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Gabriel Hawawini
Donald B. Keim

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The Wharton School
University of Pennsylvania

The Rodney L. White Center for Financial Research

The Wharton School
University of Pennsylvania
3254 Steinberg Hall-Dietrich Hall
3620 Locust Walk
Philadelphia, PA 19104-6367

(215) 898-7616
(215) 573-8084 Fax
http://finance.wharton.upenn.edu/~rlwctr

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The Cross Section of Common Stock Returns: A Review of the Evidence and Some New Findings

Gabriel Hawawini*
and
Donald B. Keim†

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*INSEAD, 77305 Fontainebleau Cedex, France.
†Wharton School, University of Pennsylvania, Philadelphia PA 19104, USA.

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Abstract

A growing number of empirical studies suggest that betas of common stocks do not adequately explain cross-sectional differences in stock returns. Instead, a number of other variables (e.g., size, ratio of book to market, earnings/price) that have no basis in extant theoretical models seem to have significant predictive ability. Some interpret the findings as evidence of market inefficiency. Others argue that the Capital Asset Pricing Model is an incomplete description of equilibrium price formation and these variables are proxies for additional risk factors. In this paper we review the evidence on the cross-sectional behavior of common stock returns on the U.S. and other equity markets around the world. We also report some new evidence on these cross-sectional relations using data from both U.S. and international stock markets. We find, among other results, that although the return premia associated with these *ad hoc* variables are significant in most international stock markets, the premia are uncorrelated across markets. The accumulating evidence prompts the following question: If these return premia occur primarily in January and are uncorrelated across major international equity markets, is it reasonable to characterize them as compensation for risk?

1. Introduction

In this paper we review the evidence on the cross-sectional behavior of common stock returns in the U.S. and other equity markets around the world.[1] Since the early 1980's, a growing number of empirical studies have documented the presence of persistent cross-sectional patterns in stock returns that do not support one of the fundamental tenets of modern finance: expected stock returns are determined by their level of beta risk through a positive and linear relation known as the capital asset pricing model, or CAPM (Sharpe (1964), Lintner (1965), Mossin (1966), Treynor (1961)).

The evidence suggests that betas of common stocks do not adequately explain cross-sectional differences in stock returns. Instead, a number of other variables that have no basis in extant theoretical models seem to have significant predictive ability. These other variables include firm size (measured by market capitalization of the firm's common stock), the ratio of book to market values (the accounting value of a firm's equity divided by its market capitalization), earnings yield (the firm's reported accounting net profits divided by price per share), and the firm's *prior* return performance.

Interpretation of the cross-sectional explanatory power of such *ad hoc* variables presents a challenge to the profession. Recall that tests of asset pricing models involve the joint null hypothesis that security markets are informationally efficient and expected returns are described by a prespecified equilibrium model (e.g., the CAPM). If the joint hypothesis is rejected, we cannot specifically attribute that rejection to one or the other branch of the hypothesis. Indeed, a lively debate continues in the literature regarding the interpretation of these results. Some interpret the findings as convincing evidence of market inefficiency: if stock returns can be predicted on the basis of historical factors such as market capitalization, book-to-market value and prior return performance, then it is difficult to characterize stock markets as informationally efficient. On the other hand, the rejection may be due to a test

[1] This paper draws on our earlier paper (Hawawini and Keim (1995)), updating its content and extending the section that reviews the recent evidence regarding the cross-sectional behavior of common stock returns. This paper, as our earlier one, is not meant to be an exhaustive compilation of the findings on the cross-sectional predictability of stock returns. The focus is on the subset of findings whose existence has proved most robust with respect to time and the stock markets in which they have been observed.

design that is based on an incorrect equilibrium model. The fact that so many of these regularities have persisted for more than thirty years suggests that perhaps our benchmark models are incomplete descriptions of equilibrium price formation.[2]

There are additional findings that challenge the notion that these *ad hoc* variables proxy for additional risk factors. First, the evidence indicates that the relation between returns and variables like firm size and book-to-market ratio is typically significant *only* during the month of January. Why would a risk-based factor manifest itself during the month of January and not the rest of the year? If investors expect higher returns for holding stocks with a particular characteristic, then it is reasonable to expect the market to deliver that premium uniformly throughout the year. If that premium is compensation for risk, is there reason to believe that the market is systematically more risky in January than the rest of the year?

Second, some new international evidence on the premia associated with size, E/P, CF/P and P/B that we report in section 6 is difficult to reconcile with an international version of the risk story. If these premia are compensation for additional risks that are priced in the context of an international asset pricing model under conditions of integrated international capital markets, then the premia should be correlated across markets in much the same way that the market risk premium is significantly correlated across markets. Inconsistent with this hypothesis, we find that the premia correlations are insignificant across the markets in our sample.

The rest of the paper is organized as follows. In the next section we briefly review the early tests of the CAPM which were generally supportive of that model. In section 3 we survey the recent empirical evidence which is at odds with the predictions of the CAPM. In section 4 we examine evidence that sheds light on whether the explanatory power of the *ad hoc* variables reflects the underlying influence of one or several underlying phenomenon.

[2]Another explanation that has not been adequately addressed in the literature is that these *ad hoc* variables proxy for betas that are not properly estimated. Put differently, the failure of the empirical evidence to provide unambiguous support for the CAPM is not necessarily proof of the model's invalidity, but may be the manifestation of our inability to accurately measure beta risk. For example, one can argue that stocks with lower ratios of P/B have higher average returns than stocks with higher ratios of P/B because they are indeed riskier in a beta sense. If we could measure beta risk with less error, then the reported negative relation between P/B and *beta-adjusted* returns may disappear.

In section 5 we discuss the research that attempts to sort through the effects to determine which ones have the greatest explanatory power. Section 6 reviews potential interpretations of the evidence on the cross-sectional behavior of returns, and reports some new international premia evidence that appears to be inconsistent with a risk-based story. Section 7 concludes the paper.

2. Tests of the (Single-Beta) Capital Asset Pricing Model

The capital asset pricing model has occupied a prominent position in financial economics for the thirty years since its origins in the papers by Treynor (1961), Sharpe (1964), Lintner (1965) and Mossin (1966). Given certain simplifying assumptions, the model states that the expected rate of return on any security is positively and linearly related to that security's systematic risk (or beta) measured relative to the market portfolio of all marketable securities. Hence, according to the CAPM, the relation between the expected return $E(R_i)$ and the systematic risk β_i of security i can be expressed as:

$$E(R_i) = a_o + a_1\beta_i \tag{1}$$

If the model is correct and security markets are efficient, stock returns should on average conform to this linear relation. Persistent departures from positive linearity would represent violations of the joint hypothesis that both the CAPM and the efficient market hypothesis are valid.

The early tests (e.g., Black, Jensen, and Scholes (1972), Blume and Friend (1973), and Fama and MacBeth (1973)) found evidence of a significant relation between average returns and estimated betas, but the estimated intercept was higher (a_o in equation (1)), and the estimated slope (a_1 in equation (1), representing an estimate of the market risk premium) was lower than predicted by the Sharpe-Lintner CAPM and only marginally important in explaining cross-sectional differences in average stock returns. The results of these studies were interpreted as being consistent with the Black (1972) version of the CAPM according to which the relation between expected returns and systematic risk should be flatter than that predicted by the standard CAPM if a risk-free security is not available to investors.

Although the early tests lend some support for the model, subsequent research was not always as accommodating. For example, in his critique of existing tests of the CAPM, Roll (1977) argued that tests performed with any "market" portfolio other than the true market portfolio are not tests of the CAPM and, therefore, cannot be interpreted as evidence for or against the model. In response to Roll's criticism of the earlier tests, Stambaugh (1982) constructed broader market indexes that included bonds, real estate and consumer durables and found that tests of the model with these broader indexes were not very sensitive to the breadth of the definition of the market proxy.

3. Evidence Inconsistent with the CAPM

The central prominence of beta in the asset pricing paradigm came into question with the first tests of ad hoc alternatives to the CAPM in the late 1970's. The earliest of these tests are those of Basu (1977) and Banz (1981). They found that the price-to-earnings ratio (P/E) and the market capitalization of common equity (firm size), respectively, provided considerably more explanatory power than beta. Other studies have extended the list of predictive factors to include, among others, the ratio of book-to-market value, price per share, and prior return performance. Combined, these studies have produced convincing evidence of cross-sectional return predictability that greatly transcends the marginal explanatory power of beta found in the earlier studies. Notably absent in this literature, though, is any supporting theory to justify the choice of factors. Nevertheless, the findings collectively represent a set of stylized facts that stand as a challenge for alternative asset pricing models.

In this section we present a sample of the more important contributions to these stylized facts. The basic set of findings have been reported in a variety of manner. For example, some researchers employ cross-sectional regression techniques similar to those originally used by Fama and MacBeth (1973)

$$R_i = a_0 + a_1\beta_i + a_2\Sigma c_{ij} + e_i \tag{2}$$

where c_{ij} represents characteristic j (size, earnings yield, price-book ratio, etc.) for stock i. Researchers have also documented these phenomenon by examining the returns of portfolios constructed on the basis of these characteristics c_{ij}.

3.1. Data and empirical methods

To maintain a unifying thread throughout the following discussion of cross-sectional return predictability, we augment much of our reporting of the original results in the literature with some basic summary statistics that document the findings with a common data set for the same time period using the same empirical methods. We believe this approach avoids some of the apples-and-oranges comparisons that bog down literature surveys of disparate studies employing widely varying samples, time periods and empirical methods. We portray our new evidence in portfolio form because we feel the returns to feasible portfolio strategies provide a useful perspective on the economic significance of the results. We report our findings using monthly value-weighted portfolio returns to avoid the potential statistical biases associated with measuring these effects with daily portfolio returns (e.g., Roll (1981), Reinganum (1982), Blume and Stambaugh (1983)). The use of monthly data also avoids biases in estmated betas due to the infrequent or nonsynchronous trading of securities (e.g., Dimson (1979), Scholes and Williams (1977)). The data are drawn from the monthly return file of the Center for Research in Security Prices (CRSP) and the Compustat annual industrial and research files. In each of the "simulation" experiments reported below, portfolios are created on March 31 of each year using prices and shares outstanding on March 31 and accounting data for the year ending on the previous December 31 (with portfolios containing only December 31 fiscal closers). Aside from new listings and delistings, which are added to or dropped from the portfolios as they occur during the year, the portfolio composition remains constant during the following twelve months over which the portfolio returns are calculated. As such, the simulated portfolios display little trading, and represent feasible strategies.

