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Is the Currency Risk Priced in Equity Markets?

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Abstract

In this paper we investigate whether the currency risk is priced in international stock markets. We suggest a parsimonious version of the international capital asset pricing model with an EGARCH-M(1,1) specification of the second moments' dynamics of stock and currency returns, assuming that the latter follow a multivariate t -distribution. This specification allows for asymmetric responses of volatility to stock and currency news, including leverage effects. Our results suggest that the currency risk is priced in international stock markets, once asymmetries in volatility are taken into account. The currency premium is found to be significant on both statistic and economic grounds. We find that a dynamic portfolio strategy that hedges against currency changes provides higher returns (as a reward for currency premium) than a strategy which ignores them.

JEL Classification: C32, C52, C53, G11, G12

Keywords: International asset pricing, currency risk, multivariate EGARCH, density forecast, dynamic hedging strategies

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1. Introduction

Although it is well known that investors can earn significant benefits from international diversification, investing in foreign markets entails exposure to currency risks for which investors need to be compensated. This source of risks can be attributed to substantial deviations from the purchasing power parity (PPP), and thus are also referred to as real exchange rate risks. If PPP does not hold, the real return of any asset differs across countries. Then, the standard (domestic) capital asset pricing model (CAPM) does not constitute a correct model to price the expected return of any stock traded in international markets. In this situation, an international CAPM (ICAPM), in addition to the market covariance risk, should include currency covariance risks [see Adler and Dumas (1983), and Dumas and Solnik (1995), *inter alia*].

Recently, there is a growing research effort to examine whether currency risk is priced in international stock markets.¹ The answer to this question has important implications for portfolio management and hedging strategies, as any source of risk which is not compensated in terms of expected returns should be hedged. Despite the plethora of empirical studies (see fn 1), evidence does not give a clear cut answer whether or not the currency risk is priced. Recently, De Santis and Gerard (1997, 1998), and Capiello, Castrén and Jääskelä (2003) based on a multivariate-generalized autoregressive conditional heteroscedasticity in mean (GARCH-M) econometric specification to estimate a conditional version of the ICAPM, show that the currency risk is priced once allowance is made for time-varying parameters of risk prices. The inadequacy of the GARCH-M model with constant parameters to price the currency risk may be attributed to a mis-specification of the GARCH-M to fully capture the dynamics of the conditional variances and covariances of stock and currency returns. In particular, the GARCH specification used in the above studies does not allow for leverage effects, or any other asymmetries, on stock and currency volatilities (and, hence, on risk premia). In addition,

¹ Jorion (1991), Chan, Karoly and Stulz (1992), Dumas and Solnik (1995), De Santis and Gerard (1997, 1998), De Santis, Gerard and Hillion (1999, 2003), Capiello and Fearnley (2000), Nilson (2002) and Capiello, Castrén and Jääskelä (2003), *inter alia*.

it relies on the assumption that the stock and currency returns are normally distributed. The latter may lead to overrejection of the ICAPM model, if the returns' distributions have fat tails [see Zhou (1993), Campbell and Zhou (1993), and Kan and Zhou (2003)]. To this end, in our analysis we will replace the normality assumption with a t -distribution.

In this paper, we re-examine the validity of the ICAPM model to price market and currency risk by adopting Nelson's (1991) exponential GARCH-M (EGARCH-M) econometric specification for the conditional second moments of the returns which allow for asymmetric effects of market news on the volatility function. Ignoring these effects may explain the evidence of time-varying currency and market prices of risk found by De Santis and Gerard (1997,1998). Using weekly data for four developed stock markets (Germany, Japan, UK and US) and the world market from 1990 to 2002, the paper provides clear cut evidence that both market and currency risk premia are priced in international markets. Our results show that a significant part of the above premia can be attributed to currency news (or crises) occurred in the nineties. These seem to have influenced both the market and currency premia.

To evaluate the performance of the ICAPM with the EGARCH-M specification of the second moments, the paper conducts two exercises. First, it examines the ability of the model to forecast the densities of future expected returns on statistic basis. This is done based on tests which account for the effects of higher dynamics of stocks and currency returns on the forecasting performance of the model. The aim of the second exercise is to examine the validity of the ICAPM on economic grounds. This is done by comparing two different portfolio investment strategies: one which hedges against currency changes and another which ignores them. The results of these two exercises support the EGARCH-M(1,1) specification of the ICAPM on both statistic and economic grounds.

The paper is organized as follows. Section 2 introduces a parsimonious version of the ICAPM, with one-index (factor) specification for the currency risk, and presents the econometric framework. Section 3 carries out the estimation and discusses the results.

Section 4 evaluates the statistic and economic performance of the model. Section 5 concludes the paper.

2. The model

In this section we present a parsimonious version of the conditional ICAPM allowing for market and currency risk premia. This model is in the spirit of Merton's (1973) two-factor asset pricing model and Adler and Dumas' (1983) version of the ICAPM, which hedges against adverse changes in exchange rates. These changes can be attributed to short run systematic purchasing power parity (PPP) deviations (real exchange rates changes), and they can affect investors' invested wealth in international stock markets.²

Under the above assumptions, in equilibrium we can write the expected excess return of a stock i , denoted as $r_{i,t}$, conditional on the current market information set Ω_{t-1} as

$$E_{t-1}(r_{i,t}) = \lambda_M Cov_{t-1}(r_{i,t}, r_{M,t}) + \lambda_C Cov_{t-1}(r_{i,t}, c_t), \quad i = 1, 2, \dots, N \quad [1]$$

where $E_{t-1}(\cdot)$ and $Cov_{t-1}(\cdot)$ denote the conditional on Ω_{t-1} mean and covariance of stock returns, respectively, $r_{M,t}$ denotes the return on the world market portfolio of all traded stocks in international stock markets (denoted as N), c_t represents the rate of return of a single currency factor (index) driving the real exchange rate changes, denoted as $c_{j,t}$, of the US dollar (reference currency) against the exchange rates of foreign countries, denoted by j .³

The asset pricing model given by equation (1) claims that the expected excess return of any stock in the international stock markets must consist of two sources of risk premia.

² The PPP is assumed to hold as a long run relationship [see Culver and Pappel (1999), for recent evidence].

