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ON THE DYNAMIC SPECIFICATION OF INTERNATIONAL ASSET PRICING MODELS*

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Résumé / Abstract

Dans ce papier, nous testons le modèle CAPM conditionnel international de Dumas et Solnik, l'APT conditionnel international de Ferson et Harvey, ainsi que plusieurs extensions de ces modèles. Ceux-ci ont habituellement été estimés par la méthode des moments généralisés et un test « J » standard n'a souvent pas permis de rejeter les spécifications retenues. Cependant, étant donnée la faible puissance de ces tests contre certaines alternatives locales, nous proposons d'autres tests de diagnostique pour approfondir l'examun empirique de ces modèles. Nous montrons que même si ces derniers n'ont pas été rejeté par le test « J », ils ne sont pas très utiles pour prévoir les premier et second moments des rendements des actions et du taux de change. Notre recherche nous mène à une spécification alternative pour modéliser le rendement des actifs internationaux qui est la formulation ARCH à facteurs. Pour cette dernière, nous trouvons beaucoup de support empirique, à la fois avec le test « J » et avec un certain nombre d'autres tests de diagnostique comme un test d'orthogonalité des résidus ou un examun systématique des erreurs sur les prix.

In this paper, we test the international conditional CAPM model of Dumas and Solnik (1993) and the international conditional APT model of Ferson and Harvey (1992), as well as various extensions of these models. These models were typically estimated by GMM and found to be valid according to the standard J-test. Given to the low power of J-tests against many specific alternatives, we propose several diagnostics to further scrutinize the empirical fit of these models. We show that although they could not be rejected on the basis of the overidentifying restrictions test, they are not very useful for consistently predicting the conditional first and second moments of equity and foreign exchange returns over time. Our specification search leads us to an alternative international conditional CAPM model with a factor ARCH formulation for modelling international returns for which we find strong support, both with the J-statistic criterion, as well as a number of other diagnostics tests, including tests for parameter stability, orthogonality of residuals and explicit analysis of pricing errors.

Keywords : generalized method of moments, diagnostc tests, international conditional CAPM, international conditional APT, factor ARCH.

Mots clé : méthode des moments généralisée, tests de diagnostique, CAPM conditionnel international, APT conditionnel international, ARCH à facteurs.

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1. Introduction

Evidence of returns predictability in both equity and foreign exchange markets has led researchers to model risk premia as time-varying and a number of conditional versions of the CAPM and APT models have been tested with some success. In the international finance literature, early unconditional international asset pricing models, such as Stulz (1981) and Adler and Dumas (1983), were re-formulated and tested in their conditional form. This approach was similar to that adopted in the field of domestic finance models. Amongst these studies, Harvey (1991), Ferson and Harvey (1992), Dumas and Solnik (1993), Ferson and Korajzcyk (1992) and others, found evidence in favor of time-varying risk premia in both equity and foreign exchange markets.

Mainly, in all these models conditional moments at time t are modelled as linear projections on various economic or financial instruments whose values are known at time t-1. The estimation technique used is the generalized method of moments procedure and the main criterion used to assess the goodness of fit of the model is the J-test for overidentifying restrictions(see Hansen (1982)).

There also exist other conditional models of asset pricing which follow more of the ARCH tradition. That is, conditional moments are specified as projections only on lags of squared returns. This formulation stresses the importance of taking into account the observed heteroskedasticity in the volatilities and covariances of returns. The estimation method used is either maximum likelihood or, again, generalized method of moments. An example of a paper in this class of models using GMM is the conditional domestic CAPM of Bodurtha and Mark (1991).

To allow for time-varying risk premia certainly yields more sophisticated asset pricing models, but the search for adequate model specifications is obviously more delicate. In particular, the dynamics of predictable returns needs to be scrutinized seriously as misspecification could be costly in terms of pricing error. In fact, to our knowledge, there has been no systematic attempt to investigate whether either of the two, the instrument method or the ARCH method, performs better than the other in predicting expected returns. Another important aspect to realize about these models is that the J-test which is frequently used to assess the overall validity of the formulation chosen has low power against particular local misspecification alternatives and is therefore not well suited for uncovering systematic mispricing due to specification errors in the model. Studies by Newey (1985) and Ghysels and Hall (1990a,b) provide examples of such situations where the test has low power.

We begin with a rigorous examination of both the CAPM model of Dumas and Solnik (1993) and the APT model of Ferson and Harvey (1992), as well as various extensions of these models. Using our own data set, we first re-estimate these models and subject them to diagnostics tests on the parameters and on the residuals of the models.

Essentially, the test on the parameters is the Andrews (1993) test for structural change with unknown change point. We show that for most of the conditional international asset pricing models that we estimated, we reject parameter stability over time. This means that even though these models cannot be rejected on the basis of the overidentifying restrictions J-test, they are not very useful for consistently predicting the conditional first and second moments of equity and foreign exchange returns over time because of parameter variation which is left unspecified.

The test on residuals, on the other hand, checks whether the model residuals are orthogonal to certain specific alternatives. We use this to analyze the orthogonality of the residuals from the model that uses one approach, say the instrument approach, to the informational content present in variables typically used in the other approach (in this case, the ARCH approach). We find that the Dumas and Solnik (1993) models are generally not rejected against the alternatives tested as opposed to the Ferson and Harvey (1992) models where we find that unexploited information remains in variables not used in the estimation; these variables being the ones typically used in the ARCH approach.

Given this evidence for the presence of pertinent information in the autoregressive elements of the model, we next propose a factor ARCH specification to explain international returns. With respect to all three tests described above, that is, the J-statistic criterion, the stability test, as well as, to some extent, the orthogonality test, our model is shown to perform well. Finally, we conclude our analysis by examining and comparing estimated pricing errors.

