The Determinants of Executive Salaries: An Econometric Survey
Author(s): David H. Ciscel and Thomas M. Carroll
Source: *The Review of Economics and Statistics*, Vol. 62, No. 1 (Feb., 1980), pp. 7-13
Published by: The MIT Press
Stable URL: http://www.jstor.org/stable/1924267
Accessed: 16/07/2013 19:24

THE DETERMINANTS OF EXECUTIVE SALARIES: AN ECONOMETRIC SURVEY

David H. Ciscel and Thomas M. Carroll*

FOR over three decades, debate has raged over the economic assumption that the large corporation, through the decisions of its managers, attempts to maximize its profits. Empirical analysis of the behavior of the corporation has led to conflicting claims. The inquiry into the determinants of executive compensation has been no exception. Statistical investigation of executive compensation has been dominated by a search for one decisive explanation. Is the size, measured by either sales or assets, or profitability, measured by net corporate income or by the rate of return on assets, the key variable in establishing the level of the executive's reward? Proponents on both sides of this issue—the managerialists who support the corporate growth hypothesis and the neoclassical economists who favor the profit maximization assumption—seem to argue that the contest can be resolved by the presentation of unambiguous evidence that will award victory to one side and vanquish the other.

This spirit of antagonism has distorted the essential element of the executive compensation question. The behavior of the corporation and the market forces that shape this behavior can be explained or illustrated only by the use of a series of intercorrelated variables. Not one of the available measures of corporate success, be it net income, sales or assets, is an exact measure of economic profits or firm size, nor is it independent of the other variables. This study focuses on the resolution of the serious econometric problems encountered in the process of estimating the determinants of executive compensation. Later we will show how the successful elimination of problems of simultaneous equations bias, multicollinearity and heteroscedasticity leads to the conclusion that the managerialist and neoclassical models of the firm are complementary, rather than substitute, explanations for the pattern of executive compensation.

I. Previous Research

The controversy over the determination of the level of executive compensation has developed as a corollary to the overall neoclassical/managerialist debate about the pattern of corporate behavior. Initially inspired by a comment by Baumol that "executive salaries appear to be more closely correlated with the scale of operations than with its, the firm's, profitability" (1959, rev. ed. 1967, p. 46), numerous studies have now appeared on one side of the debate or the other. The managerialists opened with their findings. McGuire, Chiu, and Elbing (1962), building on the research of Patton (1961) and Roberts (1959), argued that executive compensation is causally related to corporate sales. They developed correlation coefficients for the relationship between executive compensation (measured as annual salary plus bonus) and sales or net profit using several statistical methods: simple, lagged, first differences, and partial correlations. In almost every instance, they found that the managerialist hypothesis is most compatible with the data.

The managerialist position was rebutted by several studies initiated by Lewellen. In two early studies (1968, 1969), he argued that the level of executive earnings was closely tied to the financial fortunes of the corporation through ownership by, or profit-related deferred income payments to, the executive. In a final longitudinal study of several sectors of the economy, Lewellen (1971) found that so much of the corporate executive's yearly income can be traced to property income that the interest of managers and owners could not be differentiated. In a similar, although somewhat more sophisticated analysis, Masson (1971) confirmed the executive income/common stock performance hypothesis introduced by Lewellen.

However, the work of McGuire et al. was not directly answered until the publication of Lewel-

Received for publication February 13, 1978. Revision accepted for publication February 28, 1979.
* Memphis State University.
The authors express their appreciation to Ms. Patricia Byrne and Ms. Virginia Weymouth of Memphis State University for statistical assistance and to two anonymous referees for this REVIEW for helpful comments.

len and Huntsman's (1970) examination of the functional relationship between executive compensation and several measures of corporate performance, e.g., sales, assets, profits, rate of return. Two features of their analysis have persisted in all other studies of executive compensation. First, Lewellen and Huntsman found that a simple measure of executive compensation (salary plus bonus) served as an excellent proxy for total remuneration (salary, bonus plus all deferred and property income). Second, they explicitly recognized the problems of multicollinearity and heteroscedasticity between the several variables expressing corporate performance and they attempted to construct a model of the executive compensation hypothesis that would reduce these statistical problems. Lewellen and Huntsman concluded their analysis with the comment that "profits appear to have a strong and persistent influence on executive rewards, whereas sales seem to have little, if any, such impact" (1970, p. 718).

