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Pay, Performance, and Turnover of Bank CEOs

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A new data set covers chief executive officers (CEOs) of large commercial banks over the period 1982–87. For newly hired CEOs, the elasticity of pay with respect to assets is about one-third. For continuing CEOs, the change in compensation depends on performance, as measured by stock and accounting returns. The sensitivity of pay to performance diminishes with experience, but the returns are not filtered for peer-group returns. Logit regressions relate the probability of CEO departure to age and performance, as measured by stock returns filtered for peer-group returns; CEO experience does not influence this relationship.

The relation of chief executive officer (CEO) pay and turnover to performance and characteristics of companies has been the focus of a number of theoretical and empirical studies. This study extends this analysis to a new data set that covers large commercial banks over the period 1982–87.

We begin with a simple theory, motivated by Rosen's (1982) model, that generates positive matching between CEO skill, and hence compensation, and the size of banks. We then extend the analysis to consider how compensation evolves over time in response to observations on performance. We assume in the main discussion that changes in pay correspond to changes in expected marginal product. One proposition that emerges from this analysis is that the sensitivity of changes in pay to performance diminishes as CEO experience increases.

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Our empirical study begins with the relation between the levels of compensation and bank assets for newly hired CEOs. Then we study how the compensation of continuing CEOs responds to performance based on stock returns and accounting earnings. We examine how CEO experience affects the sensitivity of pay to performance and whether actual or relative performance matters.

In the next section we explore how the correlation between the levels of compensation and assets varies with CEO experience. The variations in this correlation depend on the growth of compensation, considered in the previous section, and also on the growth of assets. In addition, the correlation would be affected by CEO turnover if there is a systematic tendency to terminate CEOs who perform the poorest.

The final section uses logit regressions to relate the probability of CEO turnover to age and performance. We consider the effects from market- and accounting-based returns, the distinction between actual and relative performance, and the effects of experience. The results for turnover provide a number of interesting comparisons with those for compensation growth.

A Model of Bank Size, CEO Compensation, and Performance

We begin with a simple theoretical model that allows for matching between the size of a bank and the quality (and hence compensation) of the CEO. The production function for bank i is

$$y_i = \lambda_i F(k_i, A_i), \quad (1)$$

where λ_i is a technological or resource factor for the bank, k_i is the anticipated skill level of the CEO, and A_i represents bank assets or, more generally, an array of inputs that includes labor. In Rosen's (1982) model, CEO skill involves the quality of decisions ("general atmosphere") and the ability to supervise. The units are defined so that a CEO with twice as much supervisory talent can administer twice as many people at a given level of effectiveness. For a given quality of decision making, it is natural to assume constant returns to scale in the other inputs, including supervisory talent. Because $F(\cdot)$ exhibits constant returns to inputs aside from decision-making ability, it must show increasing returns with respect to all inputs.

The bank's net revenue is

$$\lambda_i \cdot F(k_i, A_i) - v(k_i) - rA_i, \quad (2)$$

where r , the constant cost for assets, represents payments to depositors or the opportunity cost for equity. The bank faces the CEO wage function, $v(k_i)$ with $v' > 0$, that relates CEO compensation, v_i , to the level of skill, k_i . The function $v(k_i)$ is determined in the overall population, as in Rosen (1982), by the distribution of the supply of CEO talent and by the demand

for CEO skill (from banks and also from other companies if CEO talent is substitutable across fields).

The bank chooses k_i and A_i (and other inputs) to maximize its net revenue. The second-order conditions for this maximization require $v'' > 0$; that is, at least in the neighborhood of the selected k_i , CEO pay must rise at an increasing rate with the level of skill. Since the function $v(k)$ must satisfy this condition in a full equilibrium, we assume that $v'' > 0$ applies.

The conditions for maximization of net revenue determine A_i and k_i —and therefore $v_i = v(k_i)$ —as functions of λ_i and r . For given r , an increase in λ_i implies increases in A_i , k_i , and v_i ; that is, better institutions (higher λ_i) are larger in the sense that they assemble more assets, hire a better CEO (higher k_i), and pay the CEO more (higher v_i).

An increase in λ_i lowers the ratio of CEO skill to size, k_i/A_i , because an increase in k_i raises the marginal cost of CEO talent, $v'(k_i)$, relative to the constant marginal cost, r , of assets. The behavior of the ratio of CEO pay to assets, v_i/A_i , is unclear because v_i/k_i tends to rise with k_i . Typical empirical results indicate that the elasticity of CEO pay with respect to a size variable, such as bank assets, is positive, less than one, and roughly constant at about one-third. This finding means that the ratio v_i/A_i declines as λ_i increases; a result that is possible but not inevitable within the model.

The analysis treated k_i and A_i as freely adjustable inputs. This treatment of k_i seems most appropriate at the time a CEO is installed.¹ Therefore, our initial empirical analysis deals with the relation between compensation and bank assets during a CEO's initial year in office. However, the variable v_i should be interpreted as the expected present value of compensation attached to becoming the CEO of bank i . We consider below the relation of this present value to the compensation in the initial year in office, which we denote by w_{i1} .

There are adjustment costs associated with changes in assets A_i (or numbers of employees and other inputs). In fact, variations in A_i across firms for historical reasons, rather than because of differences in current technological parameters λ_i , can also be viewed as generating cross-sectional dispersion in k_i and v_i . That is, bigger banks tend to hire better CEOs even if these banks do not currently have access to better production functions.

We consider now the change in compensation over time, assuming that the CEO remains in office. Later we allow for CEO turnover and for

¹ This freedom of choice may be limited if, as is usually the case, the promotion to CEO comes from within the company rather than as an appointment from outside. (In our sample, 46 of the 60 newly installed CEOs, or 76%, had more than 1 year of prior tenure on the corporate board.) Lazear and Rosen (1981) argue that the selection as CEO should be viewed as the final match of an extended tournament involving insiders.

growth in bank assets. In determining the growth of compensation, we assume that the bank's directors—acting in the interest of the bank's owners—use information revealed by two kinds of variables: measures of the bank's performance for year t , PERF_{it} , and corresponding measures of performance for a peer group of banks, denoted by PERF_{it}^* . In the empirical analysis, PERF_{it}^* is the average of PERF_{it} for year t and for the geographical region where bank i is located.

We begin with the hypothesis that the growth of compensation, $\log(w_{it}/w_{it-1})$, corresponds to the growth of the CEO's expected marginal product. Thus, the change in pay considers new information about the CEO's skill level, k_i , and also allows for shifts in production conditions ($F(\cdot)$ in eq. [1]) or factor supplies ($v(\cdot)$ or r in eq. [2]). This analysis abstracts from explicit or implicit labor contracts that allow for significant departures of the growth in compensation from the growth in expected marginal product.²

The main information about CEO talent comes from the observation of relative performance, $\overline{\text{PERF}}_{it} = \text{PERF}_{it} - \text{PERF}_{it}^*$ (see Holmstrom 1982). Consider the model

$$\overline{\text{PERF}}_{it} = \alpha + \beta k_i + \varepsilon_{it}, \quad t = 1, 2, \dots, \quad (3)$$

where α and β are known constants and ε_{it} has zero mean and constant variance σ_ε^2 . The formulation assumes that the skill level of the average CEO is a known constant; in particular, it is unnecessary to learn about this average value. The expected value of k_i conditioned on data through T years in office, denoted by $E(k_i)|_T$, depends on the sample mean of the $\overline{\text{PERF}}_{it}$ and on the prior information about k_i that was used in the initial hiring decision. Suppose that this prior information is equivalent in terms of information content to T_0 observations on $\overline{\text{PERF}}_{it}$, where T_0 need not be an integer. Then it is straightforward to show that the relation between $E(k_i)|_T$ and $\overline{\text{PERF}}_{it}$ involves the coefficient $1/(T_0 + T)$. Therefore, a higher level of experience, T , implies a smaller sensitivity of $E(k_i)|_T$ to $\overline{\text{PERF}}_{it}$.³

For given k_i , the CEO's value of marginal product varies with disturbances that are industry- or region- or economywide. We assume that these elements are captured by the aggregate performance variable,

² Becker and Stigler (1975), Lazear (1979), and others argue that deferred compensation has desirable incentive effects on workers. This pattern requires significant departures of wages from expected marginal products.

³ See Murphy (1986) for a similar result. The analysis is more complicated if the CEO's skill changes over time. Holmstrom (1983, sec. I.2) assumes that the skill level, k_{it} , evolves as a random walk. The sensitivity of $E(k_{it})|_T$ to performance still diminishes with T , but this sensitivity now approaches an asymptote that is positive rather than zero.

$PERF_{it}^*$. Therefore, if variations in compensation correspond to variations in expected marginal product, the changes in compensation depend on $PERF_{it}^*$ as well as \overline{PERF}_{it} . Denoting compensation for year t by w_{it} and assuming a linear functional form, the growth of compensation is given by

$$\log(w_{it}/w_{it-1}) = a + b \cdot \frac{\overline{PERF}_{it}}{(T_0 + T)} + c \cdot PERF_{it}^*, \quad t = 2, 3, \dots \quad (4)$$

The constant term a reflects growth in value of marginal product associated with greater experience of the CEO or with industry- or region- or economy-wide productivity growth.

The specification in equation (4) implies that CEOs assume compensation risk associated with uncertainty about aggregate effects, $PERF_{it}^*$, and relative performance, \overline{PERF}_{it} . The bank could insure the CEO against aggregate risk by setting compensation independently of $PERF_{it}^*$. In this case, the growth in compensation in equation (4) would depend only on relative performance, \overline{PERF}_{it} (see Holmstrom 1979).

One attraction of relative-performance agreements is that the aggregate variable $PERF_{it}^*$ cannot be manipulated significantly by a single CEO. Nevertheless, these kinds of contracts create problems associated with the presence of gaps between wages and expected marginal products: (1) if a favorable realization of $PERF_{it}^*$ results in an excess of the expected marginal product over the wage, the CEO can quit; (2) if an unfavorable realization of $PERF_{it}^*$ creates an excess of the wage over the expected marginal product, the bank can effectively renege by treating the CEO badly; and (3) insulation of w_{it} from aggregate variables gives the CEO insufficient incentive to take actions that mitigate the effects of aggregate disturbances on an individual bank's performance (see Jensen and Murphy 1988, p. 17). The benefit from insulating compensation growth from aggregate performance is also likely to be small. If the CEO cares a great deal about aggregate risk, then he can insure himself by taking the appropriate position in the stock market (e.g., by going short on a portfolio of bank stocks). Because the benefits are small and the costs are likely to be significant, we do not anticipate that CEO contracts would be sheltered from aggregate performance. In any event, the effect of $PERF_{it}^*$ in equation (4) is a test for the prevalence of contracts that reduce or eliminate the sensitivity of CEO pay to aggregate factors.