3.2. The size effect

Much of the research on cross-sectional predictability of stock returns has focused on the relation between returns and the market value of common equity, commonly referred to as the size effect. Banz (1981) was the first to document this phenomenon. For the period 1931 to 1975, he estimated a model of the form:

$$R_i = a_0 + a_1 b_i + a_2 S_i + e_i, \tag{3}$$

where S_i is a measure of the relative market capitalization ("size") for firm i. He found that the statistical association between returns and size is negative and of a greater order of magnitude than that between returns and beta documented in the earlier studies of the CAPM.

The first set of columns in Table 1 reports the average monthly returns for ten value-weighted size-portfolios of NYSE and AMEX stocks for the period April 1962 to December 1994, along with corresponding values for portfolio beta and average market capitalization of the stocks in the portfolio. The negative relation between size and average returns is clearly evident. The annualized difference in returns between the smallest and largest size deciles is 8.8 percent. Note that the portfolio betas decline with increasing size, but the differences are small. Thus, consistent with research that finds significant negative coefficients on size in equation (3) after adjusting for the explanatory power of beta, the difference in estimated OLS betas between the smallest and the largest size portfolios is insufficient to explain the difference in returns between the two extreme portfolios in Table 1.

Additional evidence in Reinganum (1990) suggests that the relative price behavior of small and large firms may differ for over-the-counter (OTC) stocks. Using data for the 1973-1988 period, Reinganum finds that small OTC shares have significantly lower returns than NYSE and AMEX firms with the same size, and that the small-firm premium for OTC stocks is much lower than for NYSE and AMEX stocks. Reinganum, motivated by earlier work by Amihud and Mendelson (1986), argues that the differences are related to differences in liquidity between the two markets, suggesting differential costs of trading small stocks in these two types of markets.[3] [4] The implication is that market structure may be an important influence on the measured size effect. If so, the analysis of the international evidence on the size effect, where we observe very different market organizations and structures, should reveal significant differences in the magnitude of the size premium across markets.

Following the discovery of a size premium in the U.S. equity markets, numerous studies have documented its existence in most stock markets around the world. Models similar to (3) have been estimated for Belgium (Hawawini, Michel, and Corhay (1989)), Canada (Calvet and Lefoll (1989)), France (Hawawini and Viallet (1987)), Ireland (Coghlan (1988)) Japan (Hawawini (1991), Chan, Hamao, and Lakonishok (1991)), Mexico (Herrera and Lockwood (1994)), Spain (Rubio (1988)), Switzerland (Corniolay and Pasquier (1991)) and the United Kingdom (Corhay, Hawawini, and Michel (1988)). In all these countries, except Mexico, there is no relation, on average, between return and beta risk when all months of the year are considered (i.e., a_1 is statistically indistinguishable from zero). There is, however, a significant negative relationship between returns and portfolio size in all countries except Canada and France (i.e., a_2 is significantly less than zero).

The portfolio evidence from international equity markets is summarized in Table 2 for the stock markets of Australia, New Zealand, Canada, Mexico, Japan, Korea, Singapore, Taiwan, and eight European countries. The monthly size premium is defined as the difference between the average monthly return on the portfolio of smallest stocks and the average monthly return on the portfolio of largest stocks. In all countries, except Korea, the size premium is positive during the reported sample periods (which, in most cases, are significantly shorter than the 32 years of data we use in Table 1 to estimate the size effect in the U.S. market). As

[3] Loughran (1993) finds, however, that of the 5.7 percent difference in returns between NYSE and NASDAQ stocks in the bottom five size deciles (based on NYSE ranking), 60 percent is due to the poor (long-run) performance of initial public offerings (IPO's) on NASDAQ. A difference of only 2.3 percent remains after purging NASDAQ returns of an IPO effect (IPO's are much more heavily concentrated on NASDAQ than on the NYSE).

[4] Fama, French, Booth, and Sinquefield (1993) show that small NYSE stocks have substantially lower ratios of price to book value than comparably- small NASDAQ stocks. They argue that the higher returns for small NYSE stocks are related to the lower price/book ratios, a relation that appears to persist independently of the size-return relation.

expected, the size premium varies significantly across markets[5]: It is most pronounced in Australia (5.73 percent) and Mexico (4.16 percent), and least significant in Canada (0.44 percent) and the United Kingdom (0.40 percent).[6] As is the case for U.S. data, differences in beta across size portfolios cannot explain differences in returns.

There are, however, significant differences across the fifteen markets in the spread between the size of the largest and smallest portfolios as indicated by the ratios of the average market capitalization of the largest portfolio to that of the smallest one, reported in Table 2. For example, in Spain, the largest size- portfolio is 228 times larger than the smallest one, whereas in the case of Taiwan it is only 17 times larger. There does not seem to be a relation between the magnitude of the size premium and the size ratio. However, because the size and number of portfolios as well as the sample periods differ across markets, it is difficult to gauge whether the magnitude of the size premium is indeed significantly different across these countries, although, as pointed out earlier, we suspect that the differences in the size premium are unlikely to be explained by these factors alone. Differences in market structures and organizations may account for some of the reported variation in the size premium across markets.

[5]Although we hypothesized that the magnitudes of the size effect across markets should reflect differences in the organizations and structures of these markets, the results reported in Table 2 are also likely to be sensitive to differences in sample dates and lengths.

[6]Although Levis (1985) finds that the size effect on the London Stock Exchange (LSE) is not statistically significant, others report a significant size premium. Banz (1985) provides evidence of a significant size effect on the LSE. His analysis is based on 29 years of monthly returns (1955-1983) taken from the London Share Price Data base (LSPD). With ten value-based portfolios, he reports a compounded annual return of 39.9 percent for the smallest portfolio versus 13.0 percent for the largest. Dimson and Marsh (1984) also report evidence of a size effect on the portfolios constructed from a sample of stocks taken from the LSPD. Over the period 1977-1983, the portfolio of smallest stocks earned a compound annual return of 41 percent and the portfolio of largest stocks realized a compound annual return of 18 percent. In Banz (1985), the compound annual return on the smallest portfolio exceeded that of the largest by 27 percent. Dimson and Marsh (1984) report that the difference is 23 percent, both before adjustment for risk. More recently, Strong and Xu (1995) report an average monthly size-premium of 0.61 percent (7.3 percent annually) for the extreme portfolios drawn from decile portfolios formed on size during the period July 1973 to July 1992 using the London Share Price Data base, a result that is not explained by differences in betas.

3.3. *The earnings-yield effect*

Earnings-related strategies have a long tradition in the investment community. The most popular of these strategies, which calls for buying stocks that sell at low multiples of earnings, can be traced back at least to Graham and Dodd (1940) who proposed that "a necessary but not a sufficient condition [for investing in a common stock is] a reasonable ratio of market price to average earnings" (p. 533). They advocated that a prudent investor should never pay as much as 20 times earnings and a suitable multiplier should be 12 or less.

Ball (1978) argues that earnings-related variables like the earnings-to-price ratio (E/P) are proxies for expected returns. In that case, if the CAPM is an incomplete specification of priced risk, then we would expect E/P to explain the portion of expected return that is in fact compensation for risk variables omitted from the tests. A valid question, then, is whether a documented relation between average returns and E/P is due to the influence of E/P, or whether E/P is merely proxying for other explanators of expected returns.

Nicholson (1960) published the first extensive study of the relation between P/E multiples (the reciprocal of the earnings yield) and subsequent total returns, showing that low P/E stocks consistently provided returns greater than the average stock. Basu (1977) introduced the notion that P/E ratios may explain violations of the CAPM and found that, for his sample of NYSE firms, there was a significant negative relation between P/E ratios and average returns in excess of those predicted by the CAPM. If one had followed his strategy of buying the quintile of lowest P/E stocks and selling short the quintile of highest P/E quintile stocks, based on annual rankings, the average annual abnormal return would have been 6.75 percent (before commissions and other transaction costs) over the 1957 to 1975 period. Reinganum (1981), analyzing both NYSE and AMEX stocks, confirmed and extended Basu's findings to 1979.

In the second set of columns in Table 1 we report the relation between average monthly returns and E/P for the 1962-94 period using the same data file of NYSE and AMEX stocks used to examine the size effect. The portfolio returns in Table 1 confirm the E/P effect documented in previous studies.[7] The difference in returns between the highest and lowest E/P portfolios is, on average, 0.39 percent per month ($T = 1.68$).[8]

There is less evidence of an E/P effect in markets outside the United States. This is partly due to a lack of computerized accounting databases available for academic research. The evidence is also more varied than that for the size effect. Countries in which an E/P effect has been examined include the United Kingdom, Japan, Singapore, Taiwan, Korea, and New Zealand. In the U.K. Levis (1989) documents a significant E/P effect for the period April 1961 to March 1985. He reports an average monthly premium of 0.58 percent (7.0 percent annually). The magnitude of the E/P effect in the United Kingdom is confirmed by the more recent work of Strong and Xu (1995). They report an average monthly premium of 0.60 percent for the extreme portfolios drawn from decile portfolios formed on E/P ratios over the period July 1973 to July 1992.[9] This premium is of the same order of magnitude as that observed in the U.S. and reported in Table 1. Adjusting portfolio returns for differences in systematic risk does not modify this conclusion.

