RISK-PREMIA SEASONALITY
IN U.S. AND EUROPEAN EQUITY MARKETS

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1. Purpose of the study

There is strong evidence indicating that average monthly stock returns exhibit seasonality; they are systematically different during specific months of the year. In particular, average stock returns have been found to be significantly higher in January than during the remaining eleven months of the year. This January effect has been observed in most stock exchanges around the world.[1]

But seasonality is not confined to monthly mean returns. At least two studies using U.S. common stocks have reported evidence of seasonality in the coefficients of the relationship between the average returns and the systematic risk. First, Rozeff and Kinney (1976) showed that (p.104) "seasonality is also a prominent feature of 'risk premiums' estimated from the two-parameter capital asset pricing model.[2] Again it is January with a relatively large risk premium which differs noticeably from all other months."

More recently, Tinic and West (1984) went a step further. They have shown that (p.562) "it is not simply that January has a larger 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. ... 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 [3] it is the only month when shareholders have consistently been paid for taking on risk!"

The purpose of this study is twofold. First, we re-examine the Tinic and West (1984) findings and extend their investigation to alternative specifications of the risk-return relationship. Second, we perform similar tests on common stocks traded in three European stock exchanges (London, Paris and Brussels) to find out if risk premia seasonality (and the uniqueness of January) is confined only to U.S. common stocks or whether this puzzling phenomeon, like that of seasonality in average monthly stock returns, occurs also in other exchanges around the world. In addition to providing evidence from countries other than the United States, the investigation of risk-premia seasonality in the United Kingdom (London Stock Exchange), France (Paris Bourse) and Belgium (Brussels Bourse) may also provide some clues which could help in understanding this phenomenon.

The rest of the paper is organized as follows. In section 2 we give a brief summary of our major findings. In section 3 we describe the properties of our sample. In section 4 we outline the methodology we employ to perform our analysis. To estimate the coefficients (risk-premia) of the relationship between average monthly stock returns and their corresponding risk we use a methodology similar to that of Fama and MacBeth (1973). We examine various specifications of the risk-return relationship. In section 5 we adopt the two-parameter capital asset pricing model in which the risk measure is systematic risk (beta coefficients) and examine the seasonality of the average estimated monthly systematic risk-premia. In section 6 the risk measure is the variance of stock returns and we examine the seasonality of the average estimated monthly total risk premia. In section 7 the risk measure is unsystematic risk and we look at the seasonality of the average estimated monthly unsystematic risk premia. In section 8 we look at seasonality in the average estimated monthly coefficients of the Fama and MacBeth (1973) four-parameter model. In section 9 we examine the pattern of seasonality in the month-to-month mean returns of stock market indexes and compare it to the pattern of seasonality in the month-to-month average estimated risk premia. The purpose is to find out whether these two phenomena are related. It is possible that risk-premia seasonality is simply another manifestation of stock returns seasonality. We conclude with section 10 in which we examine the behavior of systematic risk (beta coefficient) during each one of the 12 months of the year in order to find out if it exhibits any monthly seasonality.

2. Summary of major findings

Over the 15-year period from January 1969 to December 1983 we found the following results.

(1) In the United States and Belgium, January is the only month of the year during which there is a significant (and positive) relationship between average returns and risk regardless of how risk is measured. The result is quite robust. It holds whether we defined risk as systematic risk (beta coefficient), unsystematic risk (standard error of residual returns) or total risk (variance).

(2) In the United States January is also the only month of the year during which average stock returns are significantly positive.[4]

(3) In Belgium, average stock returns are significantly positive over several months of the year in addition to January. These are February, April, June and July. But during January, mean returns are significantly higher than during any one of the other eleven months of the year.

(4) In the United Kingdom April is the only month of the year during which there is a significant positive relationship between average returns and systematic risk. When risk is measured as total risk (variance) or as unsystematic risk then the April effect in the risk-return relationship disappears and is replaced by a unique January effect.

(5) In the United Kingdom January and April are the only months of the year during which average stock returns are significantly positive with January mean returns higher than April mean returns.

(6) In France there is not a single month of the year during which there is a significant positive relationship between average

returns, however, are significantly positive in January and July. As in the case the United States, the United Kingdom and Belgium, January returns in France are the year's highest.

(7) The United States is the only country in our sample in which significantly positive risk premia and significantly positive mean returns occur during the same unique month of January. This perfect correspondence between the pattern of risk premia seasonality and the pattern of mean returns seasonality may lead to the conclusion that the former phenomenon (risk premia seasonality) may be nothing more than an indirect manifestation of the latter phenomenon (returns seasonality). This hypothesis, however, must be rejected in light of the results we found in the European equity markets where no perfect correspondence exists between the pattern of risk premia seasonality and the pattern of mean returns seasonality.

(8) The capital asset pricing model does not provide an adequate description of the historical relationship between average returns and risk in the four markets in our sample over the 13-year period from January 1971 to December 1983. This result can be interpreted either as a failure of the two-parameter capital asset pricing model to provide an accurate representation of how securities have been priced over our sample period or, according to Roll (1977), as a proof that the capital asset pricing model cannot be tested without knowing the exact composition of the true market portfolio.

(9) We cannot reject the hypothesis that the estimated beta coefficients of 80 domestic portfolios (20 in each of the 4 countries in our sample) are the same whether they are estimated only with January returns (14 observations) or with any one of the other eleven months of the year. Apparently, seasonality in the systematic risk premia is not caused by seasonality in systematic risk.

3. Sample properties

3.1 Characteristics of the four markets

Our sample of stock markets contains the world's largest (New York) and one of the world's smallest (Brussels). Comparative statistics on the four markets are given in the upper part of table 1. The New York Stock Exchange (NYSE) is the world's largest with 1490 listed firms valued at U.S.$ 1,527 billions in December 1984 and representing 40 percent of the world capitalization [5]. The London Stock Exchange (LSE) ranks third in size after the Tokyo Stock Exchange. Although 2171 firms were listed on the LSE at the end of 1984, their total market capitalization of U.S. $236 billions was about 15 percent of the NYSE capitalization at the same date and only 6 percent of the world capitalization. The Paris Bourse is relatively small. With 504 listed firms valued at U.S. $41 billions in December 1984, it represented 1 percent of the world capitalization. The Brussels bourse with 197 listed firms worth U.S.$ 12 billions is one of the world's smallest exchange amounting to about one third of 1 percent of the world capitalization.

The French and Belgium markets are not only small in absolute terms but also relative to their respective country's Gross National Product (GNP). While the market values of NYSE and LSE stocks amount to 43 percent of their respective country's GNP, this proportion drops to 16 percent for the Brussels Bourse and to 8 percent for the Paris Bourse. The three European markets are not only smaller than the NYSE in size, they are also considerably less active as indicated by the ratio of the volume of transactions to market capitalization. On the NYSE it is 50 percent. It drops to 25 percent on the Paris Bourse, to 22 percent on the Brussels Bourse and to 20 percent on the LSE [6].

3.2 Characteristics of the data

A summary of the data and returns characteristics is given in table 1.

