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EXPOSURE TO THE WORLD AND TRADING-BLOC RISKS: A MULTIVIARIATE CAPITAL ASSET PRICING MODEL

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ABSTRACT

EXPOSURE TO THE WORLD AND TRADING-BLOC RISKS: A MULTIVIARIATE CAPITAL ASSET PRICING MODEL[1]

This paper employs a capital asset pricing model that incorporates both world and trading-bloc factors to show that the recent trend of trade regionalism has led to segmentation of world stock markets. The model is developed within a multivariate GARCH framework. The conditional time-varying betas are derived to examine the dynamics of risk exposures to the world and trading-bloc factors. The results show risk exposure behaviour that is not revealed using static risk estimates.

KEYWORDS: Multivariate GARCH, Regionalism, Systematic risks, Time-varying beta

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1. Introduction

A recent development has been the proliferation of regional trade agreements, leading to an increasing trend of regionalism.[2] As a result, the integration of world capital market might be affected as the formation of trading blocs not only promotes real convergence but also financial integration among member countries (Fratzscher, 2002). With the formation of trading blocs under the trade agreements, the future cash flows generated by corporations within a bloc are expected to be correlated, and thereby inviting possible significant trading-bloc effects in the pricing of financial assets within members of the same bloc (Heaney and Hooper, 1999). The new trend of regionalism promotes more than trade integration. Many further initiatives have been taken to deepen regional financial integration. Article 1109 in the North American Free Trade Agreement (NAFTA), for example, calls for free and quick transfers of all payments relating to equity transactions including dividends, interest and capital gains among the members. The December 1995 Association of Southeast Asian Nations (ASEAN) Summit endorsed in principle the concept of an investment area to lower and remove barriers to intra-regional investment among members of the ASEAN Free Trade Area (AFTA) (ASEAN Secretariat, http://www.aseansec.org/home.htm).

Recognizing that integration of the real sectors might have a crucial impact on economic policies that affect capital mobilization among member countries of a trading bloc, the pricing of financial assets might therefore be sensitive and highly exposed to the price movements within different financial markets in the bloc. Akdogan (1992), Heaney and Hooper (1999) and Adler and Qi (2003) have explored the use of trading-bloc asset pricing model to examine the effect of regionalism on pricing of financial assets. Each of these studies, however, focuses on a selected trading area. This paper aims to examine more extensively how the trading-bloc asset model can be used to explain the pricing of excess returns of the stock markets across five trading blocs that are different not only in geographical location, but also in terms of their level of integration. This cross trading-bloc analysis enables us to provide more general evidence of the existence of market segmentation due to regionalism that arises from the formation of trading blocs. The recently developed multivariate GARCH framework is employed to develop our asset pricing model. We exploit information from the multivariate framework to derive the time-varying risks for this study, which is a relatively new initiative. This approach has the advantage of revealing the dynamics of risk exposure behaviour which may not be discovered using conventional estimates that are static.

The rest of this paper is organized as follows. Section 2 provides a review of the trading-bloc asset pricing models that have been employed by other researchers. The section that follows outlines our empirical trading-bloc asset pricing model in the setting of a multivariate framework. The sample of study is also explained in this section. Section 4 presents the results and discussion on the findings. Concluding comments are given in the final section.

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[2] In spite of the attempt to promote multilateral trade negotiations, there is an accelerating increase in the number of trading blocs formally registered with WTO, see http://www.wto.org/.

2. A Review of the Trading-Bloc Asset Pricing Model

The international capital asset pricing model (ICAPM) can be written as:

$$E(R_{M,t}) - R_{F,t} = \beta_W [E(R_{W,t}) - R_{F,t}] \qquad \forall M \tag{1}$$

where $R_{M,t}$ and $R_{W,t}$ represent the continuously compounded returns for market *M* and world market, respectively, and $R_{F,t}$ is the international risk free rate at time period *t*. Model (1) assumes that the world stock market must be fully integrated such that the returns of each of the individual markets are systematically related to the world market returns. As world investors are assumed to be mean-variance optimizers and have access to full information, the expected returns of a market thus are priced on the world market portfolio. The empirical model for ICAPM can be written as:

$$r_{M,t} = \alpha + \beta_W r_{W,t} + e_{M,t} \qquad \forall M \tag{2}$$

where $r_{M,t} = R_{M,t} - R_{Ft}$ and $r_{W,t} = R_{W,t} - R_{F,t}$ are the excess returns over the risk free rate, while β_W is a constant parameter equal to $\dfrac{\text{cov}(r_{M,t}, r_{W,t})}{\text{var}(r_{W,t})}$.

To study the regionalism effect of the European Community (EC) on pricing of assets in the stock markets of its member countries, the world factor in ICAPM is replaced with a weighted EC portfolio of excess return in Akdogan (1992). His integrated one-factor pricing model takes the following form:

$$r_{M,t} = \alpha + \beta_T r_{T,t} + e_{M,t}; \quad \forall M \tag{3}$$

where $r_{T,t}$ is the EC market portfolio excess returns. In this study, we use $r_{T,t}$ to represent the excess returns of the portfolio of trading bloc *T* in general. In this case, β_T is the beta that measures systematic risk exposure to the trading bloc. If the stock markets are fully integrated within a trading bloc, the investment opportunity set available to investors would include all the stocks listed on the stock markets of its member countries. Therefore, the mean-variance efficient portfolio for the investors would be the trading-bloc market portfolio, and the model benchmarks the market excess returns to a portfolio constructed on the basis of trading bloc. The stock markets are grouped by blocs as a result of regionalism, leading to multiple "equity blocs" in the world stock exchange industry. Akdogan (1992) reports an increasing proportion of systematic risk for the period 1972-1990, implying an increasing integration among the EC markets within this period.

Heaney and Hooper (1999) tested a more general version of ICAPM given by:

$$r_{M,t} = \alpha + \beta_W r_{W,t} + \beta_T r_{T,t} + e_t; \quad \forall M \tag{4}$$

The model differs from equation (3) in that it also includes the excess returns of the world portfolio. Their study focuses on the integration of Asia Pacific Economic Cooperation (APEC), therefore a regional (APEC) factor is employed. Their findings suggest that most of the APEC markets are driven by both the APEC factor and the world factor. They also found that the ASEAN members are more integrated regionally than into the world network.

Models (3) and (4) are static, while $r_{W,t}$ and $r_{T,t}$ are assumed to be exogenous. The conditional density for $r_{M,t}$ could be utilized to allow the variance of the process to vary over time. Generally, the conditional version of equation (1) can be rewritten as:

$$E(r_{M,t} \mid \Omega_{t-1}) = \beta_W E(r_{W,t} \mid \Omega_{t-1}) \qquad \forall M \tag{5}$$

where $\beta_W = \dfrac{\operatorname{cov}(r_{M,t}, r_{W,t} \mid \Omega_{t-1})}{\operatorname{var}(r_{W,t} \mid \Omega_{t-1})}$ and Ω_t refers to the information set available at time t.

Assuming that the market price of risk is constant, equation (5) can be rewritten as:

$$E(r_{M,t} \mid \Omega_{t-1}) = \lambda_M Cov(r_{M,t}, r_{W,t} \mid \Omega_{t-1}) \qquad \forall M \tag{6}$$

where $\lambda_M = \dfrac{E(r_{W,t} \mid \Omega_{t-1})}{\operatorname{var}(r_{W,t} \mid \Omega_{t-1})}$. With this assumption, we can write the conditional pricing model as:

$$r_{M,t} = \lambda_M Cov(r_{M,t}, r_{W,t} \mid \Omega_{t-1}) + e_{M,t} \qquad \forall M \tag{7}$$

Adler and Qi (2003) examined the integration of Mexico to NAFTA for the period 1991-2002, using a conditional trading-bloc asset pricing model as discussed above, but did not assume exogeneity of the variables considered. They proposed a system of three equations for their model that consists of three variables, namely, excess returns for the Mexican market portfolio, excess returns for the North American market portfolio, and the exchange rate adjusted gross returns on a Mexican risk-free asset. This model allows for exposure to local idiosyncratic risk, trading bloc risk, and local bond risk, but does not consider world market risk. They used the non-constant parameter modeling approach, and reported that the integration process, related to both global as well as local events, is time-varying and achieved its peak at the end of the sample period.

Bekaert et al. (2005) also tested a similar version of conditional multi-equation model on emerging markets for the period 1980s to 1998, and the US portfolio is used as proxy for the computation of the world market risk. They studied the regional portfolio of Asia, Europe and Latin America, but not any trading bloc in particular. Their regional conditional pricing model accounts for local idiosyncratic, regional and world risks, as well as transmission of shocks from both regional and world factors to cater for contagion effects. Their results show that the trend is towards higher regional integration and lower world (the US) integration, especially in Asia, and to a lesser extent, in Latin and Europe.