Risks associated with relative performance, \overline{PERF}_{it} , in equation (4) are harder to reduce without compromising incentives for good CEO behavior. Moreover, if the CEO has superior information about skill or effort, k_i , then the CEO's risk associated with \overline{PERF}_{it} may not be great.

For a CEO in the initial year, the level of assets, A_i , relates to the expected

present value of compensation, v_i . Hence v_i corresponds to the expected present value of the w_{it} —that is, the sum of initial salary, w_{i1} , and the anticipated discounted values of future salaries, w_{i2} , w_{i3} , . . . , that are determined from equation (4). Since PERF_{it} reflects news, the date 1 expectations of $\overline{\text{PERF}}_{it}$ and PERF_{it}^* are zero; hence, $E[\log(w_{it}/w_{it-1})] = a$. Therefore, for given parameters of the distributions of $\overline{\text{PERF}}_{it}$ and PERF_{it}^* , v_i differs from w_{i1} only by a multiplicative constant. Accordingly, the empirical work uses w_{i1} —executive pay in the first year on the job—as a counterpart of v_i .

Setup of the Empirical Analysis

The empirical work uses a new panel data set on CEO compensation for large U.S. commercial banks over the period 1982–87. From the standpoint of testing theories about executive compensation, the banking industry is attractive because of the presence of a large number of firms that produce a similar product. The sample of 83 banks is a subset of the 140 banks that ranked highest in assets in 1986. Attrition of the sample occurred because of unavailable data, sometimes because banks disappeared as independent entities or were foreign owned (and therefore did not file disclosure statements with the Securities and Exchange Commission [SEC]).⁴ These considerations mean that some banks appear in the sample for some years but not others. Table 1 shows the composition of the sample by year and geographical region.

The data are from individual proxy statements, *Compuserve*, *Business Week*'s annual listing of the top 200 banks, and Standard and Poors' company reports. The information includes for each bank and year the total of salary and bonus of the highest-paid executive (usually designated as CEO), assets, accounting earnings and earnings per share, share prices, dividend yields, age of the CEO, and number of years of prior experience as CEO. The data set also includes the geographical location of the bank. Real assets are the ratio of the nominal, year-end values to the seasonally adjusted December consumer price index (CPI). Real compensation and earnings are the ratios of the nominal figures to the annual average of the CPI. Real returns to shareholders are the real dividend yield plus the growth rate of nominal share prices (year-end to year-end) less the inflation rate (December–December CPI). Appendix table A1 shows the means and standard deviations of the variables that we use in the subsequent analysis.

In a preliminary study of a dozen banks, the compensation figures were

⁴ In the case of a merger, the important consideration for our purpose is not whether the bank remains as the same entity over time but whether the CEO is continuing or new. If the CEO after the merger is the same person who was CEO of one of the original banks, we treated the CEO as continuing with past-performance characteristics corresponding to those of the old bank.

Table 1
Composition of Sample by Year and Region

	1987	1986	1985	1984	1983	1982	Total
New CEO sample:							
New England	2	0	1	1	0	1	5
New York City	2	0	1	2	0	2	7
Mid-Atlantic	1	1	1	3	2	3	11
Midwest	1	2	4	2	4	4	17
South	1	2	3	1	2	2	11
Texas/Oklahoma	1	1	0	1	0	0	3
West	2	1	0	1	1	1	6
Total	10	7	10	11	9	13	60
Continuing sample:							
New England	5	6	5	5	5	...	26
New York City	6	9	8	7	9	...	39
Mid-Atlantic	8	11	12	10	10	...	51
Midwest	19	16	14	17	15	...	81
South	16	16	16	16	14	...	78
Texas/Oklahoma	1	3	5	4	5	...	18
West	7	7	8	8	7	...	37
Total	62	68	68	67	65		330

NOTE.—Mid-Atlantic includes the District of Columbia, Maryland, New York (outside of New York City), New Jersey, and Pennsylvania. Midwest includes Indiana, Illinois, Kentucky, Michigan, Minnesota, Missouri, Ohio, and Wisconsin. South is the southeastern states, excluding Texas. West includes western and mountain states and Hawaii. States not mentioned had no banks in the sample.

expanded from salary and bonus to include the estimated value of stock options granted and some elements of deferred compensation. Because the information about stock options, including when they were granted, was often incomplete, it was not possible to value these options precisely or to assign them unambiguously to a particular year. The expanded figures on compensation differed from salary and bonus by an amount that was volatile across years and banks. However, the general nature of the relation between compensation and other variables did not depend very much on whether narrow or broad compensation was used. For this reason, the present study is limited to compensation in the form of salary and bonus.

Results for the Initial Year in Office

Over the period 1982–87 there are 60 observations on CEOs in their first year of office. For these new CEOs, the regression of $\log(\text{real compensation})$ on $\log(\text{real assets})$ is

$$\log(w_{it}) = -3.84 + .316 \cdot \log(A_{it}),$$

(0.30) (.032)

(5)

$$N = 60, \quad R^2 = .623, \quad \sigma = .267,$$

where standard errors appear in parentheses. Thus the estimated coefficient of $\log(A_{it})$ is positive and highly significant, with a t -value of 10. Equation (5) can be interpreted from the theory in terms of the joint effect of the exogenous (unobservable) bank characteristic, λ_i , on CEO compensation and assets. Thus, in equilibrium, the responses imply that compensation moves with an elasticity of about one-third with respect to assets.

The relation between executive pay and firm size shown in equation (5) is typical of findings from previous studies for various industries and time periods, although many researchers use sales, rather than assets, as the size variable (see Ciscel and Carroll [1980] for a survey). These empirical results have spurred a good deal of theoretical analysis, going back to Simon (1957) and Beckmann (1960) and including more substantial recent models by Rosen (1982) and Keren and Levhari (1983). Although these theories can rationalize the positive relation between CEO pay and size, the puzzle is why the relation is so similar across industries and time and why the estimated elasticity of pay with respect to size is usually close to one-third.

Shifts in the CEO wage function, $v(\cdot)$ —which might reflect changes in the aggregate demand for CEOs—could shift the relation between compensation and assets. The regression in equation (5) is, however, stable over time for the sample period 1982–87. Year dummies are jointly insignificant if added to the equation.

Results on Changes in Compensation for Continuing CEOs

Equation (4) brought in performance and applied to the change in pay over time for CEOs that remained in office. In the empirical analysis, the wage for year t represents partly base salary set at the beginning of the year and partly bonus set at the beginning of the next year. (The data set does not include a separation between base salary and bonus.) Performance for year t could affect this year's pay (especially through the bonus) and also next year's pay (especially through the base salary).

Two measures of performance turned out to be important empirically: real rate of return to shareholders (based on stock-market prices and dividend yields) and accounting-based real rate of return. Even if shareholders care only about market returns, the accounting returns may provide independent information about CEO ability or effort (see Holmstrom [1979], sec. 4, for a general discussion).⁵ Therefore, it is reasonable that both measures of performance matter for the change in compensation.

The variable RS_{it} is the total real rate of return (real stock-price appreciation plus real dividend yield) to stockholders of bank i in year t . This variable shows little serial correlation: a regression for 1983–87 (410 ob-

⁵The usual descriptions of CEO bonus plans also suggest an important role for accounting earnings; see Fox (1979).

servations) of RS_{it} on RS_{it-1} yields the estimated coefficient .03, SE = .06. (The mean of RS_{it} for this sample is .134.) Therefore, aside from a constant to measure the average real rate of return, the value of RS_{it} represents the news for year t that should matter for the adjustment of compensation. Empirically, it turns out that RS_{it} and RS_{it-1} each affect the change in compensation.

The accounting-based rate of return, denoted by RA_{it} , is the real earnings yield: the ratio of bank i 's real earnings per share during year t to the real price per share at the end of year $t - 1$. Aside from the use of accounting data, the real earnings yield is comparable in dimension to the market-based real rate of return, RS_{it} . Unlike RS_{it} , RA_{it} is highly positively correlated from year to year: a regression for 1983–87 (390 observations) of RA_{it} on RA_{it-1} yields the estimated coefficient 1.09, SE = .06. (The mean of RA_{it} for this sample is .113.) Therefore, the first difference of the accounting-based returns, $\Delta RA_{it} = RA_{it} - RA_{it-1}$, approximates the news in this series.

We initially neglect the role of CEO experience and use the variables RS_{it} , RS_{it-1} , and ΔRA_{it} as empirical counterparts of $PERF_{it}$. For 330 observations on continuing CEOs over 1983–87, a regression for the growth rate of real compensation is⁶

$$\begin{aligned} \log(w_{it}/w_{it-1}) = & .079 + .080 \cdot RS_{it} + .094 \cdot RS_{it-1} \\ & (.012) \quad (.031) \quad (.029) \\ & + .47 \cdot \Delta RA_{it}, \\ & (.10) \end{aligned} \quad (6)$$

$$N = 330, \quad R^2 = .146, \quad \sigma = .164.$$

Thus the estimated coefficients of RS_{it} , RS_{it-1} , and ΔRA_{it} are each significantly positive. If the lagged value of the change in the accounting return, ΔRA_{it-1} , is added to equation (6), the estimated coefficient is insignificant: $-.21$, SE = .18, and the rest of the results change little. If the current and lagged levels of the accounting returns are included separately, instead of as a first difference, the estimated coefficients are .51, SE = .11, for RA_{it} and $-.29$, SE = .19, for RA_{it-1} . This result is consistent with the specification that accounting returns enter as the first difference, ΔRA_{it} , as in equation (6). (The test statistic for this restriction is $t_{325} = 1.2$.)