TABLE 1

Description of the markets, the sample characteristics and the methology employed

COUNTRY	UNITED STATES	UNITED KINGDOM	FRANCE (FR)	BELGIUM (BR)
• **MARKET CHARACTERISTICS**				
Exchange	New York Stock Exchange	London Stock Exchange	Paris Bourse	Brussels Bourse
Market capitalization (12/84)	\$1,529,459 millions	\$236,403 millions	\$41,058 millions	\$12,342 millions
Market capitalization / World Capitalization	40%	6%	1%	0.32%
Market capitalization / Gross Nat'l. Product	43%	43%	8%	16%
Volume of transactions / Market capitalization	50%	20%	25%	22%
Number of listed firms	1490	2171	504	197
• **DATA CHARACTERISTICS**				
Data source	CRSP Tape	London Stock Price Data Base	Collected by authors	Collected by authors
Data begins on	January 1969	January 1969	January 1969	January 1969
Data ends on	December 1983	December 1983	December 1983	December 1983
Number of common stocks	782	527	112	170
Market index (including dividend yields)	Equally weighted (from the CRSP Tape)	Equally weighted (using the 527 stocks)	Equally weighted (using the 112 stocks)	Equally weighted (using the 170 stocks)
• **RETURN CHARACTERISTICS**				
Length	Monthly	Monthly	Monthly	Monthly
Number	180 months	180 months	180 months	180 months
Definition	Log of price relatives adjusted for dividend	Log of price relatives adjusted for dividends	Log of price relatives adjusted for dividend	Log of price relatives adjusted for dividend
• **FAMA-MACBETH METHODOLOGY**				
Number of portfolios	20	20	20	20
Number of stocks per portfolio[a]	40,39,...,39,40	30,26,...,26,29	6,5,...,5,6	13,8,...,8,13
Length of port. construction period[b]	12 months	12 months	12 months	12 months
Length of risk-estimation period[c]	12 months	12 months	12 months	12 months
Length of testing period[d]	12 months	12 months	12 months	12 months
Updating of port. construction period	Every year	Every year	Every year	Every year
Updating of risk-estimation period	Every year	Every year	Every year	Every year
Total number of estimated monthly parameters($\gamma_0, \gamma_1, \gamma_2$ and γ_3)	156 (from Jan. 71 to Dec. 83)	156 (from Jan. 71 to Dec. 83)	156 (from Jan. 71 to Dec. 83)	156 (from Jan. 71 to Dec. 83)

a. The middle 18 portfolios have the same size.
b. Portfolios are formed with stocks ranked according to their beta values.
c. Estimation of systematic risk (beta coefficient), total risk (variance) and unsystematic risk (standard error of residual returns).
d. Cross-sectional regressions are run for every month of the year.

The sample contains 180 monthly returns for each one of 1,591 common stocks; 782 from the NYSE, 527 from the LSE, 112 from the Paris Bourse and 170 from the Brussels Bourse [7]. The time series of monthly stock returns begins on January 1969 and ends on December 1983. The source of the data for U.S. common stocks is the tape of the Center for Research in Security Prices (CRSP tape). For U.K. common stocks we used the London Stock Price Data base (see Smithers (1977) and (1980)). Prices and dividend payments for French and Belgian stocks were collected by the authors.

The proxy for the domestic market portfolios is an equally weighted market index [8]. For U.S. common stocks the index is taken from the CRSP tape. For the other 3 markets it is computed using the stocks in the sample of securities of each country. Returns are calculated by taking the logarithm of stock price relatives adjusted for dividend payments [9].

4. Methodology

The methodology employed in this study to estimate monthly risk premia is similar to that of Fama and MacBeth (1973). A summary of the procedure we followed is given at the bottom of table 1.

The first year of monthly data (1969) is used to construct 20 equally weighted domestic portfolios ranked according to the magnitude of their beta coefficients. This was done by applying the single-index market model (Sharpe (1964), Fama (1976)) to individual securities in order to estimate the beta coefficient of each stock against its domestic market index using 12 monthly returns. For each market we ranked the individual stocks in decreasing order of the magnitude of their estimated beta coefficients and formed 20 equally-weighted domestic portfolios. The first contains the stocks with the highest betas and the twentieth the stocks with the lowest betas. The exact number of stocks in each portfolios is indicated in table 1.

The second year of monthly data (1970) is used to estimate the risk of each stock in the sample [10]. Three measures of risk are considered in this study: systematic risk (beta coefficient), total risk (variance of total returns) and unsystematic risk (standard error of the residuals of the single-index market model). Portfolios' betas (β_p) are obtained by calculating the arithmetic average of the betas of the individual stocks that make up the portfolio. The same procedure is used to calculate portfolio's total risk or variance (V_p) and portfolios' unsystematic risk (S_p). The former (V_p) is the arithmetic average of the variances of the returns of the individual securities that make up the portfolios and the latter (S_p) is the arithmetic average of the standard errors of the single-index market model residuals of the individual securities that make up the portfolios.

Finally the third year of data (1971) is used to estimate the monthly market risk-premia according to the following set of regressions:

$$R_{pt} = \gamma_{ot} + \gamma_{1t} \cdot \beta_{p,t-1} + \mu_{pt}, \tag{1}$$

$$R_{pt} = \gamma'_{ot} + \gamma_{2t} \cdot V_{p,t-1} + \mu'_{pt}, \tag{2}$$

$$R_{pt} = \gamma''_{ot} + \gamma_{3t} \cdot S_{p,t-1} + \mu''_{pt}, \tag{3}$$

$$R_{pt} = \gamma^{*}_{ot} + \gamma^{*}_{1t} \cdot \beta_{p,t-1} + \gamma^{*}_{3t} \cdot S_{p,t-1} + \mu^{*}_{pt}, \tag{4}$$

in which:

R_{pt} = realized return of portfolio p in month t,

γ_{1t} = systematic risk premia in month t,

γ_{2t} = total risk premia in month t,

γ_{3t} = unsystematic risk premia in month t,

$\beta_{p,t-1}$ = beta of portfolio p estimated over a 12-month estimation period ending on the calendar year preceding month t and updated yearly,[11]

$V_{p,t-1}$ = variance of portfolio p estimated over a 12 month estimation period ending on the calendar year preceding month t and updated yearly,

$S_{p,t-1}$ = unsystematic risk of portfolio p estimated over a 12-month estimation period ending on the calendar year preceding month t and updated yearly.

Regression (1) is the standard two-parameter capital asset pricing model (Sharpe (1964), Black (1973) and Fama (1976)) according to which security holders are compensated only for bearing systematic risk. Regressions (2) and (3) are run on the theory that investors in small markets may not be fully diversified. In this case both total risk (variance) and unsystematic risk may contribute to the pricing of securities (Levy (1978)).

Regression (4) is the Fama and MacBeth (1973) version of a four-parameter capital asset pricing model. With this specification we test for the linearity of the relationship between average returns and systematic risk (if the relationship is linear γ_{1t}^{**} should be statistically equal to zero) and for the marginal contribution of unsystematic risk to the pricing of securities (if unsystematic is diversified away it is not priced in the market and γ_{3t}^{*} should be statistically equal to zero).

For each month of the 12-month test period we calculate the realized return of each one of the 20 domestic portfolios (R_{pt}). These 20 portfolio returns are then cross-sectionally regressed on beta (regression (1)), variance (regression (2)), unsystematic risk (regression (3)) as well as on beta, beta squared and unsystematic risk (regression (4)). Recall that these measures of risk are estimated over the preceding 12-month risk-estimation period and are updated every year. From the 12 cross-sectional regressions we obtain 12 monthly estimates of the market risk premia γ_1, γ_2, γ_3, γ_1^{*}, γ_1^{**} and γ_3^{*}.

The entire procedure is then repeated using the second year of monthly data (1970) to construct portfolios, the third year of monthly data (1971) to estimate risk and the fourth year of monthly data (1972) to estimate the monthly relationship between realized returns and risk. Dropping one year of early data and adding a new one to estimate the risk premia we kept on repeating the entire procedure until we reached the year 1983. This approach provides a total of 156 monthly estimates of the risk premia γ_1, γ_2 and γ_3; 13 estimates for each of the 12 months of the year (from January 1971 to December 1983).