3. Methodology and Data

3.1. The Multivariate Trading-Bloc ICAPM and Time-Varying Beta

In this paper, we extend the model of Adler and Qi (2003) by including the world factor into the pricing process, in order to compare the effects of the world and trading bloc factors on asset pricing. The proposed model allows for mild segmentation by incorporating the local idiosyncratic risk. The model, however, focuses only on equity pricing. Given a fairly large sample of stock markets covered in this study, data constraint problems make it difficult to adopt the non-constant parameter approach of Adler and Qi (2003) which requires a large set of instrumental variables.[3] We propose an alternative that exploits information from the multivariate GARCH framework to compute the time-varying risks. This approach circumvents the data problem and thereby allows a more extensive study covering 26 markets across five trading blocs (see discussion below).

Our model includes the three factors considered by Heaney and Hooper (1999) as given in equation (4). However, no exogeneity is assumed. Extending equations (5) to (7), the model is given by a system of three equations as follows:

$$r_{M,t} = \alpha_M + \lambda_M^W \operatorname{cov}(r_{M,t}, r_{W,t}) + \lambda_M^T \operatorname{cov}(r_{M,t}, r_{T,t}^O) + \lambda_M^M \operatorname{var}(r_{M,t}) + e_{M,t} \tag{8a}$$

$$r_{W,t} = \alpha_W + \lambda_W^W \operatorname{var}(r_{W,t}) + e_{W,t} \tag{8b}$$

$$r_{T,t}^O = \alpha_T + \lambda_T^T \, var(r_{T,t}^O) + e_{T,t}^O \tag{8c}$$

where $r_{T,t}^O = \operatorname{Proj}(r_{T,t} | r_{W,t}) - r_{T,t}$ is the orthogonalised trading-bloc excess returns that are uncorrelated with the world excess returns. The orthogonalised trading-bloc excess returns are obtained following the projection procedure suggested by Cochrane (2005, p.18), and the use of this series instead of the original trading-bloc excess returns has the advantage of avoiding the pitfall of multicollinearity due to the correlation between $r_{W,t}$ and $r_{T,t}$. This predicament has not been directly addressed by Heaney and Hooper (1999).

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[3] Adler and Qi (2003) included ten instrumental variables in their system of equations for the Mexican market.

Equation (8a) shows the pricing process for the market excess returns, which is an extension to model (3) but includes an additional variable for the local idiosyncratic risk. With the additional variable, the model allows for the possibility of a mildly-segmented market structure. Equation (8b) represents the pricing process for world excess returns, and a similar pricing process is used for the excess returns to the trading bloc portfolio in equation (8c).[4]

We can rewrite our system of equations as:

$$r_t = D + G(H_t) + \Xi_t \qquad (8)$$

where $r_t = (r_{M,t}, r_{W,t}, r^O_{T,t})'$, *D* and *G* represent the intercept and slope coefficients, and the error process is $\Xi_t = (e_{M,t}, e_{W,t}, e^O_{T,t})'$. Differing from Heaney and Hooper (1999), the error process Ξ_t is assumed to follow a multivariate GARCH specification. The conditional variance for the error process is:

$$\Xi_t|\Omega_{t-1} = (e_{M,t}|\Omega_{t-1}, e_{W,t}|\Omega_{t-1}, e_{T,t}|\Omega_{t-1})' \sim (0, H_t)$$

where H_t is the conditional variance under a trivariate GARCH (1,1) setting. The conditional variance-covariance matrix H_t is given as follows:

$$H_t = \begin{bmatrix} h^2_{M,t} & h_{MW,t} & h_{MT,t} \\ h_{WM,t} & h^2_{W,t} & h_{WT,t} \\ h_{TM,t} & h_{TW,t} & h^2_{T,t} \end{bmatrix}$$

The specification of the models as given in equations (8a) to (8c) implies that:

$$H_t = \begin{bmatrix} h^2_{M,t} & h_{MW,t} & h_{MT,t} \\ 0 & h^2_{W,t} & 0 \\ 0 & 0 & h^2_{T,t} \end{bmatrix}$$

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[4] The world market portfolio is priced on its own conditional variance as

$$E(r_{W,t} \mid \Omega_{t-1}) = \lambda_W Cov(r_{W,t}, r_{W,t} \mid \Omega_{t-1}) \ = \lambda_W Var(r_{W,t} \mid \Omega_{t-1}) \ .$$

The asymmetric behaviour in financial asset pricing suggests that adverse shocks (bad news) influence the volatility of the financial asset more severely than shocks favourable to the market (good news) (see Black, 1976; Christie, 1982; French et al., 1987; Nelson, 1991; Schwert, 1989, among others).[5] Following Adler and Qi (2003), we consider a multivariate specification of Zokian's (1994) threshold GARCH (TGARCH) to capture the asymmetric responses of the conditional variances to return innovations. We employ the BEKK (Baba et al., 1990) multivariate GARCH setting to ensure that the variance-covariance matrix is positive semidefinite, and the number of parameters involved is not too large, as compared to other specifications.[6] Our GARCH specification can be written as:

$$H_t = C'C + A'\Xi_{t-1}\Xi'_{t-1}A + B'H_{t-1}B + S'_{t-1}\eta_{t-1}\eta'_{t-1}S_{t-1} \tag{9}$$

where C is the 3x3 lower triangular matrix of constants, *A*, *B* and *S* are 3x3 diagonal matrices of coefficients, and $\eta_{t-1} = \Xi_{t-1}\mathrm{I}_{t-1}$, $\mathrm{I}_{t-1} = 1$ if $\Xi_{t-1} < 0$ and 0 otherwise. To ensure that the estimation is feasible, we assume only a GARCH (1,1) specification. The individual elements of H_t in equation (9) are given by:

$$h^2_{M,t} = c^2_{11} + a^2_{11}e^2_{M,t-1} + b^2_{11}h^2_{M,t-1} + s^2_{11}\eta^2_{M,t-1}$$

$$h^2_{W,t} = c^2_{12} + c^2_{22} + a^2_{22}e^2_{W,t-1} + b^2_{22}h^2_{W,t-1} + s^2_{22}\eta^2_{W,t-1}$$

$$h^2_{T,t} = c^2_{13} + c^2_{23} + c^2_{33} + a^2_{33}e^2_{T,t-1} + b^2_{33}h^2_{T,t-1} + s^2_{33}\eta^2_{T,t-1}$$

$$h_{MW,t} = h_{WM,t} = c_{11}c_{12} + a_{22}a_{11}e_{M,t-1}e_{W,t-1} + b_{22}b_{11}h_{M,t-1}h_{W,t-1} + s_{22}s_{11}\eta_{M,t-1}\eta_{W,t-1}$$

$$h_{MT,t} = h_{TM,t} = c_{11}c_{13} + a_{33}a_{11}e_{M,t-1}e_{T,t-1} + b_{33}b_{11}h_{M,t-1}h_{T,t-1} + s_{33}s_{11}\eta_{M,t-1}\eta_{T,t-1}$$

$$h_{WT,t} = h_{TW,t} = c_{13}c_{21} + c_{22}c_{23} + a_{33}a_{22}e_{W,t-1}e_{T,t-1} + b_{33}b_{22}h_{W,t-1}h_{T,t-1} + s_{33}s_{22}\eta_{W,t-1}\eta_{T,t-1}$$

The log-likelihood function for the conditional densities of the errors process is given by:

$$L(\theta) = -\frac{n}{2}\ln(2\pi) - \frac{1}{2}\ln\left|H_T\right| - \frac{1}{2}\Xi'_t H_T^{-1}\Xi_t$$

where θ denote the vector of all the parameters.

[5] A conceivable reason for the asymmetric pricing behaviour in stock returns is the so-called leverage effect, where stock prices tend to suffer a greater drop in value with the arrival of bad news, leading to an increase in the leverage ratio of a firm's capital structure. This behavior of stock returns can also be explained by the risk-premium effect that arises due to risk aversion during a market downturn.

[6] In general, there are four variants of multivariate GARCH models – VECH model of Bollerslev et al. (1988), Conditional Constant Correlation (CCC) model of Bollerslev (1990), Factor-ARCH model of Engle et al. (1990) and BEKK model of Baba et al. (1990) and Engle and Kroner (1995).

