

An Analysis of the Use of Accounting and Market Measures of Performance in Executive Compensation Contracts

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1. Introduction

Prior research has provided useful insights into the structure of compensation plans and their incentive effects.¹ However, one important limitation of these studies is the virtual absence of any cross-sectional analyses of the attributes of compensation contracts.² This absence is related, in part, to the problems associated with controlling for "other factors" that affect compensation. That is, compensation contracts are

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¹ For example, Murphy [1985] and Coughlan and Schmidt [1985] correct some of the econometric problems of earlier research and conclude that executive compensation is, on average, strongly and positively related to the firm's stock market rate of return. Antle and Smith [1986] provide empirical evidence that executive compensation is, on average, more sensitive to the firm's return on equity relative to their industry than to the gross return on equity. Finally, Healy, Kang, and Palepu [1987] and Defeo, Lambert, and Larcker [1987] examine whether compensation contracts seem to be adjusted for accounting changes or the accounting effect of financial reorganizations made by the firm. Lambert and Larcker [1985a; 1985b] provide reviews of this research in the area of executive compensation.

² Two recent working papers provide some cross-sectional analysis of executive compensation contracts. Ely [1987] provides evidence that the relative weight placed on alternative measures of performance varies by industry. The theoretical framework employed in Adams [1987] is very similar to ours, although her empirical analysis is substantially different.

thought to be functions of characteristics of the manager, the firm, and the environment.³ In our attempt to address some of these complexities, we employ an analytical agency theory model to provide a structure for controlling for "other factors" in order to examine the "informational properties" of accounting and market measures of performance.

More specifically, we exploit the results of Holmstrom [1979] and others which suggest that the relative weight placed on a performance measure in a compensation contract is an increasing function of its "signal-to-noise" ratio with respect to the agent's actions. Using this framework, we empirically examine whether the relative use of security market and accounting measures of performance in executive compensation is related to the amount of "noise" inherent in the two signals and the "sensitivity" of these two signals to managerial actions. Our results are consistent with the hypothesis that firms place relatively more weight on market performance (and less weight on accounting performance) in compensation contracts for situations in which (i) the variance of the accounting measure of performance is high relative to the variance of the market measure of performance, (ii) the firm is experiencing high growth rates in assets and sales, and (iii) the value of the manager's personal holdings of his firm's stock is low.

Section 2 provides a theoretical framework for analyzing the use of performance measures in compensation contracts. The sample selection procedures are discussed in section 3. In section 4, we discuss the measurement of the performance variables and managerial compensation and present the results of an analysis of the time-series relation between cash compensation and accounting and market performance. In section 5, we develop a number of hypotheses regarding the informational properties of accounting and market measures of performance. In section 6, we examine whether the cross-sectional variation in the relative weights placed on security market versus accounting performance in compensation contracts is related to the signal-to-noise ratios of these performance variables. The research results are summarized in section 7.

2. Theoretical Framework

Although agency theory does not provide insights into the use of accounting or market numbers as *specific* measures of performance in

³ In contrast, many studies implicitly treat the compensation contract as if it were exogenous. For example, one line of research has examined whether the contracts (either cross-sectional differences in contracts or time-series changes in contracts) are associated with managerial decision making (e.g., Larcker [1983] and Healy [1985]). Unfortunately, it is unclear whether it is appropriate to attribute any association between the adoption of a contract and changes in managerial decisions to (i) the contractual change or (ii) other variables which are the causal determinants of the contractual change. This criticism also applies to studies which have examined the security market reaction to compensation contract changes (e.g., Lambert and Larcker [1985a; 1985b] and Brickley, Bhagat, and Lease [1985]).

compensation contracts, the theory does provide a framework for structuring our empirical analysis. In particular, agency theory can be used to specify the properties of any two generic variables that are relevant for evaluating an agent's performance. Moreover, the theory suggests both a functional form relating the informational properties of the performance variables to parameters of the agent's compensation scheme, and a means of controlling for "other" effects on the form of the compensation scheme so that the analysis can focus on informational properties.

Since most agency models examine incentive problems in a *single-period* setting, we begin our analysis of the use of performance measures in compensation contracts by discussing the implications of these single-period models. We then examine the effect of multiperiod considerations on the relation between compensation and performance.

2.1 SINGLE-PERIOD AGENCY MODELS

The "standard" agency model (see Holmstrom [1979]) analyzes incentive problems that arise when one individual, the principal, delegates decision-making tasks to another individual, the agent.⁴ The agent's output, x , is a function of his action, a , and a random state of nature. The agent's action, which is typically interpreted as the amount of effort that he supplies, is assumed to increase the cash flow; however, the effect of the state of nature on the agent's output prevents the principal from using the output to determine unambiguously the amount of effort that was supplied. As a result, the principal must rely on imperfect measures of the agent's actions, such as his output and other information (y) for both evaluation and motivation.

Holmstrom [1979] shows that the agent's compensation as a function of the performance variables x and y , denoted $c(x, y)$, is the solution to the following equation (assuming an interior solution):

$$\frac{1}{U'[c(x, y)]} = \lambda + \mu \frac{f_a(x, y | a)}{f(x, y | a)} \quad (1)$$

where $U(\cdot)$ is the agent's utility function for money (we assume that the principal is risk neutral), $f(x, y | a)$ is the density function of the variables x and y given the agent's effort, $f_a(x, y | a)$ is the derivative of the density function with respect to the agent's effort, λ is the Lagrange multiplier on the constraint that specifies the lower bound on the level of expected utility that the contract can provide to the agent, and $\mu > 0$ is the Lagrange multiplier on the constraint that ensures that the agent's choice of effort be incentive compatible.

⁴ Our analysis concentrates on incentive problems that arise from the unobservability of the agent's action (i.e., moral hazard problems). Other contracting problems, such as those that arise when the agent's skill is unknown (i.e., reputation and screening issues), are beyond the scope of our analysis. Another limitation of the agency models that we employ is that they assume that the agent is responsible only for deciding how much effort to supply; decisions involving a "risk-return" trade-off are ignored.

The performance variables x and y affect the agent's compensation through their effect on the term $f_a(x, y | a)/f(x, y | a)$. As discussed in Holmstrom [1979] and Milgrom [1981], this term is equivalent to the derivative of the logarithm of the likelihood function $f(x, y | a)$ with respect to the agent's effort, a . This provides an informational interpretation to the use of the variables x and y in the contract.⁵

Given the generality of the model, equation (1) provides little guidance regarding the functional form of the relation between the agent's compensation and the performance measures x and y . In order to derive empirical implications from this equation, we place additional structure on the model by making more specific assumptions about the form of the agent's utility function and the form of the probability distributions. In particular, we assume that the agent's utility function is a member of the power class of utility functions, which can be represented as:

$$U(c) = \frac{1}{1 - \kappa} c^{1-\kappa}. \quad (2)$$

This class of utility functions, which exhibits decreasing absolute risk aversion and constant proportional risk aversion, is commonly used in analytical work in agency theory.⁶ The parameter κ represents the coefficient of proportional risk aversion for the agent, where higher values of κ correspond to greater degrees of risk aversion. For this class of utility functions, $U'(c) = c^{-\kappa}$ and $1/U'(c) = c^\kappa$. Substituting the agent's marginal utility into equation (1) yields:

$$[c(x, y)]^\kappa = \lambda + \mu \frac{f_a(x, y | a)}{f(x, y | a)}. \quad (3)$$

We also impose some structure on the form of the probability distributions. As Banker and Datar [1987] show, the term $f_a(x, y | a)/f(x, y | a)$ is *linear* in x and y for a large class of probability distributions.⁷ In particular, we assume that:

$$\frac{f_a(x, y | a)}{f(x, y | a)} = \delta_0 + \gamma \{ \delta_x [x - E(x | a)] + \delta_y [y - E(y | a)] \}. \quad (4)$$

⁵ Note that although the principal's utility is a direct function of the firm's stock price (e.g., the variable x), agency theory does not imply that the optimal contract simply ties the agent's compensation exclusively to the firm's stock price. In particular, the random state of nature that affects security prices introduces noise into this performance measure *from the perspective of evaluating the agent's performance*. The existence of this noise implies that there are potential benefits to supplementing security prices with other measures of performance (e.g., accounting numbers) in evaluating the agent's actions.

⁶ There are also a variety of empirical studies which lend some support to the use of the power utility function (see Friend and Blume [1975] and Litzenberger and Ronn [1986]).

⁷ The class of distributions includes the exponential, normal (where the agent's effort affects the mean of x and y), and the binomial distributions, among others.

For the class of distributions considered in Banker and Datar [1987], the coefficients on the variables x and y can be interpreted as representing the “signal-to-noise ratios” of the performance variables. In particular, we have:

$$\delta_x = \frac{s(x|a)}{\text{var}(x|a)}, \quad \text{and} \quad \delta_y = \frac{s(y|a)}{\text{var}(y|a)}$$

where $s(\cdot|a)$ is the (conditional) sensitivity of the mean of the signal to the agent’s effort, and $\text{var}(\cdot|a)$ is the variance of the signal given the agent’s effort.⁸ Substituting equation (4) into equation (3) implies:

$$[c(x, y)]^\kappa = \lambda + \mu(\delta_0 + \gamma\{\delta_x[x - E(x|a)] + \delta_y[y - E(y|a)]\}). \quad (5)$$

Combining terms yields a linear expression on the right-hand side of equation (5):

$$[c(x, y)]^\kappa = \beta_0 + \beta_x[x - E(x|a)] + \beta_y[y - E(y|a)]. \quad (6)$$

The relation specified in equation (6) is complex; it is linear only if $\kappa = 1$, which corresponds to a logarithmic utility function for the agent.⁹ Moreover, the parameters λ and μ , which affect the intercept and the slopes on x and y , depend on the attractiveness of the manager’s outside employment opportunities, his disutility for effort, the “magnitude” of the agency problem, and the form of the production function. Thus, the slope coefficients (β_x and β_y) on the performance measures of interest are confounded by all of these factors, making interpretations of cross-sectional differences in the slope coefficients problematic.

However, one implication of the model is that for a given firm, if the performance measures are multiplied by their respective “signal-to-noise” ratios, the theory implies that the coefficients on the “rescaled” performance measures will be equal. To see this, equation (5) can be expressed as:

$$[c(x, y)]^\kappa = [\lambda + \mu \cdot \delta_0] + \mu \cdot \gamma\{\delta_x[x - E(x|a)] + \delta_y[y - E(y|a)]\}. \quad (7)$$

Equation (7) implies that the coefficients on the “rescaled variables” (i.e., $\delta_x[x - E(x|a)]$ and $\delta_y[y - E(y|a)]$) are equal to $\mu\gamma$. Although the

⁸ For example, if x and y are distributed as bivariate normal random variables whose means are affected by the agent’s effort, we have:

$$s(x|a) = \frac{\partial E(x|a)}{\partial a} - \frac{\text{cov}(x, y)}{\text{var}(y)} \frac{\partial(E(y|a))}{\partial a}.$$

⁹ The relation may also be nonlinear because equation (6) applies only for interior solutions for the compensation function. More generally, the optimal sharing rule will include a lower bound. In this case, equation (6) becomes:

$$c^\kappa = \max\{\underline{g}, \beta_0 + \beta_x[x - E(x|a)] + \beta_y[y - E(y|a)]\}.$$

magnitude of the coefficients on the “rescaled variables” will differ cross-sectionally as a function of $\mu\gamma$, the theory predicts that the coefficients on the “rescaled variables” will be equal for a given firm.

Another implication of the model is that the confounding effects of cross-sectional differences in the term $\mu\gamma$ can be reduced by computing the *ratio* of the slope coefficients from equation (6). That is, the ratio of the slope coefficients is:

$$\frac{\beta_x}{\beta_y} = \frac{\mu\gamma\delta_x}{\mu\gamma\delta_y} = \frac{\delta_x}{\delta_y} = \frac{s(x|a)}{s(y|a)} \frac{\text{var}(y|a)}{\text{var}(x|a)}. \quad (8)$$

Equation (8) implies that the ratio of the slope coefficients is a function of the ratio of the “signal-to-noise” ratios of the two performance variables. An increase in either the precision (i.e., the inverse of the noise) of a performance variable or its sensitivity to the agent’s actions will, *ceteris paribus*, increase the relative weight the variable receives in the compensation function.