5. Seasonality in systematic risk premia (regression (1))

The results of regression (1) are summarized in table 2 for the four markets. To save space we present only the results for the systematic risk premia (slope coefficient γ_1) and not those for the intercepts[12] γ_0. Seasonality is examined by calculating the average estimated risk premia

TABLE 2 : Market price of systematic risk (beta)

Average values of the estimated slope coefficient of the two-parameter model:

$$R_{pt} = \gamma_{0t} + \gamma_{1t} \cdot \beta_{p,t-1} + \mu_{pt}$$

(Estimated with monthly data from Jan. 1971 to Dec. 1983 - See table 1)[a]

Average of γ_1 over	U.S.	U.K.	FRANCE	BELGIUM	Sample size
All months	0.0032 0.62	-0.0048 -1.25	$\boxed{-0.0069}$ [b] $\boxed{-1.93}$	-0.0022 -1.37	156
January	$\boxed{0.0632}$ [b] $\boxed{2.46}$	0.0091 0.40	0.0193 1.21	$\boxed{0.0162}$ [b] $\boxed{2.63}$	13
February	0.0115 0.68	-0.0038 -0.35	-0.0068 -0.61	0.0003 0.05	13
March	0.0227 9.95	-0.0099 -0.68	-0.0157 -1.01	-0.0037 -0.48	13
April	0.0089 0.65	$\boxed{0.0220}$ [b] $\boxed{1.83}$	-0.0026 -0.22	-0.0026 -0.53	13
May	-0.0085 -0.63	$\boxed{-0.0211}$ [b] $\boxed{-3.33}$	$\boxed{-0.0194}$ [b] $\boxed{-1.87}$	-0.0032 -0.95	13
June	-0.0160 -1.29	$\boxed{-0.0287}$ [b] $\boxed{-1.96}$	-0.0179 -1.64	$\boxed{-0.0140}$ [b] $\boxed{-2.68}$	13
July	0.0096 0.52	0.0013 0.13	-0.0057 -0.53	0.0030 0.56	13
August	-0.0070 -0.45	0.0116 1.25	-0.0017 -0.17	$\boxed{-0.0104}$ [b] $\boxed{-2.23}$	13
September	$\boxed{-0.0167}$ [b] $\boxed{-1.88}$	-0.0088 -0.66	0.0056 0.39	$\boxed{-0.0089}$ [b] $\boxed{-2.13}$	13
October	-0.0362 -1.68	-0.0159 -1.39	-0.0067 -0.45	-0.0106 -1.75	13
November	0.0133 0.90	-0.0086 -0.68	-0.0101 -0.98	0.0040 1.02	13
December	-0.0067 -0.57	-0.0045 -0.35	$\boxed{-0.0212}$ [b] $\boxed{-1.88}$	0.0031 0.51	13
K-W Statistic[c]	18.18	14.77	6.01	20.75	
Probability	0.08	0.19[c]	0.87[c]	0.04	

a. t statistics are below the average values of the estimated coefficients.

b. significant average values at the 0.05 level.

c. Kruskal-Wallis nonparametric test statistic. The critical value at the 0.10 level for this statistic is equal to 17.27. If the calculated statistic is smaller than the critical value of 17.27 we reject the hypothesis of equal coefficients.

corresponding to each one of the 12 months of the year. Thus the 13 estimates of γ_1 for the month of January (there are 13 Januaries between January 1970 and December 1983), are averaged out and reported in table 2. Below that average is a t-statistic[13] which indicates if the average risk premium for that month is significantly different from zero at the 0.05 level. This is the case whenever the value of the t-statistic equals or exceeds 1.78. In table 2, significant estimates of the risk premia are framed. A double frame identifies positive risk premia and a single frame, negative risk premia.

Consider first the results over the entire 13-year period. The average estimated risk premia are not different from zero in the United States, the United Kingdom and Belgium. It is significantly negative in France. This means that over a period as long as 13 years holders of stocks trading on the NYSE, the LSE and the Brussels Bourse were not compensated with higher average returns for bearing higher levels of systematic risk. Investors holding stocks trading on the Paris Bourse were actually penalized rather than rewarded. They realized below average rates of return on investments with above average levels of systematic risk. Clearly, the absence of a positive relationship between average returns and systematic risk in the United States in the seventies, reported by Tinic and West (1984), is not a characteristic unique to U.S. common stocks. It is also observed in the United Kingdom and Belgium, a result that may be loosely interpreted as mild evidence of international integration among these markets[14].

We now turn to the behavior of the average systematic risk premia over each one of the 12 months of the year. The fact that the relationship between average returns and estimated systematic risk is flat over the entire 13-year period in the United States, the United Kingdom and Belgium (and negative in France) does not rule out the possibility of a significantly positive relationship between these variables over particular months of the year.

A look at table 2 shows that the average systematic risk premium is significantly positive <u>only in January</u> in the United States (6.32 percent)[15] and Belgium (1.62 percent). In the United Kingdom, however, it is significantly positive <u>only</u> in April (2.20 percent). In France it is never significantly positive although the average risk premium in January is equal to 1.93 percent. There are, however, in all four countries, a number of months during which the average systematic risk premium was significantly negative. In the United Kingdom the average risk premium is negative in May (-2.11 percent) and June (-2.87 percent). The negative contributions of these two months is sufficient to offset the positive contribution of the month of April resulting in an insignifiant relationship between average returns and systematic risk over the entire 13-year period. The same phenomenon occurs in Belgium where the negative contributions of June (-1.40 percent), August (-1.04 percent) and September (-.89 percent) offset the positive contribution of the month of January. In France, the January risk premium is equal to 1.93 percent but it is not strong enough to offset the negative contributions of May (-1.94 percent) and December (-2.12 percent) thus yielding a significantly negative relationship between average returns and systematic risk over the entire 13-year period (-.69 percent).

At the bottom of table 2 we report the Kruskall-Wallis (K-W) Statistic (Conover (1971)) which we use to test the hypothesis that the month-to-month systematic risk premia are equal. The K-W test statistic is a nonparametric test which is based on the ranking of the twelve monthly observations. It is approximately distributed as a chi-square distribution with 11 degrees of freedom. The probability values given at the bottom of table 2 indicate that the null hypothesis of equal risk-premia is rejected for the United Kingdom and France at 0.10 level. At this level of significance the critical value of the chi-square distribution is equal to 17.27.

We also tested the following two hypotheses: (1) that in the United States, France and Belgium the systematic risk premium is different in January than during the rest of the year, and (2) that in the United Kingdom the systematic risk premium is different in April than during the rest of the year. For the sake of brevity we do not report the details of these tests in this article [16]. Using an F-test, we rejected both hypotheses: the average systematic risk premium is different in January than during the rest of year in the United States, France and Belgium; and it is different in April than during the rest of the year in the United Kingdom, a country in which the systematic risk premium does not exhibit a January effect.

6. Seasonality in the total risk premium (regression (2))

The results of regression (2) are summarized in table 3 for the four markets. Again, to save space, we present only the results for the total risk premia (slope coefficient γ_2) and not those for the intercepts [17] (γ_0').

Some interesting results emerge from table 3. First, over the entire 13-year period, there is no significant relationship between average returns and total risk, even in the case of France, the only country where we observed a significant (but negative) relationship between average returns and systematic risk (see table 2). There is, however, a strong January effect in the total risk premium of U.S., U.K. and Belgian common stock. In France, although we can reject the hypothesis that month-to-month total risk premia are equal (see the K-W test at the bottom of table 3), none of these risk premia is significantly different from zero. Finally, the April effect observed in the systematic risk premium of U.K. common stocks has now been replaced by a January effect in the total risk premium.

Again, we tested the following hypothesis: that in the four countries the total risk premium is different in January than during the rest of year. Using an F-test we rejected this hypothesis in the case of the United States, the United Kingdom and Belgium but could not do so in the case of France. The details of this hypothesis testing are not reported here [18].