Since the coefficients of RS_{it} and RS_{it-1} in equation (6) are nearly equal, the 2-year average real rate of return to shareholders is a satisfactory measure

⁶ Although eq. (6) could be estimated jointly with eq. (5), there would be no gain over the separate estimation unless the error terms were substantially correlated.

of market-based performance. Defining $RS2_{it}$ to be the average of RS_{it} and RS_{it-1} , the regression becomes

$$\log(w_{it}/w_{it-1}) = .079 + .174 \cdot RS2_{it} + .46 \cdot \Delta RA_{it},$$

$$(.012) \quad (.039) \quad (.10) \quad (7)$$

$$N = 330, \quad R^2 = .146, \quad \sigma = .164.$$

In this form, the t -values for the estimated coefficients of $RS2_{it}$ and ΔRA_{it} are 4.5 and 4.8, respectively. (The test statistic associated with equality of the coefficients on RS_{it} and RS_{it-1} is $t_{326} = 0.3$.)⁷

The serial correlation of the residuals from equation (7) is negative but insignificantly different from zero. For example, a regression of the residuals at date t on those at date $t - 1$ (231 observations) yields the estimated coefficient $-.088$, $SE = .061$. Similarly, if the first lag of the dependent variable is added to equation (7), the estimated coefficient (231 observations) is negative but insignificantly different from zero: $-.079$, $SE = .059$.

The relation estimated in equation (7) is stable over time for the sample period 1983–87. In particular, year dummies are jointly insignificant if added to the equation.

Although the level of real pay in the initial year relates to the level of real assets in equation (5), the growth rate of real assets turns out not to be a performance variable that is significantly related to the growth rate of real pay. If the variable $\log(A_{it}/A_{it-1})$ is added to equation (7), the results are

$$\log(w_{it}/w_{it-1}) = .074 + .168 \cdot RS2_{it} + .46 \cdot \Delta RA_{it}$$

$$(.014) \quad (.039) \quad (.10)$$

$$+ .066 \cdot \log(A_{it}/A_{it-1}),$$

$$(.065) \quad (8)$$

$$N = 330, \quad R^2 = .149, \quad \sigma = .164.$$

The estimated coefficient of $\log(A_{it}/A_{it-1})$ is positive but insignificantly different from zero (t -value = 1.0).

⁷ The relation between growth in compensation and performance shown in eqq. (6) and (7) is consistent in a general way with Murphy's (1985) findings for 461 executives in 72 manufacturing corporations. His results for salary and bonus (app. table 9) indicate an estimated coefficient on the contemporaneous stock return of .086, $SE = .009$. He also reports a significant coefficient on the growth rate of real sales: .255, $SE = .023$. We found (see below) that an analogous variable for banks—growth rate of real assets—was insignificant once we held fixed the change in the accounting return, ΔRA_{it} , and the lagged stock return, which is included in $RS2_{it}$.

Since the sample means (for the 330 observations used in eq. [7]) of $RS2_{it}$ and ΔRA_{it} are .176 and $-.035$, respectively, equation (7) implies that real compensation for continuing CEOs grew on average by 9.4% per year (which is the sample mean of $\log(w_{it}/w_{it-1})$ —see App. table A1). From the standpoint of marginal productivity theory, this average growth rate reflects the effects on productivity from greater CEO experience and also from advances in the overall industry and economy.

Since the sample standard deviation of $RS2_{it}$ is .24, the coefficient of .174 in equation (7) means that a one-standard-deviation move in stockholders' returns generates a shift in the annual growth rate of real compensation by 4.1 percentage points—almost half of the sample mean of $\log(w_{it}/w_{it-1})$. With a standard deviation for ΔRA_{it} of .097, a one-standard-deviation change in this variable has a similar quantitative effect on $\log(w_{it}/w_{it-1})$. Thus, executive compensation is highly sensitive to performance.

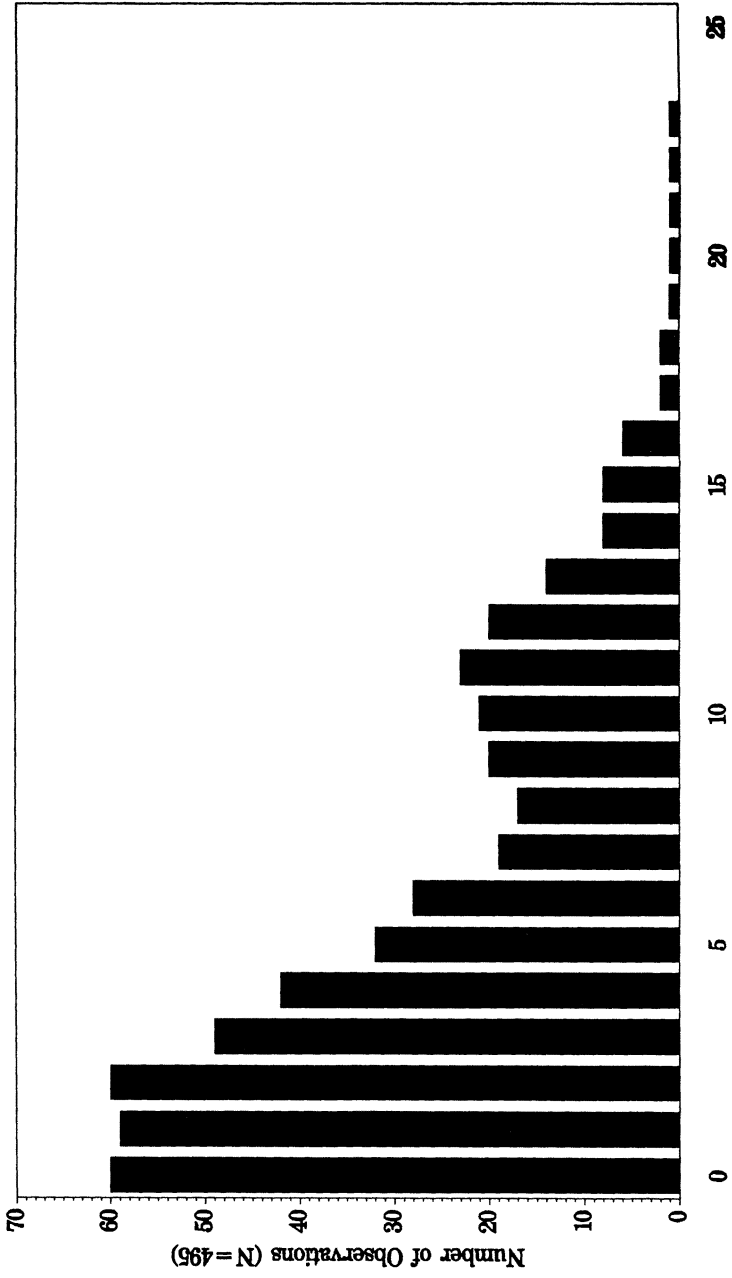
Effects of CEO Experience

The theoretical discussion implied that the response of compensation to performance diminishes in magnitude as experience increases. Among the 330 observations on continuing CEOs, the median years of prior experience as CEO, denoted $EXPER_{it}$, is 4, and the mean is 6.0 with a standard deviation of 4.6. (Note that continuing CEOs must have at least 1 year of experience.) For the full sample of 495 observations—which includes newly hired CEOs as well as cases with missing data on other variables—the median of $EXPER_{it}$ is also 4, and the mean is 5.4 with a standard deviation of 4.7. Figure 1 provides a histogram for the number of observations with each value of $EXPER_{it}$.

We first separated the two performance variables from equation (7)— $RS2_{it}$ and ΔRA_{it} —into observations with $EXPER_{it}$ below and above the median, that is, with $EXPER_{it} \leq 4$ and $EXPER_{it} \geq 5$, respectively. The results for compensation growth are then

$$\begin{aligned} \log(w_{it}/w_{it-1}) = & .083 + .212 \cdot RS2_{it}(EXPER_{it} \leq 4) \\ & (.012) \quad (.043) \\ + & .105 \cdot RS2_{it}(EXPER_{it} \geq 5) + .93 \cdot \Delta RA_{it}(EXPER_{it} \leq 4) \\ & (.060) \quad (.16) \quad (9) \\ & + .29 \cdot \Delta RA_{it}(EXPER_{it} \geq 5), \\ & (.11) \end{aligned}$$

$$N = 330, \quad R^2 = .181, \quad \sigma = .161.$$



Years of Prior Experience as CEO

FIG. 1.—Histogram for CEO experience

As predicted, the change in compensation is more sensitive to performance at lower levels of experience.⁸ A joint test of the hypothesis that the coefficients of each performance variable are the same over the two ranges of experience leads to the statistic $F_{325}^2 = 7.0$, which exceeds the 1% critical value of 4.7. The evidence that the sensitivity attenuates with experience is clearer for accounting-based performance than for market-based performance. For the variable ΔRA_{it} alone, the hypothesis of equality of the coefficients over the two ranges of experience corresponds to the statistic $t_{325} = 3.4$, which is significant at less than the 1% level. For the variable $RS2_{it}$ alone, the corresponding statistic is $t_{325} = 1.7$. This statistic is significant at the 5% level for a one-sided test (coefficient with $EXPER_{it} \leq 4$ greater than that with $EXPER_{it} \geq 5$).

The nonlinear functional form implied by the theory in equation (4) involves the interaction between performance and the term $1/(T_0 + EXPER_{it})$, where T_0 is the effective number of years of prior experience for a CEO in the initial year. We treat T_0 as a constant to be estimated. When the two performance variables— $RS2_{it}$ and ΔRA_{it} —are entered multiplicatively with the term $1/(T_0 + EXPER_{it})$, the maximum-likelihood estimate of T_0 is 2.4 years. The corresponding estimates of the other coefficients are given by

$$\begin{aligned} \log(w_{it}/w_{it-1}) = & .088 + .89 \cdot RS2_{it}/(2.4 + EXPER_{it}) \\ & (.011) \quad (.17) \\ & + 4.22 \cdot \Delta RA_{it}/(2.4 + EXPER_{it}), \\ & (0.67) \end{aligned} \quad (10)$$

$$N = 330, \quad R^2 = .173, \quad \sigma = .161.$$

The 95% confidence interval for the estimate of T_0 —determined by the likelihood ratio and the 5% critical value from the χ^2 distribution—is (0.3, 14). Although this interval is wide, the estimated value of T_0 is significantly positive.⁹ As T_0 becomes large (and the other coefficients adjust accord-

⁸ Similarly, Murphy (1986, table 3) finds that the growth in compensation is more sensitive to market-based performance for CEOs with fewer than 4.6 years of experience than for those with greater than 4.6 years.