2.2 MULTIPERIOD CONSIDERATIONS

Most multiperiod agency models suggest that compensation contracts have “memory” (e.g., Lambert [1983] and Rogerson [1985]). That is, the agent’s compensation in a period will depend not only on the realizations of the performance measures in that period but also on their realizations in *prior* periods.¹⁰ In particular, if we assume that the agent’s utility function is additively separable over time (in addition to the assumptions about the agent’s utility function and production function discussed in the previous section), Lambert [1983] shows that the optimal contract can be expressed as:

$$[c_t(\cdot)]^\epsilon - [c_{t-1}(\cdot)]^\epsilon \\ = \beta_0 + \beta_{xt}(\cdot)[x_t - E(x_t | \cdot)] + \beta_{yt}(\cdot)[y_t - E(y_t | \cdot)] \quad (9)$$

where the subscript t refers to time, the expected values for the performance measures x and y are now conditioned upon the actions in that period as well as prior realizations of the performance variables, and the slope coefficients are permitted to depend on the prior realizations of the performance variables.

From a theoretical perspective, little is known about how the realizations of the performance variables in one period affect the slope coefficients in subsequent periods. Moreover, from an empirical perspective, it is difficult to estimate the functional form of a compensation scheme whose slope coefficients vary over time as a function of the prior-period realizations of the performance measures. For these reasons, we suppress any dependence between the slope in one period and the realizations of

¹⁰ See Fellingham, Newman, and Suh [1987] for conditions on the form of the agent’s utility function for which the optimal contract does *not* involve memory.

the performance measures in prior periods. Therefore, we assume that the slope coefficients, β_x and β_y , are proportional to the time-series average of the signal-to-noise ratios of the performance measures.

Under the above assumptions, the compensation scheme can be expressed as a linear function that relates the *first difference* in compensation raised to the κ power to the "surprise" in the performance measures in the period:

$$c_t^\kappa - c_{t-1}^\kappa = \beta_0 + \beta_x[x_t - E(x_t | \cdot)] + \beta_y[y_t - E(y_t | \cdot)]. \quad (10)$$

In the empirical analysis, we estimate the slope coefficients of equation (10) for each firm using time-series data for compensation, market performance, and accounting performance. The estimated slope coefficients are then used to analyze whether the relative weights assigned to security market and accounting numbers in compensation contracts are related to the signal-to-noise ratios of these performance variables (as predicted in equations (7) and (8)).

3. Sample

Our sample of firms was primarily obtained from the *Forbes* annual compensation survey, which publishes compensation data for firms in at least one of the *Forbes* 500 listings (e.g., assets, sales, market value, or net profits). The survey includes data on cash compensation for the chief executive officer (*CEO*), years as *CEO*, years with the company, and for years 1970 to 1973 the value of shares owned by the *CEO* (and where appropriate, his family). The typical annual survey includes data for 700 to 800 firms.

We include firms in the sample if they satisfy five criteria. First, a complete history for cash compensation from 1970 to 1984 was available from various sources, including the *Forbes* data and proxy statements. Second, each firm was required to have the market value of equity ownership for the *CEO* during each year from 1970 to 1973 in the *Forbes* data. Third, the number of shares owned by the *CEO* (or his family) was reported in either the *Corporate Data Exchange (CDE) Stock Ownership Directory* or proxy statements available to the researchers.¹¹ Fourth, each firm was required to have selected financial data (discussed below) on the annual *Compustat* or *CRSP* files. Finally, each firm was required to have a constant fiscal year-end from 1970 to 1984.

The final sample of 370 firms consists of 188 industrial (not including natural resource firms) companies; 29 natural resource or petroleum-processing firms, 77 utilities or transportation companies, 25 retail or hotel firms, 48 banks or insurance companies, and 3 firms in unique industrial groups. These firms had a median of one *CEO* change (i.e., two

¹¹ For a small number of banks and finance companies, the *CEO* stock ownership data were obtained for the 1979 fiscal year.

separate *CEOs*) during the 15 years from 1970 to 1984. In addition, the executives had been employed by the firm a median of 27 years and had been the *CEO* of the firm for a median of six years.

4. *Estimation of the Slope Coefficients on the Performance Measures*

4.1 MEASUREMENT OF THE PERFORMANCE VARIABLES

Because agency theory does not identify the precise procedures for measuring the relevant accounting (corresponding to the variable y) and security market (corresponding to the variable x) performance indexes, we make this choice on the basis of observed institutional contractual arrangements and ease in interpreting the empirical results. In order to increase the comparability between the accounting and market measures of performance, we express both variables in terms of rates of return. The security market return (hereafter, *RET*), defined as the sum of the firm's capital gains and dividends divided by the stock price at the beginning of the year, has a natural interpretation as a measure of the firm's performance from the perspective of shareholders. We have chosen Return on Equity (hereafter, *ROE*), which is the firm's earnings before extraordinary items and discontinued operations divided by the average common shareholders' equity, as a "comparable" measure of accounting performance. This measure frequently appears as an explicit performance measure in bonus contracts disclosed in proxy statements.

The discussion in section 2.2 suggested that the relevant measure for each performance variable is its "surprise" component. Since the stochastic process describing the *ROE* performance variable exhibits positive autocorrelation (Foster [1986]), we use the *change in ROE* as our empirical measure of the surprise component.¹² Because the *RET* performance measure is generally considered to be uncorrelated over time, we use its level as our empirical market performance measure.¹³

¹² The evidence in Freeman, Ohlson, and Penman [1982] and Foster [1986] indicates that although the *ROE* variable exhibits positive autocorrelation, it is a stationary process. Since the use of the first difference in *ROE* implicitly assumes that *ROE* follows a random walk, we also conducted the analysis with an alternative specification of the *ROE* variable. On the basis of the evidence provided in Foster [1986, table 7.11], we used $ROE_t - .5ROE_{t-1}$ as the "surprise" in the accounting performance variable in period t . The time-series results (e.g., the average explanatory power of the regressions and the average magnitude of the t -statistics), which are not reported, were similar to the results reported in table 1. Although the statistical fit of the cross-sectional analysis (not reported) was generally weaker than the results reported in the text, the overall results were substantively similar.

¹³ Ideally, the ex ante expected value of *RET* would be subtracted from the realization of *RET* in order to derive the surprise component of *RET*. In principle, the expected rate of return on the firm's stock price could be estimated by employing an asset pricing model. For example, under the Capital Asset Pricing Model, the expected return on a firm's stock price is a function of the risk-free rate, the firm's beta, and the expected return on the market portfolio. Since estimating the *expected* return on the market portfolio is beyond

4.2 MEASUREMENT OF COMPENSATION

We use *cash compensation* (i.e., salary plus annual bonus) as our measure of executive compensation. Although cash compensation does not value the executive's entire remuneration package, salary plus bonus is almost always a significant portion of total compensation (salary plus bonus, long-term bonuses, perquisites, pensions, grants of stock, and stock options).¹⁴ For example, Benston [1985] and surveys by Booz, Allen, and Hamilton [1983] and Hay Associates [1981] report that salary plus bonus represents between 80% and 90% of total compensation.¹⁵

Consistent with most studies of compensation, we exclude changes in the value of the manager's holdings of stock and stock options from the measure of compensation.¹⁶ We recognize, however, that the choice between accounting and security market measures of performance in the salary and bonus portion of compensation may depend on the structure of the remainder of the manager's wealth. Therefore (as discussed in more detail below), we examine whether the relative weight placed on market versus accounting performance in the compensation contract is related to the extent to which the manager's other wealth is tied to the firm's stock price.

the scope of our analysis, we do not estimate the expected value of *RET* for each period. Instead, our analysis implicitly assumes that the expected value of *RET* is constant over time. Since our empirical analysis uses *real* rates of return, this assumption may not be too unrealistic. If the expected value for *RET* is constant over time, it will simply appear in the intercept of the regression.

¹⁴ See Antle and Smith [1986] and Lambert, Larcker, and Verrecchia [1987] for a discussion of the issues involved in valuing various components of compensation.

¹⁵ This discussion focuses on the level of compensation. For our purposes, the more relevant issue is whether the inclusion of other components of compensation has a significant impact on the *slope coefficients* relating compensation to performance. Jensen and Murphy [1987] provide some evidence on this issue with respect to the slope coefficient on the security market measure of performance. In particular, they find that the slope coefficient that relates salary plus bonus to changes in shareholder wealth is not significantly different from the slope coefficient that relates "total compensation" to changes in shareholder wealth. To our knowledge, there is no evidence regarding whether the slope coefficient on the accounting performance variable would be significantly affected by the inclusion of other components of compensation.

¹⁶ It is unclear whether these gains and losses should be viewed as a component of the manager's compensation (narrowly defined) or as income from personal investments. The value of these components of the manager's wealth are, by definition, directly tied to changes in the firm's stock price. For this reason, our slope coefficients will understate the magnitude of the relation between the manager's *total wealth* and his firm's stock price performance. We attempt to control for any resulting misspecification of our model by including a measure of the magnitude of the manager's stockholdings as an independent variable in the cross-sectional analysis. Consistent with most studies (see Lambert and Larcker [forthcoming] and Jensen and Murphy [1987] for exceptions), we also ignore any future-period effects on cash compensation that arise as a result of current-period performance.

4.3 BOX-COX ESTIMATION

In estimating the relation between compensation and performance, prior research has generally used either compensation or the natural logarithm of compensation as the dependent variable. These specifications are generally made without any guidance from economic theory.¹⁷ In contrast, our analytical agency model, expressed in equation (10), suggests that the relation between compensation and the accounting and market measures of performance can be expressed as:¹⁸

$$c_t^\kappa - c_{t-1}^\kappa = \beta_0 + \beta_{ROE}(ROE_t - ROE_{t-1}) + \beta_{RET}RET_t + e_t \quad (11)$$

where e_t represents a disturbance term.¹⁹ The dependent variable is the change in cash compensation raised to the power κ , where κ is the agent's coefficient or proportional risk aversion.

Since the value of κ is unknown (to the researcher), the parameters of equation (11) cannot be directly estimated using techniques such as multiple regression. However, Box and Cox [1964] have developed a procedure for estimating regression equations in which some or all of the variables are transformed by raising them to a power which is a parameter that must also be estimated. In particular, they show that the parameters κ , β_0 , β_{ROE} , and β_{RET} can be estimated via a maximum likelihood approach which utilizes iterative *OLS* procedures.²⁰

4.4 RESULTS OF BOX-COX AND MULTIPLE REGRESSION ANALYSIS

The parameters of equation (11) were estimated separately for each firm, and summary statistics for these parameter estimates are presented in table 1. The mean (median) *R*-squared of the estimation equations is

¹⁷ The choice of the logarithmic transformation is usually defended on unspecified "statistical grounds." In our model, the use of a logarithmic transformation of the dependent variable is not consistent with any member of the power utility functions.

An exception is Masson [1971] who uses compensation (more precisely, the percentage change in compensation) raised to the $\frac{2}{3}$ power. Interestingly, Masson states that this procedure is designed to reflect the decreasing marginal utility of compensation in the trade-off between income and leisure (i.e., effort).

¹⁸ Cash compensation and *RET* are expressed in 1967 dollars using the Consumer Price Index (*CPI*). The accounting return (*ROE*) is expressed in nominal terms because of our inability to estimate various asset layers and specific asset price indexes.

¹⁹ The disturbance term arises from other performance measures that may be used by shareholders in compensating the manager but which are not incorporated into our empirical model.

²⁰ More specifically, for a given value of κ , the maximum likelihood parameters estimates for β_0 , β_{ROE} , and β_{RET} can be derived by conducting a multiple regression of the transformed dependent variable on the independent variables. Following Box and Cox [1964], the dependent variable is deflated by the factor $\kappa(\bar{c})^{\kappa-1}$, where \bar{c} is the geometric mean of the compensation variable. Deflating the dependent variable by this factor leaves the magnitudes of the regression coefficients β_{ROA} and β_{RET} relatively unaffected by changes in the parameter κ and makes it easier to calculate the likelihood function.

