International Asset Pricing and World Market Integration : Evidence from a Partially Integrated ICAPM with Asymmetric Effects.

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Abstract

This paper tests a partially Segmented ICAPM using an asymmetric multivariate GARCH specification for two developed markets, two emerging markets and World market. We find that this asymmetric process provides a significantly better fit of the data than a standard symmetric process. The evidence supports the financial integration hypothesis and suggests that domestic risk is not a priced factor.

1 Introduction

Determining the extent to which a national market is integrated in the world stock market is an empirical question which has decisive impact on a number of issues affecting problems that are addressed by financial market theory. If capital markets are fully integrated, investors face common and country-specific risks, but price only common risk factors because country-specific risk is fully diversified. In this case, the same asset pricing relationships apply in all countries and expected returns should solely be determined by global risk factors. In contrast, when capital markets are segmented the asset pricing relationship varies across countries and expected returns would be determined by domestic risk factors. When capital markets are partially segmented, investors face both common and country-specific risks and price them both. In this case, expected returns should be determined by a combination of local and global risk sources. Thus, expected gains from world portfolio diversification and criteria for capital budgeting decisions will be quite different under local, global and mixed pricing.

Empirical papers investigating stock market integration have been mainly limited to developed markets. These papers include Bekaert and Harvey (1995), De Santis and Gerard (1997) and Carrieri, Errunza and Sarkissian (2002). The findings of these studies support the financial integration hypothesis. Recently, some papers have tented to focus on emerging markets, for instance De Santis and Imrohoroglu (1995) and Gerard, Thanyalakpark and Batten (2003). The results of these studies are heterogeneous.

On the other hand, if, as is argued in univariate and bivariate cases by Glosten, Jagannathan and Runkle (1993) and Kroner and Ng (1998), the conditional variances and covariances are higher during stock market downturns, the econometric specification should allow for asymmetric effects in variances and covariances.

The present paper contributes to stock market literature by testing a partially segmented international capital asset pricing model (ICAPM) using an asymmetric extension of the multivariate GARCH-in-Mean specification of De Santis and Gerard (1997). This approach, with sign and size asymmetric effects, allows to the prices of domestic and world market risks, betas and correlations to vary asymmetrically through time. The model is estimated over the period 1970-2003 simultaneously for 5 markets: the world market, 2 developed markets and 2 emerging markets.

The rest of the paper is organised as follows. Section 2 presents the model and introduces the econometric methodology. Section 3 describes briefly the data. Section 4 reports the empirical results. Concluding remarks are in section 5.

2 The Model and Empirical Methodology

The Capital Asset Pricing Model (CAPM) predicts that the expected excess return on an asset is proportional to its nondiversifiable risk measured by its covariance with the market portfolio. Under the hypothesises of stock market integration and purchasing power parity, an international conditional version of the CAPM can be written as:

$$E\left(\tilde{R}_{it} / \Omega_{t-1}\right) - R_{ft} = \delta_{t-1} Cov\left(\tilde{R}_{it}, \tilde{R}_{Wt} / \Omega_{t-1}\right), \quad \forall i \tag{1}$$

where $\tilde{R}_{it}$ is the return on asset i between time (t-1) and t, R_{ft} is the return on a risk-free asset and $\tilde{R}_{Wt}$ is the return on the market portfolio. δ_{t-1} is the price of world market risk and is equal to the world aggregate risk aversion coefficient, see Merton (1980) and Adler and Dumas (1983). All expectations are taken with respect to the set of information variables Ω_{t-1}.