The findings of the research following Lewellen and Huntsman have been mixed. Larner (1970) found that despite the growing separation of ownership and control in the large corporation, compensation is most consistently linked to profit. Ciscel (1974) found that the managerialist position was stronger, particularly when the salaries of the entire executive group were considered. Auerbach and Siegfried (1974) found that variables such as market concentration, market barriers to entry and control of the board of directors have little impact on executive compensation. Smyth, Boyes and Peseau (1975) reexamined Lewellen and Huntsman's model, improved the corrections for multicollinearity and heteroscedasticity, and concluded that executive compensation is based on a "utility function of both sales and profits" (1975, p. 79). Finally, two studies of executive compensation in Great Britain (Cosh, 1975; Meeks and Whittington, 1975) generally support the managerialist position of a relationship between compensation and sales, but not profit.

II. Theoretical and Empirical Estimation Problems

Most of the differences in previous studies of the determinants of executive compensation can be traced to obvious differences in data selection, variable transformation or hypothesized model construction. This is because testing for the determinants of executive compensation is fraught with serious theoretical and econometric problems. First, the high degree of correlation between net corporate income, sales volume and net assets means that tests for the significance of individual variables are unreliable. To eliminate multicollinearity, some transformation of the data is necessary. Unfortunately, such transformed variables often bear only the faintest relationship to the hypothesis being tested. A similar difficulty exists with the problem of heteroscedasticity. Like multicollinearity, the presence of heteroscedasticity does not bias the estimate of the coefficients themselves, but because the ordinary least squares estimate is inefficient, tests for the significance of variables are unreliable. In order to control for heteroscedasticity, some weighting system is introduced to provide a constant variance in the error term.

A third problem, although little-discussed in the executive compensation literature, is perhaps the most serious. Not only are profits and sales highly correlated in their raw form, but economic theory tells us that the definition of profit is total revenue (the actual form of the sales variable) minus total cost. Hence, sales are actually causally prior to profit; instead of being an "independent" variable, the profit variable is likely to be correlated with the error term in a one equation model. As a practical matter, the fact that total revenue (the typical reported measure of sales) is one constituent of the profit variable means that the significance of sales as an explanatory variable of executive compensation is compatible with the assumption that increasing sales tends to increase profit (the neoclassical model) and the assumption that a firm's manager seeks to increase sales as a proxy for the size of the firm (the managerial hypothesis). Below it will be shown how the indirect least squares approach to estimation (which first nets out the effect of sales on profit) not only eliminates the problem of simultaneous equations bias and multicollinearity between sales and profits, but also provides an even stronger test of both competing hypotheses.[1]

[1] It was suggested to us that profit might be considered causally prior to sales in the sense that a certain level of profit is required to maintain a certain sales volume. If valid, this

A final problem lurks in the executive compensation question—that of possible measurement error in the data. Since economists make use of published data, rather than collecting their own observations, serious dangers of measurement error may exist. However, few tests for the presence of measurement error have been developed, short of comparing the regression results of data of known accuracy with those of another data set. But such a heroic test is implausible; if one data set was known to be error free, the other data set would never be used. Hence, measurement error continues as a convenient excuse for perverse econometric results and undoubtedly explains why few, if any, econometric tests have ever conclusively disproved an economic hypothesis.

III. Econometric Results

Since at least three econometric problems—multicollinearity, heteroscedasticity and simultaneous equation bias—are likely to be encountered in testing the executive compensation model, it is important to take problems one at a time. In order of severity, it is most damaging to encounter simultaneous equation bias, since the correlation of an allegedly independent variable with the error term will make OLS estimators both biased and inconsistent. Any attempt, for instance, to base corrections for heteroscedasticity upon the observed correlation between the squared residual of the regression equation and a measure of scale will be inadequate if the estimate of the error term is itself biased.[2]

argument is more relevant in the long run, since a firm could operate at a loss for a year or more without being forced out of business. While a firm may use a profit target to determine either its product price or advertising budget, there is no reason, a priori, to expect ex post profit to exactly correspond to the ex ante profit target. In addition, the data used in this study are cross-sectional accounting data where profit is defined as a residual of revenue.

