

The presence of such outliers in the data set motivates our use of median regressions in the paper (e.g., Aggarwal and Samwick (1999a, 1999b), Milbourn (2003)). Panel C provides further evidence on the size of CEO wealth in their firms. CEO wealth is measured at the start of the year for the purposes of hypothesis testing. The median and mean CEO firm-specific wealth is $28.65 million and $237.47 million, respectively.

Our tests of the outside opportunities hypothesis and RPE rely on cross-sectional variation in the pay-for-performance sensitivity (PPS). We compute dollar returns to shareholders to assess PPS consistent with several recent papers (Aggarwal and Samwick (1999a, 1999b), Milbourn (2003), Garvey and Milbourn (2003)). In particular, we compute three measures of dollar returns: (i) Δ*Shareholder wealth-firm,* measured as the market value of the firm at the beginning of the sample year multiplied by the percent returns for the firm, where both the returns and the market value of the firm are extracted from ExecuComp and, (ii) two market-wide measures of performance, where Δ*Shareholder wealth-industry* (Δ*Shareholder wealth-market*) is measured as the market value of the firm at the beginning of the sample year multiplied by the percent returns earned by other firms in the two-digit SIC code (S&P 500 market index). In other words, Δ*Shareholder wealth-firm* is the *realized* dollar return to shareholders for a firm-year while Δ*Shareholder wealth-industry* (*ΔShareholder wealth-market*) is the *hypothetical* dollar return to shareholders if the firm had earned the return offered by the industry (S&P 500) index. Industry returns and S&P 500 index returns are calculated using CRSP tapes. Note that we use the two measures of market-wide performance, namely, the industry and S&P 500 market index, for our tests because we do not know for sure whether firms benchmark CEO pay to one or the other of these specific indexes. We acknowledge that industry and S&P 500 could potentially capture different underlying phenomena that are not explicitly modeled or discussed in our paper.

Descriptive data related to these dollar returns are presented in Panel C of Table II. The median dollar return for a firm is $591.02 million compared to the median industry-based return of $626.56 million and a market-based return of $232.47 million. As expected, market based-shareholder returns have a lower standard deviation than firm- or industry-based dollar returns. Panel D reports descriptive data on the variance of shareholder wealth and firm size. As expected, the median firm is large with a market capitalization of $5.386 billion. Panel D also provides descriptive data on the governance structure of these firms. Details of the governance data are discussed in Subsection III.B.

C. Empirical Specification

The general empirical specification employed in the paper is:

$$\begin{aligned}\Delta \textit{Total CEO Compensation}_{jt} = {} & \beta_0 + \beta_1 \Delta \textit{Shareholder Wealth-Firm}_{jt} \\ & + \beta_2 \Delta \textit{Shareholder Wealth-Industry}_{jt} \\ & + \beta_3 \Delta \textit{Shareholder Wealth-Industry} * \textit{CEO Talent}_{jt}\end{aligned}$$

$$+ \beta_4 \Delta \textit{Shareholder Wealth-Firm} * \textit{Controls}$$
$$+ 2\textit{-digit industry dummies}_j$$
$$+ \textit{Year dummies}_t + \textit{error}_{jt}. \quad (2)$$

We also estimate equation (2) after substituting the "Market Index" for the "Industry Index." In equation (2), β_1 is the pay-for-performance sensitivity, PPS. As noted earlier and consistent with Janikiraman et al. (1992), if the strong form of RPE holds in the data, we expect β_2 to be negative and equal in magnitude to β_1 (or $\beta_1 + \beta_2 = 0$). However, if the weak form of RPE describes the data, we expect β_2 to be negative but less than β_1 in absolute magnitude. These predictions are consistent with the notion that under the strong (weak) form of RPE, the compensation committee filters market-wide performance completely (partially) out of executive compensation. If the talent explanation accounts for why CEOs are not subject to RPE, we expect β_3 to be positive.[9] Our proxies for CEO talent are *Articles* and *IndadjROA*.

Extant research suggests that PPS is a function of CEO age, firm size, and variability of dollar returns. Gibbons and Murphy (1992) show that PPS increases with CEO age. Baker and Hall (2000) and Core and Guay (2001) point out that the CEO's marginal product varies with firm size, and hence PPS and firm size may be related. Aggarwal and Samwick (1999a) report that PPS decreases with the variance in shareholder dollar returns. Consistent with their work, we compute the variance of dollar returns for each firm. In particular, we multiply the market value of the firm at the beginning of the sample year with the variance of monthly returns, where monthly returns to shareholders are calculated over the 60 months preceding the sample year. We add two-digit industry dummies and year dummies to control for differences in pay levels across industries and time (Murphy (1999)). We also introduce the interaction of Δ*Shareholder Wealth-Industry* $*$ *CEO Tenure* as a control variable in the empirical tests. We do not interpret the sign on this coefficient because of two conflicting arguments. On the one hand, CEOs who have longer tenure are more likely to be talented and hence they should experience lower RPE if our CEO talent hypothesis holds. On the other hand, CEOs with longer tenure are likely to have captured their boards of directors and compensation committees. Hence, an absence of RPE, if found, could indicate managerial skimming, and therefore may not necessarily be evidence in support of the CEO talent explanation.

Thus, our main empirical specification is as follows:

$$\Delta \textit{Total CEO Compensation}_{jt}$$
$$= \beta_0 + \beta_1 \Delta \textit{Shareholder Wealth (SW)-Firm}_{jt}$$
$$+ \beta_2 \Delta \textit{Shareholder Wealth-Industry}_{jt}$$

[9] The talent explanation presumes that strong form RPE does not exist in the data. The talent explanation is of course, moot, if the strong form version of RPE holds.

$$
\begin{aligned}
&+ \beta_3 \Delta Shareholder\ Wealth\text{-}Industry_{jt} * F(Articles)_{jt} \\
&+ \beta_4 \Delta Shareholder\ Wealth\text{-}Industry_{jt} * Indadj\ ROA_{jt} \\
&+ \beta_5 \Delta Shareholder\ Wealth\text{-}Industry_{jt} * F(CEO\ Tenure)_{jt} \\
&+ \beta_6 \Delta Shareholder\ Wealth\text{-}Firm_{jt} * F(CEO\ Age)_{jt} \\
&+ \beta_7 \Delta Shareholder\ Wealth\text{-}Firm_{jt} * F(Size)_{jt} \\
&+ \beta_8 \Delta Shareholder\ Wealth\text{-}Firm_{jt} * F(Variance\ of\ shareholder\ wealth)_{jt} \\
&+ 2\text{-}digit\ industry\ dummies_j + Year\ dummies_t + error_{jt}. \qquad (3)
\end{aligned}
$$