The paper is divided in the following manner. Section 2 exposes the Dumas and Solnik(1993) and the Ferson and Harvey (1992) models in some detail. Section 3 includes a discussion on the validity of the J-statistic test for structural change in the above models, an explanation of the Andrews (1993) Sup LM test, some remarks on the asymptotic local power of GMM tests and an exposition of the form of optimal tests against specific alternatives. In section 4, we describe our data set and report estimation and diagnostics test results for the conditional CAPM and APT models. Section 5 exposes our model and includes estimations and test results. Section 6 covers the discussion on pricing errors. The last section concludes.

2. International Conditional Asset Pricing Models

What differentiates international financial theory from its domestic counterpart is essentially the presence of different nations in the former framework. Thus much depends on the definition of the concept of nation adopted. A most useful definition is one attributed initially to Solnik (1974) and where a nation is referred to as a zone of common purchasing power unit. This means that individuals of one zone use a different price index to deflate their monetary investment earnings than those in another zone. Naturally, were the hypothesis of purchasing power parity (PPP) to hold at all times, this distinction would not arise. But it is now well documented by numerous empirical studies that PPP holds at best in the long run, if at all. Furthermore, deviations from PPP are shown to be significant, of long duration and highly random.

An important such international asset pricing model is that of Adler and Dumas (1983). This is a theoretical intertemporal model of utility maximization where investors' real returns vary due to the existence of various nations, that is, the presence of various zones of common purchasing power unit. It is also assumed that the N risky security prices of the model, as well as the price index of each country, follow stationary Ito processes and that preferences are homothetic. The resulting equilibrium condition is therefore an international CAPM which relates expected nominal returns of each asset to its covariance with inflation for each country and its covariance with the market return. In addition, if one makes the assumption that, in the short run, inflation risk expressed in local currency is negligible, the covariance terms with inflation can be replaced by covariances of nominal returns with exchange rate changes. This is the starting point for the conditional model found in Dumas and Solnik (1993) which is given by

$$E[r_{jt}|\Omega_{t-1}] = \sum_{i=1}^{L} \lambda_{i,t-1}\text{cov}[r_{jt}, r_{n+i,t}|\Omega_{t-1}] + \lambda_{m,t-1}\text{cov}[r_{jt}, r_{mt}|\Omega_{t-1}] \tag{2.1}$$

The total number of assets is $m = n + L + 1$ that is, n equity portfolios, L currency deposits (other than the measurement currency), and a world portfolio. Furthermore, r_{jt} is the nominal return on asset $j, (j = 1, \ldots, m)$ in excess of the risk-free rate of the measurement currency country, $r_{n+i,t}$ is the excess return on the ith currency deposit ($i = 1, \ldots, L$), and r_{mt} is the excess nominal return on the world portfolio. Thus, both exchange rate risk and market risk are conditionally priced in this model. These prices are the time-varying coefficients, $\lambda_{k,t-1}, k = i, m$.

In order to write this model in a more parsimonious way, Dumas and Solnik estimate a restricted form of equation (1). Since the first-order condition of any portfolio choice problem can be written as

$$E[M_t r_{jt} | \Omega_{t-1}] = 0 \tag{2.2}$$

where M_t is the intertemporal marginal rate of substitution of returns, they define u_t as the unanticipated component of the relative intertemporal marginal rate of substitution and write it as

$$u_t = 1 - \frac{M_t}{E[M_t | \Omega_{t-1}]} \tag{2.3}$$

with the property that

$$E[u_t | \Omega_{t-1}] = 0. \tag{2.4}$$

Then, equation (2) implies the condition that

$$E[r_{jt} | \Omega_{t-1}] = E[r_{jt} u_t | \Omega_{t-1}], \quad j = 1, 2, \ldots, m. \tag{2.5}$$

Substituting $E[r_{jt} | \Omega_{t-1}]$ for its expression in (1), and given (4), we have the expression for u_t given by:

$$u_t = -\left[\sum_{i=1}^{L} \lambda_{i,t-1} E\left(r_{n+i,t} | \Omega_{t-1}\right) + \lambda_{m,t-1} E\left(r_{mt} | \Omega_{t-1}\right)\right] + \sum_{i=1}^{L} \lambda_{i,t-1} r_{n+i,t} + \lambda_{m,t-1} r_{mt} \tag{2.6}$$

At this point, the model still has a high number of parameters and is quite nonlinear. Empirical estimation is therefore still cumbersome. To simplify further, Dumas and Solnik further restrict equation (6) to give

$$u_t = \lambda_{0,t-1} + \sum_{i=1}^{L} \lambda_{i,t-1} r_{n+i,t} + \lambda_{m,t-1} r_{mt} \tag{2.7}$$

and estimate the model represented by equations (4), (5) and (7). The advantage of this particular formulation is that one does not need to explicitly specify the expected nominal returns. However, this same advantage can be perceived as a drawback. Given that the purpose is to test the conditional CAPM, it should be interesting to discriminate between the informational content contributed by each component of the model to the predictability of returns. With a linear projection, as in equation (7), it is no longer possible to make explicit and test the specification of the expected nominal returns separately from the market prices. In addition, much of the nonlinearity of the model is reduced which could prove to be critical. We will be returning to this point later on.

