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INVESTIGATING THE PUZZLING INTERTEMPORAL MARKET RISK-RETURN RELATIONSHIP¹

This paper investigates the conflicting results documented by the existing empirical literature on the intertemporal relationship between the expected market risk premium and the conditional market variance. We show that the previous tests are biased because they use the realized market risk premium to proxy the expected market risk premium, without accounting for the negative portion of the market risk premium distribution. The empirical evidence based on a new test, allowing up and down-market volatility to have different impacts on the market risk premium, indicates a consistent and significant risk-return relationship.

The capital asset pricing model (CAPM) developed by Sharpe (1964) and Lintner (1965) implies a linear and positive relationship between the expected return on an asset and its systematic risk (beta). Theoretically, if investors are risk-averse, a positive risk-return relationship can be expected. Several empirical studies examine the intertemporal and conditional nature of this relationship but they reach different conclusions: On one hand, Bollerslev, Engle and Wooldridge (1988) and Harvey (1989) show that this relationship is significantly positive, while French, Schwert and Stambaugh (1987), Baillie and DeGennaro (1990) and Theodossiou and Lee (1995) show that this relationship is positive, but insignificant. On the other hand, Campbell (1987) and Glosten, Jagannathan and Runkle (1993) find a significantly negative relationship. All these studies analyze the intertemporal behavior of the market risk premium based on single-factor models.

More recently, Scruggs (1998) evaluates a conditional two-factor model motivated by Merton's intertemporal capital asset pricing model (ICAPM). In this model, the conditional expected market risk premium is a linear function of the conditional market variance and the conditional market covariance with a state variable to characterize the investment opportunity set.

¹ We have benefited from comment by Narjess Boubakri, Van Son Lai, To Minh-Chau, Marie-Claude Beaulieu, Jean-François L'Her and Désiré Vencatachellum. All errors remain ours.

As a state variable, Scruggs (1998) uses the excess return on a long-term government bond, which makes it possible to hedge against shifts in the investment opportunity set. This follows Merton's (1973) argument that the changing interest rate characterizes shifts in the investment opportunity set. Empirically, Scruggs (1998) shows that the conditional market variance is significantly and positively priced, while the conditional market covariance with the excess long-term government bond returns is significantly and negatively priced. According to Scruggs (1998, p.578) :

“... estimates of the simple risk-return relation obtained from conditional single-factor models may be biased downward due to the omission of an interest rate-state variable from the conditional market risk premium equation... Including the nominal one-month Treasury-bill yield in the conditional market variance equation distorts estimates of the simple risk-return relation...”.

Thus, by estimating an ad hoc model incorporating the interest rate in both conditional mean and variance equations, Scruggs (1998) restores the positive and significant relationship between the risk premium and the conditional variance.

However, we argue that including interest rate in both conditional mean and variance equations may result in considerable incoherence. Indeed, no theoretical argument justifies the incorporation of an interest rate variable into the conditional market risk premium equation. Furthermore, the presence of the interest rate in both the conditional market risk premium and the conditional market variance equations may result in multicollinearity problems. Moreover, when the interest rate is not included in the conditional mean and variance equations, the relationship remains weak or negative. Hence, we think that the issue related to the intertemporal single-factor market risk-return relationship remains unresolved.

The purpose of this paper is to reexamine the intertemporal puzzling single-factor relationship between the conditional market risk premium and the conditional market variance found in previous studies, by testing a new hypothesis. We document that the estimates of the simple market risk-return relationship may be biased downwards due to the use of the realized market risk premium to proxy for the expected market risk premium. Therefore, not accounting for the negative portion of the realized market risk premium leads to an aggregation bias resulting from the compensation effects of up and down-market returns. This result is important since it (1) restores the pertinence of the market conditional variance as measure of risk and (2) explains why some empirical studies found a negative or weak relation between volatility and expected return.

The remainder of this paper is structured as follows: The first section presents the theoretical models and empirical procedures used for our tests. The second section discusses methodology employed. The third section describes the data and presents some empirical results, and finally the fourth section concludes the paper.

I. Models and procedures

This section describes the theoretical models and empirical procedures that we use to test the different specifications of the intertemporal single-factor relationship between the conditional market risk premium and the conditional market variance. Previous tests on this issue are essentially based on the assumption that the realized market risk premium is unbiased estimate of conditional expected market risk premium. Our main contribution here is to show the impact of this assumption on testing the intertemporal risk-return relationship.