Referring to equation (4), the betas β_W and β_T measure the exposure to systematic world and trading-bloc risks, respectively. These measures, however, are static. By considering the conditional version of the ICAPM discussed in the previous section, and using the orthogonalised trading-bloc excess returns instead of the original series, the estimates of the time-varying world and trading-bloc betas are derived as follows, respectively (see Roll, 1977, p.167):

$$\hat{\beta}_{W,t} = \frac{\operatorname{cov}(r_{Mt}, r_{Wt} \mid \Omega_{t-1}) \operatorname{var}(r_{Tt}^{O} \mid \Omega_{t-1}) - \operatorname{cov}(r_{Mt}, r_{Tt}^{O} \mid \Omega_{t-1}) \operatorname{cov}(r_{Wt}, r_{Tt}^{O} \mid \Omega_{t-1})}{\operatorname{var}(r_{Wt} \mid \Omega_{t-1}) \operatorname{var}(r_{Tt}^{O} \mid \Omega_{t-1}) - \operatorname{cov}(r_{Wt}, r_{Tt}^{O} \mid \Omega_{t-1})} \tag{10a}$$

$$\hat{\beta}_{T,t} = \frac{\operatorname{cov}(r_{Mt}, r_{Tt}^{O} \mid \Omega_{t-1}) \operatorname{var}(r_{Wt} \mid \Omega_{t-1}) - \operatorname{cov}(r_{Mt}, r_{Wt} \mid \Omega_{t-1}) \operatorname{cov}(r_{Wt}, r_{Tt}^{O} \mid \Omega_{t-1})}{\operatorname{var}(r_{Wt} \mid \Omega_{t-1}) \operatorname{var}(r_{Tt}^{O} \mid \Omega_{t-1}) - \operatorname{cov}(r_{Wt}, r_{Tt}^{O} \mid \Omega_{t-1})} \tag{10b}$$

From the system of equations in model (8), we can obtain the conditional time-varying variance and covariance for all the variables in the model. The estimates of the elements in H_t are used to compute the time-varying betas according to (10a) and (10b) where $\operatorname{cov}(r_{it}, r_{jt} \mid \Omega_{t-1}) = h_{ij}$, $i,j = M, T, W$. These estimated betas measure the exposure of an individual market to the systematic risks in the world and trading bloc, respectively. The summary point estimates of the betas can be computed as the mean of the conditional time-varying betas as follows:

$$\hat{\beta}_W \mid \Omega_{t-1} = \frac{1}{n}\left(\sum_{t=1}^{n} \hat{\beta}_{W,t} \mid \Omega_{t-1}\right) \tag{11a}$$

$$\hat{\beta}_T \mid \Omega_{t-1} = \frac{1}{n}\left(\sum_{t=1}^{n} \hat{\beta}_{T,t} \mid \Omega_{t-1}\right) \tag{11b}$$

where n is the number of observations in the sample.

3.2. Sample of Study

Stock markets of member countries of five trading blocs are selected for the analysis. The trading blocs include EU (European Union), EFTA (European Free Trade Agreement), NAFTA (North American Free Trade Agreement), CER (Australia-New Zealand Closer Economic Relations), and AFTA (Association of South-East Asia Nations (ASEAN) Free Trade Area). These trading blocs have progressed quite successfully since formation, and the stock markets of the member countries are amongst the well developed markets within the region they are located. The level of economic integration in these trading blocs is different, thereby representing different degree of regionalism. EU is a monetary union, EFTA, NAFTA and CER are free trade areas; while AFTA is established on the basis of a preferential trade agreement. Interestingly, the free trade commitment in some of these trading blocs extends beyond that suggested by their setup. For example, members of EFTA and NAFTA entered an agreement on services under a General Agreement on Trade in Services (GATS) Article Five, and this represents a higher degree of integration than a conventional free trade area.

Within the five trading blocs, a total of 26 stock markets are covered in the study. Monthly data for the period from January 1991 to August 2005 are employed. All the trading agreements went into force after the beginning of this sample period. Members of AFTA officially launched its preferential arrangements in January 1992, NAFTA started its free trade agreement under GATT Art. XXIV in February 1993, while CER, EU and NAFTA launched their services agreement under GATS Art. V in November 1995, and EFTA embarked on the agreement in June 2002 (see WTO, http://www.wto.org/). While bearing in mind that each trading bloc was set up at a different date, we chose a common sample period of analysis for all the 26 markets for comparison purposes. This sample period covers major events such as the 1997 Asian financial crisis and the market crashes that occurred in the early 2000s that affected many of the Western developed countries. Also, the sample period extends to the longest possible time series available at the time of this study was conducted.

All the stock market indices are collected from Morgan Stanley Capital International (MSCI). In the computation of excess returns, the US Treasury bill rate downloaded from the website of the EconStats (www.econstats.com) is used as the proxy for the world risk free rate. The MSCI All Country World Index is used as the proxy for the world portfolio. The trading bloc portfolios are constructed through a market capitalization weighted method. The index of the market of interest is not included in the computation of the trading-bloc index to in order to exclude the local dynamics from the trading-bloc portfolio.

4. Results and Discussion

4.1. Data and Model Estimations

Table 1 presents the summary statistics for the excess return series of the 26 markets considered in this study. The mean excess returns lie in the range from -28% to 50%. Over the 15-year sample period, only the mean excess returns for Indonesia and Thailand are negative. These are the two countries most seriously affected by the 1997 Asian financial crisis. In fact, their AFTA counterparts, Malaysia and Philippines have very low positive mean excess returns, and Singapore has the best market performance in the bloc. On the other hand, the European markets have high mean excess returns, including Finland (45%), Sweden (36%) and Switzerland (37%). The only emerging Latin market in the sample, Mexico (42%), also experienced high positive returns. The larger markets, including US, Canada, UK, Germany and France have mean excess returns within the range of 16% to 28%.

Table 1 Summary statistics for excess returns

Trading Bloc	Country	Mean	Standard Deviation	Minimum	Maximum
EU	Austria	0.1545	2.4040	-8.5379	6.2758
	Belgium	0.2174	2.1381	-9.1076	7.0069
	Denmark	0.2912	2.2425	-6.3046	5.3599
	Finland	0.4546	4.3676	-16.6533	12.1393
	France	0.2539	2.2646	-7.2468	6.1720
	Germany	0.1799	2.6646	-12.1366	8.7628
	Greece	0.1012	3.8222	-11.1794	13.7255
	Ireland	0.2079	2.3456	-6.6779	7.2180
	Italy	0.1761	2.9317	-9.1317	8.3952
	Netherlands	0.2309	2.2173	-8.5375	5.2626
	Portugal	0.1201	2.7153	-9.3599	8.4140
	Spain	0.2798	2.7275	-10.7003	8.3872
	Sweden	0.3566	3.2126	-11.1013	8.8669
	UK	0.1683	1.8127	-4.8468	4.7380
EFTA	Norway	0.2115	2.9144	-14.2548	6.6382
	Switzerland	0.3653	2.0513	-7.4793	5.8115
NAFTA	Canada	0.2723	2.2749	-10.7923	5.7867
	Mexico	0.4213	4.2221	-18.2694	10.2231
	US	0.2868	1.7886	-6.6138	4.5491
CER	Australia	0.2500	2.1794	-6.4573	5.6672
	New Zealand	0.1978	2.7326	-9.9627	6.2430
AFTA	Indonesia	-0.2806	6.0947	-22.8416	19.1567
	Malaysia	0.0242	4.1229	-15.7372	17.5420
	Philippines	0.0082	4.2316	-15.1022	15.5990
	Singapore	0.1268	3.1666	-10.0515	9.8698
	Thailand	-0.0856	5.3526	-18.1316	15.5452
World		0.1939	1.7366	-6.6765	3.7524

In terms of standard deviation, the excess returns of the markets in AFTA exhibit the most volatile behaviour, with Indonesia and Thailand having the highest standard deviation of 6.0947 and 5.3526, respectively. The excess returns for members of the other trading blocs are much less volatile. Their standard deviation is in the range from 1.7 to 3.8. The two exceptions are Finland (4.3676) and Mexico (4.2221), which were both relatively far more volatile compared to their trading-bloc counterparts. The two countries with the lowest standard deviation are US (1.7886) and UK (1.8127). The excess returns of the world portfolio are very stable with a low standard deviation of 1.7366. In general, the results suggest that the large and developed markets are relatively stable compared to the small and emerging markets.

Table 2 reports the estimations for the asset pricing model with the BEKK multivariate GARCH specification as given in equations (8) and (9). The estimated coefficients for λ_M^W and λ_M^T of the mean equation provide measures of the price of the risk associated with the world and trading-bloc factors, respectively, and they represent the risk tolerance behaviour of the individual market. A few coefficients in the mean equations are statistically significant. The estimate of λ_M^W is negative and statistically significant for Belgium and US. The estimate of λ_M^T is statistically significant for Austria, Belgium, Greece and Indonesia, but the sign of the coefficients are mixed. The estimate of λ_M^M is positive and statistically significant only for US. None of the conditional variance of the world returns is statistically significant in the mean equation of the world portfolio. The mean equation of the trading-bloc portfolio for Austria, Netherlands, UK, Switzerland and all the three NAFTA members show statistically significant conditional variance of the returns. The intercepts of the three mean equations are not statistically significant, with the exception of Mexico for the market portfolio equation, and Austria, Netherlands, UK, Switzerland and all the three NAFTA members for the trading-bloc portfolio equation. Thus, the average prices of risks are generally not different from zero. This could be the results of changing investor risk behaviour over time which turn out to average to zero when the price of risk is assumed to be static in the model. Overall, market excess returns are not significantly priced to their comovement with the world and trading-bloc returns, as well as the local market variance. Ostensibly, these results suggest that the returns of the different markets do not converge to either the world or trading-bloc factor, and the markets are highly segmented from the asset pricing point of view. This, however, must be interpreted with caution. The assumption of constant market price of risk is restrictive. Further, the dynamics of the time-varying pricing process is not revealed and may have been masked by the static estimates.