TABLE 1

Summary Statistics for the Box-Cox Estimation of the Time-Series Relationship Between Compensation (c) and the Change in Return on Equity (ROE) and the Level of Security Market Return (RET) as Expressed in Equation (11):^a

$$c^*_t - c^*_{t-1} = \beta_0 + \beta_{ROE}(ROE_t - ROE_{t-1}) + \beta_{RET}RET_t + e_t$$

	MEAN	MEDIAN	TRMEAN ^c	STDEV	Q1	Q3
κ^b	0.784	0.400	0.612	1.800	-0.900	1.800
κLO	-0.514	-1.000	-0.642	0.876	-1.000	-0.238
κUP	3.05	2.50	2.86	2.50	1.25	4.10
$B O$	2.90	2.58	2.85	5.58	-0.16	5.77
$T BO$	0.494	0.363	0.476	0.874	-0.018	0.953
$B ROE$	288.2	186.4	256.9	461.9	10.6	499.1
$T ROE$	1.283	0.939	1.156	1.955	0.069	2.044
$B RET$	9.54	6.78	8.06	34.08	-7.24	21.89
$T RET$	0.368	0.337	0.364	1.135	-0.285	1.037
$RSQD$	0.272	0.205	0.255	0.228	0.083	0.399
DW	2.168	2.164	2.172	0.537	1.782	2.563
AR	-0.040	-0.038	-0.042	0.281	-0.247	0.162

^a The parameters of equation (11) were estimated separately for each firm using 14 time-series observations. Compensation data is expressed in thousands of dollars. The table presents summary statistics for the 370 estimations.

^b κ is the power in the BOX-COX transformation (the search for κ was restricted to the interval from -1.0 to 10.0 in increments of .05.), κLO (κUP) is the lower (upper) bound for the 95% confidence interval for κ , $B O$ ($T BO$) is the estimated intercept (t -statistic), $B ROE$ ($T ROE$) is the estimated slope (t -statistic) on the ROE variable, $B RET$ ($T RET$) is the estimated slope (t -statistic) on the RET variable, $RSQD$ is the R -squared of the estimation, DW is the Durbin-Watson statistic, and AR is the estimated first-order autocorrelation in the residuals.

^c $TRMEAN$ trims the smallest 5% and the largest 5% of the observations and averages the remaining observations. $Q3$ is the third quartile and $Q1$ is the first quartile.

0.272 (0.205). The mean (median) t -statistic for the ROE slope is 1.283 (0.939), while the mean (median) t -statistic for the RET slope is only 0.368 (0.337). If we aggregate the t -statistics cross-sectionally under the assumption of cross-sectional independence, the corresponding z -statistics for the ROE slope and the RET slope are 22.70 and 6.51, respectively. Since the regressions are conducted over the same time period (i.e., 1970-84), it is unlikely that the slope coefficients are independent across firms; however, only 3 (33) independent observations out of the 370 firms are needed in order for the average t -statistic on the ROE (RET) slope to be significantly different from zero at the 0.05 level (two-tailed). These results suggest that cash compensation is statistically related to both RET and changes in ROE .

The estimates of the parameter κ in table 1 are imprecise; the mean length of the 95% confidence interval is 3.05.²¹ Nevertheless, it is instruc-

²¹ As discussed in Box and Cox [1964, pp. 214-15], the confidence intervals for κ were derived by calculating the range of κ 's which yielded values for the logarithm of the likelihood function that were within $.50\chi^2$ ($df = 1$) of the function's maximum value.

tive to compare our estimates of κ with the predictions of the theory and with the values of κ implicitly assumed in prior research. Agency theory implies that the parameter κ represents the agent's coefficient of proportional risk aversion, which should be positive. The mean (median) value of κ was 0.784 (0.400), and only 5.4% of the estimates of κ were negative and statistically significant at the 0.05 level.²² Many prior studies have analyzed compensation in first differences, which implicitly assumes a value for κ of one. Our results indicate that 74.3% of the confidence intervals for κ included one, which suggests some support for this specification.²³

Since the confidence interval for κ typically includes the value of one, we also estimated the relation between compensation and *RET* and *ROE* using standard *OLS* multiple regression. The slope coefficients, *t*-statistics, and explanatory power of these multiple regressions (not reported) are very similar to the results derived from the Box-Cox analysis. This similarity is probably due to the inability of the Box-Cox procedure to distinguish between alternative values of κ .²⁴

The results in table 1 suggest that cash compensation is more highly associated with differences in accounting returns than with levels of security market returns.²⁵ Some additional support for this statement

²² The average *t*-statistics on the performance variables and the *R*-squared of the estimation procedures for the subsample of firms for which the estimate for κ was positive (i.e., the cases that are most consistent with the theory) are virtually identical to the results reported in table 1.

²³ The inability to provide a precise estimate for κ is likely to be due to our small sample size (14 time-series observations per firm) and/or the limited range of compensation numbers. That is, the function c^{κ} can be well approximated by a linear function of c provided that the range of c is not "too large." To provide some evidence on this issue, we estimated a linear regression between the compensation numbers and the square of compensation for each firm. The mean (median) *R*-squared from these regressions was 0.982 (0.990). Similarly, a linear regression of compensation on the natural logarithm of compensation yielded a mean (median) *R*-squared of 0.970 (0.989). Finally, a regression of the square of compensation on the natural logarithm of compensation yielded a mean (median) *R*-squared of 0.914 (0.959). These results suggest that it would be difficult to determine whether the "best" specification of the dependent variable is the logarithm of compensation ($\kappa = 0$), compensation ($\kappa = 1$), or the square of compensation ($\kappa = 2$), which is consistent with the results in table 1. The finding that a linear specification of the compensation variable works "well" for most firms does not imply, however, that cross-sectional differences in the risk aversion coefficients of managers are unimportant.

²⁴ As discussed in n. 20, the dependent variable in the Box-Cox estimation was scaled so that the magnitude of the slope coefficients is relatively unaffected by the choice of κ . This will result in the magnitude of the Box-Cox slope coefficient being "close" to the magnitude of the multiple regression slope coefficients (in which κ is exogenously set equal to one).

²⁵ The magnitude of the coefficient on stock market returns is roughly consistent with the results of other studies. For example, Jensen and Murphy [1987] estimate the slope coefficient from a pooled time-series cross-sectional regression of changes in the real salary plus bonus paid to *CEOs* on the real change in the value of their firm's shareholder wealth to be 0.0000094. To transform this number into a coefficient for a regression in which real stock price *return* is the independent variable, we multiply the coefficient estimated in

can be obtained by examining the results of *univariate* regressions of the change in compensation on the change in *ROE* and the change in compensation on *RET*. In these regressions (not reported), the mean *t*-statistic on *ROE* was 1.18, compared to 0.548 for *RET*. Moreover, the average *R*-squared for the regressions involving *ROE* was 0.186, compared to 0.087 for *RET*. Therefore, the average *incremental R*-squared for the *ROE* performance variable was 0.167, whereas the average *incremental R*-squared for *RET* was only 0.068.

5. Hypotheses on the Relative Weights Assigned to the Performance Measures

We discuss two different approaches to testing hypotheses concerning the relative weights assigned to *RET* and *ROE* in compensation contracts. First, we discuss a methodology that tests the agency model prediction that, for a given firm, the slope coefficients on the performance measures scaled by their signal-to-noise ratios should be equal. Second, we discuss a methodology which is designed to address the problems associated with obtaining precise measures of the signal-to-noise ratios of *RET* and *ROE* (the statistical aspects of this approach are presented in section 6 and Appendixes A and B). Finally, we develop specific research hypotheses for empirical examination.

5.1 "EQUALITY OF COEFFICIENTS" APPROACH

Equation (7) implies that if the performance measures *ROE* and *RET* are multiplied by their respective signal-to-noise ratios, then the slope coefficients on the "rescaled" performance measures will be equal. This hypothesis can be tested by conducting a time-series estimation of changes in compensation on the rescaled performance measures, and then testing for equality of the slope coefficients. If it were possible to measure the signal-to-noise ratios for each firm without error, this "equality of coefficients" test procedure would allow for a powerful test of the agency theory relationship developed in section 2.

Although the agency theory model suggests that the "noise" in a performance measure is related to its variance, the theory provides no

Jensen and Murphy by the average market value of the firms in our sample, which is \$876,000,000 in real \$1967, and then divide by 1000.0 to adjust for the fact that we measure compensation in thousands of dollars. The resulting coefficient is 8.23, which is consistent with the mean (median) coefficient of 9.54 (6.78) reported in table 1. Of course, these two sets of results are not strictly comparable because the results in table 1 are based on a *multiple* regression in which both accounting and market measures of performance are used as independent variables, whereas Jensen and Murphy use only a market measure of performance in their regression. In order to achieve a more comparable set of results, we also conducted a simple regression in which the market measure of performance was the only independent variable. The mean slope coefficient from these regressions was 16.67, which is still roughly the same size as the coefficient reported in Jensen and Murphy.

insights into how to measure the sensitivity of the performance measures to the agent's actions. If we make some assumptions about the signal-to-noise ratios of the performance measures, we can use the "equality of coefficients" test procedure to test the *joint* hypothesis that the agency theory model and the assumption about the signal-to-noise ratios are both correct. For example, if the signal-to-noise ratios of *RET* and *ROE* are hypothesized to be equal, the theory predicts that the slope coefficients on the *RET* and *ROE* performance measures will be equal. A crude test of the equality of slope coefficients can be developed by determining whether the *ROE* slope coefficients are statistically different (in central tendency) from the *RET* slope coefficients for our sample of 370 firms. For the Box-Cox estimates, the *ROE* slope coefficient is substantially larger than the *RET* slope coefficient (matched-pairs $t = 11.52$ and Wilcoxon $z = 11.19$).

An alternative hypothesis is that the "noise" of a signal is measured by its time-series variance, and that the sensitivities of the accounting and market measures of performance are equal. After deflating the performance variables by their variances, our results fail to reject the hypothesis that the Box-Cox slope coefficients on the "rescaled" performance variables are (on average) equal (matched-pairs $t = 0.42$ and Wilcoxon $z = -0.45$).

We do not rely on these "equality of coefficients" tests because they are confounded if the signal-to-noise ratios of the two performance measures are measured with error. The resulting estimates for the regression coefficients on the "rescaled variables" would be inconsistent, and the test for equality of coefficients would not produce valid inferences. Specifically, we do not know if the slope coefficients are equal because (i) the agency model is actually a good description of contracting, or (ii) there is so much measurement error in the estimates of the signal-to-noise variables that this test has no power to reject the research hypothesis, or (iii) the coefficient on accounting performance is higher for one group of firms but lower for another group of firms, and the absence of any cross-sectional analysis obscures this fact.

5.2 MOTIVATION FOR CONDUCTING A LATENT VARIABLE CROSS-SECTIONAL ANALYSIS

The problems associated with estimating the signal-to-noise ratios of the performance measures suggest that it is desirable to employ a methodology that explicitly considers measurement error. As developed in equation (8), the agency theory model predicts that the ratio of the regression coefficients on the (unscaled) performance measures provides an estimate of the ratio of the signal-to-noise ratios of the performance measures. Our statistical procedures are designed to assess whether the ratio of these estimated regression coefficients is associated with variables that are hypothesized to have some relationship to the sensitivity and/or noise of the performance variables.

The desirability of this approach is influenced by the degree to which measurement error can be “controlled” in the variables that proxy for the sensitivity and noise of the performance measures. In the remainder of this section, we discuss the variables used to proxy for the relative weights assigned to the *RET* and *ROE* performance measures in the compensation contract, and to proxy for the sensitivity and noise of *RET* and *ROE* with respect to evaluating a manager’s performance. In section 6, we use a latent variable analysis to examine whether cross-sectional differences in the relative weights assigned to *RET* versus *ROE* are related to the variables that proxy for the sensitivity and noise of these performance measures. This methodological approach enables us to develop tests of our research hypothesis that “control” measurement error in a manner not available in the “quality of coefficients” test procedure.

5.2.1. Ratio of the Slope Coefficients. Our dependent variable, denoted *PERF MIX*, is the degree to which cash compensation is tied to *RET* relative to *ROE*. As discussed in section 2, we use the ratio of the slope coefficients on the two performance variables to control for other factors that affect the relation between compensation and performance for a particular firm.²⁶ One important disadvantage of this approach is that the ratio must be calculated from *estimates* of the slope coefficients, and the distribution of the ratio of two random variables can be complicated and possess undesirable properties.

