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Implicit contracts and the explanatory power of top executive compensation for future performance

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Recent research suggests that implicit incentive contracts may be based on performance measures that are observable only to the contracting parties. We derive and test implications of this insight for the relationship between executive compensation and firm performance. If corporate boards optimally use both observable and unobservable (to outsiders) measures of executive performance and *the unobservable measures are correlated with future firm performance, then unexplained variation in* current *compensation should predict* future *variation in firm performance. Further, compensation should be more positively associated with future performance when observable measures are less useful for contracting. Our results are consistent with these hypotheses.*

1. Introduction

■ Theoretical research on implicit or relational contracts suggests that incentive contracts may be based on performance measures that are observable only to the parties of the contract. In this article, we ask whether boards of directors use such measures to reward top executives for actions that benefit the firm, but that are not reflected in the firm's current publicly observable performance. We argue that if compensation contracts optimally incorporate both observable and unobservable (to outsiders) measures of performance *and* the unobservable measures of performance are correlated with future observable measures of performance, then variation in *current* compensation that is not explained by variation in current observable performance measures should predict *future* variation in observable performance measures. Our first hypothesis is therefore that compensation is informative about future performance.

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This reasoning also suggests that compensation should be a better predictor of future performance when unobservable measures of performance receive greater weight in the optimal incentive contract. Intuitively, when the weight placed on the unobservable performance measure is higher, there is more information about future performance contained in wages. We use this idea to test propositions generated by multiple-performance-measure agency models, such as Banker and Datar (1989), Holmström and Milgrom (1991), and Baker, Gibbons, and Murphy (1994). These analyses predict that the weight given to a performance measure in an efficient contract is a function of the properties of the other available measures. In general, the weight on one measure should be decreasing in the precision and sensitivity of other measures. This, in turn, generates our second hypothesis: when observable performance measures are less useful as contracting instruments, the strength of the relationship between current compensation and future performance should increase.

We test these hypotheses using data on firm performance and chief executive officer (CEO) compensation. Our main regression equation uses current firm performance variables and current log CEO compensation to try to predict future return on shareholders' equity. Applying several different specifications, we find strong evidence in favor of our first hypothesis; that is, we find CEO compensation is informative about future return on equity.

To explore our second hypothesis, we examine whether the informativeness of compensation for future return on equity is inversely related to the quality of observable performance measures as contracting instruments. We interact our compensation variable with various proxies for the quality of observable performance measures. Our methodology here parallels that of Bushman, Indjejikian, and Smith (1996) and Ittner, Larcker, and Rajan (1997), who use survey responses and compensation committee reports, respectively, to show that firms substitute away from firmwide financial performance measures when these measures are less indicative of managerial performance. Our primary finding is that unexplained variation in compensation is more positively related to future return on equity when the variances of market and accounting returns are higher. We then take additional proxies for the sensitivity of current observable performance measures to current managerial actions from the Bushman, Indjejikian, and Smith (1996) and Ittner, Larcker, and Rajan (1997) analyses, including the firm's market-to-book ratio, ratio of employees to sales, ratio of research and development expense to sales, and lengths of product development and product life cycles. We find some evidence that employee-to-sales ratios and lengths of product development cycles are positively related to the use of privately observed measures of performance, and we obtain mixed evidence for the other measures.

We view our results as support for the assertion that boards of directors use nonobservable (to outsiders) information to reward top executives for actions that benefit the firm but that are not reflected in currently observable performance. This finding bears directly on the current debate as to whether boards provide useful oversight of top managers. Our results are also notable in light of recent work by Hall and Liebman (1998), who report that the vast majority of variation in executive wealth associated with changes in firm value stems from executives' holdings of stock and stock options. Given this fact, one may wonder whether salary and bonus payments are superfluous as incentive instruments; that is, have boards of directors completely delegated the tasks of monitoring and rewarding top managers to capital markets as a by-product of the extensive use of equity-based pay instruments? Our analysis suggests that the answer to this question is no—boards do appear to perform a vital governance role by collecting and evaluating performance-related information that is not available to those outside the firm.

On a much broader level, our results are supportive of the hypothesis that firms use implicit or relational employment contracts to motivate and reward employees. Since information that is unobservable to outsiders is necessarily nonverifiable, enforcement of agreements based on nonobservable information cannot be accomplished through explicit, legally enforceable contracts. The theory of implicit contracts, however, suggests an alternative enforcement mechanism: if firms value reputations in the labor market for treating employees fairly, then the potential loss of this reputation can prevent firms from reneging on such agreements.[1] While the idea of implicit employment contracts underlies much recent study of incentives in organizations (see surveys by Gibbons (1998) and Prendergast (1999)), there is, to date, little empirical research documenting behavior consistent with this hypothesis.

2. Hypotheses

■ We outline our hypotheses by discussing a simple two-period agency framework. A firm hires an agent (the CEO) to take an action. The action, which is taken in the first period, generates three noisy signals, each of which is positively related to the agent's effort. The signals differ according to when and by whom they are observed. In particular, suppose x is a publicly observable signal and that x is contemporaneous with effort. (That is, x is observed in the first period when the agent takes action.) Let y be publicly observable, but suppose it is not revealed to anyone until the second period. We let z be contemporaneous with the action, but we assume z is observed by only the firm and the CEO. We assume that the partial correlation coefficients among the three signals are positive.

While those outside the firm do not observe the performance measure z, the firm and the CEO may still be able to contract on it.[2] We suppose that the payment made by the firm to the CEO in the first period depends on three factors: the first-period observable performance measure (x), the first-period unobservable (to outsiders) performance measure (z), and other individual- or firm-specific factors that are orthogonal to current and future performance (which we denote by the random variable q). Assuming the contract is a linear function of these factors, we can write

$$w = \gamma_0 + \gamma_x x + \gamma_z z + \gamma_q q.$$

Consider, then, a linear regression of wages (w) on just current observable performance (x). Letting $\hat{w} = E[w \mid x]$, we write the residual from this regression as

$$w - \hat{w}.$$

If this residual is positive, then it must be that either

$$q > E[q] \quad \text{or} \quad z > E[z \mid x].$$

That is, if compensation is unexpectedly high compared to what would be expected

[1] The theory of implicit contracts has been developed by Azariadis (1975), among others. Holmström (1983) and Bull (1987) offered the first models of reputation as an enforcement mechanism for implicit contracts.

[2] Here, we rely on an unmodelled reputational mechanism to enforce the firm's commitment to condition its payment to the agent on the privately observed signal z. Baker, Gibbons, and Murphy (1994) and Levin (1998) explicitly consider the use of reputation to enforce implicit incentive contracts based on nonverifiable information.

given current performance, then either the unobserved signal of performance is higher than its expectation given x, or the non-performance-related factors that affect compensation are higher than expected.

