TIME-VARYING EXCHANGE RATE EXPOSURE COEFFICIENTS: EVIDENCE FROM EMERGING MARKETS

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ABSTRACT

Based on the theoretical framework provided by the International Capital Asset Pricing Model (ICAPM), this paper uses time-varying second moments to investigate exchange rate exposure betas. The study is carried out at country level using stock indexes and trade-weighted exchange rates of a selected set of emerging economies. Time-varying exchange rate exposure betas are obtained with the help of a Multivariate GARCH-M model with explicit focus on the non-orthogonality between exchange rate changes and market returns. Certain aspects of the stochastic structure underlying the exposure betas are examined. Findings of the paper indicate that, although they are likely to vary over time, exchange rate exposure betas for Korea and Taiwan follow mean-reverting long-memory processes. The presence of mean-reverting exchange rate exposure coefficients has important implications for investment and hedging strategies. However, the exposure beta for Thailand is most likely to be characterized by a non-stationary unit root process

Keywords: Time-varying exchange rate exposure, Multivariate GARCH-M models, International CAPM, Fractionally integrated processes

JEL Classification: C22; F31; F37; G12; G15

1. INTRODUCTION

Most of the studies in exchange rate exposure literature implicitly assume that the exposure coefficients remain unchanged over time. Nevertheless, there are several reasons to assume that exposure coefficients are time-varying. First, a country's composition and/or shares of exports and imports may change drastically over time due to both external and internal factors[1]. Changes in demand due to the rise of new competitors in international arena is an example for the former while the introduction of trade liberalization policies is an example for the latter. Second, as substitutes are being introduced, elasticity of demand for a country's products and competitive structure of industries are likely to change over time. Allayannis and Ihrig (2001) argue that the changes in competitive structure affect the exchange rate exposure of industries. Third, financial market deregulations and liberalization attempts may lead to changes in foreign investments in local financial assets which in turn affect the exchange rate exposure of a country's stocks. Fourth, the change in location of production of MNCs in response to persistent strong currency positions may lead to the changes in the sensitivity of sectoral returns to exchange rate changes. Fifth, incidents like the 1997 currency crisis may lead to remarkable volatility changes in exchange rate markets.

In exchange rate exposure literature, there are three groups of studies that analyze the time-varying nature of exposure betas. The first group uses some primitive methods for this purpose. For instance, Dominguez and Tesar (2006) divide the sample period into a few sub-periods and estimate an exposure coefficient for each sub-period. Williamson (2001) uses dummy variables to distinguish between sub-periods and observe that

[1] Allayanis (1997) observes that the status of some US industries change from net exporters to net importers within the same sample period.

the exposure is time-variant. Alternatively, Entoff and Jamin (2003) use overlapping moving window regressions to show the time-varying behaviour of the exchange rate exposure of a bunch of German firms. Bodner and Wong (2003) also use moving window regressions with various return horizons (1, 3, 6 and 12 months)[2].

The second group of studies uses pre-specified determinants of exposure coefficients to analyze the time-variation of exposure. Allayannis (1997) suggests that exposure beta is determined by export and import shares. Using an appropriate model to accommodate this relationship, the study cites evidence for time-variation of exposure in some 4 digit level SIC industries. He reports that the same data set at industry level show significant exchange rate exposure only when the exposure is assumed to be time-varying. Allayannis and Ihrig (2001) inquire into the same phenomenon in terms of three determinants of exposure: (a) an industry's competitive structure where it sells its production; (b) the interaction of the competitive structure of the export market and the export share; (c) the interaction of the competitive structure of the imported input market and its imported input share. Mark-ups are used as a measure of the competitive structure. Bodner et al. (2002) suggest a somewhat similar model in terms of time-varying exchange rate pass-through, though they are not able to show significant evidence for time-varying exposure. Chiao and Hung (2000) use the same determinants appearing in Allayannis (1997) to examine the time variation in the exchange rate exposure of Taiwanese exporting firms. In addition, they employ dummy variables to check whether the exchange rate exposure is affected by the timing of three liberalization effects introduced within the economy. Bodnar and Gentry (1993) add a few more factors to the list of pre-specified determinants of exposure, namely, whether the relevant industry produces traded or non-traded products, the amount of internationally-priced inputs used and the industry's foreign direct investment. In an attempt of seeking the determinants of the exposure of Japanese firms, Chow and Chen (1998) use three proxies for the hedging incentives which in turn depend on the firm size. In addition to aggregate export and import shares to GDP, Entorf and Jamin (2003) use the absolute distance between exchange rates and their long-run mean as a determinant of exposure.

