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Accounting and Market-Value Measures of Profitability: Consistency, Determinants, and Uses

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This study examines the relationship between accounting and market-value measures of profitability for individual firms. Differences in measures of profitability are observed both between and within industry groups and thus cannot be explained by differences in uncontrolled industry-specific influences. We suggest that both accounting and market can be used as unique but imperfect indicators of profitability. Using a LISREL model approach, we find R & D intensity, television advertising intensity, leverage, and industry growth to be important determinants of firm profits.

KEY WORDS: Accounting; LISREL; Market value; Profitability.

1. INTRODUCTION

The use of accounting income numbers to measure firm performance is typically justified on the grounds that they are the best available data. There are measurement problems, however, caused by different accounting practices across industries, (possibly) inappropriate expensing of research and development (R & D) and advertising expenditures, a failure to reflect opportunity costs and risk, and replacement-cost accounting errors. Although many researchers argue that these problems only reduce but do not destroy the usefulness of accounting profit data, others disagree. Fisher and McGowan (1983), for example, argued that since true rates of return fundamentally depend upon the time shape of individual project revenues and costs, only by the "merest happenstance" (p. 83) will traditional firm-wide accounting averages be insightful. In fact, they suggest that firm-level accounting profit data may be meaningless.

Coincident with the rising mistrust of accounting profit data has been an increasing use of market estimates of profitability. Thomadakis (1977) and Lindenberg and Ross (1981), among others, have argued persuasively that market-value data provide new and useful perspectives on the magnitude and sources of profitability. Thus there is a promising, but still largely unexplored, area for research that jointly considers accounting and market measures of profitability. As Ross (1983) wrote,

> The prices that the markets place on the securities issued by firms and the changes in these values over time provide an ongoing assessment of the value of such firms. Accounting data, on the other hand, provide an alternative view of the same firms. The information on the balance sheets and in the income statements provides a historical record of where the firm has been and where it currently stands. One way to view the accounting data is to think of them as providing information on the resources used by the firm and on its performance. Comparing accounting data with market data, then, provides us with a comparison between two ways of looking at the same thing, and such comparisons will be valuable for both. (p. 376)

This article adopts the perspective that neither accounting nor market data provide an ideal or true measure of profitability. Instead, we argue that measures developed from both sources offer potentially unique but imperfect measures of profitability. Like Ross, and Gonedes and Dopuch (1979) before him, we believe that a comparison of accounting and market data can prove highly beneficial. Our objectives in this article are to (a) learn whether accounting and market data provide consistent or independent measures of profitability; (b) determine if any differences between accounting and market-value profit measures are caused by industry related factors, or are instead due to firm-specific influences and/or accounting conventions; and (c) consider the determinants of profitability within a framework that allows for the imperfect nature of profit measures.

Section 2 presents a correlation and factor analysis of various measures of profitability. Results suggest that accounting measures do not reflect the same underlying profitability phenomenon that is captured by market data. In Section 3 the analysis is narrowed to focus on firms within the same (two-digit) industry group to learn whether the observed variability in profit measures is due, in large part, to industry effects. The answer appears to be no. Apparently, profit variability is more sensitive to firm-specific rather than industry-specific forces. This view is supported by a linear-structural-model (LISREL) analysis of firm profit performance in Section 4. Here R & D intensity, television advertising intensity, leverage, and industry growth emerge as important determinants of firm profits. Conclusions and implications for future research are discussed in Section 5.

2. MEASURES OF PROFITABILITY

The market value of the firm can be viewed as the risk-adjusted present value of future profits and has the following major components:

$$MV(F) = MV(T) + MV(I), \tag{1}$$

where $MV(F)$ is market value of the firm and $MV(T)$ and $MV(I)$ are the capitalized values of profits attributed to tangible and intangible assets, respectively (Hirschey and Weygandt in press).

Although $MV(F)$ is observable, subcomponents $MV(T)$ and $MV(I)$ are not. Accounting book values and replacement-cost values, however, can be viewed as useful, though imperfect, measures of the market value of tangible assets. Using these accounting data, it becomes possible to isolate the profit effects of tangible assets from any additional influences of intangible assets such as market power, patents, brand loyalty, goodwill, and so on.

The most widely adopted market-value–based measure of profits is commonly referred to as the *Q ratio*, where

$$Q = \frac{MV(F)}{RC(T)} \tag{2}$$

and $RC(T)$ is the replacement cost of tangible assets (see Tobin 1978). When $Q > 1$, market value reflects profit-producing intangible assets not reflected in replacement-cost data. Since $MV(T) = RC(T) + \text{error } (e)$,

$$Q = 1 + \frac{MV(I)}{MV(T)} + e_1,$$
$$= f_1(\underline{X}) + e_1, \tag{3}$$

say, where $\underline{X}$ is a vector of intangible sources of profitability.