Aggarwal, Hiraki, and Rao (1988) provide evidence of a significant E/P effect for a sample of 574 firms listed on the first section of the Tokyo Stock Exchange during the period from 1974 to 1983. Only firms with positive earnings were included in the sample. Portfolios of high E/P stocks outperformed those with low E/P stocks even after controlling for differences in systematic risk and size across portfolios. In the case of Singapore, Wong and Lye (1990)

[7]Some have argued that because firms in the same industry tend to have similar E/P ratios, a portfolio strategy that concentrates on low E/P stocks may indeed benefit from higher than average returns, but at a cost of reduced diversification. These arguments also suggest that the E/P effect may in fact be an industry effect. For example, during the 1980's financial firms and utilities comprised anywhere from 45 to 86 percent of the highest E/P quintile constructed from our sample of firms. Peavy and Goodman (1983) address this potential bias and examine the P/E ratio of a stock relative to its industry P/E (PER). They find a significant negative relation between PER's and abnormal returns over the 1970-1980 period. A portfolio strategy that bought the quintile of lowest PER stocks and sold short the highest PER quintiles would have yielded an annualized abnormal return of 20.80 percent over the period, although this number does not account for transactions costs.

[8]The table reports total returns that are not adjusted for risk. Since the betas are not substantially different across the portfolios, inferences drawn from total returns should not diverge in a meaningful way from inferences drawn from returns adjusted for beta risk.

[9]Levis (1989) also reports a size effect (see the evidence in Table 2 for the case of a slightly different sample characteristics), but it is weaker than the P/E effect. He also finds a large degree of interdependency between the two effects with the P/E effect tending to subsume the size effect.

show that there is a significant E/P effect on that country's stock market for the same sample of firms that revealed the presence of a significant size effect in Table 2. They conclude that the E/P effect is stronger than the size effect, although not independent of firm size.

For the Taiwanese stock market, Chou and Johnson (1990) report a significant P/E effect during the period 1979-1988 for a comprehensive sample of shares with positive earnings. They show that the average monthly return of the lowest quintile P/E portfolio exceeds that of the highest quintile P/E portfolio by 2.27 percent (27.2 percent annually). Chou and Johnson find that after adjusting for differences in systematic risk, the P/E premium is still significant with an average monthly return of 1.88 percent (22.6 percent annually). Ma and Chow (1990) report a weaker but still significant Taiwanese P/E effect for a smaller sample of stocks over the period 1979 to 1986. Dividing their sample into 5 portfolios, they found a significant average risk-adjusted monthly P/E premium of 0.85 percent (10.2 percent annually).

Finally, in New Zealand, Gillan (1990) finds no evidence of a P/E effect during the period 1977 to 1984 for the same sample as the one described in Table 2 for which he reports a significant size effect. A similar conclusion is reached by Kim, Chung, and Pyun (1992) for Korea, based on the same sample of firms used to examine the size effect reported in Table 2. They find no evidence of a P/E effect on that market during the period 1980-1988 for a sample of up to 224 stocks.

In summary, the evidence from six markets outside the United States indicates that in the United Kingdom, Japan, Singapore and Taiwan there is a significant P/E effect similar to that found in the U.S. market. There is no evidence, however, of a significant P/E effect in New Zealand and Korea. Given the small size and relatively short sample period for the cases of Taiwan, New Zealand and Korea, it is difficult to draw definitive conclusions from the evidence regarding these markets.

3.4. *Variations on the E/P effect: cash flow-to-price and sales-to-price ratios*

One alternative to the E/P ratio is the ratio of cash flow to price, where cash flow is defined as reported accounting earnings plus depreciation. Its appeal lies in the fact that accounting earnings may be a misleading and biased estimate of the economic earnings with which shareholders are concerned. Cash flow per share is less manipulable and, therefore, possibly a less biased estimate of economically important flows accruing to the firm's shareholders. The distinction between reported earnings and cash flow is important when examining these effects across countries with different accounting practices regarding the reporting of earnings. In some countries, such as Japan, firms are required to use the same depreciation schedule to calculate earnings reported to shareholders and earnings subject to corporate taxes. As a result, virtually all Japanese firms use accelerated depreciation for financial reporting (to reduce their tax liability) which creates large distortions in reported earnings for firms with large capital investments. In other countries, such as the United States, firms can use accelerated depreciation for tax purposes (which reduces taxable profits) and straight-line depreciation for reporting purposes (which produces relatively higher reported earnings to shareholders). Such accounting differences explain why there is a narrower difference between Japanese and American P/CF ratios compared to the much larger difference in the P/E ratios prevailing in these countries. For example, in August 1990, the market P/CF was 7.6 in the United States and 10.6 in Japan, whereas the market P/E was 15.8 in the United States and 35.3 in Japan (Goldman Sachs Research, August 1990).[10]

There is evidence of a CF/P effect in the United States and Japan. Chan, Hamao, and Lakonishok (1991) find evidence of a significant relation between average returns and CF/P for Japanese stocks. The U.S. evidence is summarized in the third set of columns in Table 1 which reports average returns and other portfolio characteristics for ten decile portfolios based on annual rankings (at March 31) of NYSE and AMEX securities on the ratio of cash

[10] French and Poterba (1991) adjust the E/P ratio for the Japanese and U.S. markets for differences in accounting techniques and report adjusted E/P ratios of 22.8 for Japan and 14.5 for the United States. Thus, holding accounting techniques constant does not eliminate the difference between the estimates. The remaining differences may be explained by a lower level of interest rate and a faster economic growth rate in Japan compared to the United States during that period.

flow per share to price per share (CF/P) for the period 1972 to 1994. Cash flow (CF) is defined as "reported earnings plus depreciation". The average difference in returns between the two extreme CF/P decile portfolios is 0.67 percent per month ($T = 2.32$). It is larger than the 0.56 percent per month ($T = 1.93$) obtained for the E/P effect for the same time period. This translates to an average annual difference between the two effects of more than 1.25 percent.

An alternative to both the E/P and CF/P ratios is the price-to-sales (P/S) ratio. Compared to earnings and cash flow, sales revenues are probably least influenced by accounting rules and conventions. There is evidence of a P/S effect in both the United States (Senchack and Martin (1987), Jacobs and Levy (1988)) and Japan (Aggarwal, Rao, and Hiraki (1990)). For example, during the period 1968 to 1983, a portfolio of Japanese stocks with the lowest P/S ratio had an average monthly return of 1.86 percent compared with 1.13 percent for the portfolio of stocks with the highest quintile of P/S.

3.5. The price-to-book effect

Although less research, both in the United States and other countries, has examined the ability of other variables to predict cross-sectional differences in security returns, the ratio of price-per-share to book-value-per-share (P/B) deserves mention because of its significant predictive power. As is the case for the other variables discussed above, there is no theoretical model which predicts that P/B should be able to explain the cross- sectional behavior of stock returns. However, investment analysts (e.g., Graham and Dodd (1940)) have long argued that the magnitude of the deviation of current (market) price from book price per share is an important indicator of expected returns.

A succession of papers (Stattman (1980), Rosenberg, Reid, and Lanstein (1985), DeBondt and Thaler (1987), Keim (1988), and Fama and French (1992)) have documented a significant negative relation between P/B and stock returns. To provide some perspective on the magnitude of the P/B effect, the fourth set of columns in Table 1 reports average monthly returns and other portfolio characteristics for ten decile portfolios drawn from the same data we used to examine the size, E/P, and CF/P effects in the U.S. market. The average monthly returns in Table 1 indicate a significant negative relation between P/B and returns. The monthly difference in returns between the extreme P/B portfolios (0.53%, $T = 2.31$) is higher than the corresponding differential return for the E/P effect (0.38%, $T = 1.68$), but lower than that for the size effect (0.72%, $T = 2.35$) for the 1962-94 period.

There is some evidence of a P/B effect outside the United States. A P/B effect has been documented for stocks trading on the Tokyo Stock Exchange (Aggarwal, Rao, and Hiraki (1989), Chan, Hamao, and Lakonishok (1991) and Capaul, Rowley, and Sharpe (1993)), the London Stock Exchange (Capaul, Rowley, and Sharpe (1993) and Strong and Xu (1995)), and also on stock exchanges in France, Germany, and Switzerland (Capaul, Rowley, and Sharpe (1993)). Capaul, Rowley, and Sharpe (1993) report the following average monthly values for the difference in returns between lowest and highest P/B portfolios: 0.53 percent in France; 0.13 percent in Germany; 0.50 percent in Japan; 0.31 percent in Switzerland; and 0.23 in the United Kingdom.

3.6. Prior return (reversal and momentum) effect

There is evidence that prior returns can explain the cross-sectional behavior of subsequent stock returns. The literature documents two (seemingly) unrelated phenomena. The first is the existence of return reversals (past “losers” become “winners” and vice versa) over both *long-term* horizons (3 to 5 years) as well as *very short-term* periods (a month and shorter). The second is the presence of an opposite effect over horizons of *intermediate* lengths: When prior returns are measured over periods of 6 to 12 months, “losers” and “winners” retain their characteristic over subsequent periods. There is, in this case, return *momentum* rather than reversal.

DeBondt and Thaler (1985) and DeBondt and Thaler (1987) find that NYSE stocks identified as the biggest losers (winners) over a period of 3 to 5 years earn, on average, the highest (lowest) market-adjusted returns over a subsequent holding period of the same length of time. This reversal effect does not seem to disappear when returns are adjusted for size and risk (Chopra, Lakonishok, and Ritter (1992)). Ball, Kothari, and Shanken (1995) show that the abnormal returns associated with these strategies are sensitive to the portfolio formation date, which they attribute to microstructure-related biases that are most pronounced at the calendar year end. Specifically, they find negative abnormal returns when the strategy is initiated in June, in contrast to the positive abnormal returns when the strategy is initiated in December, the typical initiation date in this literature. Evidence of long-term return reversals has also been reported in a number of markets outside the United States, including Belgium (Vermaelen and Versringe (1986)), Japan (Dark and Kato (1986)), Brazil (Costa (1994)) and the United Kingdom (Clare and Thomas (1995) and Dissanaike (1996)). The reversal effect is not evident on the Toronto Stock Exchange (Kryzanowski and Zhang (1992)).