However, many studies show that expected returns in most markets are influenced by both global and local risk factors, *i.e.* most markets are neither fully integrated nor completely segmented, see Karolyi and Stulz (2002). In this partially segmented framework, expected returns should be determined by two risk factors: global market risk and residual domestic risk, see Gerard et al. (2003):

$$E\left(\tilde{R}_{it}/\Omega_{t-1}\right)-R_{ft}=\boldsymbol{d}_{t-1}Cov\left(\tilde{R}_{it},\tilde{R}_{Wt}/\Omega_{t-1}\right)+\boldsymbol{d}_{di,t-1}Var\left(\text{Res}_{it}/\Omega_{t-1}\right),\quad \forall i \tag{2}$$

where $d_{di,t-1}$ is the price of domestic risk and ($Var\left(\text{Res}_{it}/\Omega_{t-1}\right)$) captures the domestic market nondiversifiable risk uncorrelated to world risk :

$Var\left(\text{Res}_{it}/\Omega_{t-1}\right)=Var\left(R_{it}/\Omega_{t-1}\right)-Cov\left(R_{it},R_{Wt}/\Omega_{t-1}\right)^2/Var\left(R_{Wt}/\Omega_{t-1}\right)$.

Next, consider the econometric methodology. Equation (2) has to hold for every asset including the market portfolio. A benchmark system of equations can be used to test the partially integrated conditional ICAPM. For an economy with N risky assets, the following system of pricing restrictions has to be satisfied at each point in time:

$$\tilde{R}_t-R_{ft}\boldsymbol{t}=\boldsymbol{d}_{t-1}h_{Nt}+\boldsymbol{d}_{d,t-1}*q_t+\tilde{\boldsymbol{e}}_t \qquad \tilde{\boldsymbol{e}}_t/\Omega_{t-1}\sim \mathrm{N}\left(0,\mathrm{H}_t\right) \tag{3}$$

where $q_t=D(H_t)-(h_{Nt}*h_{Nt})/h_{NNt}$, and $\tilde{R}_t$ denotes the $(N\times 1)$ vector that includes $(N-1)$ risky assets and the market portfolio, $\boldsymbol{t}$ an N-dimensional vector of ones. H_t is the $(N\times N)$ conditional covariance matrix of asset returns, h_{Nt} is the N^{th} column of H_t composed of the conditional covariance of each asset with the market portfolio and h_{NNt} the conditional variance the world market portfolio. $d_{d,t-1}$ is the $(N\times 1)$ vector of time-varying prices of domestic risk, q_t is the $(N\times 1)$ vector on nondiversifiable local risk, $D(H_t)$ the diagonal components in H_t and $(*)$ denotes the Hadamard matrix product.

The dynamics of conditional moments are left unspecified by the model. However, it has been shown that securities exhibit volatility clustering and leptokurtosis. Such characteristics are taken into account by ARCH specification. To estimate the model, we develop an asymmetric extension of the multivariate GARCH process developed by De Santis and Gerard (1997). Formally, H_t can be written as follows:

$$\mathrm{H}_t=C'C+aa'*\boldsymbol{e}_{t-1}\boldsymbol{e}'_{t-1}+bb'*\mathrm{H}_{t-1}+ss'*\boldsymbol{x}_{t-1}\boldsymbol{x}'_{t-1}+zz'*\boldsymbol{h}_{t-1}\boldsymbol{h}'_{t-1} \tag{4}$$

where $x_{it}=e_{it}I_{x_{it}}$ where $I_{x_{it}}=1$ if $e_{it}<0$ otherwise $I_{x_{it}}=0$,

$\boldsymbol{h}_{it}=\boldsymbol{e}_{it}I_{h_{it}}$ where $I_{h_{it}}=1$ if $|e_{it}|>\sqrt{h_{iit}}$ otherwise 0,

$\boldsymbol{C}$ is a $(N\times N)$ lower triangular matrix, h_{iit} is the conditional variance of asset i and ***a***, ***b, s*** and ***z*** are $(N\times 1)$ vectors of unknown parameters.

This parameterisation implies that the variances in H_t depend asymmetrically only on past squared residuals and an autoregressive component, while the covariances depend asymmetrically upon past cross-products of residuals and an autoregressive component. In particular, it guarantees that the conditional variance matrix is definite and positive. We find the symmetric GARCH process of De Santis and Gerard (1997) when $s=z=0$.