Thomadakis (1977) used a market-value measure of profits called relative excess valuation, EV/S, where

$$\frac{EV}{S} = \frac{MV(F) - BV(T)}{S}, \tag{4}$$

$BV(T)$ is the book value of tangible assets, and S is sales. Since $MV(T) = BV(T) + \text{error } (e)$,

$$EV/S = \frac{MV(I)}{S} + e_2$$
$$= f_2(\underline{X}) + e_2, \tag{5}$$

say, where $\underline{X}$, again, includes various intangible sources of profitability.

The Q ratio can be thought of as the market-value analog to more traditional measures, such as accounting return divided by assets or stockholders' equity. Like each of these measures, Q can be criticized as an index of profitability in that it will be affected by varying capital intensity and leverage across firms. EV/S, calculated using historical cost data, has the attractive feature of being normalized by sales—a factor and leverage neutral measure of size. EV/S is readily calculated given widespread historical cost data, but it is more likely to be distorted by inflation than is Q and other replacement-cost measures of profitability. Perhaps the primary theoretical virtue of EV/S is that it can be thought of as the market-value analog to the return on sales or Lerner index of monopoly profits (see Hirschey in press).

On an a priori basis, we expect market-value data to provide profit information that is distinct from that provided by accounting data, given its expectational or forward-looking, as opposed to historical, perspective. This is, however, an empirical question. Table 1 shows simple correlations between traditional accounting profit measures (net income plus interest) and newer market estimates of profitability for a 386-firm sample taken from the 1977 Fortune 500. (This is a slightly reduced sample from that employed in an earlier related study—Hirschey 1982. Since we wished to explore the importance of both intraindustry and interindustry variability in profit data, we eliminated a single leather-industry firm (Standard Industrial Classification (SIC) 30) and three miscellaneous manufacturers (SIC 39) from consideration. This left 386 observations from an earlier sample of size $n = 390$. See Hirschey 1982 for a detailed description of the data.)

This sample is attractive in that it is broadly representative of industrial manufacturing. Q was estimated as the ratio of the market value of the firm divided by the replacement cost of tangible assets. (The replacement-cost data used here are the estimated annual replacement cost of inventories and productive capacity

Table 1. Correlation Matrix of Accounting Historical, Accounting Replacement, and Market-Value Measures of Profitability

Variable	π/A	π/E	π/S	π/A_R	π/E_R	π/S_R	Q	EV/S
Historical Return on Assets, π/A	1.000							
Historical Return on Equity, π/E	.738	1.000						
Historical Return on Sales, π/S	.731	.520	1.000					
Replacement Return on Assets, π/A_R	.828	.688	.652	1.000				
Replacement Return on Equity, π/E_R	.681	.831	.513	.887	1.000			
Replacement Return on Sales, π/S_R	.712	.543	.826	.867	.692	1.000		
Market Q Ratio, Q	.625	.322	.579	.639	.419	.608	1.000	
Market Relative Excess Values, EV/S	.604	.303	.617	.563	.352	.610	.937	1.000

at the end of each fiscal year for which a balance sheet is required and the approximate amount of cost of sales and depreciation based on replacement cost (Securities and Exchange Commission 1977). These are price-level adjusted historical cost numbers, which, as is typical, do not involve any adjustment to the liability side of the firm's balance sheet. Since inflation affects the real values of both assets and liabilities, these data can only be viewed as an imperfect approximation to true replacement cost numbers. See Beaver, Christie, and Griffin 1980.) Following Thomadakis, EV/S was estimated as the difference between the market value of the firm and the book value of tangible assets, all normalized by sales. Despite the differential treatment of tangible asset values, a close relation between Q and EV/S is evident. Generally weaker correlations among various accounting estimates of profitability are obvious, and the relation between accounting and market estimates of profitability seems fairly erratic.

To explore these relations more fully, a factor analysis (see Johnson and Wichern 1982) of accounting profit measures and market estimates of economic profits was conducted. An $m = 3$ factor model was analyzed, where

$$\mathbf{Y} = \Lambda \mathbf{F} + \epsilon \tag{6}$$

and $\mathbf{Y}$ is the (8×1) vector of (standardized) accounting and economic profit measures (see Table 1); Λ is an (8×3) matrix of factor loadings; $\mathbf{F}$ is a (3×1) vector of unobservable common factors with $E(\mathbf{F}) = \mathbf{0}$ and $\text{cov}(\mathbf{F}) = I$; and ϵ is an (8×1) vector of specific factors or errors with $E(\epsilon) = \mathbf{0}$ and $\text{cov}(\epsilon) = \Theta$, a diagonal positive-definite matrix. We also assume that $\mathbf{F}$ and ϵ are independently distributed.