As indicated above, we estimate equation (3) after substituting the "Market Index" for the "Industry Index." To facilitate the intuitive interpretation of the economic significance of our results, we follow Aggarwal and Samwick (1999b) and construct the cumulative distribution functions (CDFs) of certain empirical variables. In particular, any variable on which the sensitivity to dollar returns depends is first normalized according to its empirical CDF and then interacted with the appropriate dollar returns variable.[10] Normalizing the interactive raw variables to the unit interval enables ease of interpretation and reduces the impact of extreme outliers on the regression specification. To illustrate, the estimated pay for market-wide performance for a given level of Δ*Shareholder Wealth-Industry*, conditioned on *Articles*, is $\beta_2 + \beta_3\ F\ (Articles)$, where $F(\cdot)$ refers to the CDF of *Articles*. Thus, the range of coefficients is β_2 for the CEO with the least number of citations in the press and $\beta_2 + \beta_3$ for the CEO with the greatest number of citations in the press. The coefficients related to the CEO with the median number of citations in the press is $\beta_2 + 0.5\beta_3$. If the CEO talent explanation for the lack of RPE is descriptive of the data, we expect positive coefficients on the interaction between pay for market (or industry-wide) movements and the CEO talent proxies, that is, β_3 and β_4. Note that *IndadjROA* is inherently a rank measure, as described earlier, and hence is not expressed as a CDF.

Another important methodological issue is that all specifications in the paper are estimated using median regressions. The median is more robust than the mean to the presence of large outliers. Median regressions minimize the sum of absolute deviations rather than the sum of squared deviations. Moreover, because the median is also a more robust estimate of central tendency than the mean, the precision of estimates from a median regression is also higher. In untabulated sensitivity tests, we reestimate all regressions using Ordinary Least Squares (OLS) with and without winsorizing the data at the 1% and the 99% levels. Unreported results of the OLS regressions after winsorization yield inferences similar to the reported median regressions, but the results without winsorization are different. Moreover, consistent with the considerable

[10] In practice, we rank all observations for a variable in ascending order and then scale each rank by N such that each observation assumes a value between 0 and 1.

right skewness of the Δ*Total Compensation* data, the PPS estimates using the median regressions are much smaller than those obtained via OLS.[11]

Results of estimating equation (3) are presented in Table III. Columns (1) and (2) assess the existence of RPE with respect to the industry index and the market-wide S&P 500 index, respectively. As mentioned before, we report results only for the most restrictive sample of 2,343 CEO-firm-years, for which *Articles* and *IndadjROA* are both available. We reiterate that untabulated analyses confirm that all the inferences reported in the paper apply to the larger sample of 3,487 CEO-firm-years for which usable data on *Articles* exists.

Column (1) shows that the estimated PPS for a CEO in our sample with minimum *F(Articles)* and *IndadjROA* is $8.00 for a $1,000 increase in the market value of the firm. These estimates are much larger than the $3.25 per $1,000 increase in shareholder wealth estimated by Jensen and Murphy (1990). When we estimate the reduced model in column (1) with only the two terms Δ*Shareholder Wealth-Firm*$_{jt}$ and Δ*Shareholder Wealth-Industry* (and industry and year controls), we find that the estimated PPS is $3.24 per $1,000. Thus, consistent with prior work (e.g., Aggarwal and Samwick (1999a), Himmelberg and Hubbard (2000)), PPS estimates tend to be higher when controls for variance of shareholder wealth, size, and age are introduced.

In column (1), the coefficient on Δ*Shareholder Wealth-Industry* is –2.36 while the sum of the coefficients on Δ*Shareholder Wealth-Firm* and Δ*Shareholder Wealth-Industry* is not equal to zero (15.82 − 2.36, p-value = 0.00), suggesting evidence in favor of the weak form of RPE. Of particular interest is the finding that RPE usage decreases in *Articles* and *IndadjROA* in column (1), our proxies for CEO talent, consistent with Hypothesis 1.

The panel at the bottom of Table III reports PPS evaluated at median values of control variables. For example, median PPS is $8.70 per $1,000 increase in shareholder wealth under column (1). This PPS is computed by evaluating the coefficients of the regression at median values of the CDFs of the independent variables. To derive PPS at the minimum and maximum values of the talent variables, we retain all variables at median values, except for *F*(*Articles*) and *IndadjROA*, which are varied to assume minimum and maximum values. The objective is to assess the effect on PPS, including the effect of benchmarking, for executives with various ranges of talent. For example, we

[11] For example, an OLS estimation of equation (3) without data winsorization yields

$109.12 * \Delta$*Shareholder Wealth-Firm*$_{jt}$ – $163.46 * \Delta$*Shareholder Wealth-Industry*$_{jt}$ – $3.35 * \Delta$ *Shareholder Wealth-Industry*$_{jt} * F$ (*Articles*)$_{jt}$ + $101.74 * \Delta$*Shareholder Wealth-Industry*$_{jt} *$ *IndadjROA*$_{jt}$ + *Controls*$_{jt}$ (other coefficients not reported, all reported coefficients significant at the 1% level except –3.35, which has a t-statistic of –0.35).

After winsorization, we obtain

$10.63 * \Delta$*Shareholder Wealth-Firm*$_{jt}$ – $266.98 * \Delta$*Shareholder Wealth-Industry*$_{jt}$ + $63.72 * \Delta$*Shareholder Wealth-Industry*$_{jt} * F$ (*Articles*)$_{jt}$ + $143.38 * \Delta$*Shareholder Wealth-Industry*$_{jt} *$ *IndadjROA*$_{jt}$ + *Controls*$_{jt}$ (other coefficients not reported, all reported coefficients at the 1% level).

Note the big differences in the magnitude of the OLS coefficients relative to the median regression based coefficients reported in Table III.

Table III

Median Regressions of Change in CEO Compensation on Talent Proxies and Control Variables

The table presents median regression results for the final sample of 2,343 CEO-firm-years with data on both CEO citations in the press (*Articles*) and industry-adjusted average ROA over the last 3 years (*IndadjROA*), and consists of officers named as CEOs of all S&P 500 companies during the period 1993 to 2001 in the ExecuComp database after eliminating CEO firm-years with missing data from COMPUSTAT and CRSP. Median regressions minimize the sum of absolute deviations rather than the sum of squared deviations. The dependent variable is the change in total compensation (sum of total flow compensation and change in the market value of the CEO's stock and options). The independent variables include the change in the shareholder-wealth (beginning-of-the-year market value of the firm multiplied by percent returns for the firm), the change in the shareholder-wealth index (beginning-of-the-year market value of the firm multiplied by percent returns earned by other firms in the same industry for column (1) or by percent returns earned by the S&P 500 market for column (2); the change in shareholder wealth index interacted with *IndadjROA*, the empirical cumulative distribution function (CDF) for *Articles* and CEO tenure, and the change in shareholder wealth interacted with the CDF for CEO age, the market value of equity (*Size*), and beginning-of-the-year market value times the variance of monthly stock returns, expressed in decimal form over the 60 months preceding the sample year (*Variance of SW*). Column (1) shows that the estimated pay-for-performance (PPS) for a CEO in our sample with a minimum/median/maximum F(*Articles*) and *IndadjROA* is \$8.00/\$8.70/\$9.25 for a \$1,000 increase in the market value of the firm. The PPS is computed by evaluating the regression coefficients at the median values of the CDF of the control variables.

