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RISK AND RETURN

January vs. the Rest of the Year*

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This paper examines the seasonality in the basic relationship between expected return and risk during 1935–82. The results reveal that the positive relationship between return and risk is unique to January. The risk premiums during the remaining eleven months are not significantly different from zero.

1. Introduction

In recent years considerable effort has been devoted to the analysis of seasonal movements in the returns on common stocks and to the study of the relationship between these movements and other factors that are correlated with returns, most notably firm size and E/P ratios. While persuasive explanations for seasonality or other so called anomalies are wanting, the empirical evidence clearly indicates that stock returns do vary systematically with the calendar and a number of other factors, including the size of firms.[1]

Interestingly, however, relatively little is known about whether the basic relationship between risk and expected return contains any seasonality. Indeed, we are aware of only one published study of this subject. Conducted by Rozeff

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[1] See the papers in the special issue 'Symposium on Size and Stock Returns, and other Empirical Regularities', *Journal of Financial Economics*, Vol. 12 (June 1983), and Berges, McConnell and Schlarbaum (1984), Gultekin and Gultekin (1983), Roll (1983), Tinic and Barone-Adesi (1983) as well as Banz (1981) and Reinganum (1983).

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and Kinney (1976) as part of broader examination of seasonal movement in stock returns, it found that seasonality is 'a prominent feature of risk premiums estimated from the two-parameter capital asset pricing model'. More specifically, it reported that 'it is January with a relatively *large* risk premium which differs noticeably from the other months'.[2]

While we concur wholeheartedly with Rozeff and Kinney that the relationship between expected returns and risk varies systematically with the calendar, we do not believe their analysis provides an adequate picture of the relationship. As we will show below, it is not simply that January has a *larger* risk premium than other months; rather it is that January is the *only* month to show a consistently positive, statistically significant relationship between expected return and risk.[3] When data for the month of January are withdrawn from the analysis of the risk–return tradeoff, the estimates of risk premiums are not significantly different from zero. January is not simply the month in which overall stock returns have been high relative to the rest of the year, and when small firms' stocks have outperformed the market as a whole; it is the only month when shareholders have consistently been paid for taking on risk!

We have organized our presentation along the following lines. In section 2 we describe in somewhat more detail the work of Rozeff and Kinney. We present our own findings in section 3, and offer some comments and conclusions in section 4.

2. Rozeff and Kinney's study

Rozeff and Kinney investigated the seasonality in the tradeoff between risk and return by examining the behavior of Fama and MacBeth's (1973) well-known estimates of the two-parameter capital-asset pricing model (CAPM). Fama and MacBeth, it will be recalled, made monthly estimates of $\gamma_{0\tau}$, the expected return on the minimum-variance zero-beta portfolio, and $\gamma_{1\tau}$, the market's risk premium, for the period from January 1935 through June 1968 for the New York Stock Exchange. To avoid errors in the measurement of systematic risk, they based their estimates on cross-sectional data for twenty portfolios of securities instead of individual stocks. Their inferences about the tradeoff between risk and return resulted from an analysis of the average values of $\gamma_{0\tau}$ and $\gamma_{1\tau}$ over the entire period covered by their data.

[2] Rozeff and Kinney (1976, p. 400).

[3] We recently became aware of a paper written by Keim (1980) while he was a doctoral student at the University of Chicago. Like Rozeff and Kinney (1976), Keim described the seasonality in the behavior of risk premiums. However, he also failed to see that the real significance of the data is not related to the level of the risk premiums over the year, but to the presence or absence of premiums on a month-by-month basis.

For the periods 1935–67 and 1941–67, Rozeff and Kinney calculated the 75% truncated means and the standard deviations of Fama and MacBeth's estimates of $\gamma_{0\tau}$ and $\gamma_{1\tau}$ for each month of the year. They then performed simple parametric and non-parametric statistical tests of the hypothesis that the monthly truncated means for each of the two coefficients were equal. Their conclusion was that although they could not reject this hypothesis for $\gamma_{0\tau}$, they could for $\gamma_{1\tau}$. In their own words:

> Both tests reject the hypothesis of equal means for $\gamma_{1\tau}$ at the 1% level for 1941–67. The χ^2 of 15.8, which is significant at the 10% level and indicates a lack of homogeneity of variances, disagrees slightly with the low Siegel–Tukey statistic. However, given the large difference in means which is evident in the sample values of the 75% truncated means, we are quite confident of highly significant differences in mean $\gamma_{1\tau}$ by month. In particular, the January mean of 0.0450 compared to the average of 0.0056 for all months is truly extraordinary. To the extent that realizations from our somewhat small sample size of 27 monthly observations adequately measure expectations, it seems that the tradeoff of return for risk demanded and received by the market in January is much *greater* than in other months of the year [emphasis added].[4]

In spite of the insights provided by their results, Rozeff and Kinney, for whatever reasons, did not go the next step to ask whether the month-to-month variability in $\bar{\gamma}_1$ might represent something more than mere seasonality in the *level* of the market's risk premium. Put somewhat differently, their results did not cause them to question Fama and MacBeth's conclusion that the data for 1935–1968 provided support for the hypothesis that undiversifiable risk and expected return are positively related. Instead, they merely lead them to state that the market demands a systematically *higher* risk premium in the month of January. In view of Rozeff and Kinney's conclusion that abnormal returns computed from the two-parameter model 'will be free from seasonal effects, or at least much more free than the market model residuals',[5] it is evident that they were content to accept the argument that Fama and MacBeth's findings were supportive of the existence of a risk premium that merely varied in level over the year.

3. Risk and return over the year

Based on Rozeff and Kinney's findings and the growing body of data indicating that the stock market's behavior in January is different from the rest

[4] Rozeff and Kinney (1976, p. 400).

[5] Ibid.

Table 1

Average values of the Fama and MacBeth estimates of intercept and slope coefficients of the two-parameter model (estimated with monthly data and based on the equally weighted index).[a]

Averaged over	Intercept coefficient ($\bar{\gamma}_0$)	Slope coefficient ($\bar{\gamma}_1$)	Sample size
	January 1935 to June 1968		
January only	−0.000744 (−0.1480)	0.044509[b] (3.7347)	34
Rest of the year	0.006736[b] (3.3674)	0.005136 (1.5204)	368
All months	0.006104[b] (3.2447)	0.008466[b] (2.5703)	402
	January 1935 to September 1951		
January only	−0.002371 (−0.2938)	0.052459[b] (2.4872)	17
Rest of the year	0.005453 (1.5878)	0.007852 (1.2876)	184
All months	0.004792 (1.4900)	0.011624[b] (1.9668)	201
	October 1951 to June 1968		
January only:	0.000882 (0.1416)	0.036559[b] (3.1689)	17
Rest of the year	0.008020[b] (3.8962)	0.002420 (0.8300)	184
All months	0.007416[b] (3.7888)	0.005307[b] (1.8232)	201

[a] t statistics are presented in parentheses.
[b] Significant at 0.05 level.

of the year, we thought it appropriate to look much more deeply into the seasonal behavior of the estimated coefficients for γ_{0T} and γ_{1T}. What we found is both interesting and puzzling.

3.1. An analysis of Fama and MacBeth's results

In table 1 we have reproduced Fama and MacBeth's test of the two-parameter model for the entire period covered by their data, along with similar tests for their results for January and the rest of the year. As is readily apparent, the relatively high t value of $\bar{\gamma}_1$ for the entire period results primarily from what happens in January.[6] When the January data are analyzed on their own, the t

[6] When returns are leptokurtic the significance of the t statistic would be overstated. However, Fama (1976, p. 38) concludes that monthly returns are approximately normal, and Fama and MacBeth (1973) use the t distribution in testing the two-parameter model.

value for $\bar{\gamma}_1$ is even more significant; when the data for the rest of the year are analyzed, the t value for $\bar{\gamma}_1$ is insignificantly different from zero. Quite the opposite can be said about the behavior of the t value for $\bar{\gamma}_0$. While highly significant in the analysis of the data for the entire period and for the months from February through December, it is insignificant in the analysis of January's data.

