ON THE PREDICTABILITY OF COMMON STOCK RETURNS: WORLD-WIDE EVIDENCE

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1. Introduction

In this chapter we examine recent empirical findings which suggest that equity returns are predictable. These findings document persistent cross-sectional and time series patterns in returns that are not predicted by extant theory. As a result, such empirical regularities are often classified as anomalies.

The summary of research presented in this chapter is not an exhaustive compilation of the findings on predictable returns. Rather, we focus on the subset of the findings whose existence has proved most robust with respect to both time and the number of stock markets in which they have been observed. We broadly classify the findings as being cross-sectional (e.g., size and E/P effects) or time series (e.g., return autocorrelations, seasonal return patterns) in nature.

This chapter should not be viewed as an essay on the efficiency of world-wide security markets[1]. Admittedly, the nature of the evidence that we discuss is interpreted by many market observers as convincing evidence of market inefficiency. But we remind the reader of the joint hypothesis underlying such analyses that renders such a conclusion inappropriate. While we acknowledge that the strict form of market efficiency discussed in most textbooks is an unlikely description of security price determination[2], the simple fact that so many of these regularities have persisted for more than fifty years suggests that perhaps our benchmark models are less than complete descriptions of equilibrium price formation.

The chapter will proceed as follows. Section 2 discusses cross-sectional return predictability by focusing on the cross-sectional relation between returns and size, earnings-price ratios, and price-book ratios. Time series return predictability is examined in three separate sections. Section 3 discusses seasonal patterns in returns relating to calendar turning points such as the turn of the year, beginning of the week, and turn of the month. Section 4 covers the autocorrelation of individual security and portfolio returns measured over short and long horizons. Section 5 examines

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[1] See Fama (1991) for a survey of the recent evidence on market efficiency. Also see Lehman (1991) for a more selective treatise on current skepticism regarding the efficient market hypothesis, and Blume and Siegel (1992) for a review of asset pricing and market structure.

[2] In stark contrast to conclusions drawn in his earlier essay on efficient markets in 1970, and conditioned on the accumulation of empirical evidence over the past twenty years, Fama (1991) declares "The market-efficiency hypothesis, that security prices fully reflect all available information, is an extreme null hypothesis, a point on a continuum, and so almost surely false. The interesting task is not to accept or reject market efficiency but to measure the extent to which the behavior of returns departs from its predictions."

the evidence on predicting returns with ex ante observable variables. The paper concludes with a brief summary in section 6.

2. Cross-Sectional Return Predictability

2.1 The Capital Asset Pricing Model

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

$$E(R_i) = a_o + a_1\beta_i \qquad (1)$$

If the model is correct and security markets are efficient, security returns will on *average* conform to this linear relation. Persistent departures, however, represent violations of the joint hypothesis that both the CAPM and the efficient market hypothesis (EMH) are correct.

The strict set of assumptions underlying the CAPM has prompted numerous criticisms. Although any model proposes a simplified view of the world, this does not constitute sufficient basis for its rejection. The rejection or acceptance of a theory should rest on the scientific evidence. Sophisticated tests of the propositions of the CAPM became possible with the creation of the computerized data base of stock prices and distributions at the University of Chicago in the mid 1960's, on the heels of the theoretical development of the CAPM. Numerous studies were conducted in the early 1970's, the most prominent being those conducted by Black, Jensen and Scholes (1972), Blume and Friend (1973), and Fama and MacBeth (1973). These tests found that the estimated intercept was higher than the risk-free rate (the implied value of a_1 in equation (1)), and the estimated coefficient on beta (a_1 in equation (1), representing an estimate of the market risk premium) was lower than predicted by the CAPM of Sharpe (1964), Lintner (1965), Mossin (1965) and Treynor (1961) and only marginally important in explaining cross-sectional differences in average security returns. The results of these studies were interpreted as being consistent with the Black (1972) version of the CAPM.[3]

Although the early tests lend some support for the CAPM, subsequent research was not always as accommodating. For example, in his 1977 critique of existing tests of the CAPM, Roll 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 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.

Since the CAPM was not unambiguously supported by the tests, researchers formulated alternative models. Many developed equilibrium models by relaxing the CAPM assumptions. For example, Mayers (1972) allows for non-marketable assets such as human capital, and Brennan (1970) and Litzenberger and Ramaswamy (1979) relaxed the no-tax assumption. Others examined *ad hoc* alternatives to the CAPM. For example, Basu (1977) and Banz (1981) found that the ratio of price to earnings and the market capitalization of common equity, respectively, provided considerably more explanatory power than beta. Indeed, Banz found little evidence of explanatory power for beta. These two seminal studies served as a springboard for much subsequent research that has confirmed the ability of P/E and size to explain cross-sectional differences in returns. Other studies have extended the list of predictive variables to include industry-relative P/E ratios, the ratio of price to book, price per share, and other similar variables. These studies have produced far more convincing evidence of cross-sectional return predictability than any of the previous tests concerning the explanatory power of beta. Absent in this literature, though, is any supporting theory to justify the choice of variables. Nevertheless, these 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. To maintain a unifying thread through 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. Hopefully, this will avoid some of the apples-and-oranges comparisons imposed on literature surveys of

[3]Many of the tests were done using equal-weighted portfolio (asset) returns and, unknown to the researchers, their evidence was highly sensitive to the correlation of beta and size, and to the relatively high returns of small firms in January. These issues are discussed in the following sections. Also see Ritter and Chopra (1989).

research studies which employ widely varying samples, time periods and empirical methods.

2.2 The Size Effect

Most 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, Banz estimated a model of the form

$$E(R_i) = a_o + a_1\beta_i + a_2S_i \tag{2}$$

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. Similar models have been estimated for Belgium (Hawawini, Michel and Corhay (1989)), Canada (Calvet and Lefoll (1989)), France (Hawawini and Viallet (1987)), Japan (Hawawini (1991), Chan, Hamao and Lakonishok (1991)), Spain (Rubio (1988)), and the United Kingdom (Corhay, Hawawini and Michel (1987)). In all countries except France and Japan there is no relation, on average, between return and market risk when all months of the year are considered (i.e., a_1 is statistically indistinguishable from zero). There is, however, a negative relationship between returns and portfolio size in all countries except Canada and France (i.e., a_2 is significantly less than zero).[4]

Researchers have also demonstrated the existence of the size effect by examining the returns of portfolios formed on the basis of market capitalization. For example, Reinganum (1981), using daily data over the period from 1963 to 1977, showed that portfolios of small firms have significantly higher average returns than large firms. He found that the difference in returns between the smallest and the largest deciles of firms drawn from the NYSE and AMEX is about 30 percent annually. In response to Roll's (1981) conjecture that the size effect may be a statistical artifact of improperly measured risk due to the infrequent trading of small stocks (see also Hawawini (1983)), Reinganum (1982) estimated betas using methods designed to account for non-synchronous and

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[4]There is an abundance of convincing evidence that the relation between size and returns is concentrated in the month of January. We discuss this aspect of the size effect in more detail in section 2.5 below.

infrequent trading (see Scholes and Williams (1977) and Dimson (1979)).[5] He found that the magnitude of the size effect is not very sensitive to the method of estimating betas. Blume and Stambaugh (1983) demonstrate, however, that the portfolio strategy implicit in Reinganum's paper (requiring *daily* rebalancing of the portfolio to equal weights) produces upward-biased estimates of small-firm portfolio returns due to a "bid-ask" bounce that is inversely related to firm size. Blume and Stambaugh show that the measured size-related premium is halved in portfolio strategies that avoid this bias.

The bias described by Blume and Stambaugh (1983) is of substantially smaller quantitative importance in value-weighted portfolio returns and in portfolio strategies employing monthly (or quarterly, or annual) rebalancing. Thus, characterization of the size effect using monthly value-weighted portfolio returns should be reasonably free of the bid-ask bias. With this in mind, Table 1 reports average monthly returns for ten size portfolios of NYSE and AMEX stocks for the period 1951 to 1989. The size portfolios are drawn from data used in Jaffe, Keim and Westerfield (1989) and updated here to 1989.[6] The negative relation between size and expected returns is clearly evident -- the annualized difference in returns between the smallest and the largest size deciles is about 7.9 percent. The betas of the portfolios decline with increasing size, but the differences are small -- the difference between the smallest and largest size portfolio is 0.26. Thus, consistent with research that finds significant coefficients on size in equation (2) 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.

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. The

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[5]Ordinary Least Squares estimates of beta coefficients of infrequently traded stocks are lower than their 'true' beta coefficients, and since small firms tend to trade relatively infrequently, their beta coefficients are underestimated.

[6]See Jaffe, Keim and Westerfield (1989) for a detailed description of the data.

implication is that market structure may be an important influence on the measured size effect. If so, analysis of international evidence on the size effect, where we observe very different market organizations and structures, may be quite useful in understanding the cause of the size effect.[7]

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. The evidence is summarized in Table 2 for the Australian, New Zealand, Canadian, Japanese, Taiwanese, and seven European stock markets. We define the size premium 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. The monthly size premium is positive in all countries.[8] Its magnitude, however, varies across markets: It is most pronounced in Australia (5.73 percent) and Japan (1.20 percent), and is insignificant only in the United Kingdom (0.40 percent) and Canada (0.44 percent per month).[9] Note, however, that there is a wide range across the twelve markets in terms of the size (market capitalization) differential between the largest and the smallest size portfolios. For example, in Spain the average market capitalization of the stocks in the largest size portfolio is 228 times the average market capitalization of the stocks in the smallest size portfolio. But in the case of Taiwan the largest portfolio is only 17 times larger than the smallest one. Because the size and number of

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[7]Loughran (1992) 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.

[8]Contrary to the evidence of a size effect on the Taiwan Stock Exchange reported by Ma and Shaw (1990), Chou and Johnson (1990) find no evidence of a size effect for Taiwan. And according to Kim, Chung and Pyun (1992), there is no evidence of a significant size effect on the Korea Stock Exchange during the period 1980 to 1988 for a sample of up to 224 stocks.

[9] Although Levis (1985) finds that the size effect on the London Stock Exchange 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.

portfolios as well as the sample periods differ across countries, it is difficult to gauge whether the magnitude of the size premium is significantly different across countries.

Can differences in beta risk between the smallest and largest portfolios explain the size premium in these markets? The evidence is found in Table 2. As in the U.S., in Japan small firms have, on average, ***higher*** beta risk than large firms (see Table 1 for the U.S. evidence), but the higher beta risk of small firms in these two countries cannot explain the size premium -- the risk-adjusted size premium is still significantly different from zero. In the remaining countries for which data is available, the systematic risk of the smallest firms is about the same or *lower* than that of the largest firms. As mentioned above, the reason may be that the extreme illiquidity in some of these markets, especially for smaller stocks, may result in downward-biased estimates of beta -- even when betas are estimated with monthly returns. Many of these studies do not estimate betas with the methods of Scholes and Williams (1977) or Dimson (1979) which are designed to correct for this downward bias. In the countries where adjusted betas are computed, the adjusted betas of small firms are indeed higher than their standard OLS betas. But even with adjusted betas the size premium remains.

2.3 The Earnings/Price (E/P) Effect

Earnings-related strategies have a long tradition in the investment community. The most popular of these strategies, buying stocks that sell at low multiples of earnings, can be traced at least to Graham and Dodd (1940, p. 533) 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". 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 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.[10]

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[10]The question posed is whether E/P is the determinant of expected returns or whether E/P is simply a proxy for the underlying factor(s) that are the true determinants of expected returns. Of course an alternative question would ask whether E/P captures differences in equilibrium expected

Nicholson (1960) published the first extensive study of the relation between P/E multiples 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 Table 3A we update the relation between monthly expected returns and P/E to 1989 using the data file of NYSE and AMEX stocks originally constructed by Jaffe, Keim and Westerfield (1989) through 1987, and updated here through 1989[11]. The average portfolio returns reported in Table 3A confirm the E/P effect documented in previous studies. The average difference in returns between the highest and lowest E/P portfolios is on average 0.46 percent per month ($t=1.77$)[12]. For purposes of comparison, we also separately computed size portfolios for the same sample of firms. The average difference in returns between the smallest and largest size deciles for this same 1962-1989 period is 0.80 percent per month ($t=2.42$). Thus, size and E/P display similar abilities to sort firms according to expected returns.[13]

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returns, or does it simple capture misvaluations of individual securites.