Now consider the residual in a linear regression of second-period performance (y) on first-period performance (x). We can write this residual as

$$y - \hat{y},$$

where we let $\hat{y} = E[y|x]$. Since q is assumed to be orthogonal to performance, we have that

$$E[y - \hat{y}|q > E[q]] = 0.$$

However, as long as x is not sufficient for z with respect to y, we have that

$$E[y - \hat{y}|z > E[z|x]] > 0.$$

Together, these statements imply that

$$E[y - \hat{y}|w > \hat{w}] > 0.$$

In words, when compensation is unexpectedly high, this must be due either to random, individual or firm-specific factors *or* to information suggesting the executive's effort was higher than indicated by current observable performance measures alone. If it is the latter, then we should find future performance to be better than indicated by current performance alone. In a sufficiently large sample, we should find that variation in executive compensation that is not explained by current performance should be positively associated with variation in future performance.

How might one attempt to test this prediction? A sensible procedure might be to

(i) regress future performance on current performance,
(ii) regress wages on current performance, and
(iii) regress the residuals from (i) on the residuals from (ii).

This procedure would tell us how wages are related to future performance after netting out the effects of current performance, and it is equivalent to regressing future performance on current performance and wages. We therefore consider the following regression equation:

$$y = \beta_0 + \beta_1 x + \beta_2 w + \epsilon. \tag{1}$$

We expect β_2 to be positive, indicating that even controlling for current performance, wages are positively associated with future performance.

To develop our second hypothesis, we observe that when γ_z is larger, the relationship between unexplained variation in current compensation and future performance should be stronger. To see this, note that the compensation residual, $w - \hat{w}$, can be written as

$$\gamma_z(z - E[z|x]) + \gamma_q(q - E[q]).$$

When γ_z is larger, the term $z - E[z|x]$ receives more weight in the compensation residual. Higher γ_z thus implies a higher signal-to-noise ratio in the compensation residual, which means that $w - \hat{w}$ should be more highly correlated with the future performance residual $y - \hat{y}$.

Using agency theory, we then identify conditions under which we would expect γ_z to be large. In models with multiple *verifiable* measures of performance (such as Banker and Datar (1989) and Holmström and Milgrom (1991)), performance measures are substitutes in an optimal contract: the weight placed on a given measure is a decreasing function of the precision and sensitivity of the other available measures. This result arises because as a performance measure becomes more precise or more sensitive, its use places less risk onto the agent. Since the optimal contract trades off risk and incentives at the margin, performance measures that provide incentives at lowest risk cost receive the highest weights. Hence, if the firm and manager are sufficiently patient so that constraints on reputational enforcement do not affect their ability to contract on the privately observed measure z, then we expect higher choices of γ_z when the publicly observable performance measures are less precise or less sensitive.

Similar results hold, but for somewhat different reasons, when enforcement concerns do limit the parties' ability to contract on z. Baker, Gibbons, and Murphy (1994) develop a model in which reputational enforcement constraints bind and show that, subject to one caveat, verifiable and nonverifiable performance measures remain substitutes in an optimal contract.[3] In their model, a firm that considers reneging on a payment based on a nonverifiable performance measure compares the gains from reneging to the value of its reputation for making such payments. Since reputation permits the firm to contract on the nonverifiable measure, the value of the firm's reputation is simply the difference between the firm's future profits when contracting on both measures and profits when contracting on only the verifiable measure. The presence of higher-quality verifiable measures means this difference is smaller, which implies a greater temptation to renege and thus limits the firm's ability to contract on nonverifiable measures.[4] Applying this result to the setting we consider, we again expect greater weight to be placed on the privately observed (and hence nonverifiable) measure z when the publicly observable (and verifiable) performance measures are less precise or less sensitive.

To devise a test of this comparative static prediction, we interact proxies for the quality of observable measures with the variable w in (1). For example, if the variance of a firm's market returns is high, these returns are less precise as a measure of executive actions. Similarly, if a firm's product development or life cycles are long, then current performance is less sensitive to current executive actions. In these cases, we expect firms to make greater use of the privately observed measure z. Our discussion therefore predicts positive coefficients on interaction variables such as compensation/variance-of-market-returns and compensation/length-of-product-cycles.[5]

[3] While Baker, Gibbons, and Murphy (1994) refer to their verifiable and nonverifiable performance measures as "objective" and "subjective" measures, respectively, their model applies to any setting in which a firm makes use of both verifiable and nonverifiable measures.

[4] As noted above, this result is subject to one caveat. Baker, Gibbons, and Murphy (1994) show that if the firm cannot earn positive profits when contracting on only the verifiable measure, then the verifiable and nonverifiable measures are complements. We argue that this caveat is unlikely to apply in the executive compensation context, especially given the prevalence of stock and stock options, which are explicit contracts based on verifiable measures.

[5] Note that to carry this comparative static prediction over to a cross-section of firms, we need the additional assumption that proxies for the quality of publicly observable performance measures are uncorrelated with characteristics of the privately observed measure. If firms with low-quality observable measures also have low-quality unobservable measures, then our method is biased against our predicted result.

3. Empirical analysis

■ **Data.** Like most empirical analyses of agency models, we face the difficulty of translating the observable measures of performance in our model (x and y in (1) above) into the actual executive compensation context. Since one of our objectives is to assess the extent to which there is information contained in wages that is not in current publicly observable measures, we employ a vector of performance measures as our x variable. We use various measures of year-t performance, such as return on common equity (ROE_t), stock market returns (RET_t), log sales ($\Delta \log SALES_t$), and lagged values of these variables.[6] As our measure of future firm performance, we employ the firm's return on equity in year $t + 1$ (ROE_{t+1}). Data sources for these firm performance measures are CRSP and Compustat.

Our compensation data are taken from the *Forbes* Executive Compensation Surveys from 1974 through 1995. Our measure of compensation (w in (1)) is the natural logarithm of salary plus bonus. We should note that although strong incentives can be provided by an executive's holdings of stock and stock options, we are not interested in these incentives in this study. We are interested in the incentives provided by the board's ability to reward CEOs based on signals that are not publicly observable. Thus, for our purposes it is appropriate to examine only those instruments over which the board has direct control. Hence, salary plus bonus is probably a more appropriate measure than total change in CEO pay-related wealth. Of course, it would be useful to examine *grants* of stock and options to the extent that boards use these instruments rather than cash to reward managers.[7] Per SEC disclosure rules, compensation figures reflect the dollar value of salary and bonus earned during the fiscal year covered. Typically, firms' compensation committees meet early in fiscal year $t + 1$ to determine year-t bonus amounts. Our measure of year-t compensation therefore includes salary and bonus that is earned in year t but may be paid in year $t + 1$.[8] Note that for our purposes it is especially important that salary and bonus payments reported in *Forbes* be matched to the correct fiscal year of firm performance data from Compustat; accordingly, we handchecked data from all firms with non-December fiscal-year ends, referring in many cases directly to the firms' proxy statements.