Since market prices are time-varying, Dumas and Solnik make them linearly dependent upon a number of information instruments which are known at time $t-1$. These are mainly US equity market instruments and were found to have good informational content for the predictability of nominal returns(expressed in dollars) of many countries in the study by Harvey (1991). This set of instruments is denoted Z_{t-1} and is assumed to contain all relevant past information. Therefore:

$$\begin{aligned} \lambda_{0,t-1} &= -Z_{t-1}\delta \\ \lambda_{i,t-1} &= Z_{t-1}\phi_i \\ \lambda_{m,t-1} &= Z_{t-1}\phi_m \end{aligned} \tag{2.8}$$

The model is estimated by GMM for the period March 1970 till December 1991 using monthly data with the US dollar as the measurement currency. The countries considered are Germany, Japan, United Kingdom and the United States. The equity assets are the country equity indexes, the currency deposits are DM, Yen and Pound deposits, and the exchange rates are the bilateral spot exchange rates of each country with the US. The instruments include a constant, the lagged excess world return, the 30-day return on a Eurodollar deposit, the difference in yield between Moody's BAA and AAA rated bonds, the excess dividend yield on the S&P500 index, and a dummy for the month of January. These are the same instruments used in Harvey (1991) except for the 30-day Eurodollar rate which replaces the excess yield on the 90-day US T-bill.

Estimation results show that exchange risk premia are significant and time-varying with several of the ϕ and δ coefficients having high t-statistics. Also, based on the chi-squared test of the overidentifying restrictions provided by the moment conditions of the model, the authors conclude that their parsimonious representation is a satisfactory description of the international CAPM.

By relaxing the constraints imposed by Dumas and Solnik, one obtains a model similar to Ferson and Harvey's (1992) model, which is an international APT model where national equity markets are related to global risk factors. In Ferson and Harvey (1992), the expected risk premia are conditional and are linearly related to the variables constituting the information set Ω at time $t-1$ whereas the conditional betas measure the sensitivity to the global risk factors and are linearly dependent on local information variables. Although the paper reports results for estimations carried out with various numbers of factors, we will concentrate only on their 2-factor model. In this case, the world market portfolio and an aggregate of exchange rates are taken as the two factors underlying the behavior of assets based upon the theoretical justification provided by Adler and Dumas (1983). The model is given by:

$$E[r_{jt}|\Omega_{t-1}] = \sum_{k=1}^{K} \beta_{jk}(\Omega_{t-1}) E[f_{kt}|\Omega_{t-1}], \tag{2.9}$$

where f_{kt} designates a factor, $E[f_{kt}|\Omega_{t-1}]$ is the expected excess return of that factor, and $\beta_{jk}(\Omega_{t-1})$ are the conditional betas of the expected returns. Information available to investors at time $t-1$ is in Ω_{t-1} and includes global information variables as well as local ones.

Estimation of the above model is carried out for eighteen countries using monthly data extending from 1970 to 1989. Returns are in excess of the 30-day T-bill rate. The study concludes that the addition of the second factor, that is the exchange rates aggregate, shows a modest improvement over the single-factor alternative and that most of the predictability in expected returns is related to global risk premia.

Since the data extends over three different currency regimes; a fixed exchange rate period from 1970:02 till 1973:02, a dirty float from 1973:03 till 1980:12, and a more flexible float period afterwards, the authors regress pricing errors for each country on dummy variables representing each of these periods for the single-factor case and the five-factor case. They find no important misspecification related to currency regimes. Nevertheless, for both the Dumas and Solnik and Ferson and Harvey specifications, we will show that stability of coefficients proves to be an issue even if the models are judged acceptable according to the usual overidentification restrictions tests. In the next section, we present tests for structural stability and for misspecification against selected alternatives. These are applied in the sections after, in addition to the usual tests, to assess the validity of various models.

3. Diagnostics beyond overidentifying restrictions

Usually, to assess the goodness of fit of an asset pricing model that is estimated by GMM, Hansen's overidentifying restrictions test is carried out. While the test is an overall diagnostic, it is not an omnibus test against misspecification. This is because it is constructed as a test against general local misspecification alternatives and therefore has low power against some particular forms of misspecification.

In fact, Newey (1985) shows that the overidentifying restrictions test, or the J-test, which is distributed as non-central chi-squared variables under a sequence of local alternatives, has zero non-centrality parameter for some non-zero misspecification directions. In another study, Ghysels and Hall (1990b) formally show that the J-test has no power against local alternatives characterized by time-varying parameters, which means that such tests are not well suited for examining dynamic specification errors such as those related to parameter variation through time. In fact, the authors show that while the parameters that appear in the moment conditions are assumed fixed in estimations, the corresponding restrictions are not imposed in the above test. That is, if the parameters appearing in these moment restrictions are truly time-varying, but they are estimated imposing fixed coefficients instead, the overidentifying restrictions test will tend not to reject the model.

We can therefore see that in order to have a better idea about the validity of these models, one should proceed by examining explicitly both the stability of model parameters, and the soundness of the chosen specification against particular alternatives which are judged to be of possible pertinence for the model in question.

In what follows, we will explain two such explicit tests. These are the tests we will consequently use, in addition to the J-statistic, to undertake the model diagnostics described above on various asset pricing models. We also include a section discussing model pricing errors. In fact, once the models are estimated, pricing errors are easy to calculate. An examination of these is of course valuable as well for establishing overall model validity.

3.1. Testing the Parameters: a test for structural change

To test for structural change in the context of GMM, one must test the null hypothesis of constant parameters explicitly. Let us denote the parameter set as γ_t and let the alternative of a one-time change in the value of γ_t occur at the time πT, where T is the sample size and $\pi \in (0,1)$. The parameter vector γ_t either contains all coefficients of the model or a subset. The latter case is referred to as a "partial" test of structural change since only a subset of parameters are tested.