Next, turn to the price of risk. The evidence in Harvey (1991) and De Santis and Gerard (1997) suggests that the price of risk is time varying. Furthermore, Merton (1980) and Adler and Dumas (1983) show the price of world market risk to be equal to the world aggregate risk aversion coefficient. Since most investors are risk averse, the price of risk must be positive. In this paper, we follow De Santis and Gerard (1997), De Santis et al. (2003) and Gerard et al. (2003) and model the dynamics of the risk prices as a positive function of information variables: $d_{t-1} = \exp(k'_W Z_{t-1})$ and $d_{di,t-1} = \exp(k'_i Z^i_{t-1})$, where Z and Z^i are respectively a set of global and local information variables included in Ω_{t-1} and k is a set of weights that the investor uses to evaluate the conditionally expected returns.

Equations (3) and (4) constitute our benchmark model. Under the assumption of conditional normality, the log-likelihood function can be written as follows:

$$\ln L(q) = -\frac{TN}{2}\ln(2p) - \frac{1}{2}\sum_{t=1}^{T}\ln\left|H_t(q)\right| - \frac{1}{2}\sum_{t=1}^{T} e'_t(q)H_t^{-1}(q)e_t(q) \qquad (5)$$

where q is the vector of unknown parameters. To avoid incorrect inference due to the misspecification of the conditional density of asset returns the quasi-maximum likelihood (QML) approach of Bollerslev and Wooldridge (1992) is used. Simplex algorithm is used to initialize the process, then the estimation is performed using BHHH algorithm.

3 Data

The dataset includes two distinct groups of data: the returns series and the global and domestic information variables used to condition the estimation.

We use monthly returns on stock indexes for four countries plus a value weighted world market index over the period February 1970–May 2003. Given the aim of the paper, we select two large markets (the United States and the United Kingdom) and two small markets (Hong Kong and Singapore). All the indices are obtained from Morgan Stanley Capital International (MSCI) and include both capital gains and dividend yields. Returns are computed in excess of the 30-day Eurodollar deposit rate obtained from DataStream and expressed in the American dollar. Descriptive statistics for the excess returns are reported in table I.

Table I reveals a number of interesting facts. The Bera-Jarque test statistic strongly rejects the hypothesis of normally distributed returns, which supports our decision to use QML to estimate and test the model. The values of cross-correlations are relatively low. This suggests that there are still benefits from diversification across markets. The lack of autocorrelation in the return series reveals that we do not need to include an AR correction in the mean equations.

For the squared returns, autocorrelation is detected at short lags, which suggests that GARCH parameterisation for the second moments might be appropriate. Panel E of table I contains the cross-correlations of squared returns between the world and the other countries at different leads and lags. With few exceptions, only the contemporaneous correlations are statistically significant. This evidence suggests that, at least with our monthly data, the cross-market dependence in volatility is not strong and that the diagonal GARCH parameterisation for the second moments is not too restrictive.

In order to preserve the comparability between this study and others studies, the choice of global and local information variables is mainly drawn from previous empirical literature in international asset pricing, see Harvey (1991) and Bekaert and Harvey (1995). The set of global information includes a constant, the MSCI world dividend price ratio in excess of the 30-day Eurodollar deposit rate (WDY), the change in the US term premium spread measured by the yield on the ten-year US Treasury note in excess of the one-month T-Bill rate (DUSTP), the US default premium measured by the difference between Moody's Baa-rated and Aaa-rated corporate bonds (USDP) and the change on the one month Euro$ deposit rate (DWIR). The set of local information includes a constant, the local dividend price ratio in excess of the local short-term interest rate (LDY), the change in the local short-term interest rate (DLIR) and and the change in industrial production (DIP). Information variables are from MSCI, the International Financial Statistics (IFS) and DataStream and are used with one-month lag relative to the excess returns. Summary statistics for the conditioning information variables, not reported here in order to preserve space, show that the correlations among the information variables are low. Hence, our proxy of the information set contains nonredundant variables.