³ Single index models for nominal, or real, exchange rates have been considered by many authors [see Jorion (1991), Ferson and Harvey (1993), Bansal, Hsieh and Viswanathan (1993), Ng (2001), *inter alia*] in order to reduce the second moments of equation (1). This methodology simplifies the econometric

The first, known as market risk premium, is due to the conditional covariance of the stock return, $r_{i,t}$, with the world market portfolio return, $r_{M,t}$. This premium is measured by $\lambda_M \text{Cov}_{t-1}(r_{i,t}, r_{M,t})$, where λ_M is the market price of risk given in equilibrium as $\lambda_M = -\frac{J_{WW}W}{J_W}$, where J_W and J_{WW} are the first and second partial derivatives of the derived utility of wealth function $J(W(t), c(t), t)$ with respect to a representative investor's wealth, $W(t)$. The term $-\frac{J_{WW}W}{J_W}$ is the coefficient of relative risk aversion. Since risk aversion implies $J_W > 0$ and $J_{WW} < 0$, model (1) suggests a positive relationship between expected returns, $E_{t-1}(r_{i,t})$, and the market premium (i.e. $\lambda_M > 0$).

The second source of risk premia in model (1) comes from the covariance of the stock return with the real exchange rates index, c_t . The magnitude of this risk premium is given by $\lambda_C \text{Cov}_{t-1}(r_{i,t}, c_t)$, where $\lambda_C = -\frac{J_{WC}}{J_W}$ is the currency price of risk, where J_{WC} is the cross derivative of $J(W(t), c(t), t)$ with respect to $W(t)$ and the single currency factor, $c(t)$. This risk premium can be attributed to hedging motives of investors against real exchange rates changes (PPP deviations).⁴ If $J_{WC} > 0$ and $\text{Cov}_{t-1}(r_{i,t}, c_t) > 0$ [or $J_{WC} < 0$ and $\text{Cov}_{t-1}(r_{i,t}, c_t) < 0$], then investors will demand lower expected returns on holding international stocks. In these cases, the stocks can be thought of as natural hedging instruments against real exchange rates changes. If $J_{WC} > 0$ and $\text{Cov}_{t-1}(r_{i,t}, c_t) < 0$ [or $J_{WC} < 0$ and $\text{Cov}_{t-1}(r_{i,t}, c_t) > 0$], then investors will require higher compensation. If

estimation and testing procedure of (1), as it significantly reduces the number of second moments of (1), otherwise we may run in estimation and identification problems [see Dellaportas and Pourhahadi (2002)].

⁴ If the PPP (in its relative form) does not hold, then real returns on any asset differ across investors from different countries. To see this more rigorously, notice that the difference of the real return of a stock i between the domestic and a foreign country (denoted by star) at a first order approximation is given by $r_{i,t} - r_{i,t}^* \approx c_t - c_{t-1} \approx \Delta e_{t+1} - (\pi_t - \pi_t^*)$, where $c_t = \frac{e_t P_t^*}{P_t}$ is the real exchange rate, e_t is the nominal exchange rate, and π_t and π_t^* denote, respectively, the inflation rates of the domestic and foreign country, while P_t and P_t^* are their corresponding prices levels.

$J_{WC} = 0$, then the expected returns will solely depend on the market covariance risk. In this case, the ICAPM relationship (1) reduces to the standard international CAPM suggested by Solnik (1974).⁵

2.1 Econometric specification

To complete the specification of the conditional ICAPM we need to set up an econometric framework to represent the dynamics of the conditional second moments involved in the model, namely the covariances and variances. To this end, we adopt a multivariate-EGARCH-M specification. This is appropriate when the conditional variances (volatilities) and covariances of stock returns respond asymmetrically to positive (good) and negative (bad) news of stock or exchange rate market returns [see Nelson (1991), *inter alia*]. Note that a negative relationship of the volatility of stocks and/or currency returns with respect to market news is referred to as the leverage effect [see, for instance, Black (1976), Schwert (1989) and Braun, Nelson and Sunier (1990)].⁶

The multivariate EGARCH-M(1,1) system that we adopt for the estimation of the ICAPM consists of the following set of equations for the conditional mean of the variables $r_{M,t}$, $r_{i,t}$ and c_t :

$$r_{M,t} = \lambda_{M,0} + \lambda_M Var_{t-1}(r_{M,t}) + \lambda_C Cov_t(r_{M,t}, c_t) + \varepsilon_{M,t}, \quad [4a]$$

⁵ Note that the currency risk premium $\lambda_C Cov_{t-1}(r_{i,t}, c_t)$ is equal to the sum $\sum_{j=1}^{L-1} \lambda_j Cov_{t-1}(r_{i,t}, c_{j,t})$, where L denotes the total number of countries and λ_j is the currency price of risk associated with j real exchange rate changes. This can be easily seen by substituting the single index relationship

$$c_{j,t} = \gamma_{0,j} + \gamma_{1,j} c_t + v_{j,t}, \quad \text{with } v_{j,t} \sim IID(0, \sigma_v^2) \text{ and } Cov(r_{j,t}, v_{j,t}) = 0,$$

into $Cov_{t-1}(r_{i,t}, c_{j,t})$. The resulting equation implies that $\lambda_C = \sum_{j=1}^{L-1} \gamma_{1,j} \lambda_j$ represents the aggregate price of the individual countries prices of risk.

⁶ Note that the standard GARCH-M model can not capture the leverage effect since it assumes that conditional variances depends on the squared values of the past residuals. See Nelson (1991), for a more

$$r_{i,t} = \lambda_{i,0} + \lambda_M \text{Cov}_{t-1}(r_{i,t}, r_{M,t}) + \lambda_C \text{Cov}_{t-1}(r_{i,t}, c_t) + \varepsilon_{i,t} \quad ; \quad i=1,2,\dots,N, \quad [4b]$$

and

$$c_t = \lambda_{C,0} + \lambda_M \text{Cov}_{t-1}(c_t, r_{M,t}) + \lambda_C \text{Var}_{t-1}(c_t) + \varepsilon_{C,t}. \quad [4c]$$

where the intercepts $\lambda_{k,0}$ ($k = M, i, C$) are assigned to capture any remaining specific risks or market imperfections. The error terms $\varepsilon_{k,t}$ ($k = M, i, C$) in equations (4a)-(4c) have conditional on the information set \mathcal{Q}_{t-1} variance functions given by

$$\ln(\sigma_{k,t}^2) = \omega_k + g_k(z_{k,t-1}) + \beta_k [\ln(\sigma_{k,t-1}) - \omega_k] \quad (k = M, i, e), \quad [5]$$

where

$$g_k(z_{k,t-1}) = \gamma_k \left(|z_{k,t-1}| - E|z_{k,t-1}| \right) + \mathcal{G}_k z_{k,t-1}, \quad \text{with} \quad z_{k,t-1} = \frac{\varepsilon_{k,t-1}}{\sigma_{k,t-1}}, \quad [6]$$