The estimated results for the variance equations are not reported to conserve space. Most of the estimated coefficients in the variance equations are statistically significant and we perform a series of Wald tests to examine the significance of the ARCH, GARCH and asymmetric effects in H_t. The results in Table 3 show that the ARCH effects in the system are statistically significant. The GARCH effect, however, is not as significant particularly in many of the EU markets. The asymmetric effects are detected in almost all the markets. The significance of these ARCH-based effects indicates spillover of volatility from one market to the other, and therefore linkages exist between the returns of different markets. For the diagnostic checking, a multivariate normality test and two multivariate Portmanteau tests for autocorrelation are performed. The results of the Portmanteau test found autocorrelation problem for New Zealand and the five AFTA markets when the residuals are not standardized. However, no autocorrelation problem is reported if standardized residuals are used. The results of the diagnostic tests and significance of the ARCH, GARCH and asymmetric news effects justify the use of the multivariate specification given in equations (8) and (9). It must be cautioned, however, that the system residuals may not always follow a multivariate normal distribution.

Table 2 Estimation results of the BEKK multivariate GARCH model

Bloc/Country	α_M	α_W	α_T	λ_M^W	λ_M^T	λ_M^M	λ_W^W	λ_T^T	LogL	AIC	SC
EU											
Austria	1.047	0.316	-0.448	0.204	-2.750	0.257	-0.026	0.354	-956.996	11.136	11.551
	(1.154)	(0.463)	(0.209)**	(0.729)	(1.604)*	(0.160)	(0.169)	(0.166)**			
Belgium	0.276	0.411	0.053	-0.677	1.010	0.147	-0.059	-0.025	-903.888	10.533	10.947
	(0.459)	(0.464)	(0.198)	(0.377)*	(0.490)**	(0.165)	(0.155)	(0.176)			
Denmark	0.424	-0.039	0.072	0.071	-0.340	0.006	0.091	-0.051	-931.086	10.842	11.256
	(0.903)	(0.665)	(0.244)	(0.540)	(0.684)	(0.147)	(0.233)	(0.190)			
Finland	-1.696	-0.099	0.475	0.732	3.527	-0.090	0.120	-0.372	-1071.545	12.438	12.852
	(1.776)	(0.584)	(0.420)	(0.632)	(2.743)	(0.086)	(0.214)	(0.326)			
France	-1.349	-2.948	0.281	0.998	-0.180	-0.278	1.063	-0.283	-850.678	9.928	10.343
	(2.498)	(4.455)	(1.001)	(0.930)	(0.619)	(0.194)	(1.517)	(0.989)			
Germany	-0.476	0.224	1.026	0.161	1.097	-0.094	-0.010	-1.045	-902.305	10.515	10.929
	(0.633)	(0.395)	(2.142)	(0.306)	(0.795)	(0.133)	(0.140)	(2.173)			
Greece	0.687	0.316	-0.180	0.201	-0.469	-0.019	-0.024	0.034	-1154.344	13.379	13.793
	(0.806)	(0.463)	(0.190)	(0.320)	(0.266)*	(0.073)	(0.163)	(0.040)			
Ireland	0.904	0.438	0.031	-0.361	0.116	0.046	-0.078	-0.017	-934.459	10.880	11.295
	(0.757)	(0.491)	(0.251)	(0.562)	(0.398)	(0.230)	(0.172)	(0.198)			
Italy	0.604	0.457	-0.015	-0.025	-0.074	-0.030	-0.077	0.028	-986.550	11.472	11.887
	(0.659)	(0.517)	(0.212)	(0.243)	(0.477)	(0.075)	(0.174)	(0.199)			
Netherlands	0.845	0.340	0.658	-0.384	0.158	0.087	-0.057	-0.566	-838.929	9.795	10.209
	(0.596)	(0.528)	(0.376)*	(0.497)	(0.361)	(0.250)	(0.176)	(0.313)*			
Portugal	0.164	0.050	-0.187	-0.017	-0.168	0.034	0.058	0.023	-1075.794	12.486	12.901
	(1.041)	(0.530)	(0.176)	(0.946)	(0.346)	(0.282)	(0.190)	(0.041)			
Spain	-0.037	0.075	1.496	0.305	0.009	-0.046	0.046	-1.050	-931.465	10.846	11.261
	(0.916)	(0.327)	(1.290)	(0.338)	(0.550)	(0.194)	(0.119)	(0.911)			
Sweden	0.446	0.221	0.066	-0.040	0.262	-0.008	-0.009	-0.056	-967.914	11.260	11.675
	(0.899)	(0.407)	(0.255)	(0.272)	(1.077)	(0.105)	(0.145)	(0.208)			
UK	0.617	0.356	0.444	-0.324	-0.354	0.140	-0.060	-0.753	-787.747	9.213	9.627
	(0.413)	(0.514)	(0.270)*	(0.380)	(0.627)	(0.218)	(0.182)	(0.453)*			

Note: Figures in the parentheses are standard errors. *, ** and *** denote significance at the 0.10, 0.05 and 0.01 levels, respectively. **The estimated results for the variance equations are not reported to conserve space. They are available upon request.**

Table 2 (continued) Estimation results of the BEKK multivariate GARCH model

Bloc/Country	α_M	α_W	α_T	λ_M^W	λ_M^T	λ_M^M	λ_W^W	λ_T^T	LogL	AIC	SC
EFTA											
Norway	0.230	-0.010	0.148	0.258	0.487	-0.126	0.072	-0.076	-1034.881	12.021	12.436
	(2.456)	(0.493)	(0.348)	(0.544)	(3.404)	(0.234)	(0.178)	(0.159)			
Switzerland	0.984	-0.435	0.905	-0.095	0.106	-0.104	0.225	-0.188	-1030.153	11.968	12.382
	(0.931)	(0.763)	(0.368)**	(0.650)	(0.328)	(0.256)	(0.269)	(0.086)**			
NAFTA											
Canada	-0.0191	0.1411	-4.2213	0.122	0.0051	-0.0347	-0.0109	0.0231	-1316.283	15.219	15.634
	(0.733)	(0.266)	(2.223)*	(0.339)	(0.036)	(0.242)	(0.096)	(0.013)*			
Mexico	1.6183	0.2857	-4.845	-0.2182	0.0338	-0.0318	-0.0562	0.023	-1471.346	16.981	17.396
	(0.912)*	(0.244)	(2.111)**	(0.136)	(0.023)	(0.053)	(0.088)	(0.013)*			
US	0.6132	0.0358	0.0299	-0.7311	-1.7286	0.5794	0.0519	-2.9466	-367.805	4.441	4.855
	(0.410)	(0.687)	(0.016)*	(0.424)*	(2.537)	(0.343)*	(0.218)	(1.652)*			
CER											
Australia	0.3176	0.588	-0.2418	-0.1554	-0.1959	0.1267	-0.1416	0.0484	-1027.088	11.933	12.347
	(1.170)	(0.618)	(0.578)	(0.532)	(0.474)	(0.325)	(0.212)	(0.124)			
New Zealand	-0.222	0.0749	0.1035	0.1193	0.4166	-0.0734	0.0469	-0.0455	-1021.376	11.868	12.282
	(1.495)	(0.449)	(0.689)	(0.503)	(0.627)	(0.268)	(0.162)	(0.301)			
AFTA											
Indonesia	-1.482	-1.963	0.619	0.266	0.208	-0.080	0.052	0.023	-1373.098	15.865	16.279
	(2.553)	(0.643)***	(0.456)	(0.497)	(0.112)*	(0.054)	(0.198)	(0.036)			
Malaysia	-1.055	-0.257	0.020	0.350	0.198	-0.058	0.123	0.017	-1138.392	13.198	13.612
	(0.832)	(0.439)	(0.230)	(0.355)	(0.167)	(0.071)	(0.129)	(0.045)			
Philippines	-4.760	-0.099	-0.158	0.564	-0.104	0.241	0.086	0.022	-1221.037	14.137	14.551
	(3.001)	(0.772)	(0.359)	(0.590)	(0.116)	(0.148)	(0.260)	(0.036)			
Singapore	-0.088	0.348	0.221	0.190	-0.005	-0.027	-0.034	-0.024	-1080.795	12.543	12.958
	(0.548)	(0.760)	(0.271)	(0.257)	(0.059)	(0.058)	(0.254)	(0.031)			
Thailand	0.060	0.756	0.106	0.140	0.019	-0.026	-0.189	0.012	-1252.449	14.494	14.908
	(0.797)	(0.572)	(0.396)	(0.335)	(0.116)	(0.065)	(0.185)	(0.041)			

Note: Figures in the parentheses are standard errors. *, ** and *** denote significance at the 0.10, 0.05 and 0.01 levels, respectively. **The estimated results for the variance equations are not reported to conserve space. They are available upon request.**

Table 3 Wald tests for conditional time-varying effects and diagnostics tests of the BEKK multivariate GARCH model