The contribution of the second group of studies to the literature is more appealing than that of the first group as they also show the determinants of the time-variation in exposure. These include factors like time-varying export and import shares, mark-ups and pass-through. Nevertheless, those studies are not without limitations. First, the studies that analyze the time-variation in exchange rate exposure in terms of a set of pre-specified variables implicitly rely on a somewhat questionable assumption that there are no other (left out) determinants of time-variation. However, mainly due to the absence of theoretical explanations of such relationships, there may be unidentified factors which are yet important in explaining the time-variation in exposure. Second, the unavailability of data for much shorter return horizons may force the researchers to ignore some determinants or use unsuitable proxies[3]. Moreover, if return horizon in question is a day, any of such data series is not available on daily basis. Third, the above studies seem to have neglected the impact of the time-varying volatilities which is one of the major and crucial determinants of model parameters and the time-varying element of them. Fourth, the underlying stochastic structure of the exchange rate exposure betas is largely left unexamined. For instance, such studies do not answer the question whether the time-varying exposure betas are mean-reverting.

The third group of studies, which include Hunter (2005), Lim (2005) and Jayasinghe and Tsui (2008) use time-varying second moments to derive time-varying exchange rate exposure betas. While Hunter (2005) analyzes time-varying exchange rate exposure of small and large firms using Fama-French-type size-based portfolios, Lim (2005) derives time series of both market and exposure betas at country level. More importantly, Lim (2005) allows for non-orthogonality between the factors, a feature that Hunter (2005) fails to accommodate. Unlike Hunter (2005) or Lim (2005), Jayasinghe and Tsui (2008) directly use the mean structure of conditional ICAPM theorized by

[2] Dominguez and Tesar (2006) and Williamson (2001) employ Seemingly Unrelated Regressions (SUR). Bodnar and Wong (2003) rely on Generalized Method of Moments (GMM).

[3] For instance, Allayannins and Ihrig (2001) make an assumption that mark-ups vary on annual basis, though they work with monthly data, due to the unavailability of mark-ups on monthly basis.

Adler and Dumas (1983) and made econometrically feasible by De Santis and Gerard (1998) to derive time-varying exposure betas. They also look into certain aspects of the stochastic structure of exposure betas and report that time-varying exposure betas are likely to be mean reverting and follow long-memory processes.

This study focuses on the exchange rate exposure betas of country level stock returns from a few emerging economies. While using the methodology adopted in Jayasinghe and Tsui (2008), this study goes a step further by incorporating two important features. First, Jayasinghe and Tsui (2008) is based on daily data. To a certain extent, time-varying nature of the exposure betas associated with daily data is not a surprising finding. This is mainly due to the fact that a day is a very short return horizon. Alternatively, this study uses weekly data to check the robustness of the findings in Jayasinghe and Tsui (2008). Second, unlike in Jayasinghe and Tsui (2008) wherein the problem is viewed through a US investor in terms of the bilateral exchange rates between the US dollar and the relevant currency, the present study is based on the changes in trade-weighted exchange rates of the relevant currency. As discussed in Section 2, the use of trade-weighed exchange rates is more appropriate within the Adler and Duma's (1983) ICAPM framework than using bilateral exchange rates.

The present study looks into certain aspects of the underlying stochastic structure of exposure betas. An important empirical finding of the paper is that, although exchange rate exposure betas are likely to vary over time, they follow mean-reverting long-memory processes. The presence of mean-reverting exchange rate exposure coefficients has important implications for investment and hedging strategies.

The rest of this paper is organized as follows. The conceptual framework of the study that is based on a conditional international CAPM is outlined in section 2. Section 3 presents the information related to data and a preliminary analysis of the returns and exchange rate series. Section 4 describes the econometric methodology used. The main empirical findings are reported in Section 5. Some concluding remarks are included in Section 6.

2. CONCEPTUAL FRAMEWORK OF THE ANALYSIS

According to Adler and Dumas (1983), investors living in a world with purchasing power parity (PPP) violations, which is "the rule rather than exception", usually think of hedging against the purchasing power risk that would stem from unexpected inflation. The asset holding of a representative investor in such a context is characterized by two types of portfolios: (a) a world market portfolio of risky assets; and (b) "a personalized hedge portfolio which constitutes the best protection against inflation as [the investor] perceives it". As such, the expected return on an asset may consist of two parts: the market premium which depends on the asset's world market risk and an additional premium which depends on its usefulness to hedge purchasing power risk. Assuming a world with $L + 1$ number of countries (and currencies), the expected excess return on equity i is formally expressed as:

$$E_{t-1}(r_{i,t}) = \lambda_{m,t-1} Cov_{t-1}(r_{i,t}, r_{m,t}) + \sum_{l=1}^{L} \lambda_{\pi,l,t-1} Cov_{t-1}(r_{i,t}, \pi_{l,t}) \tag{1}$$

In Equation 1, $E_{t-1}(.)$ and $Cov_{t-1}(.)$ are expectations and covariances conditional on the current information set I_{t-1}; $r_{i,t}$ is excess return on a certain asset i; $r_{m,t}$ is excess return on world market portfolio; $\pi_{l,t}$ is the inflation rate in country l. The conditional covariance between $r_{i,t}$ and $r_{m,t}$ represents the world market risk and, as in the case of standard CAPM, $\lambda_{m,t-1}$ is known as the market price of risk. The conditional covariances between $r_{i,t}$ and $\pi_{l,t}$ represent both inflation and currency risk that stem from PPP violations. Specifically, $\lambda_{\pi,l,t} Cov_{t-1}(r_{i,t}, \pi_{l,t})$ is the inflation premium that the investor demands for the co-movement between the asset's nominal return and the inflation in the l^{th} country.