captures the effects of the innovation (market news) $z_{k,t-1}$ on the conditional variance $\sigma_{k,t}^2$. The terms $\mathcal{G}_k z_{k,t-1}$ and $\gamma_k \left(|z_{k,t-1}| - E|z_{k,t-1}| \right)$ in the innovation function $g_k(z_{k,t-1})$ allow the conditional variance $\sigma_{k,t}^2$ to respond asymmetrically to positive and negative returns in terms of sign and magnitude. In particular, the $\mathcal{G}_k z_{k,t-1}$ term allows for leverage effects. When $\mathcal{G}_k < 0$, $\sigma_{k,t}^2$ tends to rise (fall) following market bad (good) news. When $\mathcal{G}_k < 0$ and $\gamma_k > 0$, the term $\gamma_k \left(|z_{k,t-1}| - E|z_{k,t-1}| \right)$ implies that the magnitude of the leverage effect is larger than expected. To complete the EGARCH specification of the second moments, we assume that the conditional covariances are calculated as $\sigma_{ks} = \rho_{ks} \sigma_{k,t} \sigma_{s,t}$, for $k \neq s$. This assumption implies that the correlation coefficients of the disturbance terms $\varepsilon_{k,t}$ are constant. It is made for estimation reasons, in order to restrict the number of unknown parameters of the multivariate EGARCH, otherwise the model may be over parameterized [see Bollerslev (1990), *inter alia*].

complete discussion about the potential benefits of the EGARCH model, compared with the GARCH model.

3. Empirical Analysis

3.1 Data

We use continuously compounded weekly returns (in excess of the risk-free rate) on stock indices for the four largest markets: United States (US), United Kingdom (UK), Germany (GE) and Japan (JP), and on the world market portfolio index. All stock indices are from Morgan Stanley Capital International (MSCI) and are measured in terms of US dollar (USD), which is the reference currency. To calculate the excess returns on the indices, we use the one-month US Treasury bill rate as a risk-free rate.

Our data cover the period from January 5th, 1990 to August 16th, 2002.⁷ During this period, a number of turbulent events occurred in currency markets: the Gulf War in 1990-1991, the turmoil of the European Exchange Rate Mechanism (EMS) in 1992-1993, the Peso crisis in 1994, the Asian crisis of 1997-1998, the Russian crisis in August 1998 and the LTCM collapse in October of the same year. In addition to these, our sample covers recent turbulent stock market events, such as the burst of the equity bubble, began in the spring of year 2000, the terrorist attack of September 11th, 2001 and the collapse of the Enron and WorldCom corporations. All the above events are expected to substantially influence both market and currency risk premia.

To calculate the return on the currency index, c_t , we use log changes of an index calculated as the weighted average of the British Pound (GBP), Deutsche Mark (DEM) and Japanese Yen (JPY) nominal exchange rates, measured as the US dollar price per unit of foreign currency [see also Jorion (1991)].⁸ The weights employed to construct the exchange rate index correspond to those used to construct the New York Board of Trade

⁷ With weekly observations, we can mitigate potential biases on our estimates arisen from non-synchronous trading while, at the same time, we have high enough frequency of data to obtain a better picture of the movements of markets returns.

⁸ This can be justified by the very high estimates of the correlation coefficients between nominal exchange rates and the currency index return.

(NYBOT) US dollar index.⁹ These weights adjust the nominal exchange rates according to the trade competitiveness of each country, and thus can capture the effects of short-term real exchange changes (PPP deviations) on the currency index.

Summary statistics for all series are given in Table I, see Panels A and B. Panel A reports sample estimates of the unconditional mean, standard deviation, the coefficients of skewness and excess (over the normal) kurtosis and the LM statistic for ARCH effects, with five lags. Panel B reports the correlation coefficients among all series.

Table I: Summary statistics

Panel A: Descriptive statistics									
	World	US	UK	Germany	Japan	FXI			
Mean	-0.03	0.06	-0.01	-0.05	-0.21	-0.01			
Std. Dev	1.97	2.18	2.26	2.99	3.34	1.21			
Skewness	-0.25 (0.01)	-0.50 (0.00)	0.17 (0.08)	-0.35 (0.00)	0.33 (0.00)	0.09 (0.35)			
Kurtosis	2.19 (0.00)	3.05 (0.00)	1.90 (0.00)	2.22 (0.00)	1.21 (0.00)	0.87 (0.00)			
ARCH(5)	41.11(0.00)	48.73 (0.00)	23.50 (0.00)	91.46 (0.00)	36.86 (0.00)	15.15 (0.01)			
Panel B: Correlation coefficients									
	World	US	UK	Germany	Japan	GBP	DEM	JPY	FXI
World	1.00	0.84	0.72	0.72	0.61	0.09	0.09	0.17	0.15
USA		1.00	0.50	0.50	0.23	-0.08	-0.10	-0.07	-0.11
UK			1.00	0.62	0.34	0.35	0.16	0.10	0.23
Germany				1.00	0.33	0.20	0.30	0.13	0.28
Japan					1.00	0.12	0.16	0.51	0.34
GBP						1.00	0.68	0.22	0.75
DEM							1.00	0.37	0.90
JPY								1.00	0.69
FXI									1.00

Notes: Std. Dev stands for standard deviation, p-values are in parentheses.

With the exception of the United States, the results of the table show that the expected excess returns are negative which may be explained by the substantial falls of international stock markets over the last period of our sample and by the over-evaluation of the USD over the whole sample. The latter can also explain the negative value of the expected return of the currency index. The positive estimates of the skewness coefficient and the highest value of the standard deviation for the Japanese stock market return, compared with the other markets, may reflect the prolong depression of this market, over

⁹ The NYBOT index computes a US Dollar based exchange rate index for the following currencies: Euro,

our sample. The estimates of the kurtosis coefficient and the ARCH statistic indicate that dynamic second moments are present in the stocks and currency index returns. Finally, taking together the estimates of the skewness and kurtosis coefficients reveal that the returns are not normally distributed.

The estimates of the correlation coefficients, reported in Panel B of the table, indicate that, with exception the US market, the exposure of the stock markets to the currency changes (denoted FXI) is positive. This suggests that US investors should hold foreign stocks for hedging purposes. Conversely, the negative exposure of the US stock market return to currency index changes indicates that there may not exist hedging benefits for the US investors to hold domestic stocks. Finally, the high positive values of the correlation coefficients for the GBP, DEM and JPY exchange rate returns with the currency index return suggest that the latter can substantially capture the nominal exchange rate changes, for all countries.