Finally, note that Belgium is the only country where the total risk premium is significantly different from zero in months other than January: the total risk premium is significantly negative in June (-1.2853) and September (-1.6820). It is worth pointing out that the systematic risk premium is also significantly negative during these two months (see table 2).

7. Seasonality in the unsystematic risk premia (regression (3))

The result of regression (3) are presented in table 4. According to the capital asset pricing model the estimated unsystematic risk premia should not be significantly different from zero. A look at the results for

TABLE 3 : Market price of total risk (variance)

Average values of the estimated slope coefficient of the two-parameter model:

$$R_{pt} = \gamma'_{0t} + \gamma_{2t} \cdot V_{p,t-1} + \mu'_{p,t-1}$$

(Estimated with monthly data from Jan. 1971 to Dec. 1983 - See table 1)[a]

Average of γ_2 over	U.S.	U.K.	FRANCE	BELGIUM	Sample size
All months	0.0199 0.05	0.1357 0.77	0.4966 -1.63	-0.1396 -0.64	156
January	3.3528 [b] 2.29	1.5761 [b] 2.46	0.9384 0.83	1.5919 [b] 2.09	13
February	0.1272 0.12	0.1518 0.30	0.1367 0.12	-0.0051 -0.01	13
March	1.4424 0.99	-0.3002 -0.55	-1.1580 -1.04	0.4887 0.60	13
April	0.2486 0.28	0.4069 0.81	-0.3926 -0.37	-1.5055 -1.56	13
May	-0.3258 -0.27	-0.5414 -0.77	-1.4378 -1.63	-0.5114 -0.99	13
June	-1.2086 -1.32	-0.7391 -1.12	-0.1302 -0.19	-1.2853 [b] -3.05	13
July	1.1044 0.71	0.2139 0.38	-0.1349 -0.11	1.3210 1.33	13
August	-0.3619 -0.36	-0.4340 -1.23	-0.2536 -0.27	0.0006 0.01	13
September	-0.9346 -1.22	0.2757 0.43	0.0146 0.01	-1.6820 [b] -3.38	13
October	-2.7920 -1.36	0.3957 0.53	-1.2933 -1.04	-0.2922 -0.43	13
November	0.6047 0.46	-0.0522 -0.07	-1.1218 -1.21	1.0732 1.32	13
December	-1.0183 -1.48	0.6752 1.12	-1.1261 -1.01	-0.8698 -1.15	13
K-W Statistic[c] Probability	19.27 0.06	9.62 0.56[c]	4.36 0.96[c]	20.75 0.04	

a. t statistics are below the average values of the estimated coefficients.

b. significant average values at the 0.05 level.

c. Kruskal-Wallis nonparametric test statistic. The critical value at the 0.10 level for this statistic is equal to 17.27. If the calculated statistic is smaller than the critical value of 17.27 we reject the hypothesis of equal coefficients.

TABLE 4 : Market price of unsystematic risk

Average values of the estimated slope coefficient of the two-parameter model:

$$R_{pt} = \gamma''_{0t} + \gamma_{3t} \cdot s_{p,t-1} + \mu''_{p,t-1}$$

(Estimated with monthly data from Jan. 1971 to Dec. 1983 - See table 1)[a]

Average of γ_3 over	U.S.	U.K.	FRANCE	BELGIUM	Sample size
All months	0.0174 0.18	0.0843 1.73	-0.0891 -1.33	-0.0173 -0.42	156
January	$\boxed{1.1003}$[b] $\boxed{2.73}$	$\boxed{0.3080}$[b] $\boxed{1.79}$	0.2354 1.01	$\boxed{0.3004}$[b] $\boxed{1.79}$	13
February	0.0763 0.30	0.2158 1.44	0.0313 0.12	-0.0176 -0.12	13
March	0.4231 1.06	0.0371 0.22	-0.3678 -1.46	0.0617 0.44	13
April	0.1296 0.57	0.1110 0.74	0.0466 0.19	-0.2363 -1.24	13
May	-0.0601 -0.22	-0.0409 -0.21	-0.2498 -1.36	-0.0175 -0.15	13
June	-0.3676 -1.36	-0.0740 -0.44	0.1619 0.94	$\boxed{-0.2569}$[b] $\boxed{-2.30}$	13
July	0.1713 0.44	0.0274 0.15	-0.1811 -0.75	0.2322 1.39	13
August	-0.1338 -0.45	-0.0683 -0.50	0.0254 0.10	0.0809 0.69	13
September	-0.3596 -1.67	0.0029 0.02	-0.0450 -0.17	$\boxed{-0.1906}$[b] $\boxed{-1.98}$	13
October	-0.7315 -1.72	0.3314 1.68	-0.2359 -0.97	-0.1024 -1.001	13
November	0.2124 0.65	0.0521 0.36	-0.2255 -1.21	0.1226 1.252	13
December	-0.2519 -0.99	0.109 0.71	-0.2651 -0.96	-0.1847 -1.19	13
K-W Statistic[c] Probability	21.36 0.03	6.64 0.83[c]	6.50 0.84[c]	16.78 0.11[c]	

a. t statistics are below the average values of the estimated coefficients.

b. significant average values at the 0.05 level.

c. Kruskal-Wallis nonparametric test statistic. The critical value at the 0.10 level for this statistic is equal to 17.27. If the calculated statistic is smaller than the critical value of 17.27 we reject the hypothesis of equal coefficients.

the entire 13-year period indicates the absence of a significant relationship between average returns and unsystematic risk in the four countries. This outcome is consistent with the model's prediction.

The month-to-month analysis, however, reveals another picture. As in the case of the total risk premium (see table 3), we observe again a significantly positive unsystematic risk premium in January in the United States and the United Kingdom with no relationship between average returns and unsystematic risk during the other 11 months of the year. In France, the unsystematic risk premium is never significant. In Belgium it is significantly positive during January and significantly negative during the months of June and September.

Summing up, the results for the United States, France and Belgium are consistent regardless of the risk measure used to estimate the corresponding risk premium. In the United States and Belgium, January to the only month of the year during which the estimated risk premium is significantly positive irrespective of how we measure risk. In France there is not a single month of the year during which the risk premium is significantly different from zero irrespective of how we measure risk. Finally, in the United Kingdom we observe an April effect in the systematic risk premium which is replaced by a January effect in the total and unsystematic risk premia.

8. Seasonality in the estimated coefficients of the four-parameter model (regression (4))

Results for the estimated coefficients of the four-parameter model are found in table 5 for the United States, in table 6 for the United Kingdom, in table 7 for France and in table 8 for Belgium.

The average estimated slope coefficients ($\bar{\gamma}_1$, $\bar{\gamma}_1^*$ and $\bar{\gamma}_3$) over the 13-year period are never significantly different from zero in the four countries. Our results for the United States are consistent with those of Tinic ad West (1986) [19].

