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Human Capital and the Financial Capital Market

Employing a popular extension of the Sharpe-Linter capital asset pricing model which allows for the existence of nonmarketable human capital, this study finds that empirically the inclusion of human capital appears to have little meaningful effect upon both general capital asset pricing and individual investor portfolio composition. This is shown to arise from the fact that relationships between returns on almost all types of human capital and those of marketable financial assets are so weak as to make these two capital asset groupings effectively separable.

I. Introduction

In the past 2 decades, much progress has been made in the areas of "human capital theory" and "modern portfolio theory," with profound influence upon academic thought and practice.[1] But interestingly enough, though human capital theory recognizes human resources as part of an individual's capital asset holdings and modern portfolio theory deals with the pricing of these holdings, only recently have efforts been made to bring these two areas together.[2]

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1. For a discussion of "human capital theory" and its impact, see, e.g., Becker (1964), Ben-Porath (1967), and Schultz (1971). As for "modern portfolio theory," a discussion of its effect upon the academic community can be found in Jensen (1972); while with respect to the nonacademic financial community, see Welles (1973).

2. The implied assumption in the earlier studies which have ignored human capital is that the observed sample of risky assets used is so typical of the universe of risky assets including human capital that it captures all but insignificant effects (see Mayers 1972, p. 231). It is interesting to note that Markowitz (1959, p. 298–99), the earliest pioneer of modern portfolio theory, was aware of the human capital problem and did devote

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This paper represents an attempt at joining these two areas by exploring empirically the effect of human capital upon (1) the *pricing* of risky financial assets and (2) the *composition* of individual investor risky financial asset portfolios. Building upon recent work of David Mayers (1972, 1973) and Fama and Schwert (1977), we develop an empirical methodology by which to study the human capital effect and then test it using a number of very different sets of data.

This paper consists of four additional sections. Section II briefly reviews the theoretical framework developed by Mayers for human capital and shows that it is only the covariation between the return on human and financial capital which would be responsible for any effect upon pricing or portfolio composition. Section III discusses the empirical methodology used. Section IV describes the data and presents the test results, and Section V reviews our findings and discusses possible directions for further research.

II. The Theoretical Background

A. *Risky Financial Asset Pricing*

A primary component of modern portfolio theory is the Sharpe (1964) and Lintner (1965) capital asset pricing model:[3]

$$E(\tilde{R}_{jt}) = R_{ft} + [E(\tilde{R}_{mt}) - R_{ft}]\beta_j, \tag{1}$$

where R_{ft} is the rate of return on a risk-free security purchased at time $t - 1$ and sold at time t; $E(\tilde{R}_{jt})$ is the expected rate of return on any asset or portfolio j held from $t - 1$ to t; $E(\tilde{R}_{mt})$ is the expected rate of return on a market portfolio m consisting of an investment in every asset in the market in proportion to the total market value contributed by the asset; and

$$\beta_j \equiv \frac{\text{cov}(\tilde{R}_{jt},\tilde{R}_{mt})}{\sigma^2(\tilde{R}_{mt})} \tag{2}$$

is asset j's systematic risk measure representing its relative contribution to the risk of the market portfolio m embodied in m's variance.

The Sharpe-Lintner (SL) model assumes that the capital market is perfect and that all assets are perfectly divisible and liquid. Dropping this perfect liquidity assumption, and allowing assets to be either perfectly liquid or perfectly illiquid, David Mayers (1972, 1973) demonstrated that the SL model would remain intact except for the system-

some discussion to it. Besides the work of Mayers (1972, 1973) and Fama and Schwert (1977), which we shall discuss shortly, recent theoretical work has been done by DeBrito (1975) and Williams (1978), and recent empirical work has been done by Friend and Blume (1975), Landskroner (1977), and Cohn et al. (1975).

3. For a discussion of the development of the Sharpe and Lintner model, see Jensen (1972).

atic risk measure term. Instead of β_j, Mayers's model contains β_j^*:

$$\beta_j^* \equiv \beta_j \left\{ \left[1 + \frac{\operatorname{cov}(\tilde{R}_{jt}, \tilde{H}_t)}{V_{m,t-1} \operatorname{cov}(\tilde{R}_{jt}, \tilde{R}_{mt})} \right] \Big/ \left[1 + \frac{\operatorname{cov}(\tilde{R}_{mt}, \tilde{H}_t)}{V_{m,t-1} \sigma^2(\tilde{R}_{mt})} \right] \right\}, \qquad (3)$$

where $\tilde{H}_t$ is the total dollar return on all illiquid assets at time t; $V_{m,t-1}$ is the total market value of all liquid assets at time $t - 1$; while all other variables as well as β_j are defined as before.

If we compare expression (3) with (2), we find that the effect upon pricing of including nonmarketable human capital is only due to the presence of two additional covariance terms in (3) relating aggregate human capital to the asset being priced and the market portfolio that is used. If these covariance relationships are strong, human capital should have a meaningful effect upon capital asset pricing; if these relationships are weak, no meaningful effect should be expected. This is a question amenable to empirical verification and is addressed later in this paper.

B. *Individual Portfolio Composition*

Following the Sharpe and Lintner model, an individual's investment problem is that of choosing proportions of each asset to invest in so as to maximize his single-period expected utility of terminal wealth. Assuming the individual to be risk averse, his vector of optimal investment proportions would be:[4]

$$X_i = \frac{1}{2} \left(\frac{\partial \sigma_i^2}{\partial E_i} \right) \Sigma^{-1} [E(\tilde{V}_t) - (1 + R_{ft}) V_{t-1}], \qquad (4)$$

where X_i is the optimal investment proportions vector for investor i; E_i is the expected market value (including dividends) at time t of all investments held by investor i from $t - 1$ to t; σ_i^2 is the variance of this market value as of time t; $E(\tilde{V}_t)$ is the vector of expected individual asset values as of t; V_{t-1} is the vector of individual asset values as of time $t - 1$; Σ^{-1} is the inverse of the variance-covariance matrix of asset values as of t; and R_{ft} is the risk-free rate of return. Equation (4) contains only one element unique to investor i: $(\partial \sigma_i^2)/(\partial E_i)$, a scalar.

Since the only term unique to investor i is a scalar, the SL model implies that no differences should exist between investors in the composition of their risky financial asset portfolios. Investors should differ only in the proportion of total wealth invested in risky assets, while the proportions in each risky asset should be the same.

But if we allow for nonmarketable human capital which differs from investor to investor, this conclusion of portfolio composition uniformity does not necessarily hold true. Mayers (1972) demonstrates

4. For a derivation of expressions (4) and (5), see Mayers (1972).

that, by including human capital, instead of (4) the optimal investment proportions vector should be

$$X_i = \frac{1}{2}\ \left(\frac{\partial \sigma_i^2}{\partial E_i}\right) \Sigma^{-1}\,[E(V_t) - (1 + R_{ft})V_{t-1}] - \Sigma^{-1}\,\Sigma_i^H, \tag{5}$$

which is identical with (4) except for the addition of $-\Sigma^{-1}\,\Sigma_i^H$, where Σ_i^H is the column vector of covariances between investor i's human capital cash flow, $\tilde{H}_i$, and random market values of the individual liquid assets, V_t.

The vector of optimal investment proportions now contains not only a scalar term unique to investor i, but also an entire vector, Σ_i^H. Theoretically, then, investors should take their human capital into consideration when constructing their portfolios, and we should expect differences to exist between investor portfolios if the pattern of their respective human capital covariance vector values differ meaningfully.[5]

Since $\tilde{V}_{jt} = V_{j,t-1}\,(1 + \tilde{R}_{jt})$, we may write Σ_i^H in the following way:

$$\Sigma_i^H \equiv \begin{pmatrix} \mathrm{cov}(\tilde{H}_i, \tilde{R}_{1t}) \cdot V_{1,t-1} \\ \vdots \\ \mathrm{cov}(\tilde{H}_i, R_{Nt}) \cdot V_{N,t-1} \end{pmatrix}.$$

In this form the covariance terms which affect portfolio composition are now identical with those covariances which affect price in equation (3), except that the former represent individually held human capital while the latter represent human capital in the aggregate. The effect of human capital upon pricing and portfolio composition is, thus, only by way of its covariance relationship with the returns on marketable financial assets. The strengths of these covariances determine how important a factor human capital is. In the following sections, our primary focus shall be upon these covariances.