⁹ The standard errors shown in eq. (10) are conditional on the point estimate $T_0 = 2.4$. These standard errors are satisfactory for tests of the hypothesis that the coefficients associated with $RS2_{it}$ and ΔRA_{it} are each zero. It is possible to compute the usual asymptotic standard errors for T_0 and the other estimated coefficients based on linearization of the likelihood function about the maximum-likelihood estimates. These values are unsatisfactory because the distribution of the estimates is highly asymmetric, as indicated by the confidence interval for T_0 . For example, the standard error calculated in the usual way for T_0 is 2.0 (coefficient estimate

ingly), the effect of $EXPER_{it}$ in the form of equation (10) becomes unimportant, and the functional form approaches the linear specification (without $EXPER_{it}$) of equation (7). Therefore, $T_0 < \infty$ corresponds to the hypothesis that the sensitivity of compensation to performance diminishes with experience. The value of $-2 \cdot \log(\text{likelihood ratio})$ associated with this hypothesis is 10.6, which exceeds the 1% critical value of 6.6. Therefore, as in the results that considered just two ranges of experience in equation (9), the conclusion is that the effects of performance on compensation change attenuate with experience.¹⁰

The effects of experience in equation (10) can be compared with the results in equation (9) by calculating the implied response of compensation change to performance at various levels of experience. Table 2 shows the response coefficients implied by equation (10) for values of $EXPER_{it}$ between 1 and 20 years. These results are essentially the continuous counterpart of the results from equation (9), which allowed for only two categories of experience.

The growth of compensation may depend on the level of CEO experience, as well as on the interaction between experience and performance. If $EXPER_{it}$ is added to equation (9), its estimated coefficient is $-.0027$, $SE = .0023$. In equation (10), the result is $-.0017$, $SE = .0021$ (using the maximum-likelihood estimate, $T_0 = 2.2$). Hence, the estimated effects of $EXPER_{it}$ are statistically insignificant.¹¹ (The age of the CEO, AGE_{it} , is also insignificant in these regressions.)

Table 2
Implied Response Coefficients for Compensation Growth

$EXPER_{it}$	Response to $RS2_{it}$	Response to ΔRA_{it}
1	.26	1.24
2	.20	.96
5	.12	.57
10	.07	.34
20	.04	.19

NOTE.—The response coefficients are calculated from equation (10) at the indicated values of CEO experience, $EXPER_{it}$.

= 2.4), whereas the confidence interval derived from direct calculation of the likelihood ratio is (0.3, 14).

¹⁰ As in the previous case, the results are clearer for accounting earnings than for stock returns. For ΔRA_{it} alone, the hypothesis of irrelevance of $EXPER_{it}$ leads to a value of $-2 \cdot \log(\text{likelihood ratio})$ of 8.7. For $RS2_{it}$ alone, the corresponding value is 3.3. These values compare to the 5% critical value of 3.8.

¹¹ In some cases, the CEO's initial "year" in office represents a period of less than 12 months. (Recall, however, that we classify an individual as CEO only if

Relative Performance Evaluation

We now consider whether compensation change depends on performance filtered for peer-group performance. We measure peer-group results by the averages for the year and geographical region of stock returns and changes in earnings yield.¹² The regional breakdown used is New England, New York City, Mid-Atlantic, Midwest, South, Texas/Oklahoma, and West—see table 1. There is some arbitrariness in the selection of regions, but the breakdown should capture common regional disturbances.

The sample exhibits significant variation in performance from year to year: over the 1983–87 period (using all available data), the F -values for the joint significance of year dummies are $F_{405}^4 = 22.8$ for $RS2_{it}$ and $F_{385}^4 = 12.6$ for ΔRA_{it} . There is also significant variation in performance across regions within a year. Given year dummies, the F -values associated with dummies for region and year are $F_{375}^{30} = 5.0$ for $RS2_{it}$ and $F_{355}^{30} = 5.6$ for ΔRA_{it} .

Let $RS2_{it}^*$ and ΔRA_{it}^* be the regional averages applicable to bank i in year t . (These averages use all available data and are not limited to the sample of banks included in the regressions.) If these regional averages are added to equation (7), the results are

$$\begin{aligned} \log(w_{it}/w_{it-1}) = & .072 + .146 \cdot RS2_{it} + .45 \cdot \Delta RA_{it} \\ & (.018) \quad (.050) \quad (.12) \\ & + .071 \cdot RS2_{it}^* + .00 \cdot \Delta RA_{it}^*, \\ & (.086) \quad (.20) \end{aligned} \tag{11}$$

$$N = 330, \quad R^2 = .148, \quad \sigma = .164.$$

The coefficients on regional average performance, $RS2_{it}^*$ and ΔRA_{it}^* , are individually and jointly insignificant. Thus, the results indicate that compensation change depends on actual performance without modification for peer-group performance.

he is the highest paid executive for the year.) This initial-year effect means that the measured growth rate of compensation from year 1 to year 2 tends to be higher than otherwise. A dummy variable for exactly 1 year of prior experience as CEO has the estimated coefficient .055, SE = .029. The other results are similar to those shown in (9). Furthermore, the estimated coefficient of $EXPER_{it}$ is still insignificant if the dummy for $EXPER_{it} = 1$ is also included.

¹² Because we are dealing with a single industry, we cannot use industrywide performance as a filter. Average performance for the year is also problematic because it cannot be distinguished from time dummies in an equation for compensation growth. Time dummies are insignificant if added to eqq. (7), (9), or (10).

One possible interpretation of equation (11) is that the regional average values are not very good measures of the general performance that ought to be filtered out of individual performance. That is, if $RS2_{it}^*$ and ΔRA_{it}^* were noise, it would not be surprising that the estimated coefficients of these variables would differ insignificantly from zero. The region-year dummies do have significant explanatory power for the measures of performance; that is, there are significant common influences on banks within regions and years. Also, results reported later show that the probability of CEO turnover relates significantly to the regional average of stock returns, $RS2_{it}^*$, given the actual returns. Hence, these results indicate that $RS2_{it}^*$ does provide useful information.

Equation (11) can be rewritten as a function of relative performance, $RS2_{it} - RS2_{it}^*$ and $\Delta RA_{it} - \Delta RA_{it}^*$, and the regional average values to get

$$\begin{aligned} \log(w_{it}/w_{it-1}) = & .072 + .146 \cdot (RS2_{it} - RS2_{it}^*) + .217 \cdot RS2_{it}^* \\ & (.018) \quad (.050) \qquad \qquad (.070) \qquad (12) \\ & + .45 \cdot (\Delta RA_{it} - \Delta RA_{it}^*) + .45 \cdot \Delta RA_{it}^* \\ & (.12) \qquad \qquad \qquad (.16) \end{aligned}$$

It is clear from equation (12) that the results cannot reject a model where relative and aggregate performance each matter for compensation growth. From this perspective, it just happens that the coefficients of relative and general performance do not differ significantly so that the restricted form where only actual performance matters—equation (7)—is not rejected by the data.

The significance of the regional average values in equation (12) means that the data reject a restricted form in which only relative performance counts. The joint hypothesis for the significance of the two regional average values in equation (12) corresponds to the statistic $F_{325}^2 = 17.7$. As discussed before, the hypothesis that only relative performance matters arises if contractual arrangements fully shield CEO compensation from risks associated with variations in aggregate performance.¹³ It is this proposition that is rejected by the data.¹⁴

¹³ The hypothesis that only $PERF - PERF^*$ matters also requires enough data for each region so that the average value $PERF^*$ measures the aggregate disturbance with negligible error—see Holmstrom (1982). More generally, relative performance evaluation implies that $PERF$ and $PERF^*$ enter separately with coefficients of opposite sign but not necessarily of equal magnitude.

¹⁴ This finding is consistent with cross-industry results of Murphy (1985, table 8) and Jensen and Murphy (1988, table 2). Antle and Smith (1986, tables 4 and 5) find some evidence that accounting-based performance is filtered for industrywide outcomes. However, they find little indication of this filtering for market-based

As it stands, the weaker hypothesis that relative and general performance matter for changes in compensation does not impose restrictions on the data. In particular, the theory that changes in pay correspond to changes in expected marginal product does not dictate the relative magnitudes of the effects of relative and aggregate performance in the form of equation (12). We can generate testable hypotheses by reintroducing the effects of CEO experience. As discussed before, because the information content of an additional observation diminishes as the number of observations rises, the sensitivity of compensation change to *relative performance* falls with experience. However, the effects of aggregate performance on compensation do not interact with experience in this manner. That is, the information content of general performance has nothing to do with the experience of a particular bank's CEO.

Consider the model

$$\begin{aligned} \log(w_{it}/w_{it-1}) = & \beta_0 + \beta_1(\text{PERF}_{it} - \text{PERF}_{it}^*)/(T_0 + \text{EXPER}_{it}) \\ & + \beta_2\text{PERF}_{it}^*/(T'_0 + \text{EXPER}_{it}), \end{aligned} \quad (13)$$

where PERF_{it} refers to RS2_{it} or ΔRA_{it} and PERF_{it}^* to the year/region averages of these variables. The effects of relative performance diminish with experience if $0 < T_0 < \infty$, and the effects of general performance are invariant with experience if $T'_0 \rightarrow \infty$. Note that, as $T'_0 \rightarrow \infty$ (and $\beta_2 \rightarrow \infty$ correspondingly), the final term in equation (13) becomes linear in PERF_{it}^* .

Using RS2_{it} and ΔRA_{it} as measures of PERF_{it} , we fit equation (13) with T_0 unconstrained and with T'_0 unrestricted or set at infinity (in which case the last term is linear in PERF_{it}^*). The value for $-2 \cdot \log(\text{likelihood ratio})$ corresponding to the restriction on T'_0 is 4.7, which exceeds the 5% critical value of 3.8. Therefore, the data indicate that—holding fixed the influence of relative performance as it interacts with experience—the sensitivity of compensation change to aggregate performance diminishes with experience. In fact, the data are consistent with the hypothesis that the interaction with experience is the same for relative and general performance—that is, $T_0 = T'_0$ in equation (13). The value of $-2 \cdot \log(\text{likelihood ratio})$ corresponding to this restriction is only 0.3.

performance. In a recent study, Gibbons and Murphy (1989, tables 1, 2) report more support for the idea that individual stock returns are filtered for overall market returns in the determination of changes in CEO compensation. The results are difficult to interpret because overall market returns matter whereas various definitions of industrywide returns do not—from an informational standpoint, the industry returns seem to be more relevant.