There is also evidence of short-term return reversals. Jegadeesh (1990), Keim (1983a), and Lehmann (1990) show that a "contrarian strategy" that selects stocks on the basis of return performance over the previous week or month earn significant subsequent returns. Chang, McLeavey, and Rhee (1995) report the presence of the same phenomenon in Japan, after adjusting returns for both beta risk and size. But contrary to long-term reversals, that are often interpreted as "market overreaction" due to irrational investors, short-term reversals may possibly reflect a lack of market liquidity (Jegadeesh and Titman (1993)) and the delayed reaction of stock prices to common factors (Lo and MacKinlay (1990)).

Return momentum has been documented by Jegadeesh and Titman (1993), Chan, Jegadeesh, and Lakonishok (1995), and Asness (1995). Jegadeesh and Titman (1993) show that buying past winners and selling past losers generates significant positive returns over holding periods of 3 to 12 months, an investment strategy that does not seem to be due to differences in risk or delayed stock price reactions to common factors.

The last set of columns in Table 1 provides evidence of return momentum in our sample (we do not examine reversal strategies). Consistent with the procedures used in previous studies, we measure prior returns over the six months prior to the portfolio formation month, but exclude from this calculation the return from the last month of the six-month period to avoid possible contamination of the results from (microstructure-induced) return reversals for the adjacent monthly returns that bracket our portfolio formation date. Specifically, our prior return is measured over the 5-month period from the beginning of October to the end of March. Consistent with previous research, portfolios with the highest prior returns (the winners) earn, on average, higher subsequent returns. Also, portfolios with the lowest prior returns (the losers) earn, on average, the lowest subsequent returns. The difference in monthly returns between the extreme portfolios is 0.34 percent ($T = 1.66$) for the 1962-1994 period.

4. One Effect or Many?

4.1. Price as a Common Denominator

Size, E/P, CF/P and P/B are computed using a common variable: price per share. Blume and Stambaugh (1983) and Stoll and Whaley (1983) explored the relation between size and price and reported evidence suggesting a high rank correlation between size and price.[11] Keim (1988), Jaffe, Keim, and Westerfield (1989) and Fama and French (1992) have all recently raised this possibility regarding the other effects. We examine the association among these variables using pairwise Spearman rank correlations. The rank correlations and their associated T-values are computed as follows. Each year at the end of March, all NYSE and AMEX stocks are ranked independently on size, E/P, P/B, CF/P, prior return, and price (i.e., six separate rankings are produced). Each variable was computed using price per share at March 31 and, when applicable, accounting numbers for the previous year, with only December fiscal closers included in the rankings. Pairwise Spearman rank correlations are then computed. This procedure is repeated in each year for the period 1962 to 1994, and mean rank correlations and standard errors are computed for the entire time series of values. Table 3 reports the average rank correlations and associated T-values. The estimated rank correlations are generally large and significant. The pairwise correlations among the size, E/P, CF/P, P/B, prior return, and price rankings are significantly different from zero,

[11]Results in Blume and Husic (1973), Stoll and Whaley (1983) and Blume and Stambaugh (1983) also reveal a significant cross-sectional relation between price per share and average returns.

indicating some commonalities among the effects. Consistent with previous work (Blume and Stambaugh (1983) and Stoll and Whaley (1983)), we find that the rank correlation between market capitalization and price was by far the strongest (0.78).

4.2. Correlation of the Premia

The difference in return between the extreme portfolios in Table 1 (e.g., P/B portfolio 1 minus P/B portfolio 10) can be loosely interpreted as risk premia, if these variables are sorting out securities based on risks that are not defined by extant asset pricing models. Under the scenario that the five variables discussed above are proxies for five separate risk "factors", then the premia should be uncorrelated across variables. In Table 4 we report the pairwise correlations between the monthly premia. All of the correlations are large (in absolute value) and are significantly different from zero. Interestingly, the prior return premia are *negatively* correlated with the premia associated with the other four variables, suggesting that prior return is capturing a characteristic of stock returns that is quite different from the other variables. Note that the correlations are also significantly different from 1.0, indicating that the variables are not all proxying for the same underlying characteristic. Nevertheless, the significant correlations indicate a high degree of commonality among the effects.

4.3. The Premia are concentrated in January

The significant correlation of the premia reported in section 4.2 is in part a reflection of the fact that these effects are most pronounced in January. Specifically, the average premia during January tend to be positive and are usually significantly larger than the average premia measured during the rest of the year. This seasonality was first documented for the size premia (Keim (1983b))[12], and subsequently for the E/P premia (Cook and Rozeff (1984), Jaffe, Keim, and Westerfield (1989)), the P/B premia (Keim (1988), Fama and French (1992), and Loughran (1996)), and the momentum premium (Asness (1995)).

[12] Blume and Stambaugh (1983) show that, after correcting for an upward bias in average returns for small stocks, the size-premium is evident only in January.

Internationally, most of the evidence on January seasonality has been related to the size premium. A significant January seasonal has been reported in Belgium (Hawawini, Michel, and Corhay (1989)), Finland (Berglund (1985)), Taiwan (Ma and Chow (1990)) and Japan (Ziemba (1991) and Hawawini (1993)).[13] Countries in which the January size premium is insignificant include France (Hamon (1986)), Germany (Stehle (1992)) and the United Kingdom (Levis (1985)). Chan, Hamao, and Lakonishok (1991) report a significant January seasonal for E/P, P/B, and CF/P in Japan but, surprisingly, no January seasonal for size.

A January seasonal in the size premium that is computed with our value-weighted size portfolios is summarized in the first set of columns in Table 5 where the "all months" returns generated by the ten size portfolios in Table 1 have been separated into "January" returns and "rest-of-the-year" returns. The results in Table 5 show that average returns are significantly larger in January than the rest of the year for all size-portfolios except the largest one (portfolio ten). Further, the results show a clear negative relation between January returns and market capitalization: The smallest size portfolio has a January return of 12.11 percent whereas the largest size portfolio has an average return of only 1.94 percent. The average January size premium is thus 10.17 percent. This is significantly larger than observed in markets outside the United States and reported in the studies mentioned above. For example, the January size premium is 2.4 percent in Belgium (data from 1969 to 1983 for five size portfolios), 3.4 percent in Finland (data from 1970 to 1983 for five size portfolios) and 7.2 percent in Japan (data from 1965 to 1987 for ten size portfolios), one of the largest January size premium reported outside the U.S.

Table 5 also reports on January seasonality in the premium associated with earnings yield (E/P), cash flow yield (CF/P), price-to-book ratio (P/B) and 6-month prior return. The results are similar to those reported for the size premium, but are not as strong (see also figure 1). Average returns are higher in January than in the rest of the year, but the difference is only significant for portfolios with the highest E/P and CF/P, and lowest P/B and 6-month prior returns (the deciles for which the January effect is significant are identified

[13] Note, however, that Chan, Hamao, and Lakonishok (1991) do not find a significant January size premium in their Japanese data.

by asterisks). Also, unlike the other four variables, the relation between returns and prior returns is negative in January and positive in February to December (see also Fant and Peterson (1995)). That the relation between prior returns and subsequent returns switches signs in January precisely when the other effects are more exaggerated probably explains why the prior return premium has a negative correlation with the other premia (see section 4.2). Second, the magnitude of the January premium varies depending on the sorting variable used to construct the portfolios. The January premium is largest for portfolios based on size (10.17 percent) but is insignificant for portfolios based on 6-month prior return.

5. Can we disentangle the effects?

The evidence in section 4 suggests a great deal of commonality among the various effects. In this section we discuss the research that has attempted to disentangle the interrelation between the effects in an effort to assess whether the results are due to one or several different factors.

5.1. Interactions between the effects: A somewhat chronological perspective

The earlier studies tended to examine interactions between two effects at a time. Much of the initial research in the area focused on the interrelation between the E/P and size effects, and a variety of techniques have been used, ranging from simple analysis of average portfolio returns to sophisticated regression techniques. The disparate methods used often make comparisons difficult. In the end, the results are less than conclusive. For example, Reinganum (1981) argues that the size effect subsumes the E/P effect: Once we control for size, there is no marginal E/P effect. Recall that the same phenomenon was reported in New Zealand (Gillan (1990)). Basu (1983) argues just the opposite. So do Wong and Lye (1990) in the case of Singapore and Chou and Johnson (1990) in the case of Taiwan. Peavy and Goodman (1983) and Cook and Rozeff (1984) perform meticulous replications of, and extensions to, the methods of Basu and Reinganum, and reach surprisingly different conclusions. Peavy and Goodman's results are consistent with Basu's while Cook and Rozeff conclude that no one effect dominates the other. Banz and Breen (1986) find a size effect but no independent E/P effect, a result similar to Reinganum.

Jaffe, Keim, and Westerfield (1989) argue that the inability to disentangle the two effects may be attributable to the relatively short, and sometimes nonoverlapping, periods used in the above studies (they range from 8 to 18 years) as well as the failure of the studies (with the exception of Cook and Rozeff (1984)) to account for potential differences between January and the other months. Using data covering a 36 year period, Jaffe, Keim, and Westerfield (1989) find that after controlling for size there is a significant E/P effect in both January and the other months; controlling for the E/P effect, there is a significant size effect only in January.[14] They also conclude that the results of the earlier studies conflict because the magnitude of the two effects is period specific. Fama and French (1992) reach similar conclusions regarding the joint significance of size and E/P effects (see the regression results in their table 3).

Stattman (1980) and Rosenberg, Reid, and Lanstein (1985) were the first to examine the interaction between size and P/B for NYSE and AMEX stocks. Stattman examines average risk-adjusted portfolio returns for April 1964 to April 1979 and concludes that "even after taking account for the size effect, there remains a positive relationship between B/P [the inverse of P/B] and subsequent returns". Rosenberg, Reid and Lanstein examine market model residuals of P/B portfolios that are constructed to be orthogonal to size and other influences. They also find a significant relation between abnormal returns and P/B for the 1973-1984 period.