[2] Lewellen and Huntsman (1970) attempt to control for heteroscedasticity by dividing the variables in the regression equation by the book value of assets. In a test of the advisability of such a step the authors re-estimated the Lewellen-Huntsman equation, using the same data used for this analysis. Then, by applying both Bartlett's test (Kane, 1968, p. 374) and regressing the squared residual against assets, it was found that a statistically significant degree of heteroscedasticity existed in the weighted equations used by Lewellen and Huntsman. Since we found that the square root of assets was the appropriate deflator to use (see below), we had to conclude that this "over correction" of Lewellen and Huntsman was possibly induced by an incorrect estimate of the error term itself.

We begin the analysis by stipulating the simplest structural model:[3]

$$C_{it} = a_{0t} + a_{1t}P_{it} + a_{2t}S_{it} + u_{it} \tag{1}$$

where

C_{it} = salary plus bonus (in dollars) for the chief executive officer of the i^{th} company in year t; t = 1970, 1971, 1973–76

P_{it} = net (after tax) income (in $1,000's) for the i^{th} corporation in year t

S_{it} = total sales revenue (in $1,000's) for the i^{th} corporation in year t.

Not surprisingly, the results of using ordinary least squares to estimate equation (1) were virtually incomprehensible. Using a step-wise regression approach, it was found that the two variables, P_{it} and S_{it}, were entered more or less in random order from one year to another. Whichever variable entered first seemed to destroy the statistical significance of the other variable. This appeared to confirm suspicions of severe multicollinearity with the added problem of simultaneous equation bias.[4]

To eliminate multicollinearity and to control for the danger of simultaneous equation bias, observed profit, P_{it}, was regressed against observed sales, S_{it}. This regression yielded "profit predicted by sales," $P(S_{it})$:

$$P(S_{it}) = c + dS_{it}. \tag{2}$$

The results of the regression, not surprisingly, were very strong. They are presented in table 1.

TABLE 1.—THE RELATIONSHIP OF PROFITS AND SALES

Year	$\hat{c}_t$	$\hat{d}_t$	F-Value	R^2	n
1970	− 2193.52	.0535	357.59	.632	209
1971	−28342.80	.0668	1317.69	.795	212
1973	−19624.87	.0733	802.59	.798	219
1974	− 7382.81	.0530	547.09	.755	218
1975	− 773.56	.0469	436.81	.684	221
1976	− 5978.90	.0527	897.60	.816	219

The next step was to calculate "residual profit," $\hat{P}_{it}$, which is the observed profit variable minus "profit predicted by sales":

[3] All data used were published in the annual directories of Fortune and Forbes. The data base of corporations is the 230 largest industrial corporations in 1974 as ranked by Fortune. The executive compensation data were not published in 1972. Due to missing data, the number of firms included in any one year's analysis ranges from 209 to 221.

[4] The Pearson correlation coefficients between P_{it} and S_{it} exceeded 0.9 for all six years of data.

$$\hat{P}_{it} = P_{it} - P(S_{it}) = P_{it} - \hat{c}_t - \hat{d}_t S_{it}. \quad (3)$$

"Residual profit" can best be interpreted as profit due to reducing production cost (or increasing technical efficiency).

Since residual profit is, by definition, uncorrelated with the sales variable, replacing observed profit by residual profit handles both problems of simultaneous equation bias and multicollinearity. Furthermore, the significance of residual profit would provide extremely strong support for the neoclassical hypothesis; if chief executive officers are rewarded for increasing technical efficiency (sales revenue held constant), there is strong evidence that corporate management is interested in goals other than increasing the firm's market share or asset size.

Using residual profit, which is orthogonal to both the error term (u_{it}) and sales (S_{it}), the following equation was estimated:

$$C_{it} = b_0 + b_1\hat{P}_{it} + b_2 S_{it} + u_{it}. \quad (4)$$

The results (table 2) would appear to give, at best, only weak support to the neoclassical hypothesis. The coefficient for residual profit, $\hat{b}_1$, is statistically significant at the 1% level in only two of the six years, and even has the unexpected sign in 1973. By contrast, $\hat{b}_2$, our estimate of the effect of sales on executive compensation, was positive (the expected sign) and was significant at the 1% level, using a one-tailed test. Since the sales variable performs two roles, however, the significance of $\hat{b}_2$ cannot be interpreted as an indication that the managerialist hypothesis is correct and that the neoclassical theory is wrong.[5] Instead, rewards for increasing sales can be interpreted either as a reward for growth or as a reward for increasing the total revenue component of corporate profit.