Bloc/ Country	Multivariate ARCH Effects	Multivariate GARCH Effects	Multivariate Asymmetric Effects	Multivariate Normality Tests	Multivariate Portmanteau Tests for Autocorrelation			
					Ordinary Residuals		Standardized Residuals	
					Q (12)	Adjusted Q(12)	Q(12)	Adjusted Q(12)
EU								
Austria	8.2928	8.3981	17.8393	22.2663	97.0087	100.4977	94.7059	98.1386
	(0.0403)**	(0.0385)**	(0.0005)***	(0.6203)	(0.7670)	(0.6834)	(0.8156)	(0.7411)
Belgium	9.1668	2.8721	31.6260	48.9448	123.5142	127.9339	121.4964	126.1199
	(0.0272)**	(0.4118)	(0.0000)***	(0.0029)***	(0.1460)	(0.0925)*	(0.1768)	(0.1122)
Denmark	5.4484	3.5931	16.6351	8.6091	106.0882	110.3144	109.2584	113.5525
	(0.1418)	(0.3089)	(0.0008)***	(0.9990)	(0.5340)	(0.4201)	(0.4480)	(0.3384)
Finland	8.4498	3.0888	5.2658	22.4203	89.2361	92.8955	86.7109	90.4424
	(0.0376)**	(0.3781)	(0.1533)	(0.6114)	(0.9054)	(0.8494)	(0.9345)	(0.8888)
France	8.7445	2.6273	1.8084	39.4077	96.2458	100.3395	95.7690	99.8881
	(0.0329)**	(0.4527)	(0.6131)	(0.0335)**	(0.7838)	(0.6874)	(0.7939)	(0.6987)
Germany	15.4956	5.2967	17.6883	54.5774	105.4251	109.5681	105.6049	109.7992
	(0.0014)***	(0.1513)	(0.0005)***	(0.0006)***	(0.5522)	(0.4398)	(0.5473)	(0.4337)
Greece	3.7761	2.5255	26.8493	21.1636	98.2820	101.6270	96.3906	99.7933
	(0.2867)	(0.4707)	(0.0000)***	(0.6835)	(0.7378)	(0.6543)	(0.7806)	(0.7011)
Ireland	17.9232	7.6874	4.4253	20.1088	87.4070	91.0321	98.3082	102.4310
	(0.0005)***	(0.0529)*	(0.2190)	(0.7411)	(0.9272)	(0.8800)	(0.7371)	(0.6331)
Italy	17.1393	5.1336	43.4181	44.1160	103.7460	107.8646	112.2429	116.8004
	(0.0007)***	(0.1623)	(0.0000)***	(0.0105)**	(0.5979)	(0.4856)	(0.3706)	(0.2649)
Netherlands	6.2374	3.0906	21.3478	38.4661	102.9820	106.6902	98.7305	102.3490
	(0.1006)	(0.3779)	(0.0001)***	(0.0416)**	(0.6184)	(0.5176)	(0.7271)	(0.6353)
Portugal	7.7982	2.2712	15.0638	47.1373	98.3425	101.7819	102.7308	106.6293
	(0.0504)*	(0.5181)	(0.0018)***	(0.0047)***	(0.7363)	(0.6502)	(0.6251)	(0.5192)
Spain	3.3100	8.9259	8.0899	39.8252	87.0913	90.1711	86.1207	89.1091
	(0.3463)	(0.0303)**	(0.0442)**	(0.0304)**	(0.9306)	(0.8927)	(0.9403)	(0.9071)
Sweden	13.1373	8.2518	33.5383	25.5037	102.7738	107.0560	97.9820	102.1617
	(0.0043)***	(0.0411)**	(0.0000)***	(0.4344)	(0.6240)	(0.5076)	(0.7448)	(0.6402)
UK	10.4763	3.5041	9.2376	45.0733	106.7501	111.3694	110.2029	114.8396
	(0.0149)**	(0.3202)	(0.0263)**	(0.0082)***	(0.5159)	(0.3927)	(0.4230)	(0.3082)

Note: Figures in the parentheses are p-values. *, ** and *** denote significance at the 0.10, 0.05 and 0.01 levels, respectively. The standardized residual test is based on Lütkepohl (1991).

Table 3 (continued) Wald tests for conditional time-varying effects and diagnostics tests of the BEKK multivariate GARCH model

Country	Multivariate ARCH Effects	Multivariate GARCH Effects	Multivariate Asymmetric Effects	Multivariate Normality Tests	Multivariate Portmanteau Tests for Autocorrelation: Ordinary Residuals Q (12)	Ordinary Residuals Adjusted Q (12)	Standardized Residuals Q (12)	Standardized Residuals Adjusted Q (12)
EFTA								
Norway	9.0603	3.8601	6.1695	28.7680	111.5710	116.3718	108.1704	112.8350
	(0.0285)**	(0.2770)	(0.1036)	(0.2737)	(0.3876)	(0.2740)	(0.4773)	(0.3559)
Switzerland	3.3752	8.2677	7.2055	17.8588	114.9238	119.8649	109.8772	114.6947
	(0.3373)	(0.0408)**	(0.0656)*	(0.8483)	(0.3062)	(0.2048)	(0.4316)	(0.3115)
NAFTA								
Canada	3.5821	17.3350	14.2168	62.7468	90.8914	94.3124	109.4883	113.6995
	(0.3103)	(0.0006)***	(0.0026)***	(0.0000)***	(0.8822)	(0.8233)	(0.4419)	(0.3349)
Mexico	18.6163	10.1692	15.2175	60.2047	82.5102	85.9361	93.9434	97.8247
	(0.0003)***	(0.0172)**	(0.0016)***	(0.0001)***	(0.9676)	(0.9420)	(0.8303)	(0.7484)
US	4.2119	13.3556	17.8657	10.0204	107.7745	112.3473	103.4412	107.7596
	(0.2395)	(0.0039)***	(0.0005)***	(0.9966)	(0.4880)	(0.3680)	(0.6061)	(0.4884)
CER								
Australia	5.1810	6.6860	23.5868	10.9788	119.9980	124.7213	110.8766	115.3210
	(0.1590)	(0.0826)*	(0.0000)***	(0.9930)	(0.2024)	(0.1295)	(0.4054)	(0.2972)
New Zealand	6.1549	11.3832	10.3766	19.9446	124.1622	129.0137	110.2354	114.5098
	(0.1043)	(0.0098)***	(0.0156)**	(0.7497)	(0.1370)	(0.0821)*	(0.4222)	(0.3158)
AFTA								
Indonesia	28.1585	16.6686	3.9136	–	169.9061	175.1223	119.6726	124.1521
	(0.0000)***	(0.0008)***	(0.2709)		(0.0001)***	(0.0000)***	(0.2083)	(0.1371)
Malaysia	34.9760	38.1806	18.2188	30.0915	185.5774	192.3713	101.7211	105.8056
	(0.0000)***	(0.0000)***	(0.0004)***	(0.2209)	(0.0000)***	(0.0000)***	(0.6518)	(0.5418)
Philippines	41.3224	11.1437	9.7845	22.7292	146.6449	152.3326	106.9814	111.3251
	(0.0000)***	(0.0110)**	(0.0205)**	(0.5934)	(0.0079)***	(0.0032)***	(0.5096)	(0.3939)
Singapore	36.9820	28.5783	19.2845	72.9491	128.5234	133.1739	105.2851	108.8786
	(0.0000)***	(0.0000)***	(0.0002)***	(0.0000)***	(0.0867)*	(0.0505)*	(0.5560)	(0.4582)
Thailand	26.2739	7.6884	12.5444	48.4701	160.0107	165.7121	114.2338	118.4611
	(0.0000)***	(0.0529)*	(0.0057)***	(0.0033)***	(0.0009)***	(0.0003)***	(0.3222)	(0.2312)

Note: Figures in the parentheses are p-values. *, ** and *** denote significance at the 0.10, 0.05 and 0.01 levels, respectively. The standardized residual test is based on Lütkepohl (1991). – The multivariate normality test failed due to the need to raise a negative number to a non integer power in the computation.

4.2. World and Trading-Bloc Betas

The world and trading-bloc betas are estimated based on equation (4) with a univariate GARCH(1,1) specification for the error process. These are static measures computed for comparison purposes. The time-varying betas given in equations (10a) and (10b) are also computed. The mean of the time-varying betas are calculated according to equations (11a) and (11b) and reported in Table 4. The static estimates are also given in the same table. Generally, the estimated betas are statistically different from zero, except for a few cases. Overall, the two sets of estimates for the world and trading-bloc betas do not vary much in terms of magnitude. In nearly all cases, the world betas are larger than the trading-bloc betas. This suggests that the world market movements have a larger impact than the average market movement within a trading bloc on the pricing of market indices. The trading-bloc factor, however, is not negligible.

The trading-bloc betas are large for some of the markets in EU, including Austria, Belgium, France, Germany, Netherlands and Spain than the betas of the markets in other trading blocs. The results reported here for the EU markets are consistent with the findings of Bekaert et al. (2005) in that these markets have a large exposure to the trading-bloc factor. They found that the average betas of small EU markets with respect to the US market (proxy for the world factor) are surprisingly small, while their betas with respect to the European markets are high, mostly exceeding 0.7. Their smaller estimates of world betas could be due to the use of US data to compute the proxy for world portfolio. Further, in constructing the EU factor, we have orthogonalized the impact of the world factor, which was not the case in Bekaert et al. (2005).