On the theory that the data for other individual months might also be playing an important role, we performed similar tests for the various eleven-

Table 2

Fama and MacBeth estimates of intercept and slope coefficients of the two-parameter model (January 1935 to June 1968).[a]

Averaged over all months excluding	Average intercept coefficient ($\bar{\gamma}_0$)	Average slope coefficient ($\bar{\gamma}_1$)
January	0.006736[b] (3.3674)	0.005136 (1.5204)
February	0.006338[b] (3.1696)	0.008513[b] (2.4418)
March	0.006555[b] (3.3941)	0.009626[b] (2.7612)
April	0.005797[b] (2.9322)	0.008811[b] (2.5297)
May	0.005613[b] (2.9637)	0.009646[b] (2.7559)
June	0.006144[b] (3.1041)	0.008799[b] (2.5591)
July	0.005621[b] (2.8196)	0.007348[b] (2.1156)
August	0.005887[b] (2.9456)	0.009968[b] (2.8301)
September	0.007544[b] (3.9715)	0.007634[b] (2.4828)
October	0.006583[b] (3.3319)	0.008391[b] (2.4185)
November	0.005028[b] (2.5931)	0.009245[b] (2.6664)
December	0.005857[b] (2.9715)	0.007850[b] (2.2965)
January and July	0.006154[b] (2.8883)	0.003317 (0.9319)
Across all months	0.006104[b] (3.2447)	0.008466[b] (2.5703)

[a] t statistics are presented in parentheses.
[b] Significant at the 0.05 level.

month periods associated with omitting the month from February through December. Our results, which are summarized in table 2, show that in all of the various eleven-month periods, the t value for $\bar{\gamma}_1$ is highly significant and the t values for $\bar{\gamma}_0$ are also large.

Readers will note that at the bottom of table 2 we also present results for the ten-month period when the data for January and July are both omitted. We decided to make this calculation on the basis of the fact that the t statistic of the eleven-month period without July was somewhat lower than any of the other eleven-month periods, save the one in which January is omitted. Consistent with what might have been expected, omitting both of these months resulted in an even lower t value. The important point to remember, however, is that the January data dominate in producing this result.

Lest the reader wonder whether the January data primarily reflected a few observations that were very large, we would point out that 27 of the 34 estimates of γ_1 were positive in January. Although we did not make a thorough analysis of the month-by-month data for 1941–67, which Rozeff and Kinney identified as a period of more homogeneous stock returns, it is interesting to note that the estimates of γ_1 were positive in 24 of the 28 Januaries during this period.[7]

To obtain a somewhat better sense of the seasonal behavior of Fama and MacBeth's estimates of $\gamma_{0\tau}$ and $\gamma_{1\tau}$, we ran the following regression for the entire period covered by their data and for two subperiods, January 1935 through September 1951, and October 1951 through June 1968:

$$\tilde{\gamma}_{j\tau} = \beta_1 + \sum_{i=2}^{12} \beta_i D_i + \tilde{e}_{j\tau}, \tag{1}$$

where $j = 0$ and 1, respectively, and D_2 through D_{12} is a set of dummy variables representing the months of the year from February through December. The intercept measures the mean γ_0 or γ_1 in January. The regression equation corresponds to an analysis of variance where the differences in monthly means of γ_0 and the market risk premiums from the January average are captured by the regression coefficients. As the reader can plainly see in table 3, the intercept coefficient, β_1, representing January, is positive and significant in all three cases. In the regression for the whole period, the dummy variable for July, D_7, is the only one that is not significantly different from January, reflecting what might have been expected from table 2. In the regression for the second half of the period, however, it is significantly lower than January. It is interesting to note, by the way, that studies of the

[7] See Fama (1976, pp. 357–360).

Table 3

Seasonality in the intercept ($\gamma_{0\tau}$) and slope ($\gamma_{1\tau}$) coefficients of the two-parameter model estimated from $\tilde{\gamma}_{j\tau} = \beta_1 + \sum_{i=2}^{12} \beta_i D_i + \tilde{e}_{j\tau}$.[a]