4 Empirical Results

We first estimate the model with the symmetric GARCH process of De Santis and Gerard (1997) and then with the asymmetric GARCH process discussed earlier in the paper. Panel A of Table II reports the results of a likelihood ratio test of the symmetric versus the asymmetric process. The test rejects the symmetric specification in favour of the asymmetric one. Similar results are given by the Akaike and Schwarz criterions presented in Panel B. Residual statistics reported in Panel C show that average mean residual is closer to zero using the asymmetric specification. To sum up, our findings show that the partially integrated ICAPM with asymmetric GARCH process fits the data better than the symmetric process of De Santis and Gerard (1997).

Table III contains parameter estimates and a number of diagnostic tests for the partially segmented conditional ICAPM with asymmetric GARCH process.

The ARCH coefficients and GARCH coefficients reported in panel B are significant for all assets. This is on line with previous results in the literature. The coefficients ***a*** are relatively small in size, which indicates that conditional volatility does not change very rapidly. However, the coefficients ***b*** are large, indicating gradual fluctuations over time. One of the advantages of our approach is to authorize for asymmetric variance and covariance effects. The significant coefficients in the vector ***s*** imply that the conditional variance is higher after negative shocks for the United States, Singapore and Hong Kong. The significant coefficients in ***s*** are all positive, which implies that conditional covariances between these countries increase after common negative shocks. In the same way, the significant coefficients in vector ***z*** indicate that the conditional variance is higher after shocks large in absolute value for the U.S. the U.K. The significant coefficients in ***z*** have the same sign (negative). This result shows that conditional covariances between these countries increase after large common negative or positive shocks.

Panel A of Table III shows the mean equation parameter estimates and Panel C reports some specification tests. The constant and the coefficients of the term premium and the default premium are significant. The robust Wald test for the significance of the time-varying parameters in the price of world market risk rejects the null hypothesis at any standard level. Figure 1 plots the estimated price of World market risk. A simple visual inspection of the chart shows that the price of market risk reaches its highest values in the Seventies and the early Eighties. Between 1994 and 2000, it becomes much lower. Finally, the price of world market risk increases significantly in the last years of our sample.

Concerning the price of local residual risk, the results show that none of the estimated coefficients are significant. The robust Wald tests confirm these results and suggest that domestic risk is not a priced factor, *i.e.* over the sample period the market considered were fully integrated. In fact, the null hypothesis that the domestic risk price coefficients are jointly equal to zero cannot be rejected at any standard level. This result is confirmed by the single country tests. To sum up, no evidence of financial segmentation is detected.

Next, we consider a number of robustness tests. To address this issue, we estimate an augmented version of the model that includes, in addition to market and domestic risk, a country specific constant and the local instrumental variables Z^i:

$$E\left(\tilde{R}_{it}/\Omega_{t-1}\right)-R_{ft}=\alpha_i+\delta_{t-1}Cov\left(\tilde{R}_{it},\tilde{R}_{Wt}/\Omega_{t-1}\right)+\delta_{di,t-1}Var\left(\mathrm{Res}_{it}/\Omega_{t-1}\right)+\phi_i' Z_{t-1}^i,\quad \forall i \tag{6}$$

The inclusion of the country-specific constants can be interpreted as a measure of mild segmentation or as an average measure of other factors that cannot be captured by the model like differential tax treatment. The inclusion of local instrumental variables can be interpreted as a way to test whether any predictability is left in the local information variables after they have been used to model the dynamics of the domestic risk prices.

The test results are reported in table IV. The Wald test indicates that the country intercepts are not jointly different from zero. On the other hand, the null hypothesis that the local information variable coefficients are jointly equal to zero cannot be rejected at any standard level.

Taken together, the findings of this paper support the financial integration hypothesis and suggest that domestic risk is not a priced factor. These results are consistent with the findings of De Santis and Gerard (1997) and Gerard et al. (2003).

5 Conclusion

In this paper, we test a partially segmented ICAPM using an asymmetric multivariate GARCH specification for two developed markets (the U.S. and the U.K.), two emerging markets (Hong Kong and Singapore) and World market over the period 1970-2003. This fully parametric empirical methodology, with sign and size asymmetric effects, allows to the prices of domestic and world market risks, betas and correlations to vary asymmetrically through time. The evidence shows that this asymmetric process provides a significantly better fit of the data than a standard symmetric process. Then, we test different pricing restrictions of the model. The evidence supports the financial integration hypothesis and indicates that domestic risk is not a priced factor.