We turn first to the month-to-month results for the United States given in table 5. We note that the U.S. systematic risk premium is never

TABLE 5 : United States

Average values of the estimated coefficients of the four-parameter model:

$$R_{pt} = \gamma^*_{0t} + \gamma_{1t}\cdot\beta_{p,t-1} + \gamma^*_{1t}\cdot\beta^2_{p,t-1} + \gamma_{3t}\cdot s_{p,t-1} + \mu^*_{p,t-1}$$

(Estimated with monthly data from Jan. 1971 to Dec. 1983 - See table 1)[a]

Average over	$\bar{\gamma}_0^*$	$\bar{\gamma}_1$	$\bar{\gamma}_1^*$	$\bar{\gamma}_3$	Sample size
All months	0.003 0.39	0.011 1.145	-0.004 -0.93	-0.006 -0.08	156
January	-0.024 -0.87	0.008 0.27	0.013 1.15	0.671[b] 1.80	13
February	-0.013 -0.63	0.000 0.00	-0.000 -0.00	0.177 0.88	13
March	-0.009 -0.27	0.003 0.83	0.004 0.35	0.164 0.57	13
April	0.009 0.30	-0.038 -0.81	0.013 0.65	0.429[b] 2.18	13
May	-0.014 -0.86	0.034 1.15	-0.022 -1.55	0.061 0.22	13
June	0.021 1.30	0.015 0.69	-0.011 -1.17	-0.193 -0.77	13
July	-0.024 -1.26	0.048 1.31	-0.018 -1.30	0.035 0.13	13
August	0.008 0.28	0.028 0.80	-0.009 -0.67	-0.193 -0.69	13
September	0.026[b] 2.28	0.011 0.45	-0.004 -0.45	-0.548[b] -2.40	13
October	0.037[b] 2.08	-0.003 -0.11	-0.003 -0.37	-0.611 -1.40	13
November	-0.017 -0.93	0.037 0.87	-0.016 -0.93	0.205 1.17	13
December	0.031 1.47	-0.010 -0.30	0.009 0.58	-0.272 -0.95	13
K-W Statistic[c] Probability	14.62 0.20[c]	2.62 0.99[c]	4.45 0.96[c]	20.48 0.04	

a. t statistics are below the average values of the estimated coefficients.

b. significant average values at the 0.05 level.

c. Kruskal-Wallis nonparametric test statistic. The critical value at the 0.10 level for this statistic is equal to 17.27. If the calculated statistic is smaller than the critical value of 17.27 we reject the hypothesis of equal coefficients.

TABLE 6 : United Kingdom

Average values of the estimated coefficients of the four-parameter model:

$$R_{pt} = \gamma^*_{0t} + \gamma_{1t}\cdot\beta_{p,t-1} + \gamma^*_{1t}\cdot\beta^2_{p,t-1} + \gamma_{3t}\cdot s_{p,t-1} + \mu^*_{p,t-1}$$

(Estimated with monthly data from Jan. 1971 to Dec. 1983 - See table 1)[a]

Average over	$\bar{\gamma}_0^*$	$\bar{\gamma}_1$	$\bar{\gamma}_1^*$	$\bar{\gamma}_3$	Sample size
All months	$\boxed{0.014}$[b] $\boxed{2.34}$	-0.003 -0.35	-0.002 -0.69	0.040 0.82	156
January	0.006 0.21	0.024 0.68	-0.005 -0.66	$\boxed{0.394}$[b] $\boxed{1.79}$	13
February	0.009 0.37	0.015 0.49	-0.008 -0.71	0.140 0.85	13
March	0.030 1.41	-0.018 -0.58	-0.001 -0.14	-0.076 -0.46	13
April	0.026 1.04	0.022 0.90	-0.001 -0.06	0.084 0.44	13
May	0.028 1.12	-0.001 -0.05	$\boxed{-0.013}$[b] $\boxed{-2.42}$	-0.098 -0.51	13
June	0.031 1.38	-0.037 -0.99	-0.001 -0.08	-0.134 -0.82	13
July	0.020 1.13	$\boxed{-0.050}$[b] $\boxed{-1.84}$	$\boxed{0.029}$[b] $\boxed{2.56}$	0.026 0.16	13
August	0.011 1.11	0.004 0.14	0.002 0.13	-0.058 -0.67	13
September	0.012 0.62	-0.017 -0.53	-0.000 -0.26	-0.125 -0.75	13
October	-0.018 -0.99	0.023 0.90	$\boxed{-0.021}$[b] $\boxed{-1.86}$	0.252 1.27	13
November	0.017 0.63	-0.016 -0.51	0.002 0.24	-0.004 -0.02	13
December	-0.000 -0.01	0.015 0.17	-0.008 -0.87	0.084 0.67	13
K-W Statistic[c] Probability	6.74 0.82[c]	6.97 0.80[c]	11.57 0.40[c]	8.90 0.63[c]	

a. t statistics are below the average values of the estimated coefficients.

b. significant average values at the 0.05 level.

c. Kruskal-Wallis nonparametric test statistic. The critical value at the 0.10 level for this statistic is equal to 17.27. If the calculated statistic is smaller than the critical value of 17.27 we reject the hypothesis of equal coefficients.

TABLE 7 : France

Average values of the estimated coefficients of the four-parameter model:

$$R_{pt} = \gamma^*_{0t} + \gamma_{1t}\cdot\beta_{p,t-1} + \gamma^*_{1t}\cdot\beta^2_{p,t-1} + \gamma_{3t}\cdot s_{p,t-1} + \mu^*_{p,t-1}$$

(Estimated with monthly data from Jan. 1971 to Dec. 1983 - See table 1)[a]

Average over	$\bar{\gamma}_0^*$	$\bar{\gamma}_1$	$\bar{\gamma}_1^*$	$\bar{\gamma}_3$	Sample size
All months	0.008 1.13	0.002 0.13	-0.003 -0.27	-0.000 -0.00	156
January	-0.004 -0.17	0.040 1.00	-0.011 -0.69	0.191 0.97	13
February	-0.013 -0.67	0.041 1.38	$\boxed{-0.023}$[b] $\boxed{-1.98}$	0.121 0.42	13
March	0.029 1.06	0.018 0.34	-0.016 -0.76	-0.268 -1.28	13
April	-0.030 -1.12	$\boxed{0.092}$[b] $\boxed{1.84}$	$\boxed{-0.043}$[b] $\boxed{-2.42}$	0.155 0.56	13
May	0.011 0.29	-0.032 -0.70	0.009 0.42	0.126 0.53	13
June	0.001 0.04	$\boxed{-0.063}$[b] $\boxed{-2.67}$	0.017 1.31	0.397 1.74	13
July	$\boxed{0.060}$[b] $\boxed{2.87}$	-0.014 -0.67	0.007 0.76	-0.307 -1.46	13
August	-0.005 -0.35	0.031 0.82	-0.015 -1.14	0.141 0.64	13
September	-0.012 -0.48	-0.003 -0.07	0.007 0.32	0.034 0.15	13
October	-0.010 -0.41	-0.000 -0.01	0.002 0.08	-0.154 -0.57	13
November	$\boxed{0.047}$[b] $\boxed{2.07}$	$\boxed{-0.074}$[b] $\boxed{-1.81}$	$\boxed{0.033}$[b] $\boxed{1.81}$	-0.314 -1.34	13
December	0.025 1.08	-0.017 -0.51	-0.002 -0.11	-0.126 -0.44	13
K-W Statistic[c] Probability	13.04 0.29[c]	12.52 0.33[c]	14.07 0.23[c]	9.67 0.56[c]	

a. t statistics are below the average values of the estimated coefficients.

b. significant average values at the 0.05 level.

c. Kruskal-Wallis nonparametric test statistic. The critical value at the 0.10 level for this statistic is equal to 17.27. If the calculated statistic is smaller than the critical value of 17.27 we reject the hypothesis of equal coefficients.