The statistical distribution of the ratio of the slope coefficients is especially sensitive to either measurement error or a near-zero value for the denominator. To reduce these problems, we place the slope coefficient with the more precise (and more positive) estimate in the denominator

²⁶ One problem in interpreting empirical results concerning the ratio of the slope coefficients on the performance variables (β_x/β_y) is that this ratio is sensitive to the *scaling* of the performance variables x and y . For example, if variable x tends to be, on average, ten times as large as variable y in a given firm, then the slope coefficient on x will, ceteris paribus, be approximately $1/10$ as large as the coefficient on y . Clearly, this does not imply that variable y is a more “important” performance measure in contracting than variable x . This problem is exacerbated if the scaling differences between x and y vary cross-sectionally. For example, assume that for another firm, the variable x is only five times as large as variable y . In this situation, the ratio of the slope coefficient on x to y would be approximately $1/5$. More important, it would be inappropriate to conclude that the “informativeness” of variable y relative to variable x is twice as large for the first firm as for the second firm.

A common (statistical) solution to this type of interpretation problem is to deflate the performance measures by their standard deviations. While this will eliminate scaling problems, it also makes it difficult to test hypotheses about the amount of noise in the performance variables. That is, the transformed measures of performance will have, by construction, a variance of one, which tends to eliminate any cross-sectional variation in the relative amounts of noise in the performance measures x and y .

We repeated the analysis using *standardized* performance measures in the time-series regressions. The cross-sectional analysis then related the ratio of the slope coefficients to the *CORR*, *GROWTH*, and *OTHER* constructs. The results, not reported, were similar to the results reported in the text.

of the ratio.²⁷ As the results in table 1 indicate, the slope coefficient on *ROE* was much more likely to be positive and statistically significant than the coefficient on *RET*. For this reason, the *PERF MIX* construct is defined to be the ratio of the slope coefficient on *RET* to the slope coefficient on *ROE*.

Another problem with the ratio of the slope coefficients arises because the estimated slope coefficients on *RET* and *ROE* can be either positive or negative. As a result, this ratio can be negative if only the numerator is negative or if only the denominator is negative. This complicates the cross-sectional analysis because these outcomes have quite different interpretations for our theoretical analysis. We consider three different means of dealing with this problem, summarized in panel A of table 2. The first approach is to consider only firms with positive slope coefficients on both *ROE* and *RET*. This proxy for the *PERF MIX* variable is denoted *MR-PP* (*BC-PP*) if the slope coefficients were estimated using multiple regression (Box-Cox). A second approach is to allow the numerator of the ratio (i.e., the slope on the *RET* variable) to have any sign but only consider firms with a positive slope coefficient on *ROE*. This proxy is denoted *MR-AP* (*BC-AP*) if the slope coefficients are estimated via multiple regression (Box-Cox). Finally, we compute the ratio after reestimating the slope coefficients on *RET* and *ROE* with the constraint that each slope coefficient and the coefficient of proportional risk aversion be positive.²⁸ This proxy is denoted *MR-CC* (*BC-CC*) if the slope coefficients are estimated via multiple regression (Box-Cox).

In the subsequent discussion, analysis using *MR-PP* and/or *BC-PP* as the dependent variable is referred to as the "Positive/Positive" case. Similarly, analysis using *MR-AP* and/or *BC-AP* as the dependent variable is referred to as the "Anything/Positive" case. Finally, analysis using the *MR-CC* and/or *BC-CC* variables is referred to as the "Constrained/Constrained" case. The Positive/Positive and Constrained/Constrained cases assume that, a priori, the slope coefficients on both performance measures should be positive. The Anything/Positive case assumes that, a priori, the coefficient on the *RET* variable could be negative, perhaps to "undo" some of the manager's exposure to risk in his other wealth.

5.2.2. *Noise in the Performance Measures.* The noise in a signal re-

²⁷ The results in table 1 indicate that the precision of the estimates for the slope coefficients varies considerably across firms. This suggests that the ratio of the estimated slope coefficients will possess heteroscedasticity. If the dependent variable in the cross-sectional analysis consisted simply of a single slope coefficient, the heteroscedasticity problem could be solved by deflating the slope coefficients (and the independent variables of the cross-sectional analysis) by the standard error of the slope coefficient estimates. See Saxonhouse [1976] for additional discussion of this point. This approach is not pursued because our dependent variable is the ratio of the slope coefficients, and it is not clear how to compute the standard error of the ratio from the individual standard errors of the slope coefficients.

²⁸ The slope coefficients were constrained to lie within the interval from .01 to 5000, and the value of κ was constrained to lie in the interval from .01 to 10.

TABLE 2

Description of the Variables Used in the Cross-Sectional Analysis of the Relative Weights Placed on Performance Measures in Compensation Contracts

A. Proxies for the Relative Weight Placed on <i>RET</i> Versus <i>ROE</i> (<i>PERF MIX</i>)	
Variable	Description
<i>MR-PP</i>	The ratio of the slope coefficient on the <i>RET</i> variable to the slope coefficient on the <i>ROE</i> variable. The slope coefficients were estimated via Multiple Regression with the slopes unconstrained. The ratio is used only if both slope coefficient are Positive. The ratio is coded as "missing" if either slope coefficient is negative.
<i>BC-PP</i>	Same as <i>MR-PP</i> , except that the slope coefficients are estimated with the Box-Cox procedure.
<i>MR-AP</i>	Same as <i>MR-PP</i> , except that the slope coefficient on <i>RET</i> may take on Any value. That is, the ratio is coded as "missing" if the slope coefficient on the <i>ROE</i> variable is negative.
<i>BC-AP</i>	Same as <i>MR-AP</i> , except that the slope coefficients are estimated with the Box-Cox procedure.
<i>MR-CC</i>	Same as <i>MR-PP</i> , except that the slope coefficients are estimated with the coefficients Constrained to be positive.
<i>BC-CC</i>	Same as <i>MR-CC</i> , except that the slope coefficients are estimated with the Box-Cox procedure.
B. Proxies for the Noise in <i>ROE</i> Versus <i>RET</i> (<i>NOISE</i>)	
Variable	Description
<i>RTVAR</i>	Ratio of the Total Variance (over the period 1970 to 1984) of the change in <i>ROE</i> to the total variance of <i>RET</i> .
<i>RSVAR</i>	Ratio of the Systematic Component of the Variance of the change in <i>ROE</i> to the systematic component of the variance of <i>RET</i> .
<i>ZTRANS</i>	Fisher transformation of the coefficient of correlation for the regression of the real stock price equity return on first differences in the return on assets.
C. Proxies for the Firm's Growth (<i>GROWTH</i>)	
Variable	Description
<i>A GROW</i>	One plus the Real growth in total assets from 1970 to 1984.
<i>S GROW</i>	One plus the Real growth in sales from 1970 to 1984.

TABLE 2—continued

Variable	Description
PER-CENT	Average percentage of the firm owned by the <i>CEO</i> . The average is computed using years 1970 to 1973 and 1980.
EQ/COMP	Average ratio of the market value of equity owned by the <i>CEO</i> to his cash compensation. The average is computed using years 1970 to 1973 and 1980.
STOCK	Average level for the market value of equity for the <i>CEO</i> expressed in thousands of 1967 dollars. The average is computed using years 1970 to 1973 and 1980.

flects the degree to which it is influenced by factors other than the manager's actions. Equation (8) indicates that the appropriate measure of the amount of noise in a signal is the variance of the signal *given the agent's action*. Holding the (conditional) sensitivity of the mean of the signal to the agent's effort constant for both signals, agency theory implies a *positive* relationship between the ratio of the slope on *RET* to the slope on *ROE* and the *inverse* of the ratio of the "noise" in *RET* to the "noise" in *ROE*.

We consider three proxies for the relative amounts of *NOISE* in the two performance measures. The first measure, denoted *RTVAR*, is the ratio of the *time-series* variation in the change in *ROE* to the *time-series* variation in *RET*.²⁹ One limitation of this measure is that the *time-series*

²⁹ It might seem that the ratio of the slope coefficients is related to the inverse of the ratio of the variances of the performance variables by construction. For example, consider the case in which a dependent variable, z , is related to two independent variables, x and y , which are measured without error, in the following fashion:

$$z = \beta_0 + \beta_x x + \beta_y y + e, \quad \text{with} \quad \text{cov}(x, e) = \text{cov}(y, e) = 0.$$

If, for convenience, we assume that x and y are independent, it is well known that the ratio of the slope coefficients on the two independent variables can be written as:

$$\frac{b_x}{b_y} = \frac{\text{cov}(z, x)}{\text{var}(x)} \frac{\text{var}(y)}{\text{cov}(z, y)} = \frac{\text{var}(y)}{\text{var}(x)} \frac{\text{cov}(z, x)}{\text{cov}(z, y)}.$$

This expression makes it seem as if the hypothesized relation between the ratio of the slope coefficients and the ratio of the variances is true by construction. However, if we expand the expressions for $\text{cov}(z, x)$ and $\text{cov}(z, y)$, we have:

$$\begin{aligned} \frac{b_x}{b_y} &= \left[\frac{\text{cov}(\beta_0 + \beta_x x + \beta_y y + e, x)}{\text{var}(x)} \right] \div \left[\frac{\text{cov}(\beta_0 + \beta_x x + \beta_y y + e, y)}{\text{var}(y)} \right] \\ &= \left[\beta_x + \frac{\beta_y \text{cov}(x, y) + \text{cov}(x, e)}{\text{var}(x)} \right] \div \left[\beta_y + \frac{\beta_x \text{cov}(x, y) + \text{cov}(y, e)}{\text{var}(y)} \right]. \end{aligned}$$

variation in a signal will not only reflect the amount of "noise" in the signal but also the effect of changes in the agent's action over time. As the results in table 3 indicate, the mean (median) ratio of the variance of *ROE* to the variance of *RET* was 0.051 (0.012).

A second measure for the amount of noise in a signal can be obtained by decomposing the time-series variation in a signal into a market (or industry) component and a firm-specific component. If the market component of the performance measure is thought to be unrelated to the manager's actions, as in the relative performance evaluation literature (see Holmstrom [1982]), then the variance of the *market component* of the performance evaluation measure is a more appropriate measure of the amount of noise in a signal.³⁰

The second measure, denoted *RSVAR*, is the ratio of the variance of the systematic component of *ROE* to the variance of the systematic component of *RET*. This ratio was computed by first conducting time-series regressions of the change in *ROE* for the firm on the change in *ROE* for a value-weighted market index and the *RET* for the firm on the value-weighted market *RET*. Descriptive statistics for these "market-model" regressions are presented in table 4 (panels A and B). The market component of the change in *ROE* accounts for (on average) approximately 18% of the total variation in the change in *ROE*, whereas the market accounts for (on average) approximately 37% of the total variation in *RET*. For each firm, the results of these regressions were used to calculate the variance of the systematic component of the change in *ROE* and the variance of the systematic component of *RET*. As the results in table 3 indicate, the mean (median) value of *RSVAR* was 0.064 (0.004).

The final proxy for the *NOISE* construct is derived from discussions in the security price literature which link the magnitude of the correlation between market prices and accounting numbers with the amount of noise in accounting numbers. In fact, Salamon and Smith [1979] have suggested that the correlation between accounting numbers and market prices provides an explicit measure of the amount of "managerial misrep-

Since $\text{cov}(x, y) = \text{cov}(x, e) = \text{cov}(y, e) = 0$, it is easy to see that the apparent dependence of the ratio of the slope coefficients on the ratio of the variances disappears. While these covariances will not be exactly equal to zero in a finite sample, they are equally likely to be positive or negative, so there will be no systematic dependence between the size of the ratio of the slope coefficients and the size of the ratio of the variances.

³⁰ It is not clear that the "market" component of a performance measure constitutes noise when the executive is, in part, responsible for deciding the type of projects (and the industry) in which the firm invests. For example, Dye [1987] shows that a "pure" relative performance evaluation scheme can lead to suboptimal project selections.

Antle and Smith [1986] provide empirical evidence consistent with the hypothesis that *total* executive compensation (including the change in the value of shares of stock and stock options held by the executive) is more sensitive to the firm-specific component of performance than the industry component of performance. However, they do not provide any evidence regarding the use of relative performance evaluation in the salary plus bonus portion of compensation.