Rotated principal-component estimates of factor loadings based upon the sample correlation matrix R are presented in Table 2. Here we find that 91% of the total variance among the alternative profit measures can be captured in terms of three underlying unobservable factors. It is interesting that both market-value measures load highly and about equally on a single factor, F_1. This finding implies that either measure can be taken as an attractive indicator of the same economic phenomenon. Furthermore, given that Q and EV/S are only weakly related to F_2 and F_3, and that accounting measures have small to moderate loadings on F_1, we conclude that the market-value data provide evidence on profitability that is distinct from that provided by accounting data.

The uniqueness of accounting historical data versus accounting replacement data is less clear. Both historical and replacement measures of the return on assets and return on stockholders' equity are highly related to F_2. On the other hand, historical and replacement estimates of the return on sales are closely related to F_3. This mixing of factor loadings for accounting historical and replacement measures makes it difficult to argue that inflation adjustments make replacement profit data inherently superior to historical profit data.

Table 2. Rotated Principal Component Estimates of Factor Loadings (n = 386, m = 3 factors)

	Estimated Factor Loadings			
Variable	F_1	F_2	F_3	Communalities
Historical return on assets	.433[a]	.612[a]	.499[a]	.812
Historical return on equity	.125	.892[b]	.234	.866
Historical return on sales	.296	.238	.887[b]	.931
Replacement return on assets	.406[a]	.708[a]	.483[a]	.900
Replacement return on equity	.198	.895[b]	.283	.921
Replacement return on sales	.331	.414[a]	.789[a]	.903
Market Q ratio	.928[b]	.160	.294	.973
Market relative excess value	.910[b]	.079	.355	.961
Cumulative proportion of total variance explained	.287	.628	.908	

[a] Moderate factor loadings.
[b] Large factor loadings.

On the other hand, one can argue that returns on assets and/or equity provide fairly distinct information from that implicit in return on sales data.

From the preceding analysis, it becomes clear that important differences are present among various accounting measures of profitability, and that accounting and market measures do not capture the same aspects of profitability. Therefore, it becomes interesting to inquire as to the sources of variability among alternative profit measures. An obvious possibility is that profit variability may be caused by uncontrolled industry-specific effects. To evaluate this possibility, a more narrow analysis of firms within two-digit industry groups is considered.

3. INDUSTRY ANALYSES

If the observed variability in profit measures is attributable to uncontrolled industry-specific influences, we would expect that a factor analysis of subsamples restricted along industry lines would yield unique patterns of factor loadings. On the other hand, if the factor-loading pattern for various industry groups is compatible with that observed over the entire sample (Table 2), we would be led to believe that firm-specific influences and/or accounting conventions are responsible for profit-measure variability.

Factor analysis was performed for firms in various two-digit industry groups. (Factor analyses results for the two-digit industry groups are available from the authors upon request.) The subsamples were chosen on the grounds that they are representative of manufacturers in the consumer-nondurable (food processing), consumer-durable (transportation equipment), producer-nondurable (chemicals), and producer-durable (primary metals) goods sectors. Although there were some differences among the subsamples, market measures tend to load highly (loadings of about .8–.9) on a single factor. The loadings of the accounting profit measures were more diffuse. To some extent, the accounting return-on-assets and return-on-equity numbers define a single factor and the accounting return-on-sales numbers define a second factor. This demarcation seemed to hold for the food-processing and chemical firms. It was less apparent for the primary-metals and transportation-equipment firms. We conclude that the dichotomy between accounting and market measures is maintained both among and within industries and, therefore, that this variation is not due primarily to industry effects.

A lack of industry effects is also suggested by a K-means cluster analysis (see Johnson and Wichern 1982). A brief summary of the results is reported in Table 3. Here firms within each industry group are allocated into any of four cluster groups on the basis of the similarities between the accounting and market data. Similarities are represented by correlations. With important industry effects, we would expect to see firms within the same industry allocated to the same cluster. Again, we see from Table 3 that the pattern of profit-measure correlations is not uniquely determined along industry lines. Therefore, as in the industry-group factor analysis, the cluster analysis suggests that industry-specific influences play only a minor role, if any, in explaining profit-measure variability.