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Table I: Descriptive statistics of asset excess returns

Panel A: Summary Statistics

	Singapore	U.K.	H. Kong	U.S.	World
Mean (% per year)	8.24	7.09	12.46	5.10	4.41
Std. Dev. (% per year)	105.63	81.33	122.49	54.73	51.26
Skewness	0.50*	1.33*	-0.29**	-0.30**	-0.47*
Kurtosis (1)	5.35*	11.37*	2.25*	1.66*	1.40*
J.B.	493.85*	2268.98*	90.20*	51.98*	47.83*
Q(12)	14.52	16.87	21.46**	9.26	13.25

Panel B: Unconditional correlations of r_{it}

	Singapore	U.K.	H. Kong	U.S.	World
Singapore	1.00	0.49	0.53	0.48	0.55
U.K.		1.00	0.38	0.53	0.69
H. Kong			1.00	0.35	0.50
U.S.				1.00	0.85
World					1.00

Panel C: Autocorrelation of (r_{it})

Lag	Singapore	U.K.	H. Kong	U.S.	World
1	0.096	0.086	0.096	0.016	0.078
2	0.015	-0.099	-0.006	-0.028	-0.048
3	-0.071	0.049	-0.041	0.023	0.032
4	0.041	0.028	-0.086	-0.028	-0.020
5	0.009	-0.117**	-0.067	0.095	0.077
6	-0.062	-0.039	-0.034	-0.043	-0.033

Panel D: Autocorrelation of $(r_{it})^2$

Lag	Singapore	U.K.	H. Kong	U.S.	World
1	0.163*	0.174*	0.027	0.110**	0.056
2	0.046	0.097	0.081	0.065	0.048
3	0.038	0.062	0.099	0.120**	0.029
4	0.091	0.038	0.119**	0.013	0.019
5	0.095	0.120**	0.075	0.006	0.071
6	0.069	0.008	0.137**	0.032	0.040

Panel E: Cross-correlations of $(r_{it})^2$ - World and Country j

Lag	Singapore	U.K.	H. Kong	U.S.
-6	-0.004	-0.021	0.001	-0.032
-5	-0.011	-0.069	0.023	0.101
-4	0.020	0.001	0.025	-0.026
-3	0.036	0.081	0.074	0.022
-2	-0.011	-0.504	-0.000	-0.045
-1	0.068	0.038	0.077	0.004
0	0.553*	0.689*	0.498*	0.851*
1	0.065	0.048	0.023	0.071
2	-0.030	-0.044	-0.049	-0.007
3	-0.032	0.031	-0.059	0.068
4	-0.029	0.008	-0.030	-0.036
5	-0.015	0.036	0.003	0.071
6	-0.114**	-0.056	-0.071	-0.058

**, ** Denote statistical significance at the 1%and 5% , (1) centred on 3.*

Table II : Asymmetric versus symmetric model

$$\tilde{R}_t - R_{ft}t = d_{t-1}h_{Nt} + d_{d,t-1} * q_t + \tilde{e}_t \qquad \tilde{e}_t / \Omega_{t-1} \sim \mathrm{N}(0, \mathrm{H_t})$$

$$d_{t-1} = \exp\left(k'_W Z_{t-1}\right) ; d_{di,t-1} = \exp\left(k'_i Z^{i}_{t-1}\right)$$

Symmetric model

$$\mathrm{H}_t = C'C + aa' * e_{t-1}e'_{t-1} + bb' * \mathrm{H}_{t-1}$$

Asymmetric model

$$\mathrm{H}_t = C'C + aa' * e_{t-1}e'_{t-1} + bb' * \mathrm{H}_{t-1} + ss' * x_{t-1}x'_{t-1} + zz' * h_{t-1}h'_{t-1}$$

$$x_{it} = e_{it} I_{x_{it}} \text{ where } I_{x_{it}} = 1 \text{ if } e_{it} < 0 \text{ otherwise } I_{x_{it}} = 0 ,$$

$$h_{it} = e_{it} I_{h_{it}} \text{ where } I_{h_{it}} = 1 \text{ if } |e_{it}| > \sqrt{h_{iit}} \text{ otherwise } 0,$$