TABLE 3
Descriptive Statistics on the Variables Used in the Cross-Sectional Analysis of the Relative Weights Placed on Performance Measures in Compensation Contracts^a

	<i>N</i> ^b	MEAN	MEDIAN	TRMEAN	STDEV	Q1	Q3
<i>MR-PP</i>	166	0.172	0.062	0.104	0.508	0.023	0.170
<i>BC-PP</i>	161	0.384	0.076	0.123	1.782	0.023	0.184
<i>MR-AP</i>	277	0.041	0.012	0.028	0.508	-0.034	0.085
<i>BC-AP</i>	283	-0.032	0.011	0.035	3.851	-0.036	0.089
<i>MR-CC</i>	370	374.8	0.059	146.8	1146.3	0.0002	1.0
<i>BC-CC</i>	370	309.5	0.057	102.8	1034.5	0.0001	0.7320
<i>RTVAR</i>	370	0.051	0.012	0.024	0.180	0.005	0.031
<i>RSVAR</i>	370	0.064	0.004	0.015	0.341	0.001	0.022
<i>ZTRANS</i>	370	0.870	0.838	0.878	1.092	0.246	1.611
<i>A GROW</i>	370	2.184	1.878	1.996	1.852	1.198	2.609
<i>S GROW</i>	370	1.706	1.440	1.567	1.153	1.054	1.994
<i>PER-CENT</i>	370	1.851	0.145	0.832	5.242	0.039	0.450
<i>EQ/COMP</i>	370	82.8	4.4	25.6	292.4	1.4	12.7
<i>STOCK</i>	370	10781	742	3725	36605	197	2390

^a See table 2 for a definition of the variables.

^b *N* is the number of firms with available data. *TRMEAN* trims the smallest 5% and the largest 5% of the observations and averages the remaining observations. *Q3* is the third quartile and *Q1* is the first quartile.

resentation" in accounting numbers.³¹ However, this correlation measures noise only from the perspective of assessing the *value of the firm*. As discussed by Gjesdal [1981], an information system that is useful in valuing the firm need not be useful in assessing a manager's performance. Therefore, rankings of information systems in terms of their value in assessing (i) a manager's performance and (ii) the value of the firm need not be the same.

The correlation between the *RET* and *ROE* performance measures, which we denote *ZTRANS*, was obtained by conducting a regression of the firm's *RET* on its change in *ROE*. The results of these regressions are summarized in table 4, panel C. For example, the mean (median) *R*-squared for these regressions was 0.136 (0.084). In order to convert the correlation coefficient from this regression into a more normally distributed random variable, we defined *ZTRANS* to be the Fisher transformation of the correlation coefficient.

5.2.3. Sensitivity of the Performance Measures to the Agent's Actions. While security market prices are thought to estimate the expected present value of all future consequences of a manager's actions, accounting numbers are frequently criticized for their inability to reflect future-

³¹ The accounting literature is replete with articles on earnings manipulation (see Foster [1986] and Verrecchia [1986]).

TABLE 4

Summary Statistics for the Results of Regression Analyses of the Systematic Component of the Change in Return on Assets (ROE), the Systematic Component of Security Market Return (RET), and the Correlation Between the RET and the Change in ROE for a Sample of 370 Firms^a

A. Time-Series Regression of Change in ROE on the Change in Market ROE						
	MEAN ^b	MEDIAN	TRMEAN	STDEV	Q1	Q3
INT ^a	-0.002	0.000	-0.000	0.013	-0.004	0.003
T INT	0.055	0.058	0.038	0.818	-0.432	0.427
SLOPE	0.937	0.482	0.729	2.410	-0.014	1.541
T SLOPE	1.068	0.945	1.061	1.640	-0.042	2.220
RSQD	0.185	0.118	0.171	0.185	0.023	0.312
DW	2.023	2.010	2.030	0.551	1.636	2.456
AR	0.009	0.016	0.006	0.281	-0.212	0.207
B. Time-Series Regression of RET on the Market RET						
	MEAN	MEDIAN	TRMEAN	STDEV	Q1	Q3
INT	0.024	0.019	0.022	0.049	-0.006	0.050
T INT	0.359	0.393	0.365	0.722	-0.117	0.826
SLOPE	0.999	0.949	0.973	0.459	0.652	1.268
T SLOPE	2.890	2.790	2.840	1.220	2.052	3.525
RSQD	0.373	0.375	0.372	0.174	0.245	0.489
DW	1.845	1.829	1.845	0.443	1.534	2.141
AR	0.097	0.105	0.097	0.226	-0.054	0.255
C. Time-Series Regression of RET on Change in ROE						
	MEAN	MEDIAN	TRMEAN	STDEV	Q1	Q3
INT	0.060	0.052	0.056	0.056	0.024	0.084
T INT	0.734	0.724	0.727	0.577	0.353	1.083
SLOPE	2.351	1.810	2.321	4.332	0.438	4.114
T SLOPE	0.970	0.885	0.958	1.260	0.257	1.750
RSQD	0.136	0.084	0.122	0.146	0.022	0.218
DW	1.892	1.875	1.899	0.418	1.584	2.183
AR	0.076	0.084	0.077	0.214	-0.073	0.233

^a INT (T INT) is the estimated intercept (*t*-statistic) in the regression, SLOPE (T SLOPE) is the estimated slope (*t*-statistic), RSQD is the coefficient of determination for the regression, DW is the Durbin-Watson statistic, and AR is the estimated first-order autocorrelation of residuals for the regression.

^b TRMEAN trims the smallest 5% and the largest 5% of the observations and averages the remaining observations. Q3 is the third quartile and Q1 is the first quartile.

period consequences of current-period actions (e.g., Rappaport [1986]). While it is possible, in principle, for historical-cost-based accounting rates of return to reflect the present value of the firm's future cash flows by using "economic" depreciation (see the discussion in Beaver [1981]), it is unclear whether conventional accounting practices achieve this result. In particular, constraints imposed by Generally Accepted Accounting Principles may limit the ability of accounting numbers to reflect the cash flows that are expected to arise in the future as a result of the firm's current-period actions.

If accounting-based performance measures are not as sensitive as market-based performance measures to the actions of the manager that have future-period consequences, agency theory predicts that, *ceteris paribus*, compensation schemes will be less accounting oriented in situations in which firms are in the “early” periods (before the effects of the manager’s investments are reflected in accounting numbers), and more sensitive to accounting numbers in “later” periods (when the effects of the manager’s investments are reflected in accounting earnings).³² This hypothesis therefore predicts that the ratio of the *RET* slope to the *ROE* slope is *positively* associated with the extent to which the firm is in the “early” versus “later” stages of investment *during the same period* in which the slopes of the compensation function are estimated.

We assume that the extent to which a firm is in the early stages of investment is related to the *GROWTH* of the firm, as measured by real growth in total assets (denoted *A GROW*) and real growth in sales (*S GROW*). These growth rates were computed over the period 1970–84, the same period over which the slope coefficients in the compensation function were estimated. We assume that firms in the “early” stages of investment will be experiencing both increases in their asset bases and increasing (as opposed to constant or declining) sales.

5.2.4. Other Components of Managerial Wealth. If the other components of a manager’s wealth are highly correlated with the firm’s stock price, there may be little incentive benefit to also linking the manager’s cash compensation to stock price. Such a link may merely increase the manager’s exposure to risk. For this reason, if the manager already has a considerable amount of his wealth tied to the firm’s stock price, shareholders may desire to use other informative performance measures in the manager’s compensation contract. To test this hypothesis, we examine whether the relative weight placed on *RET* and *ROE* is *negatively* associated with the degree to which the manager’s other wealth is tied to the firm’s stock price.

We consider three proxies for this construct, which is denoted *OTHER*. The first indicator, denoted *STOCK*, is the average dollar value of stock (in 1967 dollars) owned by the *CEO* or his family. This measure assumes that the manager’s “other incentives” are proportional to the size of his equity investment in the firm. For example, the change in the manager’s wealth for a given value of *RET* is proportional to the manager’s equity investment. The second indicator, denoted *EQ/COMP*, is the average market value of equity owned by the *CEO* or his family divided by his cash compensation. This measure assumes that the importance of the manager’s equity investment is a function of the size of this investment relative to his cash compensation. The final indicator, denoted *PER-*

³² See Lambert [1981; 1986] and Ramanan [1986] for agency theory models which support this result.

CENT, is the percentage of the equity held by the *CEO* or his family. This is a commonly discussed measure of the amount of incentive the manager has to increase the firm's stock price.

For our sample of firms, the results in table 3 indicate that the mean (median) percentage of the firm's total stock held by the *CEO* and his family was 1.85% (.145%). Moreover, the mean (median) ratio of equity holdings to annual cash compensation was 82.8 (4.4). These numbers were computed by averaging the values for years 1970 to 1973 and 1980.

6. Cross-Sectional Analysis of Compensation Contracts

The preceding analysis suggests that the relative weight assigned to *RET* versus *ROE* in the cash compensation contract is:

- (i) an increasing function of the inverse of the noise ratios of the two performance measures,
- (ii) an increasing function of the degree to which the firm is in the "early" stages of investment, and
- (iii) a decreasing function of the extent to which the manager's other wealth is tied to stock price.

6.1 CORRELATION ANALYSIS

The results in table 3 indicate that the distributions of many of the variables described above are highly skewed (particularly *MR-CC*, *BC-CC*, *EQ/COMP*, and *STOCK*). In order to obtain more "normal" distributions, we applied a logarithmic transformation to each of the variables.³³ An additional justification for applying a logarithmic transformation is that it converts the multiplicative relation expressed in equation (8) into a more easily estimated additive relation. That is, equation (8) suggests that the ratio of the slope coefficients is related to the ratio of the noise measures *times* the ratio of the sensitivities of the two performance measures. Taking the logarithm of both sides of equation (8) results in an equation in which the logarithm of the ratio of the slope coefficients is related to the logarithm of the ratio of the noise measures *plus* the logarithm of the sensitivities of the performance measures.

The Pearson correlations (after applying the logarithmic transformation) among the proxies for the *PERF MIX*, *NOISE*, *GROWTH*, and

³³ Since the variables *MR-AP* and *BC-AP* contain negative values, these variables were not transformed. Moreover, since the variable *ZTRANS* had already been transformed into a normally distributed random variable, no adjustment was made to this variable either.

In addition, each of the distributions for the variables was winsorized (i.e., the "extreme" observations were reset to "less extreme" values—the extreme values were not deleted from the sample). No more than four data points per variable were affected by this procedure. The cutoff points were chosen on the basis of an examination of the univariate histogram for each variable.

OTHER constructs are presented in table 5.³⁴ These results indicate that the correlations are generally consistent with the hypotheses discussed above. The correlations between the proxies for the *PERF MIX* construct and the *NOISE* construct are generally positive (14 of 18), the correlations between the proxies for the *PERF MIX* and the *GROWTH* construct are also generally positive (9 of 12), and the correlations between the *PERF MIX* proxies and the *OTHER* proxies all have the hypothesized negative sign. Moreover, the results in table 5 also indicate that the proxies for a given theoretical construct are generally highly correlated with each other. The only exception occurs in the *NOISE* construct, where the variables *RTVAR* and *RSVAR* are highly correlated with each other but virtually uncorrelated with *ZTRANS*. This result suggests that the variable *ZTRANS* is not measuring the same underlying variable as *RTVAR* and *RSVAR*. Since the theoretical justification in section 5.2.2 for the variable *ZTRANS* is also considerably different from the justification for the *RTVAR* and *RSVAR* variables, our subsequent empirical tests will analyze the *ZTRANS* variable as a distinct construct, which we refer to as *CORR*.

Although the results in table 5 provide information concerning the relations among the *proxies* for the theoretical constructs, our real concern is with the relations among the underlying *constructs* themselves. For example, we are interested in the correlation between the constructs *PERF MIX* and *NOISE*, as opposed to the correlation between the proxies *MR-PP* and *RTVAR*. There are at least three factors which make it difficult to assess the statistical significance of the relation between the *constructs* on the basis of the correlations among the *proxies*. First, each proxy measures its underlying construct with error. The presence of this measurement error will, *ceteris paribus*, attenuate the bivariate correlation coefficients. As we discuss below, the existence of multiple proxies for a given construct can, in principle, be used to reduce this problem.