A. The theoretical model

In this paper, we based our tests on Merton's (1973) ICAPM. This model has the advantage of being flexible and testable. According to this model, the conditional expected market risk premium is given by:

$$E_{t-1}[r_{m,t}] = \left[-\frac{J_{ww}W}{J_w} \right] \sigma_{m,t}^2 + \left[-\frac{J_{wF}}{J_w} \right] \sigma_{mF,t} \quad (1)$$

where $r_{m,t}$ is the market risk premium and $E_{t-1}[\cdot]$ denotes the expectation operator conditional on information available at time $t-1$. $\sigma_{m,t}^2$ and $\sigma_{mF,t}$ are respectively the market variance and the market covariance with the state variable F conditional on information available at time $t-1$. J is the investor utility of wealth function and depends on F and wealth W . Subscripts under J denote partial derivatives with respect to the variables denoted by the subscripts. This model implies a positive and linear partial relation between the conditional expected market risk premium and the conditional market variance. In addition, when the investment opportunities are constant, equation (1) is reduced to the conditional version of the Sharpe - Lintner CAPM given by the following equation:

$$E_{t-1}[r_{m,t}] = \left[-\frac{J_{ww}W}{J_w} \right] \sigma_{m,t}^2 \quad (2)$$

Following the majority of previous studies, all our tests are based on equation (2). In this section, we present the different specifications that we test in order to verify the relationship given by equation (2).

B. The empirical procedures

First, we replicate traditional tests on Canadian data. Second, we run our empirical tests using up and down-market information decomposition.

B.1. Traditional tests. In order to test the relationship between the market conditional risk premium ($r_{m,t}$) and the conditional market variance ($\sigma_{m,t}^2$), we use the following model as derived from the previous studies².

$$\text{Model 1:} \quad r_{m,t} = \lambda_0 + \lambda_m \hat{\sigma}_{m,t}^2 + \varepsilon_{m,t} \quad (3)$$

The investors' risk-aversion hypothesis implies that the λ_m coefficient (the price of risk) is positive. Equation (3) is tested by Campbell (1987), French, Schwert and Stambaugh (1987), Glosten, Jagannathan and Runkle (1993) and Scruggs (1998).

B.2. Tests with up and down-market information decomposition. Equation (2) indicates that the conditional expected market risk premium is a positive linear function of the conditional market variance. When the expected market risk increases, the anticipated returns should adjust accordingly in order to compensate the investors for the additional risk. This relationship has crucial implications for the empirical tests because an increase in the conditional market variance must be associated with an increase in the risk premium anticipated by investors. This theoretical association led researchers to directly test for a positive relationship. However, as these tests use the realized returns rather than the expected returns, we argue that they do not explicitly test the

² We don't test a constrained version of equation (3) ($\lambda_0 = 0$) used by Merton (1980) and Harvey (1989). However, Scruggs (1998, p.589) shows that constraining the regression line to pass by the origin may result in an overestimation of the coefficient λ_m .

validity of the Merton's ICAPM. Indeed, although that there exists a positive risk-return tradeoff, there is a non-zero probability that the realized market risk premium be negative (Table 1 shows that the market risk premium is negative in 44% of the cases). Therefore, the traditional tests of the single-factor ICAPM model should account for the negative portion of the market risk premium distribution.

We present below the methodology that we use to obtain the empirical model that allows up and down-market periods volatility to have different effects on the market risk premium. Our approach is inspired by the study of Pettengill, Sundaram and Mathur (1995, p.103) on the CAPM. Indeed, according to these authors "since these tests uses realized returns instead of expected returns, we argue that the validity of the SLB (Sharpe-Lintner-Black) model is not directly examined. Indeed, recognition of a second critical relationship between the predicted market returns and the risk-free return suggests that previous tests of the relationship between beta and the returns must be modified. The need to modify previous tests results from the model's requirement that a portion of the market return distribution be below the risk-free rate".

A reasonable inference of this critical relationship may be that returns associated to high volatility are less than returns associated to low volatility when the market risk premium is negative (down-market period). To draw this inference, assume that the economy is represented by two states of the nature Ω_1 and Ω_2 characterizing up and down-market periods respectively. These two states of nature are assumed to be independent with the respective occurrence probabilities $p_1 = p(\Omega_1)$ and $p_2 = (1-p_1) = p(\Omega_2)$. In this case, we obtain the following relationship:

$$E_{t-1}[r_{m,t}] = p_1 E_{t-1}[R_{m,t} | \Omega_1] + p_2 E_{t-1}[R_{m,t} | \Omega_2] \quad (4)$$