We adjust market returns, sales, and compensation figures for inflation using the Consumer Price Index.[9] We include a firm-year in our base sample only if the CEO has been in that position for three years or more and if Compustat and CRSP data are available for the current and subsequent year, as well as the two prior years. We remove firm-years if the CEO is in the first year in office because we need to construct a first difference of compensation. We remove firm-years if the CEO is in the second year in office because the *Forbes* surveys often report compensation for a partial year for CEOs

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[6] Many studies of executive compensation have focused on these measures of firm performance. Murphy (1999) surveys this literature.

[7] Omission of these forms of compensation is troublesome for our interpretation only if firms systematically substitute away from (toward) cash compensation toward these other instruments when future performance is expected to be poor (good). Yermack (1997) finds that option grants tend to precede good performance, reporting (p. 449) that "the timing of [option] awards coincides with favorable movements in company stock prices. Patterns of companies' quarterly earnings announcements are consistent with an interpretation that CEOs receive stock option awards shortly before favorable corporate news."

[8] If firms systematically violate SEC rules by reporting compensation *paid* in year t, this would cause us to match year $t - 1$ bonuses to year $t + 1$ performance, and it would bias our analysis against finding support for our hypotheses. It is also possible that firms may use a combination of a year-t bonus and a year $t + 1$ salary increase to reward good performance in year t. Again, this would introduce a bias against our hypotheses.

[9] Because of the difficulty associated with controlling for various asset layers, we do not adjust ROE.

who are in their first year on the job. Hence, the lagged value of compensation for second-year CEOs may not reflect a full year's work. To reduce the effect of extreme values of return on equity (ROE), we drop observations if the absolute value of the firm's change in ROE is greater than .5.[10]

Our base sample contains 8,615 firm-years. We present summary statistics for this sample in Table 1. Since our regressions are run in first differences, we report changes as well as levels.

□ **Unexplained variation in compensation and future performance.** Our first objective is to test the assertion that unexplained variation in executive compensation is related to future performance. As noted above, one way to test this assertion would be the following three-step procedure: First regress future performance on currently available performance measures, then regress current compensation on currently available performance measures. Regress the residual from the first regression on the residual from the second. This would yield the relationship between future performance and current compensation, netting out the effect of currently available performance measures on both variables.

Equivalently, we can execute this procedure in one step by regressing future performance on current compensation and currently available performance measures. Our basic approach is to study an empirical model that is analogous to the first difference of (1) above. We relate changes in future ROE to changes in log compensation, stock returns, log sales, and ROE.[11] To provide as much explanatory power as possible for current compensation and future performance, we include lagged values of the firm-performance variables, as well as industry and year dummy variables. Since it is well documented that accounting returns may reflect market returns with a considerable lag (see Kothari and Sloan, 1992), we include two additional lags of stock returns.[12]

We estimate this model using ordinary least squares and report heteroskedastic-consistent standard errors. We present results in column (1) of Table 2. Future changes in ROE are shown to be negatively related to current changes in ROE and positively related to current market returns and current changes in sales. This accords with prior work on determinants of ROE (see Foster, 1986). Our analysis of the use of unobservable (to outsiders) information in compensation contracts suggests that the coefficient on $\Delta \log \text{COMP}_t$ should be positive. Our estimates indicate that, consistent with this hypothesis, variation in current compensation that is unexplained by current performance is useful in predicting future performance. Our estimate of the coefficient on $\Delta \log \text{COMP}_t$ is significant at better than the 5% level. In column (2), we add a lagged value of $\Delta \log$ COMP. While the coefficient on lagged change in compensation is not statistically different from zero, the coefficient on year-t change in compensation remains significant at the 5% level.[13]

[10] This is approximately the same as dropping the first and ninety-ninth percentiles of change in ROE.

[11] We obtained qualitatively similar results when using levels of stock returns rather than changes in returns as independent variables. Joskow and Rose (1994) compare various specifications of the relationship between log compensation and stock returns, and they report that the data support a model relating changes in log compensation to changes in stock returns.

[12] Boschen and Smith (1995) and Hallock and Oyer (1999) also examine the multiyear relationship between compensation and firm performance. While both studies present results that are consistent with our findings, neither interprets the connection as indicating the use of unobservable (to outsiders) information in measuring executive performance. In addition, neither study links the strength of this relationship to the quality of observable measures of performance as contracting instruments.

[13] An alternative explanation might be that, since compensation committees often determine fiscal year-t bonuses one or two months after the end of the year, the committee may incorporate operating results from the first few months of year $t + 1$. To test for this effect, we reestimated Table 2 using the sum of second, third, and fourth quarter earnings as the dependent variable. The results are similar to those presented.

TABLE 1 Summary Statistics

Variable	Number of Observations	Mean	Standard Deviation	Q1	Median	Q3
log $COMP_t$	8,516	6.291	.570	5.922	6.290	6.652
RET_t	8,516	.126	.314	−.066	.085	.269
ROE_t	8,516	.138	.085	.108	.142	.175
log $SALES_t$	8,516	7.467	1.115	6.740	7.467	8.182
$FROE_t$	5,735	.154	.081	.109	.144	.186
Δlog $COMP_t$	8,516	.047	.253	−.038	.040	.130
ΔRET_t	8,516	−.016	.471	−.256	−.009	.226
ΔROE_t	8,516	−.003	.070	−.020	.001	.017
Δlog $SALES_t$	8,516	.043	.150	−.024	.038	.107
$\Delta FROE_t$	5,735	.017	.074	−.011	.015	.036

COMP = salary plus bonus (in thousands). RET = firm's market return. ROE = firm's return on common equity. SALES = firm's sales (in millions). FROE = consensus analyst forecast of ROE.

We further examine our hypothesis by running separate regressions at the industry and firm levels. We first run separate OLS regressions for each two-digit SIC industry. We employ two specifications here, first using only year-t firm performance as independent variables, then using stock returns in years $t - 2$ through t and year-t values of accounting variables.[14] We construct a Z-statistic to test the hypothesis that the average coefficient on Δlog $COMP_t$ is positive. In Panel A of Table 3, we list two Z-statistics for each specification: one corresponding to industry regressions on the full sample and one corresponding to industry regressions where we drop any two-digit industry that does not contain at least 20 observations. We also estimate separate regressions for each firm. In Panel B, we present two Z-statistics for each specification: one for the full sample and one corresponding to firm-by-firm regressions where we drop any firm for which we do not have at least 15 observations. In all four sets of industry-level regressions, we can reject the null hypothesis that compensation is uninformative about future performance at better than the 1% level.[15] The firm-level regressions allow us to reject the null when limiting our sample to firms for which we have at least 15 years of data (10% level), but not when we use the entire sample.

