Did the EMS Reduce the Cost of Capital? [1]

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[1]This is a thoroughly revised version of Sentana, Shah and Wadhwani (1995), which extended the analysis of stock returns in Sentana, Shah and Wadhwani (1992) to currencies and bonds. I am very grateful to Mushtaq Shah and Sushil Wadhwani for their contributions to the earlier versions of this paper. I would also like to thank Manuel Arellano, Juan Ayuso, Philippe Bacchetta, José M. Campa, Greg Connor, Charles Goodhart, Mike Wickens and participants at seminars at AUEB (Athens), Bank of England, Bank of Spain, the Board of Governors of the Federal Reserve, Fundación BBV (Bilbao), IAE (Barcelona), LSE, UIMP (Santander), and the 1995 Royal Economic Society Conference for their comments. The editor and three anonymous referees have also helped me greatly improve the paper. Of course, the usual caveat applies. Thanks are also due to Juan Ayuso, Martin Brookes, Aelred Conelly, Simon Cook, José A. Cuenca, and Andrew Haldane for their help in collecting the data, and to Isabel Figuerola-Ferretti and Angel León for research assistance. This work was initiated at the LSE Financial Markets Group, whose financial support for the first version as part of the ESRC project "The Efficiency and Regulation of Financial Markets" is gratefully acknowledged. Address for correspondence: CEMFI, Casado del Alisal 5, 28014 Madrid, Spain (tel.: +34 91 429 0551, fax: +34 91 429 1056, e-mail: sentana@cemfi.es).

Abstract

We propose a dynamic APT multi-factor model with time-varying volatility for currency, bond and stock returns for ten European countries over the period 1977-1997. We exploit the cross-sectional dimension of the model to construct world portfolios, which when added to the original list of assets, allow us to develop simple consistent methods of estimation and testing. Our results reject the implicit asset pricing restrictions, and suggest that decreases in idiosyncratic exchange rate risk tend to lower the cost of capital, although the effect is small. Finally, we assess the potential gains from increased stock market integration.

Keywords: Currency Risk; European Monetary Union; Financial Integration; International Asset Pricing

JEL: F3, G1

1 Introduction

The European Commission argued (see *European Economy* (1990), p. 82) that *"... potentially the most important source of gains from European Monetary Union comes from the reduction in overall uncertainty EMU might provide"*. They went on to contend that a reduction in exchange rate risk would reduce the risk premium, and that this reduction in the cost of capital would stimulate corporate investment. Nevertheless, as the report acknowledged, it is not a priori obvious that intra-European exchange rate risk should affect the cost of capital. For example, firms might be able to hedge their exchange rate exposure through a variety of financial instruments. Similarly, those who diversify their investments globally may not be affected by idiosyncratic variations in a country's exchange rate. Therefore, a preliminary question that has to answered is whether country-specific exchange rate factors are priced in the context of asset pricing models applied at the international level (see Stulz (1995) for a recent survey). Although those risks should not be rewarded in a world with complete market integration, the existence of capital controls or other legal impediments to cross-border investment (such as limitations on the holdings of foreign securities by pension funds and insurance companies), informational asymmetries, illiquid markets, behavioural biases, etc. suggest that idiosyncratic exchange rate risk is likely to be priced.

Hence, whether or not EMU would reduce the cost of capital, and if so, by how much, is, ultimately, an empirical question. We attempt to throw some light on this issue within the framework of the dynamic version of the Arbitrage Pricing Theory (APT) developed in King, Sentana and Wadhwani (1994) (KSW hereinafter). Specifically, we use monthly data on currency, bond and stock returns for ten European countries over the period 1977-1997 to estimate a multi-factor model with time-varying volatility in the underlying factors, in which the idiosyncratic components of returns are (almost) uncorrelated across countries, but their correlation structure is arbitrary within each country. From the methodological point of view, we exploit the cross-sectional dimension of the model to construct diversified portfolios of European and non-European securities, which when added to the original list of assets, allow us to develop simple consistent methods of estimation and testing. In particular, we employ a GMM estimation procedure that would be efficient under conditional homoskedasticity, but which remains consistent under the more realistic assumption of dynamic heteroskedasticity.

Under the null, our model implies that country-specific risks should not be priced. But as we mentioned before, it is likely that these idiosyncratic risk components do indeed affect the required rate of return on bonds and stocks. If country-specific exchange rate volatility is associated with higher stock returns, then systems that attempt to reduce nominal exchange rate variability, such as the Exchange Rate Mechanism (ERM) of the European Monetary System (EMS), may well reduce capital costs for firms that raise funds by issuing equity. But since the stock market is not the primary source of finance in some of the countries that we look at, Germany being the prime example, we also look at bond returns.[1] In either case, though, we would still be looking at only one component of the cost of capital. Since we include currency returns in our set of assets, we can also say something about the effect of the EMS on the riskless interest rate component.

However, under a target zone system such as the ERM, the cost of capital would not necessarily fall; for example, some authors (e.g. see Batchelor (1985)) have argued that a credible system will inevitably increase interest rate volatility since the authorities are forced to defend the currency. If interest rate volatility is also positively associated with risk premia on bond and equity markets, and if ERM membership raised interest rate volatility, then, it is theoretically possible that the EMS might even have increased the cost of capital. Again, this is an

[1] But see Ando and Auerbach (1988) for a discussion of the difficulties induced by differential taxation in cross-country comparisons of the debt component of the cost of capital.

issue that we shall investigate.

More generally, testing the cross-equation asset pricing restrictions of our basic model enables us to examine the important question of whether European capital markets are integrated. In this respect, it is worth noting that an important indirect effect of EMU, in conjunction with the development of the single market for financial services, should be the elimination of many of the remaining barriers to cross-border investments in the European Union. In order to gauge the potential gains from increased market integration, we follow Stulz (1999), and compare stock market risk premia under full integration with the risk premia that would prevail in the context of completely segmented markets.

The rest of the paper is organised as follows. Section 2 discusses our basic model and its estimation procedure. Section 3 reports the empirical results, while the final section contains our conclusions. This is followed by a data appendix.

2 Theory and estimation

2.1 Asset pricing model

We base our analysis in a world with a large number of countries, and assume that for each country there are three representative assets available: a 1-period local currency deposit with safe gross return R^j_{cjt}, a long-term default-free bond portfolio, whose random gross holding return over period t in local currency is R^j_{bjt}, and a stock portfolio, whose random gross holding return in local currency is R^j_{sjt}. Let $S^{\$}_{jt}$ be the spot exchange rate for country j at the end of period t in terms of the numeraire currency (US \$ in our case), and let $R^{\$}_{c\$t}$ be the gross return on the safe asset for the US during period t in US \$. The excess returns of the three representative assets for each country in terms of the numeraire currency will be given by:

$$
\begin{aligned}
r^{\$}_{cjt} &= \log R^{\$}_{cjt} - \log R^{\$}_{c\$t} = \log S^{\$}_{jt} - \log S^{\$}_{jt-1} + \log R^{j}_{cjt} - \log R^{\$}_{c\$t} \\
r^{\$}_{bjt} &= \log R^{\$}_{bjt} - \log R^{\$}_{c\$t} = r^{\$}_{cjt} + r^{j}_{bjt} \\
r^{\$}_{sjt} &= \log R^{\$}_{sjt} - \log R^{\$}_{c\$t} = r^{\$}_{cjt} + r^{j}_{sjt}
\end{aligned} \tag{1}
$$

where r^{j}_{bjt} and r^{j}_{sjt} are the (continuously compounded) excess returns for bonds and stocks in local currency.[2] Note that under covered interest parity, $r^{\$}_{cjt}$ is the variation in exchange rates corrected for the forward premium.

Let $\mu^{\$}_{cjt}, \mu^{\$}_{bjt}$ and $\mu^{\$}_{sjt}$ be the risk premia on currency, bond and stock returns for country j, and let $\eta^{\$}_{cjt}, \eta^{\$}_{bjt}$ and $\eta^{\$}_{sjt}$ be the corresponding unanticipated (as of time $t-1$) components of returns. Given that $\eta^{\$}_{cjt}$ simply reflects the unexpected variations in exchange rates against the common numeraire currency, and these are so highly correlated across countries, a single factor model should provide a realistic representation of their covariance structure. In this respect, note that full European Monetary Union would imply that $r^{\$}_{cjt}$ would be the same for all member countries. On the other hand, while both bond and stock returns are additionally affected by interest rate movements, it is generally accepted that stock returns are also exposed to other risks. On this basis, we assume the following conditional factor structure for the innovations in returns:

$$
\begin{array}{rclcl}
\eta^{\$}_{cjt} & = & \beta^{\$}_{cje} f_{et} & + & v_{jet} \\
\eta^{\$}_{bjt} & = & \beta^{\$}_{bje} f_{et} + \beta_{bji} f_{it} & + & \delta^{\$}_{bje} v_{jet} + v_{jit} \\
\eta^{\$}_{sjt} & = & \beta^{\$}_{sje} f_{et} + \beta_{sji} f_{it} + \beta_{sjm} f_{mt} & + & \delta^{\$}_{sje} v_{jet} + \delta_{sji} v_{jit} + v_{jmt} \\
 & & \text{Systematic risk} & & \text{Specific risk}
\end{array} \tag{2}
$$

or in matrix notation $\boldsymbol{\eta}^{\$}_{jt} = \mathbf{B}^{\$}_{j}\mathbf{f}_{t} + \boldsymbol{\Delta}^{\$}_{j}\mathbf{v}_{jt}$, where f_{et} and f_{it} are common factors representing systematic exchange rate and interest rate risks, f_{mt} is a common