	1/1935–6/1968 402		1/1935–9/1951 201		10/1951–6/1968 201	
	γ_0	γ_1	γ_0	γ_1	γ_0	γ_1
β_1	−0.0012 (−0.20)	0.0432[b] (3.90)	−0.0032 (−0.30)	0.0495[b] (2.53)	0.0009 (0.13)	0.0366[b] (3.72)
β_2	0.0055 (0.60)	−0.0350[b] (−2.20)	0.0065 (0.41)	−0.0385 (−1.35)	0.0043 (0.45)	−0.0310[b] (−2.23)
β_3	0.0023 (0.26)	−0.0450[b] (−2.85)	−0.0031 (−0.20)	−0.0608[b] (2.17)	0.0077 (0.81)	−0.0288[b] (−2.08)
β_4	0.0117 (1.30)	−0.0415[b] (−2.63)	0.0171 (1.11)	−0.0529[b] (−1.89)	0.0063 (0.66)	−0.0298[b] (−2.14)
β_5	0.0114 (1.26)	−0.0478[b] (−3.03)	0.0284[b] (1.84)	−0.0681[b] (−2.43)	−0.0057 (−0.61)	−0.0271[b] (−1.95)
β_6	0.0075 (0.83)	−0.0392[b] (−2.48)	0.0096 (0.62)	−0.0349 (−1.24)	0.0054 (0.57)	−0.0431[b] (−3.11)
β_7	0.0139 (1.52)	−0.0196 (−1.23)	0.0091 (0.59)	−0.0092 (−0.33)	0.0190[b] (1.98)	−0.0306[b] (−2.17)
β_8	0.0086 (0.94)	−0.0501[b] (−3.15)	0.0084 (0.54)	−0.0514[b] (−1.83)	0.0088 (0.92)	−0.0487[b] (−3.45)
β_9	−0.0080 (−0.88)	−0.0265[b] (−1.67)	−0.0146 (−0.95)	−0.0091 (−0.32)	−0.0010 (−0.10)	−0.0452[b] (−3.20)
β_{10}	0.0040 (0.43)	−0.0327[b] (−2.06)	0.0027 (0.17)	−0.0234 (−0.82)	0.0049 (0.52)	−0.0407[b] (−2.93)
β_{11}	0.0183[b] (2.00)	−0.0455[b] (−2.86)	0.0170 (1.08)	−0.0699[b] (−2.45)	0.0193[b] (2.03)	−0.0218 (−1.57)
β_{12}	0.0133 (1.46)	−0.0350[b] (−2.20)	0.0167 (1.07)	−0.0396 (−1.39)	0.0099 (1.05)	−0.0300[b] (−2.16)

[a] $j = 0$ and 1 respectively, and D_2 through D_{12} is a set of dummy variables representing the months of the year from February through December. The t statistics are presented in parentheses.
[b] Significant at 0.05 level.

seasonality of stock prices in Australia have also identified July as a month having many of the same anomalies as January.[8]

3.2. *Updating the analysis*

On the theory that Fama and MacBeth's results might have arisen simply from some peculiarities associated with the period they studied, we updated the monthly estimates of $\gamma_{0\tau}$ and $\gamma_{1\tau}$ through December 1982, and performed the

[8] See Officer (1975) and Brown, Kleidon and Marsh (1983).

same tests described in tables 1, 2, and 3. Our results are presented in tables 4, 5, and 6. Looking first at table 4, we see that, for the entire period from January 1935 through December 1982, the value for $\bar{\gamma}_1$ for all months is slightly below the value reported by Fama and MacBeth, but that its t statistic is even higher. Quite clearly, however, the reason for this continues to be the behavior of the January coefficients, as reflected in the fact that the t statistic for $\bar{\gamma}_1$ in January is 4.6335. For the other months, in contrast, it is only 1.4145.

Table 4

Average values of the updated estimates of intercept and slope coefficients of the two-parameter model (estimated with monthly data and based on the equally weighted index).[a]

Averaged over	Intercept coefficient ($\bar{\gamma}_0$)	Slope coefficient ($\bar{\gamma}_1$)	Sample size
	January 1935 to December 1982		
January only	−0.000645 (−0.1443)	0.047052[b] (4.6335)	48
Rest of the year	0.006689[b] (4.0127)	0.003806 (1.4145)	528
All months	0.006078[b] (3.8635)	0.007410[b] (2.7966)	576
	January 1935 to December 1958		
January only	0.002686 (0.4561)	0.041207[b] (2.8710)	24
Rest of the year	0.008287[b] (3.3063)	0.004445 (1.0345)	264
All months	0.007820[b] (3.3326)	0.007509[b] (1.8111)	288
	January 1959 to December 1982		
January only	−0.003976 (−0.5754)	0.052897[b] (3.5543)	28
Rest of the year	0.005091[b] (2.3104)	0.003167 (0.9724)	264
All months	0.004335[b] (2.0654)	0.007312[b] (2.2053)	288
	January 1969 to December 1982		
January only	−0.006995 (−0.7758)	0.060880[b] (2.6394)	14
Rest of the year	0.005158[b] (1.6797)	0.001573 (0.3305)	154
All months	0.004145 (1.4239)	0.006515 (1.3340)	168

[a] t statistics are presented in parentheses.
[b] Significant at 0.05 level.