These findings on the interaction between aggregate performance and experience are not favorable to the theory of relative performance evaluation based on incomplete information about CEO skill. The results can, however, be rationalized by arguing that the sensitivity of CEO productivity to aggregate disturbances depends on experience for reasons that do not involve informational considerations. Under this interpretation, the acceptance of the hypothesis that $T_0 = T'_0$ in equation (13) would have to be viewed as a coincidence.

The Relation between Compensation and Assets

Equation (5) dealt with the relation between levels of compensation and assets for CEOs in their first year in office. The correlation between the logarithms of real compensation and real assets for this group of 60 CEOs is .79. For CEOs who continue in office, the growth of compensation depends on performance—stock returns and accounting earnings interacting with experience—as discussed before. As also noted before (eq. [8]); the growth of compensation is not significantly related to the growth in assets, given the performance measures. Because asset growth is only weakly correlated with the performance variables,¹⁵ the correlation between the logarithms of compensation and assets tends to diminish as experience increases. That is, the match between the CEO's perceived talent (as reflected in compensation) and the size of the bank worsens with tenure.

Table 3 and figure 2 show the correlation between the logarithms of

Table 3
Correlation between the Logarithms of Compensation and Assets as a Function of CEO Experience

EXPER _{it}	Number of Observations	Correlation
0	60	.79
1	58	.75
2	58	.63
3	49	.64
4	41	.52
5	29	.73
6-7	45	.88
8-9	37	.86
10-11	42	.79
12-13	34	.75
≥14	42	.82
All	495	.74

NOTE.—The correlation is between $\log(w_{it})$ and $\log(A_{it})$ at the indicated values of CEO experience, EXPER_{it}.

¹⁵ For 330 continuing CEOs, the correlation of $\log(A_{it}/A_{it-1})$ is .17 with RS_{2it} and .10 with ΔRA_{it} .

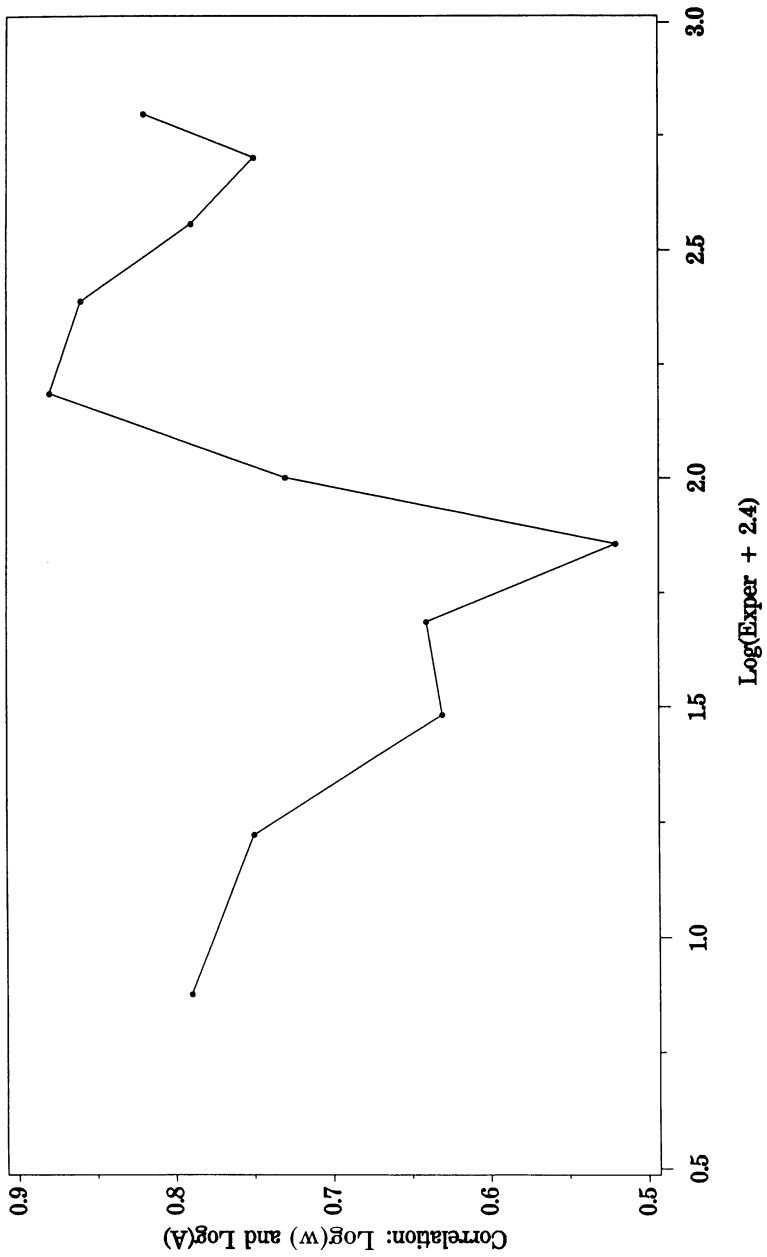


FIG. 2.—Correlation between pay and assets

real compensation and real assets as a function of CEO experience.¹⁶ As anticipated, the correlation declines from .79 at $\text{EXPER}_{it} = 0$ to .52 at $\text{EXPER}_{it} = 4$, which is the median years of experience in the sample.¹⁷ The correlation then rises, however, to .73 at $\text{EXPER}_{it} = 5$ and to values averaging .82 for $\text{EXPER}_{it} \geq 6$. In other words, after worsening initially, the match between CEO pay (perceived talent) and bank size improves as experience rises above the median.

The theory sketched at the beginning of this article, based on Rosen (1982), indicated why it would be beneficial to have a good match between CEO talent and bank size. Despite this benefit, the match deteriorates as experience rises because it is costly to adjust bank assets or to change the identity of the CEO in response to information about the CEO's talent. However, new information about talent is reflected quickly in executive pay. If compensation and assets get far out of line, the bank is motivated to make adjustments in assets or in the identity of the CEO.

The nature of the adjustment of assets to performance shows up in the regression,

$$\begin{aligned} \log(A_{it}/A_{it-1}) = & .077 + .125 \cdot \text{RS}_{it-3} + .37 \cdot \Delta\text{RA}_{it-1} \\ & (.010) \quad (.031) \quad (.11) \\ & + .29 \cdot \Delta\text{RA}_{it-2} + .123 \cdot \log(A_{it-1}/A_{it-2}), \quad (14) \\ & (.15) \quad (.054) \end{aligned}$$

$$N = 389, \quad R^2 = .088, \quad \sigma = .147,$$

where A_{it} is again the real assets of bank i . Although the third lag RS_{it-3} is significantly positive, as shown in equation (14), the earlier lags, RS_{it-1} and RS_{it-2} , are insignificant if added to the regression. Equation (14) shows that the growth of assets also relates significantly to two lags of ΔRA , as well as to the previous year's growth in assets. The general inference from these results is that asset growth responds to performance, but at substantially longer lags than those applicable to compensation growth. This behavior helps to explain why the logarithms of compensation and assets become less correlated over a range of CEO experience—0–4 years in table

¹⁶ The horizontal axis in fig. 2 plots $\log(\text{EXPER}_{it} + 2.4)$, where 2.4 is the estimate of T_0 from before. Although this specification provides a good illustration of the data, we are unsure about the proper functional form for the relation between the correlation and experience.

¹⁷ As mentioned in n. 11 above, the initial year in office corresponds in some cases to less than 12 months. Variations in the number of months tends to worsen the correlation between measured annual pay and assets. Nevertheless, the correlation in the first year exceeds that in the second year.

3 and figure 2—in which compensation adjusts more readily than assets to performance. The eventual adjustment of assets to performance tends, however, to raise the correlation between compensation and assets at higher levels of experience.

The other element that influences the correlation between compensation and assets is the selection of the sample as experience rises. In particular, when performance is especially bad—so that compensation becomes unusually low in relation to assets—the CEO is likely to be dismissed. This truncation of the sample tends to raise the correlation between compensation and assets among the CEOs that remain. To allow for this effect, we now consider CEO turnover.

CEO Turnover

If relative performance is weak and the perceived skill of the CEO is therefore less than expected initially, the bank may discharge the CEO instead of lowering pay or allowing assets to decline to match the level of skill. Dismissal avoids the costs of having a poor match between CEO skill and bank size or the costs of shrinking the bank, but it introduces costs associated with CEO turnover. These costs include the loss of specific capital associated with the incumbent CEO.

Given observed performance for T years, the bank directors estimate the CEO's skill to be $E(k_i)|_T$ as in the model that led to equation (4). The CEO is dismissed if $E(k_i)|_T$ falls below a critical value, which depends on T and the other parameters of the model. Other things being equal, a higher critical value is more likely to result in CEO dismissal. Since the variance of $E(k_i)|_T$ about the true value k_i declines with T —that is, with more information—the critical value for dismissal tends to rise with T . A high critical value for CEOs with little experience is undesirable because it results in a high frequency of CEO turnover and hence in high adjustment costs.

Recall that the sensitivity of $E(k_i)|_T$ to relative performance declines with T ; this result implies that the responsiveness of CEO pay to relative performance diminishes with experience. However, since the critical value of $E(k_i)|_T$ for dismissal rises with T , the net effect of experience on the linkage between CEO turnover and relative performance is ambiguous. Unlike the case of compensation change, the theory does not imply that the sensitivity of turnover to performance declines with experience.

The other aspect of CEO turnover that differs from compensation change concerns aggregate performance variables. Aggregate disturbances can affect values of marginal products of individual CEOs and thereby influence CEO compensation. In contrast, for banks that stay in business, the decision to dismiss a CEO is based on the desire to replace the existing head with someone else. Hence the probability of termination depends on relative performance and not on aggregate performance. Therefore, although pure

relative performance evaluation was rejected for the growth of compensation, it is interesting to reexamine this hypothesis in the context of CEO dismissal.