As a further check that unobservable information is used in contracting, we attempt to control for additional earnings-relevant information that may be publicly available before CEO bonuses are announced. Accounting literature on analysts' forecasts (see O'Brien, 1988) suggests that the consensus of analysts' forecasts is a good predictor of future earnings. From the Zacks database, we obtain forecasts of earnings per share (EPS) for year $t + 1$ that are issued before the end of year t. We use these EPS forecasts

[14] We limit our specifications in this way to economize on the number of parameters we must estimate. The Table 2, column (1) specification requires estimation of eight firm-performance parameters, while for our firm-by-firm regressions we have at most 18 observations for each firm. Using one lag of ΔROE and Δlog SALES and three lags of ΔRET in industry-level regressions, we obtain Z-statistics of 2.60 and 2.74 for the whole sample and the subsample of industries with at least 20 observations, respectively.

[15] The significance levels we attribute to the Z test depend on the assumption that the parameter estimates are independent across industries or firms. The significance level would be lower if the estimates are correlated.

TABLE 2 Estimates of Earnings-Relevant Information Contained in Executive Compensation

	Dependent Variable	
	ΔROE_{t+1}	
Independent Variable	(1)	(2)
$\Delta \log COMP_t$	.0076[b] (.0038)	.0087[b] (.0042)
$\Delta \log COMP_{t-1}$		.0039 (.0046)
ΔRET_t	.0434[a] (.0033)	.0455[a] (.0030)
ΔROE_t	−.3805[a] (.0265)	−.3886[a] (.0274)
$\Delta \log SALES_t$	.0104[c] (.0062)	.0118[c] (.0063)
ΔRET_{t-1}	.0422[a] (.0038)	.0441[a] (.0041)
ΔROE_{t-1}	−.1874[a] (.0238)	−.1887[a] (.0257)
$\Delta \log SALES_{t-1}$	−.0217[a] (.0063)	−.0242[a] (.0069)
ΔRET_{t-2}	.0243[a] (.0035)	.0253[a] (.0038)
ΔRET_{t-3}	.0116[a] (.0028)	.0110[a] (.0029)
N	8,516	7,999
Adjusted R^2	.14	.14

Fixed effects for year and two-digit industry included. Heteroskedastic-consistent standard errors in parentheses. Significance at the 1%, 5%, and 10% levels (two-tailed) is denoted by a, b, and c, respectively. COMP = salary plus bonus. RET = firm's market return. ROE = firm's return on common equity. SALES = firm's sales.

to construct a consensus forecast of ΔROE. We take the average of EPS forecasts, multiply by the number of shares outstanding at the end of year t, and then divide this by the firm's common equity at the end of year t. In Table 4 we present results from including the consensus forecast of change in ROE (which we denote by ΔFROE) in the specifications shown in Table 2. We find that, even controlling for information contained in the consensus of analysts' forecasts of year $t + 1$ earnings, year-t compensation is informative about earnings. In column (1), we reject the hypothesis that compensation is uninformative about future earnings at better than the 7% level, while in column (2) our estimate is significant at the 3% level.

This finding has implications for the nature of the information used by compensation committees in rewarding managers. If z consists of information that is public but just not contained in financial performance measures, then analysts ought to be able to examine the same information—compensation should, in this event, not contain any incremental information about future performance. Since analysts' forecasts appear not to be sufficient for compensation with respect to future earnings, the results suggest

TABLE 3 **Z-statistics Testing Hypothesis that Coefficient on Δlog COMP$_t$ Equals Zero in Industry- and Firm-level Regressions**

Panel A: Separate Regressions for Each Two-digit Industry

	Full Sample		Industries with 20 or More Observations	
Independent Variables	Z-statistic	Number of Regressions	Z-statistic	Number of Regressions
Year-t firm performance	3.12[a]	56	3.03[a]	52
Additional lags of ΔRET	2.77[a]	56	2.88[a]	52

Panel B: Separate Regressions for Each Firm

	Full Sample		Firms with 15 or More Observations	
Independent Variables	Z-statistic	Number of Regressions	Z-statistic	Number of Regressions
Year-t firm performance	1.49	366	1.73[c]	99
Additional lags of ΔRET	1.00	378	1.77[c]	133

Significance at the 1%, 5%, and 10% levels (two-tailed) is denoted by superscript a, b, and c, respectively.

that compensation committees are using some information that is not available to analysts at the time the forecasts are issued.

We take these results to be strong evidence of a positive connection between unexplained variation in current cash compensation and future firm performance. The finding is robust to a number of different specifications and inclusion of a number of different control variables. Parameter estimates on Δlog COMP$_t$ from Tables 2 and 4 range from .0076 to .0119, suggesting that a change in compensation that is 10% higher than would be predicted by currently observable performance is associated with a change in ROE that is higher by about .1 percentage points than would be predicted by current performance alone. This means inclusion of unexplained variation in compensation does not markedly improve the fit of the ROE model. We argue, however, that the economic significance of our findings does not come from the extent to which our regressions enhance our ability to forecast ROE. Rather, we interpret these results (combined with those in the next subsection) as indicating that boards of directors use information that is not available to those outside the relationship as part of an implicit incentive contract.

As further checks on the robustness of these results, we reestimate the regression of column (1) of Table 2 using three alternative approaches. First, since Hall and Liebman (1998) show that firms' use of discretionary stock option awards has increased consistently since the early 1980s, we consider whether the increasing use of option grants may have affected the relationship between current cash compensation and future performance. We allow our estimate of the coefficient on Δlog COMP$_t$ to vary monotonically over the sample by interacting compensation with a time trend. The results reported in column (1) of Table 5 show no trend in the relationship between current compensation and future performance, which suggests that increasing prevalence of stock options has not altered the relationship between current compensation and future performance. Next, we estimate our basic specification omitting all firms with non-December fiscal-year ends. This allows us to ascertain that our results are not driven by any errors in hand-matching compensation figures from *Forbes* to firm-performance

TABLE 4 Estimates of Earnings-Relevant Information Contained in Executive Compensation

Independent Variable	Dependent Variable ΔROE_{t+1} (1)	(2)
$\Delta \log COMP_t$	.0089[c] (.0048)	.0119[b] (.0052)
$\Delta \log COMP_{t-1}$		.0078 (.0050)
$\Delta FROE_t$	.3283[a] (.0299)	.3497[a] (.0282)
ΔRET_t	.0508[a] (.0038)	.0519[a] (.0040)
ΔROE_t	−.2393[a] (.0274)	−.2339[a] (.0269)
$\Delta \log SALES_t$	.0221[a] (.0081)	.0253[a] (.0081)
ΔRET_{t-1}	.0494[a] (.0044)	.0482[a] (.0047)
ΔROE_{t-1}	−.1257[a] (.0228)	−.1264[a] (.0237)
$\Delta \log SALES_{t-1}$	−.0238[a] (.0077)	−.0286[a] (.0080)
ΔRET_{t-2}	.0325[a] (.0040)	.0301[a] (.0043)
ΔRET_{t-3}	.0184[a] (.0035)	.0162[a] (.0036)
N	5,735	5,476
Adjusted R^2	.22	.22