[2]Strictly speaking, excess returns should be defined in terms of arithmetic returns, rather than geometric returns. Nevertheless, the approximation error is usually irrelevant in empirical work (see McCulloch (1975)).

residual market risk factor, v_{jet} and v_{jit} are exchange and interest rate risks specific to country j, v_{jmt} represents other risks specific to country j stocks, and the $\beta's$ and $\delta's$ are the associated factor loadings that measure the sensitivity of the assets to the different factors, which we assume time-invariant for any given unconditional normalisation of the factors. To guarantee that the $\boldsymbol{\eta}^{\$\prime}_{jt}s$ are innovations, we assume that common and specific factors are unpredictable on the basis of past information. Following KSW, we also make the assumption that the common factors are conditionally orthogonal to each other, but allow them to have time-varying conditional variances $\lambda_{et}, \lambda_{it}$ and λ_{mt}.[3] As for the idiosyncratic terms, which by definition are conditionally orthogonal to $\mathbf{f}_t$, we also assume that they are conditionally orthogonal to one another for a given j, with time-varying conditional variances $\omega_{jet}, \omega_{jit}$ and ω_{jmt}. But note that since $\left|\boldsymbol{\Delta}^{\$}_j\right| = 1$, the idiosyncratic *unconditional* covariance matrix remains totally unrestricted within a country. Finally, we assume that the idiosyncratic conditional covariance matrix has the approximate zero factor structure introduced by Chamberlain and Rothschild (1983), in which v_{jet}, v_{jit} and v_{jmt} may be correlated across countries, but only mildly so in order to guarantee that full diversification applies.[4]

Under a mild no arbitrage condition, it is possible to prove that there is a stochastic discount factor, ξ_t, which prices the available assets by discounting their uncertain payoffs across different states of the world (see e.g. Cochrane (2001)). In terms of excess returns, in particular, we will have that

$$E_{t-1}(r^{\$}_{ajt}\xi_t) = E_{t-1}(r^{\$}_{ajt})E_{t-1}(\xi_t) + cov_{t-1}(r^{\$}_{ajt}, \xi_t) = 0 \quad (a = c, b, s) \qquad (3)$$

If (and only if) the stochastic discount factor contains no idiosyncratic risk, ξ_t will be correlated with $r^{\$}_{cjt}$, $r^{\$}_{bjt}$ and $r^{\$}_{sjt}$ only through $\mathbf{f}_t$, and the assets' risk premia can

[3] See Sentana and Fiorentini (2001) for the implications of this assumption on the identification of the factors and their loadings.

[4] Their definition is asymptotic, and amounts in our case to the largest eigenvalue of the (conditional) idiosyncratic covariance matrix remaining uniformly bounded as the number of *countries*, N, goes to infinity.

be written as an exact linear combination of the volatility of the common factors, with weights proportional to the corresponding factor loadings. Specifically,

$$\begin{aligned}
\mu_{cjt}^{\$0} &= \beta_{cje}^{\$}\tau_e\lambda_{et} \\
\mu_{bjt}^{\$0} &= \beta_{bje}^{\$}\tau_e\lambda_{et} + \beta_{bji}\tau_i\lambda_{it} \\
\mu_{sjt}^{\$0} &= \beta_{sje}^{\$}\tau_e\lambda_{et} + \beta_{sji}\tau_i\lambda_{it} + \beta_{sjm}\tau_m\lambda_{mt}
\end{aligned} \tag{4}$$

where τ_k $(k = e, i, m)$ are the prices of risk corresponding to each factor, i.e. the amount of expected return that agents are willing to give away to reduce its variability by one unit. An alternative way of interpreting the above relationship can be obtained by noticing that risk premia are also linear combinations of the asset factor loadings or betas, with weights that are common to all assets. These weights can be understood as the risk premia of three factor mimicking portfolios, i.e. three unit-cost, diversified portfolios of risky assets with unit loadings on the common exchange rate, interest rate and residual stock market risk factors respectively, and zero loadings on the others. In this respect, our model coincides with a conditional version of the exact[5] APT pricing relationship.[6] Importantly, since we explicitly consider currency returns, our model will also hold for local currency excess returns on bonds and stocks, which can be understood as fully hedged returns (see e.g. Bekaert and Hodrick (1992)). Furthermore, it will also hold for stock and bond returns measured in a common (diversified) basket currency.

Note that if our asset pricing model is correct, risk prices depend on the factors, not on the assets, since otherwise there would be arbitrage opportunities. Furthermore, the model also implies that country-specific risks should not be

[5]If the stochastic discount factor ξ_t contains some idiosyncratic risk, the APT expression (4) becomes approximate, as in Ross' (1976) original formulation, and it is only possible to prove that the pricing errors would be negligible on average cross-sectionally, but not necessarily so for each asset. Connor (1984) provides conditions that guarantee exact asset pricing in a competitive equilibrium set up.

[6]Ross' (1976) results were implicitly derived for a closed economy. However, if exchange rates belong to the set of factors, then the APT can be readily generalised to an international setting (see Solnik (1983) and Ikeda (1991)).

rewarded because they can be diversified away. As we will see in the next section, these fundamental restrictions are the basis of our tests.

2.2 Alternative hypotheses

In order to assess whether EMU would reduce the cost of capital, and if so, by how much, our most general alternative is given by

$$\begin{array}{ll} \mu^{\$}_{cjt} = \mu^{\$0}_{cjt} + \rho^{\$}_{cje}\omega_{jet} & +\theta^{\$}_{cje}\lambda_{et} \\ \mu^{\$}_{bjt} = \mu^{\$0}_{bjt} + \rho^{\$}_{bje}\omega_{jet} + \rho_{bji}\omega_{jit} & +\theta^{\$}_{bje}\lambda_{et} + \theta_{bji}\lambda_{it} \\ \mu^{\$}_{sjt} = \mu^{\$0}_{sjt} + \rho^{\$}_{sje}\omega_{jet} + \rho_{sji}\omega_{jit} + \rho_{sjm}\omega_{jmt} & +\theta^{\$}_{sje}\lambda_{et} + \theta_{sji}\lambda_{it} + \theta_{sjm}\lambda_{mt} \end{array} \quad (5)$$

where $\mu^{\$0}_{cjt}, \mu^{\$0}_{bjt}$ and $\mu^{\$0}_{sjt}$ are defined in (4).

This equation enables us to test various hypotheses of interest:

1. We may ask if idiosyncratic European exchange rate variability is not priced in bond and stock markets, i.e. if $\rho_{bje} = \rho_{sje} = 0 \ \forall j$. If so, then it is difficult to see how the EMS could have affected the equity and bond components of the cost of capital by affecting exchange rate volatility. Similarly, testing if $\rho_{cje} = 0 \ \forall j$ is also very interesting, as it would throw some light on the question of whether the observed convergence of short rates in Europe is explicable by the reduction in exchange rate volatility. In both cases, the values of ρ_{aje} $(a = c, b, s)$ will allow us to measure the effects of eliminating such idiosyncratic risks.

2. If idiosyncratic interest rate volatility is priced, i.e. ρ_{bji} or $\rho_{sji} \neq 0$, then the EMS might have affected the cost of capital, *if* it also affected interest rate volatility.

3. It is possible that other sources of idiosyncratic risk are priced in the stock market, i.e. $\rho_{sjm} \neq 0$, which would suggest that European stock markets are not fully integrated.

4. We may also test for European capital market integration in the sense of asking whether the different prices of risk are common across countries. To do so, we

can ask whether freeing up the way in which the conditional variances of the common factors affect risk premia can help us explain returns better. That is, if $\theta_{bje} = \theta_{sje} = \theta_{bji} = \theta_{sji} = \theta_{sjm} = 0$.

2.3 Estimation method

As $r^{\$}_{ajt} = \mu^{\$}_{ajt} + \eta^{\$}_{ajt}$ by construction, the basic model for excess returns that we seek to estimate can be written in compact form as:

$$
\begin{array}{llll}
r^{\$}_{cjt} = \beta^{\$}_{cje} f^R_{et} & + & v_{jet} & \\
r^{\$}_{bjt} = \beta^{\$}_{bje} f^R_{et} + \beta_{bji} f^R_{it} & +\delta^{\$}_{bje} v_{jet} + & v_{jit} & \\
r^{\$}_{sjt} = \beta^{\$}_{sje} f^R_{et} + \beta_{sji} f^R_{it} + \beta_{sjm} f^R_{mt} & +\delta^{\$}_{sje} v_{jet} + \delta_{sji} v_{jit} + v_{jmt} & &
\end{array}
\tag{6}
$$

or in matrix notation $\mathbf{r}^{\$}_{jt} = \mathbf{B}^{\$}_j \mathbf{f}^R_t + \mathbf{\Delta}^{\$}_j \mathbf{v}_{jt}$, where f^R_{kt} $(k = e, i, m)$ is short-hand for $\tau_k \lambda_{kt} + f_{kt} = \pi_{kt} + f_{kt}$. Under the assumption of conditional normality, (6) could be estimated for any N countries simultaneously by maximum likelihood (see KSW for details). But with three assets per country, and a non-diagonal time-varying conditional idiosyncratic covariance matrix, this results in a very time consuming procedure even for moderately large N. However, estimation could be considerably simplified if we had data on the basis portfolios (see Sentana (1997a)). Unfortunately, we do not usually observe f^R_t directly. Nevertheless, if we construct three diversified (passive) portfolios consisting of currency deposits, bonds and stocks, with excess returns $r^{\$}_{cpt}, r^{\$}_{bpt}$ and $r^{\$}_{spt}$ respectively, a simple method of estimation and testing can still be developed, which is similar to the one used by Connor and Korajczyk (1988) for the static APT, and more recently by Forni and Reichlin (1998) and Stock and Watson (1998) for macroeconomic time series applications.