Table 4 shows logit regressions for CEO turnover.¹⁸ The dependent variable equals one if the CEO is present in year $t - 1$ but not in year t and equals zero if the CEO is in office in both years. The data do not allow us to condition departure on “reasons” such as death or illness, fires versus quits, and so on. The right side of the equation takes the form $\exp(a + bx_{it})/[1 + \exp(a + bx_{it})]$, where x_{it} is a vector of explanatory variables.

Three regressors capture the effects of the CEO’s age. The variable AGE_{it-1} is the age of the CEO in year $t - 1$ —that is, in the final year in office for departing CEOs. The data come from proxy statements that typically indicate the CEO’s age in February or March of year t . We took these numbers as measures of AGE_{it-1} , that is, as the age attained by the end of the previous year. The variable $AGESQ_{it-1}$ —the square of AGE_{it-1} —allows for additional curvature in the relation between probability of departure and age.

For many CEOs, 65 is viewed as the “normal” retirement age. Given the nature of the data, this normal behavior could correspond to AGE_{it-1} falling in a range from 63 to 66. That is, a CEO with $AGE_{it-1} = 63$ could be 64 during most of the final year in office, and one with $AGE_{it-1} = 66$ could be 65 during most of the final year. Hence we included the dummy variable $DUM6366$, which equals one if AGE_{it-1} is between 63 and 66 and zero otherwise.

Figures 3 and 4 show the numbers of departing and continuing CEOs, respectively, at various ages (AGE_{it-1}). For departing CEOs ($N = 51$), the mean age is 60.1 (SD = 6.5), and the median is 63; for continuing CEOs ($N = 407$), the mean is 55.4 (SD = 5.7), and the median is 56. Among the 51 departing CEOs,¹⁹ 27 had ages between 63 and 66.

The other explanatory variables are the performance measures used before. Column 1 of table 4 includes, aside from the age variables, only the 2-year average stock return measured relative to the region/year average, $RS2_{it} - RS2_{it}^*$. The estimated coefficient of this performance variable (-7.2 , $SE = 1.5$) is negative and significant—meaning that better relative performance as measured by the stock market reduces the probability of CEO turnover. We consider the effects of performance further below.

Each of the three age variables are statistically significant. The estimated

¹⁸ The results are essentially the same with a probit formulation. Previous studies that fit logit models for CEO turnover include Coughlan and Schmidt (1985), Warner, Watts, and Wruck (1988), and Weisbach (1988).

¹⁹ This number differs from the 60 new CEOs in the sample because of missing data on AGE_{it-1} or on other variables that enter into the logit regressions.

Table 4
Logit Regressions for CEO Turnover

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Number of observations	458	458	446	458	446	388	378	458	446
Depart; stay*	51; 407	51; 407	50; 396	51; 407	50; 396	41; 347	40; 338	51; 407	50; 396
Constant	20.5 (8.7)	20.5 (8.7)	19.7 (8.8)	20.7 (8.7)	20.2 (8.8)	18.2 (9.6)	16.1 (10.0)	20.7 (8.7)	19.8 (8.8)
AGE _{<i>t</i>-1}	-91 (.32)	-91 (.32)	-88 (.32)	-91 (.32)	-90 (.32)	-84 (.35)	-77 (.37)	-92 (.32)	-88 (.32)
AGESQ _{<i>t</i>-1}	.0087 (.0029)	.0087 (.0029)	.0084 (.0030)	.0087 (.0029)	.0085 (.0030)	.0081 (.0032)	.0076 (.0033)	.0088 (.0030)	.0085 (.0030)
DUM6366 _{<i>t</i>-1}	1.87 (.47)	1.88 (.47)	1.88 (.47)	1.86 (.47)	1.85 (.47)	1.59 (.51)	1.56 (.51)	1.87 (.47)	1.88 (.47)
RS2 _{<i>t</i>-1} - RS2 _{<i>t</i>-1}	-7.2 (1.5)	...	-7.1 (1.5)	-6.6 (1.7)	-5.7 (1.7)
RS _{<i>t</i>-1} - RS _{<i>t</i>-1}	...	-3.7 (.9)
RS _{<i>t</i>-2} - RS _{<i>t</i>-2}	...	-3.3 (1.1)
RS _{<i>t</i>-3} - RS _{<i>t</i>-3}	-7 (1.3)	-1.8 (1.6)
RS2 _{<i>t</i>-1}	-7.5 (1.6)	-7.4 (1.6)

$RS2_{it-1}^*$	6.1 (1.9)
$\Delta RA_{it-1} - \Delta RA_{it-1}^*$...	-2.0 (3.4)	-4.5 (4.0)
ΔRA_{it-1}
ΔRA_{it-1}^*
$\Delta RA_{it-2} - \Delta RA_{it-2}^*$	-7.7 (5.2)
$RS2_{it-1} - RS_{it-1}^*$ [EXPER $_{it-1} \leq 4$]	-7.6 (2.3)	-7.4 (2.3)
[EXPER $_{it-1} \geq 5$]	-7.0 (1.9)	-6.9 (1.9)
$\Delta RA_{it-1} - \Delta RA_{it-1}^*$ [EXPER $_{it-1} \leq 4$]	-1.9 (5.8)
[EXPER $_{it-1} \geq 5$]	-2.0 (4.2)

NOTE.—Asymptotic standard errors are shown in parentheses. The dependent variable equals one if the CEO departs in year t and equals zero if the CEO continues. The independent variables are described in the text and in App. Table A1.

* Number who depart and stay in each sample.