One interesting result of equation (4) is the apparently high explanatory power of $\hat{b}_0$, the constant term. Usually the intercept term in a regression equation is ignored when interpreting regression results, since this term is considered the arbitrary starting point, especially when time series analysis is involved. With cross-section analysis, the constant term can be considered the "minimum value" of the dependent variable before the effects of the independent or explanatory variables are taken into account. Since considerable skill is required to manage a major corporation, all chief executive officers must earn at least a normal return on their "general human capital" (Becker, 1964) or else the firm will lose their services to another corporation. The executive compensation controversy concerns the determinants of "incentive pay" for corporate decision-makers. What the constant term in equation (4) picks up is the equilibrium price of a chief executive officer's time, determined by the interaction of the supply and demand for management skills.

Identification of the exogenous factors that influence the base pay of executives is beyond the scope of this analysis, since the classic identification problem involved in sifting supply and demand factors is involved in such an investigation. However, the explanatory power of the constant term, when introduced as a variable in a

[5] While realizing that the procedure was not in conformity with economic theory (see note 1), we decided to determine what the results of equation (4) would have been if we had treated sales as dependent on profit, so that "residual sales" would be used in place of the sales variable in equation (4), while total profit was used in place of residual profit. The results of this regression, it could be argued, would provide as strong a test of the managerial hypothesis as equation (4) is of the neoclassical hypothesis. The results of this experiment, while dubious because of likely specification error, nevertheless provide a mirror image of the results of equation (4). Residual sales, i.e., sales independent of profit, proved significant at the 1% level in three years (1970, 1973 and 1976) and significant at the 5% level in 1971. Both coefficients (for profit and residual sales) had the expected sign in all six years, with the profit coefficient being statistically significant at the 1% level in all six years.

TABLE 2.—DETERMINANTS OF EXECUTIVE SALARIES (UNWEIGHTED): 1970–1976

Year	$\hat{b}_0$	F	$\hat{b}_1$	F	$\hat{b}_2$	F	R^2	n
1970	185,938.6[a]	752.95	.06792	1.3456	.01487[a]	38.812	.8857	209
1971	211,531.2[a]	849.14	.01821	0.0784	.00866[a]	15.920	.8711	212
1973	262,531.8[a]	1032.33	−.0045	0.0009	.01534[a]	69.222	.9026	219
1974	295,325.0[a]	1534.29	.1652[a]	17.8084	.00892[a]	53.856	.9278	218
1975	308,969.4[a]	1120.91	.1595[a]	9.6100	.00994[a]	40.323	.9018	221
1976	349,612.9[a]	1208.26	.0498	0.1161	.01318[a]	81.001	.9131	219

[a] Significant at the 1% level.

step-wise regression, and the regularity of its growth through the years investigated, indicate that attributing the value of the constant term to a market for managerial talent is not unreasonable.

One econometric problem still remained as a source of difficulty in interpreting the results of equation (4), the possibility of heteroscedasticity, or a non-constant variance in the error term. That is, letting σ_u^2 stand for the variance in the error term, heteroscedasticity will occur when σ_u^2 varies systematically with some scale factor, K_{it}, such that $\sigma_u^2 = \sigma^2 \cdot K_{it}$, where σ^2 is the homoscedastic portion of σ_u^2. The presence of heteroscedasticity was discovered by the use of Bartlett's test (Kane, 1968, p. 373).[6] The presence of heteroscedasticity in equation (4) meant that the F-values found for $\hat{b}_1$ and $\hat{b}_2$ may have been inaccurate.

Instead of following the usual procedure of the Goldfeld-Quandt test (Goldfeld and Quandt, 1965), which compares the linear specification of a model to a ratio model wherein all variables are deflated by some proxy for size (e.g., assets in Lewellen and Huntsman),[7] we decided to employ a more specific test for heteroscedasticity suggested by R. E. Park (Kane, 1968, p. 376). Since substituting residual profit for observed profit in equation (2) should eliminate the problem of simultaneous equation bias, the regression results of equation (4) provide an unbiased estimate of the disturbance term, u_{it}. Since the expected value of u_{it} is zero, the square of $\hat{u}_{it}$, obtained as the residual from equation (4), serves as an unbiased estimate of σ_u^2. This allowed us to estimate the best deflator for equation (4) to correct for heteroscedasticity. Following the suggestion of Park, we estimated

