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A NOTE ON THE SEASONALITY IN THE RISK-RETURN RELATIONSHIP*

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Esta nota presenta alguna evidencia de estacionalidad en la prima de riesgo estimada empíricamente para el mercado español de capitales. En particular, este trabajo muestra que el patrón de estacionalidad de la prima de riesgo coincide con el de las acciones. Sin embargo, esto sólo ocurre cuando se permite que las betas varíen temporalmente. También se constata cierta evidencia de estacionalidad en el riesgo beta.

1. Introduction

In recent years, several important papers have pointed out that the empirical evidence for a positive relationship between risk and return seems to be very weak [1]. For this reason, it becomes natural to extend the investigation in order to find potential seasonalities in the risk-return tradeoff [2].

This note presents further european evidence on the seasonality of the market risk premium. In particular, we argue that the evidence reported by Corhay, Hawawini and Michel (1987), and the results of Rubio (1988), in the sense that the risk premium seasonal is not a direct consequence of the stock-return seasonal, may have been caused by the procedure employed in the estimation of beta risk. In this paper, it is shown that once we recognize some stochastic process in beta, there is a striking correspondence between stock-return and risk premium seasonalities. In other words, the new empirical evidence seems to suggest that the reasons behind market risk premium and stock market return seasonalities are simply the same.

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[1] Classical examples are Gibbons (1982), Stambaugh (1982) and Shanken (1985).

[2] See Tinic and West (1984), and Corhay, Hawawini, and Michel (1987). For related evidence in the Spanish equity market, see Rubio (1988).

2. The Changing Coefficient Perspective

In the empirical literature on the risk-return relationship, it has been usually assumed that beta remains constant for some previous months[3]. Corhay, Hawawini and Michel employ the Fama and MacBeth (1973) methodology with betas estimated over the preceding (relative to the test period) twelve-month estimation period. Rubio, on the other hand, employs a contemporaneous perspective using estimates of the most up-to-date coefficients with 60 observations.

It should be recognized that both approaches assume that some average past beta is the appropriate measure of risk for a particular month in the sample. This implies that risk characteristics that may have information from several previous months are incorporated in the market risk premium of a single, well-defined month.

In this paper, we propose a simple time-varying coefficient methodology in order to estimate portfolio betas[4]. We model the shifts in betas in a stochastic but well-specified way. Thus, we recognize the possibility that betas may be time-dependent, so that it is allowed an estimation procedure of betas for each month in the sample. Once we have these estimates, we can run a cross-sectional regression in each month and have potentially more accurate estimates of market risk premiums.

Therefore, in this paper we do not require that betas are constant but we do assume that betas follow the stochastic process described by

$$R_t = X_t \beta_t + \varepsilon_t \; ; \quad t = 1, \ldots, T \qquad [1]$$

$$\beta_t = M \beta_{t-1} + \eta_t \; ; \; t = 1, \ldots, T \qquad [2]$$

where X_t is, in general, a row vector of K fixed explanatory variables, ε_t is normally and independently distributed with mean zero and variance σ^2, and η_t is a K-variate normal and independent with mean zero and covariance matrix V[5]. In our particular application, we make the additional assumption that $M = I$, where I is the, identity matrix[6]. The model is run for each month between January 1970 and December 1987. Consequently, we have 216 estima-

[3] An interesting exception is by Chan and Chen (1985).

[4] See Chow (1983) for a nice description of the time-varying coefficients literature.

[5] In our application X is a row vector composed of one explanatory variable and the intercept. The explanatory variable is the market return. Therefore, we just have a version of the traditional market model with time-dependent coefficients.

[6] It should be recognized that a major drawback in the paper is our assumption imposed by equation [2]. Given that portfolios are finally employed in the empirical application and taken into account that portfolio betas change much more slowly than individual betas, our assumption seems to be reasonable. A full investigation of the stochastic process of the betas is beyond the scope of this note. However, it clearly deserves further research. It should also be noted that when $V = 0$ and $M = I$, the model is reduced to the standard normal regression model.

tes of non-constant betas for each portfolio in the sample. Data from January 1963 to December 1969 are employed in order to estimate the first beta for the filtering procedure. This is just and OLS beta for each potfolio.