For NAFTA, there is evidence that the trading-bloc betas are not significant. This evidence is stronger for the BEKK multivariate GARCH model. As the world factor is highly dependent on the US market movements, it also remains the key driving force to the determination of the asset prices in the market in NAFTA. Bekaert et al. (2005) also reported weak regional beta in the Latin region in their emerging market sample.

The magnitude of the trading-bloc beta is very close to that of the world beta for Indonesia, Malaysia, Philippines and Thailand of the AFTA bloc. Singapore with the most developed market has a relatively lower exposure to the trading-bloc factor compared to the other markets in AFTA. Heaney and Hooper (1999) also found that Singapore shows greater openness to the world economy compared to the other markets in the AFTA bloc. Bekaert et al. (2005) did not include Singapore in their study on emerging markets. They reported larger world betas compared to the regional factor for the other Asian markets. Given that this study uses the trading-bloc portfolio for the five ASEAN countries, our beta estimates are higher than those reported by Bekaert et al. (2005) who considered a regional factor. This comparison suggests that the five ASEAN markets have a larger exposure to the AFTA trading bloc than the regional factor.

Consistent with Heaney and Hooper (1999), the estimated world betas for Australia and New Zealand are both greater than the estimated trading-bloc betas. Interestingly, the New Zealand market has a much higher exposure to the trading-bloc factor compared to Australia.

Table 4 Comparison of betas estimated from the univariate and multivariate GARCH models

Bloc/	World Beta		Trading-bloc Beta	
Country	Univariate	Multivariate	Univariate	Multivariate
EU				
Austria	0.6203***	0.6483***	0.6673***	0.8052***
Belgium	0.8151***	0.7740***	0.5901***	0.8202***
Denmark	0.8332***	0.8431***	0.5461***	0.5645***
Finland	1.5071***	1.5150***	0.0641	0.2496***
France	1.1179***	1.1507***	0.9221***	1.0571***
Germany	1.2095***	1.1333***	0.8338***	0.8385***
Greece	0.9764***	0.8931***	0.2606**	0.4691***
Ireland	0.9275***	0.9394***	0.5024***	0.4304***
Italy	0.9552***	0.9786***	0.8597***	0.9061***
Netherlands	1.0819***	1.0185***	0.6992***	0.7368***
Portugal	0.9077***	0.8458***	0.3444***	0.4743***
Spain	1.1825***	1.1583***	0.6038***	0.7795***
Sweden	1.3816***	1.3704***	0.4491***	0.5367***
UK	0.7791***	0.7882***	0.5611***	0.4512***
EFTA				
Norway	1.1395***	1.2422***	0.1988**	0.1910***
Switzerland	0.8204	0.8383***	0.0849	0.0797***
NAFTA				
Canada	1.0016***	0.8016***	0.0193**	-0.0008
Mexico	1.2637***	1.2354***	0.0138	-0.0059
US	0.9701***	0.7827	-0.8265**	-0.0814
CER				
Australia	0.8383***	0.9140***	0.3379***	0.3352***
New Zealand	0.8928***	0.9342***	0.7011***	0.7838***
AFTA				
Indonesia	1.4009***	0.7424***	0.9601***	0.7253***
Malaysia	0.9209***	0.7643***	0.8273***	0.6987***
Philippines	0.9841***	0.9329***	0.7008***	0.6689***
Singapore	1.2892***	0.9643***	0.7746***	0.5733***
Thailand	0.9353***	1.3503***	0.5003***	0.7398***

Note: *, ** and *** denote significance at the 0.10, 0.05 and 0.01 levels, respectively. The significance of the point estimates of the betas from the multivariate GARCH models are tested with *t*-test.

In short, our results suggest that the trading-bloc factor has an important role in the determination of capital asset pricing. The consistency of our results with those of other studies shows that we cannot ignore the effect of regionalism. By using trading-bloc specific portfolios, instead of regional portfolios considered in other studies, we manage to capture a larger impact due to regionalism that arises from the formation of trading blocs. Consequently, this implies that international stock markets are only partially integrated. Regionalism remains a factor that explains stock market segmentation.

The static and point estimates do not show the dynamics of risk exposure. The time-varying world and trading-bloc betas are plotted in Figure 1, which also depicts the static estimates. The world betas are plotted with solid lines while the dash lines represent the trading-bloc betas. In addition, there are two shaded areas in each graph. The first shaded area represents the period of the Asian financial crisis (July 1997 to December 1998). The second shaded area relates to the occurrence of the early 2000s market crashes (March 2000 to March 2003) that were mainly felt in the Western developed countries, which include the dot-com bubble crash, the post September-11 crash and the stock market downturn of 2002.

For most of the markets, the time-varying betas hover around the level of the static betas. This shows that the two set of estimates of betas are rather close to each other. Except Austria, the level of the static world beta is consistently higher than the level of the static trading-bloc beta. Thus, the static estimates suggest that the systematic risk exposure to the world factor is consistently higher than the exposure to the trading-bloc factor. The results remain the same with the use of time-varying betas for the markets in the EFTA bloc and two out of three markets in the NAFTA bloc. However, this is not the case for US, and many markets in AFTA, CER and EU, where the time-varying betas exhibited different behaviour. Being the major world player, the risk exposure of the US market is certainly far more erratic than that suggested by the static estimate.

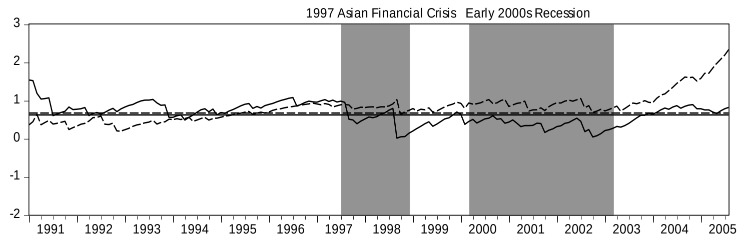
The time-varying betas for the AFTA bloc display a clear “1997 Asian crisis” effects. The estimates for the stock markets of AFTA have experienced a jump in magnitudes in 1997 but started to stabilize after year 1999. The pattern is particularly obvious for the case of Indonesia, Malaysia and Thailand; three of the countries most seriously affected by the crisis. While the world beta for Indonesia, Malaysia and Philippines experienced bigger fluctuations during this period, the world beta for Thailand has increased to a much higher level not experienced before. The trading-bloc betas, on the other hand, are relatively stable and show similar behaviour to the other periods. The Singapore market does not seem affected by the financial crisis. On the other hand, the market crashes in the early 2000s did not change the risk exposure behaviour of these markets. The beta estimates remained at about the level before the crashes.

Another interesting pattern observed in Figure 1 is that the time-varying trading-bloc betas in the EU bloc have exceeded the time-varying world beta in view of the recent strengthening and enlarging of the EU monetary union in the late 1990s. The pattern is especially obvious for Austria, Belgium, Denmark, Greece, Italy, Portugal and Spain. This shows that the trading-bloc effects are gaining more importance in the pricing of capital assets of the EU markets. Such observation is not captured through the static estimates. The findings show that the time-varying multivariate GARCH model is able to detect changes in the way that assets are priced under different market conditions. Similar evidence is observed for New Zealand. Its market exposure to the trading-bloc factor has exceeded that to the world factor since 2000.

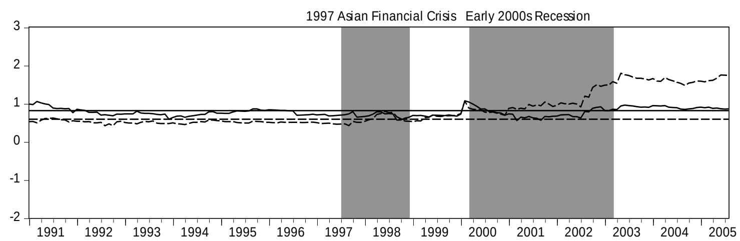
Figure 1 World and trading-bloc betas of the univariate and multivariate GARCH models