III. The Empirical Methodology

A. The Approach

A simple approach by which to assess the effect of human capital upon pricing was first used by Fama and Schwert (1977).[6] They observed

5. What we mean by individual investor portfolio composition is only the risky financial asset portion, for even if one's human capital were entirely riskless, it would surely have some effect upon the entire portfolio, at least in terms of replacing the riskless asset portion.

6. There are two alternative approaches that could be followed. First, a sample of risky asset portfolios and investor occupations could be examined to determine whether any relationship exists between portfolio composition and occupation. This method offers a direct test of the portfolio composition implication and indirectly tests the effect upon pricing. The drawback of this approach is that time-series data of the type necessary are not currently available. The second approach is to directly test for an effect upon

that, since the only difference between the SL model and the Mayers model was their systematic risk measure, all that need be done to measure the effect of human capital upon pricing is to estimate the differences between the Mayers and SL risk measures for various classes of financial securities. If these differences are found to be large, human capital should be an important factor in pricing; if not, human capital should be deemed inconsequential.

Using a measure of the return to aggregate human capital for the entire U.S. labor force, Fama and Schwert estimated these differences for various portfolios of NYSE common stock and U.S. Treasury bills and bonds and found them extremely small. They thus concluded that with regard to pricing "it is unnecessary to take account of . . . human capital."[7]

Since Fama and Schwert used aggregate human capital, they could address only the pricing question but not the portfolio composition question. We reason that with regard to the composition question the same basic approach can also be used, but in place of the return to aggregate human capital the return to individually held human capital should be used. As was demonstrated in the previous section, the difference between the Mayers and SL risk measures lies in the addition of two covariance terms in the Mayers model, $\text{cov}(\tilde{R}_{jt},\tilde{H}_t)$ and $\text{cov}(\tilde{R}_{mt},\tilde{H}_t)$. These covariance terms were shown to be the same ones affecting portfolio composition, except that, with regard to portfolio composition, the $\tilde{H}_t$'s represent the return on individually held rather than aggregate human capital. If we then duplicate the Fama and Schwert approach, using individually held human capital, and still find little difference between β_j^* and β_j, it would imply that human capital is inconsequential not only in pricing but also in portfolio composition. Small differences between β_j^* and β_j would mean that the additional two covariance terms in β_j^* which are also responsible for the human capital effect upon portfolio composition are small and insignificant and may be ignored by the individual investor.[8]

In the empirical part of the paper, we apply this approach using various sets of human capital data which are distinguished by industry affiliation, occupation, and level of educational attainment. In this way, we hope to capture the diversity that exists among the human capital

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the structure of asset prices, i.e., estimate the Mayers systematic risk measures, and test whether the postulated theoretical relationship between expected return and risk is observed. This method would test directly the effect upon asset pricing and indirectly test the effect upon an individual's portfolio composition. The method we use is similar to the second but simpler and more direct.

7. Fama and Schwert (1977, p. 120).

8. This could also happen even if the covariance values are large, in the very unlikely case where the covariance ratio, $\text{cov}(R_{jt},h_t)/\text{cov}(R_{jt},R_{mt})$ is the same for all securities. This ratio would then necessarily be equal to the ratio $\text{cov}(R_{mt},h_t)/\sigma^2(R_{mt})$, and β_j would equal β_j^* for all j. No apparent risk measure effect would then be present as well.

traits of individual investors, and possibly identify which human capital trait, if any, exerts an effect.

B. *The Statistical Technique*

To estimate the difference between Mayers's β_j^* and Sharpe-Lintner's β_j, we first need to estimate each measure separately. To estimate β_j, we simply use the slope coefficient from a linear "market model" regression of our individual portfolio rates of return on the rates of return of our market portfolio.[9] But to estimate β_j^* we were faced with a minor problem. To use equation (3), besides estimating β_j we must also estimate $\text{cov}(\tilde{R}_{jt},\tilde{H}_t)$ and $\text{cov}(\tilde{R}_{mt},\tilde{H}_t)$ from time-series data, requiring that their component distributions be stationary through time. Though distributions of R_{jt} and R_{mt} have been found to be stationary in earlier studies, H_t, when identified with labor earnings as we do here, suffers from an apparent upward trend with sizable levels of autocorrelation.[10] To deal with this problem, we use per capita earnings data for H_t when measuring the variation in the return to human capital over time so as to abstract from changes in the size of the labor force, and a modified version of (3) provided by Fama and Schwert (1977):

$$\beta_j^* = \beta_j \cdot \left\{\left[1 + \frac{H_{t-1}\,\text{cov}(\tilde{R}_{jt},\tilde{h}_t)}{V_{m,t-1}\,\text{cov}(\tilde{R}_{jt},\tilde{R}_{mt})}\right] \Big/ \left[1 + \frac{H_{t-1}\,\text{cov}(\tilde{R}_{mt},\tilde{h}_t)}{V_{m,t-1}\sigma^2(\tilde{R}_{mt})}\right]\right\}, \tag{6}$$

which contains a differenced form of H_t,h_t, the percentage change in per capita earnings over period t, with all other variables defined as before except for H_{t-1}, which is the total dollar return on human capital in period $t-1$. In place, these adjustments were found to relieve both the upward trend and high autocorrelation of the H_t variable.

Since the distributions of R_{jt}, R_{mt}, and h_t, the only random variables in equation (6), are now stationary, their parameters may be viewed as stable as well. We thus estimate β_j^* by deriving the appropriate sample covariances, solving for β_j, and plugging their estimated values into expression (6).[11] We can then compare the β_j^* estimates with their corresponding β_j's and determine whether any meaningful differences exist.

9. The term "market model" was first used by Fama (1968) to refer to the model originally referred to by Sharpe (1963) as the "diagonal model." The market model postulates that $R_{jt} = a_j + b_j R_{mt} + e_{jt}$, where R_{mt} is a market factor. Fama (1968) demonstrates that b_j is approximately equal to the Sharpe-Lintner measure of systematic risk, β_j,$\text{cov}(R_j,R_m)/\text{var}(R_m)$.

10. Fama (1965), and Fama and Schwert have demonstrated that the rates of return on financial assets appear to be generated by a stationary distribution.

11. Note that even though the variable components of eq. (6) carry time subscripts, the β^* coefficients do not. We do this because later in the paper we use mean values over time for the variable components, so that what we actually compute are mean values over time for the coefficients rather than estimates as of a particular point in time. The use of time subscripts is purely expository.

IV. The Data and Results

A. The Data

To use equation (6) to estimate β_j^*, we need time-series data of (1) per capita labor earnings for $\tilde{h}_t$, (2) rates of return on different classes of marketable securities for $\tilde{R}_{jt}$, (3) rates of return on a market portfolio of securities for $\tilde{R}_{mt}$, and (4) aggregate human capital earnings and financial asset values for $H_{t-1}/V_{m,t-1}$.