Table 3. Cross Tabulation of Firm-Profit Measures by Variable Clusters Versus SIC Codes

	Cluster Group				
SIC Industry Group	C_1	C_2	C_3	C_4	Totals
Food processing (20)	2	3	17	26	48
Transporation equipment (37)	0	4	21	14	39
Chemicals (28)	12	2	22	20	56
Primary metals (33)	0	14	6	8	28
Total	14	23	66	68	171

4. MODELING FIRM PERFORMANCE

Results reported in Section 2 suggest that accounting profit data capture an aspect of firm performance that is unique from that provided by market-value estimates. In the absence of ideal data, both accounting and market information can be analyzed, since both provide useful, though imperfect, profit information.

We take the position that accounting profits and market-value profit measures are imperfect indicators of three unobservable (latent) true performance characteristics. The (mean-corrected) measurement model is

$$\mathbf{Y} = \Lambda_y \boldsymbol{\eta} + \boldsymbol{\epsilon}, \tag{7}$$

where $\mathbf{Y}$ is an (8×1) vector of responses, $\boldsymbol{\eta}$ is a (3×1) vector of latent performance characteristics, Λ_y is an (8×3) coefficient matrix, and $\boldsymbol{\epsilon}$ is an (8×1) vector of errors with $E(\boldsymbol{\epsilon}) = \mathbf{0}$, $\operatorname{cov}(\boldsymbol{\epsilon}) = \Theta_\epsilon$, a diagonal positive-definite matrix.

A similar measurement model might be proposed for the X's (determinants of performance) and the latent variables in the two measurement models linked via a linear structural relationship as in the LISREL system (see Joreskog 1977 and Joreskog and Sorbom 1981). We find, however, that the correlations among a set of X's chosen on the basis of previous studies are extremely small; therefore, we take the X's used in our structural models to be the true determinants of firm performance. Using the LISREL notation, we set

$$\mathbf{X} \equiv \boldsymbol{\xi}, \tag{8}$$

where $\boldsymbol{\xi}$ is a (4×1) vector of the causes of performance with $E(\boldsymbol{\xi}) = \mathbf{0}$ and $\operatorname{cov}(\boldsymbol{\xi}) = \Phi$.

We assume that the η's and the ξ's are related by the linear structural model

$$\boldsymbol{\eta} = \Gamma \boldsymbol{\xi} + \boldsymbol{\zeta}, \tag{9}$$

with Γ a (3×4) matrix of structural coefficients and ζ a (3×1) error vector with $E(\zeta) = \mathbf{0}$ and $\text{cov}(\zeta) = \Psi$, a symmetric positive-definite matrix. Finally, we take the errors in models (7), (8), and (9) to be uncorrelated with each other and with the other variables in their individual equations.

We are interested in fitting the models (7), (8), and (9) to the observed (standardized) data $[\mathbf{Y}', \mathbf{X}']$. Since the η's are not observed, we cannot estimate the coefficient matrices Λ_y and Γ directly. Assuming our specifications are correct, however, certain covariance relationships must hold.

Given (7), (8), (9), and the additional assumptions concerning the errors,

$$\underline{\rho} = \text{corr}\left(\left[\frac{\mathbf{Y}}{\mathbf{X}}\right]\right) = \left[\begin{array}{c|c} \Lambda_y(\Gamma\Phi\Gamma + \Psi)\Lambda_y' + \Theta_\epsilon & \Lambda_y\Gamma\Phi \\ \hline \Phi\Gamma'\Lambda_y' & \Phi \end{array}\right]. \quad (10)$$

Partitioning the sample correlation matrix, R, constructed from observations on $\mathbf{Y}$ and $\mathbf{X}$ in a conformable manner, we obtain

$$R = \left[\begin{array}{c|c} R_{YY} & R_{YX} \\ \hline R_{XY} & R_{XX} \end{array}\right]. \quad (11)$$

Models (7), (8), and (9) can be fit to the data by estimating the parameters appearing on the right side of (10) using the information in R. The fitting can be done with the LISREL V program of Joreskog and Sorbom (1981). Denoting estimates by carets,

$$\hat{\underline{\rho}} = \left[\begin{array}{c|c} \hat{\Lambda}_Y(\hat{\Gamma}\hat{\Phi}\hat{\Gamma} + \hat{\Psi})\hat{\Lambda}_Y' + \hat{\Theta}_\epsilon & \hat{\Lambda}_Y\hat{\Gamma}\hat{\Phi} \\ \hline \hat{\Phi}\hat{\Gamma}'\hat{\Lambda}_Y & \hat{\Phi} \end{array}\right], \quad (12)$$

and one measure of model adequacy is supplied by the (average) size of the entries in the residual matrix $R - \hat{\underline{\rho}}$.