Given that the idiosyncratic errors are independent in mean from the $\delta' s$, the approximate zero factor structure assumption on the $v_{jt}'s$ implies that:

$$\begin{aligned} r^{\$}_{cpt} &= \beta^{\$}_{cpe} f^{R}_{et} \\ r^{\$}_{bpt} &= \beta^{\$}_{bpe} f^{R}_{et} + \beta_{bpi} f^{R}_{it} \\ r^{\$}_{spt} &= \beta^{\$}_{spe} f^{R}_{et} + \beta_{spi} f^{R}_{it} + \beta_{spm} f^{R}_{mt} \end{aligned} \tag{7}$$

Therefore, $\mathbf{r}^{\$}_{pt} = (r^{\$}_{cpt}, r^{\$}_{bpt}, r^{\$}_{spt})'$ is a full-rank, time-invariant transformation of the basis portfolios, whose factor loadings are a linear combination of the corresponding asset factor loadings. Since the scaling of the common factors is free, we can set $\beta^{\$}_{cpe} = \beta_{bpi} = \beta_{spm} = 1$ without loss of generality. Hence, in matrix notation, we can write $\mathbf{r}^{\$}_{pt} = \mathbf{B}^{\$}_{p}\mathbf{f}^{R}_{t}$, with $\mathbf{B}^{\$}_{p}$ unit lower-triangular and obviously $\left|\mathbf{B}^{\$}_{p}\right| = 1$. If we now add the three portfolios in (7) to the list of $3N$ assets under consideration, we can factorise the joint likelihood function into the marginal component of $\mathbf{r}^{\$}_{pt}$, and the conditional components corresponding to all the individual countries, which are given by

$$\mathbf{r}^{\$}_{jt} = \mathbf{B}^{\$*}_{j}\mathbf{r}^{\$}_{pt} + \mathbf{\Delta}^{\$}_{j}\mathbf{v}_{jt} \tag{8}$$

where the relationship between both sets of parameters is $\mathbf{B}^{\$*}_{j} = \mathbf{B}^{\$}_{j}(\mathbf{B}^{\$}_{p})^{-1}$.

Let us start with the marginal model for the portfolios. Since the factors are (conditionally) orthogonal, we can decompose the joint log-likelihood function of $\mathbf{r}^{\$}_{pt}$ (given the past) into the marginal component of $r^{\$}_{cpt}$ plus the conditional of $r^{\$}_{bpt}$ given $r^{\$}_{cpt}$ plus the conditional of $r^{\$}_{spt}$ given $r^{\$}_{cpt}$ and $r^{\$}_{bpt}$. This yields

$$\begin{array}{rcccccc} r^{\$}_{cpt} & = & (\tau_e\lambda_{et} + f_{et}) & & & & \\ r^{\$}_{bpt} & = & \beta^{\$}_{bpe} r^{\$}_{cpt} & + & (\tau_i\lambda_{it} + f_{it}) & & \\ r^{\$}_{spt} & = & (\beta^{\$}_{spe} - \beta^{\$}_{bpe}\beta_{spi}) r^{\$}_{cpt} & + & \beta_{spi} r^{\$}_{bpt} & + & (\tau_m\lambda_{mt} + f_{mt}) \end{array} \tag{9}$$

In the conditionally homoskedastic case, (9) is a recursive simultaneous equation system, and the parameter estimates are particularly simple to obtain:

(a) π_e and λ_e from the OLS regression of $r^{\$}_{cpt}$ on a constant

(b) $\beta^{\$}_{bpe}, \pi_i$ and λ_i from the OLS regression of $r^{\$}_{bpt}$ on $r^{\$}_{cpt}$ and a constant

(c) $\beta^{\$}_{spe}, \beta_{spi}, \pi_m$ and λ_m from the OLS regression of $r^{\$}_{spt}$ on $r^{\$}_{cpt}, (r^{\$}_{bpt} - \tilde{\beta}^{\$}_{bpe} r^{\$}_{cpt})$ and a constant.

However, there is a generated regressor problem in (c) which affects inferences involving $\beta^{\$}_{spe}$. Nevertheless, consistent standard errors for all the parameters in (9) can be obtained by regarding the above estimation method as GMM based on the just-identifying moment conditions implicit in (a), (b) and (c) simultaneously. A significant advantage of the GMM framework is that it is easy to see that the estimators for $\pi' s$ and $\beta' s$ remain consistent when f_{et}, f_{it} and f_{mt} follow univariate GARCH processes (and even when the prices of risk are time-varying), provided that the factor representing portfolios remain contemporaneously uncorrelated, although we must take into account that the residuals will be serially correlated in order to compute consistent standard errors.[7]

Let us now turn to the conditional models for $\mathbf{r}^{\$}_{jt}$ given $\mathbf{r}^{\$}_{pt}$. In the conditionally homoskedastic case, ML estimates of $\mathbf{B}^{\$*}_{j}$ can be obtained by using Seemingly Unrelated Regression techniques. If the idiosyncratic covariance matrix were block-diagonal, SUR applied to the $3N$ assets simultaneously would be equivalent to SUR applied to the three representative assets of each country at a time with $\mathbf{r}^{\$}_{pt}$ as regressors, and in addition, there would be no efficiency loss in estimating the model only for the N countries of interest. Besides, the triangular nature of $\mathbf{B}^{\$*}_{j}$ implies that (iterated) SUR estimates could be obtained simply as follows (see Sentana (1997b)):

(d) $\beta^{\$*}_{cje}$ and ω_{je} from the OLS regression of $r^{\$}_{cjt}$ on $r^{\$}_{cpt}$.

(e) $\beta^{\$*}_{bje}, \beta^{*}_{bji}, \delta^{\$}_{bje}$ and ω_{ji} from the OLS regression of $r^{\$}_{bjt}$ on $r^{\$}_{cpt}$ and $r^{\$}_{bpt}$ with the residual from (d) as an extra regressor.

[7] Maximum likelihood estimates of the trivariate 3-factor GARCH in mean model for $\mathbf{r}^{\$}_{pt}$ in (9) can be easily obtained by replacing OLS with univariate GARCH-M regressions in (a), (b) and (c) above, but the standard errors usually computed would be inconsistent due to the generated regressor problem.

(f) $\beta^{\$*}_{sje}, \beta^{*}_{sji}, \beta^{*}_{sjm}, \delta^{\$}_{sje}, \delta_{sji}$ and ω_{jm} from the OLS regression of $r^{\$}_{sjt}$ on $r^{\$}_{cpt}, r^{\$}_{bpt}$ and $r^{\$}_{spt}$ with the residuals from (d) and (e) as extra regressors.

Moreover, given the Frisch-Waugh theorem, it is straightforward to prove that the estimates of $\mathbf{B}^{\$}_{j}$ could be obtained directly if we ran the OLS regressions in (d), (e) and (f) with the orthogonalised portfolios, $\hat{\mathbf{f}}^{R}_{t} = (\tilde{\mathbf{B}}^{\$}_{p})^{-1}\mathbf{r}^{\$}_{pt}$ instead. But as before, the standard errors associated with $\beta^{\$}_{bje}, \beta^{\$}_{sje}, \beta_{sji}$ and $\delta^{\$}_{sje}$ would suffer from a generated regressor bias. Nevertheless, consistent standard errors for all the parameters could be obtained by recasting the above estimation method as GMM based on the just-identifying moment conditions implicit in (a), (b) and (c) together with (d), (e) and (f) for all countries simultaneously. A significant advantage of the GMM framework is that it is straightforward to conduct inferences involving parameters from different equations that remain valid when some of the assumptions made for estimation, such as block diagonality of the idiosyncratic covariance matrix, do not hold. Furthermore, it is also easy to see that the GMM parameter estimators remain consistent even if $\omega_{jet}, \omega_{jit}$ and ω_{jmt} follow GARCH processes.[8]

The different alternative hypotheses discussed above can also be easily tested in this GMM-regression framework. For instance, to see if idiosyncratic exchange rate variability affects currency risk premia, we can test whether adding the estimated conditional variance of υ_{jet}, $\tilde{\omega}_{jet}$, as an extra regressor in equation (6a) improves the explanatory power of the equation. The rationale comes from the fact that (6a) can be rewritten as $E_{t-1}(r^{\$}_{cjt} - \beta^{\$}_{cje} r^{\$}_{cpt}) = 0$. Such a testing procedure also yields as a by-product an estimate of the coefficient for ω_{jet} under the alternative, ρ_{cje}, which is of interest to measure the effects of reducing the variability of idiosyncratic risks.