TABLE 8 : Belgium

Average values of the estimated coefficients of the four-parameter model:

$$R_{pt} = \gamma_{0t} + \gamma_{1t}\beta_{p,t-1} + \gamma^*_{1t}\beta^2_{p,t-1} + \gamma_{3t}s_{p,t-1} + \mu_{p,t-1}$$

(Estimated with monthly data from Jan. 1971 to Dec. 1983 - See table 1)[a]

Average over	$\bar{\gamma}_0^*$	$\bar{\gamma}_1$	$\bar{\gamma}_1^*$	$\bar{\gamma}_3$	Sample size
All months	$\boxed{0.007}$ [b] $\boxed{1.90}$	-0.002 -0.54	-0.000 -0.27	0.030 0.63	156
January	$\boxed{0.023}$ [b] $\boxed{2.07}$	0.006 0.52	0.005 0.79	0.097 0.45	13
February	0.019 1.16	-0.001 -0.12	-0.000 -0.11	-0.029 -0.15	13
March	-0.003 -0.40	-0.002 -0.20	-0.002 -0.76	0.174 1.47	13
April	0.020 1.18	0.018 1.25	-0.009 -1.31	-0.074 -0.37	13
May	-0.002 -0.20	-0.003 -0.23	0.002 0.44	-0.002 -0.11	13
June	$\boxed{0.029}$ [b] $\boxed{3.10}$	-0.014 -1.01	-0.000 -0.09	-0.102 -0.66	13
July	0.010 0.81	-0.017 -1.52	$\boxed{0.009}$ [b] $\boxed{1.97}$	0.111 0.60	13
August	0.002 0.17	$\boxed{-0.016}$ [b] $\boxed{-1.88}$	0.000 0.03	$\boxed{0.266}$ [b] $\boxed{2.19}$	13
September	-0.003 -0.31	-0.001 -0.19	-0.004 -1.23	-0.065 -0.40	13
October	$\boxed{-0.014}$ [b] $\boxed{-1.87}$	-0.002 -0.18	-0.004 -0.96	0.043 0.31	13
November	-0.015 -1.72	0.004 0.39	-0.001 -0.13	0.047 0.56	13
December	0.015 0.82	0.006 0.62	-0.001 -0.15	-0.111 -0.58	13
K-W Statistic[c] Probability	17.92 0.08	8.02 0.71[c]	9.08 0.61[c]	4.44 0.96[c]	

a. t statistics are below the average values of the estimated coefficients.

b. significant average values at the 0.05 level.

c. Kruskal-Wallis nonparametric test statistic. The critical value at the 0.10 level for this statistic is equal to 17.27. If the calculated statistic is smaller than the critical value of 17.27 we reject the hypothesis of equal coefficients.

significantly different from zero. But the unsystematic risk premium is significantly positive in the months of January (0.671) and April (0.429) and significantly negative in the month of September (-0.548). The conclusion then is that when we examine both systematic and unsystematic risks simultaneously in a single regression, the January effect in the systematic risk premium disappears but that in the unsystematic risk premium remains [20].

Note that the January effect in the risk premia may arise from the well documented small firm effect [21]. We do not test this hypothesis in this paper but Tinic and West (1986) using U.S. data, have shown that the estimation of regression (4) with firm size (measured by the logarithm of the average equity capitalization) as an additional regression does not change the results obtained without a firm-size variable. Using their own words, "simply put, then, including firm-size in the regression model does not alter our general conclusions about the risk-return relationship in January."

We now turn to the results of the four-parameter model applied to the European markets. The picture for the United Kingdom is somewhat similar to that of the United States (compare table 6 to table 5). The unsystematic risk premium is significantly positive only in January. The systematic risk premium is significant in July but it has the wrong sign. In France (table 7) there is no January effect. The systematic risk premium is significantly positive only in April. Finally, the results for Belgium, given in table 8, show no January effect in this country. The unsystematic risk premium is significantly positive in August. The systematic risk premium is also significant in August but it has the wrong sign.

We can summarize this section's results as follows: tests of the four parameter specification of the capital asset pricing model indicate that the unsystematic risk premium is significantly positive and that systematic risk premium is not significant in both the United States and the United Kingdom. These results are not consistent with security pricing according to the capital asset pricing model or, alternatively, they support Roll's (1977) contention that the capital asset pricing model cannot be tested without the knowledge of the exact composition of the true market portfolio. Finally, no January effect was found in either the systematic or the unsystematic risk premium in France and Belgium.

9. Return seasonality

In this section we examine the pattern of seasonality in the month-to-month average returns of stock market indexes in the four countries in our sample. We also perform the same analysis on the twenty domestic portfolios of stocks in these countries. The purpose is to compare the pattern of seasonality in risk premia we discovered in the previous sections to the pattern of seasonality in stock returns and to find out whether these two phenomena are related. As we pointed out in the introductory section, it is quite possible that the risk-premia seasonality found in the previous sections is nothing more than an indirect manifestation of stock returns seasonality.

Indeed, consider the following. Suppose that portfolio average returns are equal to r during each month of the year except for month t during which they are equal to $r + e_t$ where e_t represents the difference in return between month t and all other months. Similarly suppose that portfolio risk (any measure of risk) is equal to x during each month of the year except for month t during which it is equal to $x + v_t$ where v_t represents the difference in risk between month t and all other months. We can write[22]:

$$R_t = r + e_t \tag{5}$$

and

$$X_t = x + v_t, \tag{6}$$

where R_t are portfolios' average returns and X_t their corresponding measure of risk.

The estimated risk premium given by the linear regression of X_t on R_t is:

$$\gamma_t = \frac{\text{Cov}\ (R_t, X_t)}{\text{Var}\ (X_t)}. \tag{7}$$

Given equations (5) and (6), the estimated risk premium γ_t can be rewritten as:

$$\gamma_t = \frac{Cov(r,x) + Cov(e_t,x)}{Var(x)} \cdot \frac{Var(x)}{Var(X_t)} + \frac{Cov(r,v_t) + Cov(e,v_t)}{Var(x) + Var(v_t) + 2Cov(x,v_t)} \cdot \qquad (8)$$

If there is no seasonality in the risk measure then X = x for all t and the estimated risk premium becomes:

$$\gamma_t = \frac{Cov\ (r,x)}{Var\ (x)} + \frac{Cov\ (e_t,x)}{Var\ (x)} \qquad (9)$$

$$\gamma_t = \Phi + Cov\ (e_t,x)\ /Var(x).$$

It follows that in the absence of seasonality in the risk measure (X=x) the pattern of return seasonality (e_t different from zero) may induce a pattern of risk premium seasonality (γ_t different from Φ) over the same months of the year. Hence, in this case, we should observe identical patterns of seasonality in stocks returns and in risk premia.

Month-to-month mean stock market returns and Kruskal-Wallis tests of equality of mean returns for the stock market indexes of the four countries in our sample are given in table 9. When all 168 months from January 1970 to december 1983 are considered the mean stock market monthly returns are 1.09 percent in New York, 1.13 percent in London, 0.78 percent in Paris and 0.80 percent in Brussels. Note, however, that January mean returns are 5.08 percent in New York, 5.49 percent in London, 4.10 perçent in Paris and 3.99 percent in Brussels. They are all significantly higher in January than during the rest of the year. This is the well documented January effect in mean returns (see references listed in footnote 1)[23]. Interestingly, January is the only month of the year with significantly positive returns only in the United States. In the United Kingdom we have an April effect (4.19 percent)[24], in France a July effect (3.92 percent)[25] and in Belgium a February effect (1.86 percent), an April effect (1.93 percent), a June effect (1.06 percent), a July effect (1.76 percent) and a negative October effect (-1.48 percent).

The results in table 9 are for the mean returns of the market as a whole. What is the pattern of seasonality at the level of portfolio mean returns? In table 10 we give the number of portfolios (out of 20) that have significant mean returns during specific months of the year. Again, January

TABLE 9

Month-to-month mean stock market returns and Kruskal-Wallis tests of equality of mean returns from January 1970 to December 1983; equally weighted indices for the New York, London, Paris and Brussels Stock Exchanges. T-statistics in parentheses.