In order to capture the diversity of individually held human capital, we use three different sets of per capita earnings data for $\tilde{h}_t$. To examine the effect of industry affiliation, we use Bureau of Labor Statistics data of monthly per capita production worker earnings for all eight industry classifications reported for in BLS's *Employment and Earnings: United States, 1909–72*.[12] For the effect of occupation type, we use a time series of median annual per capita labor earnings data for men classified by occupation collected from the annual issues of the Bureau of the Census's *Current Population Reports: Consumer Income* (Series P-60) beginning with 1958.[13] All occupational classifications reported for are used except for those whose content was not consistent over time or was not available for the entire period.[14] And for the effect of educational attainment, we use the University of Michigan's *Panel Study of Income Dynamics* (Morgan and Smith 1972). This consists of annual longitudinal labor earnings data by individual from a representative sample of white male Americans for the years 1967 through 1974 arranged by years of school attended.[15] Being ungrouped, the University of Michigan data should be of special interest, for they should allow us to observe more directly the effect of individually held human capital than is possible with the grouped industry and occupational data.

There are many questions that can be raised about both our use of labor earnings data as estimates of the return to human capital and our testing procedure. Since our work is essentially a time-series analysis, we are effectively assuming that the nature and quality of human

12. A description of the BLS industry data can be found in U.S. Bureau of Labor Statistics 1974. The data consist primarily of wage disbursements to predominantly blue-collar-type employees and include only workers actually employed. To adjust the data to per capita for the entire industry labor force, we used monthly industry unemployment rates as reported in the February 1973 issue of the monthly *Employment and Earnings Journal*. Though it might be objected that industry unemployment rates are unreliable because of possible crossover from one industry to another, Feldstein (1975) has found that there appears to be considerable industry attachment and little crossover.

13. A detailed description of the *Current Population Report* data can be found in U.S. Bureau of the Census (1958–72).

14. A detailed description of the general occupational classifications used can be found in U.S. Bureau of the Census (1972).

15. For a detailed description of these data and the procedures used in collecting them, see Morgan and Smith (1972).

capital are fixed over time, which they may not be. In addition, there is the fundamental question of the nonmarketability of human capital and its treatment in a one-period context. Since these questions and others have already been dealt with in depth by Fama and Schwert (1977), for the sake of brevity we refer the reader there for further discussion.

For the rates of return on marketable securities, R_{jt}, we use four different groups of marketable securities to study the industry and occupation effects and only one group for the educational attainment effect. The four are (1) a value-weighted index of NYSE stocks for the period 1947–72 constructed by Myron Scholes, (2) a government bond portfolio developed by Fama and Schwert and used in their paper (1977) to approximate a value-weighted portfolio of government debt for the period from 1947 to 1972,[16] (3) a set of 10 common stock portfolios formed according to the ranked values of their SL risk measure estimates from all stocks listed on the NYSE for the period from March 1931 to June 1970 and first used in the paper by Black, Jensen, and Scholes (1972) (BJS), and (4) a selection of industry portfolios formed according to their SEC industrial classifications from NYSE listed stocks by James MacBeth (1974) for the period of July 1945 to June 1968. The group used for the educational attainment effect is a set of 10 value-weighted common stock portfolios formed according to the ranked value of their rate of return variance estimates from all NYSE listed stocks beginning with 1963. This last set was obtained from the Center for Research in Security Prices (CRSP) of the University of Chicago, while the previous four were provided by Eugene F. Fama.

As an estimate of the rate of return on the market portfolio, R_{mt}, we use a combination of the Scholes value-weighted NYSE portfolio and the Fama-Schwert government bond index mentioned previously when studying the industry and occupation effects. For the educational attainment effect, we use instead an annual market index that was derived by Ibbotson and Sinquefield (1976) from the Standard and Poors composite.

And finally, for the $H_{t-1}/V_{m,t-1}$ ratios of equation (6), we use the mean of a time series of these ratio values estimated by Fama and Schwert, that mean being 0.68.

Fama and Schwert note that the Markowitz (1959) portfolio investment model, the foundation of modern capital asset pricing, assumes people invest in order to eventually consume. Investors thus evaluate their investment return entirely in terms of consumption goods and services obtainable, implying that investment variables should be mea-

16. Admittedly, government debt may not necessarily represent real wealth. What is assumed is that government debt closely approximates corporate debt which is real wealth (see Barro 1974).

sured in "real" rather than "nominal" terms. Accordingly, we present our tests only after having transformed the data to real terms using procedures discussed in Fama (1975).[17] Since for most of our time series little or no data are available prior to 1947, and because Fama (1975) has observed that the CPI was less accurate prior to 1953, all SL and Mayers risk measure estimates presented use only data for periods subsequent to 1952.

B. The Results

Our estimates of the Mayers-SL risk measure differences are presented in table 1 for the monthly industry worker data, in tables 4 and 5 for the occupational data, and in tables 7 and 8 for the educational attainment data. Results are reported for the Fama-Schwert government bond index, the Scholes value-weighted NYSE stock index, the 10 Black-Jensen-Scholes risk-ranked common stock portfolios, a set of 12 selected MacBeth industry portfolios, and in tables 5 and 6 for the 10 CRSP variance ranked portfolios.

Besides the difference estimates, the tables include the SL systematic risk measure estimates for each of the portfolios listed, along with their respective *t*-statistic values in brackets. Also, because there seems to be no easy way to develop comparable measures of statistical significance for the Mayers risk measures, and certainly for the difference estimates, as an aid in appreciating the human capital effect, we provide in parentheses below each risk measure difference (except for table 6) its value expressed as a fraction of its respective SL risk measure.

The question then before us is whether there are any meaningful differences between the Mayers and SL risk measures. In reviewing our results, both the magnitude and the pattern of the difference estimates should be considered.

With regard to magnitude, one should note that the arithmetic mean monthly return on our combined bond-stock market portfolio for 1953–72 was 0.0071, while it was 0.0027 for 1-month Treasury bills during the same period. If we estimate the average value for $E(\tilde{R}_{mt}) - R_{ft}$ in equation (1), we find that the average monthly risk premium per unit of Mayers's β_j^* was only 0.0044 for this period. If we then were to incorrectly use β_j for β_j^* when the difference between the two measures was, for example, 0.10, we would understate the expected return by only 0.04 of 1% per month, which seems a rather inconsequential amount. Thus we shall use 0.10 as a guide in evaluating our results. If we find the difference estimates not much above absolute 0.10, human

17. One may object to the use of real data, for they may eliminate the systematic risk due to inflation. But Fama (1975) has shown that inflation has little if any systematic risk, so that our results should be unaffected. Nevertheless, the tests were repeated on nominal data as well with much the same result.

capital would appear to be inconsequential and unimportant. If the differences far exceed 0.10, the human capital effect would then appear to be meaningful and should certainly not be ignored.

As for pattern, if human capital has a meaningful effect, one should expect it to follow some systematic pattern. Possible patterns would be whether the estimated differences are related to the portfolio's risk level, which should be observed by studying the BJS and CRSP portfolios, or whether the inclusion of individual industry production worker earnings has any greater effect upon its own respective industry's stock portfolio. If no such discernible patterns are found, it would be further evidence of the unimportance of the human capital effect.

Turning first to industry affiliation in table 1, we find that none of the risk measure differences shown exceeds absolute 0.1012. The largest risk measure difference for the government bond index portfolio was only absolute 0.036, while for the NYSE index portfolio it was only 0.0197. Of the 80 difference estimates presented for the 10 BJS portfolios, none exceeded an absolute 0.0512, and of the 96 presented for the MacBeth portfolios, only 10 surpassed absolute 0.0605. In relative terms, only four of the BJS portfolio difference estimates exceeded 2% of their respective SL risk measures, and only 10 of the MacBeth portfolio differences exceed 3.7%.

Besides the small difference levels, we find little in terms of pattern as well. Only the nondurable and durable manufacturing industries' worker earnings appear to produce risk measure differences that are related to the portfolio's SL risk level. High beta portfolios exhibit predominantly positive differences, while low beta portfolios exhibit predominantly negative differences. But all the other industry classifications produce no such pattern.

As for any greater effect upon an industry's own stock portfolio, we find no such pattern either. Only mining has any pronounced effect upon its own portfolio. But instead of increasing the systematic risk of the mining portfolio, as would be expected if there were significant positive covariation between miner earnings and mining portfolio rates of return, the inclusion of miner earnings decreases it by more than any other portfolio.