[8] Maximum likelihood estimates of the factor loadings for each country model in (8) can be obtained by replacing OLS in (d), (e) and (f) with univariate GARCH regressions, but again the standard errors usually computed would be inconsistent because of the generated regressor problem.

In view of the relatively short number of observations, we choose univariate GARCH (1,1) specifications for the conditional variances of both common and idiosyncratic factors. Such a parametrisation has been found to be a good representation of many financial time series. More efficient conditional variance estimates could be obtained by using a larger information set, or taking into account peso-type effects. However, note that even if the GARCH(1,1) model were incorrect, our tests would still be consistent, albeit less powerful, since they are simply rational expectations-type orthogonality tests, and the GARCH model generates a conditional variance estimate which, conditional on the parameters, is a function of past information. In this respect, it is important to emphasise that our tests will have the correct asymptotic size under the null despite the fact that the estimated conditional variances are generated regressors (see Pagan and Ullah (1987)).

3 Empirical application

3.1 Data

We estimate the model described above using monthly data for currency, bond and stock returns (in percentage terms) on ten European countries between October 1977 and October 1997 (i.e. 241 observations), with the first observation used for initialising the conditional variance recursions. Details of data sources and transformations can be found in the appendix. The ten countries are Belgium, Denmark, France, Germany, Italy, the Netherlands, Spain, Sweden, Switzerland and the UK. Importantly, note that the returns on both bonds and particularly stocks for each country correspond already to a well-diversified basket of domestic assets. But since it is crucial that the aggregate portfolios contain the larger non-European countries as well, we also have data on Australia, Canada, Japan and the US. In this respect, we consider equally weighted world currency, bond and stock returns as our set of portfolios in (7), which are linearly related to the

OLS representing portfolios obtained from (6) under the assumption that $\beta_{ajk} = 1$ $(a = c, b, s; k = e, i, m)$ for all the countries that we have data on (see Sentana (2000)).[9]

It is important to mention that Sweden and Switzerland have never belonged to the EMS, although their currencies were unofficially pegged to the DM for part of the sample. The UK only entered the ERM in October 1990, to leave in September 1992 together with Italy, who rejoined the system in November 1996. Finally, Spain, who joined in June 1989, was forced to realign the central parity of its currency four times between September 92 and March 95.[10] These three countries negotiated a wide margin of $\pm 6\%$ for their bilateral exchange rates, while the other countries, including Italy between 1989 and 1992, participated with bilateral limits of $\pm 2.25\%$ until August 1993, when the bands were widened to $\pm 15\%$.

3.2 Estimates of the asset pricing model

We initially estimated the 195 parameters characterizing the model under the null by maximum likelihood as explained in section 2.3. As a by-product, we obtained estimates of the conditional variances of common and specific factors. Unfortunately, the computation of consistent standard errors a la Bollerslev and Wooldridge (1992), which avoid the generated regressor problem and are robust to non-normality, is not an easy task in our case, as it involves the outer product of the gradient and the Hessian matrix of the joint log-likelihood function for

[9] It is in principle possible, albeit non-trivial, to estimate the optimal weightings of the basis portfolios together with all the other parameters (see e.g. KSW). In any case, note that the consistency of the estimation methods described in section 2.3 would not be affected if the international portfolios that we construct were not fully diversified, although it would be necessary to develop an asymptotic theory in which both N and T increase at possibly the same rate in order to robustify our inference. Both issues are left for further research.

[10] A complete list of ERM realignments between March 1979 and December 1994 can be found in Ayuso and Pérez-Jurado (1997).

the 33 asset returns. In addition, the outer-product matrix is singular under some of the alternatives because the number of parameters exceeds the number of observations. For that reason, we only present the GMM estimates discussed in section 2.3, which turn out to be rather close to the ML ones.[11] As for standard errors, we use the Newey and West (1987) formulas with a baseline bandwidth of 6 lags ($\simeq 240^{1/3}$). Qualitatively similar results are obtained by doubling or halving the lag length. Nevertheless, it is important to emphasise that the finite sample properties of the covariance matrix estimators and associated test procedures in our set up are unknown.

The results for the subsystem (7) are presented in Table 1. As one would expect from the fact that all returns are US \$ denominated, the estimated factor loadings confirm that all three portfolios are positively correlated. Besides, they also confirm that, controlling for exchange rate variations, world bond and stock returns are positively correlated. In this respect, note that since $\beta^{\$}_{pbe}$ is very close to one, the interest basis portfolio is hardly distinguishable from an equally-weighted average of excess bond returns in local currency. The estimates of the $\pi's$, though, are imprecise. The average expected returns on common risks are all positive, but insignificant at conventional levels, with the possible exception of the common interest rate risk if we consider the more relevant one-sided test.[12]

The GMM estimates of the factor loadings for each of the three asset classes can be found in Tables 2a-2c. For clarity of exposition, but without loss of generality, the results for bonds and stocks correspond hereinafter to local currency returns. Since $r^j_{bjt} = r^{\$}_{bjt} - r^{\$}_{cjt}$ and $r^j_{sjt} = r^{\$}_{sjt} - r^{\$}_{cjt}$, the only coefficients affected are those

[11] A useful way to measure the "distance" between ML and GMM parameter estimates is to look at the differences between the implied basis portfolios for common and specific factors. In this respect, we find that both sets of estimates yield very similar results, with correlations ranging from .979 to 1, with an average value of .997

[12] If we compute "average" price of risk coefficients for $k = e, i, m$ as $\hat{\pi}_k/\hat{\lambda}_k$, we obtain .001177, .140906 and .030949, which although strictly speaking inconsistent, are rather similar to the corresponding ML estimates (.001627, .121238 and .028841).

related to the effects of common and specific exchange rate risks. In particular $\beta^j_{bje} = \beta^\$_{bje} - \beta^\$_{cje}$, $\delta^j_{bje} = \delta^\$_{bje} - 1$, $\beta^j_{sje} = \beta^\$_{sje} - \beta^\$_{sje}$ and $\delta^j_{sje} = \delta^\$_{sje} - 1$.

The coefficient on the common exchange rate factor is positive and statistically significant for every currency return (Table 2a). Turning to the effect of the common exchange rate on bond returns, note that the coefficients are both positive and negative, but only significantly positive for Germany. This could be because for this country, whose central bank was traditionally concerned about imported inflation, a dollar depreciation leads to lower inflation, and hence, higher bond returns. Table 2b, column 2, shows that in all cases, increases in the common interest rate factor (i.e. the "world" bond return) are associated with higher European bond returns. Not surprisingly, this effect is always statistically significant. Finally, the idiosyncratic exchange rate factors are significantly positive for Germany, Italy, the Netherlands and the UK. Again, the intuition would be that a local currency appreciation leads to lower inflation and higher bond returns. We next examine the factor loadings for stock returns (Table 2c). The coefficients on the common exchange rate are negative in all cases, and significantly so for many countries. Hence, local currency stock returns in Europe suffer when the dollar depreciates. On the other hand, the coefficients on the common interest rate factor suggests that periods of higher world bond returns would tend to be associated with higher European stock market returns. In addition, the coefficients on the common residual market risk factor are all significantly positive. Turning to the effects of the idiosyncratic factors on stock returns, we can see that the coefficients of country-specific exchange rate factors are positive in some cases and negative in others. In particular, an idiosyncratic local currency appreciation significantly increases stock returns in France and Germany, while decreasing them in Sweden. The fact that for a given country the sensitivity of returns to common and idiosyncratic exchange rate movements is different is likely to reflect the structure of its foreign trade. Finally, the results also suggest that controlling for exchange rate variations, the idiosyncratic terms for bonds and stocks are positively and significantly correlated within countries.[13]

3.3 Direct effects of ERM on the cost of capital

Figure 1a displays the average of the conditional standard deviations of the idiosyncratic exchange rate components in ERM and non-ERM countries. For these purposes, and in line with the consensus view among both academics and practitioners, we only include Belgium, Denmark, France, Germany and the Netherlands in the ERM bloc. Notice that the temporal evolution in both groups is somewhat similar, with a fairly tranquil period between 1983 and 1992, preceded and followed by more volatile ones. Not surprisingly, the effects of the successive EMS crises that began in September 92 are noticeable in the two series, since the speculative attacks affected all European currencies but the Dutch guilder. Nevertheless, the average level and magnitude of the movements are rather different, being substantially smaller for those countries that did not modify the central parity of their currencies against the ECU. Therefore, there is clear evidence that a credible target zone system does reduce exchange rate volatility.

Figure 1b is the analogous picture for the idiosyncratic interest rate components. Except during the first part of the sample, and a brief period between March and May 1990, average interest rate volatility was actually smaller in ERM countries than in non-ERM ones. At the same time, it also seems that during turbulent periods in the foreign exchange market, idiosyncratic interest rate volatility

[13] As discussed in section 2.3, our estimation procedure is based on the maintained assumption that the covariance matrix of the idiosyncratic terms satisfies the approximate zero factor structure of Chamberlain and Rothschild (1983). We have informally assessed whether such an assumption is correct on the basis of the correlation matrix of the estimated idiosyncratic terms for all the countries in our dataset. In this respect, we find that the average of its offdiagonal elements (squared) is .017518. Similarly, we find that its largest eigenvalue only explains 10.5% of its trace, as opposed to 13% when we only consider the European countries.

increased more for those countries which insisted in maintaining their currencies within the bands. Hence, our results confirm that the reduction in idiosyncratic exchange rate volatility may sometimes be achieved at the expense of increases in local interest rate volatility.