[11] See Jaffe, Keim and Westerfield (1989) for a detailed description of the data. Although the sample in Jaffe et al. extends back to 1951, we report results here for the shorter 1962-1989 period to facilitate comparison with other results reported below for this same period. We point out that the sample includes only December 31 fiscal closers.

[12] 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 beta-risk-adjusted returns.

[13]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 the sample of firms described in section 2.3. 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

There are only a few studies which examine the P/E effect in markets outside the United States largely due to a lack of computerized accounting databases which can be used for academic research. In one example, Levis (1989) reports evidence documenting the presence of a significant P/E effect on the London Stock Exchange over the period April 1961 to March 1985. He reports an average monthly premium of 0.58 percent, 7.0 percent annually. This is similar to the U.S. results reported in table 3A, and also to results in Basu (1977,1983) and Reinganum (1981). Adjusting portfolio returns for differences in systematic risk does not modify this conclusion.[14]

Aggarwal, Hiraki and Rao (1988) provide evidence of a significant P/E 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 low P/E stocks outperformed those with relatively higher P/E stocks even after controlling for differences in systematic risk *and* size across portfolios.

For the Taiwan Stock Exchange, 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 Shaw (1990) report a weaker but still significant 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, for the New Zealand Stock Exchange, Gillan (1990) finds no evidence of a P/E effect for the same sample as described in Table 2 for which he reports a significant size effect. Portfolios based on low P/E ratios do not earn significantly higher risk-adjusted returns than portfolios based on high P/E ratios during the period 1977 to 1984. A similar conclusion is reached by Kim, Chung and Pyun (1992) for Korea. They find no evidence of either a size effect or a P/E

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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.

[14]Levis 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 reports a large degree of interdependency between the two effects with the P/E effect tending to subsume the size effect.

effect on that market during the period 1980-1988 for a sample of up to 224 stocks.

In summary, the evidence from five markets outside the United States indicates that in the United Kingdom, Japan and Taiwan there is a significant P/E effect similar to that found in the U.S. 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 Taiwan, New Zealand and Korean markets, however, it is difficult to draw definitive conclusions from the evidence regarding these three markets.

2.4 A Variation on the E/P Effect: the Ratio of Cash-Flow-to-Price

An 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).[15] Chan, Hamao and Lakonishok

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[15]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.

(1991) find evidence of a significant relation between expected returns and cash-flow yield (CF/P) for Japanese stocks.

Consider the evidence on the CF/P effect summarized in Table 3B. The Table 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-flow per share to price per share (CF/P) for the period 1972 to 1989. Cash flow (CF) is defined as reported earnings plus depreciation. The similarity of the average values of E/P and CF/P for the portfolios shows there is much overlap between the composition of the CF/P portfolios in Table 3B and the E/P portfolios reported in Table 3A. There appears to be some marginal explanatory power provided by CF/P relative to E/P in the evidence on average returns reported in the first column, although the results are for two different time periods. To make the comparison for the 1972-1989 period, the average difference in returns between the two extreme E/P decile portfolios is 0.72 percent per month ($t=2.05$) and the difference in returns between the two extreme CF/P portfolios is 0.89 percent per month ($t=2.44$). This translates to an average *annual* difference between the two effects of about 2.0 percent.[16]

2.5 The Price/Book (P/B) Effect

Although less research 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 ability. As is the case for the other variables discussed above, there is no theoretical model that predicts P/B should have predictive power. However, investment analysts (e.g., Graham and Dodd) 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 (1991)) document a significant negative relation

[16]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 indeed 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)). In these markets, low P/S ratio portfolios have, on average, higher returns than portfolios with relatively higher P/S ratios. 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.

between P/B and returns. To provide some perspective on the magnitude of the P/B effect, Table 4 reports average monthly returns and other portfolio characteristics for ten decile portfolios drawn from the same data used above for the E/P and CF/P results. The average monthly returns for the 1962 to 1989 period in Table 4 show a significant negative relation between P/B and returns. Further, the monthly difference in returns between the extreme P/B portfolios (0.78 percent, t=2.35) for the 1972 to 1989 period (not reported in table 4), is comparable to the corresponding differential return for the E/P effect (0.72 percent, t=2.05), CF/P effect (0.89 percent, t=2.44) and larger than the size premium (0.56 percent, t=1.35) measured over the same period.[17]

2.6 One Effect or Many?

The research discussed above documents significant cross-sectional relations between abnormal returns and size, E/P (or CF/P) and P/B. Few would argue that these separate findings are entirely independent phenomena. Several characteristics of the portfolios support such a conclusion: First, the ranking variables all share a common variable -- price per share of the firm's common stock. Second, the various effects all display similar return patterns through time.

2.6.1 Price as a Common Denominator

Market capitalization, E/P, P/B and CF/P 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.[18] Keim (1988), Jaffe, Keim and Westerfield (1989) and Fama and French (1991) have all recently raised this possibility regarding the other effects. That E/P and P/B also produce rankings based, to some extent, on price per share, is evident in the rightmost columns in Tables 3A and 4 where the average share price of the stocks in the size, E/P and P/B portfolios decline monotonically with the respective variables.

To demonstrate the association among these variables, Table 5 reports pairwise rank correlations between each of the variables described above and also for dividend yield, a variable

[17]The international evidence of a P/B effect is limited. One exception is the case of the Tokyo Stock Exchange for which Aggarwal, Rao and Hiraki (1989) report the presence of a P/B effect. See also Chan, Hamao and Lakonishok (1991).

[18]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.

that has also been shown to have cross-sectional return predictability[19]. These rank correlations and the 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, D/P, and price (i.e., five separate rankings are produced). Each variable is computed using price per share at March 31 and, when applicable, accounting numbers for the previous year[20]. Pairwise Spearman rank correlations are computed. This procedure is conducted in each year for the period 1962 to 1989, and mean rank correlations and standard errors are computed for the entire time series of values. Consistent with the above conjecture, the rank correlations are generally large and significant. With the exception of the rank correlations for size-E/P and size-dividend yield, the correlations are significantly different from zero, indicating some commonalities among the effects [21]. As suggested in previous work (Blume and Stambaugh (1983) and Stoll and Whaley (1983)), the rank correlation between market capitalization and price is significant and by far the largest correlation in the table. The correlation between price and P/B is also quite large (0.33).

2.6.2 Common Variation Through Time: Within-Year Seasonality

The fact that all these effects are most pronounced in the month of January suggests that the effects are associated with some common underlying factor. This January seasonal has been demonstrated for the size effect (Keim (1983b)), E/P effect (Cooke and Rozeff (1984) and Jaffe, Keim and Westerfield (1989)), and the P/B effect (Keim (1988) and Fama and French (1991)).[22] There is a burgeoning literature on the January effect that is discussed in more detail below in section 3.2. We merely report on the general stylized facts at this juncture.

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[19] The relation between stock returns and dividend yields is most often attributed to the differential taxation of capital gains and ordinary income as described in the after-tax asset pricing models developed by Brennan (1970) and Litzenberger and Ramaswamy (1979). For additional evidence, both pro and con, regarding the relation between stock returns and dividends, see Black and Scholes (1974), Blume (1980), Keim (1985), Miller and Scholes (1982) and Rosenberg and Marathe (1979).

[20] Note that only December fiscal closers are included in the rankings.

[21] This lack of correlation between size and E/P is also evident in Table 3A.

[22] Interestingly, the P/E effect in the Tokyo market manifests itself in all months of the year *except* January. During that month, there is a significant *reverse* P/E effect -- that is, high P/E portfolios outperform low P/E portfolios during January.

Figure 1 summarizes the January seasonal for the size, E/P and P/B effects using portfolios of NYSE and AMEX securities. Considering each effect separately, the difference in return between the two extreme decile portfolios (e.g., smallest size portfolio minus largest size portfolio) is measured separately in each month of the year and the average for each month is computed. The figure, which reports the month-by-month averages for each of the three effects, clearly shows a January seasonal in the effects.

2.6.3 Common Variation Through Time: Long-Term Variation

In addition to the within-year variation, the magnitude of these effects have been shown to vary over longer periods of time. For example, Brown, Kleidon and Marsh (1983) find that the size effect reverses itself for sustained periods: while for most periods there is a small-firm premium, there are a few periods (e.g., 1969-1973) when there is a *large*-firm premium.[23] Stated differently, there are extended periods when small capitalization portfolios underperform large-capitalization portfolios on a risk-adjusted basis. Jaffe, Keim and Westerfield (1989) report similar findings for the E/P effect. Thus, we must distinguish between the unconditional and conditional expected values for the effects. The relatively short period (1972-1989) used in the comparisons reported in section 2.4 may not be long enough to capture the "long-run" magnitudes of these effects.

To illustrate the time-varying nature of these effects, Figure 2 plots the annual size, E/P and P/B effects over the 1962 to 1989 period. The magnitude of the effects is measured as the respective differences in returns between the extreme decile portfolios composed of NYSE and AMEX securites. We make several observations. First, and consistent with previous findings, the magnitudes of the effects change substantially over time.[24] Hence, the estimated magnitudes of the effects are quite sensitive to the period in which they are measured. For example, there are extended periods when the effects reverse, especially for the size effect (Brown, Kleidon and Marsh (1983)). As a case in point, the size effect was inverted for a large portion of the 1980's, which comprises a long segment of the time period used in the 1972-1989 comparison above. As a result,

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[23]Note, however, that even in periods when there is, on average, a large-firm premium, there is a significant small-firm premium in January (see Keim (1983b)).

[24]Keim and Stambaugh (1986) examine the *conditional* size effect and show that an *ex ante* observable variable (based on the level of small firm prices), chosen to proxy for variation in expected returns, predicts the variation in the size effect. Keim (1989) and Fama and French (1991) note similar capabilities to predict the E/P effect and the P/B effect, respectively.

it is not clear that the relative differences in the magnitudes of the effects noted is section 2.5 are unbiased estimates of the long-run differences in returns arising from portfolio strategies employing these different variables. Second, it also appears that while the annual P/B and E/P effects are significantly related (the simple correlation, r, between the annual P/B and E/P effects is 0.76 and significant), the size effect is less correlated with both the E/P effect (r=0.31) and the P/B effect (r=0.36).

2.7 Disentangling the Effects

The discussion in section 2.6 concerns two stylized facts. First, the variables that have been shown to predict cross-sectional variation in returns are all significantly correlated with price per share, which itself has been shown to significantly predict cross-sectional differences in returns. Second, the ability of these variables to predict cross-sectional differences in returns is significantly greater in January. Indeed, it is very weak outside of January. The question, then, is whether these separate findings are all a manifestation of the same underlying phenomenon; and if so, which is the cleanest proxy for the underlying effect.

2.7.1 Size and E/P

Most of the research in this area has examined the interrelation between the E/P and size effects. 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 (i.e., once we control for size, there is no marginal E/P effect). Basu (1983) argues just the opposite. Peavy and Goodman (1983) and Cook and Rozeff (1984), after performing meticulous replications of and extensions to the methods of Basu and Reinganum, reach surprisingly different conclusions. Peavy and Goodman's results are consistent with Basu's, but Cooke and Rozeff conclude that no one effect dominates the other. Banz and Breen (1988) 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 time periods used in the above studies (ranging from 8 to 18 years) that do not overlap, and the failure of the studies (with the exception of Cook and Rozeff (1984)) to account for potential differences between January and the other months (see Figure 1). Using data covering a 36-year period, Jaffe, Keim and Westerfield 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[25]. They also conclude that the results of the earlier studies conflict because the magnitude of the two effects is period specific (see Figure 2). Fama and French (1991) reach similar conclusions regarding the joint significance of size and E/P effects (see the regression results in their Table 3).