As most practitioners do, this model assumes that inflation in a certain country is non-stochastic[4] (see Dumas and Solnik (1995) and De Santis and Gerard (1997), for instance). In such a world, PPP deviations are precisely reflected in exchange rate changes. This is a plausible assumption given the

[4] If inflation is stochastic then the model can be expressed in such a way that the expected returns are dependent on three premiums, namely, market, currency and inflation. See Moerman and van Dijk (2006), for such a variation of the model.

fact that the fluctuations of inflation are negligible as compared to exchange rate fluctuations (Cappiello et al., 2003). Then, the only random component in $\pi_{x,t}$ is currency risk and the changes in PPP deviations are identified with exchange rate changes. Accordingly, $\lambda_{\pi,l,t-1}$ can be denoted as $\lambda_{x,l,t-1}$ and identified as the currency price of risk associated with the currency in l^{th} country. Thus seen, Equation 1 consists of L number of currency premiums ($\sum_{l=1}^{L}\lambda_{x,l,t-1}Cov_{t-1}(r_{i,t}, r_{x,l,t})$) that stem from the covariances between the asset's returns and the changes in the exchange rates in L number of other countries. Empirically, this large number of currency premiums may be limited to a few (see De Santis and Gerard, 1998, among others).

A more parsimonious version of the above model can be obtained by adopting the L number of exchange rates between the country to which the asset belongs and L number of countries in terms of a single trade-weighted exchange rate which may represent a weighted average of the relationships between the currency in question and the other currencies (Giurda and Tsavalis, 2004). Then L number of currency premiums can be replaced by a single currency premium as follows:

$$E_{t-1}(r_{i,t}) = \lambda_{m,t-1}Cov_{t-1}(r_{i,t}, r_{m,t}) + \lambda_{x,t-1}Cov_{t-1}(r_{i,t}, r_{x,t}) \quad (2)$$

where $r_{x,t}$ is the change in the relevant trade-weighted exchange rate in time t. The ICAPM relationship represented by Equation 2 can also be expressed as follows:

$$E_{t-1}(r_{i,t}) = \beta_{m,t-1}E_{t-1}(r_{m,t}) + \beta_{x,t-1}E_{t-1}(r_{x,t}). \quad (3)$$

where $\beta_{m,t-1}$ and $\beta_{x,t-1}$ are market beta and the exchange rate exposure beta, respectively. $\beta_{m,t-1}$ measures the asset's exposure to market risk while $\beta_{x,t-1}$ measures its exposure to currency risk. Viewed from this perspective, the time-varying nature of the second moments makes both betas time-varying. The intuition is that, while the expected returns on an asset is proportional to market returns and exchange rate changes, depending on the conditioning information that is publicly available at time t-1, the proportionality factors (market and exchange rate exposure betas) themselves are also time-varying. In other words, the investors are sensitive to "the new information that periodically becomes available to [them], who then use it to adjust their investment strategies".

Within the above framework, we focus on the return on financial assets in a few emerging markets. Return on the relevant country stock index is assumed to be a reasonable proxy for the return on an asset in that country. ICAPM relationship represented by Equation 2 is applied to each country. Viewed from this perspective, return on a certain country stock index can be explained in terms of the covariance between the returns and the return on the world market portfolio and the covariance between the returns and the changes in the selected trade-weighted exchange rate. Obviously, this is a too simplified specification of the original Adler and Dumas (1983) model. However, we may neglect the information loss that may stem from making the model simple, as our main objective is to derive exchange rate exposure betas, but not to test the validity of Adler and Dumas (1983) version of ICAPM[5].

Assuming that the market returns and exchange rate changes are not necessarily orthogonal, we suggest the following parsimonious version of ICAPM-related mean structure for the purpose of deriving time-varying exchange rate exposure betas:

$$E_{t-1}(r_{i,t}) = \lambda_{X,t-1}Cov_{t-1}(r_{i,t}, r_{x,t}) + \lambda_{M,t-1}Cov_{t-1}(r_{i,t}, r_{m,t}) \quad (4)$$

$$E_{t-1}(r_{x,t}) = \lambda_{X,t-1}Var_{t-1}(r_{x,t}) + \lambda_{M,t-1}Cov_{t-1}(r_{x,t}, r_{m,t}) \quad (5)$$

$$E_{t-1}(r_{m,t}) = \lambda_{X,t-1}Cov_{t-1}(r_{m,t}, r_{x,t}) + \lambda_{M,t-1}Var_{t-1}(r_{m,t}) \quad (6)$$

where $r_{i,t}$ is return on country i's stock index at time t; $r_{m,t}$ is return on the world market portfolio at time t; $r_{x,t}$ is the change in bilateral nominal exchange rate between the US dollar and the currency of country i at time t; $\lambda_{M,t-1}$ is market price of risk; and $\lambda_{X,t-1}$ is the currency price of risk. Since we allow for non-orthogonality between market returns and exchange rate changes, a non-

[5] See De Santis and Gerard (1998) and Cappiello et al (2003), for attempts to test the validity of ICAPM using a set of countries and a number of relevant exchange rates.

zero $Cov_{t-1}(r_{m,t}, r_{x,t})$ term enters into mean equations 5 and 6.