The null and alternative of the test are then formulated as

$$H_0 : \gamma_t = \gamma_0$$

$$H_1 : \gamma(\pi) : \begin{cases} \gamma_1(\pi)\,, & \text{for } t = 1, 2, \ldots, \pi T \\ \gamma_2(\pi)\,, & \text{for } t = \pi T + 1, \ldots, T \end{cases}$$

From Andrews and Fair (1988), for the case where π is known, it is possible to construct Wald, LM, or LR-like statistics to test H_0 against H_1. However, when π is unknown, or known to belong to a subset Π of $(0, 1)$, one can calculate test statistics, for instance, based on the supremum of $W_T(\pi)$, $LM_T(\pi)$ or $LR_T(\pi)$, for $\pi \in \Pi$ as proposed by Andrews (1993). Amongst these test statistics, we will concentrate only on the Sup LM test as it only requires estimation of the model under the null, which considerably reduces computation costs. This test statistic is a quadratic form based on the score function obtained from the minimization of the GMM criterion function evaluated at a given restricted estimator. The quadratic form is assigned a weight matrix such that the statistic has a chi-squared distribution under the null for each fixed π. The statistic is therefore calculated for each π and the maximal value is designated as the Sup LM test.[1]

This Sup LM test has been used by Ghysels (1994) to test the stability of the model coefficients of the conditional CAPM of Harvey (1991), the conditional APT of Ferson and Korajczyk (1992), and the nonlinear APT of Bansal and Viswanathan (1993). The results showed, among other things, that most of the linear projection parameters of these models were unstable.

3.2. Testing the Residuals: orthogonality tests

To test for the orthogonality of the model residuals to particular information variables, we will use an optimal GMM test, as described in Newey (1985) and Tauchen (1985). Here, optimality refers to the fact that the value of the non-centrality parameter of the test statistic under the local alternative is maximal for all misspecification directions, and that the test has the smallest possible degrees of freedom amongst those with the aforementioned property. It should be noted that the J-test has a non-centrality parameter which is the largest amongst GMM tests with r-q degrees of freedom (where r is the number of orthogonality conditions and q is the dimension of the parameter vector to be estimated) and

[1] For the full definition of the statistic and more details, we refer the reader to Andrews (1993).

for all misspecification directions. However, Newey shows that when particular alternatives are chosen, one can find test statistics which have, in addition to the largest possible value of the non-centrality parameter, degrees of freedom smaller than r-q. Such tests have therefore more power than the J-test.

Since asset pricing models choose only certain information variables from a set of such variables to assess the predictability of expected returns, it is always interesting to check whether those instruments not used in the estimation could have increased the explanatory power of the model had they been used. In the context of a model estimated via GMM, this amounts to testing whether other orthogonality conditions could have been used in the estimation. The test is described as follows:

Given the moment conditions $f_{1t}(b)$ for $t = 1, 2, .., T$, at the true parameter value b_0 we have that $E[f_{1t}(b_0)] = 0$ and b is then a consistent estimator of b_0. Using the sample moments we can define the terms

$$g_{1T}(b) = (1/T)\sum_{t=1}^{T} f_{1t}(b) \quad \text{and} \quad S_{11,T} = (1/T)\sum_{t=1}^{T} f_{1t}(b) f_{1t}(b)' \tag{3.1}$$

then the GMM estimator b_T is the parameter set that minimizes the quadratic form

$$\Phi(b_T) = g_{1T}(b)' \left[S_{11,T}\right]^{-1} g_{1T}(b) \tag{3.2}$$

Now, if $f_{2t}(b)$ is an $(l\text{x}1)$ vector of orthogonality conditions which were not used in the estimation and which we suspect should have been included in the model, an optimal GMM statistic can be formulated to test for the omission of these moments. This statistic is given by

$$CS = T[L_T g_T(b_T)]' \left[Q_T\right]^{-1} \left[L_T g_T(b_T)\right] \tag{3.3}$$

and is distributed as a χ^2 with l degrees of freedom. The matrices which define this test are detailed as follows[2]

$$g_T(b) = [g_{1T}(b)' \;\; g_{2T}(b)']', \;\; \text{with} \;\; g_{2T}(b) = (1/T)\sum_{t=1}^{T} f_{2t}(b) \tag{3.4}$$

and $L_T = [0 : I_l]$, so that $L_T g_T(b_T)$ is a linear combination of the estimated sample moments $g_T(b_T)$.

[2]For a more detailed exposition of the test, we refer the reader to Newey (1985) and Bodurtha and Mark (1991)

In addition,

$$S_{ij,T} = (1/T)\sum_{t=1}^{T} f_{it}(b) f_{jt}(b)', \quad i = 1,2, \quad j = 1,2 \tag{3.5}$$

$$H_{iT} = (1/T)\sum_{t=1}^{T} \partial f_{it}(b_T)/\partial b, \quad i = 1,2 \quad \text{and} \quad B_T = \left[H'_{1T}\left[S_{11,T}\right]^{-1} H_{1T}\right]^{-1} H'_{2T} \tag{3.6}$$

Finally,

$$Q_T = S_{22,T} - S_{21,T}\left[S_{11,T}\right]^{-1} H_{1,T} B_T - B_T\prime H_{1,T}\prime \left[S_{11,T}\right]^{-1} S_{12,T} + H_{2T} B_T \tag{3.7}$$

Thus, for a given significance level, a large value of the statistic means rejection of the null in favor of the specified alternative moments.

3.3. Comparing Ex-Post Pricing Errors

A fairly easy and natural way of assessing overall model validity is by examining ex-post pricing errors of various models. A Tx1 vector of pricing errors, e_j, for returns on an asset j is defined as the difference between the observed excess returns data for that asset and the returns predicted by the estimated model. For a single observation this is given by:

$$e_{jt} = r_{jt} - E[r_{jt}|\Omega_{t-1}] \tag{3.8}$$

where r_{jt} is the excess return on asset j at time t, and $E[r_{jt}|\Omega_{t-1}]$ is the estimated excess return at time t implied by the model.