EU

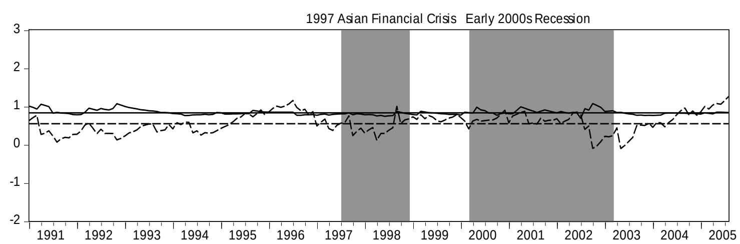
Austria



Belgium



Denmark



Finland

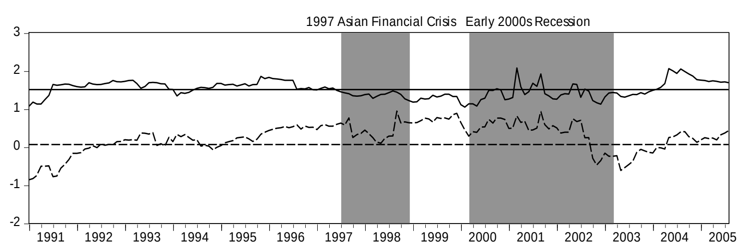
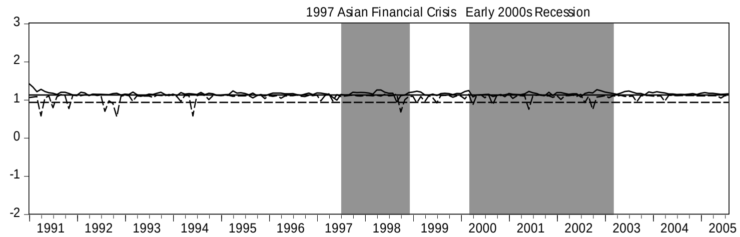
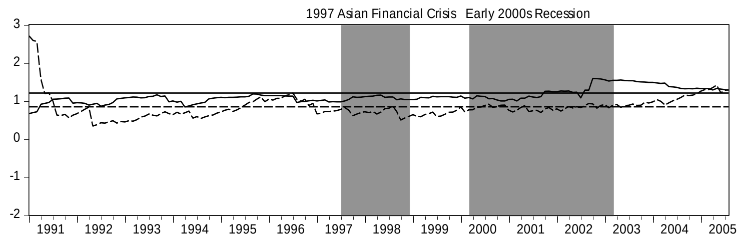


Figure 1 World and trading-bloc betas of the univariate and multivariate GARCH models (continue)

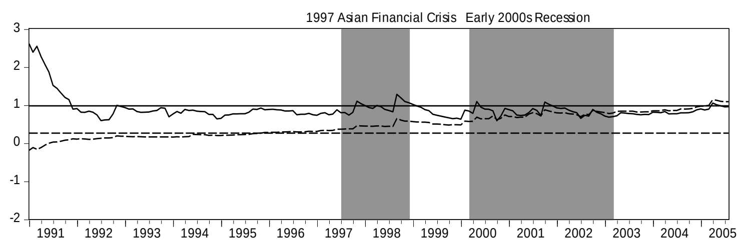
France



Germany



Greece



Ireland

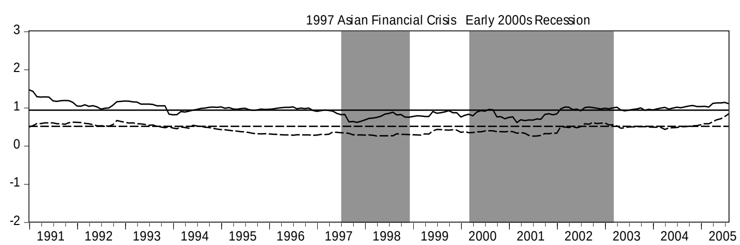
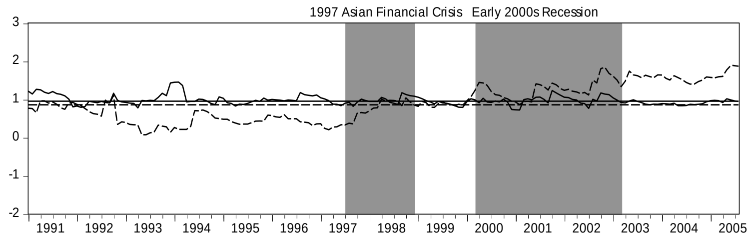
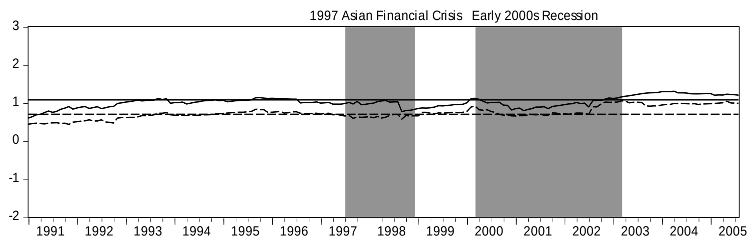


Figure 1 World and trading-bloc betas of the univariate and multivariate GARCH models (continue)

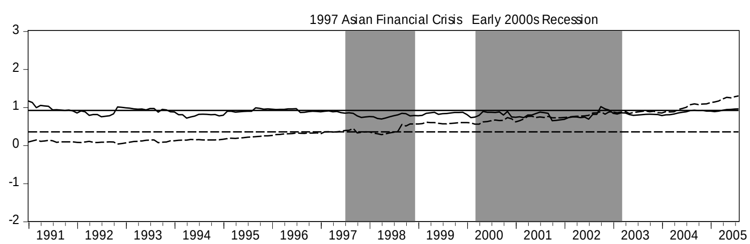
Italy



Netherlands



Portugal



Spain

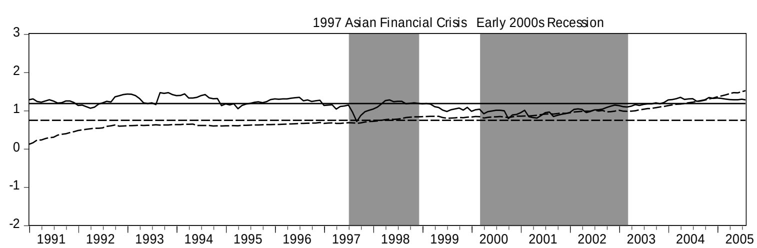
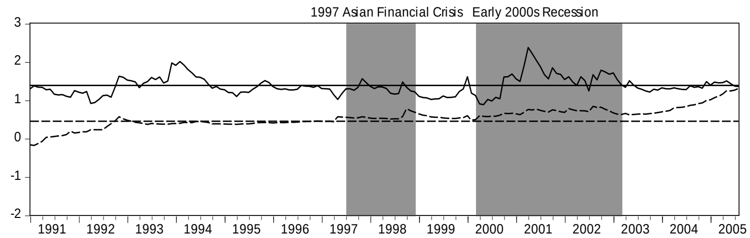
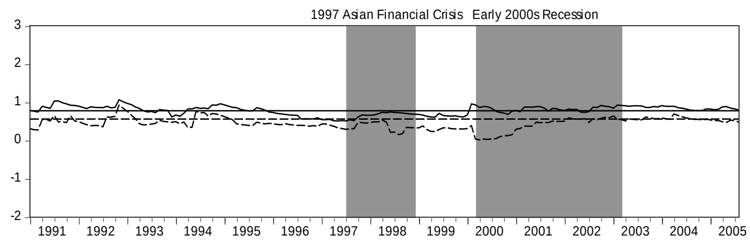


Figure 1 World and trading-bloc betas of the univariate and multivariate GARCH models (continue)

Sweden

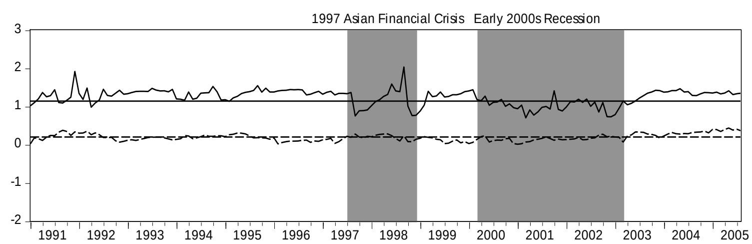


UK



EFTA

Norway



Switzerland

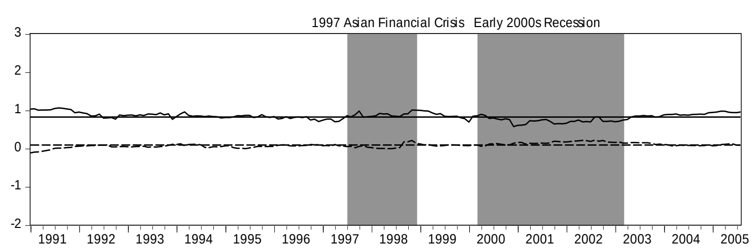
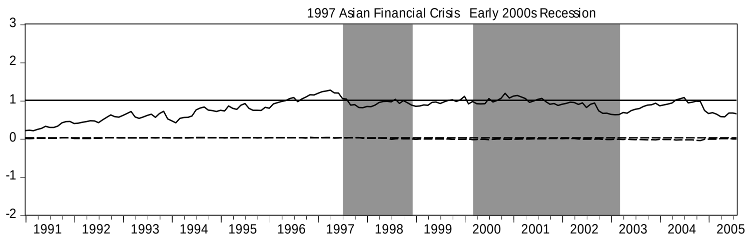


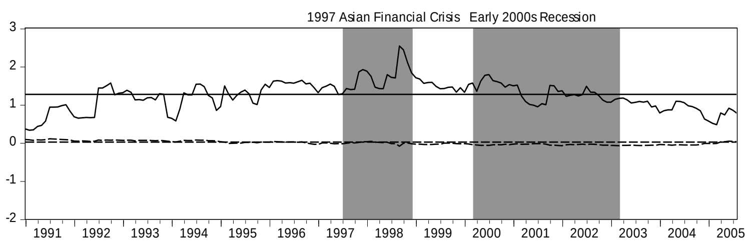
Figure 1 World and trading-bloc betas of the univariate and multivariate GARCH models (continue)