2.7.2 Size and P/B

Stattman (1980) and Rosenberg, Reid and Lanstein (1985) (RRL) were the first to examine the possible interaction between size and P/B for NYSE and AMEX stocks. Stattman examines average beta-risk-adjusted portfolio returns for the 4/1964 to 4/1979 period 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". RRL 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. Fama and French (1991) use data for the longer period of 1962-1990. Their sample includes NYSE and AMEX stocks for the entire period and OTC stocks for the 1973-1990 subperiod. Fama and French estimate an extension of cross-sectional model given in equation (2) with the addition of P/B, using data for individual stocks.[26] Based on their findings, they conclude that size and P/B are sufficient to characterize cross-sectional differences in expected returns.

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[25] 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.

[26] 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 for the test assets in equation (2), to avoid estimation error in individual beta estimates, unduly forfeit the valuable information in the cross section of individual security characteristics such as market values or E/P's. This same point was emphasized earlier in Litzenberger and Ramaswamy (1979) in their analysis of the relation between stock returns and dividend yields.

2.7.3 E/P and P/B

The consensus from the research detailed above is that the relation between market capitalization and expected returns is quite robust.[27] There is strong evidence, though, that other variables, especially E/P and P/B, provide additional explanatory power for expected returns. Fama and French (1991) investigate whether these two variables are proxying for the same additional influence by estimating equation (2) with both P/B and E/P as additional independent variables. Based on their results, they conclude that size and P/B together subsume any additional explanatory power of E/P. It is difficult to compare their results to earlier findings on the E/P effect since Fama and French compute ratios of E/P and P/B using market prices that occur at a point in time before the market knows the value of earnings or book value. Most studies use a market price several months after the end of the firm's fiscal year-end to insure that the accounting information in the ratio is known at the time the price is recorded. Such a "look-ahead bias" in the research design can significantly affect the results (e.g., see Banz and Breen (1986)).[28]

To facilitate comparison of the interaction between the E/P and P/B effects to previous studies of the E/P effect, we use the data described above in section 2.3 to compute returns for portfolios created on the basis of both E/P and P/B, where the market price used in the ratio occurs three months after the end of the firm's fiscal year. Briefly, at the end of March in each year from 1962 to 1989 we sort all NYSE and AMEX stocks by P/B and form five groups of equal numbers of securities based on the P/B ranking. Within each of these groups we again rank the stocks by E/P and create five subgroups within each of the P/B groups.[29] Individual stock returns are adjusted for the influence of the size effect by simply subtracting the return for 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

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[27] 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.

[28]Fama and French do not start measuring returns for their portfolios until six months after the portfolio formation date. Thus, it is not clear how the look-ahead bias will ultimately affect their results. However, there is loss of power in the tests due to the use of "stale" prices.

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

computed each month.

The results, reported in Table 6 seem to tell a story similar to Fama and French. First, there is slight evidence of a relation between E/P and returns, although this relation is largely confined to the lowest P/B category, and extends only through the first four quintiles. Second, there appears to be a relation between P/B and subsequent returns, although this relation is often not monotonic and does not appear in all E/P categories. Curiously, the P/B effect occurs in the E/P categories in which the average price of the securities in the portfolios increases with P/B. (In this regard, recall the significant rank correlations between price and P/B in Table 5.) Although the experiment controls for the influence of market capitalization on expected 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 prospect that low P/B stocks are simply stocks whose prices have dropped relative to book values that vary little through time. As discussed below in section 4.1, 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.[30] As a result, stocks that have recently declined substantially 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* higher levels of risk, the subsequent higher average returns that compensate for this risk appear anomalous.[31]

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[30]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 occuring 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 occuring after the analysis interval -- 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.

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

3. Time Series Return Predictability: Seasonal Patterns

Consider a model of stock prices in which expected returns are constant through time. In the next three sections we entertain evidence relating to the validity of such a model. In this section, we examine the evidence regarding "seasonal" patterns in returns relating to calendar turning points such as the beginning of the week, the month, and the year.

3.1 Patterns in Daily Returns Around Weekends

Consider an exchange where trading takes place Monday through Friday. If the process generating stock returns operates continuously, then the expected return for Monday should be three times the expected return for the other days of the week to compensate for a three-day holding period, given that no trading takes place over the two-day weekend. This is known as the calendar-time hypothesis. The alternative is the trading-time hypothesis according to which returns are generated only during active trading and the expected return is the same for each of the five days of the week. The evidence we discuss in this section finds that stock returns, in many Western countries, are on average negative on Monday. This is inconsistent with both the calendar-time and the trading-time hypotheses since the former predicts a *larger* return, and the latter an *equal* return, for Monday relative to the other days of the week.

Cross (1973) and French (1980) document significant negative Monday returns using the S&P Composite Index beginning in 1953, and Gibbons and Hess (1981) find the same pattern for the Dow Jones Industrial Index of 30 stocks (1962-78). Keim and Stambaugh (1984) extend the findings for the S&P Composite to include the period 1928-1982, and also find the same pattern in actively-traded OTC stocks. Lakonishok and Smidt (1988) extend the finding for the Dow Jones Industrial Index to include the period 1897-1986.

The weekend effect has also been documented in many other stock markets. The international evidence is summarized in Table 7. The research reported in the table covers periods of various lengths, generally in excess of fifteen years. In each case the study reporting the most extensive data is used, and returns are computed using the closing (end-of-trading) value of the index. In particular, Monday returns are computed from Friday close to Monday close and hence include the non-trading weekend period (Friday close to Monday open) as well as Monday's trading hours (Monday open to Monday close). The exchanges in most countries are closed on Saturdays and Sundays. Exceptions are: Japan, where the exchange was open every second, fourth and fifth (if any) Saturday of the month for morning trading (this practice was discontinued in January 1989, see Ziemba (1992)); Korea, where the market is open every Saturday; and the U.S., where the NYSE was open for trading on Saturdays (generally from 10:00 A.M. until noon) during most of the period prior to June 1952. Starred returns are statistically different from zero.

Several observations can be made :

(1) There are significant differences in average daily returns across days of the week in nearly all countries (statistical tests in the corresponding studies confirm this for all the countries in the sample, except for the equal-weighted index in Germany).

(2) In general, average daily returns during the last three days of the week (Wednesday, Thursday and Friday) are positive. In contrast, average daily returns during the first two days of the week (Monday and Tuesday) are often negative. (However, the Greek index displays significantly negative average returns on Wednesday.)

(3) In most non-Asian markets, average returns on the first day of the trading week -- usually Monday -- are significantly negative. Note that in Spain, Tuesday returns are negative, corresponding to Tuesday being the beginning of the trading week for the Madrid market during the period examined.

(4) Tuesday returns are significantly negative for the Pacific rim countries of Japan, Korea, Singapore and Australia, sometimes in conjunction with significantly negative Monday returns (Japan and Korea).

(5) In France and Germany, the negative Monday effect manifests itself in the value-weighed index rather than the equally weighted index which means that it is mostly caused by relatively larger firms (large firms are over-represented in the value-weighted index compared to the equal-weighted index). Thus, the Monday effect may be related to firm size.

Related to this last point, Keim and Stambaugh (1984), noting results in Gibbons and Hess that suggest that Friday returns vary cross-sectionally with market value, find that the return differential between small and large firms increases as the week progresses, and is largest on Friday. In addition, Keim (1986) shows that, controlling for the large average returns in January, the "Friday effect" and the "Monday effect" are not different in January than in the other months.

All of the studies which document negative Monday returns use Friday-close-to-Monday-close returns, and thus cannot distinguish whether the negative returns are due to the weekend non-trading period or to active trading on Monday. Research that has tried to sort out this issue is not in complete aggreement. For example, Harris (1985), for all NYSE stocks for the period 1981-1983, and Smirlock and Starks (1986), for the Dow Jones 30 for the period 1963 to 1973, examine intra-daily returns and show that negative Monday returns accrue both during the nontrading hours over the weekend (Friday close to Monday open) and during trading on Monday. On the other hand, Rogalski (1984) examines intra-daily data for the period 1974-1984 and concludes that negative Monday returns accrue entirely during the weekend non-trading period.

We do not yet have a satisfactory explanation for the weekend effect. The fact that the pattern is robust across so many different markets argues persuasively against many institutionally-motivated explanations. For example, one potential explanation that has been examined in the United States (Gibbons and Hess (1981) and Lakonishok and Levi (1982)) and the United Kingdom (Theobald and Price (1984)) is based on the delay between trading and settlement (actual transfer of funds) due to check clearing. On the NYSE there is a five-business-day settlement period to which an additional day is added for check clearing. This means that for stocks purchased on a business day other than Friday, the buyer will have eight *calendar* days before transferring funds. For stocks purchased on Friday, he will have ten calendar days and thus two more days of interest. In an efficient securities market the buyer should be willing to pay more for stocks purchased on Friday by an amount not exceeding two days of interest. Consequently, observed returns on Friday should be higher than those on other days of the week and those of Monday should be lower. Gibbons and Hess and Lakonishok and Levi showed that adjusting daily returns for the interest rate reduces the Monday effect, but comes nowhere close to explaining it.

On the London exchange trading takes place over consecutive account periods of two weeks' length beginning every other Monday. Settlement, however, is made on the second Monday after the end of the account. This means that for stocks purchased on the first Monday of the account, the buyer will have 21 calendar days before losing funds; whereas for stocks purchased the preceding Friday, he will have only 10 calendar days. For stocks purchased on the second Monday of the account, the buyer will have 14 calendar days but 17 calendar days for stocks purchased the preceding Friday. According to the settlement-delay hypothesis, the first Monday returns for the account should be higher than the returns on the other days of the week and the second Monday returns should be smaller. However, the fact that stocks generally go ex-dividend on the *first* Monday of the account (Theobald and Price (1984)) will partly offset the rise in price and also the return predicted for that day by the settlement-delay hypothesis. Indeed, returns on non-ex-dividend Mondays are generally negative, whereas returns on ex-dividend Mondays are generally positive.[32] This result is qualitatively consistent with the settlement-delay hypothesis but the magnitude of the Monday effect on the NYSE and the LSE cannot be fully explained by the settlement-delay hypothesis.

The Tuesday effect in Australia, Korea, Japan and Singapore is conjectured as resulting from time zone differences relative to New York (Jaffe and Westerfield (1985a, 1985b)). These countries are all one day ahead of New York; hence, the Tuesday effect in these countries may reflect the earlier Monday effect in New York.

We also know that the weekend effect is not completely explained by measurement error in recorded prices (Gibbons and Hess (1981), Keim and Stambaugh (1984), Smirlock and Starks (1984)), specialist trading activity (Keim and Stambaugh (1984)), and systematic patterns in investor buying and selling behavior (Keim (1989) and Lakonishok and Maberly (1990)).

What are the implications of the above findings for investment managers? The evidence indicates that the major stock exchanges around the world are part of a global market in which the price movements of individual exchanges are interrelated. Lags in the correlation structure of stock returns around the world seem to reflect differences in time zones. The Monday effect reported in North America and the United Kingdom is replaced by a Tuesday effect in many countries outside the New York time zone. There may be some advantages to be gained from that knowledge: investors *planning* to buy stocks should do so preferably on Monday for Canadian, U.S. and U.K. stocks and on Tuesday for most Far Eastern stocks. Similarly, a stock sale should be preferably carried out on a Wednesday, Thursday or Friday.