$$\ln (u_{it})^2 = \ln \alpha + \delta \ln (X_{it}).$$

Two alternative scale proxies were used, sales and book value assets. The sales variable, S_{it}, showed no relationship with the squared residual, while the book value of assets, A_{it}, was correlated with the squared residual at the 1% level of confidence.[8] However, the estimated value of δ, $\hat{\delta} = 0.979$, not only allowed us to reject the null hypothesis that $\delta = 0$, but also allowed us to reject the ratio hypothesis, i.e., $\delta = 2$. Hence, similar to the procedure used by Smyth, Boyes and Peseau (1975),[9] we found that the best correction for heteroscedasticity was not the simple ratio model employed by Lewellen and Huntsman, but, in our case, the best deflator was $(A_{it})^{\frac{1}{2}}$. By dividing each variable in equation (4) by the square root of assets, $(A_{it})^{\frac{1}{2}}$, this equation becomes

$$\frac{C_{it}}{(A_{it})^{\frac{1}{2}}} = b_0 \frac{1}{(A_{it})^{\frac{1}{2}}} + b_1 \frac{\hat{P}_{it}}{(A_{it})^{\frac{1}{2}}} + b_2 \frac{S_{it}}{(A_{it})^{\frac{1}{2}}} + \frac{u_{it}}{(A_{it})^{\frac{1}{2}}}. \tag{5}$$

The results of the estimation of equation (5), reported in table 3, did indeed increase the significance of the residual profit variable (now weighted by the square root of corporate assets) without reducing the significance of the other variables. These results show all coefficients having the expected signs, with $\hat{b}_0$ and $\hat{b}_2$ being significantly different from zero at the 1% level in all years, while $\hat{b}_1$, the coefficient for residual profit, is significant at the 1% level for five years and at the 5% level for one year, 1971. The constant term, although reduced somewhat in equation (5) relative to equation (4), nevertheless is of a mag-

[6] Using this test we were able to reject the null hypothesis of homoscedasticity with 95% confidence, employing a one-tailed test. A one-tailed test was deemed more appropriate in light of Lewellen and Huntsman's findings of positive heteroscedasticity in their paper (1970). Had we employed a two-tailed test, we would have been able to reject the homoscedasticity hypothesis with only 90% confidence.

[7] Given the presence of heteroscedasticity in the linear form of the model, the use of a ratio model assumes that the heteroscedastic element of σ_u^2 is equal to the square of some scale proxy, i.e., that $\sigma_u^2 = \sigma^2 K_{it} \approx \sigma^2 (X_{it})^2$, where X_{it} is the scale proxy. If the heteroscedastic element varies with some other transformation of X_{it}, $K_{it} = \alpha (X_{it})^{\delta}$, with $\delta \neq 2$, the ratio model suggested by the Goldfeld-Quandt method could contain an even greater degree of heteroscedasticity than the linear specification (see note 2).

[8] Regressing the natural logarithm of the squared residual, $\ln(u_{it})^2$ against the natural logarithm of sales yielded a regression coefficient estimate of $\hat{\delta} = 0.000003$ and an F-value of 0.0000, which was obviously not statistically significant. The regression of $\ln(u_{it})^2$ against $\ln(A_{it})$ yielded a value of $\hat{\delta} = 0.979$ and an F-value of 5.20, which was statistically different from both zero and two, but not significantly different than one. For computational ease, only data from 1974 were used since all data for all years were recorded in the order of book value assets for 1974.

[9] Smyth, Boyes and Peseau (1975) found a deflator of $(A_{it})^{.8}$ would minimize heteroscedasticity. This difference from our results could be due to two differences in techniques. First, they employed an iterative method, rather than the Park method for discovering the scale proxy. Second, they did not employ residual profit, but used observed profit and sales as independent variables. If the causal priority of sales to profits introduces problems of simultaneity, their determination of the "best" deflator would be based on a biased estimation of the residual term.