It would thus appear that there is little important covariation between the various industries' worker earnings and the monthly return on financial asset portfolios contributing to so small a systematic risk effect.[18] This "weak" relationship may be seen even more strikingly

18. Since industry worker earnings data are drawn primarily from blue-collar-type workers whose wage rates are negotiated periodically, it may be argued that monthly data do not represent the proper time synchronization between portfolio returns and earnings-level movements and also miss important lag effects that may be present. To test this possibility, we replicated our tests using annualized versions of our industry data and found there to be little change in our results.

in table 2, which summarizes the covariance values for industry worker percentage earnings changes, h_t, and MacBeth industry portfolio rates of return. If important relationships did exist between industry worker earnings and financial asset portfolios, they should certainly be most pronounced with their own industry's stock portfolio. But in table 2 we find no such marked relationship.

Similarly, if we turn to table 3, which summarizes the slope coefficient values from a linear market model regression of h_t on the market index, R_{mt}, which we denote β_{hm}, we find only one coefficient that is significant at the 10% confidence level.

Since we find no meaningful covariation by individual industry, the industry affiliation of one's human capital, at least with the eight industry classifications shown, should have little to do with the composition of that individual's portfolio. There should thus be no meaningful human capital portfolio composition effect.[19]

And with regard to pricing, these results indicate that human capital in the aggregate has little effect here as well. For besides the smallness of the difference estimates, we also find in table 1 that none of the portfolios exhibit any pervasively one-sided effects across industries, implying in the aggregate an even smaller place for human capital.

Turning to the effect of occupational identification as reported in tables 4 and 5, we find that, although most of the occupation classifications presented show results similar to those of the previously discussed industry worker categories, five occupation classifications do appear to exert some major systematic risk effects. Most striking are the professional, technical, and kindred self-employed classifications with an effect of 0.7665 on the government bond index, −0.5672 on the NYSE index, and effects in excess of an absolute 1.3304 for eight of the 10 BJS portfolios and of 0.18 for 10 of the 12 MacBeth portfolios. The farm laborers and foremen, laborers except farmers, farmers and farm managers, and managers, officials, and proprietors self-employed classifications exhibit effects that are less dramatic but are in excess of absolute 0.1632.

It would thus appear that occupational association with any one of the five above classifications may very well influence portfolio composition, even though the other classifications should not. But this apparent inconsistency with the results using industry worker data should be expected for at least two reasons.

19. It should be noted that, since we use industry worker per capita earnings, which may not necessarily represent any single worker's earnings pattern, there may be some degree of "washing out" of possible effects from aggregating. Thus our implications concerning individual portfolio composition should be viewed with some measure of reservation.

TABLE 1 Comparisons of Mayers and Sharpe-Lintner (SL) Risk Estimate Differences across Industry Groups—Monthly Industry Worker Real Earnings Data

		Production Worker Industry Affiliation							
Marketable Security Portfolios	SL Risk Measure	Mining (1)	Const. (2)	ND (3)	D (4)	W & R (5)	T & U (6)	Fin. (7)	Services (8)
		January 1953–June 1972					February 1964–June 1972		
1. Government bond index	.0632	−.0157	−.0360	−.0091	−.0145	−.0049	−.0024	−.0018	−.0015
	[3.89]	(−.248)	(−.570)	(−.144)	(−.230)	(−.078)	(.023)	(.107)	(−.015)
2. NYSE index	1.4946	−.0196	.0197	.0072	.0044	.0070	.0024	.0008	.0013
	[86.83]	(−.013)	(.013)	(.005)	(.003)	(.005)	(−.002)	(.001)	(.001)
		January 1953–June 1970					February 1964–June 1970		
1. BJS 1	2.4375	−.0136	−.0032	.0177	.0080	.0441	.0301	.0142	.0057
	[29.89]	(−.006)	(−.001)	(.007)	(.003)	(.018)	(.013)	(.006)	(.002)
2. BJS 2	2.1126	−.0270	−.0365	.0168	−.0013	.0395	.0015	.0135	−.0029
	[34.47]	(−.013)	(−.017)	(.008)	(−.001)	(.019)	(.001)	(.007)	(−.001)
3. BJS 3	1.9356	−.0366	−.0103	.0057	−.0031	.0226	.0028	.0162	−.0041
	[35.89]	(−.019)	(.003)	(.003)	(−.001)	(.012)	(.001)	(.009)	(−.002)
4. BJS 4	1.8157	−.0231	−.0044	.0123	.0052	.0006	−.0138	.0078	−.0052
	[40.18]	(−.013)	(−.002)	(.007)	(.002)	(.000)	(−.008)	(.004)	(−.003)
5. BJS 5	1.6648	−.0120	.0212	.0066	−.0103	.0175	−.0138	.0044	−.0134
	[41.445]	(−.007)	(.013)	(.004)	(.003)	(.011)	(−.009)	(.003)	(−.009)
6. BJS 6	1.6175	−.0030	.0211	.0070	−.001	.0112	−.0237	−.0014	−.0017
	[46.26]	(−.002)	(.013)	(.004)	(−.006)	(.007)	(−.017)	(−.001)	(−.001)
7. BJS 7	1.4832	−.0297	−.0241	.0069	−.0121	.0075	.0017	.0054	.0010
	[43.96]	(−.020)	(−.016)	(.005)	(−.001)	(.005)	(.001)	(.004)	(.001)
8. BJS 8	1.369	−.0139	−.0260	.0021	−.0106	.0005	.0002	.0031	.0022
	[42.17]	(−.010)	(−.019)	(.002)	(−.008)	(.000)	(.000)	(.002)	(.002)
9. BJS 9	1.1683	−.0224	−.0497	−.0032	−.0215	.0111	−.0011	.0015	.0050
	[35.931]	(−.019)	(−.043)	(−.003)	(−.008)	(.010)	(−.001)	(.001)	(.005)
10. BJS 10	1.1751	−.0132	−.0264	−.0214	−.0398	−.0512	−.0485	.0146	−.0033
	[21.85]	(−.011)	(−.022)	(−.018)	(−.018)	(−.044)	(−.040)	(.012)	(−.008)

		January 1953–June 1968					February 1964–June 1968		
1. Mining (10–14) (mining)	1.7469	−.0756	−.0911	.0348	.0077	.0519	.0795	.0747	.0074
	[19.18]	(−.043)	(−.052)	(.020)	(.004)	(.030)	(.050)	(.047)	(.005)
2. Food (20) (nd)	1.3987	−.0482	−.0954	−.0084	−.0403	.0258	−.0021	.0203	.0144
	[23.541]	(−.034)	(−.068)	(−.006)	(−.029)	(.018)	(−.001)	(.014)	(.010)
3. Textile (22, 23) (nd)	1.8022	−.0135	.0505	.0200	−.0214	.0635	.0578	.0437	−.0229
	[19.091]	(−.008)	(.028)	(.011)	(−.012)	(.035)	(.030)	(.023)	(−.012)
4. Chemicals (28) (nd)	1.9049	−.0192	−.0263	.0156	.0100	−.0011	.0209	.0041	.0041
	[37.23]	(−.010)	(−.014)	(.008)	(.005)	(−.001)	(.012)	(.002)	(.002)
5. Primary metals (33) (d)	2.1698	.0600	.0868	.0216	.0211	.0522	−.1012	−.0305	−.0020
	[24.69]	(.028)	(.040)	(.010)	(.010)	(.024)	(−.052)	(−.016)	(−.001)
6. Machinery (35) (d)	2.0252	−.0379	.0300	.0067	−.0037	.0162	.0378	−.0138	−.0198
	[29.80]	(−.019)	(.015)	(.003)	(−.002)	(.008)	(.019)	(−.007)	(−.010)
7. Mfg. misc. (38, 39) (d)	1.9647	−.0285	−.0003	.0037	.0146	−.0196	.0124	.0206	.0114
	[23.44]	(−.014)	(−.000)	(.002)	(.007)	(−.010)	(.006)	(.010)	(.006)
8. Transportation (41, 42, 44, 45, 47) (t & u)	1.9403	−.0032	−.0292	.0260	−.0088	.0600	.0681	.0047	.0140
	[18.36]	(−.002)	(−.015)	(.013)	(−.005)	(.031)	(.035)	(.002)	(.007)
9. Utilities (49) (t & u)	1.0437	−.0509	−.0525	−.0204	−.0642	.0193	−.0179	.0075	.0163
	[16.50]	(−.049)	(−.050)	(−.020)	(−.062)	(.018)	(−.020)	(.009)	(.018)
10. Dept. stores (53) (w & r)	1.5146	−.0127	.0081	−.0009	−.0330	.0129	.0597	.0277	.0284
	[18.63]	(−.008)	(.005)	(−.001)	(−.022)	(.008)	(.037)	(.017)	(.018)
11. Other retail (50, 54, 56–59) (w & r)	1.5296	.0115	−.0352	−.0098	−.0113	.0116	.0371	.0247	−.0058
	[22.11]	(.008)	(−.023)	(−.006)	(−.007)	(.008)	(.025)	(.016)	(−.004)
12. Banking, fin., insur., and r.e. (61, 63, 65, 67)	1.7310	−.0484	−.0605	−.0112	−.0353	−.0436	−.0039	.0105	.0177
	[26.34]	(−.028)	(−.035)	(−.006)	(−.020)	(−.025)	(−.002)	(.006)	(.010)