We can then express the conditional market variance as:

$$\sigma_{m,t}^2 = p_1^2 \sigma_{mu,t}^2 + p_2^2 \sigma_{md,t}^2 \quad (5)$$

where $\sigma_{mu,t}^2$ and $\sigma_{md,t}^2$ are conditional market variances associated respectively to up and down-market periods. Using equation (5), equation (2) can be re-written as:

$$E_{t-1}[r_{m,t}] = \left(p_1^2 \left[-\frac{J_{ww} W}{J_w} \right] \right) \sigma_{mu,t}^2 + \left(p_2^2 \left[-\frac{J_{ww} W}{J_w} \right] \right) \sigma_{md,t}^2 \quad (6)$$

A testable version of the relation (6) is therefore:

$$\underline{\text{Model 2:}} \quad r_{m,t} = \lambda_0 + \lambda_{m1} \delta \hat{\sigma}_{m,t}^2 + \lambda_{m2} (1 - \delta) \hat{\sigma}_{m,t}^2 + \varepsilon_{m,t} \quad (7)$$

where $\delta = 1$ if the market-risk premium is positive (up-market) and $\delta = 0$ if the market risk premium is negative (down-market). Knowing that λ_{m1} (λ_{m2}) is estimated during up (down) market periods, the expected sign of this coefficient is positive (negative).

II. Methodology

A. The empirical model of the conditional variance

The estimation of the models previously described requires modeling the conditional volatility of the market risk premium. The ARCH models pioneered by Engle (1982) permit the measurement and prediction of the time-varying conditional volatility. Following empirical evidences on market volatility, we assume that the conditional market volatility is time-varying and follows an exponential generalized autoregressive conditional heteroskedastic (EGARCH) process, as developed by Nelson (1991). The use of EGARCH model is motivated by the fact that it captures the volatility clustering which is a characteristics of high-frequency asset returns (see Mandelbrot (1963) and Fama (1965)). Also, its formulation is well suited to accommodate asymmetric effects in the evolution of volatility process. Since most studies find that one period is enough to capture the characteristics of most financial data series, we have considered the EGARCH (1,1) model³. According to this model, the conditional market variance depends on the amplitude as well as the past innovation sign. The conditional market variance is given by:

$$\begin{aligned}\ln(\hat{\sigma}_{m,t}^2) &= \omega + \alpha g(e_{t-1}) + \beta \ln(\hat{\sigma}_{m,t-1}^2) \\ e_t &= \varepsilon_{m,t} / \hat{\sigma}_{m,t} \quad e_t \sim N(0,1) \\ g(e_{t-1}) &= \theta e_{t-1} + (|e_{t-1}| - \sqrt{2/\pi})\end{aligned}\tag{8}$$

What distinguishes the EGARCH model is that it is suited to accommodate asymmetric effects since it incorporates the news response function $g(e_{t-1})$ with coefficient α and allows to measure the sign and the past innovation amplitude. The coefficient θ measures the asymmetry of the response of the conditional market variance to signs of past return shocks. A negative (positive) coefficient θ implies that negative (positive) return shocks have more impact on the conditional volatility than positive (negative) return shocks of the same magnitude. When $\theta = 0$, the news response function $g(e_{t-1})$ is then symmetric and depends only on lagged returns shocks. According to Engle and Ng (1993, p.1753) “The EGARCH model differs from the standard GARCH model in two main respects: (1) the EGARCH model allows good news and bad news to have different impacts on volatility, while the standard GARCH model does not, and (2) the EGARCH model allows big news to have a greater impact on volatility than the standard GARCH model”.

B. Maximum likelihood methodology

The models are estimated by using the maximum likelihood method. The maximum likelihood estimates are obtained by searching for values of parameters that maximize the likelihood function (L), calculated from the products of all conditional densities of the prediction errors.

$$\ln L = \sum_{t=1}^N \frac{1}{2} \left[-\ln(2\pi) - \ln(\hat{\sigma}_{m,t}^2) - \frac{\varepsilon_{m,t}}{\hat{\sigma}_{m,t}} \right]\tag{9}$$

The likelihood function is maximized by using the dual Quasi-Newton algorithm. The starting values for the regression parameters are obtained by using the ordinary least squares estimates.

³ See Bollerslev, Chou and Kroner (1992) for a survey.

III. Empirical results

A. Data description

We use the value-weighted monthly returns of all traded stocks in the Toronto Stock Exchange as a proxy for the market returns. The three-month Treasury-bill returns are used as risk-free returns. These data are from the TSE-Western file and cover the period from March 1950 to December 1995.