3.2 Estimation of the ICAPM

In Table II(a), we present the estimation results for the ICAPM with the multivariate EGARCH-M(1,1) econometric specification, given by equations (4)-(5).¹⁰ To capture the degree of excess kurtosis appeared in the data (see Table I), in the estimation procedure we assume that the standardised errors $\varepsilon_{w,t}$, $\varepsilon_{i,t}$ and $\varepsilon_{c,t}$ follow a multivariate t -distribution with degrees of freedom, denoted as DF, which are estimated by the maximum likelihood procedure. In Table II(b), we present estimates of the ICAPM based on a GARCH-M(1,1) specification of the variance functions (5), often used in practice to estimate the second moments of the ICAPM [see De Santis *et al* (1997, 1998)]. The comparison of the results of Table II(a) with those of Table II(b) enable us to evaluate whether the EGARCH-M(1,1) model can better represent the dynamics of the

Japanese Yen, British Pound, Canadian Dollar, Swedish Krona and Swiss Franc.

¹⁰ Our estimates are obtained based on Broyden, Fletcher, Goldfarb and Shanno's (BFGS) algorithm [see RATS 5.0 manual].

second moments of the returns. To this end, in the tables we report estimates of the Akaike and Schwarz information criteria.

The results of the tables indicate that the EGARCH-M(1,1) specification of the ICAPM constitutes a more accurate description of the data than the standard GARCH-M(1,1) model. This can be justified in terms of the maximum log-likelihood function value, which is found to be higher for the EGARCH-M(1,1) specification, and on the estimates of the information criteria. The parameter estimates and their standard errors (reported in parentheses) of the EGARCH-M(1,1) specification show that the price of risk coefficients λ_M and λ_C are different from zero, thus implying that both market and currency risks are priced in international financial markets. The sign of the market price of risk coefficient is positive, which is in accordance with the predictions of the theory [see Section 2].

Table II(a): Estimation of the multivariate EGARCH-M(1,1) with *t*-student distribution

	World	US	UK	Germany	Japan	FXI
$\lambda_{i,0}$	-0.288 (0.103)	-0.371 (0.122)	-0.145 (0.101)	-0.114 (0.126)	-0.279 (0.102)	0.329 (0.085)
λ_M			0.090 (0.024)			
λ_C			-0.201 (0.036)			
ω_i	1.551 (0.065)	1.759 (0.054)	1.862 (0.065)	2.329 (0.077)	4.718 (0.570)	-3.225 (0.827)
β_i	0.971 (0.005)	0.903 (0.021)	0.894 (0.045)	0.909 (0.039)	0.998 (0.000)	1.002 (0.001)
ϑ_i	-0.023 (0.004)	-0.046 (0.011)	-0.062 (0.027)	-0.064 (0.026)	-0.050 (0.003)	-0.017 (0.010)
γ_i	0.009 (0.004)	-0.030 (0.009)	0.022 (0.033)	0.069 (0.026)	-0.006 (0.002)	0.069 (0.012)
DF						4.260 (0.049)
Akaike						12901.6
Schwarz						13117.1
Log-Lik						-6402.81

Notes: The estimates of the correlation coefficients are omitted for reasons of space. These are found to be close to their unconditional estimates, presented in Table I.

Given that there is a positive correlation between stock and currency returns for the foreign countries [see Table I], the negative sign of the currency price of risk coefficient implies that $J_{WC} > 0$, which means that foreign stocks constitute natural hedgers against

real exchange rates movements [see Section 2]. This is consistent with evidence provided by other studies [see De Santis (1998) and Cappiello, Castrén and Jääskelä (2003), *inter alia*]. It means that US investors will demand a lower risk premium for holding foreign stocks. Finally, the small value of the estimated DF coefficient justifies the need to estimate the conditional second moments of the returns based on the t -distribution.

In contrast to Table II(a), the results reported in Table II(b) (based on the GARCH-M(1,1) specification) indicate that both estimates of λ_M and λ_C are not different from zero at 5%, even though their sign is consistent with that of the EGARCH-M(1,1) specification.

Table II(b): Estimation of the multivariate GARCH-M(1,1) with t -student distribution

	World	US	UK	Germany	Japan	FXI
$\lambda_{i,0}$	-1.540 (1.021)	-1.611 (1.000)	-1.135 (0.873)	-1.299 (1.149)	-1.576 (1.104)	0.321 (0.298)
λ_M			0.377 (0.224)			
λ_C			-0.270 (0.145)			
ω_i	0.954 (0.208)	2.816 (0.418)	1.898 (0.809)	2.082 (1.457)	7.852 (2.921)	0.063 (0.040)
a_i	0.024 (0.013)	0.022 (0.014)	0.049 (0.029)	0.040 (0.021)	0.094 (0.052)	0.028 (0.017)
b_i	0.773 (0.048)	0.509 (0.062)	0.672 (0.135)	0.766 (0.133)	0.329 (0.240)	0.949 (0.028)
DF						4.259 (0.044)
Akaike						13199.9
Schwarz						13388.4
Log-Lik						-6557.95

Notes: The GARCH-M(1,1) variance functions are given by $\sigma_{k,t}^2 = \omega_k + a_k \varepsilon_{k,t-1}^2 + b_k \sigma_{k,t-1}^2$ (for $k = M, i, C$).

These differences can be attributed to the fact that the GARCH-M(1,1) does not adequately capture the dynamics of the second moments of the data, as argued before. The estimates of the EGARCH-M(1,1) specification indicate that there exist strong leverage effects in the variance functions, as the estimates of \mathcal{G}_k , for $(k = M, i)$, are negative and different from zero. Note that, although the positive sign of \mathcal{G}_k for the currency index return reveals an asymmetric effect of news on the currency volatility (for

instance, a currency depreciation), this is not theoretically justified as reflecting leverage effects. The positive and statistically different from zero estimates of the γ_k , for $(k = M, i)$ coefficients, capturing the magnitude effect of the market news on the variance functions, reveal that the leverage effects are larger than expected for most markets, with the exception of US and Japan.

Apart from affecting the estimates of the risks price coefficients, λ_M and λ_C , ignoring the leverage effects seems to influence the degree of persistency of the markets' volatility. The comparison of the estimates of the persistency coefficients β_i for the EGARCH-M(1,1) specification and $(a_i + b_i)$ for the GARCH-M(1,1) model indicates that the former is much higher. This should be expected because the GARCH-M(1,1) specification does not capture the component of the stock markets' volatility coming from the leverage effects, or the other asymmetries in the variance function. Note that for the currency index return the estimate of the persistency coefficient, β_i , is not statistically different from unity, which implies that the currency volatility is an integrated process of order one. This can be attributed to the currency crises occurred during our sample.