A list and short description of 13 potential causes of firm performance included for analysis is provided in Appendix A. (A theoretical discussion concerning the appropriateness of these measures is provided in Hirschey 1982 and Ravenscraft 1983.)

A preliminary correlation analysis and several univariate (in the response variable) regressions reveal that, at most, only 4 or 5 of the 13 potential X's are indeed related to the measures of firm performance, Y's, discussed before. Coefficients of determination for the univariate regressions were typically in the 30%–40% range. Potentially important X's were selected because of their (highly) significant t values. The X's selected for further study include $X_1 = RD/S$, $X_2 = TV/S$, $X_3 = LEV$, and $X_4 = IG$.

The correlation matrices R_{XY} and R_{XX}, for all $n = 386$ firms in our sample, are displayed in Tables 4 and 5, respectively. (Recall that R_{YY} is given in Table 1.)

As noted before, our model is a special case of Joreskog and Sorbom's LISREL model. We used the LISREL V program to generate least squares (LS) and maximum likelihood estimates of the model parameters. The maximum likelihood estimates are derived from a multivariate normal likelihood function. The fitting criterion for LS is

$$\min_{\theta} F(\theta) = \min_{\theta} (\tfrac{1}{2})\text{trace}[(R - \underline{\rho})^2], \quad (13)$$

where θ denotes the vector of all unknown parameters.

There is little difference between the maximum likelihood and the LS estimates. The LS estimates of the model parameters are displayed in Appendix B and Appendix C. We present the LS estimates because the LISREL V program does not constrain parameter estimates to admissible regions and the maximum likelihood estimates of $\text{var}(\epsilon_5)$ and $\text{var}(\epsilon_7)$ are negative ($-.002$ in both cases; see the corresponding LS estimates in Appendix B). The standard errors displayed in Appendix B are those obtained from the normal theory information matrix. We note that these sample standard errors are not much different from $1/\sqrt{n} = .05$. (Our data do not appear to be normally distributed. Thus the estimates that maximize the normal theory likelihood function might be called quasi-maximum likelihood estimates. The standard errors given in Appendix B must be interpreted with some caution, although the significance of the estimates displayed does not appear to be in doubt.)

It is apparent from Appendix B and Appendix C that our model very nearly reproduces the observed correlation matrix R and offers a meaningful explanation for our data. (We note that the indeterminancy in the model caused by the scales of the unobserved η's is removed by fixing one of the parameters in each column of Λ_Y to be .9. In addition, we actually consider a constrained version of our model obtained by setting some of the elements of Λ_Y and Γ equal to zero. The choices were suggested by the correlations between the Y's and X's and previous runs of the LISREL program. Finally, we note that $\hat{\Phi} = R_{XX}$. All unknown parameters in our model appear to be identified.) The fit of the measurement model is particularly impressive. Concentrating on $\hat{\Lambda}_Y$, we find that the market measures have large "loadings" on latent variable η_1. For the most part, the accounting measures return on assets and return on equity measures latent variable η_2. Finally, accounting return-on-sales measures load highly on latent variable η_3. We conclude that the three groups of variables are measuring, generally speaking, three different but correlated performance constructs. As expected, this conclusion is roughly consistent with our factor analysis in Section 2.

Examining the structural model, we see that all three η's are related to X_3 (leverage). As leverage increases, the values of the true performance variables decrease, ceteris paribus. In addition, all of the η's are positively

Table 4. Correlation Matrix R_{XY} (n = 386 firms)

	$\pi/A(Y_1)$	$\pi/E(Y_2)$	$\pi/S(Y_3)$	$\pi/A_R(Y_4)$	$\pi/E_R(Y_5)$	$\pi/S_R(Y_6)$	$Q(Y_7)$	$EV/S(Y_8)$
RD/S (X_1)	.209	.084	.380	.341	.243	.414	.422	.437
TV/S (X_2)	.232	.099	.120	.169	.106	.115	.348	.305
LVG (X_3)	−.597	−.310	−.531	−.505	−.377	−.505	−.401	−.343
IGR (X_4)	.096	.097	.241	.104	.092	.195	.090	.116

related to X_1 (R & D intensity), although the coefficient of η_2 on X_1 is fairly small. Variable X_2 (TV advertising intensity) seems to affect only η_1, and variable X_4 (industry growth) seems to affect only η_3. We conclude that two potentially strong determinants of firm performance are leverage and R & D intensity. Television advertising and industry growth also appear to determine performance, but to a lesser extent than the other two variables.