3.2 Patterns in Returns Around the Turn of the Year

Rozeff and Kinney (1977) found that indexes containing all the stocks listed on the NYSE displayed significantly higher returns in January than in the other eleven months over the period 1904-1974. Gultekin and Gultekin (1983) examined the monthly stock returns from 17 countries

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[32] In smaller markets, daily returns, even for frequently traded stocks, are usually correlated. This phenomenon may prevent the detection of a Monday effect in the data. An additional problem may arise from the fact that the returns of market indices containing stocks with low trading frequency are generally serially correlated even if no serial correlation exists in the individual underlying securities returns (Hawawini (1978, 1980 a,b)). Theobald and Price (1984) have shown that in this case seasonal patterns in the daily returns of the individual securities that make up the index will be "diffused". Indeed, they report a stronger mean seasonality in more regularly "traded" indices. Another characteristic of thinly traded stocks which has to be taken into account is the significant deviation of their return distribution from normality.

during the period January 1959 through December 1979 and found that all the countries in their sample exhibited a large and positive January mean return. Average January returns were significantly larger than returns in other months for 13 of the 17 countries analyzed.

Table 8 summarizes the turn-of-the-year effect by reporting average monthly returns of eighteen market indices by month of the year. Several observations can be made. First, there is considerable variation in average returns across months of the year. In particular, average returns during January are always positive and generally significantly higher than during the rest of the year (a notable exception is Korea). Second, the magnitude of the January seasonal depends on the composition of the stock market index. Broader and equally-weighted indices, which emphasize smaller stocks, exhibit a stronger January seasonal than narrower or value-weighted indices. The equal-weighted index of all NYSE shares has an average January return of 5.08 percent compared to 1.04 percent for a value-weighted index. The same is true in Japan where the index reported in table 8 is equally-weighted and exhibits an average January return of 5.58 percent compared to 4.36 percent for a value-weighted index containing the same stocks and measured over the same period (see Hawawini (1991)).

Keim (1983b) documented that the magnitude of the size effect varied by month of the year. He found that fifty percent of the annual size premium was concentrated in the month of January. Subsequent research by Blume and Stambaugh (1983) demonstrated that, after correcting for an upward bias in average returns for small stocks that was common to the experimental design in of the early studies on the size effect, the size effect is evident only in January. The worldwide evidence of monthly seasonality in the size premium is summarized in Table 9 for the subsample of the countries presented in Table 8 for which evidence is available. The size premium is measured using the same method as the one reported in Table 2 but instead of taking all months of the year into consideration, the size premium is first measured during the month of January and then during the rest of the year (from February through December). In all countries except France and the United Kingdom the size premium is significantly larger during January than during the rest of the year, although it generally remains positive after January.

What explains this phenomenon? The most popular hypothesis attributes the effect to year-end tax-loss selling. The tax-loss selling hypothesis can be summarized as follows:

> The hypothesis maintains that tax laws influence investors' portfolio decisions by encouraging the sale of securities that have experienced recent price declines so that the (short term) capital loss can be offset against taxable income. Small firm stocks are likely candidates for tax-loss selling since these stocks typically have higher variances of price

changes and, therefore, larger probabilities of large price declines. Importantly, the tax-loss argument relies on the assumption that investors wait until the tax year-end to sell their common stock 'losers.' For example, in the U.S., a combination of liquidity requirements and eagerness to realize capital losses before the new tax year may dictate sale of such securities at year-end. The heavy selling pressure during this period supposedly depresses the prices of small firm stocks. After the tax year-end, the price pressure disappears and prices rebound to equilibrium levels. Hence small firm stocks display large returns in the beginning of the new tax year. (Brown, Keim, Kleidon and Marsh (1983), p.107)

Reinganum (1983) and Roll (1983) both examine the hypothesis and their tests suggest that part, but not all, of the abnormal returns in January is related to tax-related trading. On the other hand, Schultz (1985) finds that prior to 1917 -- before the U.S. tax code created incentives for tax-motivated selling -- there was no evidence of a January effect for his sample of low-price stocks. This finding is confirmed in a recent paper by Jones, Lee and Apenbrink (1991) who find no evidence of a January effect prior to 1917 for the individual stocks in the Cowles Industrial Index.

Others have tested the hypothesis by examining the month-to-month behavior of returns in countries with tax codes similar to the U.S. code but with different tax year-ends. The tax-loss selling hypothesis predicts that, in the month immediately following the tax year-end, returns of small firms will be large relative to both other months and other firms. The hypothesis makes no predictions regarding the behavior of returns during the other months. The international evidence is far from conclusive. The studies that examine returns in countries with similar tax codes to the U.S. but with different tax year-ends (for example, the U.K. has an April tax year-end and Australia has a June tax year-end) find seasonals after the tax year-end, but often find large returns in January that are not predicted by the hypothesis. There is no January size premium in France, a country that taxes capital gains, while a January size-premium is reported in Japan, a country that did *not* tax capital gains for individual investors until 1989. Of course, the January size-premium in countries that neither tax capital gains nor have a December 31 tax year-end may be induced by foreign investors who pay capital gains taxes in their home country. Some evidence in favor of the tax-induced hypothesis is that no seasonality was detected in the United Kingdom *prior* to the introduction of capital gains taxes in this country in April 1965 (Reinganum and Shapiro (1987)). After April 1965, seasonality appeared both in January and April.[33] Further, Berges, McConnell

[33]It is worth noting that in the United Kingdom, the size premium has been shown to be the largest during the month of May when it is equal to 2.45 percent (1.29 percent for the smallest portfolio minus -1.16 percent for the largest portfolio). This represents an annualized return of almost 30 percent. Interestingly, there is an old British maxim that says "sell in May and go away".

and Schlarbaum (1984) find a January seasonal in Canadian stock returns prior to 1972, a period when Canada had no taxes on capital gains.

The inconsistent evidence regarding the tax-loss selling hypothesis has led to other potential explanations. One possibility concerns the impact of institutional 'window dressing' at the end of the year -- the practice of selling off 'loser' stocks at year-end so that they don't appear on the year-end statements sent to constituent shareholders (Haugen and Lakonishok (1987))[34]. Although there is some evidence that institutions behave in this fashion, the resulting impact on stock prices is difficult to distinguish from the impact of tax-loss selling. Similar in spirit to the 'price pressure' necessary in the tax-loss-selling and window-dressing stories is the notion that liquidity constraints of market participants may influence security returns, and these effects may have seasonal patterns. For example, periodic infusions of cash into the market as a result of, say, institutional transfers for pension accounts or proceeds from bonuses or profit-sharing plans, may impact the market. For example, Kato and Schallheim (1985) and Hawawini (1991), in an examination of the size effect in Japan, find January *and* June seasonals in small firm returns that coincide with traditional Japanese bonuses paid at the end of December and in June. Further, Rozeff (1986) finds a substantial *upward* shift in the ratio of sales to purchases of common stocks by investors (who are not members of the NYSE) at the turn of the year that coincides with the small firm returns in January (although Rozeff interprets this as evidence of tax-loss selling). Ritter (1988) documents a similar pattern in the daily sales to purchases ratio for retail customers of a large brokerage firm, and argues, in conjunction with buying pressure at the beginning of the new year, that the turn-of-the-year price behavior is a result of price pressure. Finally, Ariel (1987) finds a daily pattern in daily stock returns in *every month* but February that parallels precisely the pattern that occurs at the turn of the year (see section 3.3). Such a consistent pattern at month-ends is likely related to investor buying and selling behavior that is motivated by reasons other than taxes.

The above arguments rely on the notions of price pressure and irrational market participants to translate the buying and selling behavior of market participants into abnormal returns. There is an alternative. Most (all) of the above studies use daily *closing* transaction prices to compute returns for the analyses. These closing prices may be equivalent to dealer (specialist) bid or ask prices depending on whether the trade was seller- or buyer-initiated. Now, as a case in point, if

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Obviously the maxim applies to large rather than small firms.

[34]See also the related evidence on market "overreaction" below in section 4.1.

there is a preponderance of seller-initiated transactions at the end of the year (because of tax-loss selling, perhaps) and an abundance of buyer-initiated transactions in the beginning of the new year (as investors rebalance their portfolios with substitute securities for the losers they sold in December), and this behavior tended to be particularly pronounced for small stocks (because the losers tend to be heavily populated in that sample), then the large returns we tend to see at the turn of the year could be due to a systematic movement from transactions occurring closer to the bid toward transactions that are closer to the ask. In other words, the turn-of-the-year effect for small stocks may be nothing more than a reflection of the bid-ask spread for these stocks -- an artifact of the data. Keim (1989) investigates and confirms this conjecture for OTC National Market System stocks for the 1983-1988 period, and for NYSE and AMEX stocks for the 1988-89 turn-of-the-year period. However, even after adjusting for this 'bias' in returns, Keim still finds evidence of a significant turn-of-the-year effect for the sample period. That is, the prices on small stocks at the end of January are substantially higher than at the end of December.

3.3 Patterns in Returns around the Turn of the Month

The monthly effect was found by Ariel (1987) who showed that for the period 1963 to 1981 the average returns for common stocks on the NYSE and AMEX are positive only for the last trading day of the month and for the trading days during the first half of the month. This statement is true even if the large returns around the turn of the year are removed from the sample. During the latter half of the month returns are indistinguishable from zero. Ariel concludes that during his sample period "all of the market's cumulative advance occurred around the first half of the month, the second half contributing nothing to the cumulative increase." Results in Lakonishok and Smidt (1988) for the 30 stocks in the Dow Jones Index for the period extending back to 1897 tend to confirm Ariel's finding. Although Ariel is unable to explain the finding, one potential explanation involves portfolio rebalancing of individual and institutional investors related to cash infusions from month-end salaries and contributions (see Ritter (1988) and Ritter and Chopra (1989) for a discussion of these effects at the turn of the year). Results from Japan lend some credence to this story. Ziemba (1991) shows that the index of the 225 stocks listed on the first section of the Tokyo Stock Exchange have significantly higher average returns between day -5 (zero being the beginning of the month) and day +2. The turn-of-the-month effect may begin earlier in Japan because most salaries in Japan are paid on days 20 to 25 of the month with the 25th being the most popular. In the United States most salaries tend to be paid on day -1.

3.4 Patterns in Returns Around Holidays

Ariel (1990) finds that over one-third of the return accruing to the index of all stocks on the NYSE and AMEX over the period from 1963 to 1982 was earned on the trading days preceding the eight holidays that result in market closings each year.[35] The finding holds even if New Years day is excluded. Lakonishok and Smidt (1988) confirm Ariel's findings back to 1897 for the 30 stocks in the Dow Jones Industrial Average. For the Japanese index described in section 3.3, Ziemba (1991) finds that the average return is +0.31 percent (t=3.10) before holidays, significantly larger than the average return on the other days. There is a strong holiday effect in Japan on April 28 (with an average daily return of 0.22 percent); May 2 (0.54 percent) and May 4 (0.33 percent). These are the three days preceding the three holidays occuring during the so-called Golden Week on April 29, May 3 and May 5.

4. Time Series Return Predictability: Return Autocorrelations

Most early work studying the efficiency of the stock market presumed an equilibrium model of stock prices in which expected returns are constant through time (prices follow a "random walk"). The constant expected returns model implies that if the market is efficient price changes are unpredictable so that the past history of price changes is not informative for the prediction of future price changes. Researchers have examined this notion of market efficiency by testing whether return autocorrelations are equal to zero. We examine the evidence on stock return autocorrelations first for individual securities and then for portfolios.

4.1 Individual Security Return Autocorrelations

In an early examination of return autocorrelations, Fama (1965) examined the autocorrelation of *daily* returns for the individual Dow Jones 30 industrial stocks. He found that 75 percent of the Dow 30 stocks had significantly positive autocorrelations in the 1957-1962 period. Foerster and Keim (1992) update these results for the 1963 to 1990 period and find that 80 percent are significantly positive. Although the Dow 30 is a limited sample of relatively homogeneous stocks, they are nevertheless an interesting sample because (1) they represent the stocks most widely followed by analysts and other market observers, (2) they are among the most actively traded of all

[35] These holidays are New Years, President's Day, Good Friday, Memorial Day, Independence Day, Labor Day, Thanksgiving, and Christmas.

stocks, and, as a result, (3) they have relatively tight bid-ask spreads. Since the Dow 30 stocks also have relatively high prices, they are less subject to potential bid-ask-related biases that might influence estimated returns and, therefore, autocorrelations.[36] Based on the evidence for these very liquid, actively-researched stocks, we are left with the impression of an underlying positive serial dependency in stock returns.