Overall, the results of our empirical analysis indicate that both the market and the currency sources of risk are priced in equilibrium expected returns. The strong leverage effects, which are present in both stock and currency markets, seem to critically affect the dynamics, the persistency and the asymmetry of markets volatility, and hence the market and currency risk premia.

Next, we present in-sample, point t estimates of the market and currency premia, as well as the total premia, with the aim of investigating the effects that currency episodes had on equity and currency markets. These estimates are plotted in Figures 1-6. The market premia (measured on the left vertical axis) are calculated as $MP_{k,t-1} = \lambda_M Cov_{t-1}(r_{k,t}, r_{M,t})$, while the currency premia (measured on the right vertical axis inverted) are calculated as $CP_{k,t-1} = \lambda_C Cov_{t-1}(r_{k,t}, c_t)$, for $k = (M, i, C)$. These are based on the EGARCH-M(1,1)

estimates, reported in Table II(a). In Figures 7-12, we present the total risk premia, which is computed as $TP_{k,t-1} = \lambda_{k,0} + MP_{k,t-1} + CP_{k,t-1}$.

Figure 1: World

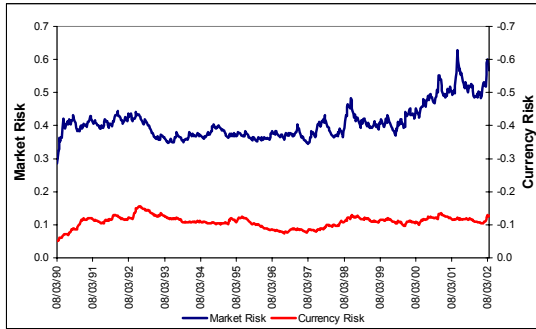


Figure 2: US

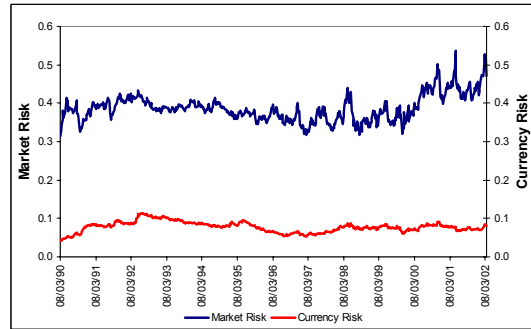


Figure 3: UK

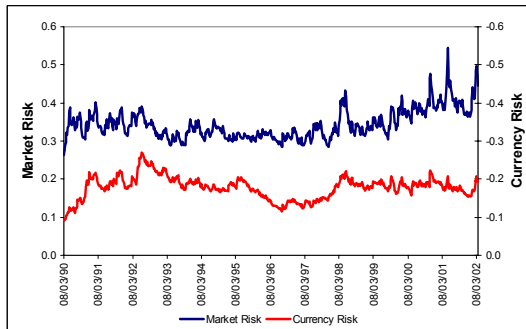


Figure 4: Germany

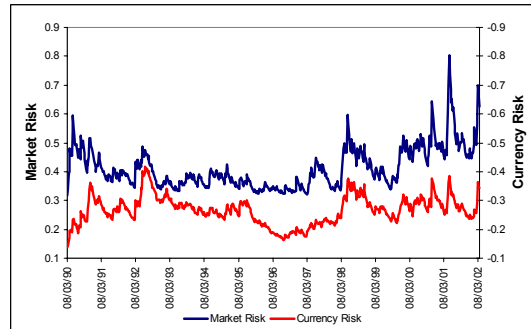


Figure 5: Japan

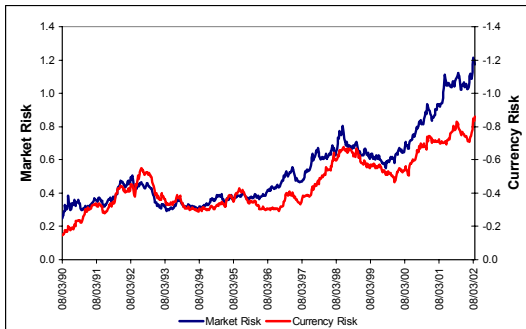
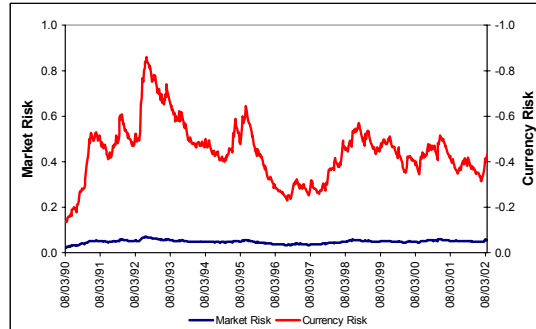


Figure 6: FXI



Inspection of the above figures leads to the following conclusions. First, the currency premium constitutes an important component of the total premium (i.e. the expected excess return). The big variations of the currency premia seem to be connected with the currency crises occurred within our sample, namely the 1992-1993 ERM crisis, the 1995 peso crisis and 1998 currency crises in Russia and Japan [see Figure 6].