	Dependent Variable: ΔTotal Compensation		
Independent Variables	Predicted Sign	Industry Index (1)	Market Index (2)
ΔSW − firm	+	15.82*	14.30*
ΔSW − index	$\beta_2 < 0$, weak RPE, $\beta_1 + \beta_2 = 0$, strong RPE	−2.36*	−6.83*
ΔSW − index ∗ *F*(Articles)	+H1	0.77*	1.56*
ΔSW − index ∗ IndadjROA	+H1	0.55*	3.75*
Controls			
ΔSW − index ∗ *F*(Tenure)	?	3.22*	4.93*
ΔSW − firm ∗ *F*(CEO age)	?	−2.67*	−0.81*
ΔSW − firm ∗ *F*(Size)	?	−27.66*	−28.34*
ΔSW − firm ∗ *F*(Variance of SW)	?	16.10*	17.80*
Sample Size		2,343	2,343
Psuedo-R^2		0.092	0.091
Test Strong RPE: $\beta_1 + \beta_2 = 0$ (*p*-value)		0.000	0.000
Estimated Pay Sensitivities Assuming Media Control Variables			
Minimum *F*(Articles) and IndadjROA		\$8.00	\$4.58
Median *F*(Articles) and IndadjROA		\$8.70	\$7.66
Maximum *F*(Articles) and IndadjROA		\$9.25	\$9.56

* represents statistical significance at *p*-value < 0.1, and is one tailed when the coefficient sign is predicted, and two tailed otherwise. Coefficients on intercept, industry, and year dummies are suppressed. Since the cumulative distribution functions (CDFs) of variables are computed on a yearly basis, estimated pay sensitivities are evaluated at median values of approximately 0.50 for the CDFs of the variables tenure, CEO age, size, and variance of shareholder wealth (SW). Market Index (2) is evaluated at the 0.49 quantile.

observe that the sensitivity of pay to the industry index for a CEO with median *Articles* and *IndadjROA* under column (1) is \$8.70, and that sensitivity ranges from \$8.00 for a CEO with minimum *Articles* and *IndadjROA* to \$9.25 for a CEO with maximum *Articles* and *IndadjROA*.

A similar picture emerges from column (2), where the proxy for peer performance is the S&P 500 market index. The coefficient on Δ*Shareholder*

Wealth-Market Index in column (2) is –6.83 and the sum of the coefficients on the Δ*Shareholder Wealth-Firm* and Δ*Shareholder Wealth-Market* index is reliably different from zero (14.30 − 6.83, p-value = 0.00), consistent with weak form RPE. Note that the magnitude of the coefficient on the Δ*Shareholder Wealth-Market* index (−6.83) is larger than the magnitude of the coefficient on the Δ*Shareholder Wealth-Industry* index (−2.36), consistent with Gibbons and Murphy (1990).

As before, the extent of RPE diminishes with proxies for CEO talent, namely, *Articles* and *IndadjROA*. In fact, as shown in the lower panel in the table corresponding to column (2), the spread in the absence of RPE due to talent proxies is higher when the market index is the peer performance measure used. We observe that the sensitivity of pay to the market index for a CEO with median *Articles* and *IndadjROA* is \$7.66, and that sensitivity ranges from \$4.58 for a CEO with minimum *Articles* and *IndadjROA* to \$9.56 for a CEO with maximum *Articles* and *IndadjROA*.[12] Thus, the range of extreme PPS values, scaled by the median-talented CEO's PPS, is reasonably large, varying from 59.8% (\$4.58/\$7.66) to 124.8% (\$9.56/\$7.66). This reasonably wide range suggests that the impact of CEO talent on the absence of RPE is likely to be economically significant. Thus, the empirical evidence is consistent with the argument that the absence of RPE is positively related to CEO talent.

Incidentally, it is interesting to note that CEOs with longer tenure have lower RPE in their compensation contracts. However, assigning an unambiguous interpretation to this finding is difficult, as discussed earlier. Turning to the control variables, PPS decreases with firm size and CEO age, but contrary to prior work (Aggarwal and Samwick (1999a)), PPS appears to increase with the variance of shareholder returns in our sample.[13] In summary, the evidence presented in Table III is consistent with Oyer (2004) and Himmelberg and Hubbard's (2000) theory that RPE is optimally lower for talented CEOs.

III. Alternate Explanations

A. CEO Talent and PPS

Milbourn (2003) finds that CEO talent, also proxied by a count of citations in the press for the CEO, is related to the observed pay-for-performance sensitivity, PPS. Although his paper addresses a different research question (the link

[12] As a benchmark, note that Aggarwal and Samwick (1999a) report PPS estimates ranging from \$6.59 to \$14.52 per \$1,000 for the sample period 1993 to 1996 while Milbourn (2003) reports PPS ranging from \$1.35 for the largest firm to \$25.44 for the smallest firm. Somewhat smaller PPS estimates in our study are not surprising because our sample consists of S&P 500 firms and PPS estimates tend to be smaller for larger firms, whereas Aggarwal and Samwick (1999a) and Milbourn (2003) focus on a much broader sample of S&P 1500 firms.