3. Empirical Results

The results reported in Table 1 are based on nine equally weighted industry portfolios. The first column contains the stock market returns averaged over each month from January 1970 to December 1987. The returns on all securities in the sample were used to compute an estimate of the monthly return on the market portfolio. It corresponds to a value-weighted index, where the weights are the market values of each security at the end of the preceding year. The Madrid Stock Exchange prepares a weighted index composed of, approximately, 70 stocks. This index is also rebalanced every year. It was decided to use the returns on all stocks in the sample given that, in most months, the number of stocks was well above the number of securities included in the Madrid Stock Exchange. Moreover, our index is computed with stocks from the three major stock exchanges in the country[7]. As expected, given the previous empirical evidence, both January and February have positive and significant average returns.

The second column of Table 1 reports the market risk premium estimated with non-constant betas. It should be noted the similarity between this estimate and the stock returns of the first column. Again, only January and February have positive and significant average risk premium. It is also interesting to point out that, contrary to previous evidence, the estimate of γ_1 over all months is positive and significantly different from zero. The estimate is equal to 1.41% per month.

Finally, the third column of Table 1 contains the market risk premium estimated with constant betas. The industry portfolio betas are the estimates of the most up-to-date coefficients. This implies that each beta is obtained with data from the particular month in which we want to estimate γ_1, and the previous 59 months. The results are consistent with the empirical evidence reported by Corhay, Hawawini, and Michel and Rubio. When betas are not estimated by a time-varying coefficients methodology, the monthly risk premium seasonal is not simply a reflection of the monthly return seasonal. In particular, note the risk premium of October, April and May. Moreover, γ_1 over all months is not statistically different from zero. Of course, there is a rather striking difference with the second column, where the reasons behind stock return seasonal and risk premium seasonal seem to be exactly the same.

[7] The total number of securities contained in the sample varied from 98 to 140. Data on prices for shares trading on more than one stock exchange were obtained from the exchange on which the share had highest trading volume. Moreover, given the stocks finally employed and taken into account that we use monthly observations, the potential biases from thin trading are mostly controlled.

TABLE 1
Seasonality in the empirical market risk premium and stock market returns

Market Risk Premium and Stock Market Returns Averaged Over	Stock Market Returns	Market Risk Premium with Non-Constant Betas	Market Risk Premium with Constant Betas
January	0.0477* (2.95)	0.0721* (3.53)	0.0685* (2.62)
February	0.0339* (2.98)	0.0575* (3.71)	0.0464* (2.80)
March	0.0105 (0.58)	0.0311 (1.36)	0.0181 (0.94)
April	0.0160 (1.15)	0.0379 (1.55)	0.0545 (1.75)
May	0.0044 (0.44)	−0.0043 (−0.25)	−0.0350 (−1.67)
June	0.0148 (1.04)	0.0123 (0.63)	0.0122 (0.55)
July	0.0179 (1.37)	0.0216 (1.08)	−0.0050 (−0.16)
August	0.0120 (0.99)	0.0141 (0.71)	0.0019 (0.67)
September	−0.0165 (−1.32)	−0.0279 (−1.41)	0.0018 (0.08)
October	−0.0172 (−0.83)	−0.0260 (−0.97)	−0.0539* (−2.27)
November	0.0071 (0.60)	0.0067 (0.28)	−0.0212 (−0.60)
December	−0.0005 (−0.05)	−0.0259 (−1.50)	−0.0148 (−0.74)
All months	0.0108* (2.63)	0.0141* (2.26)	0.0061 (0.81)
All months except January	0.0076 (1.82)	0.0088 (1.37)	0.0005 (0.06)

* Significant at the 0.05 level.

Monthly estimates of the market risk premium from 1970 to 1987. Estimators for each month are obtained from a regression of a 9-vector of industry portfolio returns realized at month t on a nine by two matrix of betas and ones. The returns on all securities in the sample were used to compute a value-weighted estimate of the monthly return on the market portfolio. t-values in parenthesis.