Figure 7: World

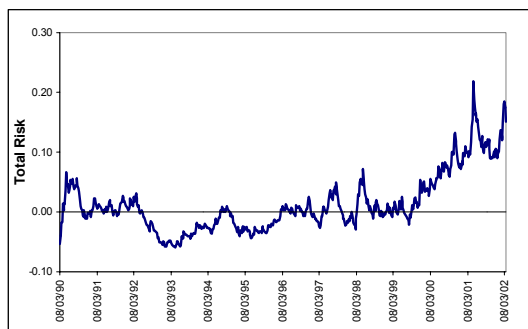


Figure 8: US

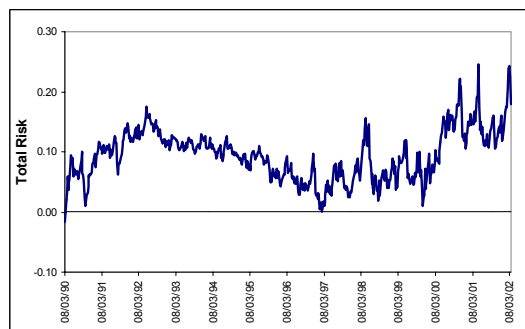


Figure 9: UK

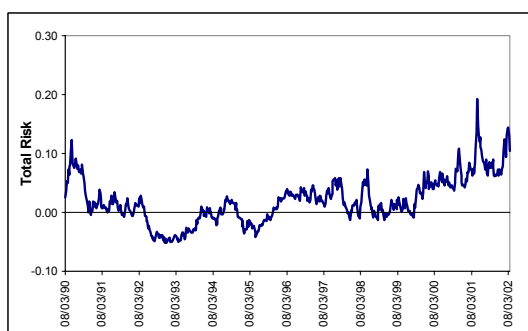


Figure 10: Germany

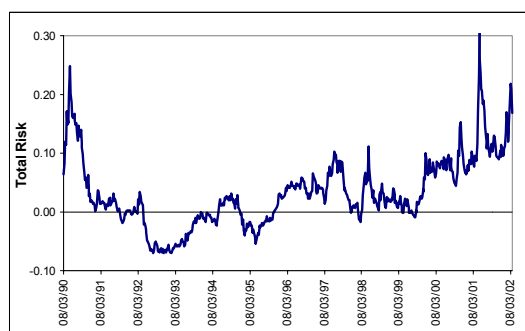


Figure 11: Japan

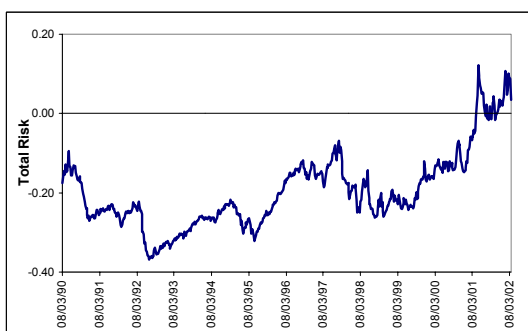
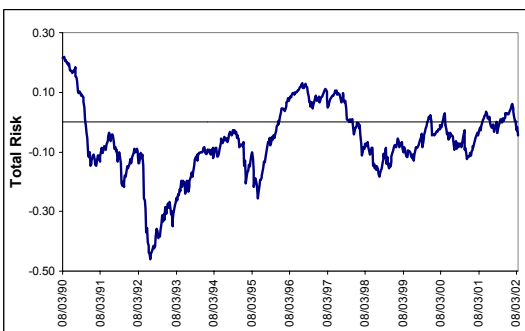


Figure 12: FXI



These crises have also critically influenced the volatility of the stock markets. The effects of these crises were more profound for the UK, German and Japanese stock markets volatility, compared with the US market. Note that, for Germany and Japan, the market premia move very closely with the currency ones. The above results suggest that the currency crises have significantly influenced the stock market premia.

The second conclusion that can be drawn from the figures is that the total premium dramatically changes, over our sample. It fluctuates between negative and positive values. The negative sign of the total premia for the non-US markets, driven by the

currency hedging attitude of the investors, corresponds to periods in which the US dollar was strong (e.g. during currency crises). Finally, the positive, upward sloping movements of the total premia towards the end of the sample may be attributed to the stock markets crises, e.g. the September 11th terrorist attack and the Enron and WorldCom corporations collapses.

4. Economic evaluation of the ICAPM

The results of Section 3 reveal that the currency premium is priced in international financial markets and constitutes an important component of foreign expected returns. If this is the case, then Merton's modern portfolio approach suggests that investors should hold, in addition to the market portfolio, another portfolio which will hedge their invested wealth for adverse effects from currency changes. The goal of this section is to evaluate the above theoretical prediction on economic grounds. To this end, we will compare the economic profits (reflecting risk premia effects) of two portfolios: one which counts for currency effects and another which does not.

Specifically, modern portfolio theory in a dynamic set up implies the following optimal portfolio allocation

$$\mathbf{w}_{t|t-1} = \left(-\frac{J_W}{WJ_{WW}} \right) \boldsymbol{\Sigma}_{t|t-1}^{-1} E_{t-1}(\mathbf{r}_t) + \left(-\frac{J_{WC}}{WJ_{WW}} \right) \boldsymbol{\Sigma}_{t|t-1}^{-1} \boldsymbol{\sigma}_{t|t-1}, \quad [7]$$

where $\mathbf{w}_{t|t-1}$ denotes the t -period ($N \times 1$)-vector of the optimal weights for N risky stocks (here $N=4$) at time $t-1$, $\boldsymbol{\Sigma}_{t|t-1}$ stands for the t -period ($N \times N$) conditional variance-covariance matrix of the stock returns, with elements $Cov_{t-1}(r_{i,t}, r_{j,t})$ $\{i, j\} = 1, 2, \dots, N$, $\boldsymbol{\sigma}_{t|t-1}$ stands for the ($N \times 1$) vector of the conditional covariances of the stocks returns with the changes of the currency index return, c_t , with elements $Cov_{t-1}(r_{i,t}, c_t)$, and $E_{t-1}(\mathbf{r}_t)$ is

the (NXI) vector of the conditional expected excess returns, with elements $E_{t-1}(r_{i,t})$.

Writing equation (7) as

$$\mathbf{w}_{t|t-1} = \gamma_M \mathbf{w}_{M,t|t-1} + \gamma_C \mathbf{w}_{C,t|t-1}, \quad [8]$$

where $\mathbf{w}_{M,t|t-1} = \Sigma_{t|t-1}^{-1} E_{t-1}(\mathbf{r}_t)$, $\mathbf{w}_{C,t|t-1} = \boldsymbol{\sigma}_{t|t-1} E_{t-1}(\mathbf{r}_t)$, $\gamma_M = \left(-\frac{J_W}{WJ_{WW}} \right) = \frac{1}{\lambda_M}$ is the

inverse of the market price of risk coefficient and $\gamma_C = \left(-\frac{J_{WC}}{WJ_{WW}} \right) = -\frac{\lambda_C}{\lambda_M}$ ¹¹, we can see

that the ICAPM predicts that investors in equilibrium should hold a portfolio $\mathbf{w}_{t|t-1}$ which consists of two other portfolios: the portfolio $\mathbf{w}_{M,t|t-1}$ (the market portfolio), compensating investors for bearing the market source of risk and the portfolio $\mathbf{w}_{C,t|t-1}$ (the hedge portfolio) compensating investors for wealth losses coming from currency changes.

The optimal portfolio allocation implied by equation (8) should generate higher realised profits (excess returns) per unit of risk than a portfolio allocation which assumes only the market source of risk, i.e. $\lambda_C = 0$. Before assessing the validity of the above statement, we need to evaluate the forecasting performance of the EGARCH-M(1,1) specification of the ICAPM. This is necessary because the economic evaluation of the model requires that its statistical specification provides unbiased forecasts of the realised returns which are used to determine the optimal weights $\mathbf{w}_{t|t-1}$.