Second, it is difficult to "combine" the correlations among alternative indicators for the same proxy. For example, although the correlations among *all* of the proxies for the *PERF MIX* and the *OTHER* constructs have the predicted signs, the statistical significance of the individual correlations (not presented) varies considerably across proxies. In the case of the *PERF MIX* and the *GROWTH* constructs, this problem is even worse because the signs of the correlations are different for the individual proxies. Inconsistencies in the significance levels and signs of the bivariate correlations make it difficult to draw conclusions concerning the research hypotheses regarding the relations among the underlying constructs.

³⁴ We also computed the Spearman rank-order correlations, which should be less affected by the "nonnormality" of the data, for the untransformed variables. The results (not presented) were very similar to the Pearson correlations applied to the transformed variables presented in table 5.

TABLE 5
Matrix of Pearson Correlation Coefficients for the Variables Used in the Cross-Sectional Analysis of the Relative Weights Placed on Performance Measures in Compensation Contracts^a

	MR-PP	BC-PP	MR-AP	BC-AP	MR-CC	BC-CC	RTVAR	RSVAR	ZTRANS	A GROW	S GROW	PERCENT	EQ/COMP
BC-PP	0.882												
MR-AP	0.787	0.714											
BC-AP	0.681	0.831	0.837										
MR-CC	0.999	0.645	0.507	0.611									
BC-CC	0.795	0.859	0.549	0.705	0.928								
RTVAR	0.359	0.222	0.062	0.063	-0.023	0.023							
RSVAR	0.157	0.063	0.014	0.004	-0.031	-0.040	0.662						
ZTRANS	0.067	0.139	-0.048	0.111	0.204	0.176	0.029	-0.028					
A GROW	-0.001	0.139	0.038	0.080	0.113	0.117	-0.328	-0.320	0.035				
S GROW	-0.031	0.043	-0.009	0.013	0.017	0.043	-0.337	-0.328	-0.006	0.802			
PER-CENT	-0.292	-0.233	-0.086	-0.075	-0.078	-0.115	0.034	0.028	-0.067	-0.050	0.096		
EQ/COMP	-0.308	-0.265	-0.072	-0.086	-0.090	-0.110	-0.028	-0.013	-0.032	0.019	0.133	0.900	
STOCK	-0.299	-0.266	-0.064	-0.080	-0.113	-0.136	0.011	0.020	-0.043	-0.041	0.140	0.868	0.980

^a See table 2 for a definition of the variables. The sample size for the correlations is 370, except for comparisons involving MR-PP ($n = 166$), BC-PP ($n = 161$), MR-AP ($n = 277$), and BC-AP ($n = 283$).

Finally, the individual elements of the correlation matrix are *simple* correlations not *partial* correlations. Although this problem could be solved by conducting a multiple regression of each proxy for *PERF MIX* on all of the proxies for the remaining constructs, this would not solve the other problems discussed above. Moreover, because the proxies for a given construct are typically highly correlated, it may be difficult to obtain precise estimates of the individual coefficients.

6.2 LATENT VARIABLE ANALYSIS

In order to mitigate the measurement and interpretation problems associated with standard multivariate analysis (such as the correlation coefficients in table 5), we estimate each underlying theoretical construct, or *latent variable*, by "combining" each of its proxies.³⁵ Because such a "combination" should possess less measurement error than the individual proxies, the "measurement error" bias in estimating the relations among the latent variables should be less serious. Moreover, this procedure allows the calculation of a *single* correlation coefficient (or regression coefficient) between any two latent variables, and this avoids difficulties in interpreting several regression coefficients for variables which attempt to measure the same construct.

To implement this procedure, we employ a latent variable structural equation model to estimate the relations among the latent or unobservable variables. Similar to most econometric analyses, we assume that the endogenous latent variable, *PERF MIX*, is related to the exogenous latent variables, *NOISE*, *CORR*, *GROWTH*, and *OTHER*, by the following linear expression:

$$\begin{aligned} \text{PERF MIX} = \gamma_0 + \gamma_1 \text{NOISE} + \gamma_2 \text{CORR} \\ + \gamma_3 \text{GROWTH} + \gamma_4 \text{OTHER} + \zeta \quad (12) \end{aligned}$$

where the γ_i are regression coefficients and ζ is the structural equation residual or disturbance term. The signs and significance levels of the γ_i are used to examine our research hypotheses. The statistical theory underlying this methodological approach and the computational algorithms are described in Joreskog [1969; 1971; 1978], Browne [1984], and Bentler [1983a; 1983b], among others, and are briefly discussed in Appendix A.

6.3 RESULTS OF THE LATENT VARIABLE ANALYSIS

As discussed in Appendix A, our data do not strictly conform to a multivariate normal distribution. This distributional assumption can be

³⁵ Another alternative would be to select a single proxy for each construct and then estimate a multiple regression using the selected proxies. This alternative is not pursued because the selection of the "best" proxy would be arbitrary. Moreover, there is likely to be a loss of information associated with dropping the other proxies for the constructs. Finally, the "best" proxy is likely to be subject to the measurement error problems discussed in the text.

critical to the properties of maximum likelihood parameter estimates of equation (12). In order to assess the sensitivity of our estimates to the absence of multivariate normality, we also provide estimates using elliptical distribution theory (which allows for a common kurtosis for each variable) and unweighted least squares (which makes no distributional assumptions). In addition, several assessments of the degree to which our model characterizes the observed data are presented in Appendix B.

The estimates for the parameters of the structural model are presented in table 6. In general, the maximum likelihood, elliptical distribution, and unweighted least squares procedures provide similar results in terms of coefficient estimates, statistical significance of the estimates, and explanatory power of the structural model. Moreover, the signs of the coefficients are generally the same across the three cases (Positive/Positive, Anything/Positive, and Constrained/Constrained). In general, the slope coefficients on the *NOISE*, *GROWTH*, and *OTHER* variables have the predicted signs (i.e., positive, positive, and negative, respectively), and each variable has a slope coefficient that is statistically significant (at the 0.05 level, two-tailed) in at least one of the cases. The coefficient on the *CORR* variable was positive in every case, which implies that a lower correlation between *ROE* and *RET* is associated with a lower relative weight being placed on *RET* (and a higher relative weight placed on *ROE*), ceteris paribus. It should be noted that the statistical significance of the coefficients varies across the three cases. The significance of the *GROWTH* and *OTHER* constructs was the most consistent across the cases, and the significance of the *NOISE* construct was the least consistent.

Finally, the correlation matrix among the constructs (not presented) indicates that two of the exogenous constructs, *NOISE* and *GROWTH*, are significantly *negatively* correlated, while the *GROWTH* and the *OTHER* constructs were significantly *positively* correlated. Although the magnitudes of the correlations vary from case to case, the correlation between *NOISE* and *GROWTH* is approximately -0.35 , and the correlation between *GROWTH* and *OTHER* is approximately 0.14 . These results suggest that, ceteris paribus, high-growth firms tend to have a lower variance of *ROE* relative to *RET* and high amounts of "other incentives" tied to stock price.³⁶

7. Discussion and Summary

This paper examines the usage of accounting return on equity (*ROE*) and security market return (*RET*) as performance measures in the cash compensation (salary plus annual bonus) contracts of chief executive

³⁶ As in multiple regression analysis, a high correlation between independent variables makes it difficult to obtain a precise estimate for the slope coefficients on these variables. As a result, the standard error associated with the estimate of each coefficient becomes large.

TABLE 6
Estimates of the Structural Model in Equation (12):
PERF MIX = γ_1 NOISE + γ_2 CORR + γ_3 GROWTH + γ_4 OTHER

	γ_1	γ_2	γ_3	γ_4	R-squared
A. Positive/Positive Case ($n = 148$)—both the RET and ROE slope are positive					
Maximum Likelihood Coefficient	.346	.083	.135	-.297	21.9%
Maximum Likelihood z-statistic	4.061	1.121	1.750	-3.982	
Elliptical Distribution Coefficient	.345	.083	.134	-.296	22.0%
Elliptical Distribution z-statistic	3.771	1.044	1.625	-3.707	
Unweighted Least Squares Coefficient	.324	.098	.196	-.314	21.9%
Unweighted Least Squares t-statistic	3.593	1.369	2.002	-3.756	
B. Anything/Positive Case ($n = 270$)—RET slope can be positive or negative, but ROE slope is positive					
Maximum Likelihood Coefficient	.037	.033	.100	-.107	1.9%
Maximum Likelihood z-statistic	.494	.538	1.433	-1.742	
Elliptical Distribution Coefficient	.037	.033	.100	-.107	1.9%
Elliptical Distribution z-statistic	.420	.457	1.214	-1.477	
Unweighted Least Squares Coefficient	.049	.006	.131	-.091	2.1%
Unweighted Least Squares t-statistic	.577	.087	1.833	-1.221	
C. Constrained/Constrained Case ($n = 370$)—both the slope for RET and ROE were constrained to be positive					
Maximum Likelihood Coefficient	-.015	.200	.043	-.096	5.2%
Maximum Likelihood z-statistic	-.226	3.909	.705	-1.847	
Elliptical Distribution Coefficient	-.016	.200	.040	-.095	5.2%
Elliptical Distribution z-statistic	-.230	3.552	.601	-1.657	
Unweighted Least Squares Coefficient	.022	.190	.096	-.114	5.8%
Unweighted Least Squares t-statistic	.321	3.597	1.427	-2.193	

officers.³⁷ Our analysis relies on analytical agency models, particularly the results of Holmstrom [1979] and Banker and Datar [1987], for a theoretical framework relating executive compensation to *ROE* and *RET* based on the “informational properties” of the two performance measures. We control for the influence of “other factors” on the form of the compensation function (such as differences in the form of the agent’s utility function and external opportunities) in order to concentrate on the “informational properties” of *ROE* and *RET*.

The theoretical model specifies a linear relation between compensation (raised to a power that is unknown to the researcher) and the performance variables *ROE* and *RET*. We estimate the parameters of this relation for each firm using the Box-Cox power transformation estimation procedure. Our results suggest that cash compensation exhibits a strong positive time-series relation with *ROE*, but only a modest time-series relation with *RET*. Our estimates of the appropriate power transformation for the compensation number, which can be interpreted as an estimate of the manager’s coefficient of proportional risk aversion, are imprecise. This result is consistent with the conclusions of prior researchers who indicate that their results are insensitive to different ad hoc specifications of the compensation variable.

Previous researchers have also suggested that the magnitude of the correlation and/or magnitude of the slope coefficient between compensation and *RET* provides evidence on the “rationality” of compensation

³⁷ We also conducted the analysis using only the “firm-specific” component of the accounting and security market measures of performance. In order to construct the firm-specific components of performance, we defined the accounting performance index to be the change in the value-weighted *ROE* for all firms on the *Compustat* tape (excluding the firms in our sample), and the security market performance index to be the (real) value-weighted *CRSP* index of security market returns. For each firm, we estimated a time-series regression of the firm’s performance on the performance index over the period 1970–84 (a separate regression was conducted for the accounting and the security market measures of performance). The results of these regressions are summarized in table 4 (panels A and B). We then defined the “firm-specific” components of the firm’s performance measure to be the residuals from these regressions.

The firm-specific components of performance were then used in the time-series regressions of compensation on performance. The results of these regressions (not reported) were similar to the results reported in table 1. In particular, the average *t*-statistic was much higher for the coefficient on the accounting measure of performance than for the market measure of performance. In conducting the cross-sectional analysis, the “independent variables” were constructed in a manner consistent with the time-series analysis. For example, since the time-series analysis regresses compensation on only the firm-specific component of performance, we defined the *NOISE* variable to be the variance of the firm-specific component of the change in *ROE* divided by the variance of the firm-specific component of *RET* in the cross-sectional analysis. The results of the cross-sectional analysis (not reported) were similar to the results reported in the text. For example, the coefficients on the *NOISE*, *CORR*, *GROWTH*, and *OTHER* variables were generally positive, positive, positive, and negative, respectively. The coefficients were most significant for the *NOISE* and *OTHER* variables. Finally, the explanatory power of the cross-sectional analysis was greatest for the Positive/Positive case.

packages (e.g., see Coughlan and Schmidt [1985], Murphy [1986], and Jensen and Murphy [1987]). However, from an agency theory perspective, there is no a priori reason to expect that a high correlation between compensation and *RET* is indicative of a "good" contract or a "bad" contract. Instead, our analysis suggests that the usage of *RET* in the contract will depend on the magnitude of its "signal-to-noise" ratio with respect to evaluating the agent's performance relative to the "signal-to-noise" ratio of other performance measures (such as *ROE*).