TABLE 3.—DETERMINANTS OF EXECUTIVE SALARIES: 1970–1976
(corrected for heteroscedasticity)

Year	$\hat{b}_0$	F	$\hat{b}_1$	F	$\hat{b}_2$	F	R^2	n
1970	164,068[b]	514.08	.39320[b]	12.47	.02917[b]	25.75	.8954	209
1971	200,063[b]	458.39	.20511[a]	2.77	.01547[b]	4.56	.8608	212
1973	219,490[b]	599.95	.34815[b]	13.57	.03092[b]	38.16	.9002	219
1974	259,759[b]	832.42	.27065[b]	6.20	.01995[b]	24.34	.9138	218
1975	268,898[b]	670.57	.42315[b]	17.89	.02083[b]	22.17	.8953	221
1976	315,723[b]	698.72	.44302[b]	11.79	.02157[b]	22.35	.8946	219

[a] Significant at the 5% level.
[b] Significant at the 1% level; all one-tailed tests.

nitude that supports the designation of this coefficient as the opportunity cost of a chief executive's time, or the money payment for his general human capital.

The significance of the coefficient $\hat{b}_1$, the estimated effect of a change in residual profit (i.e., profit due to increases in technical efficiency) on the level of executive compensation, provides a strong vindication of the neoclassical model of management behavior. In all years we can confidently reject the null hypothesis that increasing the firm's profits by means other than increasing sales (e.g., by reducing production costs) does not increase executive rewards. The significance of the coefficient $\hat{b}_2$, the effect of an increase of sales (revenue) on executive salary plus bonus, supports both the neoclassical and managerial theories of the firm, since increasing sales also tends to increase profit.

Unfortunately, the indirect squares approach of equations (4) and (5), made necessary to avoid problems of multicollinearity and simultaneous equations bias, does not provide a decisive test of the managerial hypothesis, that managers are rewarded for achieving a steady growth in sales, which provides both stability and a steady growth in corporate resources. That executive compensation increases with sales is consistent with both the neoclassical and the managerialist positions, and thus leaves the controversy unresolved.

Because of the dual explanation of the effect of sales on executive compensation, we decided that the inclusion of assets as a separate variable would provide an indirect test of the validity of the managerialist hypothesis. If the level of assets, independent of both sales and the residual profit, influences the level of executive compensation, then the argument that managers are compensated for firm growth, independent of profit, gains support. Two variants of equation (5), each including book value assets, A_{it}, as an additional explanatory variable, were examined.[10] The results were mixed. The inclusion of A_{it} in the equation reduced the significance of S_{it} in most instances. In equation (7), where the influence of both the level of assets and the square root of assets on executive compensation was tested, three of the explanatory variables are significant in five out of six years, while sales are significant in four years. The apparent pattern that emerges from inclusion of assets in the equation is that executive compensation (unweighted) rises with the square root of assets, but falls slightly with the level of assets. We interpret this to mean that there are diminishing returns to firm growth in the reward structure of chief executive officers.

[10] One variant constrained the equation to pass through the origin, but added the explanatory variable "book value of assets," A_{it}:

$$\frac{C_{it}}{(A_{it})^{\frac{1}{2}}} + b_0 \frac{1}{(A_{it})^{\frac{1}{2}}} + b_1 \frac{\hat{P}_{it}}{(A_{it})^{\frac{1}{2}}} + b_2 \frac{S_{it}}{(A_{it})^{\frac{1}{2}}} + b_3 \frac{A_{it}}{(A_{it})^{\frac{1}{2}}} + \frac{u_{it}}{(A_{it})^{\frac{1}{2}}}. \tag{6}$$

In addition, inclusion of both book value of assets and an intercept term (which measures the effect of the square root of A_{it} on C_{it}) allowed us to examine the possibility that executive compensation may increase at a decreasing rate with the level of assets:

$$\frac{C_{it}}{(A_{it})^{\frac{1}{2}}} = b_0 \frac{1}{(A_{it})^{\frac{1}{2}}} + b_1 \frac{\hat{P}_{it}}{(A_{it})^{\frac{1}{2}}} + b_2 \frac{S_{it}}{(A_{it})^{\frac{1}{2}}} + b_3 \frac{A_{it}}{(A_{it})^{\frac{1}{2}}} + b_4 \frac{(A_{it})^{\frac{1}{2}}}{(A_{it})^{\frac{1}{2}}} + \frac{u_{it}}{(A_{it})^{\frac{1}{2}}}. \tag{7}$$

However, the inclusion of assets may reintroduce problems of multicollinearity, which the creation of the residual profit variable originally eliminated (albeit, as a consequence of eliminating the more serious problem of simultaneous equations bias). This multicollinearity could explain the decline in the significance of the sales variable when assets are included as a separate variable. Detailed statistical results of equations (6) and (7) are available from the authors upon request.