4.1 Forecasting performance

In this subsection, we evaluate the forecasting performance of the EGARCH-M(1,1) specification of the ICAPM based on density forecast testing procedures [see Diebold, Gunther and Tay (1998), *inter alia*]. These methods can account for the effects of higher

order dynamics on evaluating the ability of the model to accurately predict many percentiles of the empirical distributions of the expected stock and currency returns. Since, in general, tests for forecasting ability can be used as tests for the structural stability of a model, the above methods can be also thought of as testing whether the estimates of the parameters of the EGARCH-M(1,1) specification (4)-(6) are subject to structural changes.

The general idea behind density forecast evaluation is that the conditional density probabilities of the returns, denoted by $p(r_{k,t}|\Omega_{t-1})$ $k = (M, i, c)$, should correspond to the true conditional density implied by the EGARCH(1,1)-M model (4), denoted as $f(r_{k,t}|\Omega_{t-1})$, i.e.

$$p(r_{k,t}|\Omega_{t-1}) = f(r_{k,t}|\Omega_{t-1}).$$

Then, the probability integral transforms of the actual realisations of the returns over the forecast period with respect to the model's forecast densities ($p(r_{k,t}|\Omega_{t-1})$) should be *IIDU*[0,1], i.e.

$$z_{k,t} = \int_{-\infty}^{r_{k,t}} p_{k,t}(u) du \sim IIDU[0,1], \quad [9]$$

where $z_{k,t}$ denotes the probability transform variable and $U[.]$ stands for the uniform distribution.

The result of equation (9) implies that the cumulative distribution of $z_{k,t}$ should lie on the 45⁰ line (which is the theoretical cumulative distribution function - CDF) and that the inverse function of $z_{k,t}$ [say $y_{k,t} = \Phi^{-1}(z_{k,t})$] is *IIDN*(0,1). These two implications can

¹¹ $\gamma_C = \left(-\frac{J_{WC}}{WJ_{WW}} \right) = \left(-\frac{J_{WC}}{J_W} \frac{J_W}{WJ_{WW}} \right) = -\frac{\lambda_C}{\lambda_M}$

be exploited to assess the forecasting performance of the EGARCH(1,1)-M specification of the ICAPM (1). To this end, we carried out an in-sample-forecasting exercise for the period from 21st of May 1999 to the end of the sample, 16th August 2002. This period covers two regimes of the international stock markets: the bull and the bear, started in the spring of year 2000. Thus, it allows us to see if the forecasting performance of our model remains robust to the above market regime changes.

Figure 13: World

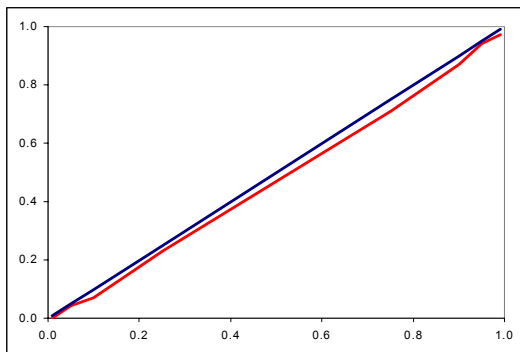


Figure 14: US

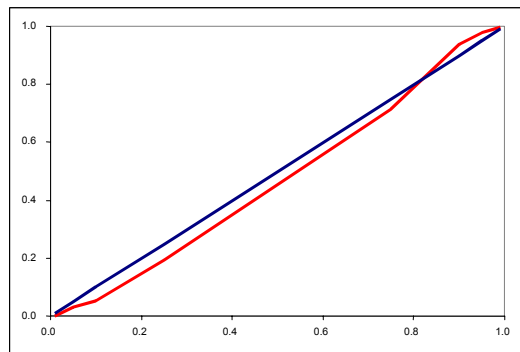


Figure 15: UK

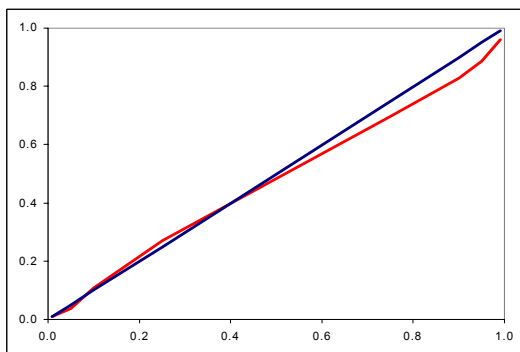


Figure 16: Germany

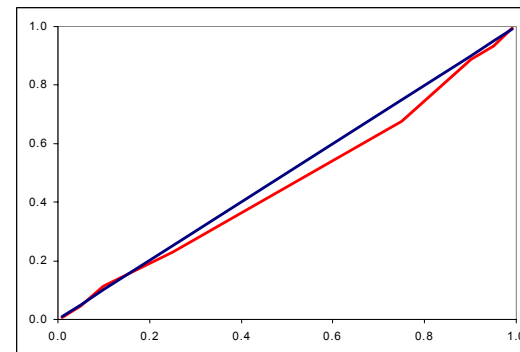


Figure 17: Japan

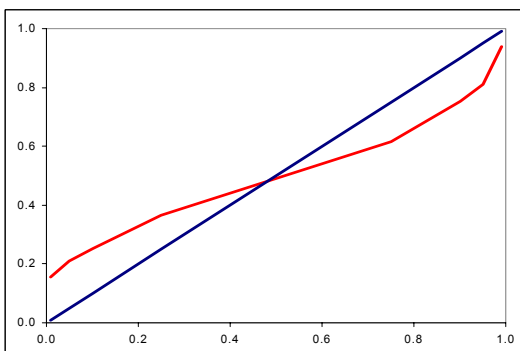
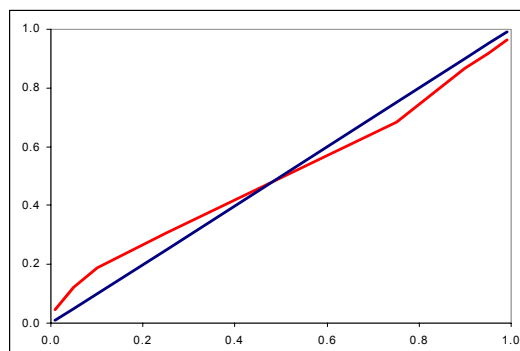


Figure 18: FXI



In Figures 13-18 we graphically present the empirical distribution of $z_{k,t}$ vis-à-vis the 45^0 line, for all the realised returns. The figures indicate that the empirical cumulative distributions are very close to the 45^0 line. The only exception is for the Japanese stock return, where the empirical distribution substantially deviates from its theoretical CDF. This can be attributed to the fact that the conditional variance of the Japanese stock return seems to follow a non-stationary, explosive pattern [see Figure 5].