Contrary to the common practice, we do not convert the returns into a common/reference currency. Returns on each country index are measured in the relevant local currency. We also select a value-weighted world market index which is not converted into a common/reference currency. The purpose of this exercise is to obtain country level portfolios and a world market portfolio which represent "the theoretical performance of an index without any impact from foreign exchange fluctuations" (MSCI, 1998). Our reluctance to convert returns on country indexes and the world market index into a common currency is due to a few theoretical and empirical reasons. First, it helps us separate market risk from currency risk. As Giannopoulos (1995) argues, these two risks are not additive and conversion of various country stock index returns into a common currency will have an adverse impact on their volatility. Second, conversion of country index returns (the dependent variable) into US dollars using the exchange rate chosen would lead to inaccurate exposure coefficients because the changes in the same exchange rate is an independent variable in the regression[6]. Third, conversion of the returns on a world market index denominated in a common currency (mostly in US dollar) into local currency might have resulted in an unaffordable degree of multicolinearity between the two regressors. In addition to the resultant inefficient parameter estimates, it would also lead to unrealistic estimates of exchange rate exposure beta.

3. ECONOMETRIC METHODOLOGY

In order to derive time-varying exchange rate exposure betas with time-varying second moments, we turn to multivariate GARCH-type models. More specifically, assuming *constant prices* for market risk and currency risk[7] and that the market returns and exchange rate changes are not necessarily orthogonal, we employ a trivariate BEKK-GARCH (*p, q, K*)-M model. Following Jayasinghe and Tsui (2008), we use the following model structure to derive time-varying exposure betas:

$$r_{j,t} = \lambda_{0,j} + \lambda_X h_{X,t} + \lambda_M h_{M,t} + \theta_i \varepsilon_{i,t-1} + \varepsilon_{i,t} \qquad (7)$$

$$j = i, m, x$$

$$z_t = \varepsilon_t H_t^{-1/2} \qquad (8)$$

$$\varepsilon_t \mid I_{t-1} = (\varepsilon_{i,t} \varepsilon_{m,t} \varepsilon_{x,t})' \mid I_{t-1} \sim N(0, H_t)$$

$$H_t = C'C + A'\varepsilon_{t-1}\varepsilon'_{t-1}A + B'H_{t-1}B \qquad (9)$$

$$H_t = \begin{bmatrix} H_t^{uu} & H_t^{uv} \\ H_t^{vu} & H_t^{vv} \end{bmatrix} \qquad (10)$$

$$\mathrm{B} = \left[H_t^{vv}\right]^{-1} H^{vu} \qquad (11)$$

where $r_{j,t}$ is 3 x 1 vector that consists of three elements: return on country index at time t ($r_{i,t}$), return on world market portfolio at time t ($r_{m,t}$) and changes in bilateral nominal exchange rate between the US dollar and the currency of country i at time t ($r_{x,t}$)[8]. An intercept and a MA(1) term is included in each of the three mean equations in order to capture any remaining risk or market inefficiencies. ε_t is a 3 x 1 vector of residuals from mean equations in (6), which are assumed to be normally distributed with mean 0 and variance H_t, which is 3 x 3 variance covariance matrix. And, $\varepsilon_t \mid I_{t-1}$ denotes the vector of random shocks at time *t* given all available information at time (*t*-1). In addition, $h_{X,t}$ and $h_{M,t}$ are 3 x 1 vectors that consist of the elements in the second and the third

[6] Strictly speaking, this exercise is done using the exchange rates moderated by the base year rate. However, there exists a strong correlation between a series converted using the moderated exchange rates and a series converted using the current rates.

[7] Constant prices can be justified on the grounds that the suggested model is just a data generating process to obtain time-varying market and exchange rate exposure betas in terms of time-varying second moments.

[8] The exchange rate is expressed as the local currency price of foreign currency and an increase implies a depreciation of the relevant currency relative to the other currencies.

columns of H_t, respectively[9]. Finally, z_t denotes the standardized residuals that are assumed to be identically and independently distributed with mean 0 and variance 1.

Parameters λ_M and λ_X are market price of risk and currency price of risk, respectively. C is an upper triangular 3 x 3 matrix that contains the constants in conditional variance and covariance equations. Both A_{kl} and B_{kn} are 3 x 3 parameter matrixes. Although the use of a trivariate model offers the opportunity to capture the interdependence between the volatilities in terms of non-zero off-diagonal terms in parameter matrices, we make the restrictive assumption that parameter matrixes A_{kl} and B_{kn} are diagonal for two reasons. First, the full BEKK formulation is less parsimonious and computationally tedious[10]. Second, as the results of diagnostic tests reported in Section 5 show, the suggested diagonal version of the model sufficiently captures the non-linearities in stock returns and exchange rate changes. For parsimony, we set $K = 1$. As a residual analysis based on Ljung-Box statistic reveals, the optimal lag orders for GARCH and ARCH terms are as follows: $p = 1$, $q = 1$.