[13] Whether PPS is related positively or negatively to firm risk and whether such an inference is due to the interaction of PPS and size is a subject of much debate in the literature (Aggarwal and Samwick (1999a), Core and Guay (2001), Prendergast (2000)). We do not take a position on this issue because the interaction between PPS and (i) size and (ii) variance of shareholder returns (risk) are control variables in our specifications.

between CEO talent and PPS) from that of our paper (the link between CEO talent and the absence of RPE), one can ask whether the empirical relation between the proxy for CEO talent and the absence of RPE that we document in Table III is incremental to Milbourn's findings. To address that question, we interact our proxies for CEO talent (*Articles* and *IndadjROA*) with both Δ*Shareholder Wealth-Firm* (PPS) and Δ*Shareholder Wealth-Industry* (Market) (RPE), and we introduce these interaction terms simultaneously in an abridged version of equation (2) and estimate the resultant specification. Results reported in Table IV show that the interactions of the CEO talent proxies with either Δ*Shareholder Wealth-Industry* or Δ*Shareholder Wealth-Market* are positive and statistically significant in three out of four cases (columns 1, 3, and 4), consistent with less RPE for talented CEOs. We are also able to replicate

Table IV

Median Regressions of Change in CEO Compensation on Talent Proxies to Control for the Link between CEO Talent and PPS

The table presents median regression results for the final sample of 2,343 CEO-firm-years with data on both CEO citations in the press (*Articles*) and industry-adjusted average ROA over the last 3 years (*IndadjROA*), and consists of officers named as CEOs of all S&P 500 companies during the period 1993 to 2001 in the ExecuComp database after eliminating CEO firm-years with missing data from COMPUSTAT and CRSP. Median regressions minimize the sum of absolute deviations rather than the sum of squared deviations. The dependent variable is the change in total compensation (sum of total flow compensation and change in the market value of the CEO's stock and options). The independent variables include the change in shareholder-wealth firm (beginning-of-the-year market value of the firm multiplied by percent returns for the firm), the change in shareholder-wealth index (beginning-of-the-year market value of the firm multiplied by percent returns earned by other firms in the same industry for columns (1) and (2) or by percent returns earned by the S&P 500 market for columns (3) and (4)), the change in the shareholder-wealth index interacted with *IndadjROA* and the empirical cumulative distribution function (CDF) for *Articles,* and the change in shareholder-wealth interacted with *IndadjROA* and the CDF for *Articles*.

		Dependent Variable: ΔTotal Compensation			
Independent Variables	Predicted Sign	Industry Index (1)	Industry Index (2)	Market Index (3)	Market Index (4)
ΔSW − firm	+	2.65*	2.21*	2.49*	3.31*
ΔSW − index	$\beta_2 < 0$, weak RPE, $\beta_1 + \beta_2 = 0$, strong RPE	−0.38*	−0.05	−1.87*	−6.42*
ΔSW − index ∗ *F*(Articles)	+H1	0.46*		1.01*	
ΔSW − index ∗ IndadjROA	+H1		−0.23#		6.43*
Milbourn's variables					
ΔSW − firm ∗ *F*(Articles)	+	0.50*		0.80*	
ΔSW − firm ∗ IndadjROA	+		2.02*		0.80*
Sample size		3,487	2,343	3,487	2,343
Psuedo-R^2		0.061	0.069	0.061	0.069
Test Strong RPE: $\beta_1 + \beta_2 = 0$ (p-value)		0.000	0.000	0.000	0.000

* and # represent statistical significance at p-value < 0.1 and 0.10, respectively, and is one tailed when the coefficient sign is predicted, two tailed otherwise. Coefficients on intercept, industry, and year dummies are suppressed.

Milbourn's findings that PPS increases with CEO talent, as the coefficients on the interaction of Δ*Shareholder Wealth-Firm* and CEO talent variables are positive and significant in all cases. Thus, results in Table IV assure us that the interaction of CEO talent and RPE is a different empirical phenomenon from the one documented by Milbourn.[14]

B. Skimming Explanation

The empirical results in Table III show that CEOs who are more talented are paid more when the performance of the industry or market index is good. However, corporate governance activists claim that this result may be consistent with rent extraction by CEOs who exploit lax governance mechanisms. Bertrand and Mullainathan (2001) suggest that departures from RPE are pure rents to CEOs. We investigate whether the sensitivity of pay for market or industry-wide performance is indeed a manifestation of such pure rents by including proxies for the quality of corporate governance in the model in equation (3). In particular, we introduce interactions between Δ*Shareholder Wealth-Industry* (or Δ*Shareholder Wealth-Market*) and five governance proxies: (i) a measure of shareholder rights, *g score,* compiled by Gompers, Ishii, and Metrick (2003), where a larger (smaller) *g score* indicates less (more) bargaining power for the shareholder vis-à-vis the manager; (ii) a dummy variable, *CEO-Chair*, that is set to one if the chairman of the board is the CEO and zero otherwise, as per ExecuComp; (iii) the proportion of the executive team that serves on the board, *Onboard,* as per ExecuComp; (iv) the proportion of the executive team subject to an interlocked relation (*Interlock*), as per ExecuComp;[15] and (v) *Meetings*, the number of board meetings held during the year, as per ExecuComp. Greater values of the first four proxies and a smaller number of *Meetings* are assumed to indicate poor governance. If the skimming explanation accounts for our results, we would expect the interaction of Δ*Shareholder Wealth-Industry* (or Δ*Shareholder Wealth-Market*) and the governance proxies to indicate less RPE in the presence of poor governance. More important, if the skimming explanation were to dominate the CEO talent explanation, the coefficients on the

[14] Note that we do not present estimated pay sensitivities for minimum and maximum levels of the CEO talent variables here because it would be hard to isolate the effect of CEO talent on RPE as this effect is commingled with the effect of CEO talent on PPS in the specification shown in Table IV.

[15] ExecuComp codes a dummy variable *"EXEC_DIR"* as *"TRUE"* if an executive officer served as director during the indicated fiscal year. In the paper, the *Onboard* variable captures the proportion of officers who are on the board of directors (calculated as the average of cases in which the dummy variable is set to "TRUE" across officers of a firm for each firm-year). Furthermore, an officer is said to have an interlocked relation if that officer (i) serves on the compensation committee or (ii) serves on the board (or compensation committee) of another company that has an executive officer serving on the board (or compensation committee) of his company. ExecuComp captures this information for each officer of the firm by setting an indicator variable, "PINTRLOC," to "TRUE." In the paper, the *Interlock* variable captures the proportion of officers who have an interlocked relationship (calculated as the average of cases in which "PINTRLOC" is set to "TRUE" across officers of a firm for each firm-year).

interaction of Δ*Shareholder Wealth-Industry* (or Δ*Shareholder Wealth-Market*) and the CEO talent proxies should be zero.

Panel D of Table II reports descriptive data on the governance proxies used in this paper. To conserve sample size, we set missing *g scores* to zero. We obtain a *g score* for 90% of the firm-years, indicated by the dummy variable D_{gscore}, that is set equal to one (zero) when we have (do not have) a *g score*. The CEO is also the chair of the board 86% of the time. Approximately 34% of the executive team serves on the board, while 1% of the executive team is subject to an interlocked relation. The board meets seven times a year in the median firm. In untabulated analyses, we find that our talent proxies are not highly correlated with the governance variables. In fact, none of the correlations between either *Articles* or *IndadjROA* with the corporate governance variables exceeds |0.18|, suggesting that the CEO talent and RPE results are not likely due to governance issues.