Table I presents the different characteristics of returns. Panel A reports the descriptive statistics. The results show that the realized risk premium is positive in only 56% of cases. The risk premium time series presents negative skewness (-0.59). Besides, the p-value associated with the Shapiro-Wilk (1965) test is 0.018 that indicates the rejection of the risk premium normality hypothesis at 5% level. Panel B reports the autocorrelation coefficients of orders 1, 2, 3, 4, 6, 9 and 12. The results indicate weak autocorrelation coefficients for the market risk premium while the autocorrelation coefficients are very strong for the risk-free returns. Besides, the amplitudes of these autocorrelation coefficients are always higher than two standard deviations. Finally, Panel C presents the McLeod and Li (1983) parametric portmanteau test (Q^2). This test is based on the squares of residuals and cover shifts of order 1, 2, 3, 4, 6, 9 and 12 of the autocorrelation function. These tests clearly indicate the presence of heteroskedasticity in the series of returns and justify the use of EGARCH models.

Table I: Descriptive statistics

This table summarizes statistics regarding the market return (R_m), the risk-free rate of return (R_f) and the market risk premium ($r_m = R_m - R_f$) from March 1950 to December 1995 (550 observations). The market return corresponds to the return of the Canadian value-weighted index. Returns on Canadian three-month Treasury-bill proxy for the risk-free rate of return. Panel A presents the usual descriptive statistics of these three variables (R_m , r_m and R_f). The last column (S-W test) provides the p-value associated with the Shapiro-Wilk (1965) normality test. The p-value indicates the error probability of rejecting the null of normality when it is true. Panel B presents the autocorrelation coefficients of orders 1, 2, 3, 4, 6, 9 and 12 for these three series. Panel C presents the parametric portmanteau (Q^2) tests of McLeod and Li (1983). This test is based on squared residuals and cover shifts of orders 1, 2, 3, 4, 6, 9 and 12 of the autocorrelation function.

Panel A: Univariate statistics							
	Mean (x 100)	Standard Deviation (x 100)	Median (x 100)	Skewness	Kurtose	Positive Values	S-W test (p-value)
R_m	0.9593	4.3350	1.0900	0.4752	2.8855	61%	0.0973
r_m	0.4425	4.3444	0.5301	0.5852	3.0675	56%	0.0188
R_f	0.5168	0.3345	0.4777	0.8730	1.1047	100%	0.0001
Panel B: Autocorrelation coefficients (ρ_j)							
	ρ_1	ρ_2	ρ_3	ρ_4	ρ_6	ρ_9	ρ_{12}
R_m	0.0708	0.0554	0.0637	0.0075	0.0074	0.0276	0.0069
r_m	0.0820	-0.0423	0.0750	0.0195	0.0169	0.0313	0.0059
R_f	0.9127 ^a	0.8874 ^a	0.8514 ^a	0.8334 ^a	0.7985 ^a	0.7880 ^a	0.7289 ^a
Panel C: Parametric portmanteau tests							
	$Q^2(1)$	$Q^2(2)$	$Q^2(3)$	$Q^2(4)$	$Q^2(6)$	$Q^2(9)$	$Q^2(12)$
R_m	3.5913 ^b	12.4645 ^c	12.8692 ^c	17.6097 ^c	21.8078 ^c	32.1836 ^c	33.076 ^c
r_m	3.4563 ^b	10.1908 ^c	10.5118 ^b	15.0729 ^c	19.3903 ^c	28.9424 ^c	29.7908 ^c
R_f	262.585 ^c	450.515 ^c	556.439 ^c	626.387 ^c	754.877 ^c	1024.9 ^c	1188.4 ^c

^a Higher than two standard deviations.

^b Indicates presence of Heteroskedasticity with p-value lower than 10%

^c Indicates presence of Heteroskedasticity with p-value lower than 1%

B. The results

B.1. The traditional one-factor risk-return relationship. Table II presents results from Model 1 estimation with a variance modeled using an EGARCH (1,1) process as described by equation (8). Model 1 is replicated in order to facilitate comparison to the model 2. However, it is comparable with ones estimated by Glosten et al. (1993) and Scruggs (1998) in the US context. The results show that the estimate of the conditional variance coefficient (λ_m) in the market risk premium equation is negative (-0.157), but insignificant. This result confirms some findings in the US context, and is compatible with the conclusion of Baillie and DeGennaro (1990, p.211) “that traditional two-parameter models relating portfolio means to variances are inappropriate and indicate the need for research into other measure of risk”. Figures 1 and 2 plot respectively Canadian monthly market risk premium and residuals from the predicted values of market risk premium based on both the structural and time-series parts of Model 1. Similarities between these two figures indicate that Model 1 fails to capture the market risk premium changes. This is confirmed by the very low coefficient of determination obtained.