Following Lim (2005), time-varying betas can be obtained through Equations 10 and 11. In Equation 10, H_t^{uu}, H_t^{vv} and H_t^{vu} are the conditional variance-covariance matrixes of the assets to be priced, the factors with which the assets are priced, and between the assets and factors, respectively.

Assuming that the standardized residuals of the suggested trivariate GARCH model are conditionally normally distributed, the conditional log-likelihood of residual vector ε_t at time *t* can be defined as follows:

$$\ell(\varphi)_t = -\frac{1}{2}\ln(2\pi) - \frac{1}{2}\ln|H_t| - \frac{1}{2}\varepsilon_t' H_t^{-1}\varepsilon_t \quad (12)$$

The log-likelihood function of the sample is obtained as $L(\varphi) = \sum_{t=1}^{T} \ell(\varphi)_t$, where T is the number of observations. The parameter vector φ of the trivariate BEKK-GARCH(1,1,1)-M model is estimated by maximizing L with respect to φ. In order to accommodate the non-normal features reflected in the basic statistics of country returns and the exchange rate changes, all estimates of the parameters are obtained through the quasi-maximum likelihood (QML) estimation method proposed by Bollerslev and Wooldridge (1992). Under certain regularity conditions, the QML estimate is assumed to be consistent and asymptotically normal. Therefore, statistical inference can be drawn due to robust standard errors. Required computer programs are coded in GAUSS and use BHHH algorithm to compute QML estimates.

4. DATA

The sample consists of three emerging economies: Korea, Taiwan and Thailand. We use weekly closing stock prices obtained on Wednesdays for the period from 30th December 1998 to 30th Dec 2006[11]. The resultant sample period provides us with 418 observations. All stock indexes are from Morgan Stanley Capital International (MSCI) and extracted from Datastream. Country level portfolios are represented by MSCI country indexes measured in relevant local currency. World market portfolio is represented by the MSCI world market index MSWRLDL. It is a value-weighted world market index which is not converted into a common/reference currency and, therefore, free from exchange rate fluctuations (Giannopoulos, 1995; MSCI, 1998). All trade-weighted exchange rates are from J P Morgan and extracted from Datastream.

Continuously compounded weekly returns and exchange rate changes are calculated as follows:

[9] Since we allow for non-orthogonality between market returns and exchange rate changes, a non-zero $h_{xm,t}$ term enters into mean equations for market returns and exchange rate changes.

[10] In our initial round of regressions, we found that the full BEKK model did not converge in some cases.

[11] The currency crisis period is excluded from the sample in order to avoid the impact of unusual currency moments. Though we attempted to include Singapore and Australia in the sample, the exchange rate exposure coefficients of the country level stock returns of these economies are not significant for weekly return horizons.

$$r_{j,t} = \ln\left(R_{j,t} \big/ R_{j,t-1} \right) * 100 \qquad j = i, m, x$$

where $R_{j,t}$ and $R_{j,t-1}$ are the stock prices/exchange rates for the week t and (t-1) respectively. *i, m* and *x* denote the country in question, world market portfolio and the relevant exchange rate, respectively.

All return series show excess kurtosis which ranges from the lowest 0.4755 (Korea) to the highest 1.9398 (Thailand)[12]. Jarque-Bera statistic is high in all cases except for Korea. Exchange rate changes show somewhat higher excess kurtosis ranging from 2.5770 (Taiwan) to 5.6381 (Thailand). In all three cases, excess kurtosis of the exchange rate changes is greater than that of the returns on the relevant country index. The high Jarque-Bera statistic together with excess kurtosis in some cases implies that the exchange rate changes are not normally distributed. The non-normal features of both country stock returns and exchange rate changes justify the use of QML method of estimation.

As evidenced by the augmented Dicky-Fuller test, continuously compounded returns on all country indexes and the world market index and exchange rate changes are stationary. The Ljung-Box test for returns evaluated at 20 lags ($Q\ (20)$) reveals that, except for the US, there are no linear dependencies. Exchange rate changes for all four economies show the same pattern and are free from linear dependencies. In addition, the Ljung-Box test for squared returns evaluated for 20 lags ($Q^2(20)$) displays that all return and changes in exchange rate series possess a great deal of non-linear dependencies. This provides some empirical support for the use of GARCH-type models to derive time-varying exchange rate exposure betas.

5. EMPIRICAL FINDINGS

In this section, we first report the results of a few tests that show the likeliness of time-varying or unstable parameters in the selected sample. After this pre-estimation assessment, we move onto deriving time-varying exchange rate exposure betas and some diagnostic checks for adequacy of the proposed model to derive exposure betas. It is followed by a brief investigation of the stochastic structure of time-varying exposure betas.