IV. Conclusions

The primary thrust of our findings is negative. The empirical estimates of the determinants of executive compensation, after correcting for multicollinearity, the causal priority of sales to profits, and eliminating the inefficiency stemming from heteroscedasticity, are in conformity with both a neoclassical and a managerialist interpretation of firm behavior. Executives are paid for increasing profits, whether through sales growth or cost control. However, since the sales variable may also serve as a measure for firm size, and since asset size of the corporation also bears an important influence on executives' salaries, there is a strong indication that decisions concerning executives' salaries are influenced by several aspects of corporate performance.

In addition, the level of executive compensation is basically determined in a market for executives. There is a consistent pattern in the regression results from year to year that implies that the base salary received by management is determined through the interaction of supply and demand.

An important econometric trade-off exists between the results of the various equations investigated here. When only sales and residual profits are used to explain increments in executive compensation, the compensation formula, through the intercept term, reveals the influence of the market for managerial talents on the base salary of chief executives. For successive years, an executive salary, independent of firm performance, is determined by exogenous factors. However, the introduction of the weight, the square root of assets, to reduce the problem of heteroscedasticity, muddled the meaning of the market salary term. The second formulation of the compensation formula indicates that both sales, as a measure of size and profitability, and residual profits, as a measure of technical efficiency, have a strong influence on the levels of executive compensation in the large corporation.

An econometric analysis of the determinants of the level of executive compensation stands apart from the neoclassical/managerialist controversy. The accounting data available for analysis of the issue do not allow the construction of a model robust enough to differentiate between the contending schools of thought. However, three influences—the market for managerial talent, the external performance of the firm and the internal technical efficiency of production—can be identified for successive years as significant and regular influences on the level of executive compensation.

REFERENCES

Auerbach, Paul R., and John J. Siegfried, "Executive Compensation and Corporate Control," *Nebraska Journal of Economics and Business* 13 (Summer 1974), 3–16.

Baumol, William J., *Business Behavior, Value and Growth* (rev. ed.) (New York: Harcourt, Brace and World, 1967).

Becker, Gary S., *Human Capital: A Theoretical and Empirical Analysis* (New York: Columbia University Press for NBER, 1964).

Ciscel, David H., "Determinants of Executive Compensation," *Southern Economic Journal* 40 (Apr. 1974), 613–617.

Cosh, Andrew, "The Remuneration of Chief Executives in the United Kingdom," *Economic Journal* 85 (Mar. 1975), 75–94.

Goldfeld, Stephen M., and Richard E. Quandt, "Some Tests for Homoscedasticity," *Journal of the American Statistical Association* 60 (1965), 539–547.

Kane, Edward J., *Economic Statistics and Econometrics* (New York: Harper and Row, 1968).

Larner, Robert J., *Management Control and the Large Corporation* (New York: Dunellen, 1970).

Lewellen, Wilbur G., *Executive Compensation in Large Industrial Corporations* (New York: Columbia University Press for NBER, 1968).

———, "Management and Ownership in the Large Firm," *Journal of Finance* 24 (May 1969), 299–332.

———, *The Ownership Income of Management* (New York: Columbia University Press for NBER, 1971).

Lewellen, Wilbur G., and Blaine Huntsman, "Managerial Pay and Corporate Performance," *American Economic Review* 60 (June 1970), 710–720.

McGuire, Joseph W., John S. Y. Chiu, and Albar O. Elbing, "Executive Incomes, Sales and Profits," *American Economic Review* 52 (Sept. 1962), 753–761.

Masson, Robert T., "Executive Motivations, Earnings, and Consequent Equity Performance," *The Journal of Political Economy* 79 (Nov. 1971), 1278–1292.

Meeks, Geoffrey, and Geoffrey Whittington, "Directors' Pay, Growth, and Profitability," *The Journal of Industrial Economics* 24 (Sept. 1975), 1–14.

Patton, Arch, *Men, Money and Motivation* (New York: McGraw-Hill, 1961).

Roberts, David R., *Executive Compensation* (Glencoe: Free Press, 1959).

Smyth, David J., William J. Boyes, and Dennis E. Peseau, *Size, Growth, Profits, and Executive Compensation in the Large Corporation* (New York: Holmes & Meier, 1975).