Results of regressions after introducing the governance terms are reported in Table V. These results suggest mixed evidence in support of the skimming hypothesis. We observe less RPE when the CEO is also the chair of the board and when the board holds fewer meetings in a year in both the specifications reported in the table, consistent with skimming. The results are somewhat mixed across the two columns when the interaction of the index and *Onboard* or *Interlock* are considered. Most important, however, the CEO talent explanation for the lack of RPE for talented managers survives the skimming explanation. In particular, the coefficients on the interaction of Δ*Shareholder Wealth-Industry* (or Δ*Shareholder Wealth-Market*) and the CEO talent proxies are positive and significant in both specifications. To point to one case, the sensitivity of CEO pay to the market index for CEOs with minimum and maximum *Articles* and *IndadjROA* ranges from \$5.59 to \$12.04, as per column (2).

C. Oligopolistic Industries

Aggarwal and Samwick (1999b) suggest that the absence of RPE is driven by strategic interaction among firms in the product markets. In particular, as in RPE, a negative weight on the rival firm's performance in compensating managers might create incentives for managers to compete too aggressively in the product market and thus hurt the firm's profitability. Thus, in highly concentrated industries, lower emphasis on RPE may be optimal to avoid excessive competition among rival managers and hence to maximize joint (industry-wide) returns. To assess whether that explanation could affect our results, we compute the *Herfindahl-Hirschman Index* and interact this index with market-wide or industry-wide performance. The *Herfindahl–Hirschman Index* (*HHI*) is equal to the sum of an industry's (two-digit SIC) squared market shares (in percentages). Following Defond and Park (1999), the *HHI* value assigned to each sample year equals its industry average *HHI* over the 5 years prior to the event year. High (low) levels of *HHI* imply high (low) industry concentration and high competition. We interact *HHI* with Δ*Shareholder Wealth-Industry* (or Δ*Shareholder Wealth-Market*) and introduce the resultant interaction term in

Table V

Median Regressions of Change in CEO Compensation on Talent Proxies: Skimming Explanation

The table presents median regression results for the final sample of 2,343 CEO-firm years with data on both CEO citations in the press (*Articles*) and industry-adjusted average ROA over the last 3 years (*IndadjROA*), and consists of officers named as CEOs of all S&P 500 companies during the period 1993 to 2001 in the ExecuComp database after eliminating CEO firm-years with missing data from COMPUSTAT and CRSP. Median regressions minimize the sum of absolute deviations rather than the sum of squared deviations. The dependent variable is the change in total compensation (sum of total flow compensation and change in the market value of the CEO's stock and options). The independent variables include the change in shareholder-wealth firm (beginning-of-the-year market value of the firm multiplied by percent returns for the firm), the change in shareholder-wealth index (beginning-of-the-year market value of the firm multiplied by percent returns earned by other firms in the same industry for column (1) or by percent returns earned by the S&P 500 market for column (2)), the change in shareholder-wealth-index interacted with *IndadjROA* and the empirical cumulative distribution function (CDF) for *Articles* and CEO tenure, the change in shareholder wealth-firm interacted with the CDF for CEO age, the market value of equity (*Size*) and the beginning-of-the-year market value times the variance of monthly stock returns, expressed in decimal form over the 60 months preceding the sample year (*Variance of SW*), and the change in shareholder-wealth index interacted with each of the following skimming variables: *g score*, a measure of shareholder rights compiled by Gompers et al. (2003), a dummy variable coded one if the CEO is the chairman of the board (*CEO-Chair*), the proportion of the executive team serving on the board (*Onboard*), the proportion of the executive team subject to interlocked relation (*Interlock*), and the number of board meetings held during the year (*Meetings*). Column (1) shows that the estimated pay-for-performance (PPS) for a CEO in our sample with a minimum/median/maximum F(*Articles*) and *IndadjROA* is \$8.19/\$9.18/\$9.89 for a \$1,000 increase in the market value of the firm. The PPS is computed by evaluating the regression coefficients at the median values of the CDF of the control and skimming variables.

	Dependent Variable: ΔTotal Compensation		
Independent Variables	Predicted Sign	Industry Index (1)	Market Index (2)
ΔSW − firm	+	15.90*	14.50*
ΔSW − index	$\beta_2 < 0$, weak RPE, $\beta_1 + \beta_2 = 0$, Strong RPE	−0.75*	−3.06*
ΔSW − index ∗ *F*(Articles)	+H1	0.85*	1.61*
ΔSW − index ∗ IndadjROA	+H1	0.95*	5.28*
Skimming explanation	If skimming:		
ΔSW − index ∗ g score	+	−0.02*	−0.08*
ΔSW − index ∗ CEO-Chair	+	1.10*	3.14*
ΔSW − index ∗ Onboard	+	0.36*	−4.70*
ΔSW − index ∗ Interlock	+	2.64*	2.48
ΔSW − index ∗ Meetings	−	−0.32*	−0.58*
Controls			
ΔSW − index ∗ *F*(Tenure)	?	2.69*	5.09*
ΔSW − firm ∗ *F*(CEO age)	?	−3.17*	−0.83*
ΔSW − firm ∗ *F*(Size)	?	−27.24*	−28.12*
ΔSW − firm ∗ *F*(Variance of SW)	?	16.10*	17.31*
Sample size		2,343	2,343
Psuedo-R^2		0.096	0.093
Test strong RPE: $\beta_1 + \beta_2 = 0$ (*p*-value)		0.000	0.000
Estimated Pay Sensitivities Assuming Median Control and Skimming Explanation Variables			
Minimum *F*(Articles) and IndadjROA		\$8.19	\$5.59
Median *F*(Articles) and IndadjROA		\$9.18	\$9.65
Maximum *F*(Articles) and IndadjROA		\$9.89	\$12.04

* represents statistical significance at *p*-value < 0.1, and is one tailed when the coefficient sign is predicted, two tailed otherwise. Coefficients on intercept, industry, and year dummies are suppressed. Since the cumulative distribution functions (CDFs) of variables are computed on a yearly basis, estimated pay sensitivities are evaluated at median values of approximately 0.50 for the CDFs of the variables tenure, CEO age, size, and variance of shareholder wealth (SW). The median values used for D_{gscore}* g score, CEO-Chair, Onboard, Interlock, and Meetings are 9, 1, 0.30, 0, and 7.0, respectively.

equation (3).[16] This interaction term would assume a positive coefficient if it were optimal to soften RPE incentives in concentrated industries. Moreover, if our results are driven by strategic interplay in product markets, then the associations between pay for market-wide or industry-wide performance and CEO talent should disappear. A priori, we do not expect our Table III results to be substantially affected by the introduction of the *HHI* interaction because the (untabulated) correlation between *HHI* and the CEO talent proxies is not statistically significant. None of the results reported in Table VI are consistent with the hypothesis of less RPE in oligopolistic industries. As expected, the inclusion of these *HHI*-based interactions has virtually no effect on the CEO talent interactions.