Panel A: Likelihood ratio test

Null hypothesis	c^2	***df***	***p-value***
H₀: $s = z = 0$	26.130	10	0.003

Panel B : Information criterions

	Symmetric model	***Asymmetric model***
AIC	-11860.70	-11870.70
SBC	-11745.09	-11775.02

Panel C: Residual diagnostics

	Singapore	**U.K.**	**H. Kong**	**U.S.**	**World**
Symmetric GARCH					
Mean($\times$ *100*)	0.24	0.05	0.32	-0.04	-0.08
Skewness	0.50*	1.17*	0.29**	-0.39*	-0.41*
Kurtosis$^{(1)}$	5.33*	10.68*	4.61*	1.58*	1.18*
J.B.	489.88*	1992.03*	360.47*	52.08*	35.07*
Q(12)	13.36	18.73	13.96	9.82	13.91
Asymmetric GARCH					
Mean($\times$ *100*)	0.04	-0.00	0.05	-0.01	0.00
Skewness	0.10	0.33*	-0.24**	-0.32*	-0.41*
Kurtosis$^{(1)}$	5.05*	3.91*	2.17*	1.63*	1.17*
J.B.	425.18*	261.77*	82.77*	51.36*	34.40
Q(12)	12.11	17.76	17.38	8.47	13.43

**, ** Denote statistical significance at the 1%and 5% , (1) centred on 3.*

Table III : Quasi-maximum likelihood estimates of the partially integrated conditional ICAPM

$$\tilde{R}_t - R_{ft}\boldsymbol{t} = \boldsymbol{d}_{t-1} h_{Nt} + \boldsymbol{d}_{d,t-1} * q_t + \tilde{\boldsymbol{e}}_t \qquad \tilde{\boldsymbol{e}}_t / \Omega_{t-1} \sim \mathrm{N}(0, \mathrm{H}_t)$$

$$\boldsymbol{d}_{t-1} = \exp\left(\boldsymbol{k}_W' Z_{t-1}\right) ; \ \boldsymbol{d}_{di,t-1} = \exp\left(\boldsymbol{k}_i' Z_{t-1}^i\right)$$

$$\mathrm{H}_t = C'C + aa' * \boldsymbol{e}_{t-1}\boldsymbol{e}_{t-1}' + bb' * \mathrm{H}_{t-1} + ss' * \boldsymbol{x}_{t-1}\boldsymbol{x}_{t-1}' + zz' * \boldsymbol{h}_{t-1}\boldsymbol{h}_{t-1}'$$

$$\boldsymbol{x}_{it} = \boldsymbol{e}_{it} I_{x_{it}} \text{ where } I_{x_{it}} = 1 \text{ if } \boldsymbol{e}_{it} < 0 \text{ otherwise } I_{x_{it}} = 0,$$

$$\boldsymbol{h}_{it} = \boldsymbol{e}_{it} I_{h_{it}} \text{ where } I_{h_{it}} = 1 \text{ if } |\boldsymbol{e}_{it}| > \sqrt{h_{iit}} \text{ otherwise } 0,$$

A: parameter estimates-mean equations

(a) ***Price of world market risk***

	Const.	WDY	DUSTP	USDP	DWIR
Price of market risk	0.448*	1.581	-0.543**	0.868*	-0.641
	(0.063)	(1.684)	(0.203)	(0.290)	(0.612)

(*b*) ***Price of domestic risk***

	Const.	LDY	DLIR	DIP
Singa. Domestic Price	-0.925	-2.517	0.401	1.378
	(2.584)	(5.811)	(2.263)	(3.635)
British Domestic Price	1.561	4.361	-0.434	-4.466
	(2.413)	(4.898)	(1.903)	(8.957)
Hong K. Domestic Price	0.486	1.723	-1.577	-1.961
	(1.402)	(4.489)	(1.556)	(4.254)
American Domestic Price	-0.326	-0.578	-0.036	0.364
	(1.759)	(4.731)	(2.299)	(4.271)