There are two typical measures which are useful for evaluating such errors. These are the absolute mean error and the root mean squared error. For an asset j, the absolute mean error is expressed as:

$$AME_j = (1/T)\sum_{t=1}^{T} |e_{jt}|, \tag{3.9}$$

and the root mean squared error is given by:

$$RMSE_j = \sqrt{1/(T-1)\sum_{t=1}^{T}(e_{jt} - ME_j)^2} \tag{3.10}$$

where

$$ME_j = (1/T)\sum_{t=1}^{T} e_{jt} \tag{3.11}$$

By comparing these measures across models, in addition to their autocorrelations, one can determine the accuracy of a model's in-sample performance relative to the other.

4. Data, Estimations and Test Results

In this section, we test the Dumas and Solnik (1993) and the Ferson and Harvey (1992) models after re-estimating them with our own data set. Where feasible, the data and instruments were kept as similar to the original model as possible to allow for meaningful comparisons. A subsection is devoted to each model.

4.1. The Dumas and Solnik Model

First, we estimate and test the Dumas and Solnik (1993) model. The data is basically the same if only for a longer holding period; 30-day holding period data being unavailable to us, we work with data on 90-day investment periods. Having said this, our data is therefore monthly, measured in US dollars, for Germany, Japan, the UK and the US, spanning the period of September 1978 till February 1994. Equity returns are constructed using MSCI country indexes (with dividend reinvestment) and returns on the foreign exchange market are calculated using returns on a currency deposit compounded by the exchange rate variation relative to the US dollar. The instruments used are a constant, the lagged excess world return, a dummy for the month of january, the excess US junk bond spread, the excess US dividend yield and the 30-day return on a Eurodollar deposit. Excess returns are taken with respect to the 90-day US T-bill rate.

The estimations and testing are carried out using GMM and the Sup LM tests. It should be noted that with respect to the Sup LM test, 'full' testing indicates that all the model parameters are tested as opposed to 'partial' testing where only selected parameters are tested.[3] The econometric specification for the 4-country model of Dumas and Solnik (1993) that we estimate is given as follows:

$$u_t = \lambda_{0,t-1} + \sum_{i=1}^{3} \lambda_{i,t-1} r_{n+i,t} + \lambda_{m,t-1} r_{mt}$$

[3]The interval adopted for the LM tests here and in the rest of the paper is $(0.2, 0.8)$. Significance of the test in the tables is reported as: * at 10%, ** at 5%, *** at 1%. Note that for parameter numbers greater than 20, asymptotic critical values for Sup LM test were obtained by extrapolation. Nevertheless, these are probably good approximations since the increase in the critical values with the number of parameters is quite linear.

with

$$
\begin{aligned}
\lambda_{0,t-1} &= -Z_{t-1}\delta \\
\lambda_{i,t-1} &= Z_{t-1}\phi_i \\
\lambda_{m,t-1} &= Z_{t-1}\phi_m
\end{aligned}
$$

and where equations (4) and (5) are conditioned relative to Z_{t-1}.

Defining h_{jt} as the unanticipated error term from equation (5), that is $h_{jt} = r_{jt} - (r_{jt}u_t)$, we obtain 4 residuals for each of the 4 countries' equity portfolios, 3 residuals for the currency deposits (DM, Yen and Pound deposits), and a residual for the world portfolio. These, in addition to u_t, constitute the residual set to which Z_{t-1} is orthogonal, and together form the orthogonality conditions which are then exploited in the GMM estimation. Therefore, this model has a total of 30 parameters and 54 orthogonality conditions. This means there are 24 degrees of freedom for the J-statistic.

Estimation results show a value of 20.75 for this statistic with a P-value of 65%, compared with 28.80 and a P-value of 23% obtained by Dumas and Solnik. According to this criterion, our data yields an even better fit than the original study and the model cannot be rejected. However, once explicit stability tests are carried out on the model and the Sup LM test applied, the hypothesis of stable coefficients is rejected at the 5% level, both when all the parameters of the model are jointly tested, and also when only the six parameters of λ_0 are tested. Table 1 summarizes these results.

At this point, it is interesting to see if this rejection is due to the specific choice of instruments above given that the Ghysels (1994) study found the parameters of certain instruments to be specially unstable regardless of the model specification tested. Accordingly, we divide the instrument set into subsets of five instruments each: Instrument set A includes a constant, the lagged excess world return, a January dummy, the excess junk spread, and the excess dividend yield. Set B contains the first four instruments of set A as well as the short term rate. Finally, set C contains the first three instruments of set A, the excess dividend yield and the short term rate. We then estimate the 4-country model with each of these sets and test again for stability. For each of these cases, total parameter count is 25 and the number of orthogonality conditions is 45, implying 20 degrees of freedom for the J-statistic. Also, since the number of instruments is now five, this is also the number of parameters in λ_0.

The results are reported in Table 2. They indicate that regardless of the set of instruments employed, and despite high P-values of the J-statistic, parameter stability is strongly rejected when the full parameter vector is tested. It therefore does not seem as though the particular instrument combination used is the cause of model rejection.

We also investigate whether the rejection is caused by the behavior of excess returns of one of the countries in particular or if it is a phenomenon general to all countries. This is carried out by estimating and testing the stability of 2-country versions of the above model (that is, with the US and another country) with the three instrument sets A, B, and C. The results are found in Table 3 and indicate that except for the UK in the case of set A and Japan in the case of set B, parameter stability is rejected in the remaining seven cases.

From Tables 1-3 we notice that partial parameter stability is also majoritarily rejected at the 5% level. It is interesting to see whether the particular linear constraint imposed by Dumas and Solnik is the cause of this, given that it does not allow the manifestation of the nonlinear dynamics in returns which is implied by the main model found in equation (1). We therefore replace equation (7) with equation (6) and test again. For this purpose, however, we need to impose a specific formulation for the expressions of the conditional expected returns and we choose to model these as linear projections on the instruments. The projection equations for $\lambda_{i,t-1}$ and $\lambda_{m,t-1}$ remain unchanged. Again, we use the three instrument sets A, B, and C defined above.