Table 4 also provides results for a number of subperiods, including subperiods based on dividing the data into two equal parts, as well as the post Fama-MacBeth subperiod beginning in January 1969. As readers can see, the same basic pattern prevails, irrespective of the subperiod studied. In all subperiods the t statistic for $\bar{\gamma}_1$ during January is highly significant. For the rest of the year, however, it is positive but insignificant. For all of the months combined, it is significant in the two long subperiods but insignificant in the shorter subperiod beginning in January 1969.

Table 5

Updated estimates of intercept and slope coefficients of the two-parameter model (January 1935 to December 1982).[a]

Averaged over all months excluding	Average intercept coefficient ($\bar{\gamma}_0$)	Average slope coefficient ($\bar{\gamma}_1$)
January	0.006689[b]	0.003806
	(4.0127)	(1.4145)
February	0.006752[b]	0.007182[b]
	(4.0690)	(2.5713)
March	0.006673[b]	0.007665[b]
	(4.0935)	(2.7506)
April	0.006082[b]	0.007615[b]
	(3.7372)	(2.7198)
May	0.006113[b]	0.008468[b]
	(3.7771)	(3.0048)
June	0.005688[b]	0.008382[b]
	(3.4425)	(3.0022)
July	0.005882[b]	0.006451[b]
	(3.5636)	(2.3393)
August	0.006121[b]	0.008016[b]
	(3.7051)	(2.8399)
September	0.006874[b]	0.007310[b]
	(4.2041)	(2.8334)
October	0.005618[b]	0.008163[b]
	(3.4096)	(2.9539)
November	0.005123[b]	0.008053[b]
	(3.1698)	(2.8933)
December	0.005318[b]	0.007812[b]
	(3.2024)	(2.7769)
January and July	0.006535[b]	0.002391
	(3.7140)	(0.8543)
Across all months	0.006078[b]	0.007410[b]
	(3.8635)	(2.7966)

[a] t statistics are presented in parentheses.
[b] Significant at the 0.05 level.

Table 6

Seasonality in the intercept ($\gamma_{0\tau}$) and slope ($\gamma_{1\tau}$) coefficients of the two-parameter model estimated from $\tilde{\gamma}_{jt} = \beta_1 + \Sigma_{i=2}^{12} \beta_i D_i + \tilde{e}_{jt}$ for sample sizes 576, 288 and 168.[a]

	1/1935–12/1982 576		1/1935–12/1958 288		1/1959–12/1982 288		1/1969–12/1982 168	
	γ_0	γ_1	γ_0	γ_1	γ_0	γ_1	γ_0	γ_1
β_1	−0.0006 (−0.12)	0.0470[b] (5.20)	0.0027 (0.33)	0.0412[b] (2.90)	−0.0040 (−0.55)	0.0529[b] (4.75)	−0.0070 (−0.71)	0.0609[b] (3.71)
β_2	−0.0007 (−0.09)	−0.0371[b] (−2.90)	−0.0003 (−0.03)	−0.0332[b] (−1.65)	−0.0011 (−0.10)	−0.0410[b] (−2.61)	−0.0017 (−0.12)	−0.0506[b] (−2.18)
β_3	0.0002 (0.023)	−0.0424[b] (−3.31)	0.0007 (0.06)	−0.0505[b] (−2.51)	−0.0004 (−0.04)	−0.0344[b] (−2.18)	−0.0006 (−0.04)	−0.0370 (−1.60)
β_4	0.0067 (0.87)	−0.0419[b] (−3.27)	0.0105 (0.92)	−0.0450[b] (−2.24)	0.0028 (0.27)	−0.0388[b] (−2.47)	0.0007 (0.05)	−0.0467[b] (−2.01)
β_5	0.0063 (0.82)	−0.0513[b] (−4.00)	0.0142 (1.24)	−0.0497[b] (−2.47)	−0.0016 (−0.15)	−0.0528[b] (−3.36)	0.0043 (0.31)	−0.0664[b] (−2.86)
β_6	0.0110 (1.43)	−0.0503[b] (−3.93)	0.0086 (0.75)	−0.0333[b] (−1.66)	0.0134 (1.32)	−0.0673[b] (−4.28)	0.0205 (1.46)	−0.0755[b] (−3.25)
β_7	0.0089 (1.16)	−0.0291[b] (−2.27)	0.0076 (0.66)	−0.0100 (−0.50)	0.0101 (0.99)	−0.0481[b] (−3.06)	0.0033 (0.23)	−0.0510[b] (−2.20)
β_8	0.0062 (0.81)	−0.0463[b] (−3.62)	0.0039 (0.34)	−0.0496[b] (−2.47)	0.0086 (0.84)	−0.0430[b] (−2.74)	0.0093 (0.67)	−0.0445[b] (−1.92)
β_9	−0.0020 (−0.27)	−0.0385[b] (−3.01)	−0.0134 (−1.16)	−0.0158 (−0.79)	0.0093 (0.91)	−0.0612[b] (−3.89)	0.0174 (1.24)	−0.0728[b] (−3.13)
β_{10}	0.0118 (1.54)	−0.0479[b] (−3.74)	−0.0002 (−0.02)	−0.0283 (−1.41)	0.0237[b] (2.33)	−0.0675[b] (−4.29)	0.0329[b] (2.35)	−0.0846[b] (−3.64)
β_{11}	0.0172[b] (2.25)	−0.0467[b] (−3.65)	0.0151 (1.32)	−0.0510[b] (−2.54)	0.0193[b] (1.90)	−0.0424[b] (−2.70)	0.0200 (1.43)	−0.0553[b] (−2.38)
β_{12}	0.0151[b] (1.97)	−0.0441[b] (−3.44)	0.0148 (1.29)	−0.0378[b] (−1.88)	0.0153 (1.51)	−0.0503[b] (−3.20)	0.0275[b] (1.96)	−0.0679[b] (−2.92)