According to our basic asset pricing model, though, such volatility movements should have no impact on expected returns in a world of complete financial market integration. For that reason, we begin by testing whether contrary to the theory, idiosyncratic exchange rate and interest rate risks are priced. The results for the pricing of idiosyncratic exchange rate volatility in the currency, bond and stock markets of ERM and non-ERM countries are presented in Table 3 (rows 1 to 3 of panels a and b, respectively). On the currency markets, the joint Wald tests clearly indicate a rejection of the null hypothesis, both in core ERM countries, and especially non-ERM ones. As for local currency bond and stock returns, the tests would reject in the case of non-ERM countries, but not in the case of core ERM ones.

We also find that idiosyncratic interest rate seems to be priced in bond markets (see row 4 of panels a and b), but in stock markets the evidence against the null is only significant for non-ERM countries (see row 5). In this respect, it is important to bear in mind that all tests have, ceteris paribus, lower power in the stock return equations because the standard deviation of risks specific to stocks in each country are much higher (compare the idiosyncratic variances in Tables 2a, 2b and 2c). Our overall negative results are perhaps not totally surprising in view of the existence of exchange controls (at least for part of the sample) and other barriers to cross-border investment, as well as informational asymmetries, behavioural biases, liquidity premia, etc. (see also section 3.4 below).

Given that the individual coefficients show great dispersion, being even negative sometimes, and that the effects of exchange rate and interest rate volatility movements on risk premia may partly compensate each other, we have attempted to get a measure of the net effects of idiosyncratic exchange rate and interest rate volatility on each asset by computing the differences in fitted values between alternative and null. Note that in this way each country acts as its own control. Figure 2a presents the average net effect on currency returns across ERM and non-ERM countries, while Figures 2b and 2c display the analogous effects for bond and stock returns respectively. The corresponding sample means and relevant t-statistics are reported in Table 4. Please note the changes in scale, which reflect the differences in average risk premia across asset classes.

If we ignore the negative spikes, which correspond to devaluations in the Swedish krona, the evidence on currency returns suggests that for both group of countries, idiosyncratic exchange rate variability significantly increases short interest rate differentials for a given expected depreciation. This could explain the convergence of short rates in Europe during periods of low intra-European exchange rate volatility. Therefore, a system that reduced both the expected level and the volatility of idiosyncratic exchange rate movements would seem to reduce the riskless component of the cost of capital. The evidence for bond and stock markets is less clear cut, with a small negative effect in the ERM bloc on average over time, and a significantly larger, positive effect in the other group. Nevertheless, a comparison of Figures 2b and 2c with Figure 1a suggests that periods in which local exchange rate volatility has risen have tended to be associated with increases in the required rate of return on bonds and equities. If with all the usual caveats in place, we were to extrapolate these results to a currency system such as EMU, in which country-specific exchange rate uncertainty has almost vanished, and the interest rate policy of the member countries is not only common, but more importantly, released from the obligation of maintaining intra-European exchange rate stability, our tentative conclusion would be that, ceteris paribus, the cost of capital in euroland is likely to be lower than it would have been otherwise, although the gains would probably be fairly small.

3.4 The gains from globalisation

The international asset pricing model described in section 2.1 implicitly assumes that European financial markets are integrated. Specifically, we have assumed that idiosyncratic risk is not priced, and that the price of risk associated with each of the underlying common factors is the same for all countries. Therefore, one way of testing for market integration is to examine if idiosyncratic factors are priced. We have already done this in the previous section for exchange rate and interest rate risks. The results of the tests for the pricing of country-specific residual market risk in the equity markets show that the coefficients are not jointly significantly different from zero (see Table 5, row 1), although again this is probably due to lack of power.

The other way to assess if European markets are integrated is to test whether the prices of risk associated with the common factors are the same across countries in each of the three asset classes. The null hypothesis of common exchange rate risk evaluation is not rejected in the currency markets (Table 5, row 2), but it is strongly rejected in the bond and stock markets (rows 3 and 4). As for the common interest rate factor, the assumption of common valuation is not rejected in European bond markets (see row 5), but it is rejected in stock markets (row 6). Finally, we also find significant differential pricing for common residual market risks (see row 7).

One attractive reinterpretation of our lack of integration results is to say that international investors are incompletely diversified, which is consistent with the well known fact that most investors show a home bias, in the sense that they hold only a relatively small proportion of their assets outside their country (see e.g.

Kang and Stulz (1997) for a discussion and references). In this respect, an important indirect effect of EMU would be the elimination of some of the elements that limit cross-border investments in Europe, such as the restrictions on foreign asset holdings by pension funds, insurance companies and other financial institutions, or the behavioural biases traditionally shown by individual investors in favour of assets denominated in their own domestic currency. In fact, there is ample anecdotal evidence that such a process is already taking place at a very rapid rate. In order to gauge the potential gains from increased market integration, we are going to follow Stulz (1999) in comparing stock market risk premia under full integration with the risk premia that would prevail in the context of fully segmented markets. In this sense, note that although the distinction between common and specific risk becomes irrelevant in the latter context, it is still possible to write the unanticipated components of stock returns in terms of three orthogonal sources of risk:

$$\begin{array}{rlll}\eta^j_{sjt} = & \beta^j_{sje} f_{et} + \delta^j_{sje} \upsilon_{jet} & +\beta_{sji} f_{it} + \delta_{sji} \upsilon_{jit} & +\beta_{sjm} f_{mt} + \upsilon_{jmt} \\ & \text{Exchange rate risk} & \text{Interest rate risk} & \text{Residual market risk}\end{array} \tag{10}$$

Since the stock portfolio for each country corresponds to a diversified basket of domestic stocks, an argument similar to the one presented in section 2.1 would then result in the following domestic version of the APT:

$$\mu^{js}_{sjt} = \varphi_{je}(\beta^{j2}_{sje}\lambda_{et} + \delta^{j2}_{sje}\omega_{jet}) + \varphi_{ji}(\beta^{2}_{sie}\lambda_{it} + \delta^{2}_{sie}\omega_{jit}) + \varphi_{jm}(\beta^{2}_{sjm}\lambda_{mt} + \omega_{jmt}) \tag{11}$$

where φ_{je}, φ_{ji} and φ_{jm} are the prices of risk in country j. Note that (11) is a special case of our general alternative hypothesis (5), with the restrictions $\rho_{sjk} = \varphi_{sjk} = \tau_k$ and $\theta_{sjk} = \varphi_{sjk}\beta_{jsk}(\beta_{jsk}-1)$ for $k = e, i, m$. Assuming as in Stulz (1999) that financial market integration does not affect the prices of residual market risk, so that $\varphi_{jm} = \tau_m$, we can assess whether there would be gains (on average over time) to each country from stock market integration by comparing the following quantities:[14]

$$\beta_{sjm}E(\lambda_{mt}) \lesseqgtr \beta^2_{sjm}E(\lambda_{mt}) + E(\omega_{jmt})$$

Table 6 reports the difference between the right and left hand sides of the above expression for the European countries in our database. Apart from the fact that all the differences are highly significantly positive, the other striking result is that there is substantial variation across countries. In particular, the largest average gains correspond to countries such Italy, Sweden or Spain, which tend to have both large $\beta'_{sjm}s$, and especially, very large idiosyncratic variances. As a result, those countries are the ones that a priori would benefit most from an increase in stock market integration. In contrast, the benefits for countries such as the Netherlands, Switzerland or the UK, whose stock markets have significantly smaller residual market risk, probably because they already have closer links with world markets, would be smaller. If we multiply those differences by .03, which is roughly our implicitly estimate of τ_m (see footnote 10), our results suggest that the potential gains from stock market integration could be rather large. Nevertheless, it is important to emphasise that those gains should only be taken as indicative, in view of the large standard error associated with the price of risk coefficient, and the fact that we are comparing a situation of full segmentation with another of complete integration. In practice, of course, markets are neither fully segmented, nor fully integrated, and moreover, the transition from one state to the other is typically a gradual process, whose effects are partly anticipated by investors. In this respect, Hardouvelis, Malliaropulos and Priestley (2000) find that the degree of integration of European stock markets increased substantially after 1995, when forward interest differentials vis-à-vis Germany began to narrow in anticipation

[14] Given the definition of beta as the ratio of covariance to variance, and the fact that for innovations the unconditional variance coincides with the average of the conditional one, we can equivalently write this expression in terms of covariances, as in Stulz (1999), yielding $E\left[cov_{t-1}(r^j_{sjt}, f^R_{mt})\right] \lesseqgtr E\left[V_{t-1}(\beta_{sjm}f^R_{mt} + v_{jmt})\right]$

of EMU membership.