NAFTA

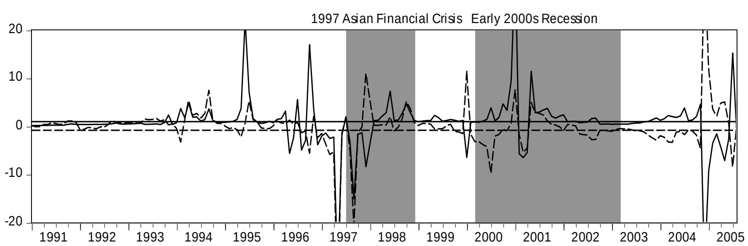
Canada



Mexico



US



CER

Australia

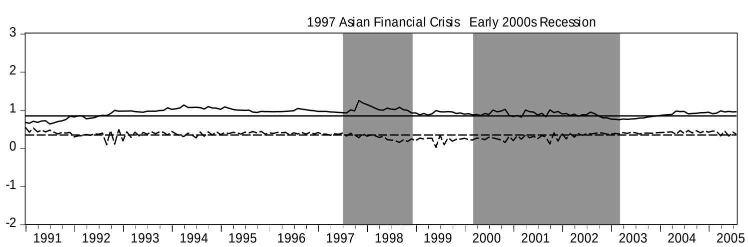
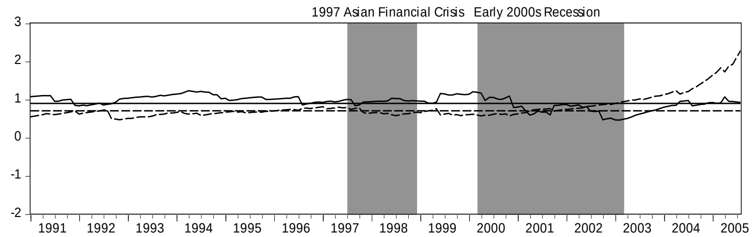


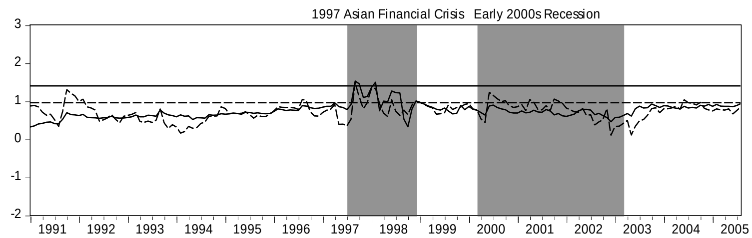
Figure 1 World and trading-bloc betas of the univariate and multivariate GARCH models (continue)

New Zealand

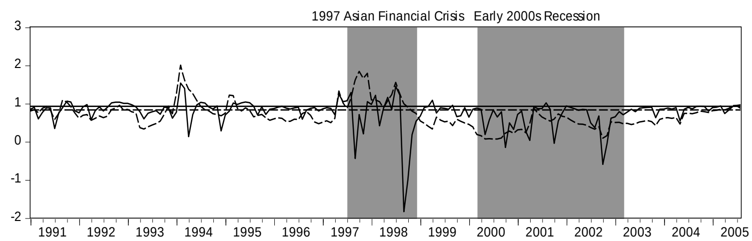


AFTA

Indonesia



Malaysia



Philippines

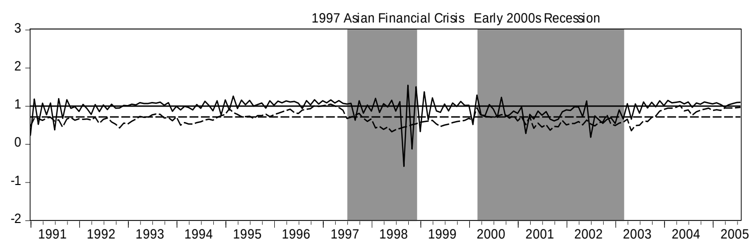
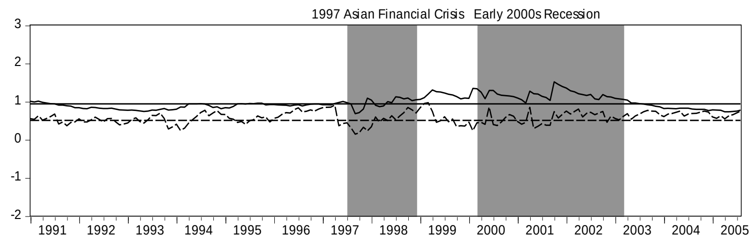
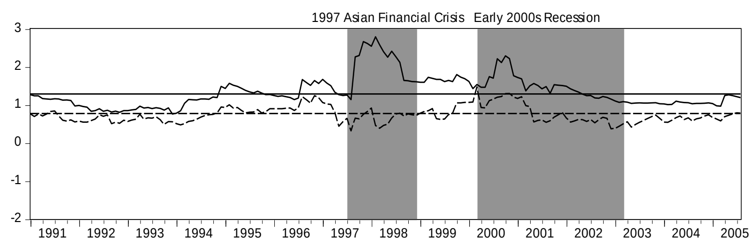


Figure 1 World and trading-bloc betas of the univariate and multivariate GARCH models (continue)

Singapore



Thailand



Note: Solid lines are the world betas while dash lines are the trading-bloc betas.

5. Concluding Remarks

The findings of this study suggest that international stock markets are only partially integrated using conditional capital asset pricing models. The results indicate that the effect of trade regionalism that leads to world market segmentation cannot be ignored. By examining regionalism due to formation of trading blocs, in contrast to regional factor attributed to geographical location considered in other works, this study captures a stronger impact of regionalism on segmentation. The systematic risk exposure to movements in stock markets within a trading bloc therefore remains an important factor in the pricing of capital assets.

This study proposes using estimates from the multivariate GARCH asset pricing models for computing time-varying betas that have the advantage of tracking the changing behaviour of the world and trading-bloc risk exposure over time. Most of the stock markets are found to have different risk exposure behaviour across time, especially those in the AFTA and EU bloc. During the period of the Asian financial crisis, all the stock markets in AFTA except Singapore exhibited erratic asset pricing behaviour which is not observed for other periods. The results also show that the recent monetary unification of the EU has led to an increased importance of the trading-bloc factor in asset pricing determination for many of the markets in EU.

References

Adler, M. and Qi, R., 2003. Mexico's integration into the North American capital market. Emerging Market Review 4, 91-120.

Akdogan, H., 1992. Behaviour of systematic risk in a regionally integrated model for stock prices. Economics Letters 39, 213-216.

Baba, Y., Engle, R.F., Kraft, D.F. and Kroner, K.F., 1990. Multivariate simultaneous generalized ARCH. Department of Economics Mimeo, University of California, San Diego.

Bekaert, K.G., Harvey, C.R. and Ng, A., 2005. Market integration and contagion. Journal of Business 78, 39-69.

Black, F., 1976. Studies of stock price volatility changes. Proceedings of the 1976 Meetings of the American Statistical Association, Business and Economics Statistics Section, American Statistical Association, Washington DC, pp.177-181.

Bollerslev, T., 1990. Modelling the coherence in short-run nominal exchange rates: a multivariate generalized ARCH model. Review of Economics and Statistics 72, 498-505.

Bollerslev, T., Engle, R.F. and Wooldridge, J.M., 1988. A capital asset pricing model with time-varying covariances. Journal of Political Economy 96, 116-131.

Christie, A. 1982., The stochastic behavior of common stock variances: value, leverage, and interest rate effects. Journal of Financial Economics 10, 15-36.

Cochrane, J.H., 2005. Asset Pricing, revised edn. Princeton University Press, New Jersey.

Engle, R.F. and Kroner, K.F., 1995. Multivariate simultaneous generalized ARCH. Econometric Theory 11, 122-150.

Engle, R.F., Ng, V.K. and Rothschild, M., 1990. Assets pricing with a factor ARCH covariance structure: empirical estimates for Treasury Bills. Journal of Econometric 45, 213-238.

Fratzscher, M., 2002. Financial market integration in Europe: on the effects of EMU on stock markets. International Journal of Finance and Economics 7, 165-193.

French, K.R., Schwert, G.W. and Stambaugh, R.F., 1987. Expected stock returns and volatility. Journal of Financial Economics 19, 3-29.

Heaney, R. and Hooper, V., 1999. World, regional and political risk influences upon Asia Pacific equity market returns. Australian Journal of Management 24, 131-142.

Nelson, D.B., 1991. Conditional heteroskedasticity in asset returns: a new approach. Econometrica 59, 347-370.

Pagan, A., 1984. Econometric issues in the analysis of regressions with generated regressors. International Economic Review 25, 267-290.

Roll, R., 1977. A Critique of the asset pricing theory's test; Part I: an equilibrium theory of value for market forecasts. Journal of Business 54, 363-406.

Schwert, G.W., 1989. Why does stock market volatility change over time? Journal of Finance 44, 1115-1153.

Zokian, J.M., 1994. Threshold heteroskedastic models. Journal of Economic Dynamics and Control 18, 931-955.