See notes to Table 2. FROE = consensus analyst forecast of ROE.

information from Compustat. As shown in column (2) of Table 5, the coefficient on $\Delta \log COMP_t$ remains significant at the 10% level and the point estimate is very close to that shown in Table 2. Finally, we apply a more sophisticated ROE forecasting model developed by Fairfield, Sweeney, and Yohn (1996). This model breaks current earnings down into accounting classifications and allows the various components of earnings to have different predictive content. To estimate this model, we replace ΔROE_t and ΔROE_{t-1} with components of current and lagged earnings divided by common equity.[16] Using this model, the estimated coefficient on $\Delta \log COMP_t$ is .0099, which is comparable in magnitude to our other estimates and significant at better than the 5% level.

We performed, but do not report, several additional robustness checks. Conducting our analysis using return on assets rather than return on equity does not change the main findings. Using the level of ROE as a dependent variable also leads us to similar

[16] Specific components of earnings used are gross margin, selling, general and administrative expenses, depreciation expense, interest expense, minority income, nonoperating income, income tax expense, and special items.

TABLE 5 Robustness Checks

	Dependent Variable		
Independent Variable	Time Trend	ΔROE_{t+1} December Fiscal-Year Ends Only	ROE Components Model
$\Delta \log COMP_t$	.0070[b] (.0035)	.0072[c] (.0043)	.0099[b] (.0048)
$\Delta \log COMP_t \cdot TREND$	.0002 (.0006)		
ΔRET_t	.0432[a] (.0032)	.0478[a] (.0042)	.0356[a] (.0037)
ΔROE_t	−.3802[a] (.0265)	−.3906[a] (.0294)	
$\Delta \log SALES_t$	.0107[c] (.0062)	.0183[a] (.0071)	−.0030 (.0078)
ΔRET_{t-1}	.0420[a] (.0038)	.0463[a] (.0047)	.0345[a] (.0044)
ΔROE_{t-1}	−.1873[a] (.0238)	−.1823[a] (.0264)	
$\Delta \log SALES_{t-1}$	−.0219[a] (.0063)	−.0189[b] (.0075)	−.0283[a] (.0074)
ΔRET_{t-2}	.0242[a] (.0035)	.0265[a] (.0043)	.0219[a] (.0039)
ΔRET_{t-3}	.0116[a] (.0028)	.0113[a] (.0034)	.0122[a] (.0030)
N	8,516	6,744	6,458
Adjusted R^2	.14	.14	.20

See notes to Table 2. TREND = linear time trend taking value zero for 1985.

conclusions.[17] We obtained similar results when using generalized least squares accounting for firmwise heteroskedasticity. We also applied a procedure suggested by Hatanaka (1974) to correct for autocorrelation in the residuals. We estimated a separate autocorrelation parameter for each firm and then used Cochrane-Orcutt differencing to remove the autocorrelation. The results are not qualitatively different from those presented. Finally, we included in our sample those observations for which the magnitude of ΔROE_{t+1} is greater than .5 and estimated parameters using median regression. These estimates were also comparable to those presented.

□ **Factors affecting the strength of the relationship between unexplained variation in current compensation and future performance.** We next examine our second hypothesis, that unexplained variation in compensation is a better predictor of future earnings when current publicly observable measures are less valuable for contracting purposes. To test this assertion, we first develop proxies for the precision and sensitivity of current publicly observable performance measures. When public measures are less

[17] We also experimented with longer leads of ROE as a dependent variable. While most parameters retained the same signs as in the results presented here, the estimates were not generally statistically significant.

precise or less sensitive measures of an executive's actions, we expect contracts to depend more heavily on privately observed measures of performance. We expect the relationship between unexplained variation in current compensation and future performance to be stronger in this case. Our empirical strategy is to interact these proxy variables with Δlog COMP. This methodology allows the coefficient on Δlog COMP to vary monotonically with the precision and sensitivity of observable performance measures.

As measures of the precision of public firm performance measures, we compute the variances of the firm's market and accounting rates of return. We denote by VRET_t the log of the variance of the firm's monthly stock market returns over the 60 months before the beginning of the firm's fiscal year *t*. We denote by VROE_t, VROA_t, and VROS_t the log of the variances of the firm's returns on equity, assets, and sales over the five years prior to fiscal year *t*.

We take several proxies for the sensitivity of observable performance measures to current executive actions from prior research. Bushman, Indjejikian, and Smith (1996) obtain survey responses from large firms on the extent to which executive compensation is based on "individual" performance measures (as opposed to firmwide financial measures), while Ittner, Larcker, and Rajan (1997) examine compensation committee reports contained in firms' proxy statements for evidence of use of "nonfinancial" measures of performance. Both studies attempt to test the agency-theoretic prediction that such measures should be used more extensively when there is less performance-relevant information contained in financial measures of performance. These authors suggest a number of possible proxies for the sensitivity of currently observable performance measures, and we take five variables from their analyses. We let MTB_t, EMPS_t, and RDS_t represent the firm's market-to-book, employees-to-sales, and R&D-expense-to-sales ratios, respectively, where each ratio is computed as the average over the five fiscal years prior to year *t*. Following Bushman, Indjejikian, and Smith (1996), we take measures of product development and life cycles from the National Academy of Engineering's industry-level categorization of product development cycle time and product life cycle time. We classify cycle times as either long or short, where long cycle times are those exceeding four years, and define the dummy variables DEV and LIFE to be one if the cycle time is long. Bushman, Indjejikian, and Smith (1996) and Ittner, Larcker, and Rajan (1997) argue that high values for these five variables may indicate that current executive actions are directed more toward affecting the firm's future performance, and thus that current firm performance is not as sensitive to current actions. If

TABLE 6 Descriptive Statistics for Explanatory Variables

Variable	Definition	Number of Observations	Mean	Standard Deviation	Q1	Median	Q3
VRET	Log(variance of firm's market return)	8,213	−5.01	.59	−5.38	−5.02	−4.64
VROA	Log(variance of ROA)	8,213	−9.92	2.62	−11.49	−9.57	−7.99
VROE	Log(variance of ROE)	8,213	−7.48	2.04	−8.83	−7.52	−6.16
VROS	Log(variance of ROS)	8,213	−9.07	2.13	−1.29	−8.92	−7.70
MTB	Market-to-book ratio	7,052	1.67	1.22	.92	1.30	2.01
EMPS	Number of employees/sales	6,879	.011	.014	.005	.009	.014
RDS	R&D/sales	7,149	.011	.026	.000	.000	.009