5.1. SOME PRE-ESTIMATION RESULTS

We use a battery of tests to show that countries selected in the sample are more likely to possess time-variant (unstable) exchange rate exposure betas. All tests are based on OLS estimation of the conventional augmented market model that is widely used to estimate exchange rate exposure[13]. The first such test is the cumulative sum of squared recursive residuals (CSSRR) test suggested by Brown et al. (1975). The CSSRR test is performed at the 5% level of significance. During the sample period, the CSSRR crosses the critical value boundaries in all cases, thus suggesting the underlying parameter instabilities[14]. White's (1980) test for unconditional heteroskedasticity and ARCH-LM test for conditional heteroskedasticity are also used to diagnose possible parameter instabilities. White's test statistic is significant in all cases except for Korea at 5 degrees of freedom at the significance level 1%, suggesting the presence of unconditional heteroskedasticity. ARCH-LM test statistic for 4 lags is significant for all the cases at the significance level 1%. Results from all three tests suggest that the constant parameters in the specification represented by Equation 2 are highly improbable.

5.2. TIME-VARYING EXCHANGE RATE EXPOSURE BETAS

The maximum likelihood estimates for the suggested trivariate BEKK-GARCH(1,1,1)-M model are reported in Table 1. According to ICAPM reasoning, the market price of risk (λ_M) must be positive and the same for all countries. However, there is no such restriction for the currency price of risk (λ_X). As the estimation results reported in Panel A indicate, λ_M is positive and does not vastly vary across countries. More specifically, λ_M is not statistically significant in any of those cases. Unlike market price of risk, λ_X varies remarkably across countries between the range 0.1623 (Korea) and 0.5657 (Taiwan). Since

[12] To conserve space, results are not shown here.

[13] This refers to the constant parameter version of the regression equation in (2).

[14] To conserve space, results are not shown here.

the relevant exchange rate varies across countries, this variation in the parameter can be understood. In all cases, currency price of risk is also not statistically significant[15].

Furthermore, all GARCH terms (denoted by b_j for $j = i, m, x$) are highly significant, suggesting that the conditional variances are highly correlated to the past conditional variances. All ARCH terms (denoted by a_j for $j = i, m, x$) is also significant except in one case suggesting the presence of volatility clustering in both stock and exchange rate markets of all countries.

Table 1: Maximum likelihood estimates for the trivariate diagonal BEKK GARCH(1,1,1)–M model

Coeff	Korea	Taiwan	Thailand
λ_M	0.0094	0.0091	0.0218
	(0.31)	(0.20)	(0.63)
λ_X	0.1623	0.5657	0.2659
	(0.72)	(0.82)	(1.00)
b_i	0.9812*	0.9661*	0.9776*
	(211.38)	(5.94)	(66.03)
a_i	0.1706*	0.2406	0.1783*
	(6.45)	(0.43)	(2.70)
b_x	0.8840*	0.8266*	0.9265*
	(13.45)	(4.02)	(12.910)
a_x	0.2874*	0.2108*	0.2697*
	(2.88)	(2.89)	(2.88)
b_m	0.9640*	0.9623*	0.9636*
	(75.17)	(65.50)	(64.84)
a_m	0.2415*	0.2501*	0.2423*
	(5.31)	(5.03)	(5.25)

Notes: Reader is referred to the set of Equations 7 - 9 for the relevant model; *t*-values are in parenthesis; * indicates the significance at %5 level.

[15] These results are consistent with the previous findings in the literature. For instance, De Santis and Gerard (1998) and Cappiello et al. (2003) also find that both market and currency premiums are insignificant as long as the prices are not allowed to be time-variant.

The diagnostic checks for the estimated model, results of which are not shown here to conserve space, reveal that linear and non-linear dependencies have been adequately captured by the proposed trivariate BEKK GARCH(1,1,1)-M model. Ljung-Box statistics for standardized and squared standardized residuals evaluated for 20 lags ($Q(20)$ and $Q^2(20)$, respectively) are not only well below the critical value of 31.481 at the 5 % level, but are significantly low as compared to those of the return series. These results imply that the suggested model adequately filters linear and non-linear dependencies and is appropriate in deriving reliable estimates of time-varying exchange rate exposure betas.

5.3. THE STOCHASTIC STRUCTURE OF EXCHANGE RATE EXPOSURE BETAS

Table 2 compares the mean values of time-varying exchange rate exposure betas and their OLS point estimates. The mean value of each series is reasonably close to the relevant OLS point estimate.

Exposure beta associated with the relevant trade-weighted exchange rate is positive in all three emerging markets. The intuition is that importers or investors whose consumption basket consists of a lot of imported goods from these countries cannot hedge against currency risk by investing in stocks in those markets. This is because returns on stocks in those countries are positively correlated with the depreciation of local currency (or appreciation of the importers' currency). However, the exporters to these economies can hedge against currency risk by investing in stocks in these emerging markets for the same reason.