Average return over	Sample size	STOCK EXCHANGES			
		NYSE	LSE	PSE	BSE
January	14	0.0508 [a]	0.0549 [a]	0.0410 [a]	0.0399 [a]
		(2.10)	(2.04)	(2.40)	(4.08)
February	14	0.0078	0.0221	0.0041	0.0186 [a]
		(0.74)	(1.58)	(0.36)	(2.73)
March	14	0.0151	0.0073	0.0159	0.0037
		(0.94)	(0.42)	(0.75)	(0.49)
April	14	0.0057	0.0419 [a]	0.0171	0.0193 [a]
		(0.36)	(3.31)	(1.19)	(2.71)
May	14	-0.0070	-0.0048	-0.0069	-0.0008
		(-0.50)	(-0.41)	(-0.45)	(-0.13)
June	14	0.0070	-0.0139	-0.0156	0.0106 [a]
		(0.59)	(-0.96)	(-1.11)	(2.06)
July	14	0.0092	0.0122	0.0392 [a]	0.0176 [a]
		(0.67)	(1.18)	(2.76)	(2.98)
August	14	0.0106	0.0113	0.0196	0.0028
		(0.66)	(0.83)	(1.52)	(0.36)
September	14	0.0031	-0.0104	-0.0040	-0.0094
		(0.23)	(-0.52)	(-0.21)	(-1.17)
October	14	-0.0091	-0.0007	-0.0176	-0.0148 [a]
		(-0.42)	(-0.06)	(-1.24)	(-2.48)
November	14	0.0225	-0.0006	-0.0027	-0.0070
		(1.22)	(-0.029)	(-0.22)	(-0.98)
December	14	0.0148	0.0162	0.0037	0.0158
		(1.129)	(1.06)	(0.41)	(1.58)
All months <u>except</u> January	154	0.0072	0.0073	0.0048	0.0051
		(1.61)	(1.63)	(1.07)	(2.25)
All months	168	0.0109	0.0113	0.0078	0.0080
		(2.33)	(2.37)	(1.78)	(3.41)
K-W test: statistics[b]		8.45	16.16	17.53	33.93
K-W test: probability		0.6728	0.1352	0.0932	0.0004

a. Significant at the 0.05 level.

b. Kruskal-Wallis nonparametric test statistic. The critical value at the 0.10 level for this statistic is equal to 17.27. If the calculated statistic is smaller than the critical value of 17.27 we reject the hypothesis of equal coefficients.

TABLE 10

Number of portfolios out of 20 with statistically significant mean returns at the 0.05 level (a t statistics equal or higher than 1.78) in specific months of the year (from 1970 to 1983). Negative returns are indicated with the number of portfolios in parentheses. All other returns are positive.

Month of the year	Sample size	US	UK	FR	BE
January	14	10	17	12	19
February	14	-	3	(1)[a]	10
March	14	-	-	-	1
April	14	-	20[b]	1	10
May	14	-	-	-	(2)[a]
June	14	-	-	2	3
July	14	-	-	16	9
August	14	-	-	7	1
September	14	-	-	(2)[a]	(6)[a]
October	14	-	-	(6)[a]	(9)[a]
November	14	-	-	-	(8)[a]
December	14	-	-	1	3

a. Mean returns are negative.

b. All 20 mean returns have t statistics higher than 2.2, that is, they are significant at the 0.025 level.

emerges as a month where mean returns are different from those during the rest of the year. Note, however, that the January effect in mean returns does not occur in all 20 portfolios. For example, only 10 portfolios exhibit a January effect in mean returns in the United States [26]. Note, also, the strength of the April effect in the United Kingdom. All 20 U.K. portfolios have mean returns that are highly positively significant that month of the year.

Is there a perfect correspondence between the seasonal pattern of mean returns and the seasonal pattern of risk premia? In the case of the United States the answer is yes. January emerges in a consistent manner as the only month of the year during which both risk premia (regardless of how risk is measured) and mean returns are significantly different from zero. But this perfect correspondence does not exist in European equity markets, particularly in the case of the United Kingdom. There is an April effect and no January effect in the systematic risk premium of U.K. common stocks. There is a January effect and no April effect in the total and the unsystematic risk premia of U.K. common stocks. And there are both a January effect and an April effect in the mean returns of U.K. Common Stocks.

We showed that as long as the risk measure does not exhibit seasonality we should expect identical patterns of seasonality in mean returns and risk premia. This is what we found in the United States but not in the United Kingdom. The next logical step is to examine the monthly behavior of beta coefficients to see if they exhibit any seasonality.

10. Does beta exhibit any seasonality?

Our results are summarized in table 11. We run the single-index market model for each of the 20 domestic portfolios in each one of the four countries and for each months of the year (14 observations per regression covering the period January 1970 to December 1983). To save space we only report the estimated betas of 5 portfolios (the first, the fifth, the tenth, the fifteenth and the twentieth).[27] All estimated betas are significantly positive except for some of the December betas in France. We

Table 11: Testing for seasonality in betas

OLS estimates of beta coefficients based on a single index market model for portfolios number 5, 10, 15 and 20 (see table 1) using 14 monthly returns observations which are all the same month of the year over the period January 1970 to December 1983.

	United States					United Kingdom					France					Belgium				
	β_1	β_5	β_{10}	β_{15}	β_{20}	β_1	β_5	β_{10}	β_{15}	β_{20}	β_1	β_5	β_{10}	β_{15}	β_{20}	β_1	β_5	β_{10}	β_{15}	β_{20}
All months [a]	1.35	.99	.78	.71	.71	1.11	1.13	1.08	1.02	.65	.98	1.10	.76	1.04	1.06	1.12	.95	1.32	.93	.84
January [b]	1.39	.94	.76	.75	.71	1.11	1.19	1.12	1.07	.55	1.07	.98	.86	1.18	1.06	1.31	.83	1.03	1.10	.79
February	1.50	.99	.64	.70	.66	1.03	1.24	1.01	.85	.54	.75	1.10	.48	1.11	1.02	1.08	.97	1.02	1.00	.86
March	1.27	.99	.67	.57	.56	1.16	1.13	1.12	1.05	.80	.88	1.08	.90	.99	1.08	.90	.92	1.85	.61	.50
April	1.35	.98	.81	.73	.73	1.12	1.16	.96	1.09	.57	.95	1.04	.75	.94	1.05	1.27	1.03	1.61	1.46	.74
May	1.52	.95	.75	.80	.80	1.00	1.05	1.06	1.01	.68	1.08	1.08	.91	1.28	1.07	.86	.96	1.52	.76	.35
June	1.24	1.01	.61	.78	.74	1.19	1.23	1.05	1.02	.40	1.13	1.35	.74	.75	.82	.60	.79	1.14	1.16	1.43
July	1.41	1.09	.77	.59	.59	1.02	1.02	1.20	1.08	.84	1.00	1.24	.69	.90	1.23	1.60	.67	.84	.59	.72
August	1.27	.88	.91	.76	.72	1.13	1.04	1.07	.98	.73	1.03	1.00	.67	1.40	1.26	.90	.97	1.38	.85	.91
September	1.48	1.19	.81	.86	.80	1.13	1.12	1.10	.87	.65	.80	1.22	.61	1.21	1.05	1.08	1.10	.92	1.17	.96
October	1.20	1.04	.89	.71	.75	.80	1.08	.94	1.07	.53	1.01	1.10	.78	1.03	1.08	1.43	1.40	1.86	.99	1.06
November	1.41	.95	.90	.72	.71	1.13	1.19	1.12	1.13	.65	1.18	1.13	.67	.94	1.13	1.16	1.24	1.73	.72	.73
December	1.05	.93	.78	.65	.77	1.16	1.07	1.05	1.01	.90	.56 [c]	1.04	.18 [c]	.67 [c]	.85	1.11	.79	1.28	.83	.48
F-Statistic [d]	1.47	1.25	1.67	1.23	1.00	.66	.66	.65	1.42	1.31	.96	.31	1.30	1.34	.81	.95	.74	1.98	1.21	1.01
Probability	.15	.26	.09	.27	.43	.78	.78	.78	.17	.22	.49	.98	.23	.20	.63	.49	.70	.03	.18	.44

[a] all-month betas are estimated with 168 monthly return observations.