[a] $j = 0$ and 1 respectively, and D_2 through D_{12} is a set of dummy variables representing the months of the year from February through December. The t statistics are presented in parentheses.

[b] Significant at 0.05 level.

Turning to table 5, we see that excluding any month except January from the analysis still yields a t statistic for $\bar{\gamma}_1$ which is greater than 2.0. When the January data are excluded, however, the t statistic for $\bar{\gamma}_1$ drops to only 1.4145. Once again, then, we see that the key month continues to be January.[9]

Finally, in table 6 we show the results of running regression, eq. (1), for the entire period covered by the data and for the three subperiods described above. Readers will note that the intercept, which represents January, is positive and highly significant in all cases and that the remaining coefficients are uniformly

[9] The estimates of $\bar{\gamma}_0$ and $\bar{\gamma}_1$ for the subperiods are available from the authors on request.

negative and predominantly significant. In other words, the values of $\bar{\gamma}_1$ for the months from February through December are significantly lower than $\bar{\gamma}_1$ for January. Indeed, they are virtually equal to zero!

3.3. *Results using a value weighted index*

In his 1980 paper, Keim observed that the parameters of the time series of excess returns used in analyzing CAPM anomalies seem to be sensitive to the type of index employed.[10] Since all of the results presented thus far are based on the use of an equally weighted index, we thought it appropriate to test whether the findings would continue to hold when a value weighted index was used instead. The results are shown in table 7. Simply put, they indicate that using a value weighted index has virtually no impact on the conclusions one can draw. January continues to be the month in which there is a systematic, positive relationship between the realized returns and systematic risks of portfolios.[11]

4. Comments and conclusions

In a recent article dealing with anomalies in stock returns, William Schwert (1983) observed that the empirical support for a positive relationship between risk and expected returns is 'surprisingly weak'. As evidence to support his conclusion Schwert noted that 'in Fama and MacBeth the t statistic testing the hypothesis that the slope of the risk–return relation is zero is 2.57 for the 1935–68 sample period, but it is only 1.92, 0.70 and 1.73 for the 1934–45, 1946–55 and 1956–68 subperiods, respectively'.[12] When the seasonal behavior of the Fama and MacBeth results is considered, however, the t statistic for the 1935–1968 period becomes highly suspect and the basic tradeoff between risk and expected return virtually disappears.

Updating the monthly estimates through 1982 does nothing to change this conclusion. As we saw, although the t statistic for $\bar{\gamma}_1$ for the entire 1935–82 period is even higher than what Fama and MacBeth reported, the impact of the January data is even more pronounced. When these data are withheld, the t statistic for $\bar{\gamma}_1$ is insignificant and lower than for the 1935–68 period. In short, nothing that has happened since the middle of 1968 suggests that the seasonality we found in Fama and MacBeth's data was unique to the period they studied.

[10] Keim (1980, p. 12).

[11] The results of the regression equation (1), which are based on the value weighted index, are not reported to conserve space. They are available from the authors on request.

[12] Schwert (1983, p. 4).