A number of studies have examined the interaction between (contrarian) return reversals and both size and beta risk. Keim (1983a) and Jegadeesh (1990) find that short-term (monthly) reversals provide explanatory power for average returns beyond the influence of size. Chan (1988) and Ball and Kothari (1989) argue that the abnormal risk-adjusted returns reported for long-term (3 to 5 year) contrarian investment strategies are due to inadequate adjustment for risk. That is, a loser firm whose stock price (and therefore market capitaliza-

[14]This result is consistent with other studies which have found that when examined alone, the size effect is significant only in January (e.g., Blume and Stambaugh (1983)). See section 3.2.

tion) has declined, in the absence of a concomitant decline in the value of the debt, becomes more leveraged and, other things equal, more risky. Traditional methods for computing betas may underestimate the CAPM-implied risk for these firms and, therefore, overstate the abnormal return. In this vein, Zarowin (1989) shows that loser firms tend to be small firms and winner firms tend to be large firms. After controlling for the size effect, Zarowin finds insignificant evidence of contrarian profits. More recently, Ball, Kothari, and Shanken (1995) show that contrarian results are sensitive to the timing of portfolio formation in the empirical simulations of contrarian portfolio strategies. They show that when portfolios are formed by grouping based on past performance periods ending in mid-year rather than in December, the reversal effect is significantly reduced. In contrast to the above studies, Chopra, Lakonishok, and Ritter (1992) find that after adjusting for size and beta there is an economically important overreaction effect, especially for small stocks. Chopra, Lakonishok, and Ritter (1992) find, though, that these contrarian profits are "heavily concentrated in January," suggesting they are related to the January effect. They conclude that their findings are not due to tax-loss-selling, but the prominent role played by small stocks in their findings suggests that the buying and selling behavior of individual investors may be important. Clearly more work is needed to sort out these issues.

The consensus from the research detailed above is that the relation between market capitalization and average returns is quite robust.[15] In addition, variables such as E/P and P/B seem to provide explanatory power for cross-sectional differences in average returns beyond the influence of size. Fama and French (1992) attempt to sort out the relative explanatory power of the three variables. They perform their test on a sample of individual stocks[16] trading on the NYSE and AMEX over the 1963-1990 period and the NASDAQ

[15] An exception to the above findings is the recent study by Chan, Hamao, and Lakonishok (1991) who find that in the Japanese stock market P/B and CF/P are sufficient to characterize cross-sectional differences in expected returns. That size is unimportant in explaining expected returns appears to be unique to the Japanese and Korean markets.

[16] Most previous research has demonstrated that estimates of beta do not enter significantly into models like the one represented in equation (2) in the presence of other explanatory variables such as size and E/P (an exception is Chan and Chen (1988)). Thus, Fama and French argue that methodologies that use portfolios to avoid estimation error in individual beta estimates unduly forfeit the valuable information in the cross section of individual security characteristics such as market value or E/P. This latter point was

market for the 1973-1990 subperiod. They estimate equation (2) with beta, market cap, book-to-market ratio (B/P) and earnings yield (E/P). Their regression results indicate that beta and E/P are insignificant while both size and B/P have predictive power. Based on their findings, they conclude that size and B/P are sufficient to characterize cross-sectional differences in expected returns. They conclude that three "risk" factors - market, size, and P/B - are sufficient to explain the cross section of expected stock returns (see also Fama and French (1993)). However, in a subsequent paper (Fama and French (1996)), they report that size and B/P cannot explain the momentum effect. In the next subsection, we reexamine the interrelation between size, B/P and momentum with the data from section 3.

5.2. Interactions between the size, P/B and prior-return effects: Another look

To facilitate the comparison of the interaction between size, the P/B effect and the 6-month prior-return performance, we use our sample of NYSE and AMEX stocks to compute size-adjusted returns for portfolios created on the basis of both P/B and prior returns. Briefly, at the end of March in each year from 1962 to 1994 we sort all NYSE and AMEX stocks by P/B and form five groups of equal numbers of securities based on their P/B ranking.[17] The book value in our ratio is a December 31 value, the market cap is from March 31. Within each of these groups we again rank the stocks according to the magnitude of their prior returns (calculated from the beginning of October to the end of February) and create five subgroups within each of the five P/B groups. Individual stock returns are adjusted for the influence of the size effect by simply subtracting the return corresponding to the size decile (see Table 1) of which that security is a member. The composition of the portfolios remains constant over the next twelve months and value-weighted size-adjusted portfolio returns are computed each month.

The results are reported in Table 6. First, there is a positive relation between prior returns and subsequent returns within each of the five P/B quintiles, except for the highest

emphasized earlier in Litzenberger and Ramaswamy (1979) in their analysis of the relation between stock returns and dividend yields.

[17]For purposes of this experiment, we eliminate from the sample all stocks with negative P/B values.

P/B quintile in which subsequent returns drops from 0.28 percent to 0.02 percent when we move from the fourth to the fifth prior-return quintile. Second, there is evidence of a negative relation between P/B and subsequent returns within each of the five prior-return quintiles, although this relation is strongest for the quintile of stocks with negative prior returns. As prior returns increase, the negative relation between P/B and subsequent returns is confined to the first three P/B quintiles.

Note that the P/B effect occurs in the prior-return categories in which the average price of the securities in the portfolios increases with P/B[18] (except for the highest prior-return category in which average price first rises to $40.7 and then drops to $32.8 to rise again to $39.3). Although the experiment controls for the influence of size (market capitalization) on average returns, it does not explicitly control for price, which has also been shown to influence returns. Thus, the high average returns for low P/B stocks may reflect some underlying relation between returns and low price.

An alternative hypothesis involves the possibility that low P/B stocks are simply stocks whose prices have dropped relative to book values that vary little through time. Firms whose stocks have recently declined in price, in the absence of a concomitant decline in the value of the debt, have become more leveraged and, other things equal, more risky. Traditional estimation methods may underestimate "true" beta risk for such firms and, therefore, overstate "risk-adjusted" returns. [19] As a result, stocks that have recently suffered a substantial decline in price will tend to have underestimated betas and low ratios of P/B. Hence, P/B may be a more accurate proxy for "true" beta risk than traditional estimates of beta due to the measurement error in the traditional estimates. Given their unobserved

[18]In this regard, recall that we found a significant rank correlation between price and P/B (0.33).

[19]Traditional methods (e.g., OLS) that have been used for estimating betas in most cross-sectional analyses use four or more years of monthly returns data and (implicitly) apply equal weights to all observations in the time series. Clearly, the most relevant observations, the ones that should be given the most weight in the estimation, are those occurring closest to the period of analysis (e.g., portfolio formation date). Thus, the betas used in such studies are "stale" in that they are estimated using information that, in large measure, is not relevant. This estimation shortfall also applies to studies that use "future" betas estimated from data occurring after the analysis interval since structural changes that impact firm risk can also affect the post-analysis observations, thereby rendering them less relevant for assessing the firm's risk in the analysis interval.

higher levels of risk, the subsequent higher average returns that compensate for this risk appear anomalous.[20]

Previous research described above shows that the size and P/B effects are most concentrated in the month of January. In Table 7, therefore, we report the average returns for the 25 portfolios in Table 6 separately for January and for the remaining eleven months. Consider the P/B effect first. The P/B effect in January (Panel A) is more pronounced than when computed over all months as in Table 6. In February-December, though, the relation between P/B and returns is flat. Thus, after controlling for size and momentum effects, the P/B effect is evident in the data only during the month of January. The story is quite different for the momentum effect. After controlling for both size and P/B, the relation between prior and subsequent returns has the same significant positive relation as noted in Table 6 for all months. In January, though, the "momentum" effect has the appearance of a reversal effect in that subsequent returns increase as prior returns decrease. That is, stocks that have recently declined (and therefore more likely candidates for trading at the end of the year based on taxes or window dressing) have the largest returns in January, particularly if they are low-price stocks. Thus, in January, it appears that end-of-year trading patterns tend to offset the momentum effect that, unlike the other effects, persists throughout the rest of the year.

6. In Search of an Explanation

Fama and French (1992) argue that the significant relation between returns and variables like size, P/B and prior return is evidence of compensation for additional sources of risk that are not included in extant asset pricing models. A plethora of papers have appeared recently that question this interpretation. They fall into four categories.

[20] DeBondt and Thaler (1987) show that such price reversals are most extreme for low P/B stocks.

6.1. Beta-adjusted excess returns are evidence of market inefficiency

Several studies interpret the reported excess returns as evidence of market inefficiency. For example, Lakonishok, Schleifer, and Vishny (1994) argue that investors are irrational because they avoid "value" stocks that they mistakenly consider too risky, even though the evidence indicates that this does not appear to be the case (at least for conventional measures of risk). According to Haugen (1995), institutional investors avoid buying "value" stocks because these investors' performance is measured against indexes of mostly large, "glamour" stocks. Again, individuals who buy these neglected "value" stocks outperform the indexes of large, "glamour" stocks in the long run.

6.2. The results are due to statistical biases or peculiarities in the data sample

Recent papers by Kothari, Shanken, and Sloan (1995) and Breen and Korajczyk (1994) suggest that the P/B results may be due to survivor biases in samples of stocks drawn from the Compustat files. Compustat tends to include (and add) stocks in their files only after the stock has demonstrated a successful track record. Thus, small firms with low P/B ratios that subsequently perform poorly (or fail) are unlikely to appear on the files. However, Davis (1994) and Kim (1997) provide evidence that appears contrary to this hypothesis. Davis (1994) finds evidence of P/B, cash flow and E/P effects for a survivor-bias-free sample in the 1940-1963 period prior to the Compustat-specific period used in many studies, although his estimates of the P/B effect are half those estimated by Fama and French (1992). Kim (1997) shows that survivor bias does not significantly reduce the P/B effect.