The econometric model estimated in this case is therefore

$$u_t = -\left[\sum_{i=1}^{L} \lambda_{i,t-1} E\left(r_{n+i,t}|Z_{t-1}\right) + \lambda_{m,t-1} E\left(r_{mt}|Z_{t-1}\right)\right] + \sum_{i=1}^{L} \lambda_{i,t-1} r_{n+i,t} + \lambda_{m,t-1} r_{mt}$$

with

$$\begin{aligned} E\left(r_{n+i,t}|Z_{t-1}\right) &= Z_{t-1}\delta_i \\ E\left(r_{mt}|Z_{t-1}\right) &= Z_{t-1}\delta_m \\ \lambda_{i,t-1} &= Z_{t-1}\phi_i \\ \lambda_{m,t-1} &= Z_{t-1}\phi_m \end{aligned}$$

along with equations (4) and (5), again conditional to Z_{t-1}.

Here, the total number of parameters is 40 and a full LM test is undertaken. The results are tabulated in Table 4. Once again, P-values for the overidentification restrictions test are very high and the model would not have been rejected on the basis of this statistic alone. Nevertheless, the stability test overwhelmingly rejects invariance of coefficients over time indicating that the added nonlinearity is insufficient to improve the model specification. For the sake of comparison, we also estimated and tested the 2-country versions of the above model (see Table 5). The conclusions remain unchanged.

Next, we go back to the constrained Dumas and Solnik model (equations 4,5 & 7) and examine the orthogonality conditions for each country pair. We test whether each of the residuals of the model is orthogonal to various lags of the model returns. These lags are grouped into subsets and are designated as alternative instrument sets $AZ_{t-1}^1, AZ_{t-1}^2, AZ_{t-1}^3$, and AZ_{t-1}^4. The first one includes lags 1, 2 and 3 of the excess equity returns of the country paired with the US, the second includes lags 1, 2 and 3 of the excess returns to a deposit made in the currency of the country other than the US, the third set is comprised of lags 1, 2 and 3 of the excess equity returns of the US, and the last includes lags 2 and 3 of the world excess returns as well as lag 1 of the US excess equity returns. Orthogonality tests against AZ_{t-1}^1, AZ_{t-1}^2 and AZ_{t-1}^3 revealed very high P-values in all cases and for all the residuals, varying from 0.95 to 1.00 and implying that the model could not be rejected against these alternatives. Nevertheless, for the last subset, AZ_{t-1}^4, low P-values for a few of the residuals indicated that there might be some misspecification present. The results against this last alternative instrument set are tabulated in Table 8 for all three country pairs. They show that only for the Germany-US and Japan-US pairs is the issue of misspecification a concern. Despite this fact, in general, one can conclude that the Dumas and Solnik model fares well against all four alternative instrument subsets.

4.2. The Ferson and Harvey Model

We now return to equation (1) which is the most general form of this international asset pricing and can therefore be seen as a multi-beta (APT) model if the λ coefficients are interpreted as risk premia (see equation (9)). In fact, the 2-country version of such a model is then closely comparable to the international APT model of Ferson and Harvey (1992) which is estimated by GMM . This is the model examined next.

We estimate the international APT 2-factor 2-country model, with the US as one of the countries and either Germany, Japan or the UK as the other. The two factors are the excess return on the world portfolio and the excess return on the bilateral exchange rate. Instead of this latter factor, Ferson and Harvey use an aggregate of exchange rates, as explained previously. However, we chose to adopt the bilateral exchange rate because it is probably more helpful in capturing the evolution of the various shocks in the economies of the nations involved. This is specially valid in a two-country case model where one country's instruments predominate in the information set, in this case, the US's.

We start by estimating a constant beta version of the APT with all time-variation captured by the conditional risk premia. This is based on the conclusion of the Ferson and Harvey study proper, that most of the dynamics is found in the risk premia rather than in the betas.

The Ferson and Harvey(1993) econometric model specification is described as

$$\varepsilon_t = (u_{1t}\; u_{2t}\; u_{3t}) = \begin{pmatrix} (r_t - Z_{t-1}\delta)' \\ (f_t - Z_{t-1}\gamma)' \\ (u_{2t}u'_{2t}\beta - f_t u_{1t})' \end{pmatrix}', \tag{4.1}$$

$$E(\varepsilon_t | Z_{t-1}) = 0, \tag{4.2}$$

where f_t is the 2×1 vector of factors, r_t is the excess equity return of the country considered other than the US, and Z_{t-1} is a 5-instrument information set (either A,B or C defined previously). The residuals u_{2t} and u_{3t} are both of dimension 1×2 as they include a term for each of the two factors in the model.

Table 6 contains the results of GMM estimations and the various LM tests for each of the 2-country multi-beta models. The number of parameters, in each case, is as follows: total 17, delta 5, gamma 10, beta 2. The instrument sets are successively A, B, and C. We can see that the P-values for these models are high varying from 39 to 99 percent and that for Japan and the UK, the model is not rejected at the 5% level in 2 cases. This provides us with valuable information in selecting the desirable instrument set for each country pair in the context of these types of models. Nevertheless, we also note that, for the rejected cases, much of the instability is coming from the γ coefficients which are the linear projections of the factors on the instruments chosen.