From Table 1, we also tested the hypothesis that the January market return and the January risk premium differ from the mean returns and risk premiums during the rest of the year. Both hypothesis could not be rejected:

$$\hat{\gamma}_{1t} = \underset{(2.57)}{0.0721} - \underset{(-2.16)}{0.0633}\, D_{RY} \;;\; t = 1, \ldots, 12$$

$$\bar{R}_{mt} = \underset{(3.20)}{0.0477} - \underset{(-2.58)}{0.0402}\, D_{RY} \;;\; t = 1, \ldots, 12$$

where $\hat{\gamma}_{1t}$ and $\bar{R}_{mt}$ represent the average risk premium and mean returns over the twelve months. D_{RY} is a dummy variable representing the rest of the year.

We also tested the same hypothesis for February. In this case, both hypothesis could be rejected:

$$\hat{\gamma}_{1t} = \underset{(1.87)}{0.0575} - \underset{(-1.47)}{0.0474}\, D_{RY} \;;\; t = 1, \ldots, 12$$

$$\bar{R}_{mt} = \underset{(1.92)}{0.0339} - \underset{(-1.36)}{0.0252}\, D_{RY} \;;\; t = 1, \ldots, 12$$

Note that the intercepts are the mean risk premium and mean stock returns for January and February respectively. The slope coefficients represent the difference between the rest of the year and the mean in January or February depending on the regression.

Finally, we tested the hypothesis that the month-to-month risk premium are equal over all months. We run the following regression:

$$\hat{\gamma}_{1t} = a_1 + \sum_{i=2}^{12} a_i D_{it} + e_t$$

where $t = 1, \ldots, 216$. D_{it} are dummy variables representing each month from February ($i = 2$) to December ($i = 12$). a_1 is the average risk premium in January and a_i is the difference between the average risk premium in January and the average risk premium in month i.

It was found that the January risk premium differ from the risk premium of all months except February, March and April. The tests above used non-stationary betas.

4. On the Seasonality of Beta Risk

Even if the previous evidence were not subject to measurement errors and it may be accepted as valid, we would like to have a clear explanation of both, stock return and market risk premium January seasonal. One obvious possibility could be that there exists the same seasonal pattern in beta risk.

TABLE 2
Seasonality in market betas

Industries / Months	Banks	Chemicals and Textile	Construction	Electric Utilities	Food and Beverages	Real Estate	Siderurgy, Mining and Automobile	Insurance	Communications
January	1.045	1.279	0.996	0.589	1.100	0.898	1.337	0.634	1.140
February	1.104	0.545	1.204	0.961	0.402	0.796	0.650	0.214	0.814
March	1.483	0.820	0.785	0.774	0.522	0.610	0.826	1.146	0.939
April	1.140	1.700	1.327	0.597	0.591	0.933	1.612	0.665	0.563
May	0.889	1.014	1.135	1.349	0.775	0.816	1.310	0.815	1.066
June	0.574	1.203	1.175	0.737	0.741	0.725	1.159	0.290	1.447
July	1.361	1.308	0.866	0.396	0.604	1.028	1.090	0.748	1.048
August	1.113	1.721	1.013	1.255	0.864	0.527	1.389	0.459	0.727
September	1.245	0.984	0.999	0.608	0.884	0.895	1.064	0.265	1.271
October	0.901	1.093	1.027	0.697	1.152	0.857	1.148	1.309	0.917
November	1.201	1.043	0.881	0.661	0.856	1.147	1.153	0.939	0.925
December	1.073	0.754	1.132	0.537	0.583	0.902	0.614	0.166	0.651
All Months	1.121	1.151	1.041	0.705	0.772	0.834	1.146	0.741	0.955
F-statistic (11,192)	2.780	2.877	0.763	1.837	2.014	0.818	1.503	1.760	1.571
(p-value)	(0.002)	(0.002)	(0.677)	(0.050)	(0.029)	(0.622)	(0.133)	(0.063)	(0.110)

OLS estimates of beta coefficients for 9 industries using a value-weighted market portfolio. The regressions are run with 18 monthly return observations which are all from the same month of the year from January 1970 to December 1987.