The total effects of the X's on the Y's are given by the entries in the matrix product $\hat{\Lambda}_Y\hat{\Gamma}$. These effects are reported in Appendix C. The marked coefficients have absolute values that exceed twice their asymptotic standard errors. Standard errors for the coefficients in $\hat{\Lambda}_Y\hat{\Gamma}$ are not available directly from the LISREL V program. Approximate standard errors might be calculated from approximation formulas for the sum of products of correlated random variables using the (estimated) variances and covariances of the individual estimates in $\hat{\Lambda}_Y$ and $\hat{\Gamma}$. These approximate standard errors may be somewhat crude and are, therefore, omitted from Appendix C. (Sample calculations for the approximate standard errors yield values in the neighborhood of .04). Chamberlain (1982) gave standard errors for minimum-distance estimators constructed from data with distributions admitting four finite moments. These estimators are equivalent to the LISREL LS estimators for a direct regression of **Y** on **X**. The LISREL LS estimates of the coefficients in a direct regression of **Y** and **X** are of slightly different magnitudes, but they are generally consistent with the estimates (including the structural zeros) displayed in Appendix C. The direct-regression results and the approximate calculations discussed earlier form the basis for the market entries in Appendix C.

The results in Appendix C provide a description of the linkage between the observable causes and the imperfect indicators of profits. Again, leverage and R & D intensity are seen to be important determinants (from our initial list of possibilities) of firm performance. The leverage effect on profit performance seems quite reasonable, since higher leverage connotes greater risk, which will tend to reduce the discounted present value of the future profit stream as represented by Q and EV/S. The role of R & D as a cause of profitability supports earlier findings of an important effect of R & D on both accounting and market-based estimates of profitability (e.g., see Manfield et al. 1977 and Hirschey and Weygandt in press). We note, however, that leverage (X_3) has a relatively smaller effect on market indicators, whereas R & D intensity (X_1) has a relatively smaller effect on accounting indicators of performance.

Industry growth (X_4) does not affect the market indicators (because of zeros in the relevant positions in $\hat{\Lambda}_Y$ and $\hat{\Gamma}$) and has a small, but significant, effect on the accounting return-on-sales indicators. Others have observed a profitability effect of industry growth, and it has often been interpreted as evidence of short-run disequilibrium conditions in which unanticipated changes in prices and costs lead to short-run profits (Strickland and Weiss 1976).

Television advertising (X_2) appears to only affect the market indicators. The superior profitability of television, as opposed to nontelevision, advertising has also been a subject of interest in the literature. This finding is usually attributed to the importance of advertising as a successful marketing strategy when low-price, repeat-purchase consumer nondurable goods are considered. As in the case of R & D, profitability effects of television advertising also undoubtedly reflect the fact that these expenditures are inherently risky and accompanied by substantial required rates of return (Hirschey 1982).

We note that the signs and magnitudes of the estimates of the total effects (Appendix C) of **X** on **Y**, derived from our restricted structural model, are consistent with the LS estimates of these effects obtained from a multivariate regression of the Y's on the X's. The latter corresponds to the unrestricted reduced form of our initial model. (The results of the regression analysis are available from the authors.)

We find it interesting that no important influence of traditional market-structure variables emerged in the analysis. Although perhaps surprising, this result is consistent with findings reported by Ravenscraft (1983) and others. A schematic representation of the estimated relationship among the performance variables and the determinants of performance is shown in Figure 1.

Our LISREL model was also fit to the data from the four industry groups introduced in Section 3. We find that for the chemicals and transportation-equipment industries, the model fit is comparable to the fit for all firms described in Appendix B. Of course, there are

Table 5. Correlation Matrix R_{XX} (n = 386 firms)

	RD/S	TV/S	LVG	IGR
RD/S	1.000		(symmetric)	
TV/S	.068	1.000		
LVG	−.184	−.165	1.000	
IGR	.127	−.076	−.143	1.000

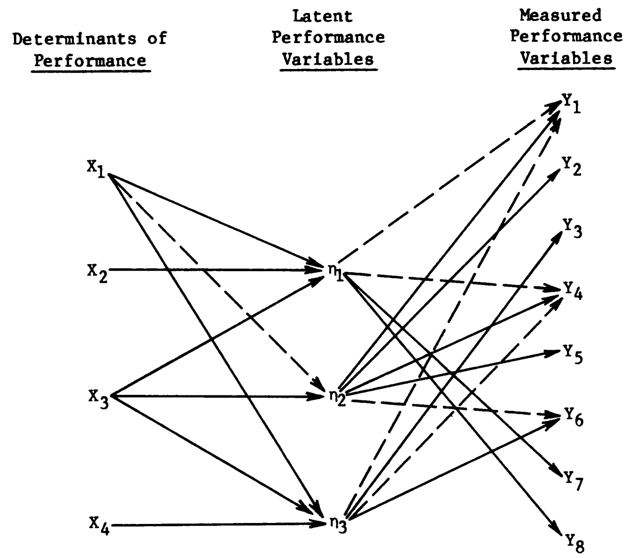

Figure 1. A Schematic Representation of the Estimated Relationships Among Performance Variables and Determinants of Performance. — → indicates weaker effects than ——→.

some changes in the estimated coefficients.