TABLE 7 **Correlations between Explanatory Variables**

	VRET	VROA	VROE	VROS	MTB	EMPS	RDS	DEV	LIFE
VRET	1.00								
VROA	.33[a]	1.00							
VROE	.38[a]	.67[a]	1.00						
VROS	.12[a]	.36[a]	.64[a]	1.00					
MTB	.11[a]	.27[a]	.09[a]	−.07[a]	1.00				
EMPS	.11[a]	.05[a]	−.07[a]	−.12[a]	.03[a]	1.00			
RDS	.16[a]	.30[a]	.13[a]	.11[a]	.30[a]	.04[a]	1.00		
DEV	−.24[a]	.20[a]	.02[c]	.10[a]	−.04[a]	−.15[a]	.08[a]	1.00	
LIFE	−.23[a]	.24[a]	.01	.04[a]	.02[b]	−.12[a]	.04[a]	.86[a]	1.00

Significance at the 1%, 5%, and 10% levels is denoted by superscript a, b, and c, respectively.

this is the case, then we expect a positive relationship between these variables and the coefficient on $\Delta\log \text{COMP}_t$. Summary statistics and correlations for these variables are given in Table 6 and Table 7, respectively. Limits on availability of this set of variables reduce our sample to 8,213 observations.

We begin by exploring whether the strength of the relationship between unexplained variation in current compensation and future performance is affected by the precision of current public measures of firm performance. To the basic specification estimated in column (1) of Table 2, we add the interaction terms $\Delta\log \text{COMP}_t \cdot \text{VRET}_t$ and $\Delta\log \text{COMP}_t \cdot \text{VROE}_t$. We also include VRET and VROE in the regression by themselves to pick up any direct relationship between these variables and ΔROE_{t+1}. Since it is the interaction terms that are of primary interest, we omit the coefficients on the direct effects of VRET and VROE from our tables.

We estimate using OLS and report results along with heteroskedastic-consistent standard errors in column (1) of Table 8. In columns (2) and (3), we replace VROE with VROA and VROS, respectively. In column (1), the coefficients on the interaction terms with VRET and VROE are both positive, as expected, and the coefficient on the VROE interaction is significantly different from zero at the 1% level. Note that since VROE_t and VRET_t are correlated (see Table 7), the individual effects may be imprecisely estimated. We therefore assess the joint significance of $\Delta\log \text{COMP}_t \cdot \text{VROE}_t$ and $\Delta\log \text{COMP}_t \cdot \text{VRET}_t$ by performing a Wald test. We reject the hypothesis that both are equal to zero at better than the 1% level. In column (2), the effect again appears to load solely onto $\Delta\log \text{COMP}_t \cdot \text{VROA}_t$. This coefficient is significant at better than the 2% level, while the coefficient on $\Delta\log \text{COMP}_t \cdot \text{VRET}_t$ is not significantly different from zero. However, we can again reject the hypothesis that both are zero at better than the 1% level. In column (3), the coefficients on $\Delta\log \text{COMP}_t \cdot \text{VRET}_t$ and $\Delta\log \text{COMP}_t \cdot \text{VROS}_t$ are both significant at the 5% level, and we again reject (at the 3% level) the hypothesis that both $\Delta\log \text{COMP}_t \cdot \text{VRET}_t$ and $\Delta\log \text{COMP}_t \cdot \text{VROS}_t$ are equal to zero.

We interpret these results in Table 8 as being consistent with our hypothesis. It appears that unexplained variation in current compensation is a stronger predictor of future performance when public performance measures are more variable. This is consistent with the agency-theoretic prediction that firms and employees will substitute away from noisier performance measures in employment contracts.

TABLE 8 **OLS Estimates of Earnings-Relevant Information Contained in Executive Compensation. Compensation Variable Interacted with Proxies for Precision of Publicly Observable Measures of Performance**

Independent Variable	Dependent Variable ΔROE_{t+1} (1)	(2)	(3)
Δlog $COMP_t$	.0880[b] (.0378)	.1067[a] (.0385)	.1120[a] (.0430)
Δlog $COMP_t \cdot VRET_t$	.0083 (.0077)	.0118 (.0074)	.0143[b] (.0071)
Δlog $COMP_t \cdot VROE_t$	.0060[a] (.0023)		
Δlog $COMP_t \cdot VROA_t$		.0045[b] (.0017)	
Δlog $COMP_t \cdot VROS_t$			.0041[b] (.0020)
ΔRET_t	.0473[a] (.0034)	.0472[a] (.0034)	.0472[a] (.0033)
ΔROE_t	−.3911[a] (.0274)	−.3900[a] (.0273)	−.3893[a] (.0273)
Δlog $SALES_t$	.0118[c] (.0067)	.0108 (.0067)	.0105 (.0067)
ΔRET_{t-1}	.0454[a] (.0040)	.0453[a] (.0040)	.0450[a] (.0040)
ΔROE_{t-1}	−.1872[a] (.0240)	−.1863[a] (.0239)	−.1860[a] (.0239)
Δlog $SALES_{t-1}$	−.0187[a] (.0064)	−.0191[a] (.0064)	−.0191[a] (.0065)
ΔRET_{t-2}	.0264[a] (.0037)	.0265[a] (.0036)	.0262[a] (.0037)
ΔRET_{t-3}	.0111[a] (.0028)	.0111[a] (.0028)	.0109[a] (.0028)
N	8,213	8,213	8,213
Adjusted R^2	.15	.15	.15

See notes to Table 2. VRET = log(variance of firm's market return). VROA = log(variance of ROA). VROE = log(variance of ROE). VROS = log(variance of ROS).

We next explore the proxies for the sensitivity of current firm performance to current executive actions. In column (1) of Table 9, we add interactions of RDS, MTB, EMPS, and DEV with Δlog COMP to the specification shown in column (1) of Table 8. In column (2), we replace DEV with LIFE. As above, we include the direct effects of these variables in our regression but omit them from our table.