4 Conclusions

In this paper, we use monthly data on currency, bond and stock returns for ten European countries over the period 1977-1997 to estimate a dynamic multi-factor APT model with time-varying volatility in both common and idiosyncratic factors. From the methodological point of view, our main contribution is to exploit the cross-sectional dimension of the model to construct diversified world portfolios, which when added to the original list of assets, can be used to develop simple, but nevertheless consistent GMM-based methods of estimation and testing.

We find that, controlling for exchange rate variations, world bond and stock returns are positively correlated. We also find that the average expected returns on the common risks are all positive, but not very precisely estimated. Moreover, our results show that increases in the common exchange rate, interest rate and residual market factors are associated with higher European currency, bond and stock returns respectively. They also suggest that a local currency appreciation generally leads to higher (local currency) bond returns, while (local currency) stock returns in Europe suffer when the dollar depreciates. In addition, we find that controlling for exchange rate variations, the idiosyncratic terms for bonds and stocks are positively and significantly correlated within countries.

Furthermore, our findings indicate that a target zone system such as the EMS does reduce exchange rate volatility as long as it remains credible, and that average interest rate volatility has actually been smaller in ERM countries than in non-ERM ones. At the same time, our results confirm that reductions in idiosyncratic exchange rate volatility may sometimes be achieved at the expense of increases in local interest rate volatility. Importantly, our evidence also suggests that a system that reduces both the expected level and the volatility of idiosyncratic exchange rate movements is likely to reduce the riskless component of the cost of capital. The evidence for bond and stock markets is less clear cut, although periods in which local exchange rate volatility has risen have tended to be associated with increases in the required rate of return on bonds and equities. Nevertheless, the effects that we uncover are small.

We also find overall negative results on market integration. For that reason, and given that an important indirect effect of EMU would be an increase in the degree of integration of European financial markets, we compare stock market risk premia under full integration with the risk premia that would prevail in the context of fully segmented markets. Our results suggest that such an upper bound on the potential gains from stock market integration could be rather large.

Finally, it is worth bearing in mind that there are other important channels through which the removal of financial market segmentation by means of regional arrangements such as the EMS could lead to a decrease in funding costs. Given the separation of ownership and control that exists in the quoted corporate sector, the most important mechanism will be the reduction in the agency costs faced by firms raising outside capital, which are not directly related to the required rate of return used in present discounted value calculations (see Stulz (1999)). At the same time, the creation of a wider, deeper and more liquid capital market is likely to intensify the pressure on European firms to pay more attention to shareholders, and to seek to raise the value of their stocks. In this respect, cross-border mergers and acquisitions show a clear upward trend. In addition, increased competition by banks and other financial institutions may also result in a reduction in the cost of capital for smaller unquoted companies, which account for a significant proportion of European investment. Therefore, there is little doubt that the measuring of the effects of "globalisation" on the cost of capital would continue to be an area of active research.

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Data Appendix

Details of the data series used are as follows (where appropriate the name of the series is followed by its Datastream code).

Stock prices and dividend yields:

* Morgan Stanley Capital International Perspectives

Exchange rates:

* US$-Aus$, end of period (AUOCEXCH)
* US$-BelFr, end of period (BGOCEXCH)
* US$-CAN$, end of period (CNOCEXCH)
* US$-DENKr, end of period (DKOCEXCH)
* US$-FraFr, end of period (FROCEXCH)
* US$-DM, end of period (BDOCEXCH)
* US$-ItaLit, end of period (ITOCEXCH)
* US$-Yen, end of period (JPOCEXCH)
* US$-NetFl, end of period (NLOCEXCH)
* US$-SpPta, end of period (ESOCEXCH)
* US$-SweKr, end of period (SDOCEXCH)
* US$-SwiFr, end of period (SWOCEXCH)
* US$-Stg, end of period (UKOCEXCH)

Short interest rates:

* Belgium, Canada, Denmark, France, Germany, Italy, Japan, Netherlands, Sweden, Switzerland, UK, USA: 1-month Eurocurrency rate from the Bank of England.

* Australia: Yield on 3-month Treasury Bill (AUTRSBL%)

* Spain: pre 1983:1: 3-month interbank rate from the Bank of Spain; post 1983:1: 1-month Eurocurrency rate.

Bond returns:

* Canada, France, Germany, Japan, Netherlands, Switzerland, UK, USA: pre 1978:1: Shiller (1979) approximation on 10-year bond yields from Goldman Sachs; post 1978:2: Salomon Brothers bond return indices.

* Australia: pre-1984:10: Shiller (1979) approximation on 10-year bond yield; post 1984:11: Salomon Brothers bond return indices

* Italy: pre 1990:12: Shiller (1979) approximation on 10 year bond yields from Goldman Sachs; post 1991:1: Salomon Brothers bond return indices.

* Belgium, Denmark, Sweden: pre 1985:3: Shiller (1979) approximation on 10 year bond yields from Goldman Sachs; 1985:4-1990:12: Datastream bond return indices (Belgium (ABGGYG4(RI)), Denmark (ADKGYG4(RI)), Sweden (ASDGYG4(RI))); post 1991:1: Salomon Brothers bond return indices.

* Spain: pre 1982:1: Shiller (1979) approximation on bond yields from Bank of Spain; 1982:2-1985:3: Shiller (1979) approximation on bond yields from Goldman Sachs; 1985:3:1990:12: Datastream bond return indices (AESGYG4(RI)); post 1991:1: Salomon Brothers bond return indices.

Table 1

Prices of risk and factor loadings for equally-weighted world portfolios

Sample period 1977:11-1997:10

World portfolio	Common exchange rate risk	Common interest rate risk	Common stock market risk	Unconditional variances
Currencies	$\beta^{\$}_{cpe} = 1$			$\lambda_e = 7.18554$ (.746466)
Bonds	$\beta^{\$}_{bpe} = .979692$ (.038047)	$\beta_{bpi} = 1$		$\lambda_i = 1.54193$ (.203489)
Stocks	$\beta^{\$}_{spe} = .706310$ (.115252)	$\beta_{spi} = 1.33681$ (.212905)	$\beta_{spm} = 1$	$\lambda_m = 11.3880$ (3.39880)
Unconditional risk premia	$\pi_e = .00849632$ (.198522)	$\pi_i = .219093$ (.126910)	$\pi_m = .356913$ (.243211)	

GMM estimates of equation (7): $\mathbf{r}^{\$}_{pt} = \mathbf{B}^{\$}_{p}\mathbf{f}^{R}_{t}$, $\mathbf{r}^{\$}_{pt} = (r^{\$}_{cpt}, r^{\$}_{bpt}, r^{\$}_{spt})'$, $\mathbf{f}^{R}_{t} = (f^{R}_{et}, f^{R}_{it}, f^{R}_{mt})'$, $f^{R}_{kt} = \tau_k\lambda_{kt} + f_{kt} = \pi_{kt} + f_{kt}$, $\pi_k = E(\pi_{kt})$, $\lambda_k = E(\lambda_{kt}) = V(f_{kt})$ $(k = e, i, m)$.

Newey-West (1987) heteroskedasticity and autocorrelation robust standard errors in parenthesis.

Table 2a

Factor loadings for currency returns ($)

Sample period 1977:11-1997:10

Country	Common exchange rate risk ($\beta^{\$}_{cje}$)	Idiosyncratic variance (ω_{je})
Belgium	1.26526 (.025079)	.856507 (.160474)
Denmark	1.22027 (.023217)	.894656 (.125631)
France	1.20585 (.024049)	.987137 (.152557)
Germany	1.24651 (.023489)	.848515 (.092211)
Italy	1.04731 (.047859)	2.17514 (.606642)
Netherlands	1.26122 (.025452)	.817188 (.105427)
Spain	1.05542 (.049952)	2.40062 (.602722)
Sweden	1.00790 (.056559)	2.87831 (1.06253)
Switzerland	1.33764 (.036824)	2.61699 (.384713)
UK	1.02758 (.063866)	4.21787 (.609914)

GMM estimates of equation (6a): $r^{\$}_{cjt} = \beta^{\$}_{cje} f^{R}_{et} + v_{jet}$, $\omega_{je} = E(\omega_{jet}) = V(v_{jet})$.

Newey-West (1987) heteroskedasticity and autocorrelation robust standard errors in parenthesis.

Table 2b

Factor loadings for bond returns (local currency)

Sample period 1977:11-1997:10

Country	Common exchange rate risk (β^j_{bje})	Common interest rate risk (β_{bji})	Specific exchange rate risk (δ^j_{bje})	Idosyncratic variance (ω_{ji})
Belgium	.012869	.746426	-.010701	1.03904
	(.040965)	(.063506)	(.099932)	(.134862)
Denmark	-.031760	1.15055	.05528	4.34104
	(.052475)	(.098457)	(.135574)	(1.02396)
France	-.004692	1.02727	.082677	1.40630
	(.047316)	(.072686)	(.091778)	(.211966)
Germany	.143256	.929245	.270939	.882759
	(.056360)	(.065022)	(.074528)	(.147316)
Italy	-.047515	.962191	.347770	2.88643
	(.053223)	(.104847)	(.078079)	(.471932)
Netherlands	.069076	.908067	.225466	.863978
	(.042638)	(.059902)	(.077839)	(.153695)
Spain	-.086345	.819648	-.014769	2.71115
	(.053599)	(.133285)	(.079147)	(.248375)
Sweden	-.045976	.788694	.060973	2.70035
	(.052172)	(.115046)	(.122362)	(.353565)
Switzerland	.049296	.590225	.067478	1.04408
	(.039099)	(.075243)	(.041963)	(.123455)
UK	.056706	1.36035	.277993	3.94258
	(.056990)	(.109731)	(.092055)	(.564751)

GMM estimates of equation (6b): $r^j_{bjt} = \beta^j_{bje} f^R_{et} + \beta_{bji} f^R_{it} + \delta^j_{bje} v_{jet} + v_{jit}$, $\omega_{ji} = E(\omega_{jit}) = V(v_{jit})$.
Newey-West (1987) heteroskedasticity and autocorrelation robust standard errors in parenthesis.