French and Roll (1986) compute autocorrelations for all NYSE and AMEX stocks and find that the *daily* autocorrelations are on average negative for exchange-traded stocks. They find, however, that the estimated autocorrelations are inversely related to the market capitalization of the stock: smallest stock autocorrelations are the most negative, and the stocks in the largest decile of market capitalization have positive autocorrelations on average. This latter finding confirms the results above for the Dow 30, while the former result may be a reflection of the influence of a bid-ask bounce discussed in footnote 36.[37]

In most of these studies, the predictable component of returns explains a trivial percentage of total return variability (typically less than 1 percent). Lo and MacKinlay (1988), who find that *weekly* returns are negatively autocorrelated across all firm sizes, also conclude that the serial correlation is both economically and statistically insignificant.[38] They concur with previous studies that the idiosyncratic variability of individual stocks overwhelms the predictable component.

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[36] In addition to the inferential problems associated with the joint nature of such tests, there are biases and other microstructure-related effects that may induce "artificial" serial dependencies into returns. For example, Niederhoffer and Osborne (1966) find that successive trades tend to occur alternatively at the bid and then the ask, resulting in negative serial correlation in returns at short frequencies. Indeed, it is this bid-ask bounce that motivated Roll (1984) and others to exploit this negative serial dependency to estimate the (unobservable) effective bid-ask spread. Thus, this negative serial dependency at short frequencies may mask the true underlying process governing price behavior. Such effects will have a larger impact on estimated returns and autocorrelations for smaller stocks which have lower prices and, consequently, for which the bid-ask spread represents a larger percentage of price.

[37] Solnick (1973) finds significant negative autocorrelation of daily returns for the majority of stocks trading on exchanges in Belgium, France, Germany, Italy, Netherlands and the United Kingdom. Lawrence (1986) reports significant negative daily autocorrelations for 31 percent of the stocks trading on the Kuala Lumpur exchange and 79 percent of the stocks trading on the Singapore exchange. Butler and Malaikah (1992) report that 36 percent and all of the stocks trading on the Kuwaiti and Saudi stock markets, respectively, exhibit significant negative daily autocorrelations.

[38] Cootner (1964) earlier found evidence of negative autocorrelation for weekly stock price changes.

Lehman (1990) directly tests the economic significance of the negative weekly autocorrelations by examining the profitability of a trading rule based on this price pattern. Based on his evidence he concludes that the trading rule is profitable for certain market participants, but his assumption regarding the magnitude of the trading costs necessary to implement the strategy may not fully account for the costs relevant for most market participants. In addition, Conrad, Gultekin and Kaul (1991) replicate Lehman's experiment with closing bid quotes rather than with transaction prices and find that the profits are indistinguishable from zero.

Blume and Friend (1978, p.170-71), Keim (1983a) and Jagadeesh (1990) find evidence of significant negative autocorrelation using *monthly* returns. In contrast to the above studies that estimate serial correlations, however, these studies estimate cross-sectional regressions like (2) using individual securities, but with the lagged returns as independent variables. Keim and Jagedeesh also control for the size effect when estimating the coefficients. The return "reversals" they document are, therefore, unlikely to be due to biases related to bid-ask spreads.

The research on return reversals that has had the largest impact on both academics and practitioners is the work of De Bondt and Thaler (1985,1987). De Bondt and Thaler examine longer-horizon returns and find that NYSE stocks identified as the biggest losers (winners) over a 3- to 5-year period have the highest (lowest) market-adjusted returns on average over the following period. De Bondt and Thaler attribute the predictability to market "overreaction" in which stock prices diverge from fundamental value because of (irrational) waves of optimism or pessimism before returning eventually to fundamental values.[39] These findings have generated many contrarian investment strategies whose profitability is predicated on such negatively autocorrelated price changes.

There may be no free lunch, though. Chan (1988) and Ball and Kothari (1989) argue that the abnormal risk-adjusted returns reported for contrarian investment strategies are due to inadequate adjustment for risk. That is, a loser firm whose stock price (and therefore market capitalization) 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 risk may underestimate the priced risk for these firms and, therefore, overstate the abnormal return. In this vein, Zarowin (1990) 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

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[39] See also Poterba and Summers (1988) and DeLong, Schleifer, Summers and Waldman (1990).

contrarian profits. In contrast to these 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, et al. 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 the findings suggests that the buying and selling behavior of individual investors may be important. Clearly, more work is needed to sort out these issues.

4.2 Portfolio Return Autocorrelations

Because of variance reduction obtained from diversification, portfolio returns provide more powerful tests of the ability of past returns to predict future returns. However, this increased power may be offset by biases and induced autocorrelation caused by the non-trading of securities contained in the portfolios. Fisher (1966) shows that infrequent trading will result in significant portfolio autocorrelation, which will be more serious for portfolios that contain less-frequently traded stocks (see also Hawawini (1978, 1980b)). Scholes and Williams (1977), Dimson (1977) and Cohen, Hawawini, Maier and Schwartz (1983) develop models of the return process in which OLS estimates of beta computed with returns measured over a short (such as daily) interval will contain a bias related to the degree of nontrading of the stocks in the index.

The extent to which such effects surface in the data became evident in the early 1980's with the accumulation of evidence on the size effect. Reinganum (1981), Roll (1981), Keim (1983b) and others report significant positive autocorrelations (in the vicinity of 0.4) for the daily returns of portfolios composed of small, infrequently-traded stocks. It is plausible that such autocorrelation may merely reflect information being revealed in prices of infrequently-traded shares with a lag due to their infrequent trading, thereby imparting serial correlation in the returns of portfolio that contain small, infrequently-traded shares. However, Lo and MacKinlay (1990a) develop a model of nontrading and conclude that, given plausible levels of nontrading, the model cannot explain the level of autocorrelation found in the data. In addition, recent evidence in Foerster and Keim (1992) on actual levels of nontrading finds that periods when nontrading is highest do not correspond to periods during which daily autocorrelations are largest.

There is also evidence of significant and mostly postive autocorrelations in the daily returns of stock market indexes around the world. For example, in the United Kingdom, the Financial Times All Share Index (a value-weighted index) exhibited significantly positive lag-one autocorrelation in daily returns (0.19) but insignificant positive lag-one monthly autocorrelations (0.13) during the period 1965-1989 (Poon and Taylor (1992)). Baily, Stulz and Yen (1990) provide evidence of lag-one autocorrelations for indexes of several Pacific-Basin stock markets during the period 1977 to 1985: 0.21 in Australia; 0.06 in Hong Kong; 0,21 in Singapore; 0.30 in Thailand; -0.08 in Taiwan; and -0.09 in Korea (all significantly different from zero). In the case of Japan, Kishimoto (1990) reports a significant daily autocorrelation for the Topix Index (a composite index of all stocks traded on the Tokyo Stock Exchange) for the period 1949-1988.

Recent evidence indicates that daily index autocorrelations have declined during the 1980's. Froot and Perold (1990) and Foerster and Keim (1992) report that that daily autocorrelations of returns of portfolios of both large and small stocks are insignificantly different from zero in the last half of the 1980's. In fact, for several short subperiods daily autocorrelations were negative. Froot-Perold and Foerster-Keim suggest that the introduction of index futures contracts and the initiation of new institutional trading practices (program trading, the proliferation of index portfolio management) during this period led to higher correlation of closing prices across stocks (due to increased trading in baskets of stocks) resulting in *portfolio bid-ask bounce*. This induced negative autocorrelation may offset any underlying autoregressive component in returns, resulting in lower (negative) daily autocorrelations. Brenner, Subrahmanyam and Uno (1990) and Kishimoto (1990) find a similar decline in the daily autocorrelations of the Nikkei Stock Average and the Topix index after September 1988 when futures contracts on these indexes were introduced.

Because of the potential statistical problems associated with measuring autocorrelations of daily portfolio returns, many researchers shifted their focus to longer-interval returns. Since even the smallest firms on the NYSE or AMEX rarely go more than several days without trading, such biases will unlikely contaminate returns measured over intervals of a week or longer. Lo and MacKinlay (1988) examine weekly returns and find evidence of positive autocorrelation that is strongest for the portfolio of smallest stocks (0.42) and weakest for the portfolio of largest stocks (0.14). To determine whether the autocorrelation results for weekly returns are influenced by non-trading biases, Conrad and Kaul (1988) compute portfolio autocorrelations using weekly returns that were computed only with prices that were the result of actual transactions (i.e., stocks that didn't trade were excluded). Their results are very similar to Lo and MacKinlay, suggesting that significant positive weekly autocorrelations are not attributable to non-trading. That the relation between autocorrelations and the size of the firms in the portfolio is not due to nonsynchronous or infrequent trading is reinforced by the results for monthly returns in Keim and Stambaugh (1986).

They find the same relation between market capitalization and autocorrelations for monthly portfolio returns, although it is weaker than the evidence for shorter-interval returns. Over the 1928-1978 period the return autocorrelation for the portfolio comprising the smallest (largest) quintile of market capitalization was 0.17 (0.13).

Cohen, et al. (1983) and Lo and MacKinlay (1990) develop models of the return-generating process in which positive autocorrelation in portfolio returns is in large part due to significant cross-autocorrelations between securities exhibiting differing degrees of non-trading. Like Roll (1981), Reinganum (1982) and other early research on the size effect that finds lagged values of the return on the market predict subsequent returns on small firm portfolio returns, Lo and MacKinlay find that returns of large-stock (very liquid) portfolios lead the returns of small-stock (less liquid) portfolios. The important contribution of their paper though is the reconciliation of the seemingly paradoxical coexistence of positive portfolio autocorrelations and negative individual stock autocorrelations. They demonstrate that portfolio returns may be significantly autocorrelated even if individual security return autocorrelations are negative (or zero): since the autocorrelation of portfolio returns is the sum of individual security autocovariances and cross-autocovariances, if the cross-autocovariances are sufficiently large relative to the autocovariances, they will overshadow the contribution of the autocovariances (see also Hawawini (1978, 1980b)).

The evidence suggests that, in contrast to individual security autocorrelations, short-horizon portfolio return autocorrelations are statistically significant. However, the estimated autocorrelations are small enough (for example, at the monthly level) that they may not be economically significant. Is the evidence sufficient to reject the joint hypothesis of market efficiency and the model of constant expected returns? If we reject, we are again faced with the standard inferential dilemma: do we interpret the rejection as a rejection of the model in favor of an alternative of time-varying expected returns? or as a violation of the efficient market hypothesis? However, Summers (1986) argues that such tests lack power to reject the null against interesting (controversial) alternative models. For example, Summers (1986) presents a model featuring irrational behavior of market participants where security prices (in aggregate) take long and wide swings from fundamental values, yet this irrationality will not be detectable in tests using short horizon returns (as described above). Stambaugh (1986) concurs with Summers that a uniformly powerful test exists only in the dreams of statisticians. More importantly, though, Stambaugh points out that such security price behavior would manifest itself in significant negative autocorrelations at longer return horizons and he provides some supporting evidence of negative first-order autocorrelations (-0.31) for 5-year S&P 500 returns.