The two regressions offer similar results. The coefficients on Δlog $COMP_t \cdot VRET_t$ and Δlog $COMP_t \cdot VROE_t$ are positive and significant at better than the 10% level. In both regressions, we reject the hypothesis that both coefficients equal zero at better than the 1% level. The signs of the coefficients on the sensitivity proxies are, for the most part, consistent with our hypothesis. In both regressions, three of the four are positive, with MTB presenting the lone exception. The ratio of employees to sales has

TABLE 9 OLS Estimates of Earnings-Relevant Information Contained in Executive Compensation. Compensation Variable Interacted with Proxies for Sensitivity of Publicly Observable Measures of Performance

Independent Variable	Dependent Variable ΔROE_{t+1} (1)	(2)
$\Delta\log COMP_t$	.1309[a]	.1260[a]
	(.0478)	(.0472)
$\Delta\log COMP_t \cdot VRET_t$	.0186[c]	.0168[c]
	(.0102)	(.0099)
$\Delta\log COMP_t \cdot VROE_t$	.0049[c]	.0052[c]
	(.0027)	(.0027)
$\Delta\log COMP_t \cdot MTB_t$	−.0041	−.0045
	(.0042)	(.0041)
$\Delta\log COMP_t \cdot EMPS_t$	.3507[c]	.3220[c]
	(.1931)	(.1844)
$\Delta\log COMP_t \cdot RDS_t$	.0145	.0344
	(.1658)	(.1641)
$\Delta\log COMP_t \cdot DEV$	.0192[c]	
	(.0105)	
$\Delta\log COMP_t \cdot LIFE$		.0140
		(.0096)
ΔRET_t	.0497[a]	.0497[a]
	(.0036)	(.0036)
ΔROE_t	−.4155[a]	−.4161[a]
	(.0286)	(.0287)
$\Delta\log SALES_t$	.0114	.0115
	(.0074)	(.0074)
ΔRET_{t-1}	.0502[a]	.0501[a]
	(.0044)	(.0044)
ΔROE_{t-1}	−.1914[a]	−.1912[a]
	(.0242)	(.0242)
$\Delta\log SALES_{t-1}$	−.0165[b]	−.0165[b]
	(.0070)	(.0070)
ΔRET_{t-2}	.0297[a]	.0296[a]
	(.0039)	(.0040)
ΔRET_{t-3}	.0141[a]	.0141[a]
	(.0032)	(.0032)
N	6,794	6,794
Adjusted R^2	.15	.15

See notes to Table 2. VRET = log(variance of firm's market return). VROE = log(variance of ROE). MTB = Market-to-book ratio. EMPS = employees-to-sales ratio. RDS = Research & development expenses divided by sales. DEV, LIFE = dummy variables for long product development cycle and life cycle, respectively.

the strongest effect; its coefficient is positive and significant at better than the 10% level in both regressions. Product cycle times appear to have some explanatory power as well, with the variable DEV entering significantly in column (1). LIFE, however, is not significant ($p = .15$) in column (2). Coefficients on RDS and MTB are positive and negative, respectively, but statistically indistinguishable from zero.

We interpret the results using the sensitivity proxies as being moderately supportive of our hypothesis, although they are perhaps not as strong as those involving the precision proxies in Table 8. With the exception of market-to-book ratio, the sensitivity proxies appear with the expected sign in both specifications. The ratio of employees to sales is significantly positive in both regressions, while one of the two product cycle time variables is significant.

These results have some similarities to findings of previous authors. Bushman, Indjejikian, and Smith (1996) report that firms with longer product cycles (as measured by DEV and LIFE) are likely to rely more heavily on individual performance measurement, whereas we find some evidence that longer product cycles lead to a stronger positive relationship between current compensation and future performance. However, they also find that high market-to-book ratios are indicative of increased use of individual performance measurement, whereas our coefficient on the interaction of this variable with $\Delta\log$ COMP is opposite to expectations (but insignificant). In addition, they do not detect a relationship between the variance of market returns and the the use of individual performance measurement. Ittner, Larcker, and Rajan's (1997) strongest result is a positive relationship between a latent variable they call STRATEGY, of which the employee-to-sales and R&D-to-sales ratios are two components, and the use of nonfinancial performance measures. Comparably, we find that higher employees-to-sales ratios appear to be associated with stronger connections between current compensation and future performance. They also report a significant positive coefficient on a second latent variable that depends on the industry standard deviations of accounting returns.[18]

We repeat the robustness checks shown in Table 5 by applying similar procedures to the specification in column (1) of Table 8. In Table 10, we first interact the time trend with $\Delta\log$ COMP, $\Delta\log \text{COMP}_t \cdot \text{VRET}_t$, and $\Delta\log \text{COMP}_t \cdot \text{VROE}_t$ to examine whether the relationship between the variance of observable performance measures and the predictive content of current compensation was affected by the growing use of stock options through the 1980s and early 1990s. None of the trend interaction terms are significantly different from zero. We also estimate this model using only firms with December fiscal-year ends and using Fairfield, Sweeney, and Yohn's (1996) accounting-classification-based ROE-forecasting model. In all cases, the results are in line with expectations. Coefficients on $\Delta\log \text{COMP}_t \cdot \text{VROE}_t$ are significantly different from zero, and the hypothesis that the coefficients on both $\Delta\log \text{COMP}_t \cdot \text{VROE}_t$ and $\Delta\log \text{COMP}_t \cdot \text{VRET}_t$ are zero is rejected.[19]

We performed several additional robustness checks for this section. We estimated the models of this section using ROA rather than ROE as a dependent variable, using the level of ROE as a dependent variable, adding additional lags of performance, using

[18] Ittner, Larcker, and Rajan (1997) find some additional explanatory power in a latent variable based on the correlation between market and accounting returns, but only when restricting attention to a subsample of manufacturing firms. When limiting our sample to manufacturing firms, we found this correlation variable approaches significance ($p = .13$). Including this variable did not materially affect our other estimates.

[19] Estimating these specifications using the sensitivity proxies yields results comparable to those in Table 9. Estimates of coefficients on interactions between $\Delta\log \text{COMP}_t$ and EMPS, DEV, and LIFE are similar in magnitude, with somewhat higher standard errors.

TABLE 10 Robustness Checks

Independent Variable	Dependent Variable		
	Time Trend	ΔROE_{t+1} December Fiscal-Year Ends Only	ROE Components Model
$\Delta\log COMP_t$	.0784[b] (.0368)	.0897[b] (.0403)	.0996[b] (.0472)
$\Delta\log COMP_t \cdot VRET_t$	.0075 (.0069)	.0073 (.0081)	.0106 (.0106)
$\Delta\log COMP_t \cdot VROE_t$	.0052[b] (.0023)	.0072[a] (.0026)	.0059[b] (.0026)
$\Delta\log COMP_t \cdot TREND$	.0018 (.0068)		
$\Delta\log COMP_t \cdot VRET_t \cdot TREND$	−.0002 (.0013)		
$\Delta\log COMP_t \cdot VROE_t \cdot TREND$	.0005 (.0004)		
ΔRET_t	.0472[a] (.0034)	.0510[a] (.0043)	.0391[a] (.0037)
ΔROE_t	−.3910[a] (.0274)	−.4092[a] (.0303)	
$\Delta\log SALES_t$	.0117[c] (.0067)	.0208[a] (.0078)	−.0026 (.0080)
ΔRET_{t-1}	.0453[a] (.0040)	.0501[a] (.0048)	.0373[a] (.0045)
ΔROE_{t-1}	−.1874[a] (.0239)	−.1848[a] (.0266)	
$\Delta\log SALES_{t-1}$	−.0189[a] (.0065)	−.0149[c] (.0078)	−.0219[a] (.0073)
ΔRET_{t-2}	.0264[a] (.0037)	.0283[a] (.0044)	.0237[a] (.0038)
ΔRET_{t-3}	.0111[a] (.0028)	.0113[a] (.0034)	.0118[a] (.0030)
N	8,213	6,502	6,210
Adjusted R^2	.15	.15	.21