Table 2c

Factor loadings for stock returns (local currency)

Sample period 1977:11-1997:10

Country	Common exchange rate risk (β^j_{sje})	Common interest rate risk (β_{sji})	Common market risk (β_{sjm})	Specific exchange rate risk (δ^j_{sje})	Specific interest rate risk (δ_{sji})	Idiosyncratic variance (ω_{jm})
Belgium	-.322000	1.54156	.869533	.128528	.752030	11.0754
	(.1168345)	(.225872)	(.072319)	(.244034)	(.203083)	(2.26265)
Denmark	-.374027	1.30474	.642234	-.247583	.627680	13.9370
	(.120046)	(.212141)	(.086603)	(.298787)	(.132589)	(1.65446)
France	-.117952	1.66850	1.19002	.790233	1.19843	14.4385
	(.183294)	(.300995)	(.069961)	(.378364)	(.268907)	(2.19152)
Germany	-.239521	1.50536	1.07267	.774702	.695732	12.4569
	(.148164)	(.298862)	(.078756)	(.254677)	(.217266)	(1.87028)
Italy	-.453854	1.37731	1.15685	.458841	.575473	30.5493
	(.195227)	(.418395)	(.191447)	(.238861)	(.225265)	(4.73570)
Netherlands	-.509754	1.56830	.979499	-.239149	.333657	6.80658
	(.143979)	(.227616)	(.064407)	(.176940)	(.165672)	(.932920)
Spain	-.264740	1.31521	1.08326	-.057328	.724307	19.1805
	(.147007)	(.400857)	(.101797)	(.195961)	(.179790)	(2.55937)
Sweden	-.493719	1.29952	1.19243	-.644218	.570327	21.0370
	(.171656)	(.367956)	(.121713)	(.287745)	(.231128)	(2.55007)
Switzerland	-.286744	1.56240	.895178	-.035626	.755266	7.66734
	(.146489)	(.260817)	(.079937)	(.137907)	(.158261)	(.856350)
UK	-.221163	1.46592	1.01946	-.122607	.689536	8.18090
	(.142131)	(.262333)	(.080639)	(.120984)	(.114595)	(1.04401)

GMM estimates of equation (6c): $r^j_{sjt} = \beta^j_{sje} f^R_{et} + \beta_{sji} f^R_{it} + \beta_{sjm} f^R_{mt} + \delta^j_{sje} v_{jet} + \delta_{sji} v_{jit} + v_{jmt}$, $\omega_{jm} = E(\omega_{jmt}) = V(v_{jmt})$.

Newey-West (1987) heteroskedasticity and autocorrelation robust standard errors in parenthesis.

Table 3

Joint tests for pricing of idiosyncratic exchange rate and interest rates risks

Sample period 1977:11-1997:10

Panel a: Core ERM countries				
Null hypothesis	Risk	Asset	Joint Wald test	p-value
$\rho^{\$}_{cje} = 0\ \forall j$	exchange rate	currencies ($)	12.7396	0.02594
$\rho^{j}_{bje} = 0\ \forall j$	"	bonds (l.c.)	7.8618	0.16402
$\rho^{j}_{sje} = 0\ \forall j$	"	stocks (l.c.)	2.2047	0.82016
$\rho_{bji} = 0\ \forall j$	interest rate	bonds	11.7580	0.03826
$\rho_{sji} = 0\ \forall j$	"	stocks	8.3415	0.13839

Panel b: Non-ERM countries				
Null hypothesis	Risk	Asset	Joint Wald test	p-value
$\rho^{\$}_{cje} = 0\ \forall j$	exchange rate	currencies ($)	47.4100	0.0
$\rho^{j}_{bje} = 0\ \forall j$	"	bonds (l.c.)	11.3007	0.04573
$\rho^{j}_{sje} = 0\ \forall j$	"	stocks (l.c.)	16.3413	0.00593
$\rho_{bji} = 0\ \forall j$	interest rate	bonds	17.2587	0.00403
$\rho_{sji} = 0\ \forall j$	"	stocks	11.8358	0.03711

Computed using Newey-West (1987) heteroskedasticity and autocorrelation consistent covariance matrix estimator.

Table 4

Net effect of idiosyncratic exchange rate and interest volatility on returns

Sample period 1977:11-1997:10

Asset	Countries	Average	t-statistic	p-value
	core ERM	.029019	35.3667	0.
Currencies ($)	non-ERM	.022681	5.0649	0.
	difference	-.006338	-1.4525	0.146
	core ERM	-.022311	-1.5830	0.113
Bonds (l.c.)	non-ERM	.049634	5.1728	0.
	difference	.071945	5.1430	0.
	core ERM	-.020553	-2.5034	0.012
Stocks (l.c.)	non-ERM	.065158	2.6136	0.009
	difference	.085711	3.4089	0.001

Computed using Newey-West (1987) heteroskedasticity and autocorrelation consistent covariance matrix estimator.

Table 5

Additional joint tests for market integration

Sample period 1977:11-1997:10

Null hypothesis	Risk	Asset	Joint Wald test	p-value
$\rho_{sjm} = 0 \ \forall j$	residual market	stocks	13.1805	0.21376
$\theta^{\$}_{cje} = 0 \ \forall j$	exchange rate	currencies ($)	16.1045	0.09668
$\theta^{j}_{bje} = 0 \ \forall j$	"	bonds (l.c.)	31.3994	0.00050
$\theta^{j}_{sje} = 0 \ \forall j$	"	stocks (l.c.)	103.3475	0.0
$\theta_{bji} = 0 \ \forall j$	interest rate	bonds	11.2748	0.33651
$\theta_{sji} = 0 \ \forall j$	"	stocks	26.1813	0.00350
$\theta_{sjm} = 0 \ \forall j$	residual market	stocks	142.2266	0.0

Computed using Newey-West (1987) heteroskedasticity and autocorrelation consistent covariance matrix estimator.

Table 6

Average gains from increased stock market integration (local currency)

Sample period 1977:11-1997:10

Country	$\beta^2_{sjm}\lambda_m + \omega_{jm}$	$\beta_{sjm}\lambda_m$	Difference
Belgium	19.6857	9.90221	9.78350
	(4.43808)	(3.42981)	(2.30608)
Denmark	18.6341	7.31373	11.3204
	(2.06554)	(1.74589)	(2.02143)
France	30.5656	13.5519	17.0136
	(5.52645)	(3.74877)	(2.84319)
Germany	25.5600	12.2155	13.3445
	(5.68106)	(4.08642)	(2.40860)
Italy	45.7898	13.1742	32.6156
	(6.33277)	(2.51710)	(5.46835)
Netherlands	17.7324	11.1545	6.57790
	(3.73677)	(3.32912)	(1.21043)
Spain	32.5438	12.3362	20.2077
	(5.89328)	(3.97726)	(3.08184)
Sweden	37.2293	13.5793	23.6500
	(5.21221)	(3.48773)	(3.14164)
Switzerland	16.7930	10.1943	6.59876
	(4.33630)	(3.78708)	(1.06078)
UK	20.0163	11.6095	8.40678
	(5.49114)	(4.02881)	(1.70797)

Newey-West (1987) heteroskedasticity and autocorrelation robust standard errors in parenthesis.