Panel B: parameter estimates-Multivariate GARCH process

	Singapore	U.K.	Hong Kong	U.S.	World
a	0.325*	0.223*	0.308*	0.163*	0.260*
	(0.038)	(0.029)	(0.032)	(0.037)	(0.042)
b	0.428*	0.827*	0.908*	0.764*	0.729*
	(0.306)	(0.035)	(0.020)	(0.077)	(0.049)
s	0.140**	0.013	0.032*	0.025**	-0.009
	(0.068)	(0.015)	(0.004)	(0.009)	(0.007)
z	0.037	-0.025**	-0.011	-0.026**	0.032
	(0.050)	(0.010)	(0.010)	(0.013)	(0.020)

Panel C: Specification tests

Null hypothesis	c^2	*df*	*p-value*
Is the price of world market risk constant? $H_0: \ \boldsymbol{d}_{m,j} = 0 \quad \forall j > 1$	20.28	4	0.000
Is the price of American domestic risk equal to zero? $H_0: \ \boldsymbol{d}_{dUs,j} = 0$	0.30	4	0.989
Is the price of Singa domestic risk equal to zero? $H_0: \ \boldsymbol{d}_{dS,j} = 0$	0.37	4	0.984
Is the price of Hong K. domestic risk equal to zero? $H_0: \ \boldsymbol{d}_{dHK,j} = 0$	1.08	4	0.897
Is the price of British domestic risk equal to zero? $H_0: \ \boldsymbol{d}_{dUK,j} = 0$	1.65	4	0.797
Are the prices of domestic risk jointly equal to zero? $H_0: \ \boldsymbol{d}_{d,j} = 0 \quad \forall j,k$	3.91	16	0.999
Are the s coefficients jointly equal to zero? $H_0: \ s_i = 0 \quad \forall i$	139.68	5	0.000
Are the z coefficients jointly equal to zero? $H_0: \ z_i = 0 \quad \forall i$	67.50	5	0.000

**, ** Denote statistical significance at the 1% and 5%l, QML standard errors are reported in parentheses, (a) equal to 0 for the normal distribution. In order to preserve space, estimates of the intercept matrix C is not reported.*

Tableau IV : Robustness tests

$$E\left(\tilde{R}_{it}/\Omega_{t-1}\right)-R_{ft}=a_i+d_{t-1}Cov\left(\tilde{R}_{it},\tilde{R}_{Wt}/\Omega_{t-1}\right)+d_{di,t-1}Var\left(\mathrm{Res}_{it}/\Omega_{t-1}\right)+f_i' Z_{t-1}^{i}, \quad \forall i$$

$$d_{t-1}=\exp\left(k_W' Z_{t-1}\right)\ ;\ d_{di,t-1}=\exp\left(k_i' Z_{t-1}^{i}\right)$$

$$\mathrm{H}_t = C'C + aa' * e_{t-1}e_{t-1}' + bb' * \mathrm{H}_{t-1} + ss' * x_{t-1}x_{t-1}' + zz' * h_{t-1}h_{t-1}'$$

$$x_{it}=e_{it}I_{x_{it}} \text{ where } I_{x_{it}}=1 \text{ if } e_{it}<0 \text{ otherwise } I_{x_{it}}=0,$$

$$h_{it}=e_{it}I_{h_{it}} \text{ where } I_{h_{it}}=1 \text{ if } \left|e_{it}\right|>\sqrt{h_{iit}} \text{ otherwise } 0,$$

Null hypothesis	c^2	*df*	*p-value*
Are country-specific constants all equal to zero? H₀ : $a_i=0 \quad \forall i$	1.98	4	0.739
Are the local information variable coefficients jointly equal to zero? H₀ : $f_i=0 \quad \forall i$	12.98	12	0.370

Figure1 : World price of risk