Orthogonality tests are also carried out for each of the model residuals with respect to the same alternatives as described in the previousl section. Tables 9, 10 and 11 report P-values for the three country pairs for the tests that were carried out against the first three of the four alternative subsets. In them there is some indication of misspecification as all three tables include results which are insignificant at the 5 % level. More specifically, this misspecification seems to be concentrated around the residual from the projection equation for the exchange rate factor, specially with the alternative instrument set AZ^2_{t-1}. This indicates that the projection equation for this factor should have included a number of own lags. Furthermore, test results against the fourth set, AZ^4_{t-1}, which are not tabulated, proved to be more dramatic as all P-values for all the residuals and for all the country pairs hovered around values of 10^{-4}. Clearly, this is a strong indication of the presence of misspecification against the alternative represented by lags 2 and 3 of world excess equity returns and lag 1 of excess US equity returns.

In conclusion to this section, we see that although the models were judged satisfactory according to the J-criterion, they are either unstable over time, or contain misspecification, or exibit both symptoms. Having pin-pointed the variables which still contain pertinent information, and undertanding the importance of avoiding instruments which could lead to unstable model coefficients, it seems appropriate to propose a factor-ARCH model which formulates conditional moments as functions of own lagged variables. This is the model presented in section 5.

5. Autoregressive Factor Models

It seems apparent, at this point, that simple linear projection equations with constant betas do not adequately describe the process of conditional excess returns. On the one hand, the instruments used proved to yield unstable parameter values in linear projection equations, and, on the other hand, the models did not take into account certain nonlinear phenomena observed for financial and monetary series, such as persistence in volatility. We propose a factor-ARCH model which takes both these facts into account. Following along the lines of Bodurtha and Mark (1991), our model avoids external instruments altogether. In addition, the autoregressive components in the variance and covariance terms of the model will help capture much of the observed behavior of volatility in financial and monetary series.

Our purely autoregressive factor ARCH model is given by:

$$E[r_{jt}|\Omega_{t-1}] = \sum_{k=1}^{K} \beta_{jk}(\Omega_{t-1}) E[f_{kt}|\Omega_{t-1}], \tag{5.1}$$

where

$$E[f_{kt}|\Omega_{t-1}] = \sum_{l=1}^{4} \alpha_{kl} f_{k,t-l}\,, \qquad f_{kt} = r_{xt},\, r_{mt} \tag{5.2}$$

and

$$\beta_{jk}(\Omega_{t-1}) = \frac{Cov[r_{jt}, f_{kt}|\Omega_{t-1}]}{Var[f_{kt}|\Omega_{t-1}]} \tag{5.3}$$

We adopt an AR(4) for the expected value of the factors because, in a separate regression, this formulation yielded significant coefficients and the residuals did not appear to be correlated. In addition, it is a relatively parsimonious representation.

The remaining econometric specification of the model is given as follows. The unanticipated components of factors, that is, excess foreign exchange returns (r_{xt}) and excess market returns (r_{mt}), are given by:

$$u_{xt} = r_{xt} - E[r_{xt}|Z_{t-1}] \tag{5.4}$$

$$u_{mt} = r_{mt} - E[r_{mt}|Z_{t-1}] \tag{5.5}$$

and that for excess equity returns as:

$$u_{jt} = r_{jt} - E[r_{jt}|Z_{t-1}] \tag{5.6}$$

where Z_{t-1} is the set of information variables available to the investor.

From here, we obtain that

$$\begin{aligned}
\text{Var}[r_{xt}|Z_{t-1}] &= E[u_{xt}^2|Z_{t-1}] \\
\text{Var}[r_{mt}|Z_{t-1}] &= E[u_{mt}^2|Z_{t-1}] \\
\text{Cov}[r_{xt}, r_{jt}|Z_{t-1}] &= E[u_{xt}u_{jt}|Z_{t-1}] \\
\text{Cov}[r_{mt}, r_{jt}|Z_{t-1}] &= E[u_{mt}u_{jt}|Z_{t-1}]
\end{aligned} \tag{5.7}$$

We assume the conditional variance to be an ARCH(1) and the conditional covariance an autoregression of order one. The unanticipated components of these elements are given by

$$\begin{aligned}
\eta_{xt} &= u_{xt}^2 - E[u_{xt}^2|Z_{t-1}] \\
\eta_{mt} &= u_{mt}^2 - E[u_{mt}^2|Z_{t-1}] \\
\eta_{xjt} &= u_{xt}u_{jt} - E[u_{xt}u_{jt}|Z_{t-1}] \\
\eta_{mjt} &= u_{mt}u_{jt} - E[u_{mt}u_{jt}|Z_{t-1}]
\end{aligned} \tag{5.8}$$

therefore we obtain,

$$
\begin{aligned}
E[u_{xt}^2|Z_{t-1}] &= \delta_{x0} + \delta_{x1}(u_{x,t-1}^2) \\
E[u_{mt}^2|Z_{t-1}] &= \delta_{m0} + \delta_{m1}(u_{m,t-1}^2) \\
E[u_{xt}u_{jt}|Z_{t-1}] &= \delta_{xj0} + \delta_{xj1}(u_{x,t-1}u_{j,t-1}) \\
E[u_{mt}u_{jt}|Z_{t-1}] &= \delta_{mj0} + \delta_{mj1}(u_{m,t-1}u_{j,t-1})
\end{aligned}
\tag{5.9}
$$

Substituting the corresponding terms in (15), we obtain the following orthogonality conditions:

$$
E(u_{xt}Z_{t-1}^1, u_{mt}Z_{t-1}^1, \eta_{xt}Z_{t-1}^2, \eta_{mt}Z_{t-1}^2, \eta_{xjt}Z_{t-1}^3, \eta_{mjt}Z_{t-1}^3, u_{jt}Z_{t-1}^3) = 0 \tag{5.10}
$$

where

Z_{t-1}^1 includes a constant, r_{xt} lagged one period,r_{mt} lagged one period,
Z_{t-1}^2 includes a constant, u_{xt}^2 lagged one period, u_{mt}^2 lagged one period,
Z_{t-1}^3 includes a constant, $u_{xt}u_{jt}$ lagged one period, $u_{mt}u_{jt}$ lagged one period.