The original model does not seem to be right for the food-processing and primary-metals industries. There is some evidence from a correlation analysis that a slightly different set of determinants of performance (X's) may be appropriate for these industries. For example, for primary metals, diversification and industry size appear to be more important determinants of performance than TV-advertising intensity and R & D intensity.

We interpret our model fitting by industry results to suggest that there is some homogeneity (as measured by our model for all firms) among industries, but enough differences exist to suggest that there is also a slight industry effect. That is, differences among firms, with respect to the overall relationship between performance and determinants of performance, depend to some extent on industry as well as on firm-specific factors. This conclusion is consistent with the evidence presented in Sections 2 and 3.

5. CONCLUSIONS

This study analyzes the consistency, determinants, and uses of accounting and market-value measures of profitability. Whereas accounting-income numbers have a primarily historical interpretation, market-value measures of profitability are expectational or forward looking. Comparing accounting and market data provides us with a useful dual perspective on the firm-profit performance phenomenon.

In the first part of the analysis, a factor analysis was conducted to evaluate the consistency of accounting versus market estimates of profitability. It is interesting that both the Q and (EV/S) market-value measures load highly and about equally on a unique underlying unobservable firm-profitability characteristic. Furthermore, these observed differences between accounting and market estimates cannot be explained as reflecting uncontrolled industry-specific influences. A factor analysis conducted over firms in various two-digit industry groups reveals the same pattern of differences between accounting and market estimates of profitability as was evident in the overall sample. A lack of industry effects was also suggested by a K-means cluster analysis. When grouped according to similarities between various profit measures, there was no marked tendency towards firms grouping along industry lines.

These differences between accounting and market measures of profitability suggest the validity of cautioning remarks concerning the use (misuse) of accounting data. Researchers primarily concerned with the causes and consequences of monopoly profits would seem well advised to adopt the Q ratio or (preferably) EV/S instead of more traditional accounting-income data. If one takes the reasonable view that accounting and market data are, together, attractive but imperfect profit measures, a multivariate approach to the study of firm performance is suggested. Taking accounting and market estimates of profitability as imperfect indicators of three unobservable performance characteristics, we find a significant explanatory role for R & D intensity, television advertising, leverage, and industry growth as determinants of profitability. It is interesting that no important role for traditional market-structure variables was discovered. These findings suggest that firm performance may be closely tied to firm-specific influences rather than to structural measures of competition.

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APPENDIX A: POTENTIAL CAUSES OF FIRM PERFORMANCE

Investment Characteristics Variables

S = Size measured by total net sales.

D = Diversification measured as $D = 1 - \sum_{j=1}^{m} S_j^2$, where S_j represents the share of firm sales in the jth four-digit census industry.

RD/S = R & D intensity.

TV/S = Television advertising intensity.

NTV/S = Nontelevision advertising intensity.

Risk Variables

LEV = Leverage measured by the book value of debt to market value of common ratio.

B = Stock-price beta measure of systematic risk.

Market Structure Variables

MS = Market share weighted to reflect firm sales in m four-digit census industries, where MS

$= \sum_{i=1}^{m} \sum_{j=1}^{m} S_i S_j$ and S_i is firm market share in the ith industry.

CR = Concentration weighted to reflect firm sales in m four-digit census industries, where $CR = \sum_{j=1}^{m} S_j CR_j$ and CR_j is the four-firm concentration ratio in the jth industry.

MS/CR = Relative firm size.

Demand Characteristics Variables

IS = Industry size measured by total value of shipments.

IGR = Industry growth measured as the 10-year average annual rate of growth in primary product industry value of shipments.

CD = (0, 1) consumer demand dummy variable.