NOTE.—Numbers in brackets are t-statistics, while numbers in parentheses are the systematic risk measure differences as a fraction of their respective SL risk measure. Numbers within parentheses adjacent to the portfolio titles are their respective SEC industrial classification numbers, while the words within the parentheses are the general industrial classifications within which the portfolio falls. Abbreviations: const., Construction; nd, nondurable; d, durable; w & r, wholesale and retail; t & u, transportation and utilities; fin., finance; mfg. misc., manufacturing miscellaneous; dept., department; insur., insurance; r.e., real estate.

TABLE 2 **Comparisons of the Respective cov(R_j, h)'s Distinguished by Portfolio and Industry—Monthly Industry Worker Real Earnings Data**

	Production Worker Industry Affiliation							
	January 1953–June 1968					February 1964–June 1968		
Marketable Security Portfolios	Mining (1)	Const. (2)	ND (3)	D (4)	W & R (5)	T & U (6)	Fin. (7)	Service (8)
1. Mining (10–14) (mining)	−.000106	−.000014	.000023	−.000023	.000013	−.000013	.000157	.000013
2. Food (20) (nd)	−.000077	−.000027	−.000006	−.000049	.000000	−.000068	.000102	.000018
3. Textile (22, 23) (nd)	−.000068	.000087	.000013	−.000043	.000020	−.000044	.000149	−.000010
4. Chemicals (28) (nd)	−.000075	.000036	.000010	−.000024	−.000024	−.000064	.000105	.000011
5. Primary metals (33) (d)	−.000033	.000123	.000014	−.000021	.000008	−.000166	.000088	.000007
6. Machinery (35) (d)	−.000091	.000079	.000004	−.000035	−.000014	−.000064	.000106	−.000007
7. Mfg. misc. (38, 39) (d)	−.000083	.000056	.000002	−.000022	−.000037	−.000082	.000134	.000018
8. Transportation (41, 42, 44, 45, 47) (t & u)	−.000066	.000035	.000017	−.000037	.000016	−.000039	.000119	.000020
9. Utilities (49) (t & u)	−.000067	−.000007	−.000014	−.000059	.000000	−.000054	.000088	.000017
10. Dept. stores (53) (w & r)	−.000058	.000049	−.000001	−.000046	−.000010	−.000029	.000118	.000030
11. Other retailing (50, 54 56–59) (w & r)	−.000043	.000019	−.000007	−.000032	−.000011	−.000041	.000109	.000002
12. Banking, fin. insur., and r.e. (61, 63, 65, 67)	−.000058	.000049	−.000001	−.000046	−.000010	−.000029	.000118	.000030

NOTE.—For an explanation of the contents of the parentheses and list of abbreviations, see table 1.

TABLE 3 $\beta_{h,m}$ **Summarized across Industries**

Industry	Monthly
January 1953–June 1972:	
1. Mining	.0317
	(.542)
2. Construction	−.0409
	(−.48)
3. Nondurable manufacture	.0137
	(.49)
4. Durable manufacture	.0144
	(.34)
5. Wholesale and retail	−.0127
	(−.44)
February 1964–June 1972:	
6. Transportation and public utilities	.0209
	(.412)
7. Finance, insurance, and real estate	.0516
	(1.781)
8. Service	.0114
	(.443)

NOTE.—The numbers in parentheses are t-statistics. Also note that the t-statistic, two-tailed test, 10% level of significance for the monthly data is an absolute 1.645.

First, these five classifications are unique in that they are not subject to binding contractual wage agreements. The self-employed professional, etc., and managers, etc., classifications surely are not subject to such agreements, while farm-related occupations probably are not. The fortunes of these particular classifications would then be more affected by economic factors that also affect capital market movements, meaning stronger labor earnings—financial asset return covariation, contributing to the larger risk measure effects just seen. The other occupational classifications, as well as our industry production workers, are more subject to binding wage agreements, insulating them from these general capital market movements and diminishing their effect upon systematic risk.

Second, these self-employed classifications may actually be specializing in asset management, such as the equity of one's own business venture and the like. The earnings then received would be primarily generated by investments in financial rather than human capital, and would be expected to relate strongly with general capital market movements.[20] The other classifications represent more properly the return to human capital, which does relate weakly with the financial market.

Focusing more directly upon the relationship between occupational earnings and market return, we present, in table 6, the β_{hm} coefficients

20. For a further discussion of this possibility, see the article by Ehrlich and Ben-Zion (1976).

TABLE 4 Comparisons of Mayers and SL Risk Estimate Differences across Occupations for the Government Bond, NYSE, and BJS Portfolios

		Occupational Classifications						
		Professional, Technical, and Kindred			Managers, Officials, Proprietors			
Marketable Security Portfolios	SL Risk Measure	Total (1)	Self-employed (2)	Salaried (3)	Total (4)	Self-employed (5)	Salaried (6)	Clerical (7)
1958–71:								
1. Government bond index	−.0532	.0376	.7665	.0306	.0372	−.0027	.0563	.0785
	[−.43]	(−.707)	(−14.404)	(−.576)	(−.700)	(.050)	(−1.058)	(−1.476)
2. NYSE index	1.294	.0160	−.5672	.0169	.0405	.0090	.0528	−.0674
	[9.37]	(.012)	(−.4383)	(.013)	(.031)	(.007)	(.041)	(−.051)
1958–69:								
1. BJS 1	2.716	−.1437	−5.3413	−.1262	−.0242	.1291	−.1593	−.1410
	[5.96]	(−.053)	(−1.967)	(−.046)	(−.009)	(.048)	(−.059)	(−.052)
2. BJS 2	2.460	−.0052	1.3937	−.0071	.0094	.1372	−.0914	−.0416
	[6.61]	(−.002)	(.567)	(−.003)	(.004)	(.056)	(−.037)	(−.017)
3. BJS 3	2.238	.0213	.0383	.0214	.0068	.0788	−.0471	.0063
	[6.66]	(.010)	(.017)	(.010)	(.003)	(.035)	(−.021)	(.003)
4. BJS 4	2.171	−.0238	−2.1589	−.0177	−.0241	.0963	−.1004	−.0436
	[7.95]	(−.011)	(−.994)	(−.008)	(−.011)	(.044)	(−.046)	(−.020)
5. BJS 5	2.001	.0252	−.6944	.0247	.0268	.0899	−.0221	−.0260
	[10.67]	(.013)	(−.347)	(.012)	(.013)	(.045)	(−.011)	(−.013)
6. BJS 6	1.901	−.0321	−1.3304	−.0233	−.0060	.0582	−.0446	−.0381
	[12.24]	(−.017)	(−.700)	(−.012)	(−.003)	(.031)	(−.023)	(−.020)
7. BJS 7	1.646	−.0121	−2.1753	−.0059	−.0213	.0059	−.0238	−.0048
	[13.76]	(−.007)	(−1.322)	(−.004)	(−.013)	(.004)	(−.014)	(−.003)
8. BJS 8	1.561	−.1107	−4.0513	−.0984	−.1052	−.0435	−.1057	−.0450
	[8.02]	(−.071)	(−2.595)	(−.063)	(−.067)	(−.028)	(−.068)	(−.029)
9. BJS 9	1.242	−.1228	−7.6377	−.1025	−.0770	−.0538	−.0662	−.0796
	[5.89]	(−.099)	(−6.150)	(−.083)	(−.062)	(−.043)	(−.053)	(−.064)
10. BJS 10	1.545	−.0684	−4.1020	−.0615	−.0751	−.0048	−.1023	−.0641
	[7.28]	(−.044)	(−2.655)	(−.040)	(−.049)	(−.003)	(−.066)	(−.041)