Figure 1a: Average conditional standard deviation of idiosyncratic exchage rate factors

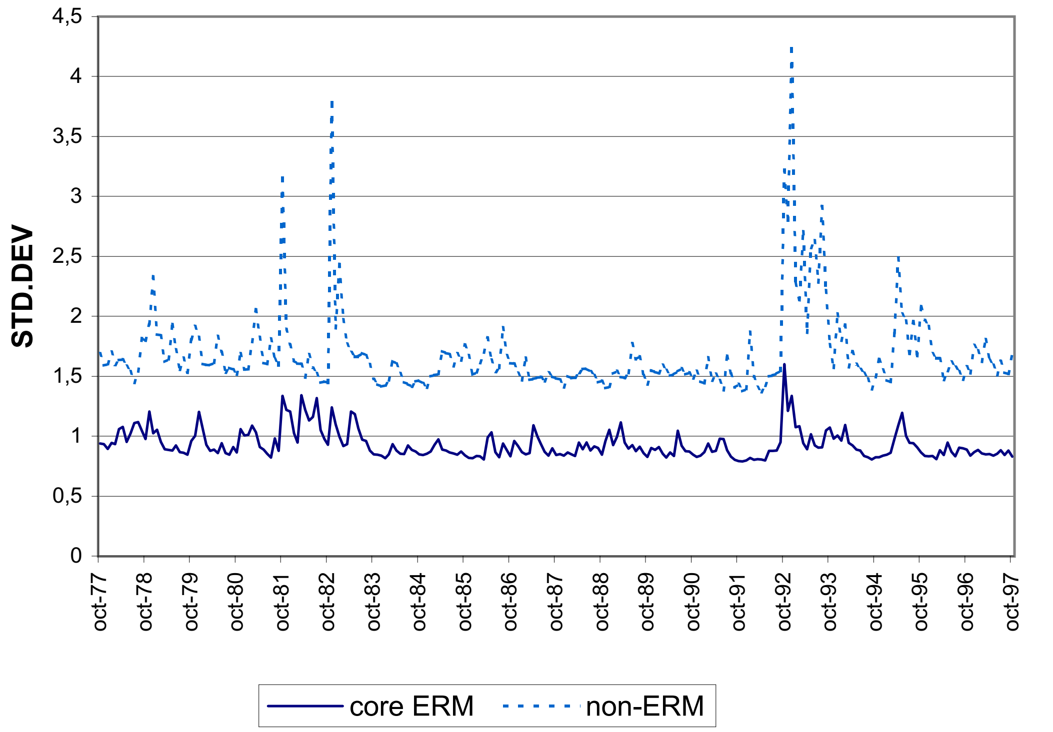

Figure 1b: Average conditional standard deviation of idiosyncratic interest rate factors

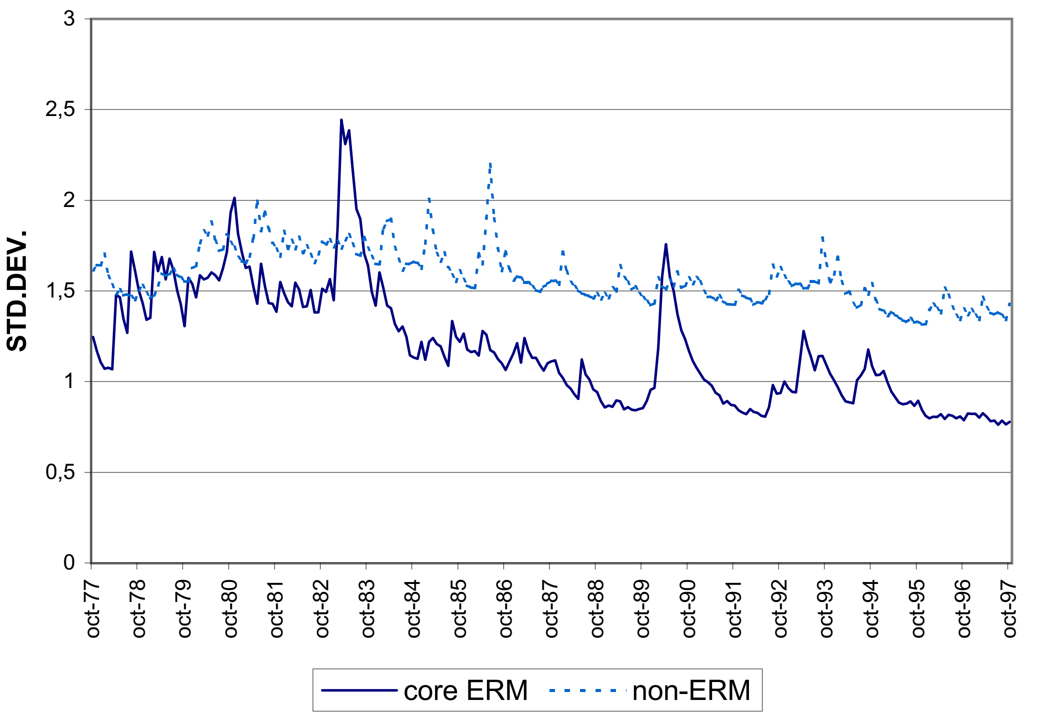

Figure 2a: Net effect of idiosyncratic exchange rate volatility on currency returns

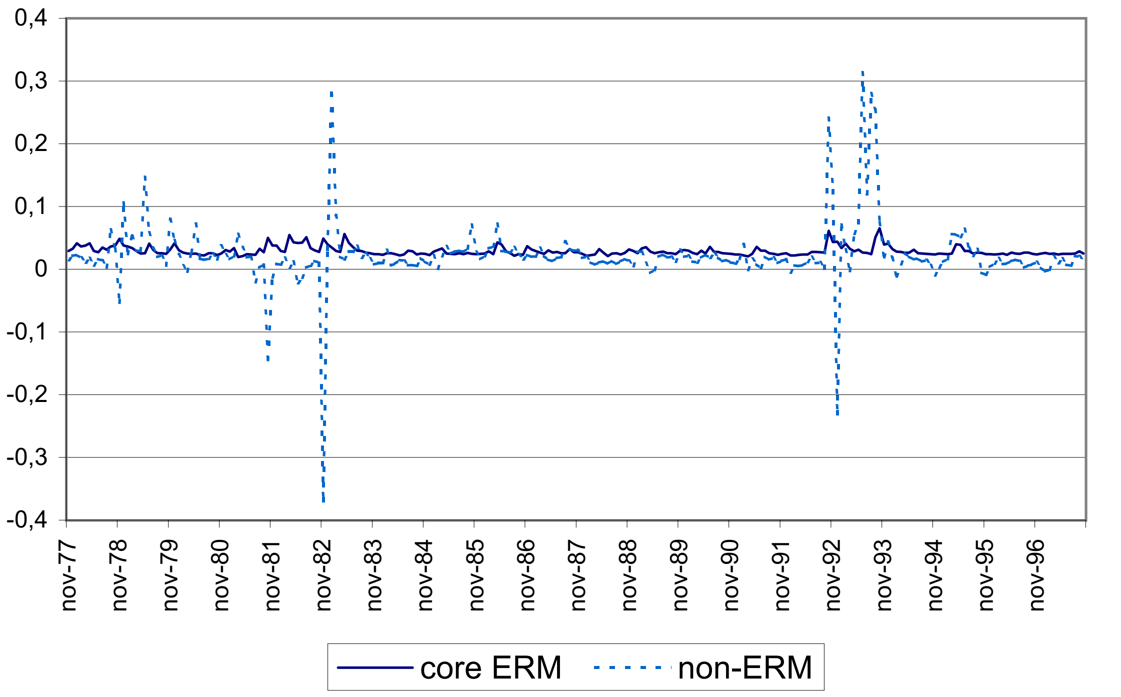

Figure 2b: Net effect of idiosyncratic exchange rate and interest rate volatility on bond returns

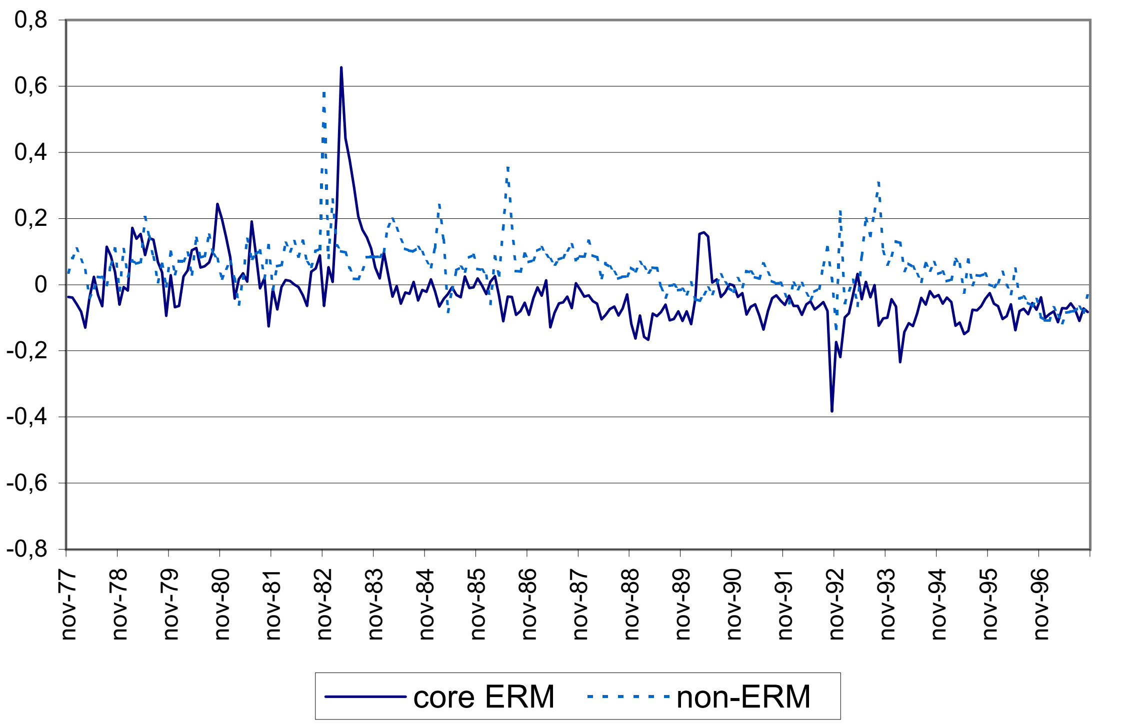

Figure 2c: Net effect of idiosyncratic exchange rate and interest rate volatility on stock returns