APPENDIX B: A SUMMARY OF THE LINEAR STRUCTURAL MODELING

Measurement Model

$$\mathbf{Y} = \Lambda_Y \eta + \epsilon.$$

$$\hat{\Lambda}_Y = \begin{bmatrix} .255 & .583 & .224 \\ (.033) & (.040) & (.044) \\ 0 & .9^* & 0 \\ 0 & 0 & .9^* \\ .204 & .699 & .227 \\ (.025) & (.036) & (.033) \\ 0 & 1.034 & 0 \\ & (.038) & \\ 0 & .207 & .799 \\ & (.039) & (.036) \\ .9^* & 0 & 0 \\ .861 & 0 & 0 \\ (.018) & & \end{bmatrix};$$

$$\hat{\text{cov}}(\epsilon) = \hat{\Theta}_\epsilon$$
$$= \text{diag}[.218\ .295\ .189\ .105\ .070\ .130\ .032\ .113].$$

Asterisks indicate parameter values fixed to determine a scaling for the η's. Standard errors for parameter estimates are in parentheses.

Square multiple correlations for Y's; $1 - ((\hat{\text{var}}(\epsilon_i)/\hat{\text{var}}(Y_i)) = 1 - (\hat{\theta}_{ii}/1)$:

Y_1	Y_2	Y_3	Y_4	Y_5	Y_6	Y_7	Y_8
.782	.705	.811	.895	.930	.870	.968	.887.

Total coefficient of determination for Y's; $1 - (|\hat{\Theta}_\epsilon|/|R_{YY}|) = .996$ ($|A|$ denotes determinant of matrix A).

Structural Model

$\eta = \Gamma\xi + \zeta$, with $\xi \equiv \mathbf{X}$, $\text{cov}(\zeta) = \Psi$, $\text{cov}(\xi) = \text{cov}(\mathbf{X}) = \Phi$, $\text{cov}(\eta) = \Gamma\phi\Gamma' + \Psi$.

$$\hat{\Gamma} = \begin{bmatrix} .398 & .301 & -.325 & 0 \\ (.046) & (.040) & (.046) & \\ .091 & 0 & -.384 & 0 \\ (.045) & & (.046) & \\ .327 & 0 & -.511 & .126 \\ (.042) & & (.043) & (.035) \end{bmatrix};$$

$$\hat{\Phi} = R_{XX}.$$

Standard errors for parameter estimates are in parentheses.

$$\hat{\Psi} = \hat{\text{cov}}(\zeta) = \begin{bmatrix} .745 & \text{(symmetric)} & \\ .186 & .702 & \\ .360 & .360 & .527 \end{bmatrix};$$

$$\hat{\text{cov}}(\eta) = \begin{bmatrix} 1.195 & \text{(symmetric)} & \\ .401 & .870 & \\ .754 & .583 & 1.001 \end{bmatrix}.$$

Squared multiple correlations for η's; $1 - (\hat{\text{var}}(\zeta_i)/\hat{\text{var}}(\eta_i))$:

η_1	η_2	η_3
.377	.194	.474.

Total coefficient of determination for structural equations:

$$1 - (|\hat{\Psi}|/|\hat{\text{cov}}(\eta)|)$$
$$= 1 - (|\hat{\Psi}|/|\hat{\Gamma}\hat{\Phi}\hat{\Gamma}' + \hat{\Psi}|) = .597.$$

Measure of goodness of fit of the whole model: goodness-of-fit index;

$$1 - \left(\frac{\text{trace}(R - \underline{\hat{\rho}})^2}{\text{trace } R^2}\right) = .995.$$

Root mean square residual;

$$\left[2 \sum_{i=1}^{12} \sum_{j=1}^{i} (r_{ij} - \hat{\rho}_{ij})^2/12(13)\right]^{1/2} = .036.$$

[Note: $R = \{r_{ij}\}$; $\underline{\hat{\rho}} = \{\hat{\rho}_{ij}\}$.]

APPENDIX C: ESTIMATED TOTAL EFFECTS MATRICES FOR LINEAR STRUCTURAL MODEL

Total effects of X's on η's: $\hat{\Gamma}$ (see Appendix B).
Total effects of η's on Y's: $\hat{\Lambda}_Y$ (see Appendix B).
Total effects of X's on Y's: $\hat{\Lambda}_Y\hat{\Gamma}$.

$\hat{\Lambda}_Y\hat{\Gamma}$:	X_1	X_2	X_3	X_4
Y_1	.228*	.007	−.421*	.028
Y_2	.082	0	−.345*	0
Y_3	.294*	0	−.460*	.113*
Y_4	.219*	.061	−.450*	.028
Y_5	.094	0	−.397*	0
Y_6	.280*	0	−.488*	.101*
Y_7	.358*	.271*	−.292*	0
Y_8	.343*	.259*	−.280*	0

Asterisks indicate that absolute value exceeds twice the asymptotic standard error.

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