TABLE 4 *(Continued)*

Marketable Security Portfolios	Occupational Classifications: Sales Workers (8)	Craftsmen-Foremen (9)	Operatives (10)	Service Workers (11)	Farm Laborers & Foremen (12)	Farmers & Farm Managers (13)	Laborers except Farmers (14)	Engineers (Salaried) (15)
1958–71:								
1. Government bond index	.0436	.0472	.0305	.0474	−.0675	.0347	.1078	.0348
	(−.820)	(−.886)	(−.574)	(−.891)	(1.268)	(−.653)	(−2.025)	(−.654)
2. NYSE index	.0007	−.0096	.0074	−.0275	.0243	.0082	−.0288	.0271
	(.001)	(−.007)	(.006)	(−.021)	(.0188)	(.006)	(−.0222)	(.021)
1958–69:								
1. BJS 1	−.0415	−.0512	−.0951	.0646	.6699	.5453	.2668	−.1040
	(−.015)	(−.019)	(−.035)	(.024)	(.247)	(.201)	(.0982)	(−.038)
2. BJS 2	−.0452	−.0043	−.0441	.1195	.3854	.0341	.2614	−.0135
	(−.018)	(−.002)	(−.018)	(.049)	(.157)	(.014)	(.106)	(−.005)
3. BJS 3	−.0477	.0096	−.0451	.0833	.2268	−.2168	.2046	−.0080
	(−.021)	(.004)	(−.020)	(.037)	(.101)	(−.097)	(.091)	(−.004)
4. BJS 4	−.0529	−.0293	−.0638	.0893	.1881	−.0958	.1152	−.0257
	(−.024)	(−.014)	(−.029)	(.040)	(.087)	(−.044)	(.053)	(−.012)
5. BJS 5	−.0138	.0017	−.0188	.0774	.1248	−.0014	.0781	.0261
	(−.007)	(.001)	(−.009)	(.039)	(.062)	(−.001)	(.039)	(−.012)
6. BJS 6	−.0099	−.0147	−.0131	.0426	.1573	−.0767	.1242	−.0232
	(−.005)	(−.008)	(−.007)	(.022)	(.083)	(−.040)	(.065)	(−.012)
7. BJS 7	−.0117	−.0231	−.0236	.0086	−.1034	−.1510	−.0913	−.0249
	(−.007)	(−.014)	(−.014)	(.005)	(−.063)	(−.092)	(−.055)	(−.015)
8. BJS 8	−.0438	−.0747	−.0684	−.0575	−.1722	.0794	−.1998	−.0835
	(−.028)	(−.048)	(−.044)	(−.037)	(−.110)	(.051)	(−.128)	(−.053)
9. BJS 9	−.0209	−.0800	−.0655	−.0737	−.1327	.1311	−.2111	−.0855
	(−.017)	(−.064)	(−.053)	(−.059)	(−.107)	(.106)	(−.170)	(−.069)
10. BJS 10	−.0621	−.0854	−.0901	.0056	−.1386	−.0417	−.1904	−.0717
	(−.040)	(−.055)	(−.058)	(.004)	(−.090)	(−.027)	(−.123)	(−.046)

NOTE.—The numbers in brackets are t-statistics, while the numbers in parentheses are the systematic risk measure differences as a fraction of their respective SL risk measures.

TABLE 5 Comparisons of Mayers and SL Risk Estimate Differences across Occupations for the MacBeth Industry Portfolios

		Occupational Classifications						
		Professional, Technical, and Kindred			Managers, Officials, Proprietors			
Marketable Security Portfolios	SL Risk Measure	Total (1)	Self-employed (2)	Salaried (3)	Total (4)	Self-employed (5)	Salaried (6)	Clerical (7)
1958–67:								
1. Mining (10–14)	2.951	.2808	−3.0729	.1916	.0121	−.0426	.0498	.2243
	[4.75]	(.095)	(−1.041)	(.065)	(.004)	(−.014)	(.017)	(.076)
2. Food (20)	1.952	.0397	.0471	.0384	.0219	−.0444	.0619	.0514
	[6.35]	(.020)	(.024)	(.020)	(.011)	(−.023)	(.032)	(.026)
3. Textile (22, 23)	2.955	−.0752	.2010	−.0498	−.0102	.2904	−.1434	.0009
	[6.89]	(−.025)	(.068)	(−.017)	(−.003)	(.098)	(−.049)	(.000)
4. Chemicals (28)	2.037	.0653	.4821	.0830	.1115	.0926	.1318	.0497
	[7.99]	(.032)	(.237)	(.040)	(.055)	(.045)	(.065)	(.024)
5. Primary metals (33)	2.271	.0794	−.0771	.0861	.2060	.2541	.1364	−.0359
	[4.79]	(.035)	(−.034)	(.038)	(.091)	(.112)	(.060)	(−.016)
6. Machinery (35)	2.985	.1490	−1.2373	.1288	.1141	.2209	.0588	.1932
	[6.03]	(.050)	(−.415)	(.043)	(.038)	(.074)	(.020)	(.065)
7. Mfg. misc. (38, 39)	2.728	.2158	.3490	.2380	.2587	.1746	.3232	.2816
	[3.44]	(.079)	(.128)	(.087)	(.095)	(.064)	(.118)	(.103)
8. Transportation (41, 42, 44, 45, 47)	1.672	−.0312	−1.2291	−.0546	.0239	.1279	−.1147	−.0847
	[2.83]	(−.019)	(−.735)	(−.033)	(.014)	(.076)	(−.069)	(−.051)
9. Utilities (49)	5.322	−.1905	1.3478	−.1624	−.1587	−.1632	−.1332	−.1372
	[1.28]	(−.358)	(2.532)	(−.305)	(−.298)	(−.307)	(−.250)	(−.258)
10. Department stores (53)	3.118	.0447	−.1876	.0383	−.0625	.0756	−.0470	.0986
	[7.25]	(.014)	(−.060)	(.012)	(−.020)	(.024)	(−.015)	(.032)
11. Other retailing (50, 54, 56–59)	2.852	.0861	−.1767	.0750	.0421	.1021	.0621	.0768
	[6.97]	(.030)	(−.062)	(.026)	(.015)	(.036)	(.022)	(.027)
12. Banking, finance, insurance, real estate (61, 63, 65, 67)	2.099	.0725	−.3039	.0656	.0342	.0482	.0477	.0614
	[4.94]	(.035)	(−.145)	(.031)	(.016)	(.023)	(.021)	(.029)