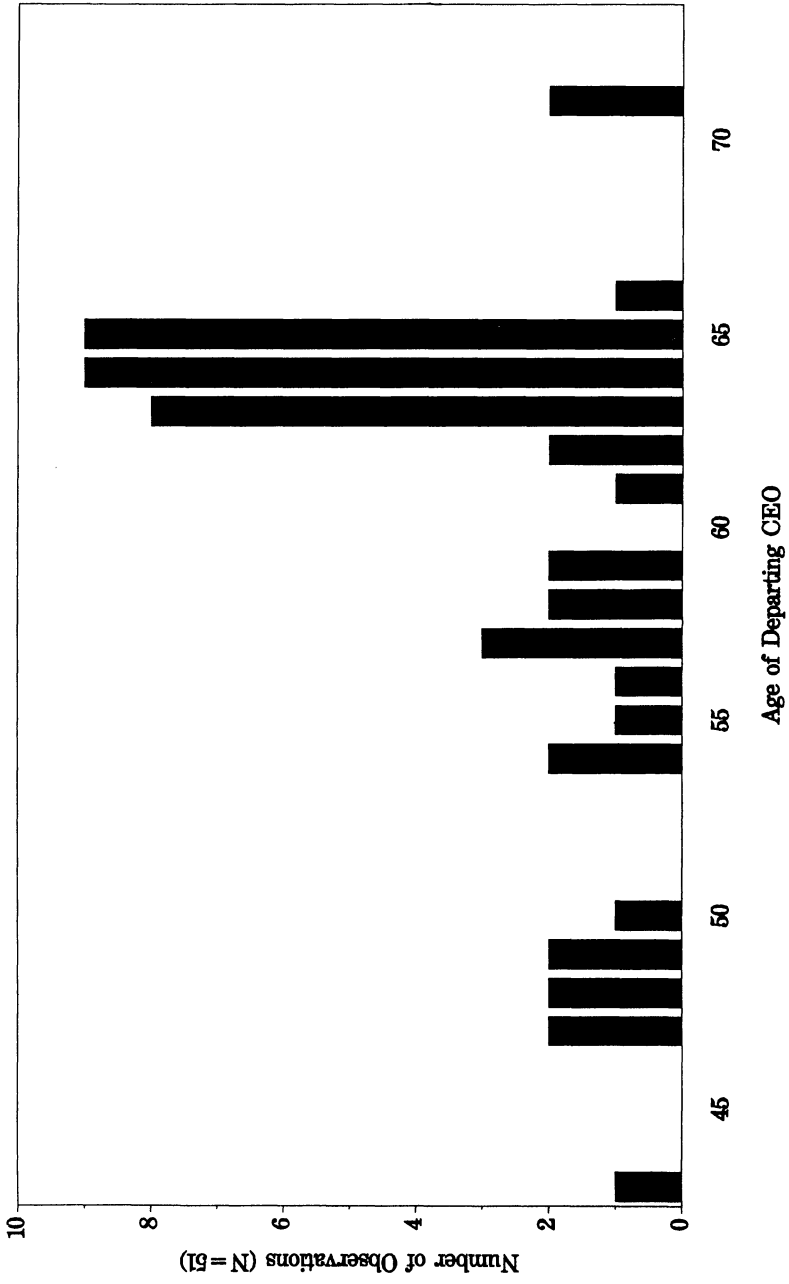


FIG. 3.—Histogram for ages of departing CEOs

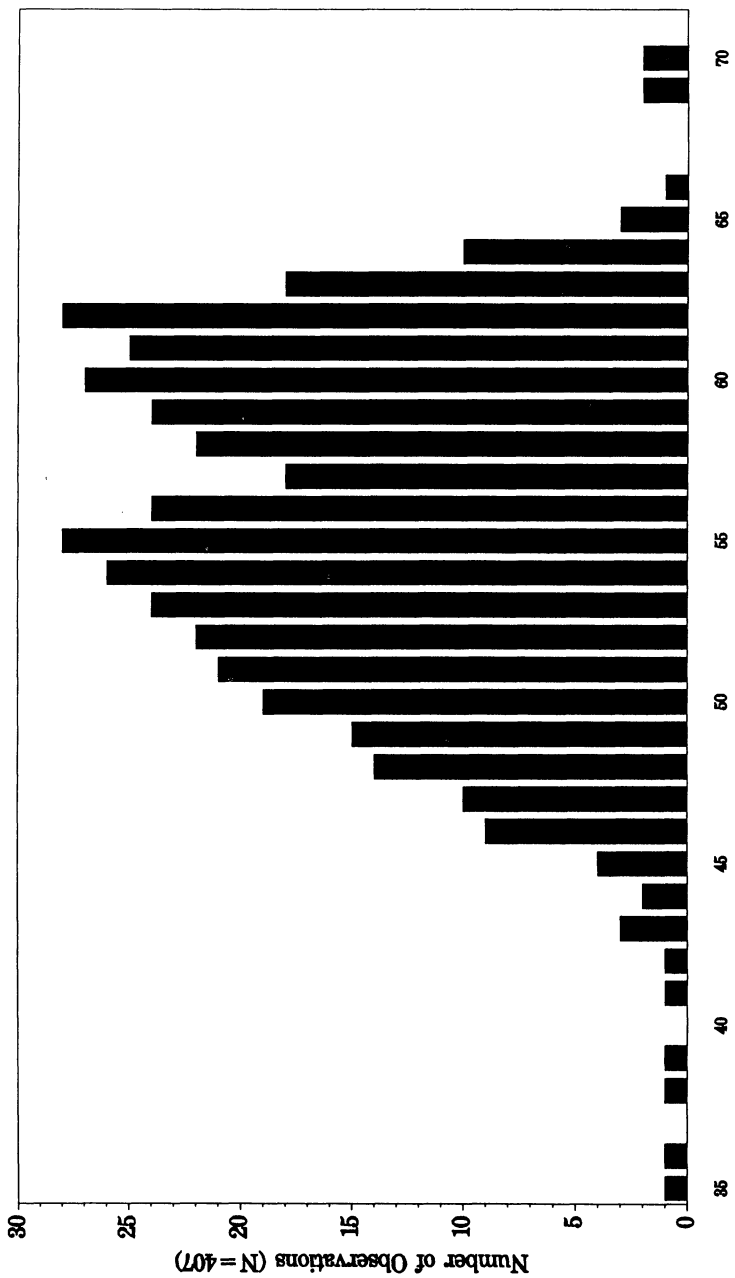


FIG. 4.—Histogram for ages of continuing CEOs

coefficients (col. 1 of table 4) of $-.91$ ($SE = .32$) on AGE_{it-1} and $.0087$ ($SE = .0029$) on $AGESQ_{it-1}$ imply that the probability of departure falls with age for $AGE_{it-1} \leq 52$, and rises with age beyond that. The estimated coefficient on $DUM6366$ of 1.87 ($SE = .47$) means that, given the other effects of age, there is an especially high probability of departure at the normal retirement ages between 63 and 66.

The solid line in figure 5 shows the estimated probability of CEO departure as a function of age for a CEO with average stock-market performance ($RS2_{it} - RS2_{it}^* = 0$). The estimated values come from the regression shown in column 1 of table 4. The dotted line in the figure is the frequency of departures found in the sample at the various ages. Note, however, that the numbers of observations are small at the extremes of young and high ages. The frequency shown in the figure is a 3-year moving average of observed values (number of CEOs departing relative to the total number in the age group) for ages between 41 and 65. The value shown for 66 is for that age only, and the constant value shown for 67–71 is the average of values over that entire interval (consisting of two observations each for 69, 70, and 71, and zero for 67 and 68). The dramatic rise in departure probability around age 63 is clear in the data. Whether there is a fall in the probability after age 66 (and then a subsequent rise with age) is unclear because of the small number of observations in that interval. A dummy variable for $AGE_{it-1} \geq 67$ is, however, insignificant if added to the regression in column 1 of table 4 (estimated coefficient of 0.3 , $SE = 1.7$).

We tested the hypothesis that the coefficients of AGE_{it-1} , $AGESQ_{it-1}$, and $RS2_{it} - RS2_{it}^*$ were the same over the age range 63–66 as for all other ages. (Given the small number of observations, it does not matter which group contains the values with ages above 66.) The test statistic is $-2 \cdot \log(\text{likelihood ratio}) = 4.2$, which is less than the 5% critical value of 7.8. Therefore, a different intercept (the variable $DUM6366$) is sufficient to account for the differing behavior around the normal retirement age of 65. In particular, the estimated coefficient of $RS2_{it} - RS2_{it}^*$ is significantly negative when estimated only over the subsample of CEOs aged between 63 and 66 (59 observations, of which 27 are of CEO departures). The estimated coefficient for this subsample is -7.3 , $SE = 3.2$. Thus, even around age 65, CEOs who perform better are significantly less likely to depart.²⁰

Consider now further aspects of the relation between CEO departure and performance. Table 5 shows the nature of the relation between turnover frequency and relative stock returns in the underlying data. The table re-

²⁰ Coughlan and Schmidt (1985, table 6) find a significantly negative effect of abnormal stock returns on the probability of CEO departure for CEOs aged ≤ 63 , but not for older CEOs. Similarly, Weisbach (1988, table 3) reports significant effects only for CEOs aged ≤ 63 or ≥ 67 .

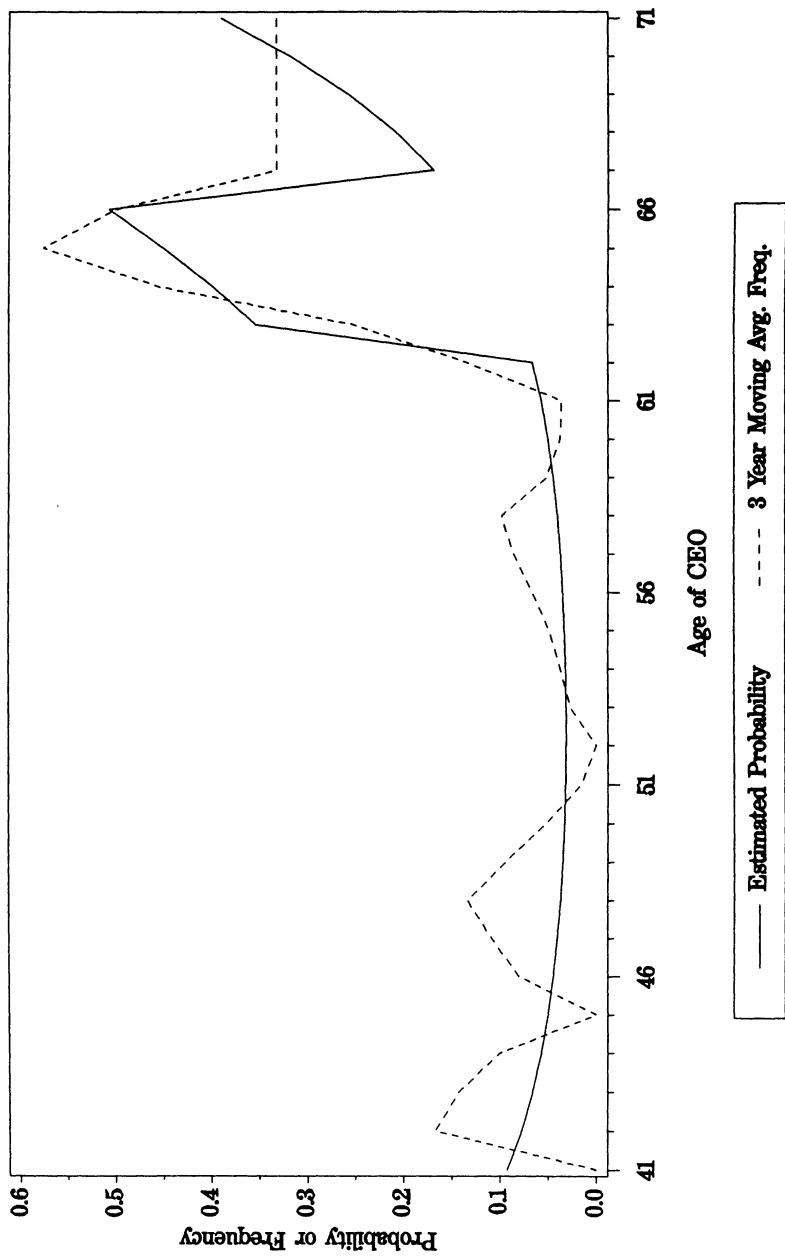


FIG. 5.—CEO departure probability and frequency

Table 5
Frequency of CEO Departure

Range of $RS2_{it-1} - RS2_{it-1}^*$	Range of AGE_{it-1}			All
	≤ 55	56-62	≥ 63	
$> .08$.020(51)	.000(49)	.100(10)	.018(110)
$(0, .08)$.018(57)	.022(45)	.333(15)	.060(117)
$(-.08, 0)$.055(55)	.106(47)	.600(20)	.164(122)
$< -.08$.118(51)	.132(38)	.550(20)	.202(109)
All	.051(214)	.061(179)	.446(65)	.111(458)

NOTE.—The entry in each cell is the ratio of CEO departures to the total number of CEO observations. The number in parentheses is the number of observations.

ports the annual frequency of CEO departures for categories of ages (≤ 55 , between 56 and 62, and ≥ 63) and relative returns, $RS2_{it-1} - RS2_{it-1}^*$ ($> .08$, $(0, .08)$, $(-.08, 0)$, and $< -.08$). Since the mean of $RS2_{it-1} - RS2_{it-1}^*$ is close to zero and the standard deviation, σ , is .16, the categories for returns are $> \sigma/2$, $(0, \sigma/2)$, $(-\sigma/2, 0)$, and $< -\sigma/2$. The table shows a marked tendency in all age ranges for the departure frequency to rise as the stock return worsens. For example, for all ages combined, the frequency goes from .02 for returns above $\sigma/2$ to .06 between zero and $\sigma/2$, .16 between zero and $-\sigma/2$, and .20 below $\sigma/2$. (For the entire sample, the average annual departure frequency is .11.)

Column 2 of table 4 allows for separate coefficients on $RS_{it-1} - RS_{it-1}^*$ and $RS_{it-2} - RS_{it-2}^*$ in the logit regression. As in the case of compensation growth, the two estimated coefficients are nearly equal, so that the 2-year average variable $RS2_{it} - RS2_{it}^*$ —used in column 1 of the table—is satisfactory.

Column 3 of the table adds the region-relative change in the earnings yield, $\Delta RA_{it-1} - \Delta RA_{it-1}^*$. The estimated coefficient is negative but insignificant: -2.0 , $SE = 3.4$. Thus, unlike compensation growth, the probability of CEO departure is not significantly related to accounting-based performance.