Hypotheses regarding cross-sectional differences in the relative amounts of "signal-to-noise" in *ROE* and *RET* were examined using a latent variable analysis. This methodology reduces problems that arise in cross-sectional regression analysis when the dependent and independent variables specified by the theory cannot be directly observed. The results of our cross-sectional analysis suggest that the degree to which compensation is related to *RET* versus *ROE* is positively related to the inverse of the degree of "NOISE" in the two performance measures. This result is consistent with the agency theory notion that the influence of factors other than the agent's action (i.e., noncontrollable factors) on a performance measure can decrease the relative weight that it receives in the compensation contract.

In addition, we find that high-"GROWTH" firms tend to place relatively more emphasis on *RET* rather than *ROE* as a performance measure. We interpret this result as consistent with the hypothesis that accounting numbers provide a less useful measure of the agent's performance when the consequences of the agent's current-period actions tend to occur in the future and are not reflected in current-period accounting numbers. We also find that firms place less importance on *RET* (relative to *ROE*) in the cash compensation contract when *RET* is of more importance in the *OTHER* components of the agent's wealth. This suggests that the overall structure of the agent's wealth plays a role in the design of each component of the agent's compensation.

Finally, our results suggest that there is a positive association between the relative influence of *RET* versus *ROE* in compensation and the correlation between *RET* and *ROE*. Alternatively stated, this result suggests that a low correlation between *RET* and *ROE* is associated with less relative weight on *RET* and more relative weight on *ROE* in the compensation function. This is consistent with the hypothesis that a lack of correlation between *RET* and *ROE* does not imply that accounting earnings contain "measurement error" from the perspective of evaluating the agent's performance. As discussed in Gjesdal [1981], an information system that does not provide very informative signals regarding the value of the firm can provide signals that are valuable in evaluating the agent's performance.

One important limitation to our analysis arises from the "undesirable" distributional properties of the *ratio* of the estimated slope coefficients

of compensation on *RET* versus *ROE*. As discussed in section 2, calculating the ratio of the two slope coefficients, in principle, allows us to control for “other factors” that affect the relation between compensation and performance. However, it is not clear how to handle negative slope coefficients, or how to assess the moments of the distribution of the ratio.

A related limitation is that our cross-sectional results are affected by the operational definition of the ratio of the slope coefficients. The results were strongest for the subsample of firms for which the slopes on both *RET* and *ROE* were positive (i.e., the Positive/Positive case). However, including firms with a negative slope coefficient on *RET* (i.e., the Anything/Positive case) substantially reduced the statistical significance of the results.

Finally, our results are limited by the restrictive structural model relating the constructs of interest. Specifically, we assume a simple recursive structure in which *PERF MIX* is an endogenous variable and *NOISE*, *CORR*, *GROWTH*, and *OTHER* are exogenous variables. If some of these constructs are endogenous, it would be desirable to develop a more complicated nonrecursive model in order to avoid the bias in the coefficients that results from ignoring the simultaneous equations aspects of the system.

APPENDIX A

Overview of Latent Variable Analysis

In a latent variable analysis, each observed variable, or proxy, is assumed to be composed of the “true score” for its underlying construct and “measurement error.” In particular, it is assumed that the observed variables have a common factor analysis structure. For the Positive/Positive case, this structure can be represented as:

$$\begin{bmatrix} MR-PP \\ BC-PP \\ RTVAR \\ RSVAR \\ ZTRANS \\ A GROW \\ S GROW \\ PERCENT \\ EQ/COMP \\ STOCK \end{bmatrix} = \begin{bmatrix} \lambda_1 & 0 & 0 & 0 & 0 \\ \lambda_2 & 0 & 0 & 0 & 0 \\ 0 & \lambda_3 & 0 & 0 & 0 \\ 0 & \lambda_4 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & \lambda_6 & 0 \\ 0 & 0 & 0 & \lambda_7 & 0 \\ 0 & 0 & 0 & 0 & \lambda_8 \\ 0 & 0 & 0 & 0 & \lambda_9 \\ 0 & 0 & 0 & 0 & \lambda_{10} \end{bmatrix} \begin{bmatrix} PERF MIX \\ NOISE \\ CORR \\ GROWTH \\ OTHER \end{bmatrix} + \begin{bmatrix} \delta_1 \\ \delta_2 \\ \delta_3 \\ \delta_4 \\ 0 \\ \delta_6 \\ \delta_7 \\ \delta_8 \\ \delta_9 \\ \delta_{10} \end{bmatrix} \tag{A1}$$

where the λ_i are regression parameters, and the δ_i are “measurement errors” or disturbance terms. The disturbance terms in equation (A1) are assumed to be uncorrelated with the underlying constructs and uncorrelated with each other. Since the *CORR* construct has only a single variable

(i.e., *ZTRANS*), the relation between *CORR* and *ZTRANS* is expressed as an identity.³⁸

Statistical procedures for estimating the relations among the latent variables, such as those implied by the structure imposed by equation (A1), have been developed by Joreskog [1969; 1971; 1978], Browne [1984], and Bentler [1983a; 1983b], among others. These procedures exploit the fact that the covariance matrix among the *observed* variables can be expressed as a function of the parameters λ_i and $\text{var}(\delta_i)$ from equation (A1) and the parameters of the variance-covariance matrix among the *latent* variables.³⁹ Specifically, let $\Sigma(\theta)$ denote the variance-covariance matrix implied by a given set of basic parameters θ (i.e., the parameters λ_i , $\text{var}(\delta_i)$, and the parameters of the variance-covariance matrix of the latent variables). Intuitively, the estimation procedure can be thought of as choosing the parameters θ whose corresponding variance-covariance matrix, $\Sigma(\theta)$, most closely reproduces the empirical variance-covariance matrix for the observed variables, S . Given the estimates for the variance-covariance matrix for the latent variables, the partial correlation coefficients (and their associated standard errors) between the latent variables, which correspond to the parameters γ_i in equation (12), can then be obtained in a straightforward manner.

More formally, the basic parameters θ are chosen to minimize the fit function:

$$(\mathbf{s} - \boldsymbol{\sigma}(\theta))' \mathbf{W}(\mathbf{s} - \boldsymbol{\sigma}(\theta)), \quad (\text{A2})$$

where \mathbf{s} is a (55×1) vector of the variances and covariances of the observed variables, and $\boldsymbol{\sigma}$ is the corresponding (55×1) vector of variances and covariances implied by the basic parameters θ .⁴⁰ The vector $(\mathbf{s} - \boldsymbol{\sigma}(\theta))$ therefore represents the *residual* variances and covariances of the observed variables (i.e., the portion of the variances and covariances that cannot be explained by the underlying parameters θ).

The matrix \mathbf{W} determines the weight that each *residual* variance or covariance receives in selecting the basic parameters θ to minimize the fit function. The choice of the matrix \mathbf{W} depends on the assumptions made about the distributions of the observed variables. For example, if the observed variables are independent and follow a multivariate normal distribution, the weight matrix can be specified so that the estimation procedure provides full information *maximum likelihood* (*FIML*) esti-

³⁸ We also constrain the variances of the latent variables to be equal to one in order to identify the system (see Long [1976] for a discussion of this issue).

³⁹ For example, the covariance between the variables *MR-PP* and *RTVAR* is $\lambda_1 \lambda_3 \cdot \text{cov}(\text{PERF MIX}, \text{NOISE})$, and the covariance between the variables *BC-PP* and *RSVAR* is $\lambda_2 \lambda_4 \cdot \text{cov}(\text{PERF MIX}, \text{NOISE})$. As discussed in n. 38, since the variances of the latent variables are constrained to equal one, the covariances among the latent variables are equal to the correlations between the latent variables.

⁴⁰ In particular, s is the lower triangular elements (including the diagonal elements) of the estimated variance-covariance matrix of the observed variables.

mates and standard errors for the coefficients relating the latent variables to other latent variables and for the coefficients relating the latent variables to their observed proxies.⁴¹ Under these conditions, the *FIML* estimates are consistent, asymptotically normally distributed, and asymptotically efficient.

To provide some evidence on how well our data satisfy the assumption of multivariate normality, table 7 presents the skewness and kurtosis of each variable used in the cross-sectional analysis. The results indicate that, even after the data have undergone a logarithmic transformation, the assumption of multivariate normality can be rejected (for an overall test of the assumption of multivariate normality, we use the Mardia [1970] test, which is distributed as a standard normal variable in large samples). Multivariate normality is least seriously violated for the Positive/Positive case, followed by the Constrained/Constrained case and the Anything/Positive case. For the proxies for the exogenous constructs, the *EQ/COMP* and *STOCK* variables are positively skewed, and the *S GROW* variable has a more positive kurtosis than would be expected in a normal distribution. For the proxies for the endogenous constructs, the Positive/Positive proxies are closest to satisfying the normality assumption, the Anything/Positive proxies are leptokurtic, and the Constrained/Constrained proxies are positively skewed.

Bentler [1985, pp. 53–54] reports that “there is little empirical or theoretical guidance available as to when a statistically significant variation from normality becomes large enough to affect structural modeling conclusions.” The maximum likelihood *parameter* estimates are generally robust to violations of normality. However, the standard errors must be interpreted with caution when the variables are not normally distributed. In order to provide some evidence on the appropriateness of the maximum likelihood procedures for our problem, we also conduct the estimation using procedures that allow for distributions that are more general than multivariate normal.

For situations in which the primary violation of normality results from the kurtosis of the variables, Browne [1984] has developed estimation procedures (i.e., the specification of the form of the *W* matrix) based on *elliptical distribution theory*. In particular, the elliptical distribution estimation procedures are based on the assumption that the variables have a *common* coefficient of kurtosis. Under these conditions, the resulting elliptical estimators are consistent and asymptotically efficient. The estimation procedure also provides standard errors for the parameter estimates which allow significance tests to be performed. For our data

⁴¹ In particular, if the weight matrix is iteratively updated during estimation, minimizing equation (A2) is equivalent to minimizing the function $\log[\det(\Sigma)] + \text{trace}[S \Sigma^{-1}] - \log[\det(S)] - p$, where \log is the natural logarithm, \det is the determinant of a matrix, trace is the trace of a matrix, and p is the number of observed variables ($p = 10$ for our model). See Browne [1984, p. 65] for a discussion of this point. This function corresponds to the standard likelihood function for latent variable estimation.

TABLE 7
Descriptive Statistics for the Variables Used in the Cross-Sectional Analysis of the Performance Measures Used in Compensation Contracts

A. Positive/Positive Case ($n = 148$ firms)—both the RET and ROE slopes are positive											
Variable ^a	MR-PP	BC-PP	RTVAR	RSVAR	ZTRANS	A GROW	S GROW	PERCENT	EQ/COMP	STOCK	
SKEWNESS	-.4009	-.1263	.4288	-.3222	-.0063	.0990	.0115	.5183	.9607	.6918	
KURTOSIS	-.1808	-.1386	-.0223	-.0357	.0917	.1832	1.2314	-.0166	.4909	.2400	
Standard Error for Univariate Tests of Skewness: .201 ^b											
Standard Error for Univariate Tests of Kurtosis: .401 ^c											
z-statistic for Mardia Test for Multivariate Normality: 7.27											
B. Anything/Positive Case ($n = 270$ firms)—RET slope can be positive or negative, but ROE slope is positive											
Variable:	MR-AP	BC-AP	RTVAR	RSVAR	ZTRANS	A GROW	S GROW	PERCENT	EQ/COMP	STOCK	
SKEWNESS	-.2267	.6440	.6162	-.2691	-.2268	.1945	.0561	.4016	.7843	.5225	
KURTOSIS	6.8142	6.7658	.1873	-.1834	.1654	.2926	.8706	-.1191	.2615	.0973	
Standard Error for Univariate Tests of Skewness: .149											
Standard Error for Univariate Tests of Kurtosis: .298											
z-statistic for Mardia Test for Multivariate Normality: 24.64											
C. Constrained/Constrained ($n = 370$ firms)—both the slopes for RET and ROE were constrained to be positive											
Variable:	MR-CC	BC-CC	RTVAR	RSVAR	ZTRANS	A GROW	S GROW	PERCENT	EQ/COMP	STOCK	
SKEWNESS	.9621	1.0719	.8384	-.1590	-.0656	.0144	.0242	.5128	.8477	.5748	
KURTOSIS	-.1641	.0425	.8560	-.1925	.3006	.1846	.6602	-.1307	.3334	.0385	
Standard Error for Univariate Tests of Skewness: .127											
Standard Error for Univariate Tests of Kurtosis: .254											
z-statistic for Mardia Test for Multivariate Normality: 16.15											

^a See table 2 for a definition of the variables.