D. Firm-Level Hedging

Garvey and Milbourn (2003) argue that compensation committees do not filter out market-wide changes from executive compensation (i.e., the absence of RPE) because executives can filter out such changes on their own by accessing the capital market. In particular, they argue that younger executives and those with less firm-specific wealth (a proxy for the CEO's unobservable total wealth) cannot undo the filtering of market effects easily. Hence, the sensitivity of their pay to market-wide factors would be smaller—that is, these executives will face more RPE than older executives and those with more firm-specific wealth. To assess the robustness of our results to this explanation, we interact Δ*Shareholder Wealth-Industry* (or Δ*Shareholder Wealth-Market*) with two variables, *CEO age* and the *level* of firm-specific wealth held by the CEO at the beginning of the year. We introduce these interaction terms into equation (3) and investigate whether the coefficients on Δ*Shareholder Wealth-Industry* and CEO talent proxies are still positive and significant. Results reported in Table VII indicate mixed support for the hedging explanation, as the interaction of industry and market indexes and the level of CEO wealth is positive and significant while the interaction of the indexes with CEO age is negative and significant throughout. These results suggest that CEOs with more firm-specific wealth face less RPE, consistent with the hedging explanation. However, inconsistent with the Garvey and Milbourn (2003) hypothesis, older CEOs in our sample face more RPE.

More importantly, the interaction of the industry and market indexes with the *Articles*-based proxy for CEO talent continues to be positive and significant for the industry index but is positive and insignificant for the market index. In untabulated results, we find that the interaction of *Articles* and *market index* is positive and statistically significant (coefficient = 0.68, p-value < 0.01) when the larger sample of 3,487 observations with usable *Articles* observations is used. However, the interaction of the *market index* and *IndadjROA* is negative in column (1) and insignificant in column (2). Further investigation

[16] The oligopolistic industry explanation is a natural fit with the industry index but not with the market index. We present results related to market index here for completeness.

Table VI

Median Regressions of Change in CEO Compensation on Talent Proxies: Oligopoly Explanation

The table presents median regression results for the final sample of 2,343 CEO-firm years with data on both CEO citations in the press (*Articles*) and industry-adjusted average ROA over the last 3 years (*IndadjROA*), and consists of officers named as CEOs of all S&P 500 companies during the period 1993 to 2001 in the ExecuComp database after eliminating CEO firm-years with missing data from COMPUSTAT and CRSP. Median regressions minimize the sum of absolute deviations rather than the sum of squared deviations. The dependent variable is the change in total compensation (sum of total flow compensation and change in the market value of the CEO's stock and options). The independent variables include the change in shareholder-wealth firm (beginning-of-the-year market value of the firm multiplied by percent returns for the firm), the change in shareholder-wealth index (beginning-of-the-year market value of the firm multiplied by percent returns earned by other firms in the same industry for column (1) or by percent returns earned by the S&P 500 market for column (2)), the change in shareholder-wealth index interacted with *IndadjROA* and the empirical cumulative distribution function (CDF) for *Articles* and CEO tenure, the change in shareholder-wealth firm interacted with the CDF of CEO age, market value of equity (*Size*), and beginning-of-the-year market value times the variance of monthly stock returns, expressed in decimal form over the 60 months preceding the data year (*Variance of SW*) and the change in shareholder-wealth index interacted with the following oligopoly variable: sum of the industry's squared market share in percentages as in Defond and Park (1999) (*Herfindahl–Hierschman Index*). Column (1) shows that the estimated pay-for-performance (PPS) for a CEO in our sample with a minimum/median/maximum *F(Articles)* and *IndadjROA* is \$8.03/\$8.65/\$9.19 for a \$1,000 increase in the market value of the firm. The PPS is computed by evaluating the regression coefficients at the median values of the CDF of the control and oligopoly variables.

		Dependent Variable: ΔTotal Compensation	
Independent Variables	Predicted Sign	Industry Index (1)	Market Index (2)
ΔSW − firm	+	15.73*	14.13*
ΔSW − index	$\beta_2 < 0$, weak RPE, $\beta_1 + \beta_2 = 0$, strong RPE	−2.39*	−6.37*
ΔSW − index ∗ *F*(Articles)	+H1	0.87†	1.41*
ΔSW − index ∗ IndadjROA	+H1	0.35†	3.99*
Oligopoly explanation			
ΔSW−index ∗ Herfindahl–Hierschman Index	+ if less RPE in concentrated industry	−0.28*	−1.63*
Controls			
ΔSW − index ∗ *F*(Tenure)	?	3.48*	4.83*
ΔSW − firm ∗ *F*(CEO age)	?	−2.60*	−0.84*
ΔSW − firm ∗ *F*(Size)	?	−27.64*	−28.21*
ΔSW − firm ∗ *F*(Variance of SW)	?	16.16*	17.92*
Sample size		2,343	2,343
Psuedo-R^2		0.094	0.092
Test Strong RPE: $\beta_1 + \beta_2 = 0$ (*p*-value)		0.000	0.000
Estimated Pay Sensitivities Assuming Median Control and Oligopoly Explanation Variables			
Minimum *F*(Articles) and IndadjROA		\$8.03	\$4.69
Median *F*(Articles) and IndadjROA		\$8.65	\$7.86
Maximum *F*(Articles) and IndadjROA		\$9.19	\$9.76

* and † represent statistical significance at *p*-value < 0.1 and 0.05, respectively, and is one tailed when the coefficient sign is predicted, two tailed otherwise. Coefficients on intercept, industry, and year dummies are suppressed. Since the cumulative distribution functions (CDFs) of variables are computed on a yearly basis, estimated pay sensitivities are evaluated at median values of approximately 0.50 for the CDFs of the variables tenure, CEO age, size, and variance of shareholder wealth (SW). The median value used for the Herfindahl–Hierschman-Index is 0.15.