A number of subsequent papers examined the autocorrelation of long-horizon portfolio returns in more detail. For example, Fama and French (1988) find a U-shaped pattern of autocorrelations as a function of return horizon extending out to 10 years. The largest departures from zero are the estimated autocorrelations of -0.25 to -0.45 for 3- to 5-year returns. However, these strong negative autocorrelations are largely due to the 1926-1936 period which is centered on the stock market crash of 1929. Poterba and Summers (1988) also find evidence of negative autocorrelation in long horizon returns for the longer 1871-1985 period for the annual S&P Composite index, backdated to 1871 using the Cowles data as reported in Wilson and Jones (1987). However, Richardson (1991) shows that data generated by a random walk will naturally induce the type of U-shaped pattern in the autocorrelation observed in the stock price data.

In the end, although the autocorrelation tests for both short- and long-horizon returns are suggestive of time variation in expected returns, the statistical shortcomings of the tests prevent clean inferences. Small sample sizes impair the power of the tests using long horizon returns. In the tests using short-horizon returns, lagged returns have marginal explanatory power. However, there is a more fundamental problem: the variation in expected returns that we try to predict represent only a small component of the total variation in returns. As Fama (1991) points out, past returns do indeed contain information about expected returns, but they are a very noisy signal. A more powerful test should exploit explanatory variables that contain more precise information about expected returns. We turn now to such tests.

5. Time Series Return Predictability: Forecasting with ex ante Observable Variables

There is a rapidly-growing body of research that has documented the ability of predetermined variables to predict returns. In the earliest work along these lines, researchers (e.g., Bodie (1976), Jaffe and Mandelker (1976), Nelson (1976) and Fama and Schwert (1977)) uncovered a negative relation between short-horizon (monthly) stock returns and expected inflation during the post-1953 period.[40] However, the explanatory power of expected inflation does not generalize to other assets, and even for stocks it is quite low (less than 3 percent).

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[40] Fama and Schwert (1977) use the T-Bill yield as a measure of expected inflation and find a particularly strong correlation with stock returns in the post-1953 period. This correlation may be period-specific, however. Keim and Stambaugh (1986) show that this relation between stock returns and T-Bill yields is insignificant in the 1926-1952 period (see their footnote 2).

More recent research has focused on variables designed to proxy for the expected risk premium in the stock and bond markets using variables related to the current level of asset prices. The idea is that since asset prices are the discounted values of expected future cash flows, where the discount rate is the expected risk premium, changes in asset price levels convey information about the market's expectations about changing expected risk premiums. Keim and Stambaugh (1986) and Campbell (1987) use such variables to forecast monthly stock and bond returns. For example, Keim and Stambaugh (1986) use a variable that captures the level of small stock prices and a January dummy variable to forecast the returns on long-term government bonds, low- and high-grade corporate bonds and portfolios of NYSE stocks corresponding to quintiles based on market capitalization. They find that the combination of these variables reliably predicts returns across all the asset categories, explaining as much as 14 percent of the variation in returns in the 1953-1978 period. Keim and Stambaugh also point out that the coefficient on the price level variable is larger for corporate bonds than government bonds, for low-grade bonds than high-grade bonds, for stocks than bonds, and for small stocks than large stocks. This accords with intuition about generally increasing risk across these asset categories. Broadening the set of assets under consideration, Harvey (1991) finds that U.S. dividend yield and term structure variables predict monthly returns on a wide array of foreign common stock portfolios. Campbell and Hamao (1991) find similar evidence for Japanese and U.K. stocks. The common variation in expected returns across the asset categories and the ordering of the coefficients argues more for a time-varying expected returns explanation of the results than an inefficient market explanation.[41]

Other research has examined longer-horizon returns. For example, Rozeff (1984) and Shiller (1984) investigate the explanatory power of dividend yields on annual stock returns. Rozeff finds that dividend yields explain 14 percent of the variation in the S&P composite index over the 1926-1981 period. Shiller also examines the predictability of annual S&P Composite returns and finds that dividend yields explain nearly 16 percent of the variation in the 1946-1983 period. Shiller also finds predictive ability of ratios of earnings to price in the 1946-1983 period. He reports an R-squared of .106 for this period. Shiller reports, however, that both dividend yields and earnings yields have very little predictive power in the 1898-1945 period. Additional research finds that the predictive power of dividend yields and earnings yields increases with the length of the return

[41] Fama and French (1989) find similar results for long-horizon returns across similar asset categories and draw similar inferences.

horizon (Campbell and Shiller (1988) and Fama and French (1988b)). For example, Fama and French (1988b) report that dividend yields explain about 25 percent of the variation in 2- to 4-year returns.

6. Concluding Remarks

Research in finance over the past ten to fifteen years has revealed stock price behavior that is inconsistent with the predictions of familiar models. Some of the evidence argues reasonably convincingly for alternatives to existing paradigms -- in particular the evidence on cross-sectional predictability based on variables like size, E/P and P/B. Indeed, many authors argue that the findings provide support for multi-factor alternatives to the CAPM like the models of Ross (1976) or Merton (1973). One of the largest contributions of this entire line of research is that it has perhaps sharpened our focus on relevant alternative sources of risk, and future theoretical work should certainly benefit. On the other hand, there are at least two reasons why it is difficult to argue that the evidence constitutes proof that the CAPM is 'wrong.' First, there is Roll's conjecture that such evidence cannot be considered a violation of the CAPM since we can never measure the market portfolio and, hence, cannot test the model. Second, no one has yet conclusively shown that variables like size, E/P and P/B are not simply proxies for measurment error in betas. Are we certain, for example, that variation in ratios of P/B is not picking up variation in leverage which is not reflected in OLS betas that are typically estimated with sixty months of prior (stale) prices? The book is not closed -- we think that more research is necessary to resolve these issues.

The research on time series predictability, as a whole, is convincing evidence that expected returns are not constant through time. That most asset pricing models allow for time-varying expected returns renders these results quite appealing; anyway, constant expected returns is a rather confining characterization of stock returns. Some of the temporal patterns in returns -- in particular those relating to calendar turning points -- are more troubling than others since they defy economic (rational) interpretations. For example, we can sensibly tell a business conditions story about time variation in expected returns, but the institutionally-oriented stories that have been advanced for the calendar-related patterns are harder to swallow since they often (always) require irrational market participants.

Finally, there is 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 described above 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 (Merton (1985), Lakonishok and Smidt (1988), and Lo and MacKinlay (1990b)) have warned about adjusting tests of significance for these lost degrees of freedom. Also, that many of these effects have persisted for nearly 100 years in no way guarantees their persistence into the future. How many years of data are necessary to construct powerful tests? Research over the next 100 years will hopefully settle many of these issues.

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Table 1

Monthly Returns and Other Characteristics for Value-Weighted Portfolios of NYSE and AMEX Stocks Formed on the Basis of Market Capitalization

(April 1951–December 1989)

Size Portfolio	Mean Return	Std. Dev. Return	Beta	Mkt. Cap. ($Mill.)	E/P (%)	Price ($)
Smallest	1.65	6.76	1.17	9.7	−6.6	11.13
2	1.48	6.20	1.19	23.2	4.6	16.71
3	1.34	5.77	1.15	41.4	4.5	21.15
4	1.28	5.65	1.17	68.0	7.3	25.31
5	1.33	5.17	1.11	109.8	7.6	29.39
6	1.21	4.82	1.05	178.9	7.6	31.95
7	1.20	4.68	1.04	291.4	7.4	36.83
8	1.23	4.58	1.03	502.3	7.5	40.23
9	1.08	4.41	1.01	902.1	8.0	49.72
Largest	0.99	4.09	0.95	3983.0	7.7	66.92

Prior to 1962, the portfolios contained only NYSE stocks. The portfolios are created on March 31 of each year using March 31 shares outstanding and prices. 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 over the following twelve months. Portfolios contain only December 31 fiscal closers.

Table 2

International Evidence of a Size Premium[1]

Country:	Australia	Belgium	Canada	Finland	France	Germany	Ireland	Japan	New Zealand	Spain	Switzerland	Taiwan	United Kingdom
Test period	1958–81	1969–83	1973–80	1970–81	1977–88	1954–90	1977–86	1965–87	1977–84	1963–82	1973–88	1979–86	1958–82
No. of securities[2]	281 to 937	170	391	50	529–460	All FSE	40	1st Section TSE	about 100	98 to 140	153	53 to 72	All LSE
No. of size portfolios	10	5	5	10	5	9	5	10	5	10	6	5	10
Market value of largest portfolio of firms divided by market value of smallest portfolio of firms[3]	N.A.	188	67	113	N.A.	N.A.	N.A.	N.A.	60	228	99	17	182
Average monthly return on the smallest portfolio of firms[4]	6.75%	1.17%	1.67%	1.65%	1.20%	1.54%	3.10%	2.57%	0.69%[7]	0.58%[8]	0.94%	0.47%[8]	1.33%
Average monthly return on the largest portfolio of firms[4]	1.02%	0.65%	1.23%	0.89%	0.30%	1.05%	2.63%	1.37%	0.18%[7]	0.02%[8]	0.42%	−0.10%[8]	0.93%
Size premium (small minus large)	5.73%	0.52%	0.44%	0.76%	0.90%	0.49%	0.47%	1.20%	0.51%[7]	0.56%	0.52%	0.57%[8]	0.40%
Average risk of smallest portfolio													
- standard beta coefficient[5]	1.04	1.01	N.A.	0.36	N.A.	0.80	N.A.	1.12[9]	N.A.	N.A.	N.A.	0.79	0.31
- adjusted beta coefficient[6]	N.A.	N.A.	N.A.	0.52	N.A.	N.A.	N.A.	1.22[9]	0.90	N.A.	N.A.	0.55	0.64
Average risk of largest portfolio													
- standard beta coefficient[5]	0.95	0.98	N.A.	1.00	N.A.	1.08	N.A.	0.81[9]	N.A.	N.A.	N.A.	0.99	1.01
- adjusted beta coefficient[6]	N.A.	N.A.	N.A.	0.95	N.A.	N.A.	N.A.	0.77[9]	0.99	N.A.	N.A.	0.72	1.02

[1] Sources: *Australia*, Brown et al. (1983); *Belgium*, Hawawini et al. (1989); *Canada*, Berges et al. (1984); *Finland*, Wahlroos and Berglund (1986); *France*, adapted from Louvet and Taramasco (1991); *Germany*, Stehle (1992); *Ireland*, adapted from Coghlan (1988); *Japan*, Ziemba (1991); *New Zealand*, Gillan (1990); *Spain*, Rubio (1986); *Switzerland*, Cornioley and Pasquier-Dorthe (1991); *Taiwan*, Ma and Shaw (1990); *United Kingdom*, Levis (1985).

[2] TSE = Tokyo Stock Exchange; LSE = London Stock Exchange; NYSE = New York Stock Exchange.

[3] The ratio is based on average market value over the sample period, except for the United States where the ratio is calculated in 1975 and Finland where it is calculated in 1970.

[4] All returns are significantly different from zero at the .05 level.

[5] Standard beta coefficients are estimated using Ordinary Least Square regression; N.A. = not available.

[6] Adjusted betas are estimated using the Scholes and Williams (1977) method in the case of France and Taiwan and the Dimson (1979) method in all other cases. These two methods adjust the estimated beta coefficient for the thin trading that characterizes smaller firms; N.A. = not available.

[7] These are abnormal returns calculated with respect to a Capital Asset Pricing Model.

[8] For Spain the average returns on the small and the large portfolios are returns in excess of those predicted by the Capital Asset Pricing Model; that is, they are risk-adjusted returns. For Taiwan average portfolio returns are calculated using excess returns on individual securities which, in turn, are estimated by subtracting from each individual security return, the return of the portfolio to which a security belongs (five control portfolios are constructed according to the level of systematic risk). Note that another study of the size effect on the Taiwan Stock Exchange (Chou and Johnson (1990)) finds no evidence of a significant size effect after controlling for a P/E effect.

[9] These beta coefficients are from a different sample given in Nakamura and Terada (1984).