Table 2: Comparison between OLS point estimates of betas and the mean values of time-varying betas

Country	Market beta		Exchange rate exposure beta	
	OLS β_m	Mean of $\beta_{m,t}$	OLS β_x	Mean of $\beta_{x,t}$
Korea	0.8807	1.0516	1.2112	0.9318
Taiwan	0.7084	0.8813	2.0817	1.7347
Thailand	0.6227	0.7616	1.6996	1.4858

Table 3: Preliminary statistics of time-varying betas

Panel A: Exchange rate exposure betas

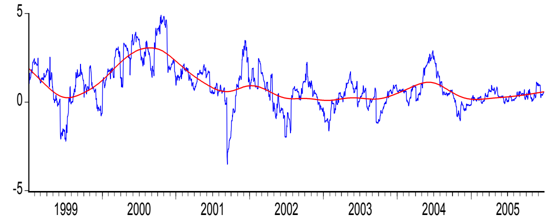
Coefficient	Korea	Taiwan	Thailand
Mean	0.9318	1.7347	1.4858
Maximum	2.2004	5.3367	4.9305
Minimum	-0.0916	0.0310	0.0439
S D	0.3885	0.7447	0.7195
Skewness	0.3480	2.1149	1.7795
Kurtosis	3.1422	9.4108	9.0372
J-B stat	8.74	973.33	851.34

Panel B: Market betas

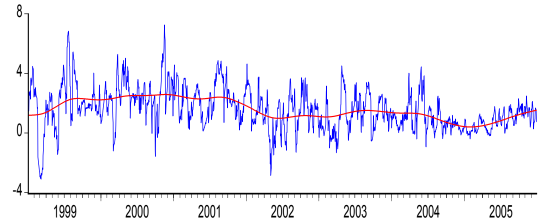
Coefficient	Korea	Taiwan	Thailand
Mean	1.0517	0.8823	0.7616
Maximum	1.6052	1.3209	1.6454
Minimum	0.4760	0.2796	0.2501
S D	0.2322	0.2019	0.3049
Skewness	-0.2576	-0.3678	0.5402
Kurtosis	2.4129	2.8611	2.7460
J-B stat	10.57	9.24	21.35

Summary statistics of exchange rate exposure betas are reported in Panel A of Table 3. Standard deviation of exposure beta series ranges from 0.3885 for Korea to 0.7447 for Taiwan. All exposure beta series are positively skewed and leptokurtic. A comparison between Panel A and B in Table 3 indicates that the standard deviation of exchange rate exposure beta is usually higher than that of market beta. In all three cases, kurtosis of each exposure beta distribution is also always higher than the kurtosis of its counterpart market beta distribution. This suggests that an exposure beta distribution tends to have more outliers than the outliers in its counterpart market beta distribution.

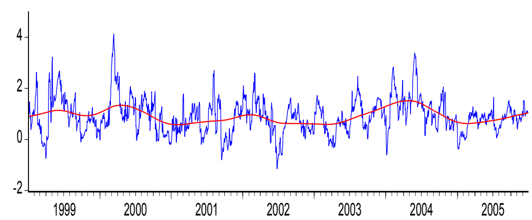
Figure 1 provides a visual glimpse of all time-varying exposure betas. As we use weekly data, the estimates of time-varying betas may be "still volatile and inevitably subject to estimation error" (De Santis and Gerard, 1998). As such, the Hodrick-Prescott filtered trends of betas are also included in Figures 1. Each exposure beta seems to fluctuate within a wide range.



(a) Korea



(b) Taiwan



(c) Thailand

Figure 1: Time-varying exchange rate exposure betas

Next, we examine whether the exchange rate exposure betas are mean-reverting and stationary. Widely used semi-nonparametric Gewek and Porter-Hudak (1983) test is employed for this purpose[16]. In order to see the sensitivity of the

[16] The test is based on the following spectral regression equation: $\ln I(\omega_s) = c + \phi \ln\left(4\sin^2(\omega_s/2)\right) + \zeta$ for $s = 1,2,......n(T)$ where T is the number of observations in the series concerned; $I(\omega_s)$ is the periodogram of a series at harmonic frequency $\omega_s = (2\pi\, s/T)$ with $s = 1,2,......T-1$; ζ is random error; n represents the number of low frequency ordinates and is usually determined as $n = T^{\alpha}$. OLS estimation of ϕ provides a consistent estimate of $-d$ in the ARFIMA process $\Phi(L)(1-L)^d y_t = \Theta(L)\upsilon_t$ where $\upsilon_t \sim (0,\sigma^2)$.

estimates of the fractional difference parameter d to the choice of α, three values of α are used here: 0.50, 0.55 and 0.60. First, we perform a one-sided test to check the validity of the null hypothesis of $d = 0$ against the alternative of $d > 0$. The null is rejected at the 5 % level for all cases suggesting that all exposure beta series are more likely to be represented by an ARFIMA process. Then a second one-sided test is performed for the null hypothesis of $d = 1$ against the alternative of $d < 1$. Test results for exchange rate exposure beta are reported in Table 4. The null is rejected at the 5 % level for Korea and Taiwan under all three α values. However, null is accepted under all three α values for Thailand suggesting the presence of unit roots. Korea shows a difference parameter d that is less than 0.5 when $\alpha = 0.5$ and $\alpha = 0.55$. Taiwan shows a difference parameter d that is less than 0.5 only when $\alpha = 0.5$.