Table VII

Median Regressions of Change in CEO Compensation on Talent Proxies: Hedging Explanation

The table presents median regression results for the final sample of 2,343 CEO-firm years with data on both CEO citations in the press (*Articles*) and industry-adjusted average ROA over the last 3 years (*IndadjROA*), and consists of officers named as CEOs of all S&P 500 companies during the period 1993 to 2001 in the ExecuComp database after eliminating CEO firm-years with missing data from COMPUSTAT and CRSP. Median regressions minimize the sum of absolute deviations rather than the sum of squared deviations. The dependent variable is the change in total compensation (sum of total flow compensation and change in the market value of the CEO's stock and options). The independent variables include the change in shareholder-wealth firm (beginning-of-the-year market value of the firm multiplied by percent returns for the firm), the change in shareholder-wealth index (beginning-of-the-year market value of the firm multiplied by percent returns earned by other firms in the same industry for column (1) or by percent returns earned by the S&P 500 market for column (2)), the change in shareholder-wealth-index interacted with *IndadjROA* and the empirical cumulative distribution function (CDF) for *Articles* and CEO tenure, the change in shareholder wealth-firm interacted with the CDF for CEO age, market value of equity (*Size*) and beginning-of-the-year market value times the variance of monthly stock returns, expressed in decimal form over the 60 months preceding the data year (*Variance of SW*), and the change in shareholder-wealth index interacted with each of the following hedging variables: CEO age and CEO firm specific wealth at the start of the year. Column (1) shows that the estimated pay-for-performance (PPS) for a CEO in our sample with a minimum/median/maximum *F(Articles)* and *IndadjROA* is \$9.17/\$8.31/\$8.01 for a \$1,000 increase in the market value of the firm. The PPS is computed by evaluating the regression coefficients at the median values of the CDF of the control and hedging variables.

		Dependent Variable: ΔTotal Compensation	
Independent Variables	Predicted Sign	Industry Index (1)	Market Index (2)
ΔSW − firm	+	18.19*	14.80*
ΔSW − index	$\beta_2 < 0$, weak RPE, $\beta_1+\beta_2 = 0$, strong RPE	−4.79*	−7.98*
ΔSW − index ∗ *F*(Articles)	+H1	0.39*	0.13
ΔSW − index ∗ IndadjROA	+H1	−1.65*	0.37
Hedging explanation	If hedging:		
ΔSW − index ∗ *F*(CEO age)	+	−0.87*	−5.80*
ΔSW − index ∗ *F*(CEO firm specific wealth at the start of the year)	+	7.93*	11.16*
Controls			
ΔSW − index ∗ *F*(Tenure)	?	1.96*	5.62*
ΔSW − firm ∗ *F*(CEO age)	?	−2.90*	−0.39*
ΔSW − firm ∗ *F*(Size)	?	−26.26*	−28.28*
ΔSW − firm ∗ *F*(Variance of SW)	?	11.95*	16.79*
Sample size		2,343	2,343
Psuedo-R^2		0.098	0.094
Test Strong RPE: $\beta_1 + \beta_2 = 0$ (*p*-value)		0.000	0.000
Estimated Pay Sensitivities Assuming Median Control and Hedging Explanation Variables			
Minimum *F*(Articles) and IndadjROA		\$9.17	\$6.34
Median *F*(Articles) and IndadjROA		\$8.31	\$6.63
Maximum *F*(Articles) and IndadjROA		\$8.01	\$6.81

* represents statistical significance at *p*-value < 0.1, and is one tailed when the coefficient sign is predicted, two tailed otherwise. Coefficients on intercept, industry, and year dummies are suppressed. Since the cumulative distribution functions (CDFs) of variables are computed on a yearly basis, estimated pay sensitivities are evaluated at median values of approximately 0.50 for the CDFs of the variables tenure, CEO age, firm specific wealth at the start of the year, size, and variance of shareholder wealth (SW).

reveals that the lowest correlation between the interaction of any index and *F*(firm-specific wealth at the start of the year) and the interaction of that index and *IndadjROA* is 0.94 (p-value < 0.01). Hence, disentangling the CEO talent explanation from the hedging hypothesis is difficult when *IndadjROA* is the talent proxy. However, when *Articles* is the talent proxy, the CEO talent explanation still generally stands.

E. Asymmetric Benchmarking

Bertrand and Mullainathan (2001) and Garvey and Milbourn (2004) note that CEO pay is benchmarked to the market index only when the market seems to be up. In other words, CEO pay is shielded from bad luck (or negative exogenous shocks) but is sensitive to good luck (positive exogenous shocks). The authors characterize these findings as evidence of rent extraction by managers. It is ambiguous whether the retention explanation proposed by Oyer (2004) and Himmelberg and Hubbard (2000) is symmetric with respect to up and down markets—is a talented CEO's pay as exposed to the down market as it is in the up market?

In contrast to Bertrand and Mullainathan (2001) and Garvey and Milbourn's (2004) thesis that asymmetric benchmarking indicates rent extraction by managers, one can posit at least three conceptual arguments for why shielding talented CEOs in a down market downturn might be optimal from the firm's perspective. First, more talented CEOs would continue to have greater outside opportunities in a bad market relative to less talented CEOs. Moreover, they can pursue opportunities in the non-profit sector or they may decide to take time off during a downturn to write a memoir or pursue other personal interests. In fact, retaining talented executives in a market downturn might be even more important for the firm than retaining them in an up market. Second, firms often weed out less talented managers in a bad market, when it is relatively easy to hire away talented managers from rival firms. Hence, firms with talented managers have to be even more careful to not lose such managers in a bad market. Third, shielding pay of talented managers in a down market has the same economic flavor as repricing stock options in a down market. Several researchers argue that option repricing might be optimal (e.g., Saly (1994), Acharya et al. (2000), Carter and Lynch (2001, 2004), Chidambaram and Prabhala (2003)).

To explore whether asymmetric benchmarking occurs, especially for talented executives, we separate the sample into an up (down) sample depending on whether the industry index or the S&P 500 market index records negative (non-negative) buy-and-hold annual returns. For the up and down subsamples, we estimate the basic equation (3) and present the results related to the up (down) markets in Panels A (B) of Table VIII. Results reported in Panel A are very similar to those reported in Table III in that talented managers are subject to less relative performance evaluation in an up market. Turning to the down market results in Panel B, we find that the coefficient on Δ*Shareholder Wealth-Industry* is negative and significant while the one on Δ*Shareholder Wealth-Market* is negative but insignificant, suggesting that firms generally

Table VIII

Median Regressions of Change in CEO Compensation on Talent Proxies: Asymmetric Benchmarking Explanation

The table presents median regression results for the final sample of CEO-firm years with data on both CEO citations in the press (*Articles*) and industry-adjusted average ROA over the last 3 years (*IndadjROA*), and consists of officers named as CEOs of all S&P 500 companies during the period 1993 to 2001 in the ExecuComp database after eliminating CEO firm-years with missing data from COMPUSTAT and CRSP. Median regressions minimize the sum of absolute deviations rather than the sum of squared deviations. The dependent variable is the change in total compensation (sum of total flow compensation and change in the market value of the CEO's stock and options). The independent variables include the change in shareholder-wealth firm (beginning-of-the-year market value of the firm multiplied by percent returns for the firm), the change in shareholder-wealth index (beginning-of-the-year market value of the firm multiplied by percent returns earned by other firms in the same industry for column (1) or by percent returns earned by the S&P 500 market for column (2)), the change in shareholder-wealth index interacted with *IndadjROA* and the empirical cumulative distribution function (CDF) for *Articles* and CEO tenure, and the change in shareholder-wealth firm interacted with the CDF for CEO age, market value of equity (*Size*) and beginning-of-the-year market value times the variance of monthly stock returns, expressed in decimal form over the 60 months preceding the sample year (*Variance of SW*). Column (1) shows that the estimated pay-for-performance (PPS) for a CEO in our sample with a minimum/median/maximum F(*Articles*) and *IndadjROA* is \$7.28/\$8.78/\$9.85 for a \$1,000 dollar increase in the market value of the firm. The PPS is computed by evaluating the regression coefficients at the median values of the CDF of the control variables. Panel A presents the median regression results for when the industry or market index was up using 1,747 and 1,629 CEO-firm years, respectively, and Panel B presents the median regression results for when the industry or market index was down using 596 and 714 CEO-firm years, respectively.