Table 3A

Monthly Returns and Other Characteristics for Value-Weighted Portfolios of NYSE and AMEX Stocks Formed on the Basis of the Ratio of Earnings/Price

(April 1962–December 1989)

E/P Portfolio	Mean Return	Std. Dev. Return	Beta	Mkt. Cap. ($Mill.)	E/P (%)	Price ($)
Negative	1.55	8.68	1.46	141.8	−37.7	10.57
Lowest	0.79	5.61	1.09	888.4	2.7	37.55
2	0.90	5.06	1.05	938.4	5.0	38.02
3	0.91	4.94	1.03	849.1	6.4	35.38
4	0.87	4.60	0.95	913.9	7.5	31.91
5	0.79	4.81	0.98	913.5	8.5	31.22
6	0.97	4.65	0.94	808.0	9.4	29.68
7	1.05	4.69	0.96	682.3	10.5	28.11
8	1.13	4.64	0.87	777.8	11.7	27.47
9	1.34	4.65	0.89	656.5	13.3	25.07
Highest	1.25	5.48	1.01	569.8	18.9	21.82

Portfolios are created on March 31 of each year using year-end accounting values and March 31 prices. 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 over the following twelve months. Portfolios contain only December 31 fiscal closers.

Table 3B

Monthly Returns and Other Characteristics for Value-Weighted Portfolios of NYSE and AMEX Stocks Formed on the Basis of the Ratio of Cash Flow/Price

(April 1972–December 1989)

CF/P Portfolio	Mean Return	Std. Dev. Return	Beta	Mkt. Cap. ($Mill.)	CF/P (%)	E/P (%)
Negative	1.32	10.38	1.53	114.4	−53.5	−77.2
Lowest	0.73	5.73	1.06	969.2	5.1	1.0
2	0.92	5.47	1.05	1159.3	8.9	5.2
3	1.12	5.55	1.07	861.6	11.6	7.0
4	1.14	5.27	1.03	897.5	13.8	8.2
5	0.90	5.37	1.02	1042.1	16.1	9.1
6	1.38	5.49	1.05	800.0	18.5	10.7
7	1.05	5.54	1.03	832.3	21.2	11.1
8	1.32	5.11	0.94	899.5	24.5	12.0
9	1.35	5.20	0.92	1182.2	29.6	13.0
Highest	1.62	5.81	1.03	821.7	50.5	12.9

Portfolios are created on March 31 of each year using year-end accounting values and March 31 prices. 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 over the following twelve months. Portfolios contain only December 31 fiscal closers.

Table 4

Monthly Returns and Other Characteristics for Value-Weighted Portfolios of NYSE and AMEX Stocks Formed on the Basis of the Ratio of Price/Book

(April 1962–December 1989)

P/B Portfolio	Mean Return	Std. Dev. Return	Beta	Mkt. Cap. ($Mill.)	P/B	Price ($)
Negative	1.66	8.37	1.29	118.6	−7.07	13.62
Lowest	1.49	5.63	1.04	260.5	0.57	15.98
2	1.46	4.98	0.95	401.2	0.83	19.99
3	1.08	4.66	0.90	619.5	1.00	23.56
4	1.12	4.32	0.83	667.7	1.15	25.72
5	0.96	4.56	0.90	641.1	1.31	26.94
6	0.82	4.45	0.91	834.6	1.52	29.97
7	0.86	4.79	0.98	752.3	1.80	31.45
8	0.93	4.87	1.02	813.0	2.20	33.93
9	0.85	5.25	1.11	1000.8	2.95	37.48
Highest	0.91	5.19	1.05	1429.8	8.17	46.66

Portfolios are created on March 31 of each year using year-end accounting values and March 31 prices. 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 over the following twelve months. Portfolios contain only December 31 fiscal closers.

Table 5

Average Rank Correlations (t-statistics)
Between Several Predetermined Characteristics for NYSE and AMEX Stocks, and also Between These Characteristics and Returns During the Following Year

(1962–1989)

	Market Capitalization	Earnings/ Price	Price/ Book	Dividend Yield	Price
Earnings/Price	0.05 (1.70)				
Price/Book	0.30 (15.09)	−0.29 (−10.99)			
Dividend Yield	−0.01 (−0.51)	0.36 (14.50)	−0.47 (−31.03)		
Price per Share	0.78 (84.94)	0.11 (4.02)	0.33 (17.40)	−0.13 (−7.12)	
Annual Return $(t+1)$	0.03 (0.92)	0.12 (6.09)	−0.07 (−3.45)	0.04 (1.46)	0.04 (1.15)

*Correlations are computed annually using ranks of individual stocks. All rankings are conducted at the end of March, using prices at that time and accounting numbers for the previous fiscal year.

Table 6

Size-Adjusted Monthly Returns (in percent) and Other Characteristics for Twenty-Five Portfolios of NYSE and AMEX Stocks Ranked First by P/B Ratio and Then by E/P Ratio over the Period April 1962–December 1989[a]

	Earnings/Price				
	Lowest	2	3	4	Highest
Returns					
Low P/B	0.23	0.22	0.48	0.50	0.09
	(0.22)[b]	(0.16)	(0.14)	(0.15)	(0.17)
2	0.14	−0.17	0.15	0.12	0.23
	(0.16)	(0.15)	(0.13)	(0.14)	(0.16)
3	−0.06	−0.26	−0.02	0.07	0.06
	(0.14)	(0.13)	(0.12)	(0.13)	(0.14)
4	−0.07	0.02	−0.09	−0.18	0.08
	(0.15)	(0.11)	(0.11)	(0.12)	(0.16)
High P/B	−0.04	−0.10	−0.01	−0.03	0.04
	(0.18)	(0.13)	(0.11)	(0.11)	(0.15)
Price/Book					
Low P/B	0.69	0.74	0.74	0.74	0.68
2	1.08	1.09	1.08	1.07	1.06
3	1.43	1.44	1.42	1.40	1.40
4	2.05	2.04	2.00	1.97	1.95
High P/B	7.24	5.18	4.52	4.03	5.43
Earnings/Price (%)					
Low P/B	4.27	8.85	11.49	13.98	20.61
2	5.56	9.06	10.81	12.38	16.11
3	5.41	8.25	9.68	11.15	14.86
4	4.52	7.02	8.21	9.43	13.24
High P/B	2.80	4.68	5.76	6.98	10.49
Price ($)					
Low P/B	14.25	19.58	21.13	22.05	20.86
2	22.07	27.11	27.22	27.27	23.61
3	29.50	29.73	30.66	29.99	26.40
4	35.27	35.91	35.61	33.23	28.39
High P/B	55.48	49.10	45.08	39.06	31.17
Market Capitalization ($Mill.)					
Low P/B	154.97	253.64	445.70	631.84	493.98
2	493.02	727.21	704.08	737.39	764.06
3	651.23	770.02	850.76	859.78	745.37
4	705.64	951.78	921.58	839.79	661.38
High P/B	1519.06	1472.63	1314.98	1282.69	914.91

[a] We define a size-adjusted monthly return for security i as the return for that security minus the monthly portfolio return for the size decile in which security i is a member. P/B and E/P 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 values of either P/B or E/P are excluded from the sample.

[b] Standard deviations of returns in parentheses. ($N = 333$).

Table 7

International Evidence of Daily Seasonality in Stock Market Returns[1]

(Average Daily Percentage Returns)

Country[2]	Period	Index	Monday	Tuesday	Wednesday	Thursday	Friday	Saturday
Australia	1975–84	EWI	0.044	−0.116*	0.045	0.198*	0.157*	–
Belgium	1977–85	VWI	0.098*	−0.032	0.041*	0.111*	0.130*	–
Belgium	1977–85	EWI	0.080*	0.026*	0.046*	0.044*	0.062*	–
Canada	1969–84	TC[3]	−0.157*	−0.003	0.073*	0.075*	0.094*	–
Finland	1977–82	VWI	0.086*	0.066*	0.030	0.070*	0.074*	–
France	1977–89	VWI	−0.050	0.139*	0.126*	0.133*	0.100*	–
France	1977–89	EWI	0.083*	0.117*	0.122*	0.117*	0.013	–
Germany	1970–85	VWI	−0.085*	0.008	0.044	0.060*	0.099*	–
Germany	1970–85	EWI	0.005	−0.026	0.031	0.066	0.107*	–
Greece	1978–86	VWI	0.070*	−0.058	−0.073*	0.041	0.160*	–
Japan	1949–88	VWI[4]	−0.071*	−0.044*	0.115*	0.081*	0.042*	0.133*
Korea	1980–84	VWI	−0.072*	−0.087*	0.087*	0.014*	0.120*	0.230*
Singapore	1969–84	ST[5]	−0.036	−0.107*	0.079*	0.121*	0.100*	–
Spain	1979–83	VWI	–	−0.072*	0.003	0.037	0.071*	–
United Kingdom	1969–84	FT-A[6]	−0.095*	0.106*	0.090*	0.011	0.044	–
United States	1928–52	S&P[7]	−0.223*	0.076*	0.084*	0.066	0.029	0.147*
United States	1952–82	S&P[7]	−0.154*	0.026	0.103*	0.036*	0.092*	–

[1]Starred returns are significantly different from zero at the 0.05 level. Non-starred returns are not significantly different from zero. VWIndex = Value-Weighted Index; EWIndex = Equally-Weighted Index.

[2]**Sources:** *Australia*, Ball and Bowers (1987); *Belgium*, Corhay (1991); *Canada*, *Singapore*, *United Kingdom*, Condoyanni, O'Hanlon and Ward (1987); *Finland*, Berglund (1985); *France*, Hamon and Jacquillat (1991); *Germany*, Frantzmann (1988); *Greece*, Condoyanni, McLeay and O'Hanlon (1989); *Japan*, Ziemba (1991); *Korea*, Kim (1988); *Spain*, Santesmases (1986); *United States*, Keim and Stambaugh (1984).

[3]TC = Toronto Composite Index.

[4]The Tokyo Stock Exchange is closed on Saturdays since February 1989.

[5]ST = Straits Times Index.

[6]FT-A = Financial Times All Share Index.

[7]S&P = Standard & Poor's Composite Index. The New York Stock Exchange is closed on Saturdays since June 1952.

Table 8

International Evidence of Monthly Seasonality in Stock Market Returns[1]

(Average Monthly Percentage Returns)

Country	Jan.	Feb.	March	April	May	June	July	Aug.	Sept.	Oct.	Nov.	Dec.	All Mths.
1. Australia	2.65	−0.58	0.51	0.84	0.97	0.43	0.66	−0.37	−2.39	2.13	−0.85	3.99	0.67
2. Belgium	3.20	1.09	0.40	1.48	−1.36	−0.84	1.44	−1.17	−1.87	−0.69	0.42	−0.09	0.17
3. Canada	2.90	0.07	0.79	0.41	−0.96	−0.30	0.69	0.60	−0.06	−0.82	1.44	2.61	0.61
4. Finland	3.62	2.09	2.63	0.95	−0.02	1.55	2.53	0.97	−0.76	0.77	0.61	2.51	N.A.[2]
5. France	3.72	−0.18	1.98	0.94	−0.66	−1.90	1.53	1.03	−1.21	−0.72	0.43	0.15	0.43
6. Germany	1.62	0.55	2.17	0.80	0.15	1.62	2.11	2.18	−0.50	−0.46	0.93	1.42	1.09
7. Hong Kong	7.98	4.22	−2.34	1.86	2.34	2.40	2.74	−0.86	−1.69	3.21	−2.49	5.45	1.90
8. Japan	5.58	1.00	2.90	0.51	0.18	2.51	0.65	0.92	0.13	0.52	1.81	2.48	1.60
9. Korea	0.42	2.10	3.72	0.73	1.65	2.54	2.91	−0.13	−0.01	−0.67	3.02	3.34	1.64
10. Malaysia	1.70	0.21	−0.07	0.11	0.49	0.04	−0.34	−0.38	−0.10	0.45	−0.25	0.54	0.20
11. Netherlands	3.74	−0.53	0.62	2.49	0.54	−1.43	1.02	−1.69	−3.34	−0.96	−0.80	1.48	0.38
12. Singapore	7.81	0.69	0.28	0.54	3.45	1.76	0.71	−1.12	−0.11	−0.23	−0.07	1.91	1.31
13. Spain	3.04	1.99	0.14	0.14	−0.81	0.59	1.47	1.00	−2.02	0.18	0.36	−0.45	N.A.[2]
14. Taiwan	6.26	3.41	2.40	4.97	3.49	1.24	0.88	3.83	4.01	−3.70	0.47	1.63	2.41
15. United Kingdom	3.40	0.69	1.25	3.13	−1.21	−1.69	−1.11	1.88	−0.24	0.80	−0.61	2.06	0.70
16. U.K. FT-A[3]	3.06	0.79	1.15	3.57	−1.00	−0.85	−0.22	2.62	0.03	1.26	0.16	2.34	1.08
17. United States	1.04	−0.41	1.27	0.96	−1.38	−0.56	0.14	0.34	−0.79	0.78	1.03	1.42	0.32
18. U.S. EW[4]	5.08	0.55	1.55	0.44	−1.42	−1.00	0.73	0.72	−0.42	−0.79	1.79	1.37	N.A.[2]