Column 4 separates the stock-market performance into the actual return, $RS2_{it-1}$, and the regional average return, $RS2_{it-1}^*$. The estimated coefficients are opposite in sign and of similar magnitude: -7.5 (1.6) and 6.1 (1.9), respectively. A test of the pure relative performance hypothesis—that the coefficients are of equal magnitude but opposite in sign—leads to the value of $-2 \cdot \log(\text{likelihood ratio})$ of 1.0, which is well below the 5% critical value of 3.8. Thus, pure relative performance evaluation is accepted here.²¹

²¹ Warner et al. (1988, table 5) report a logit regression for the probability of CEO change. They find that the coefficient of the contemporaneous own stock return is negative and that of the market return is positive and of comparable magnitude. The results involving lagged returns are less clear. They do not provide formal tests of the hypothesis of pure relative performance evaluation. Gibbons

The acceptance of the hypothesis of relative performance evaluation for the probability of CEO departure contrasts with the results for compensation growth. Thus, the indication is that CEO turnover depends on relative performance, whereas compensation growth depends on relative and aggregate performance. These results are consistent with the theory in which compensation growth corresponds to the change in expected marginal product, but turnover involves a comparison of the existing CEO with alternative executives.

Column 6 of table 4 shows that an additional lag of the stock-return variable, $RS_{it-3} - RS_{it-2}^*$, is insignificant. Hence the main response of CEO turnover to market-based performance occurs over a 2-year period. Column 7 adds another lag of accounting-based performance. The estimated coefficient of $\Delta RA_{it-2} - \Delta RA_{it-2}^*$ is negative but insignificant: -7.7 , $SE = 5.2$. The introduction of this second lag raises the magnitude of the estimated coefficient of the first lag, $\Delta RA_{it-1} - \Delta RA_{it-2}^*$, to -4.5 , $SE = 4.0$. The two lags of accounting-based relative performance are, however, jointly insignificant: the value of $-2 \cdot \log(\text{likelihood ratio})$ is 2.6, which is below the 5% critical value of 6.0. Hence the conclusion again is that the probability of CEO turnover does not relate significantly to accounting-based performance.²²

Our finding is that market- and accounting-based performances are each important for compensation growth, whereas only the market-based measure is significant for turnover probability. A possible explanation involves the idea of Gibbons and Murphy (1988) that accounting earnings are prone to manipulation by the CEO in the short run. For CEOs who are close to the margin of being dismissed—because they have performed badly—the horizon is short, and the incentive to manipulate the accounting numbers is great. For this reason, a decision to terminate the CEO gives little weight to accounting earnings and relies instead on stock returns or other data that are relatively immune from manipulation.

Column 8 of table 4 divides $RS_{it-1} - RS_{it-1}^*$ into ranges in which CEO experience is below or above the median ($EXPER_{it} \leq 4$ and ≥ 5 , respectively). The estimated coefficients in the two ranges are very close; therefore the results are consistent with the hypothesis that the sensitivity of turnover to performance is independent of experience. (The test statistic is 0.1 with a 5% critical value of 3.8.) This result on experience is another contrast with the findings for compensation growth; in that case theoretical reasoning and empirical evidence showed that compensation change was more sensitive to performance at lower levels of experience. For CEO turnover,

and Murphy (1989, table 6) report results that are similar to those of Warner et al.

²² In contrast, Weisbach (1988, table 5) reports a significantly negative relation between the probability of CEO turnover and changes in accounting earnings, given the behavior of stock returns.

the theoretical effect of experience is ambiguous, and the empirical effect turns out to be indistinguishable from zero.

It is possible that CEO turnover reflects mismatches in either direction—the CEO is either too bad or too good for the bank—rather than poor performance, *per se*. If mismatches in either direction are important, the probability of CEO departure would rise with the magnitude of relative performance. We added the absolute value of $RS2_{it-1} - RS2_{it-1}^*$ to the regression in column 1 of table 4. The estimated coefficient of this absolute value has the “wrong” sign and differs insignificantly from zero: -3.4 ($SE = 3.4$), whereas that of the algebraic value, $RS2_{it-1} - RS2_{it-1}^*$, remains significantly negative: -9.2 ($SE = 2.7$). Thus the results indicate that CEOs who perform much better than expected are especially likely to remain with the bank, rather than tending to move to another (larger) bank that is a better match for their unexpectedly high skill. One reason that this type of move tends not to occur is that the match between CEO talent and bank size can be improved by expanding the size of the bank, as in equation (14).

Given the results shown in table 4, the main effects of performance on the probability of CEO turnover are captured by the logit regression in column 1, which includes $RS2_{it-1} - RS2_{it-1}^*$ as the only performance variable. To evaluate the performance effects quantitatively, note that the logit form implies that the derivative of the logarithm of the departure probability with respect to the relative stock return is $\beta(1 - p)$, where β is the regression coefficient (-7.2) and p is the probability of departure.²³ For example, if $RS2_{it-1} - RS2_{it-1}^* = 0$, the derivative at age 55 is -7.0 , which means that an increase by .01 in the stock return reduces the departure probability by 7%—from .033 to .031. At age 65, the derivative is -4.0 , so that an increase by .01 in the return lowers the probability by 4%—from .45 to .43.

Table 6 shows the estimated probability of departure (based on the logit regression in col. 1 of table 4) at ages 50, 60, and 65 and for relative stock returns between .32 and $-.32$. Since the sample standard deviation of $RS2_{it-1} - RS2_{it-1}^*$ is .16, the range for relative returns is a two-standard-error band about the mean. The estimated values in table 6 are basically the fitted values corresponding to the observed frequencies of departures that were shown before in table 5.

Summary of Major Findings and Conclusions

We studied compensation for bank CEOs by examining the match between levels of pay and bank size for newly hired chief executives. The elasticity of about one-third for compensation in relation to assets is in line with previous estimates for other industries and time periods. For CEOs who continue in office, the growth of compensation varies positively

²³ This result follows from the formula $p = \exp(\alpha + \beta r) / [1 + \exp(\alpha + \beta r)]$, where r represents the relative stock return.

Table 6
Estimated Probability of CEO Departure

RS2 _{<i>t</i>-1} - RS2* _{<i>t</i>-1}	AGE _{<i>t</i>-1}		
	50	60	65
.32	.003	.005	.075
.24	.006	.009	.126
.16	.011	.016	.205
.08	.019	.029	.315
.00	.033	.050	.450
-.08	.057	.086	.594
-.16	.097	.144	.723
-.24	.161	.230	.823
-.32	.255	.348	.892

NOTE.—Each entry shows the estimated probability of CEO departure at the indicated values of AGE_{*t*-1} and RS2_{*t*-1} - RS2*_{*t*-1}, based on the logit regression in column 1 of table 4. The range of values shown for RS2_{*t*-1} - RS2*_{*t*-1} is a two-standard-error band about the mean of zero.

with performance measures based on stock returns and accounting earnings. The sensitivity of compensation change to performance declines significantly as CEO experience increases. We interpreted this effect in terms of the declining information content of additional observations on performance.

There is no indication that individual bank performance is filtered for regional average performance in the relation with compensation growth; in particular, the data reject the hypothesis that only relative performance affects the change in compensation. The results are consistent with a theory in which the growth in pay equals the growth in expected marginal product; in this case, CEO pay responds to relative and aggregate performance. In contrast, the findings are inconsistent with the existence of agreements that fully shield CEO compensation from aggregate risks.

Since compensation growth reacts to stock returns and accounting earnings, but not to growth in assets, the correlation between the levels of compensation and assets—which reflects the match between the quality of the CEO and the size of the organization—tends to worsen as tenure increases. Empirically, this correlation declines as experience rises from 0 to 4 years (the sample median) but subsequently increases to a level comparable to that for new CEOs. One mechanism that raises the correlation at higher levels of experience is the lagged response of bank assets to performance. Another element is the truncation of the sample, via CEO departure, to eliminate the executives whose performance is especially bad.

We estimated logit regressions to relate the probability of CEO departure to age and performance. The probability of departure rises with age (for ages above the early 50s) and becomes particularly high in the normal retirement span around age 65. Even around age 65, the probability of departure declines significantly with better performance.

The main findings for the relation between CEO turnover and performance are that, first, the departure probability falls significantly with stock returns but not with accounting earnings; second, the effects of stock returns enter relative to regional average returns; and third, the sensitivity of departure probability to stock returns does not vary significantly with CEO experience. The potential for manipulation of the accounting results may explain why accounting-based performance is unimportant for turnover but is significant for compensation growth (CEOs who are close to the margin of termination have short horizons and are therefore more likely to engage in earnings manipulation). The success of relative performance evaluation in the context of CEO turnover accords with a model in which dismissal involves a comparison of the incumbent with alternative chief executives. This model is also consistent with the result that the sensitivity of CEO departure to performance does not vary systematically with experience.

Appendix

Table A1
Means and Standard Deviations of Variables

	Sample of New CEOs (<i>N</i> = 60)	Sample of Continuing CEOs (<i>N</i> = 330)	Sample for Logit Regressions		
			All (<i>N</i> = 458)	Depart (<i>N</i> = 51)	Stay (<i>N</i> = 407)
w_{it}	.430(.198)	.497(.197)
Dw_{it}094(.177)
A_{it}	19.6(34.0)	15.4(22.8)
DA_{it}101(.140)
RS_{it}^a134(.332)	.193(.300)	.067(.307)	.208(.295)
$RS_{it} - RS_{it}^*{}^a$002(.255)	.010(.165)	-.083(.115)	.022(.167)
RA_{it}^a121(.106)	.158(.083) ^b	.141(.163) ^c	.160(.066) ^d
ΔRA_{it}^a	...	-.035(.097)	-.019(.071) ^b	-.029(.125) ^c	-.018(.061) ^d
$\Delta RA_{it} - \Delta RA_{it}^*{}^a$002(.075)	-.002(.053) ^b	-.008(.070) ^c	-.002(.050) ^d
$EXPER_{it}$	0	6.0(4.6)
AGE_{it}^a	52.9(6.6)	56.5(5.7)	55.9(6.0)	60.1(6.5)	55.4(5.7)

NOTE.—Standard deviations are shown in parentheses. The logit sample refers to the regressions in table 4. The total of 458 observations breaks down into 51 CEOs who depart and 407 who stay. The symbol *D* on a variable denotes the growth rate from $t - 1$ to t ; w is real compensation in millions, A is real assets in billions, RS is the real rate of return on stocks, RA is the real earnings yield, ΔRA is the change in RA from $t - 1$ to t , $EXPER$ is prior years as CEO, and AGE is the age indicated on the next proxy statement (usually February or March).

^a Values for the logit sample refer to year $t - 1$.

^b $N = 446$.

^c $N = 50$.

^d $N = 396$.

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