	Dependent Variable: ΔTotal Compensation		
Independent Variables	Predicted Sign	Industry Index (1)	Market Index (2)
Panel A: Asymmetric Benchmarking Explanation When the Index Is Up			
ΔSW − firm	+	16.11*	15.43*
ΔSW − index	$\beta_2 < 0$, weak RPE, $\beta_1 + \beta_2 = 0$, strong RPE	−3.09*	−10.70*
ΔSW − index ∗ *F*(Articles)	+H1	1.30*	4.34*
ΔSW − index ∗ IndadjROA	+H1	1.42*	6.88*
Controls			
ΔSW − index ∗ *F*(Tenure)	?	3.31*	10.79*
ΔSW − firm ∗ *F*(CEO age)	?	−3.68*	−1.91*
ΔSW − firm ∗ *F*(Size)	?	−26.30*	−29.44*
ΔSW − firm ∗ *F*(Variance of SW)	?	14.97*	17.90*
Sample size		1,747	1,629
Psuedo-R^2		0.096	0.082
Test Strong RPE: $\beta_1 + \beta_2 = 0$ (p-value)		0.000	0.000
Estimated Pay Sensitivities Assuming Median Control Variables			
Minimum *F*(Articles) and IndadjROA		\$7.28	\$4.04
Median *F*(Articles) and IndadjROA		\$8.78	\$10.37
Maximum *F*(Articles) and IndadjROA		\$9.85	\$14.60

(*continued*)

Table VIII—*Continued*

Independent Variables	Predicted Sign	Dependent Variable: ΔTotal Compensation Industry Index (1)	 Market Index (2)
Panel B: Asymmetric Benchmarking Explanation When the Index Is Down			
ΔSW − firm	+	13.30*	13.87*
ΔSW − index	$\beta_2 < 0$, weak RPE, $\beta_1 + \beta_2 = 0$, strong RPE	−2.74*	−0.70
ΔSW − index ∗ *F*(Articles)	+H1	1.10*	−2.32*
ΔSW − index ∗ IndadjROA	+H1	0.83*	−1.07
Controls			
ΔSW − index ∗ *F*(Tenure)	?	2.93*	2.36*
ΔSW − firm ∗ *F*(CEO age)	?	0.77*	−0.41*
ΔSW − firm ∗ *F*(Size)	?	−27.94*	−30.60*
ΔSW − firm ∗ *F*(Variance of SW)	?	17.33*	20.28*
Sample size		596	714
Psuedo-R^2		0.103	0.141
Test Strong RPE: $\beta_1 + \beta_2 = 0$ (p-value)		0.000	0.000
Estimated Pay Sensitivities Assuming Median Control Variables			
Minimum *F*(Articles) and IndadjROA		$7.21	$8.81
Median *F*(Articles) and IndadjROA		$8.24	$7.07
Maximum *F*(Articles) and IndadjROA		$9.04	$7.91

* represents statistical significance at $p < 0.1$, and is one tailed when the coefficient sign is predicted, two tailed otherwise. Coefficients on intercept, industry, and year dummies are suppressed. Since the cumulative distribution functions (CDFs) of variables are computed on a yearly basis, estimated pay sensitivities are evaluated at median values of approximately 0.50 for the CDFs of the variables tenure, CEO age, size, and variance of shareholder wealth (SW). Industry Index (1) is evaluated at the 0.49 quantile.

shield CEO pay from an industry but not a market downturn. However, the sign on the interaction of Δ*Shareholder Wealth-Industry* (or Δ*Shareholder Wealth-Market*) and *F*(*Articles*) is mixed. Note that in column (2), the sign on the interaction term Δ*Shareholder Wealth-Market* ∗ *F(Articles)* is negative, suggesting that talented executives are shielded more in a downturn while in column (1), the sign on the interaction term Δ*Shareholder Wealth-Industry* ∗ *F(Articles)* is positive, suggesting that talented executives are shielded less in a downturn. Furthermore, the coefficient on Δ*Shareholder Wealth-Industry* (or Δ*Shareholder Wealth-Market*) and *IndadjROA* is also mixed—positive in column (1) and insignificant in column (2). Hence, the results reveal no systematic pattern of asymmetric benchmarking for either talented or less talented executives, and thus do not suggest that rent extraction is the dominant explanation for asymmetric benchmarking.

IV. Conclusions

A widespread concern among both practitioners and academics is that executive pay lacks relative performance evaluation (RPE). In recent years,

significant increases in CEO compensation concurrent with the bull markets has prompted several critics to argue that CEO compensation for market-wide shocks is windfall pay consistent with systematic governance failures. A few economists (Oyer (2004) and Himmelberg and Hubbard (2000)) propose that what looks like windfall pay might reflect the special conditions in the labor markets for CEOs. In particular, if CEO talent is scarce, then the supply of talented CEOs is relatively inelastic, in which case it may be optimal to reward CEOs for market-wide shocks if such shocks raise the firm's market value and the CEO's outside employment opportunities. Empirical testing of this hypothesis has been hampered by the absence of readily available proxies for CEO talent.

We contribute in this regard by constructing proxies for CEO talent based on a CEO's visibility in the financial press and the CEO firm's past industry-adjusted ROA. Relying on emerging research (Milbourn (2003) and Francis et al. (2004)) and our own work, we validate the press visibility proxy in several ways. Our empirical results based on these proxies show that the absence of RPE is systematically related to CEO talent. These results are robust to several checks against alternate hypotheses such as skimming and strategic consideration in oligopolistic industries. We observe asymmetric benchmarking (less RPE when the market is up but more RPE when the market is down), but there is no systematic pattern of asymmetric benchmarking for talented executives and thus our results do not support rent extraction as the dominant explanation. Taken together, we believe our results are consistent with the CEO talent explanation for the relative absence of RPE in executive compensation contracts.

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