Table II: Traditional test of conditional models of the market risk premium

This table presents the results on the conditional model of the relation between the market risk premium ($r_{m,t}$) and the conditional market variance ($\sigma_{m,t}^2$) using Canadian monthly data for the period of 1950:3 to 1995:12 (550 observations). Parameters are estimated within the following EGARCH (1,1) – M system

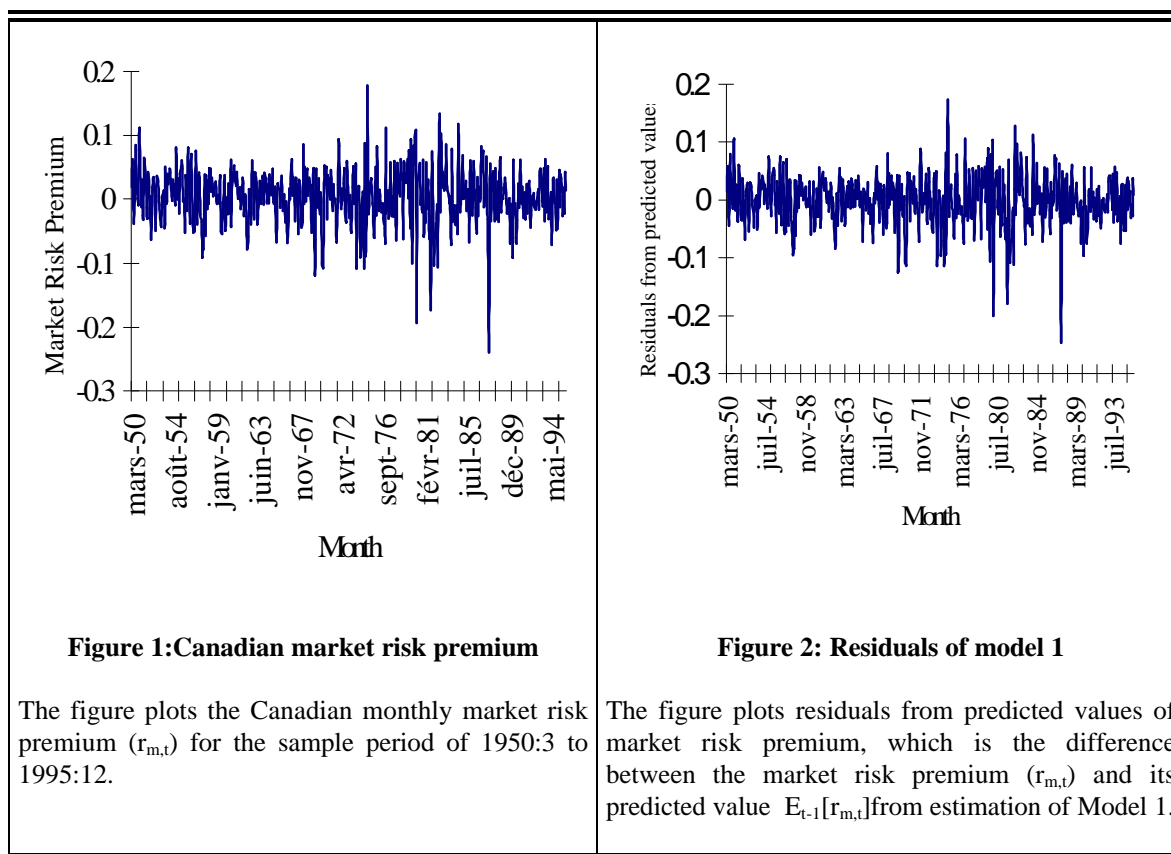
$$r_{m,t} = \lambda_0 + \lambda_m \hat{\sigma}_{m,t}^2 + \varepsilon_{m,t}$$

$$\varepsilon_{m,t} = \hat{\sigma}_{m,t} e_t \quad e_t \sim N(0,1)$$

$$\ln(\hat{\sigma}_{m,t}^2) = \omega + \alpha g(e_{t-1}) + \beta \ln(\hat{\sigma}_{m,t-1}^2)$$

$$g(e_{t-1}) = \theta e_{t-1} + (|e_{t-1}| - \sqrt{2/\pi})$$

	λ_0 (x100)	λ_m	ω	α	β	θ	ln L	R ²
Coefficient	0.5448	-0.1570	-1.3188	0.3169	0.7915	-0.3857	968.62	0.0004
t-test	1.350	-0.069	-2.713	4.248	10.372	-2.751		
p-value	0.1772	0.9446	0.0067	0.0001	0.0001	0.0059		