^b In a large sample, the coefficient of skewness is distributed normally with a standard error that is approximately equal to $\sqrt{6/(N+1)}$.

^c In a large sample, the coefficient of kurtosis is distributed normally with a standard error that is approximately equal to $\sqrt{24/(N+1)}$.

(see the results in table 7), the assumption of equal coefficients of kurtosis is more likely to be met for the Positive/Positive and the Constrained/Constrained cases than for the Anything/Positive case.

In order to provide estimates under even more general assumptions regarding the distributions of the variables, we also estimate the coefficients using *unweighted least squares*.⁴² While the coefficients provided by this method are consistent, they are not asymptotically efficient. One further disadvantage of the unweighted least squares estimation procedure is that the standard errors of the coefficients cannot be directly calculated. Joreskog [1981] suggests that the standard errors can be calculated using jackknifing methods. Therefore, our empirical results present jackknifed standard errors for the coefficients estimated using the unweighted least squares method. It should be noted, however, that jackknifing methods can be sensitive to skewness in the data (see Mosteller and Tukey [1977]).

APPENDIX B

Assessment of the Latent Variable Model

Before the parameter estimates can be interpreted, it is important first to examine how well the estimated model implied by the theory characterizes the data. In particular, it is critical to demonstrate that each proxy measure has a positive and statistically significant relationship with its construct. This examination is done by determining whether each λ is positive and has a "large" z -statistic.⁴³ The results in table 8

⁴² In this case, the estimation procedure is equivalent to minimizing $(\mathbf{s} - \sigma(\theta))' \mathbf{W}(\mathbf{s} - \sigma(\theta))$, where \mathbf{W} is an identity matrix. This is equivalent to minimizing the sum of the squared covariance residuals.

An alternative estimation procedure is to use the Arbitrary Distribution Function (*ADF*) estimators developed by Browne [1984]. These estimators impose no distributional restrictions except that the first four moments and the eighth moment of the distributions exist. This procedure is computationally intensive relative to the ones that we employ and sensitive to the starting values provided to the algorithm. The *ADF* estimators provided results similar to those reported in tables 6 and 8 for the Positive/Positive case. However, we experienced convergence problems in attempting to estimate the parameters of the other two cases. This result may be due to the fact that the relations between the variables appear to be much stronger in the Positive/Positive case than in the other two cases. Moreover, the simulation results in Browne [1984] seem to suggest that the *ADF* estimates can be biased in small samples. Despite these problems, the distribution-free nature of the *ADF* estimators may make them attractive in research studies in which multivariate normality is a tenuous assumption.

⁴³ For ease of comparison, the parameters reported in tables 6 and 8 are *standardized* (i.e., computed as if the latent variables have unit variances and the observed variables also have unit variances). Since the scales for the variables are irrelevant for our purposes, this does not affect the substantive inferences made regarding the research hypotheses. However, the z -statistics reported in the tables are based on the unstandardized solution (because the input into the estimation algorithm is the covariance matrix not the correlation matrix). The covariance matrix is used as the input because the mathematical statistics that form the basis of the estimation are developed from knowledge of the distribution of the covariance matrix (e.g., if the data are multivariate normal, the covariance matrix is Wishart distributed). The distribution of the correlation matrix is less obvious.

TABLE 8
Estimation of the Parameters of the Latent Variable Model Expressed in Equation (A1)

CONSTRUCT: PROXY: PARAMETER:	PERF MIX		NOISE		CORR		GROWTH		OTHER		GOODNESS-OF-FIT INDEXES ^c				
	λ_1 MR-PP	λ_2 BC-PP	λ_3 RTVAR	λ_4 RSVAR	λ_5 ZTRANS	λ_6 A	λ_7 GROW	λ_8 S	λ_9 EQ/COMP	λ_{10} STOCK	BENT-BON	CHISQ	prob.	% sig.	
A. Positive/Positive Case ($n = 148$ firms)—both the RET and ROE slopes are positive															
ML ^b	0.991	0.889	0.979	0.716	1.000*	0.797	1.000	0.883	1.000	0.980	.960	50.948	.002	1.8%	
ML	z-statistic:	13.728	11.963	9.318	7.517	—	11.701	—	13.716	—	16.484	—	—	—	
EL ^b	0.992	0.889	0.980	0.714	1.000	0.797	1.000	0.883	1.000	0.980	.939	42.091	.024	1.8%	
EL	z-statistic:	12.546	10.978	8.612	6.944	—	10.736	—	12.519	—	14.918	—	—	—	
ULS ^b	0.975	0.900	0.930	0.734	1.000	0.825	0.912	0.871	1.000	0.975	.985	N/A	N/A	0.0%	
ULS	t-statistic:	17.412	15.497	6.940	5.243	—	3.986	3.948	34.701	—	87.865	—	—	—	
B. Anything/Positive Case ($n = 270$ firms)—RET slope can be positive or negative, but ROE slope is positive															
ML	Coefficient:	0.837	1.000	0.828	0.827	1.000	0.815	0.982	0.888	1.000	0.979	.965	75.188	<.001	3.6%
ML	z-statistic:	17.046	—	10.967	10.967	—	12.316	14.357	18.696	—	22.237	—	—	—	
EL	Coefficient:	0.837	1.000	0.828	0.827	1.000	0.814	0.982	0.888	1.000	0.979	.948	48.172	.005	3.6%
EL	z-statistic:	14.017	—	9.183	9.181	—	10.276	11.922	15.278	—	17.900	—	—	—	
ULS	Coefficient:	0.880	1.000	0.822	0.826	1.000	0.813	.978	0.875	1.000	0.969	.991	N/A	N/A	0.0%
ULS	t-statistic:	21.015	—	12.422	9.741	—	12.357	12.256	47.831	—	155.815	—	—	—	

C. Constrained/Constrained ($n = 370$ firms)—both the slopes for *RET* and *ROE* were constrained to be positive

<i>ML</i>	Coefficient:	0.989	0.938	0.823	0.804	1.000	0.820	0.979	0.900	1.000	0.980	.956	146.496	<.001	5.4%
<i>ML</i>	<i>z</i> -statistic:	17.163	16.435	12.810	12.611	—	15.017	17.593	22.403	—	26.105				
<i>EL</i>	Coefficient:	0.991	0.936	0.823	0.804	1.000	0.818	0.981	0.900	1.000	0.980	.932	106.668	<.001	5.4%
<i>EL</i>	<i>z</i> -statistic:	15.350	14.664	11.496	11.324	—	13.421	15.711	19.767	—	22.820				
<i>ULS</i>	Coefficient:	0.928	0.999	0.818	0.803	1.000	0.843	0.947	0.891	1.000	0.977	.987	N/A	N/A	5.4%
<i>ULS</i>	<i>t</i> -statistic:	48.353	230.396	15.212	12.360	—	12.730	12.829	65.044	—	227.200				

^a Parameter was constrained.

^b *ML* refers to maximum likelihood estimates and standard errors, *EL* refers to elliptical distribution estimates and standard errors, and *ULS* refers to unweighted least squares jackknifed estimates and standard errors.

^c *BENT-BON* is the Bentler-Bonnet index described in Appendix B, *CHISQ* is likelihood ratio test statistic, prob. is the probability of observing a chi-squared value at least as large as *CHISQ* under the null hypothesis that the theoretical model describes the observed data, and % sig. refers to the percentage of standardized residuals that are statistically significant at the 0.05 level (two-tailed).

indicate that for each case and for each estimation method, each λ is close to one and is highly significant.⁴⁴ It is also desirable to observe that the structural equation has a nontrivial coefficient of determination.⁴⁵ The explanatory power of the structural model varied considerably across the three cases (table 8). The Positive/Positive case had the highest *R*-squared (21.9%), followed by the Constrained/Constrained (5.2%) and the Anything/Positive (1.9%) cases.⁴⁶

Several common indexes are used to assess the "goodness of fit" for the estimated theoretical specification. One index is the maximum likelihood chi-square statistic, which is computed by multiplying the sample size by the minimum of the maximum likelihood fit function (see n. 41). If the theoretical model is able to reproduce the observed variance-covariance matrix, the chi-squared value will be small (with an associated large probability level). The probability value associated with the chi-squared is typically compared to a standard "cutoff" of 0.01 or 0.05. For each of our three cases, we can reject (at the 0.01 level) the hypothesis that our model explains the covariance structure of the observed variables. However, as discussed by Joreskog and Sorbom [1984], among others, the chi-square is very sensitive to sample size and departures from multivariate normality. In particular, large sample sizes and lack of multivariate normality tend to increase the chi-square above what can be expected from specification error in the theoretical model.

An alternative goodness-of-fit statistic has been developed by Bentler and Bonett [1980]. This index compares the value (denoted *Q*) of the fit

⁴⁴ Several of the loadings were constrained to be equal to one in table 8. This was done because the unconstrained estimation exhibited an "improper solution" (i.e., a standardized loading, λ , greater than one and an associated negative measurement error variance). Following the procedures recommended in Van Driel [1978] and Gerbing and Anderson [forthcoming], the cause of the improper solution appeared to be sampling variation, as opposed to model misspecification. Therefore, the loading for the improper solution was constrained at one, and the respecified model was estimated. For the constrained λ 's, we do not report *z*-statistics.

It is important to note, however, that this type of respecification can produce biased estimates for the loadings and correlations between latent variables. Specifically, the simulation results in Gerbing and Anderson [forthcoming] suggest that the λ 's for the factor containing the improper solution are "too large" and the correlations between latent variables are "too small." In addition, the standard errors for the parameter estimates are "too large." This type of bias is a limitation to our empirical analysis.

⁴⁵ See Fornell and Larcker [1981], Bagozzi, Fornell, and Larcker [1981], and Joreskog and Sorbom [1984] for a discussion of various indexes of explanatory power in a latent variable context.

⁴⁶ In psychometric terms, the large, positive λ 's suggest that the proxies exhibit *convergent validity* (i.e., different measures of the same construct are highly correlated). In addition, the results also suggest that the proxies exhibit *discriminant validity* (i.e., different measures of different constructs are less highly correlated than different measures of the same construct). Discriminant validity can be demonstrated if none of the confidence intervals for the bivariate correlations between constructs includes one. Although not reported in table 8, none of our confidence intervals includes the value of one.

function in equation (A2) for the theoretical model to the value (denoted Q_0) of the fit function for a "null model" (i.e., a model where each proxy variable is its own latent variable and each of these latent variables is constrained to be independent). The Bentler and Bonett index is computed as $[1 - Q/Q_0]$. This index can be thought of as providing some indication of how much better the theoretical model fits the data relative to the severely constrained null model. Bentler [1985] suggests that the Bentler and Bonett index should exceed 0.90 in reasonable models. For each of our cases, the Bentler-Bonett index exceeds this suggested "cutoff".

Finally, we also examine the standardized residual covariance (or correlation) matrix, $(s - \sigma(\theta))$, for large and statistically significant values. Many large residuals are indicative of an inability of the theoretical model to reproduce the observed variance-covariance matrix. Therefore, we report the percentage of standardized residual covariances that are statistically significant at the 0.05 level (two-tailed). For the Positive/Positive case, less than 2% of the standardized residual covariances were significant. For the Anything/Positive case, less than 4% were significant, and for the Constrained/Constrained case, less than 6% were significant. Overall, we conclude that our model has a "reasonable" ability to reproduce the observed variance-covariance matrix.

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