See notes to Table 2. VRET = log(variance of firm's market return). VROE = log(variance of ROE). TREND = linear time trend taking value zero for 1985.

generalized least squares accounting for firmwise heteroskedasticity, controlling for autocorrelation using the method of Hatanaka (1974), and using median regression on a sample that includes observations where the magnitude of ΔROE_{t+1} is greater than .5. Our findings of fairly strong evidence that strength of compensation/future-performance relationship is positively related to the variability of observable performance measures and somewhat weaker evidence that this link is inversely related to the sensitivity of current performance to executive actions are robust to these changes.

4. Discussion

■ Before concluding, we discuss two potential alternative explanations for our results, both of which are closely related to our basic hypothesis that privately observed information is used in contracting. First, we note that while our discussion in Section 2 treats the properties of the privately observed measure as exogenous, endogenizing the properties of z yields a complementary interpretation of our results. Consider a firm for which the publicly observable measures are of exogenously low quality. The value of investments that enhance the quality of z (for example, attracting board members who are more skilled at monitoring managers) will be relatively high for such firms. Hence, one might expect firms with low-quality observable performance measures to *both* obtain better privately observed measures *and* place greater weight on such measures. We emphasize the contracting story in the text because it is present whether the properties of z are endogenous or exogenous. If a firm with exogenously low-quality observable performance measures makes investments that enhance the quality of z, then the firm should substitute toward use of z in contracting with its manager. Bushman et al. (1999) explore this issue by analyzing the relation between board composition and properties of observable measures of managerial performance.

A specific version of an income-smoothing story yields another potential explanation. Consider an executive who manipulates accounting accruals to smooth year-to-year fluctuations in earnings. Suppose "true" earnings are high in the current year and the executive defers some income until the next year. Then if the current year's bonus payment is based on "true" earnings, current compensation will be unexpectedly high relative to reported earnings. In addition, next year's reported earnings will be higher than expected given current performance, since income is being shifted from the present to the future. This generates a link between current compensation and future performance like that predicted by our first hypothesis. If this behavior is more common among firms whose earnings are more variable (and the earnings *remain* more variable after the smoothing), then results similar to our second hypothesis may apply. Note, however, that since this version of earnings smoothing requires that compensation be based on the "true" realization of earnings (as opposed to reported earnings), it is still fundamentally a story in which information that is not publicly available is used to measure and reward managerial performance.

Most academic research on income smoothing emphasizes a markedly different idea from that described here. This literature focuses on the role of explicit earnings-based bonus contracts or managers' career concerns in inducing managers to manipulate accounting figures, and typically assumes either that it is difficult for the board of directors to undo the effects of such manipulations, or that the board wishes the manager to smooth fluctuations in income. (See, for example, Dye (1988), Gaver, Gaver, and Austin (1995), and Fudenberg and Tirole (1995).) If boards are unable to undo the effects of income smoothing or wish to induce managers to undertake income smoothing, then current bonuses should match current earnings, and there does not arise a link between current compensation and future performance. While we are unable to rule out the version of the income smoothing described in the previous paragraph as a possible alternative explanation for our results, we note that this story requires that managers engage in income-smoothing behavior even though their compensation reflects "true" (and hence privately observed) earnings.

5. Conclusion

■ In this article we have developed a framework for analyzing the use of unobservable (to outsiders) information in executive compensation contracts. We argued that if firms use performance measures that are observable only to those inside the firm to reward top executives *and* these measures of performance are correlated with future indicators of performance, then variation in *current* compensation that is not explained by variation in current observable performance measures should predict *future* variation in observable performance measures. This assertion led us to examine the question of whether earnings-relevant information is contained in executive compensation.

We then applied the results of agency models that consider settings with multiple performance signals to argue that as publicly observable measures of performance become more noisy, the parties to the contract should rely more heavily on the privately observed performance measure. We showed that as the parties shift more weight onto the privately observed performance measure, compensation ought to become more positively related to future earnings.

We tested these two hypotheses using executive compensation data taken from the *Forbes* Executive Compensation Surveys. We found strong evidence to suggest that unexplained variation in current compensation is related to future performance. Controlling for current performance and analysts' forecasts of future performance, our regression estimates allowed us to reject the hypothesis that compensation is unrelated to future performance. We showed this finding to be robust to many alternative specifications. This finding is consistent with the hypothesis that boards of directors use information that is not available to those outside the relationship as part of an implicit incentive compensation contract.

We find additional evidence suggesting that when the variance of publicly observable measures of performance is higher, the relationship between unexplained variation in current compensation and future performance is stronger. This is consistent with firms substituting away from public performance measures toward measures that are unobservable to outsiders as the public measures become noisier. Our estimates suggest that unexplained variation in current compensation and future performance are more closely linked as the variances of both market and accounting returns increase. From prior work on firms' use of individual or nonfinancial performance measures, we take several additional proxies for the sensitivity of currently observable measures of performance to executive actions. Our findings that compensation is more closely related to future earnings when employees-to-sales ratios are higher and when product development cycles are longer are similar to those reported by Bushman, Indjejikian, and Smith (1996) and Ittner, Larcker, and Rajan (1997).

The primary caveat that must be applied to our analysis is the potential presence of omitted firm-performance-related variables. While we have experimented with a large variety of different explanatory variables, including contemporaneous and lagged performance variables and various measures of accounting returns, there is no way to systematically rule out the possibility that any connection we find between unexplained variation in compensation and future performance is due to the omission of an observable performance variable that is correlated with both current compensation and future performance.

Future work could usefully proceed in three directions. First, it may be possible to extend this work to analyze market measures of firm performance using an event-study methodology. Such an analysis would require some way of assessing when news about compensation amounts is revealed to the market. Dates of firms' proxy disclosures may suffice, but board compensation committees typically determine compensation amounts well in advance of the release of annual proxy statements. Second, our analysis suggests that disclosure of compensation amounts may convey information to market participants. Given this, it may be useful to consider whether firms have an incentive to choose compensation amounts strategically to affect outsiders' perceptions of the value of the firm. Third, it may be possible to link Bushman et al.'s (1999) analysis with ours by examining how board characteristics affect the relationship between current compensation and future performance.

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