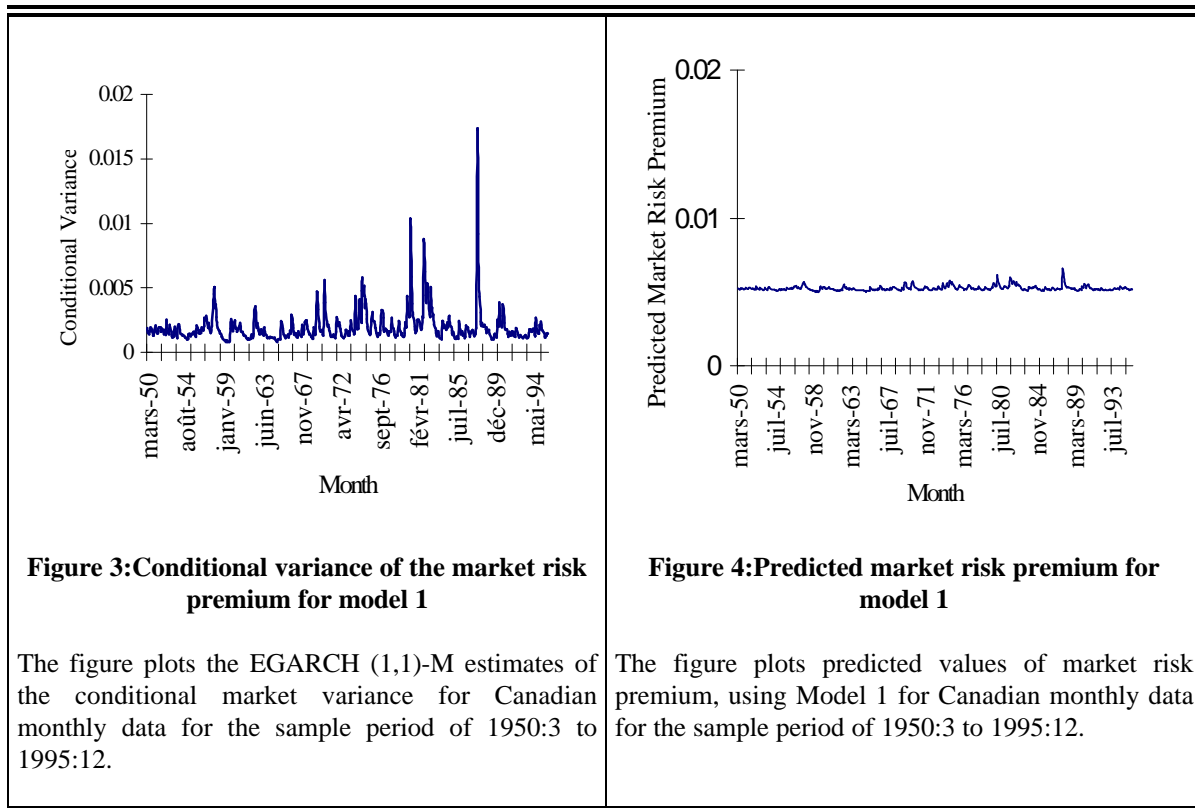


Like Scruggs (1998) and contrary to Glosten et al. (1993), the parameters in the variance equation are significant. Estimates of the coefficient of lagged conditional volatility (β) are significantly positive at 0.1% level and are associated with a half-life of 2.97 months⁴. Comparing this result with that found by Scruggs ($h = 6.43$ months) reveals that shocks are more persistent on

⁴ The half-life (h) of a market shock is given by the following expression: $h = \ln(0.5)/\ln(\beta)$, see Nelson (1991).

the American market than on the Canadian market. The impact of past return shocks measured by α is significant at 0.1% level. These two parameters (α and β) indicate a strong *hysteresis* of shocks. These results highlight also the asymmetric effect in the evolution of conditional variance. Indeed, the θ coefficient (-0.386) is significantly negative ($t = -2.751$) indicating that negative returns shocks have more impact on conditional volatility than positive returns shocks. This may be explained by leverage effect first reported by Black (1976) and theorized by Christie (1982).