TABLE 5 *(Continued)*

Marketable Security Portfolios	Occupational Classifications: Sales Workers (8)	Craftsmen-Foremen (9)	Operatives (10)	Service Workers (11)	Farm Laborers & Foremen (12)	Farmers & Farm Managers (13)	Laborers except Farmers (14)	Engineers (Salaried) (15)
1958–67:								
1. Mining (10–14)	−.1219	.1403	.0489	.1105	−.4438	−.6693	.1057	.1449
	(−.041)	(.048)	(.017)	(.037)	(−.150)	(−.227)	(.036)	(.049)
2. Food (20)	.0064	.0227	−.0087	−.0086	−.1663	−.3294	−.1803	−.0410
	(.003)	(.012)	(−.004)	(−.004)	(−.085)	(.169)	(−.092)	(−.021)
3. Textile (22, 23)	.0386	.0113	−.0031	.1209	.3345	.3707	.1208	−.0400
	(.013)	(.004)	(−.001)	(.041)	(.113)	(.125)	(.041)	(−.014)
4. Chemicals (28)	.0779	.1139	.0672	.0416	.1926	−.2775	.2721	.0588
	(.038)	(.056)	(.033)	(.020)	(.095)	(−.136)	(.134)	(.029)
5. Primary metals (33)	.1395	.1641	.1416	.1366	.6785	.4419	.5942	.1279
	(.061)	(.072)	(.062)	(.060)	(.299)	(.195)	(.262)	(.056)
6. Machinery (35)	.0302	.2444	.0738	.1473	.4159	−.0921	.5194	.0713
	(.010)	(.082)	(.025)	(.049)	(.139)	(−.031)	(.174)	(.024)
7. Mfg. misc. (38, 39)	.1407	.3948	.1363	.0885	.5572	−.8433	.7944	.1012
	(.052)	(.145)	(.050)	(.032)	(.204)	(−.309)	(.291)	(.037)
8. Transportation (41, 42, 44, 45, 47)	.0153	.0258	.0613	.0956	.6826	.8703	.5077	.0255
	(.009)	(.015)	(.037)	(.057)	(.408)	(.521)	(.304)	(.015)
9. Utilities (49)	−.0496	−.2370	−.1304	−.1419	−.4226	.1771	−.6677	−.1479
	(−.093)	(−.445)	(−.245)	(−.267)	(−.794)	(.333)	(−1.254)	(−.278)
10. Department stores (53)	−.0656	−.0032	−.0559	.0261	−.4970	−.2594	−.3076	.0445
	(−.021)	(−.001)	(−.018)	(.008)	(−.159)	(−.083)	(−.099)	(.014)
11. Other retailing (50, 54, 56–59)	−.0069	.0677	−.0200	.0389	−.2930	.0408	−.2641	.0621
	(−.002)	(.024)	(−.007)	(.014)	(−.103)	(.014)	(−.093)	(.022)
12. Banking, finance, insurance, real estate (61, 63, 65, 67)	.0117	.0200	.0055	.0621	−.2636	−.4050	−.2668	−.0285
	(.006)	(.010)	(.003)	(.030)	(−.126)	(−.193)	(−.127)	(−.014)

NOTE.—The numbers in the parentheses which are adjacent to the portfolio titles are their respective SEC industrial classification numbers.

TABLE 6 β_{hm} Summarized across Occupations

Professional, technical, kindred:	
Total	−.1151
	(−.83)
Self-employed	−1.330
	(−2.45)
Salaried	−.0532
	(−.40)
Managers, officials, proprietors:	
Total	−.0182
	(−.13)
Self-employed	.2476
	(1.36)
Salaried	−.1105
	(−.67)
Clerical	−.0473
	(−.41)
Sales workers	.1300
	(1.18)
Craftsmen-foremen	−.0890
	(−.87)
Operatives	−.0452
	(−.47)
Service workers	.1078
	(.63)
Farm laborers & foremen	.5009
	(.93)
Farmers & farm managers	−.4098
	(−1.20)
Laborers except farmers	−.0148
	(−.06)
Engineers (salaried)	−.0667
	(−.49)

NOTE.—The numbers within parentheses are the respective t-statistics. Also note that the t-statistic, two-tailed test, 10% level of significance is an absolute 1.771.

for all the occupation classifications. As expected, all of the previously mentioned occupations, except laborers, have large coefficients which are significantly different from zero, the self-employed professional, etc., being the largest at −1.330. All the other classifications, with the exception of sales workers, have small coefficients that are not significant. These results again indicate that, except for the self-employed and farm-related categories, occupation too should have little to do with portfolio composition.

As for pricing, only two of the portfolios, manufacturing miscellaneous and utilities, exhibit any predominant one-sided risk effects across all occupational classifications. It would thus also appear from the occupational data that human capital in the aggregate should have little effect upon capital asset pricing as well.

Finally, turning to educational attainment, we find again little in terms of magnitude or pattern to the human capital effect. Looking at table 7, which presents the Mayers-SL risk measure differences using

TABLE 7 The Mayers and SL Risk Estimate Differences for the Individual Data Aggregated by Educational Attainment

		Years of Schooling						
1968–74 Portfolios	SL Estimates	0–7 (N = 42)	8 (N = 69)	9–11 (N = 131)	12 (N = 307)	13–15 (N = 134)	16 (N = 108)	17+ (N = 93)
CRSP 1	1.956	−.0694	−.0118	−.0138	.0164	−.0360	.0213	.0755
	[3.20]	(−.035)	(−.006)	(−.007)	(.008)	(−.018)	(.011)	(.039)
CRSP 2	1.687	−.0931	−.0045	.0096	.0175	−.0320	.0132	.0330
	[3.94]	(−.055)	(−.003)	(.006)	(.010)	(−.019)	(.008)	(.020)
CRSP 3	1.507	−.0401	−.0032	−.0787	−.0106	−.0859	.0124	.0760
	[3.43]	(−.027)	(−.002)	(−.052)	(−.007)	(−.057)	(.008)	(.050)
CRSP 4	1.582	−.0181	−.0553	−.0624	−.0259	−.0659	.0465	.0250
	[4.45]	(−.011)	(−.035)	(−.039)	(.016)	(−.042)	(.029)	(.016)
CRSP 5	1.470	−.0119	−.0126	−.0305	.0014	−.0218	.0107	.0449
	[5.22]	(−.008)	(−.009)	(−.021)	(.001)	(−.015)	(.007)	(.031)
CRSP 6	1.320	.0104	−.0003	−.0250	.0000	−.0293	.0575	.0365
	[5.88]	(.008)	(.0)	(−.019)	(.0)	(−.022)	(.044)	(.028)
CRSP 7	1.135	.0081	−.0232	−.0525	−.0096	−.0234	.0716	.0386
	[5.90]	(.007)	(−.020)	(−.046)	(−.008)	(−.021)	(.063)	(.034)
CRSP 8	1.063	.0435	.0100	−.0413	.0004	−.0202	.0607	.0530
	[6.68]	(.041)	(.009)	(−.039)	(.0)	(−.019)	(.057)	(.050)
CRSP 9	.9797	.0249	.0088	−.0396	.0024	−.0173	.0635	.0496
	[7.73]	(.025)	(.009)	(−.040)	(.002)	(−.018)	(.065)	(.051)
CRSP 10	.929	.0195	.0152	−.0394	.0034	−.0289	.0446	.0475
	[7.62]	(.021)	(.016)	(−.042)	(.004)	(−.031)	(.048)	(.051)

NOTE.—The numbers in parentheses are the systematic risk measure differences expressed as a fraction of their respective SL risk measures, while the numbers in brackets are the respective t-statistics.