[b] betas for any particular month of the year are estimated with 14 monthly return observations.

[c] not significantly different from zero at the 0.05 level; all other coefficients are significant at the 0.05 level.

[d] calculated F-Statistic. The critical value of the F-Statistic at the 0.10 level is 1.70 with (11, 144) degrees of freedom. If the calculated F-Statistic is smaller than the critical value we accept the hypothesis of equal betas.

tested the hypothesis that for a given portfolio the 12 month-to-month betas are the same. Performing an F-test (Dunn and Clark (1974,p.322)) we could not reject the hypothesis that month-to-month betas are the same at the 0.10 level. At the bottom of table 11 we report the F-statistics and the corresponding probabilities. The critical value for the F statistic at the 0.10 level is 1.70. In all but 4 out of 80 cases (20 portfolios x 4 countries) the computed F-statistic is less than the critical value. Thus we accept the null hypothesis of equal betas across the 12 months of the year.

The above results are interesting but, unfortunately, they cannot explain why there is a January effect in stock returns in the United Kingdom and no January effect in the systematic risk premium. More research is needed in this area before definite conclusions can be drawn on the nature and causes of risk premia seasonality.

FOOTNOTES

1. There is a growing literature dealing with the issue of seasonality in the monthly returns of stock market indexes around the world. For the case of Canada see Berges, McConnell and Schlarbaum (1984) as well as Tinic, Barone-Adesi and West (1983). For the Australian evidence see Officer (1975) and Brown, Keim, Kleidon and Marsh (1984). For the case of Japan see Jaffe and Westerfield (1985) and Kato and Schallheim (1985). The evidence for European countries is found in Hamon (1986) for the case of France; Reinganum and Shapiro (1984), Dimson and March (1984), Beckers, Rosenberg and Rudd (1983) and Levis (1985) for the case of the United Kingdom; Wahlroos and Berglund (1983) for the case of Finland; and van den Bergh and Wessels (1985) for the case of the Netherlands.

2. According to the capital asset pricing model of Sharpe (1964) and the modified version of Black (1970), in equilibrium, the expected returns on risky assets are linearly related to systematic risk. This model has been tested with French data by Hawawini, Michel and Viallet (1983) and with Belgian data by Hawawini and Michel (1982). Both data set tend to support the model. Tests of the model with U.S. data are found in Fama and Mac Beth (1973) and Fama (1976).

3. The small-firm effect (the fact that, on average, small firms out perform large firms on a risk-adjusted bases) discovered by Banz (1981) has been shown to be related to the January effect. Keim (1983) has shown that the superior performance of small firms occurs mostly during the month of January.

4. One frequently cited explanation of the January effect is the so-called tax-loss selling hypothesis. According to that hypothesis, as the end of the fiscal year approaches (December in the United States, France and Belgium and March in the United Kingdom), investors can reduce their taxes by selling the stocks on which they lost money during the year. In doing so they realize capital losses that are deductible from their taxable income. The sale of securities at the end of the fiscal year depresses prices which recover at the beginning of the next fiscal year as stocks move back toward their equilibrium value. For countries with a fiscal year ending in December (United States, France, and Belgium) this trading implies that mean returns in December will be smaller than in the other months of the year and that mean returns in January will be significantly larger. In the United Kingdom individuals close their fiscal year at the beginning of April but corporations and partnerships may select the last day of December as the end of their fiscal year. Reinganum and Shapiro (1984) find that the behavior of British stock returns is consistent with the tax-loss selling hypothesis; and Beckers, Rosenberg and Rudd (1983) conclude that their statistical analysis of the British data provide weak support for the tax-loss selling hypothesis.

5. These figures are from the 1984 statistics published by the International Federation of Stock Exchanges (Paris). The world capitalization in December 1984 in US$ 3,805 billions.

6. The lower figure for the LSE in partly due to the fact that there is a significantly larger number of listed shares on that exchange compared to other countries' exchanges. Some of these shares are infrequently traded.

7. These are stocks for which we have the complete set of 180 monthly returns. The elimination of firms that do not have an unbroken returns record between January 1969 and December 1983 may introduce in our sample a "survivalship" bias since we only examine the behavior of firms which were successful over the entire sample period.

8. Performing our tests with value weighted indexes did change the conclusions reached with tests based on equally-weighted indexes.

9. Note that for returns less than 10 percent the logarithm of price relatives is equal to the logarithm of one plus the rate of return which is in turn approximately equal to the rate of return. We have log $(P_t + D_t/P_{t-1}) = \log (1 + R_t) \approx R_t$.

10. Twelve monthly observations may not be the "optimal" number of observations to estimate the risk of common stocks. But taking a longer risk-estimation period would have reduced the length of the testing period.

11. Risk measures are updated yearly. This means that the subscript (t - 1) in the risk measures should be interpreted as the preceeding year rather than the preceeding month.

12. Complete results are available on request from the second author. Some of these results can also be found in Corhay, Hawawini and Michel (1985).

13. This t-statistic is calculated by dividing the average value of the n estimated monthly coefficients by the ratio of their standard error to the square root of the number of observations (n).

14. It is worth recalling that the period covered by this study includes the two major oil crises of the seventy.

15. Our results are similar to those reported by Tinic and West (1984). They found an average estimated monthly risk premium of 6.09 percent in January (with a t-statistic of 2.64) for the period January 1969 to December 1982.

16. See Corhay, Hawawini and Michel (1985).

17. Complete results are available on request from the second author. Some of these results can also be found in Corhay, Hawawini and Michel (1985).

18. See Corhay, Hawawini and Michel (1985).

19. For the period January 1969 to December 1982, Tinic and West (1986) report the following results:

	$\bar{\gamma}_0$	$\bar{\gamma}_1$	$\bar{\gamma}_2$	$\bar{\gamma}_3$
January only	0.0340 (0.69)	-0.1517 (-1.74)	0.0825 (1.73)	0.8990 (1.90)
All months	0.0008 (0.09)	0.0026 (0.16)	-0.0010 (-0.13)	0.0866 (0.97)

20. Note that the systematic risk premium has the correct (positive) sign. It has the wrong (negative) sign in the regression reported by Tinic and West (1986). See footnote 19.

21. See footnote 3.

22. We dropped the subscript p for convenience.

23. Our results do not seem to confirm those reported in Gultekin and Gultekin (1983) except for the case of Belgium. They found that in France the month-to-month returns do not differ at the 10 percent level of significance and that in the United Kingdom the month-to-month returns do differ at the 10 percent level of significance. We found the opposite; in France month-to-month returns are significantly different and in the United Kingdom they are not.

24. Note that contrary to Levis (1985) we do not find a negative May effect. Such an effect could have been expected in view of the old LSE adage "Sell in May and go away". According to the Investors Chronicle (1986) there is some logic behind this adage since the market winds down for the holiday season. Parliament goes quiet and many companies with March year-ends have published their final results by mid-June. Although we do not report a significantly negative May effect in mean returns on the LSE, we do report a significantly negative May effect in this market's systematic risk minimum (-2.11 percent). See table 2.

25. The July effect may be partly attributed to the fact that, in France, roughly two thirds of all dividend payments occur in July (Hamon (1985)). Litzenberger and Ramaswamy (1979, 1982) have shown that shareholders' expected returns are positively related to dividend yields. It follows that in months where dividends are paid, average returns should be higher relative to those months in which no dividends are paid; hence the observed July effect in France.

26. These are portfolios 1, 4, 5, 7, 8, 9, 14, 16, 19 and 20.

27. Complete results are available on request from the second author.

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