Figures 3 and 4 plot respectively the EGARCH (1,1)-M estimates of the conditional market variance and the monthly-expected risk premium estimated from Model 1. It is clear from visual inspection of Figure 3 (confirmed by the statistical tests) that the market risk premium is not i. i. d through time. The plot of Figure 3 is very similar to plot of conditional market volatility presented in Scruggs (1998) and shows several distinct periods of volatility clustering. Figure 4 reveals that the predicted market risk premium does not exhibit the same patterns as the realized market risk premium (Figure 1) and exhibits a weak variation, even if it is represented to the tenth of the scale of Figure 1. Moreover, the predicted risk premium is positive throughout the sample period and consequently fails to fit the negative risk premium reported in Table I (44% of the cases). This induces a great dispersion of errors, particularly in the down-market periods, which results in a non-constant error variance (heteroskedasticity) and non-significant estimates of Model 1.



In summary, our results confirm those obtained for the American market. This is not surprising, knowing the findings of Eun and Shim (1989) that shocks in the American market are quickly disseminated to the rest of the world and of Theodossiou and Lee (1993) that the conditional volatility in the Canadian market is imported from outside, particularly from the US market. We find a negative and insignificant relationship between the Canadian market risk premium and the conditional variance when the conditional volatility is estimated with an asymmetry effect and without incorporating the risk-free return.

B.2. The segmented one-factor risk-return relationship. Table III presents the results of estimating Model 2. As we discussed previously, we transformed Model 1 in order to allow for different reactions (in sign and amplitude) depending on up and down-market periods. We modelled the volatility using an EGARCH (1,1) process.

As expected, the coefficient λ_{mu} (16.64) is significantly positive at the 0.1% level ($t = 11.04$). Thus, during up-market periods, increases in volatility results in a rise of the market risk premium. The coefficient λ_{md} (-16.01) is significantly negative at the 0.1% level ($t = -10.95$). This implies a negative relationship between the market risk premium and the conditional variance during the down-market periods. Thus, increases of the market conditional variance results in an increase of losses during the down market periods. These results go against those of traditional tests. Indeed, the amplitudes of the price of risk for up and down-market periods are very close in absolute value but in opposite sign. Consequently, we argue that aggregating these two-segmented conditional relations results in a misspecification which may explain the weakness and the absence of consistency (both economically and statistically) of traditional empirical tests.

Table III: Estimation of the conditional relation between the market risk premium and the conditional market variance at up and down-market periods

This table presents the results on the conditional relation between the market-risk premium ($r_{m,t}$) and the conditional market variance in up ($\sigma_{mu,t}^2$) and down ($\sigma_{md,t}^2$) market periods (Model 2) using Canadian monthly data for the period of 1950:3 to 1995:12 (550 observations). The dummy variable δ [$\delta = 1$ if $r_{m,t} \geq 0$, and $\delta = 0$ if $r_{m,t} < 0$] is used to separate the up and down-market patterns. Parameters are estimated within the following EGARCH (1,1) – M system