[1] Sources: *Finland*, Berglund (1985; the period covered is from January 1970 to December 1983); *United Kingdom*, Levis (1985; the period covered is from January 1958 to December 1982 for Financial Times All Share Index); *Spain*, Rubio (1986; the period covered is January 1963 to December 1982); *Netherlands*, van der Bergh and Wessels (1985; the period covered is January 1966 to December 1982); *Malaysia*, Wong et al. (1990; the period covered is 1970 to 1985; Index is value-weighted); *Germany*, Stehle (1992; the period covered is 1954 to 1990; value-weighted index of all stocks listed on the Frankfurt Stock Exchange); *Japan*, Hawawini (1991; the period covered is from January 1955 to December 1985 for an equally-weighted index of up to 566 stocks); *Hong Kong, Korea, Taiwan* and *Singapore*, Lee (1992; the period covered is from January 1970 to December 1989, except *Korea* that begins in 1975; *Hong Kong* index is Hang Seng; *Korea* index is Korea Composite Stock; *Taiwan* index is Taiwan Stock Exchange; *Singapore* index is Straits Times). All other data are from Gultekin and Gultekin (1983) and are based on stock market indexes from Capital International Perspective (value-weighted indexes). The period covered is January 1959 to December 1979.

[2] Not Available.

[3] FT-A = Financial Times All Share index.

[4] Equally-weighted index of New York Stock Exchange shares.

Table 9

International Evidence of Seasonality in the Size Premium[1]

Country:	Belgium	Finland	France	Germany	Japan	Taiwan	United Kingdom
Test period	1969–1983	1970–1983	1968–1980	1954–1990	1965–1987	1979–1986	1958–1982
Number of securities[2]	170	40	201	All FSE	First Section TSE	53 to 72	All LSE
Number of size portfolios	5	5	5	9	10	5	10
$\frac{\text{Market value largest portfolio}}{\text{Market value smallest portfolio}}$	188	113	83	N.A.	N.A.	17	182
January return on:[3]							
- Smallest portfolio	5.4%	5.9%	3.7%	2.8%	9.5%	N.A.	2.3%
- Largest portfolio	3.0%	2.5%	4.4%	1.3%	2.3%	N.A.	3.6%
January size-premium[4]	2.4%	3.4%	−0.7%	1.6%	7.2%	3.4%[6]	−1.3%
Rest-of-the-year monthly return on:[5]							
- Smallest portfolio	0.8%	1.8%	1.4%	1.4%	1.9%	N.A.	1.2%
- Largest portfolio	0.4%	1.0%	0.7%	1.0%	1.3%	N.A.	0.7%
Rest-of-the-year size premium	0.4%	0.8%	0.7%	0.4%	0.6%	0.3%[6]	0.5%

[1]Sources: *Belgium*, Hawawini et al. (1988); *Finland*, Berglund (1985); *France*, Hamon (1986); *Germany*, Stehle (1992); *Japan*, Ziemba (1991); *Taiwan*, Ma et al. (1990), and *United Kingdom*, Levis (1985). Note that except for the cases of Finland and France, the sources used in Table 9 are the same as those used in Table 2.

[2]TSE = Tokyo Stock Exchange; LSE = London Stock Exchange; N.A. = Not Available.

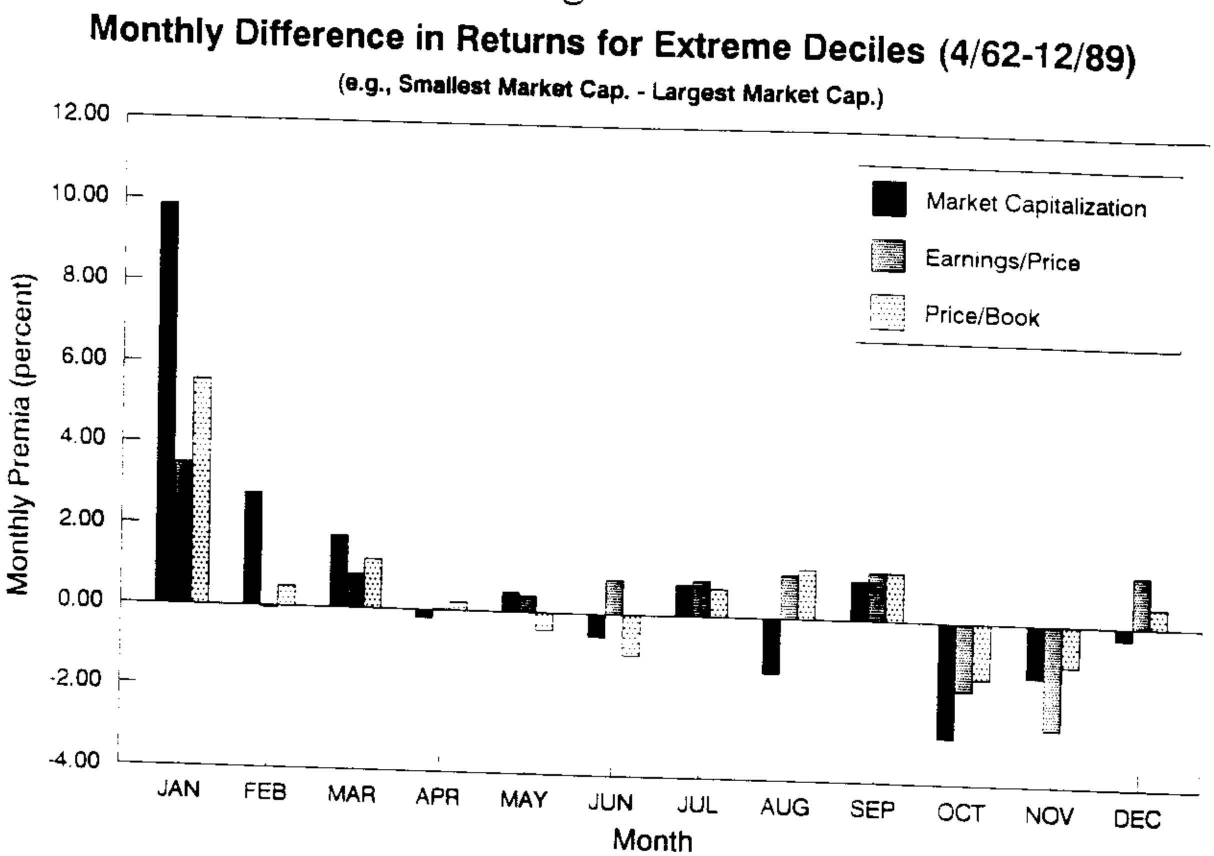
[3]All monthly mean returns are significantly different from zero.

[4]The January size-premium is significantly different from zero only in Belgium, Finland, Taiwan and Japan.

[5]Monthly mean returns are not significantly different from zero for the largest portfolio.

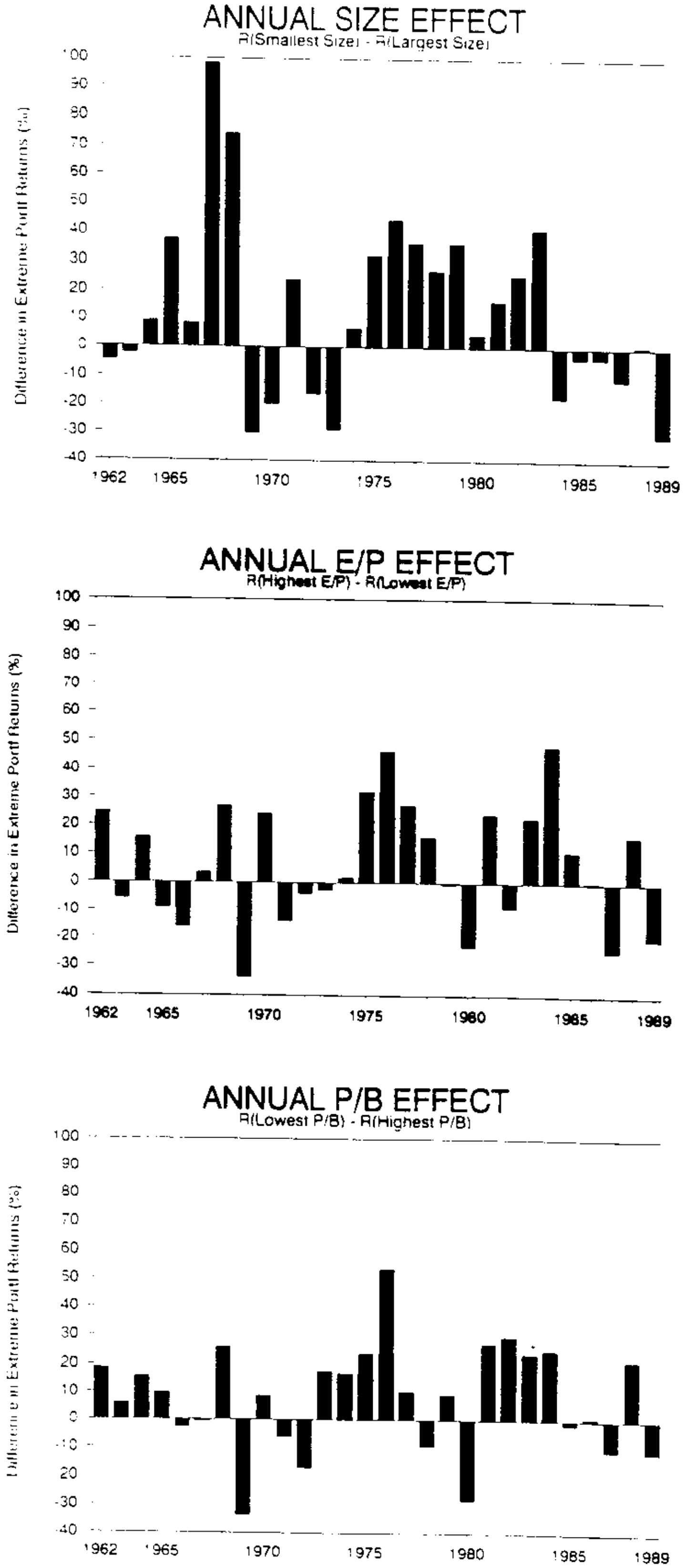
[6]The size-premium is measured with *excess* return estimated with OLS beta coefficients.

Figure 1



Based on monthly returns of value-weighted decile portfolios of NYSE and AMEX stocks. Portfolios are (independently) constructed on March 31 of each year using March 31 shares outstanding and prices, and prior-year-end accounting values. Aside from new listings and delistings, which are added to or dropped from the portfolio as they occur during the year, the portfolio composition remains constant over the following twelve months. The portfolios contain only December 31 fiscal closers.

Figure 2



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