Table 2 contains the empirical results. An OLS regression with 18 observations was run for each industry and for each month of the year. The hypothesis that, for a given industry portfolio, the 12 month-to-month betas are the same was tested by an F-test. It should be pointed out that the restricted regression of this test, imposes the same slope coefficient over the twelve months allowing, at the same time, for a different intercept.

The evidence suggests that we can reject the hypothesis of equal betas across the twelve months for Banks, Chemicals and Textile, Electric Utilities and Food and Beverages. Moreover, some evidence of beta risk seasonals are also found in Siderurgy, Mining and Automobile, Insurance and Communications. The hypothesis is accepted for Construction and Real Estate.

Unfortunately, except for Chemicals and Textile, and Siderurgy, Mining and Automobile, there is no evidence of a January beta seasonal. This seems to imply that beta risk seasonality is not the reasons behind the market risk premium January seasonal. However, the empirical evidence of Table 2 provides a reasonable argument in favor of time-varying coefficient methodologies in estimating betas.

5. Conclusion

This note has investigated the monthly behavior of the CAPM-based market risk premium in the Spanish capital market. The well-known January seasonal is found. However, the main point becomes the rather strong similarity between the monthly stock return seasonal and the market risk premium seasonal once we allow for non-constant betas. Evidence of beta risk seasonality is also presented.

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Abstract

This note presents some new evidence of seasonality in the empirically estimated market risk premium in the Spanish Capital market. In particular, the paper shows that the pattern of risk premium seasonality coincides with the pattern of stock return seasonality. However, this is only the case when betas are allowed to be time-dependent. Some evidence on beta risk seasonality is also reported.

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INFORME DEL DIRECTOR

Siguiendo con nuestra pauta informativa, ofrecemos a continuación algunos datos básicos de la evolución de la revista en los años 1988 y 1989.

1. Artículos recibidos

	1988	1989
• Total de artículos recibidos	34	40
• Aceptados	24 (70,5 %)	23 (57,5 %)
• Retirados	2 (5,9 %)	2 (5 %)
• Denegados	8 (23,6 %)	9 (22,5 %)
• En proceso de evaluación	—	6 (15 %)

En 1989 se registró un incremento en la recepción de artículos con relación a 1988 de 17,6 %.

2. Artículos publicados

	1988	1989
• Bloque monográfico	—	1 (4 art.)
• Panorama	1	—
• Artículos	17	20
• Información y Documentación	1	1
• Notas	4	1
• Suplemento*	3.ª Jornadas	4.ª Jornadas

3. Duración del proceso de evaluación

Período medio transcurrido desde la recepción hasta la aceptación de los artículos aparecidos cada año.

	1988	1989
• Menos de 3 meses	2	5
• Entre 3 y 4 meses	4	4
• Entre 4 y 6 meses	8	4
• Entre 6 y 12 meses	7	11
• Más de 12 meses	2	1
• Período medio	6,1	5,8

* El suplemento recoge los resúmenes de las ponencias presentadas en las Jornadas de Economía Industrial que anualmente patrocina la Fundación Empresa Pública.

4. Evaluadores

A parte de los miembros de los Consejos de Redacción y Asesor han colaborado en el proceso de evaluación un total de 25 personas a los que la Revista quiere dejar constancia de su agradecimiento por su generosa y cualificada ayuda.

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Continuando con la aplicación del acuerdo tomado en la reunión del Consejo Asesor de 1988 referente a su renovación regular por quintas partes, se han producido los siguientes cambios en el mismo:

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Alvaro Cuervo (Universidad Complutense de Madrid).
Amadeo Petitbó (Universidad de Barcelona, Centro de Estudios y Planificación).
Clemente Polo (Universidad Autónoma de Barcelona).

Entradas

Javier Andrés (Universidad de Valencia).
Jaume García (Universidad Autónoma de Barcelona).
José Manuel González-Páramo (Universidad Complutense de Madrid).
Vicente Salas (Universidad de Zaragoza).

El Consejo de Redacción quiere dejar constancia de su agradecimiento a los miembros salientes por su desinteresada y activa participación en la marcha de la revista.