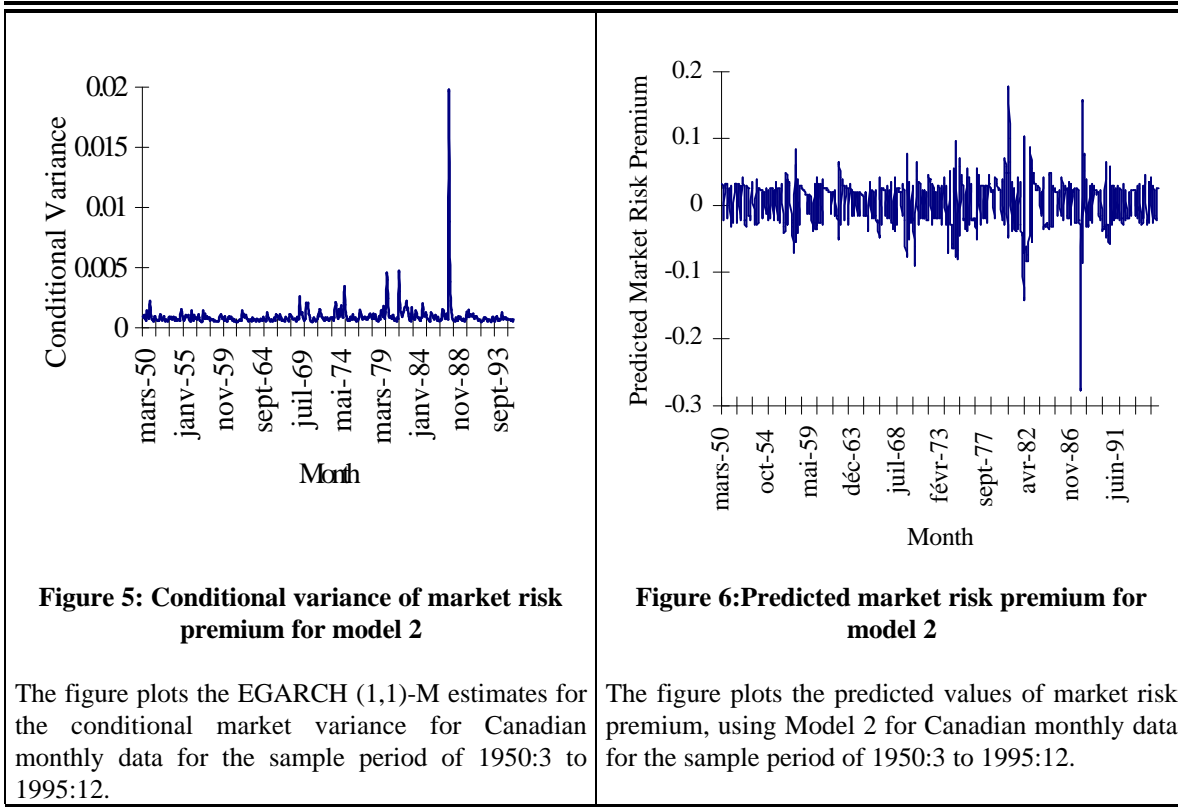
$$\begin{aligned}
 r_{m,t} &= \lambda_0 + \lambda_{mu} \hat{\sigma}_{mu,t}^2 + \lambda_{md} \hat{\sigma}_{md,t}^2 + \varepsilon_{m,t} \\
 \hat{\sigma}_{mu,t}^2 &= \delta \hat{\sigma}_{m,t}^2 = \hat{\sigma}_{m,t}^2 - \hat{\sigma}_{md,t}^2 \\
 \varepsilon_{m,t} &= \hat{\sigma}_{m,t} e_t \quad e_t \sim N(0,1) \\
 \ln(\hat{\sigma}_{m,t}^2) &= \omega + \alpha g(e_{t-1}) + \beta \ln(\hat{\sigma}_{m,t-1}^2) \\
 g(e_{t-1}) &= \theta e_t + \left(|e_t| - \sqrt{2/\pi} \right)
 \end{aligned}$$

	λ_0 (x100)	λ_{mu}	λ_{md}	ω	α	β	θ	ln L	R ²
Coefficient	0.1042	16.6398	-16.0125	-2.8886	0.4503	0.5888	0.0257	1175.47	0.4941
t-test	0.386	11.040	-10.951	-3.036	6.742	4.406	0.200		
p-value	0.6998	0.0001	0.0001	0.0024	0.0001	0.0001	0.8414		

The estimation of the conditional variance equation shows the presence of heteroskedasticity. The coefficient for past volatility shocks (α) and past conditional variance (β) are statistically significant, indicating that the volatility of the Canadian market risk premium is predictable using past information. The asymmetry of response parameter (θ) is statistically insignificant. This shows that unexpected change of the market risk premium have a symmetric impact on volatility, when the asymmetrical effect is taken into account in the mean equation (Model 2).

Figures 5 and 6 plot respectively EGARCH (1,1)-M estimates of conditional market variance and monthly-expected risk premium estimated from Model 2. The conditional volatility plot exhibits, in some cases, extreme volatility as shown in Figure 5. Figure 6 reveals that model 2 also predict down-market patterns and give a plot that exhibits a better fit of realized market risk

premium. This is confirmed by the appreciable increase of the likelihood function and the coefficient of determination of Model 2 compared to Model 1.



IV. Conclusion

This paper investigates the relation between the market risk premium and the conditional market variance and attempts to resolve the conflicting results reported by previous studies. Results from traditional test reveal a negative and insignificant relation between the market risk premium and the conditional market variance and confirm weak relations reported in the US market. However, we show that these traditional tests employed by previous studies are biased because they aggregate the risk premium associated with up and down-market periods when they use of realized market risk premium to proxy for conditional expected market risk premium.

We also conduct a new test allowing up and down-market volatility to have different effects on the market risk premium. The empirical results indicate a strong relationship between the market risk premium and the conditional market variance whatever the sign of the market risk premium. We obtain a positive (negative) and significant relationship between the market risk premium and the conditional market variance in bull (bear) market context. These results support well the Merton's ICAPM.

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