Also in this second category are papers that show that the magnitude of the P/B effect is sensitive to the types of securities analyzed. For example, Loughran (1996) argues that the relation between P/B and returns is a manifestation of the low returns for Amex and Nasdaq growth stocks used in those studies. Loughran shows that the P/B effect is not significant for large-cap stocks or small-cap NYSE stocks during the 1963-1991 period. Based on his results, Loughran concludes that "for the majority of money managers, the empirical findings of Fama and French (1992) are of little economic importance for predicting future portfolio returns".

Finally, Berk (1995) and Berk (1997) claim that the reported relation between firms' returns and their market value is not at all a "size" effect. He makes the claim in light of his findings that there is no significant relation between average returns and four measures of firm size (book value of assets, book value of gross fixed assets, annual sales, and number of employees). Berk concludes that his "results are evidence in favor of the hypothesis that the size effect is due to the endogenous identity relating the market value of a firm to its discount rate".[21]

6.3. *Three factors are not enough*

Some argue that more than three variables are required to characterize the multidimensional nature of risk. This notion has some precedence in the literature. A number of earlier papers (e.g., Rosenberg and Marathe (1979), Sharpe (1982), Jacobs and Levy (1988)) estimate cross-sectional regression models similar to Fama and French but using a larger and richer set of independent variables. The renewed willingness of the academic community to entertain such analyses since the publication of Fama and French has resulted in much activity in this area. For example, Barbee, Mukherji, and Raines (1996) examine the explanatory power of the ratio of sales to price when added to size, book-to-market ratio and leverage (debt-to-equity ratio) in a multiple regression, and find that it has significant explanatory power. Jagannathan and Wang (1996) suggest that the lack of empirical support for the CAPM may be due to: (1) the exclusion of human capital in the market portfolio; and (2) the systematic variability of beta-risk over the business cycle. They account for human capital by estimating the sensitivity of portfolio returns to the growth rate in per capita labor income. The variation in beta is taken into account by estimating the sensitivity of portfolio returns to the yield spread on low- and high-grade corporate bonds When added to beta and

[21] Berk argues that since "a firm's market value is endogenously determined in equilibrium as the discounted value of expected future cash flows, it depends on the discount rate. For example, if two firms have the same expected cash flow, the one with the larger discount rate will have the lower market value. Consequently, according to this view expected returns will always be negatively correlated with market value, *ceteris paribus*".

size, the two new variables emerge as the only variables with significant predictive power. Asness (1995) examines the marginal contribution of (intermediate-term) price momentum and (long-term) price reversals to the size and B/P effects in a multiple regression. He finds that both prior return variables are significant in explaining the cross-sectional behavior of subsequent returns, beyond the contribution of size and B/P.

6.4. The findings are real, but cannot be attributed to risk

Loughran (1996) shows that for NYSE stocks, the size and P/B ratio explain none of the cross-sectional variation in returns if January is excluded from the sample (see also Keim (1988) and table 5 above). It is difficult to characterize the relation between *ad hoc* factors and returns as risk-based when it manifests itself primarily in January. In a similar vein, Daniel and Titman (1995) show that the significant coefficients on size and P/B in regressions like (2) are simply correlation between returns and those characteristics, and not implied risk premia associated with risk factors. Similarly, in the following section we report some new evidence on size, P/B, E/P and CF/P premia across several developed equity markets that appear to be inconsistent with a risk story.

6.4.1. Are the premia compensation for risk: Some more international evidence

Previous sections surveyed evidence of size, E/P, CF/P and P/B effects on many international stock markets. In this section we present some new evidence to confirm these findings. In the same spirit as the evidence we presented for the U.S. market, we present findings for a common time period using the same empirical methods across all markets. We again report our evidence in portfolio form because we feel the returns to feasible portfolios provide a useful perspective on the economic significance of the results.

The data for the international equity markets are from Morgan Stanley Capital International (MSCI) for the period January 1975 to December 1994.[22] The data include all stocks in the MSCI indexes for France, Germany, Japan and the U.K. The stocks in the sample are

[22] We are grateful to Brandywine Asset Management, especially Steve Tonkovich, for providing access to these data.

the most actively traded and popular stocks, and represent the bulk of the market capitalization, in their respective markets. In addition, stocks with negative or missing values for the ranking variables and stocks that have extremely small market capitalization relative to the local market are excluded.[23] Returns include dividends and are adjusted for the local investor tax credit by country.[24] Finally, returns are computed in home currencies. Note that for the premia results, in which we are most interested, any currency influences will be effectively canceled, allowing for clean cross-country comparisons.

In each market we construct four sets of portfolios based on independent rankings on size (market capitalization), E/P (ratio of earnings to price), CF/P (ratio of cash flow to price), and P/B (ratio of price per share to book value per share). Rankings are conducted at December 31 of each year using end-of-year market cap and price and the most recently-reported accounting values (not restated). Note that all of these values are recorded in the published versions of the MSCI data as of the portfolio formation date so that the values used to construct the portfolios are known to market participants. Thus, there is no look-ahead, fill-in or survivor bias in the portfolios or the associated tests. The portfolios are buy-and-hold portfolios that are equal-weighted at the time of portfolio formation.[25] The portfolios are held for twelve months, at which time the sorting and portfolio rebalancing procedure is repeated. The process is repeated each year from December 1973 to December 1994 resulting in a 20-year time series of monthly returns. The U.S. portfolios are value-weighted (quintile) portfolios of NYSE and AMEX stocks based on the U.S. data described in previous sections.

Monthly premia associated with each of the sorting variables are constructed as the difference between the extreme quartile portfolio returns in each month. For example, the

[23]For example, a stock was excluded from the portfolios formed at December 1991 if its market cap, stated in millions of U.S. dollars, was \$56 in France, \$64 in Germany, \$91 in the U.K., and \$449 in Japan. These lower bounds changed from year to year as market conditions varied.

[24]The adjustment is conservative, implying the maximum that foreign investors might pay: 25% in U.K.; 36% in Germany; 33% in France; and 0% in Japan.

[25]Since there is not substantial diversity in market capitalization values within quartiles in the MSCI data, the equal weighting of the MSCI portfolios at the portfolio formation date will not produce results that differ in a meaningful way from value-weighted portfolios.

E/P premium is computed as the difference in return between the highest E/P quartile stocks and the lowest E/P quartile stocks. Table 8 contains the average and standard deviation of the premium for each ranking variable within each country. The premia are economically and statistically significant for all the variables across all the countries (one exception is the size effect in Germany). Further, the magnitude of the effects in the four international markets is comparable to that computed for the U.S and reported in previous sections.

Fama and French (1992) argue that premia like those reported in table 8 are compensation for risk. If that is the case, under conditions of integrated international capital markets, one expects that the premia will be correlated across markets (e.g., the size premia should be correlated across markets). To this end, Table 8 reports, within each panel, the correlations across countries for a particular variable. Our objective is to assess the degree of correlation among premia *across* countries.[26] The correlations are mostly insignificantly different from zero.[27] This lack of correlation of the premia across markets contrasts with the significant correlation of the *market* risk premium across international markets that has been widely documented. The lack of correlation in the size, E/P, P/B, and CF/P premia across markets calls into question risk-based explanations for these premia, at least in the context of extant international asset pricing models.[28]

7. Concluding Remarks

A large body of evidence has accumulated that collectively suggets that betas do not adequately explain cross-sectional differences in average stock returns. Instead, several variables that have no basis in theory sem to have significant explanatory power. In addition, many papers have attempted to explain these findings. In this paper we provide a synthesis of the evidence and explanations. One interpretation is that the beta-adjusted excess returns

[26]We have also examined whether *different* ranking variables produce similar results *within* a country. The within-country correlations are low, suggesting that the different variables seem to be picking up different sources of variability in returns within each country. The correlations tend to be highest among the P/E, P/B, and P/CE premia. The size premia displays a lower correlation with the other premia within a particular market. These results are consistent with existing research for U.S. equities.

[27]Evidence from two ten-year subperiods show that this lack of correlation is not period specific.

[28]The low correlations also suggest the possibilty of substantial benefits from international diversification of value and size portfolio strategies.

represent evidence that a multidimensional model of risk and return is necessary to explain the cross section of stock returns. That is, beta by itself is insufficient to characterize the risks of common stocks. Prominent in this category are Fama and French (1992) who suggest a three-factor equity-pricing model to replace the CAPM. Their three-factor model adds two empirically-determined explanatory factors: size (market capitalization) and financial distress (B/M). Others proposes an additional factor, prior return.

Our findings suggest that such conclusions may be premature. Aside from the absence of a theory that predicts such variables have a place in the risk-return paradigm, the evidence strongly indicates that the statistical relation between returns and variables like size and B/M derives primarily from the month of January. It is difficult to tell an asset pricing story where risk manifests itself only during one month of the year. In addition, we find that the premia are uncorrelated across international equity markets. It is difficult to reconcile this finding with a risk-based story in which, under conditions of integrated international capital markets, the risk premia from a multidimensional international asset pricing model are correlated across markets in the same way the market risk premium is correlated across markets.

These points notwithstanding, a significant contribution of this line of research is that is has sharpened our focus on potential alternative sources of risk, and future theoretical work should certainly benefit. However, the evidence, in and of itself, does not constitute proof that the CAPM is "wrong". For example, no one has yet conclusively shown that variables like size and P/B are not simply proxies for measurement error in betas. Are we certain that cross-sectional variation in P/B is not picking up variation in leverage that is not reflected in betas that are typically estimated with sixty months of prior - and arguably stale - prices? The book is not closed; we think more research is necessary to resolve these issues.

There is also the question of believability. That is, is the evidence as robust as the sheer quantity of results would lead us to believe? First, there is the issue of data snooping - many of the papers we have cited in this article were predicated on previous research that documented the same findings with the same data. Degrees of freedom are lost at each turn and several authors have warned about adjusting tests of significance for these lost degrees of freedom. Also, the existence of these patterns in our experiments does not necessarily imply that they exist in the returns of implementable portfolios - that is, returns net of transactions costs. (e.g., market illiquidity and transactions costs may render a small stock strategy infeasible). Finally, that many of these effects have persisted for nearly 100 years in no way guarantees their persistence in the future. How many years of data are necessary to construct powerful tests? Research over the next 100 years will, we hope, settle many of these issues.