The evaluation of the density forecasts made above is informal. Therefore, in Table III we present the results of two formal test statistics for density forecast evaluation. The first is the well known Kolmogorov-Smirnov test statistic, denoted as KS. This statistic measures if the maximum distance of the empirical cumulative distribution of $z_{k,t}$ from its theoretical (45^0 line) is not statistically significant. The second is a parametric test statistic suggested by Berkowitz (2001), denoted as BK, build up on the result that the inverse function of $z_{k,t}$ [see $y_{k,t} = \Phi^{-1}(z_{k,t})$ above] is *IIDN*(0,1). This has the following testable implications: there will be no systematic deviations of $y_{k,t}$ (and hence $z_{k,t}$) and that the unconditional mean of $y_{k,t}$ will be zero. These can be jointly tested using the following auxiliary regression

$$y_{k,t} = a_{k,0} + a_{k,1}y_{k,t-1} + \xi_{k,t}, \quad [10]$$

which can be used to test for the null hypothesis $H_0 : a_0 = a_1 = 0$.

Table III: Formal tests

	World	US	UK	Germany	Japan	FXI
KS	0.08	0.10	0.08	0.09	0.17	0.09
BK	2.95 (0.23)	6.93 (0.03)	3.08 (0.21)	5.49 (0.06)	4.61 (0.10)	0.15 (0.93)

Notes: The critical value for the KS test is 0.1048. BK is a Wald test statistic of the null hypothesis $H_0 : a_0 = a_1 = 0$. p-values in parentheses.

The results of the table indicate that the overall density forecasting performance of the EGARCH-M(1,1) specification (4) is satisfactory and remains robust across the two market

regimes of the forecasting interval. Both the KS and BK statistics show that the deviations of the empirical distribution of the returns series $r_{k,t}$ are not significant and systematic, at 5% level. Note that for US the BK statistic cannot reject the null at 3% level. For Japan, the two test statistics lead to different conclusions. The BK statistic cannot reject the null, whilst the KS clearly reject it. Given the substantial deviation between the empirical and theoretical CDFs of $z_{k,t}$ in the case of Japan [see Figure 17], this difference may be attributed to the low power of the BK statistic in the presence of the non-stationary behaviour of the second moments, documented for Japan [see Figure 5].

4.2 Economic performance

Having found that the EGARCH(1,1)-M specification of the ICAPM can provide satisfactory density forecasts of the returns, we next turn into evaluating its economic performance. In Figure 19 we present estimates of the cumulative profits implied by the ICAPM of a \$1 investment in May 21st 1999 under two dynamic investment strategies: first, when the market portfolio is hedged against currency changes (referred to as H-strategy) and, second, when currency changes are ignored, i.e. $\lambda_C = 0$, (referred to as NH-strategy). In Figure 20, we present the difference of the two cumulative profit series.

To calculate the profits, we work as follows. At any point of our forecast interval, we estimate the expected returns and their conditional second moments for one-period ahead. These estimates are used to compute the optimal portfolio weights and the cumulative profits, under each strategy. For the H-strategy, the optimal weights are calculated based on equation (8), where γ_M and γ_C are estimated using the values of λ_M and λ_C reported in Table II(a). For the NH-strategy, the optimal weights and λ_M are estimated based on a EGARCH-M(1,1) specifications of the ICAPM which does allow for currency risk. The results of the figures indicate that the cumulative profits of the H-strategy outperform the ones of the NH-strategy, almost at each point of the forecast interval.

This can be formally confirmed by summary statistics presented in Table V, reporting the mean, the standard deviation and the test statistic of difference in means for the two cumulative profit series. These statistics clearly show that the H-strategy has a higher mean and lower volatility than the NH-strategy, and that the mean-difference between the two profit series is different from zero.

Figure 19: Cumulated Profits

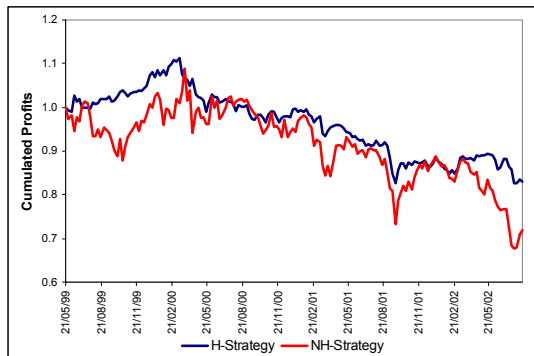


Figure 20: H-Strategy Outperformance

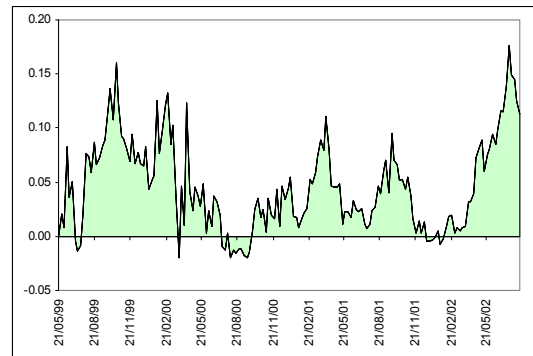


Table IV: Statistics on the cumulative profits

	H-Strategy	NH-Strategy
Average	0.962	0.912
St. deviation	0.073	0.082
Statistic of difference in mean		5.494

Note that, under both strategies, Figure 19 reveals that the cumulative profits start declining after the beginning of year 2000. This can be attributed to the burst of the international stock markets bubble in spring of 2000 and to a series of exogenous events that affected the the markets, such as the terrorists attack in September 2001 and the collapse of the Enron and WorldCom corporations. These effects can not be predicted and, thus, hedged under the ICAPM. However, even for this period of events, Figures 19-20 indicate that the H-strategy outperforms the NH-strategy.

5. Conclusions

In this paper we suggest a parsimonious version of the ICAPM in the spirit of a two-single factors asset pricing model of Merton with the aim of examining whether or not

the currency risk is priced in international stock markets. To estimate the first and second conditional moments of the model, we use an EGARCH-M(1,1) specification for stocks and currency returns. This specification enables us to better represent the dynamics of the returns' volatility, as it can capture leverage effects or any other asymmetries due to stock and/or currency market news.

Our analysis provides a number of interesting results. First, it shows that the currency premium is priced in international stock markets and that it constitutes an important component of expected stock returns. Second, it shows that the volatility of both market and currency premia critically depend on currency news. Third, it shows that the EGARCH-M(1,1) econometric specification of the ICAPM provides accurate density forecasts of the stock and currency returns. Finally, in assessing the economic implications of our results, it finds that a dynamic portfolio investment strategy hedging against currency risk outperforms one which ignores currency changes, thus implying an economically significant reward for the currency premium.

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