TABLE 8 Distributional Characteristics of the Mayers and SL Risk Estimate Differences for Individuals by Educational Grouping

		Years of Schooling											
		0–7 (N = 42)			8 (N = 69)			9–11 (N = 131)			12 (N = 307)		
1968–74 Portfolios	SL Estimates	Q_1	Median	Q_3	Q_1	Median	Q_3	Q_1	Median	Q_3	Q_1	Median	Q_3
CRSP 1	1.956 [3.20]	−.550	−.0118	.641	−.650	.0262	.5572	−.650	−.0649	.6420	−.4982	−.0119	.6557
CRSP 2	1.687 [3.94]	−.503	.1955	−.2490	−.406	.1322	.5082	−.346	−.0375	.4152	−.3101	.0240	.4540
CRSP 3	1.507 [3.43]	−.621	−.1255	.3522	−.368	.0665	.5062	−.482	−.0501	.4347	−.3757	.0220	.5563
CRSP 4	1.582 [4.45]	−.551	−.1466	.2737	−.260	.0146	.3622	−.356	−.0272	.2776	−.3090	−.0200	.3685
CRSP 5	1.470 [5.22]	−.539	−.0320	.283	−.219	.0357	.3449	−.337	−.0146	.3163	−.2777	.0132	.3171
CRSP 6	1.320 [5.88]	−.315	−.0587	.2635	−.191	.0003	.2007	−.280	−.0200	.1996	−.1959	.0327	.2607
CRSP 7	1.135 [5.90]	−.340	−.0583	.2126	−.198	−.0048	.1933	−.245	−.0597	.1198	−.1876	−.0050	.2008
CRSP 8	1.063 [6.68]	−.178	.0327	.2228	−.137	.0165	.1858	−.2013	−.0363	.1374	−.1597	.0043	.1849
CRSP 9	.9797 [7.73]	−.129	.0313	.177	−.140	−.0006	.1500	−.1816	−.0361	.0784	−.1442	.0057	.1668
CRSP 10	.929 [7.62]	−.105	.0543	.168	−.104	.0259	.2022	−.1577	−.0181	.0839	−.1075	.0163	.1563

TABLE 8 *(Continued)*

		Years of Schooling								
		13–15 (N = 134)			16 (N = 108)			17+ (N = 93)		
1968–74 Portfolios	SL Estimates	Q_1	Median	Q_3	Q_1	Median	Q_3	Q_1	Median	Q_3
CRSP 1	1.956 [3.20]	−.4309	−.0336	.7676	−.3606	.0898	.7030	−.3864	.0443	.5199
CRSP 2	1.687 [3.94]	−.3360	.0084	.4407	−.2988	.0770	.4537	−.2648	−.0057	.3332
CRSP 3	1.507 [3.43]	−.3228	−.0133	.4605	−.2819	.1354	.5602	−.2265	.0611	.4421
CRSP 4	1.582 [4.45]	−.2662	−.0016	.3859	−.2141	.0903	.3754	−.1922	.0366	.2748
CRSP 5	1.470 [5.22]	−.1727	−.0273	.2845	−.1627	.0700	.3578	−.1833	.0267	.3111
CRSP 6	1.320 [5.88]	−.1672	−.0184	.2319	−.1507	.0526	.3129	−.1208	.0376	.2356
CRSP 7	1.135 [5.90]	−.1305	.0031	.1886	−.1082	.0598	.2693	−.0969	.0590	.2083
CRSP 8	1.063 [6.68]	−.1442	−.0092	.1718	−.0805	.0503	.2745	−.0658	.0390	.2184
CRSP 9	.9797 [7.73]	−.1358	−.0102	.1508	−.0715	.0385	.2021	−.0543	.0335	.1729
CRSP 10	.929 [7.62]	−.1396	−.0104	.1306	−.0650	.0226	.1448	−.0758	.0351	.1393

NOTE.—The numbers in brackets are the SL estimate t-statistics.

the individual longitudinal data after first grouping them by schooling level, we find the largest difference to be a −0.0931, while almost 60 of the 70 differences shown do not go beyond absolute 0.055. The effects are mixed with the first four schooling intervals exhibiting both positive and negative signs, the 13–15 interval only negative, and the 16 and 17+ intervals only positive. In addition, we find no pervasively one-sided effects when viewing each CRSP portfolio across schooling intervals.

To test whether these effects are not simply the result of aggregation, we present in table 8 the distributional characteristics of the Mayers-SL differences for each CRSP portfolio when the individual ungrouped data are used. The median and the first and third quartiles are shown for each portfolio by schooling level.[21] Again the results are much the same, with few medians in excess of absolute 0.10 and no discernible pattern.

Clearly, then, educational attainment, as with industry affiliation and most occupational classifications, should appear to have little effect upon portfolio composition and, in the aggregate, upon pricing as well.

V. Summary and Concluding Remarks

This study has attempted to examine empirically the effect of nonmarketable human capital upon both capital asset pricing and individual portfolio composition. Following the theoretical framework first proposed by David Mayers (1972, 1973), we developed an empirical methodology by which to study this effect and applied it to a number of different sets of labor earnings data distinguished by industry, occupation, and schooling.

We find with regard to portfolio composition that only occupational classification seemed to matter, and then only if characterized as self-employed or farm related. Nonfarm salaried workers appear to be able to ignore their own human capital when constructing their portfolios, regardless of industry affiliation, particular occupation, or level of schooling.

And with regard to capital asset pricing, our results appear to strongly confirm those of Fama and Schwert (1977) that human capital in the aggregate has little to do with capital market pricing as well. Empirical work on the pricing of financial assets may then justifiably ignore the existence of human capital, as has been the practice in the past.

21. The means are not reported here, since there were certain outliers which affected their computation.

The reason for these results, we find, is the fact that human capital, both in the whole or in the part, is "weakly" related with the financial market—so weakly, in fact, that no meaningful covariation appears to exist overall between changes in labor earnings and the rate of return on financial assets. Why this should be so is a matter for conjecture, but at least two suggestions may be offered.

First, employees are typically the first in line of all creditors to be compensated by the firm. They would thus be the least affected by the periodic fluctuations in the fortunes of the firm, while creditors farther down the line would bear the major burden of the risk. Labor earnings would therefore exhibit little covariance with the market.[22]

And second, as discussed earlier, most salaried workers have multiperiod contracts which would smooth their earnings stream and protect them from the periodic fluctuations in the fortunes of the market and the firm. This would then also contribute to the little covariance that we see.[23]

Admittedly, we have only scratched the surface of the research required on human capital and the financial market. Much still can be done.

First, in line with the Mayers model, we have assumed throughout only the two-period world of Sharpe and Lintner. A more appropriate framework might be that of a multiperiod model, since human capital is essentially a multiperiod commitment. Thus, as a possible next step, the implications of introducing human capital into a multiperiod pricing model might be studied and testable implications might be developed.

Second, this study has viewed human capital as a purely exogenously determined asset which is given and must be dealt with by the investor. A preferable alternative might be to develop an investment model explicitly allowing for endogenously determined human capital and then drawing implications for pricing and portfolio composition.[24]

22. This approach was suggested to me by Eugene Fama.

23. Recent work has been done by a number of economists on implicit labor contracts which tries to show why it pays for firms to offer workers a long-term contract that involves relatively little fluctuation in their wages. Among the authors of such papers are Baily (1974) and Azariadis (1975). Our results are consistent with such contracts, since we do find not only little covariance but little fluctuation. Of course, from a longer-run perspective, the riskiness on human capital would be positively related to that of other capital as firms go under or prosper. But in the context of shorter-run fluctuations, including year-to-year fluctuations, our results are consistent with the implications of the Azariadis and Baily models. (These remarks were suggested to me by Gary Becker.) In a related work, Liberman (1979) found that not only are the absolute earnings streams quite smooth over time but so is the wage structure as well. It is demonstrated there that the rate of return on educational investment, a major component of human capital, exhibits remarkable stability and durability in the face of major movements in the financial market rate, which would be consistent with the results of this paper.

24. Work in this direction has been done by Levhari and Weiss (1974) and Williams (1978).

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