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Table 1

Monthly percentage returns (Standard errors) and betas for value-weighted portfolios of NYSE and AMEX stocks formed on the basis of size (market capitalization), earnings-to-price ratios, cash flow-to-price ratios, price-to-book ratios, and prior return[1]
(April 1962 to December 1994)

Portfolio[2]	Size (Market cap.)			Earnings-to-price ratio			Cash flow-to-price ratio[3]			Price-to-book ratio			Prior return (%)		
	Size ($ million)	Return (%)	Beta	E/P	Return (%)	Beta	CF/P	Return (%)	Beta	P/B	Return (%)	Beta	Prior Return (%)	Return (%)	Beta
1	$10	1.56 (0.37)	1.11	19.39	1.21 (0.27)	1.01	52.08	1.47 (0.33)	1.00	0.57	1.43 (0.28)	1.04	53.1	1.18 (0.29)	1.13
2	$26	1.41 (0.34)	1.14	12.88	1.25 (0.23)	0.93	27.75	1.32 (0.29)	0.90	0.84	1.42 (0.25)	0.97	24.9	1.24 (0.26)	1.05
3	$48	1.25 (0.31)	1.10	11.26	1.08 (0.23)	0.88	23.02	1.17 (0.29)	0.91	1.02	1.06 (0.23)	0.92	16.7	1.09 (0.24)	1.02
4	$83	1.23 (0.31)	1.15	10.09	1.02 (0.23)	0.95	19.91	0.94 (0.31)	0.99	1.18	1.05 (0.21)	0.84	11.2	1.03 (0.23)	1.02
5	$104	1.22 (0.28)	1.10	9.08	0.96 (0.23)	0.94	17.37	1.14 (0.31)	1.01	1.35	1.00 (0.22)	0.90	6.5	0.88 (0.22)	0.96
6	$239	1.12 (0.26)	1.04	8.14	0.77 (0.23)	0.99	15.05	0.87 (0.30)	0.99	1.56	0.79 (0.22)	0.91	2.3	0.91 (0.22)	0.93
7	$402	1.09 (0.25)	1.06	7.19	0.83 (0.22)	0.96	12.96	1.12 (0.30)	1.03	1.86	0.84 (0.23)	0.98	−1.9	0.85 (0.23)	0.95
8	$715	1.09 (0.24)	1.05	6.13	0.89 (0.24)	1.04	10.85	1.05 (0.32)	1.08	2.30	0.91 (0.24)	1.03	−6.6	0.92 (0.24)	0.96
9	$1,341	1.03 (0.23)	1.03	4.78	0.88 (0.25)	1.06	8.40	0.89 (0.31)	1.06	3.10	0.82 (0.25)	1.11	−13.1	0.62 (0.27)	1.05
10	$5,820	0.83 (0.21)	0.95	2.49	0.82 (0.26)	1.08	4.77	0.80 (0.33)	1.07	10.00	0.90 (0.25)	1.07	−29.6	0.83 (0.28)	1.15

[1] Portfolios are created on March 31 of each year using prices and shares outstanding on March 31 and accounting data for the year ending the previous December 31 (with portfolios containing only December 31 fiscal closers). Prior return is calculated from the beginning of October to the end of February immediately preceding the March formation month. Aside from new listings and delistings, which are added to or dropped from the portfolios as they occur during the year, the portfolio composition remains constant during the following twelve months over which portfolio returns are calculated. The data reported in the table are average values over the sample period. Betas are computed relative to the S&P 500 Index.

[2] Portfolio 1 (portfolio 10) is the portfolio with the smallest (largest) market capitalization, highest (lowest) E/P and CF/P, lowest (highest) P/B, and largest (smallest) 6-month prior returns.

[3] All results for the Cash Flow-to-Price portfolios are for the period April 1972–December 1994.

Table 2

The Size Effect: International Evidence

(Size Premium $= \bar{R}_{\text{Small}} - \bar{R}_{\text{Large}}$)

	Monthly Size Premium (%)	Test Period	Number of Portfolios	Largest Size/[2] Smallest Size
Australia	1.21	1958–81	10	NA
Belgium	0.52	1969–83	5	188
Canada	0.44	1973–80	5	67
Finland	0.76	1970–81	10	133
France	0.90	1977–88	5	NA
Germany	0.49	1954–90	9	NA
Ireland	0.47	1977–86	5	NA
Japan	1.20	1965–87	10	NA
Korea	−0.40	1984–88	10	62
Mexico	4.16	1982–87	6	37
New Zealand[3]	0.51	1977–84	5	60
Singapore[4]	0.41	1975–85	3	23
Spain[3]	0.56	1963–82	10	228
Switzerland	0.52	1973–88	6	99
Taiwan[5]	0.57	1979–86	5	17
U.K.	0.61	1973–92	10	182
U.S.	0.61	1951–94	10	490

[1] **Sources**: *Australia*: Brown et al. (1983); *Belgium*: Hawawini et al. (1989); *Canada*: Berges et al. (1984); *Finland*: Wahlroos et al. (1986); *France*: Louvet et al. (1991); *Germany*: Stehle (1992); *Ireland*: Coghlan (1988); *Japan*: Ziemba (1991); *Korea*: Kim et al. (1992); *Mexico*: Herrera et al. (1994); *New Zealand*: Gillan (1990); *Singapore*: Wong et al. (1990); *Switzerland*: Cornioley et al. (1991); *Taiwan*: Ma et al. (1990); *United Kingdom*: Strong and Xu (1995).

[2] Ratio based on average market values (median in Singapore) over sample period, except for Great Britain where it is calculated in 1975 and Finland in 1970. NA = Not Available.

[3] Returns, for this country, are not raw returns but risk-adjusted, *excess* returns estimated with the Capital Asset Pricing Model.

[4] Note that Wong et al. (1990) report that the size effect in Singapore appears to be of secondary importance when compared with the E/P effect.

[5] *Excess* returns estimated by subtracting from each individual security the return of the portfolio to which they belong. Note that another study of the size effect in Taiwan (Chou and Johnson (1990)) finds no evidence of a significant size effect after controlling for an E/P effect.

Table 3
Average Rank Correlations (t-statistics)
Between Several Predetermined Characteristics
For NYSE and AMEX Stocks

(1962–1994)

	Market Capitalization	Earnings/ Price	Cash Flow/ Price	Price/ Book	Prior Return
Earnings/Price	−0.10 (−4.14)				
Cash Flow/Price	−0.11 (−4.72)	0.68 (45.00)			
Price/Book	0.32 (17.31)	−0.43 (−25.23)	−0.48 (−24.32)		
Prior Return	0.06 (1.72)	−0.13 (−6.42)	−0.14 (−7.27)	0.16 (7.54)	
Price per Share	0.78 (104.21)	−0.07 (−3.80)	−0.15 (−6.44)	0.34 (19.49)	0.16 (5.01)

*Correlations are computed annually using ranks of individual stocks. All rankings are conducted at the end of March, using prices at the time and accounting numbers for the previous fiscal year. Stocks with negative earnings, cash flows, or book values are excluded. Correlations computed with CF/P are for 1972-1994.

Table 4

Premia Correlations

(April 1962–December 1994)

	E/P	CF/P	P/B	Prior Return
Size	0.265	0.444	0.472	-0.017
E/P		0.727	0.590	-0.230
CF/P			0.760	-0.212
P/B				-0.172

[1] Correlations involving CF/P are for 1972-94.

Table 5

Monthly percentage returns during January and the rest of the year for value-weighted portfolios of NYSE and AMEX stocks formed on the basis of size (market capitalization), earnings-to-price ratios, cash flow-to-price ratios, price-to-book ratios, and prior return[1]
(April 1962 to December 1994)

Portfolio[2]	Size (Market cap.)		Earnings-to-price ratio		Cash flow-to-price ratio[3]		Price-to-book ratio		Prior Return	
	January	Rest of year	January	Rest of year	January	Rest of year	January	Rest of year	January	Rest of year
1	12.11*	0.63	5.37*	0.84	6.28*	1.04	7.23*	0.92	2.82	1.03
2	8.87*	0.75	4.30*	0.98	4.94*	1.00	5.38*	1.07	1.72	1.20
3	7.58*	0.69	3.59*	0.86	4.68*	0.87	4.36*	0.77	1.73	1.04
4	6.80*	0.74	2.73*	0.86	2.91	0.76	3.69*	0.82	2.19	0.92
5	5.74*	0.82	2.60*	0.81	2.65	1.01	3.13*	0.81	2.91*	0.70
6	4.67*	0.81	1.91	0.67	2.74	0.71	2.03	0.68	2.86*	0.73
7	3.77*	0.86	2.21	0.70	2.84	0.97	1.90	0.75	3.31*	0.64
8	3.52*	0.87	1.78	0.81	2.64	0.91	2.58*	0.76	3.29*	0.71
9	3.01*	0.85	1.61	0.82	2.23	0.78	1.91	0.72	3.65*	0.36
10	1.94	0.75	2.28	0.70	1.01	0.78	1.37	0.86	3.37*	0.61

[1]Portfolios are created on March 31 of each year using prices and shares outstanding on March 31 and accounting data for the year ending the previous December 31 (with portfolios containing only December 31 fiscal closers). Prior return is calculated from the beginning of October to the end of February immediately preceding the March formation month. Aside from new listings and delistings, which are added to or dropped from the portfolios as they occur during the year, the portfolio composition remains constant during the following twelve months over which portfolio returns are calculated. The data reported in the table are average values over the sample period. An asterisk indicates that January average returns are significantly larger than the average returns during the rest of the year at the 0.05 level.

[2]Portfolio 1 (portfolio 10) is the portfolio with the smallest (largest) market capitalization, highest (lowest) E/P and CF/P, lowest (highest) P/B, and largest (smallest) 6-month prior returns.