These results from the GPH test have a few important implications. First, all time-varying exchange rate exposure beta series consistently reject the hypotheses $I(0)$. It suggests that all exposure betas in the sample are characterized by a $I(d)$ process with $0 < d < 1$ or a unit root process.

Table 4: GPH test results for time-varying exchange rate exposure betas

Country	Value of difference parameter d		
	$\alpha = 0.50$	$\alpha = 0.55$	$\alpha = 0.60$
Korea	0.3331*	0.3314*	0.5932*
	(-2.74)	(-3.74)	(-2.51)
Taiwan	0.4638*	0.6836*	0.5896*
	(-4.07)	(-2.28)	(-3.68)
Thailand	0.8189	0.8275	0.7914
	(-0.89)	(-1.03)	(-1.67)

d represents ϕ in the regression equation $\ln I(\omega_j) = c + \phi \ln\left(4 \sin^2(\omega_j/2)\right) + \zeta$; *t*-statistics are in parentheses; * indicates the significance at least at the 5% level

Second, in the case of Korea and Taiwan, hypothesis $I(1)$ is rejected. The fact that the difference parameter d for exposure beta series in those two countries is greater than 0 and less than 1 implies that they are mean-reverting. However, the impact of a shock on exposure betas is likely to decay hyperbolically, which is much slower than a rapid geometric decay represented by a standard ARMA process. Mean-reverting exchange rate exposure betas have both theoretical and empirical implications. First, since returns are linear functions of betas (that represent exposure to market risk, currency risk or any other risk), it is argued that mean reverting betas is an essential element in assuring the stationarity of returns. Second, the absence of mean reversion makes the notion of equilibrium have little relevance even in the long-run (Lai, 1997). Third, mean reverting exposure betas imply that these coefficients can be used for forecasting purposes. This may be extremely important news in hedging against currency risk. However, the hypothesis $I(1)$ is accepted for Thailand. Implication is that the exposure beta of Thailand is characterized by a unit root process and is non-stationary.

Third, though exposure betas of Korea and Taiwan are mean-reverting, they are more likely to show covariance non-stationary dynamics.

6. CONCLUDING REMARKS

We have used a trivariate BEKK-GARCH(1,1,1)-M model based on a conditional ICAPM framework to obtain time-varying exchange rate exposure betas. Our approach does not require some prior understanding of the determinants of the time-variation of exposure beta to obtain the estimates of the same. Also, the suggested approach is more appropriate than the GARCH-based methods that use inappropriate mean structures in deriving time-varying betas. As the mean structure in such models does not represent a relevant ICAPM, the resultant information loss may lead to inaccurate estimates.

In deriving time-varying exchange rate exposure betas, we emphasize the necessity of taking the non-orthogonality between exchange rate changes and market returns into account. A portion of exchange rate exposure is always captured by market beta and is also priced under the label of market risk. What counts for the decisions of firms and investors is the portion of exchange rate exposure that is not captured by market beta and hence not priced under the market risk. As such, the estimated time-varying exchange rate exposure betas are more reliable than those in the studies that does not take this non-orthogonality into account.

The examination of the stochastic structure of the time-varying exchange rate exposure betas offers some useful insights. As evidenced by their basic statistics, exposure betas are usually more volatile than market betas. And, as compared to the market beta series for each economy, exposure beta series tend to have more outliers. Results from the GPH test reveal that exposure betas of Korea and Taiwan are long-memory processes characterized by fractional integration. Exposure beta in both cases turn out to be mean-reverting, though their mean reverting dynamics could be highly persistent and display a slow hyperbolical decay. As for the covariance stationarity, however, we do not obtain unambiguous results and the matter is left for future research. Exposure beta of Thailand is characterized by a unit root process and non-stationary.

Findings of this study support and strengthen the findings of Jayasinghe and Tsui (2008) which also conclude that exchange rate exposure betas in a sample of nine emerging and developed markets are more likely to be long memory processes though they are mean-reverting. However, the present study goes a step further than the former due to its selection of more appropriate proxies for the variables in question. More specifically, exchange rate exposure beta of a certain country in the present study is based on weekly data and a trade-weighted exchange rate. The use of a trade-weighted exchange rate is more appropriate in the context of the conceptual framework provided by the ICAPM. This is because, in ICAPM, the return on a certain asset partly depends on the covariance between the return and the change in a large number of exchange rates, which can be replaced with the change in a single currency index that represents the weighted average of all exchange rates. Exchange rate exposure beta of a certain country in Jayasinghe and Tsui (2008) is based on daily data and a single bilateral exchange rate between the US dollar and the currency in the country to which the asset belongs. There may be a possible information loss in such a framework as the covariance between return and changes in other exchange rates are not taken into account.

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