Table 7

Average values of the estimated intercept and slope coefficients of the two-parameter model (estimated with monthly data and based on the value weighted index).[a]

Averaged over	Intercept coefficient ($\bar{\gamma}_0$)	Slope coefficient ($\bar{\gamma}_1$)	Sample size
	January 1935 to December 1982		
January only	0.009830 (1.5380)	0.030210[b] (3.3180)	48
Rest of the year	0.005802[b] (3.2629)	0.003574 (1.5582)	528
All months	0.006138[b] (3.5849)	0.005794[b] (2.5737)	576
	January 1935 to December 1958		
January only	0.015397 (1.4981)	0.022520[b] (1.9791)	24
Rest of the year	0.007839[b] (3.0372)	0.003377 (1.0130)	264
All months	0.008469[b] (3.3774)	0.004972 (1.5522)	288
	January 1959 to December 1982		
January only	0.004262 (0.5408)	0.037899[b] (2.6082)	24
Rest of the year	0.003765 (1.5367)	0.003771 (1.1919)	264
All months	0.003807 (1.6320)	0.006615[b] (2.0841)	288
	January 1969 to December 1982		
January only	−0.001565 (−0.1599)	0.044729[b] (1.9985)	14
Rest of the year	0.002957 (0.8455)	0.003267 (0.7016)	154
All months	0.002580 (0.7831)	0.006722 (1.4323)	168

[a] t statistics are presented in parentheses.
[b] Significant at 0.05 level.

But is it just possible that what we have presented is nothing more than a statistical artifice? A positive answer to this question cannot be rejected out of hand. CAPM, after all, is phrased in terms of an ex ante relationship. Hence, a completely clean test of it, based on studying ex post data, must assume the expectations are being realized on the average, and that the sample size is large enough for the averages to be meaningful. Unfortunately, no one knows for certain what constitutes a sufficient sample size. Fama and MacBeth reasoned that a period of 402 months provided enough data to make some meaningful tests. If they were right, our analysis of their results, not to mention our analysis of the data for an even longer period, would seem to cast serious doubt on the validity of the two-parameter model. But if they were wrong, we may simply have found that not even ex post data from a period approaching a half a century can pick up an ex ante relationship.

In reflecting on whether Fama and MacBeth's results are really telling us anything we think it important to recall that many of the studies conducted since the publication of their paper also point to January as a month that is 'different' from the rest of the year. If average returns can be higher in January and small-firm effects appear in January, why isn't it possible for risk and returns to have a unique relationship in January? Indeed, since Fama and MacBeth's high-beta portfolios are more populated by small capitalization stocks, firm size and beta are negatively correlated. Hence, it may be that the significantly positive risk premiums estimated for January partially reflect the small-firm effect.

In any event, to the extent that the risk–return tradeoff shows up only in January, much of what now constitutes the received version of modern finance is brought into question. While, it can be argued that an investor who holds over a long period of years can be expected to reap the benefits from owning risky securities in January, all of the research dealing with asset valuation assumes that the relationship between risk and expected return is not merely a reflection of what happens in a single month and that investors are being compensated for taking risk throughout the year. Recent textbooks dealing with portfolio management and investment selection, for example, build heavily on the idea of a relatively consistent risk–return tradeoff – one that does not tell investors they must be in the market in January to reap the rewards from taking risk. Similarly, estimates of the cost of equity capital based on CAPM incorporate the notion that increases in a stock's systematic risk result in higher expected returns for shareholders throughout the year.

The results reported here may also have important implications for the empirical tests of alternative asset pricing models, particularly the Arbitrage Pricing Model. While the results of these tests thus far are somewhat ambiguous and highly controversial, they are being interpreted as providing support for a four-factor model which explains the covariance structure of security returns.[13] To our best knowledge, however, none of these studies have examined the seasonalities in the estimated risk premiums. Based on what we have found, it is quite conceivable that the factors are priced only in January. In any event, it seems relevant to investigate whether the Arbitrage Pricing Model is also subject to the type of January effect we have found.

[13] For example, see Roll and Ross (1980, 1983) Brown and Weinstein (1983), Chen, Roll and Ross (1983), and Dhrymes, Friend and Gultekin (1982).

For the time being at least, we do not propose to try to provide a rationale for the results we have presented in this paper. Many other researchers are busy trying to explain what makes January 'different'. We will leave it to them to consider why the risk–return tradeoff only shows up consistently in the first month of the year.

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