[3]Results for the Cash Flow-to-Price portfolios are for the period April 1972–December 1994.

Table 6
Size adjusted monthly returns (%) and other characteristics for 25 portfolios of NYSE and AMEX stocks ranked first by price-to-book ratio (P/B) and then by prior return

April 1962 to December 1994 (393 observations)

	Prior Return (beginning of October to end of February of preceding year)				
	Lowest	2	3	4	Highest
A. Monthly percentage returns (Standard error)					
Low P/B	0.15 (0.22)	0.35 (0.15)	0.27 (0.15)	0.48 (0.13)	0.44 (0.16)
2	−0.12 (0.15)	0.04 (0.15)	0.01 (0.13)	0.13 (0.12)	0.39 (0.15)
3	−0.12 (0.14)	−0.09 (0.13)	−0.08 (0.10)	−0.10 (0.11)	0.16 (0.14)
4	−0.26 (0.13)	−0.28 (0.12)	−0.08 (0.10)	−0.05 (0.11)	0.32 (0.15)
High P/B	−0.40 (0.14)	−0.09 (0.12)	0.11 (0.11)	0.28 (0.13)	0.02 (0.17)
B. Price-to-book ratio					
Low P/B	0.66	0.69	0.71	0.71	0.70
2	1.08	1.09	1.09	1.09	1.09
3	1.44	1.44	1.44	1.45	1.45
4	2.04	2.05	2.04	2.06	2.07
High P/B	7.36	5.78	5.83	7.17	6.79
C. Six-month prior returns (%)					
Low P/B	−25.66	−8.74	−0.08	8.92	29.47
2	−18.32	−4.16	3.07	10.87	32.16
3	−18.14	−3.19	4.37	12.93	34.24
4	−19.41	−2.68	6.33	16.11	40.58
High P/B	−20.98	−1.01	9.00	19.50	50.40
D. Price ($)					
Low P/B	13.5	16.6	17.9	18.9	17.9
2	21.0	24.9	26.3	26.0	23.4
3	22.5	44.9	30.0	29.4	40.7
4	27.1	32.2	34.5	34.3	32.8
High P/B	30.4	40.8	46.9	48.3	39.3
E. Market capitalization ($ million)					
Low P/B	321	299	360	418	290
2	638	758	888	830	535
3	660	1144	1079	1082	715
4	811	1403	1195	1173	693
High P/B	1075	1847	2210	1947	872

The size-adjusted monthly return for a security is defined as the return for that security minus the monthly portfolio return for the size decile in which the security is a member. P/B and prior return portfolios in the table are value-weighted combinations of these monthly size- adjusted returns. All portfolios are formed on March 31 of each year using year-end accounting values and March 31 market prices. Stocks with negative P/B values are excluded from the sample.

Table 7
Size-adjusted monthly returns (%)
for 25 portfolios of NYSE and AMEX stocks
ranked first by price-to-book ratio (P/B)
and then by prior return

April 1962 to December 1994 (393 observations)

	Prior Return				
	Lowest	2	3	4	Highest
A. January					
Low P/B	3.71	3.13	2.13	2.15	1.34
	(1.32)	(0.69)	0.76	0.53	0.55
2	0.96	1.28	1.41	1.10	0.38
	(0.57)	(0.50)	(0.53)	(0.54)	(0.52)
3	0.35	0.66	0.05	−0.16	−1.07
	(0.63)	(0.71)	(0.36)	(0.44)	(0.52)
4	0.92	−0.36	−0.03	−0.92	−0.83
	(0.53)	(0.56)	(0.40)	(0.47)	(0.74)
High P/B	−0.08	−0.06	−1.03	−1.02	−0.84
	(0.57)	(0.44)	(0.45)	(0.55)	(0.55)
B. February–December					
Low P/B	−0.16	0.10	0.10	0.33	0.36
	(0.20)	(0.15)	(0.15)	(0.13)	(0.16)
2	−0.21	−0.06	−0.11	0.04	0.39
	(0.15)	(0.15)	(0.13)	(0.12)	(0.15)
3	−0.16	−0.16	−0.09	−0.10	0.27
	(0.15)	(0.13)	(0.11)	(0.11)	(0.14)
4	−0.36	−0.28	−0.08	0.03	0.42
	(0.14)	(0.12)	(0.10)	(0.12)	(0.15)
High P/B	−0.43	−0.09	0.21	0.40	0.10
	(0.15)	(0.12)	(0.11)	(0.13)	(0.17)

The size-adjusted monthly return for a security is defined as the return for that security minus the monthly portfolio return for the size decile in which the security is a member. P/B and prior return portfolios in the table are value-weighted combinations of these monthly size-adjusted returns. All portfolios are formed on March 31 of each year using year-end accounting values and March 31 market prices. Stocks with negative P/B values are excluded from the sample.

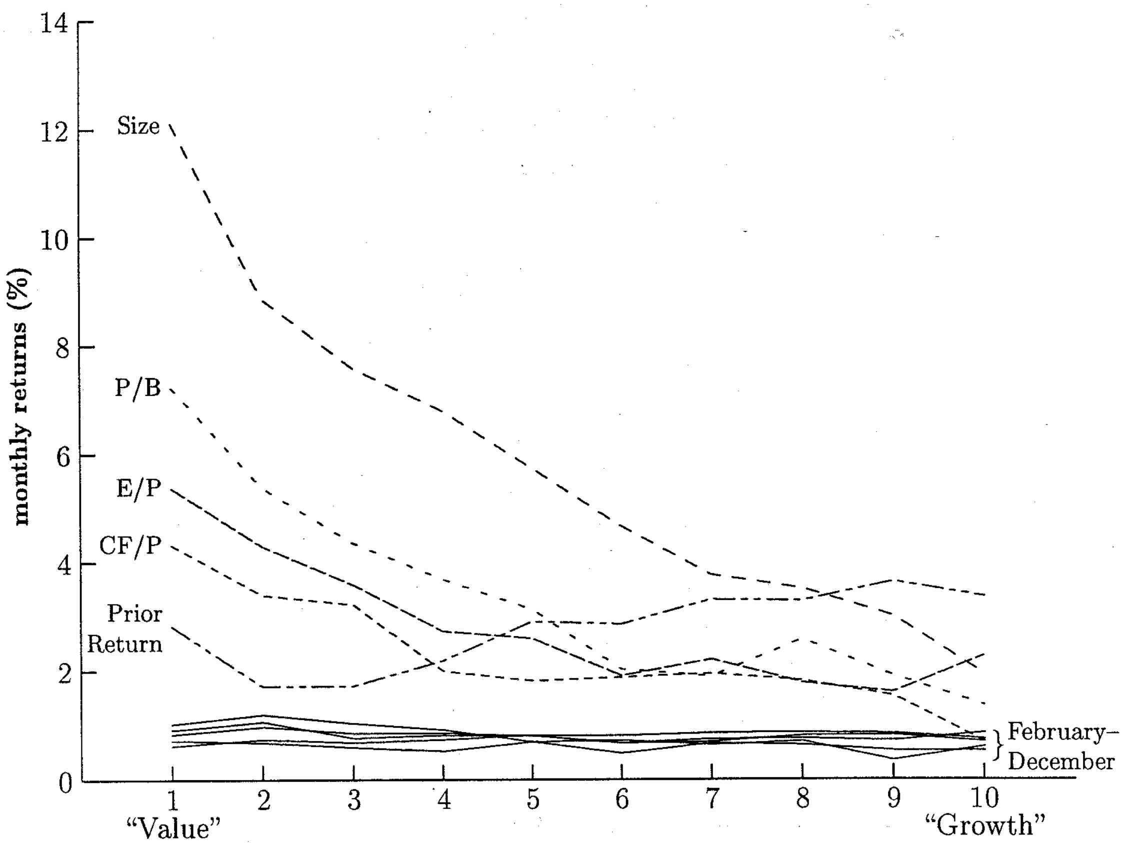
Table 8
Mean Monthly Premia (%) and Correlations Between Premia Across Five International Markets
1975–1994

	Mean Premia	Standard Deviation	Correlations			
			U.K.	Germany	Japan	U.S.
A. Size Effect						
France	0.47	3.66	0.07	0.15	−0.00	0.07
U.K.	0.60	3.26		0.02	0.12	0.24
Germany	−0.01	3.03			0.09	0.07
Japan	0.35	4.11				−0.02
U.S.	0.76	4.62				
B. E/P Effect						
France	0.75	3.17	0.12	0.11	0.07	0.15
U.K.	0.47	2.35		0.02	0.07	0.10
Germany	0.34	2.90			0.02	0.10
Japan	0.49	2.88				0.01
U.S.	0.49	3.33				
C. CF/P Effect						
France	0.77	3.96	0.22	0.24	0.08	0.03
U.K.	0.57	2.89		0.11	0.04	0.14
Germany	0.45	2.71			0.02	0.11
Japan	0.40	4.00				0.10
U.S.	0.50	3.45				
D. P/B Effect						
France	0.80	4.04	0.14	0.13	−0.00	0.16
U.K.	0.57	3.10		0.20	0.17	0.29
Germany	0.19	2.96			0.09	0.25
Japan	0.84	3.30				0.10
U.S.	0.57	3.60				

The premia are computed independently in each market. The premia are defined as the difference in returns between equal-weighted portfolios containing stocks from, respectively, the smallest and largest quartiles of market cap in panel A, the highest and lowest quartiles of E/P in panel B, the highest and lowest quartiles of CF/P in panel C, and the lowest and highest quartiles of P/B in panel D. Quartile cutoffs are determined separately for each market.

Figure 1
The January Effect

Monthly percentage returns, computed separately in January and in February-December, for portfolios constructed on the basis of market cap, E/P, CF/P, P/B, and prior return.
(April 1962 – December 1994)



The portfolio labeled "Value" ("Growth") is the portfolio with the smallest (largest) market capitalization, highest (lowest) E/P and CF/P, lowest (highest) P/B, and largest (smallest) prior return. Stocks with negative values for P/B, E/P or CF/P are excluded from the sample.