There are, therefore, 18 parameters and 21 moment conditions implying 3 overidentifying restrictions. Estimation and full parameter LM test results are found in Table 7. For all cases results are very satisfactory both according to the χ^2 statistic and with respect to stability. The χ^2 P-values are at 99% for the UK and Germany and at 58% for Japan. As for stability, none of the models is rejected even at the 10% level.
Next, we apply the CS test to our model to check the validity of our residuals against certain alternatives. This time, the alternative subsets are comprised of the instruments used in the models of Dumas and Solnik (1993) and Ferson and Harvey (1992). The first of these subsets is denoted XSJUNK and includes 3 lags of the excess US junk bond spread, the second, XSDIV, contains 3 lags of the excess US dividend yield, and the last, STRATE, includes 3 lags of the 30-day Eurodollar return.
Results of the CS tests are found in Tables 12, 13 and 14. They indicate that, generally speaking, the model is robust against the external instrument alternatives. Nevertheless, they also reveal that it will be suitable to modify the specification of the covariance terms to include possibly more own lags, since P-values of covariance residual tests turn out to be generally lower than 5 %.

6. Pricing Errors

In this section, we examine pricing errors of the equity portfolios for the Dumas and Solnik constrained model, the Ferson and Harvey model and the factor-ARCH model, in the case of two countries. For the first model, the pricing error is determined from equation (5). As defined in section 4.1, this is given by the variable h_{jt} which we re-define here as e_{jt} and write it as

$$e_{jt} = r_{jt} - (r_{jt} u_t) \tag{6.1}$$

For the Ferson and Harvey model, e_{jt} is given by:

$$e_{jt} = r_{jt} - \sum_{k=1}^{2} \beta_{jk} E[f_{kt} | \Omega_{t-1}] \tag{6.2}$$

and for the factor-ARCH, this is given by

$$e_{jt} = r_{jt} - \sum_{k=1}^{2} \frac{Cov[r_{jt}, f_{kt} | \Omega_{t-1}]}{Var[f_{kt} | \Omega_{t-1}]} \sum_{l=1}^{4} \alpha_{kl} f_{k,t-l} \tag{6.3}$$

as in equation (28). From these we obtain the absolute mean and the root mean square errors which are tabulated in Table 15. From these we can see both the absolute mean error and the root mean square error yield qualitatively similar results. Quantitatively, the Ferson and Harvey and the factor-ARCH values are generally close. Also, for all three country pairs, the Dumas and Solnik model has the highest pricing error statistics. For instance, the absolute mean error varies between 5.38 % per month (this is the case of Germany and the US estimated with instrument set C) and 8.72 % per month (the UK-US country pair estimated with instrument set A). Similarly, the root mean squared error ranges from 8.16 % to 12.72 % p.m. for the same two cases. Amongst the remaining models, the Ferson and Harvey model yields the smallest statistics for the UK-US pair with 4.92 % p.m. for the absolute mean and 6.22 % for the root mean squared error. However, for Germany and the US and Japan and the US, the factor-ARCH model outperforms the others with absolute mean errors of 4.87 % for the first and 5.57 % for the second. To summarize the findings from this table, we have that, in 2 cases out of 3, the factor-ARCH yields the best results. The Ferson and Harvey model performs the best in the remaining case.

We also calculated 12 autocorrelations for each of the e_{jt}. From amongst these we selected lags 1, 2, 3, 6, and 12 which we report in Table 16. Ljung-Box white noise tests were also run on the 12 autocorrelations, the P-values of which we also included in the same table. The results show that, generally, all the autocorrelations are low and that only in two cases is the null hypothesis of white noise autocorrelations rejected at the 5% level. These are the Dumas and Solnik model results for the Germany-US pair, once estimated with instrument set A, and another time, with instrument C.

7. Conclusion

In conclusion, we can say that based on the J-test criterion, we would not have rejected any of the models. What allows us to distinguish the more desirable model is the explicit testing for stability on the one hand, and for orthogonality of errors, on the other. As it turns out, the factor-ARCH formulation seems to be the surest to adopt amongst the three models. It was found to be sound structurally, to hold well versus misspecification against various specific alternatives, and yielded pricing errors which could not be rejected when tested for white noise. In opposition, the Dumas and Solnik model, although more parsimonious, was shown to be structurally unstable, and the Ferson and Harvey model, although relatively more stable, exhibited misspecification against some alternatives.

In choosing an econometric description for the international conditional asset pricing model, one has the option of using the latent approach, leaving first or second moments of returns unspecified, or, one can parameterize these moments more particularly. The decision to use either a linear or a nonlinear formulation should be directed according to observed facts and empirical regularities. Thus, although the first option has the advantage of being parsimonious and is easier to handle from a numerical point of view, it seems important, given the outcome of the tests above, not to ignore the observed persistence in volatility and not to oversimplify by using linear structures. In this respect, the factor-ARCH formulation, with second moments specified nonlinearly, albeit in an ad hoc manner, seems to be the more adequate and viable specification for the purpose described above.

Table 1

Dumas and Solnik model - equations (4), (5) and (7)

GER,JAP,UK,US; six instruments

MODEL	χ^2(P-value)	Deg. Freedom	Full Sup LM Test		Partial Sup LM Test
Our data	20.75(0.65)	24	108.5	***	20.76
D&S data	28.80(0.23)	24	N.A.		N.A.

Significance of tests is reported as: * at 10%, ** at 5%, *** at 1%.

Sup LM critical values, all parameters 1%=65, 5%=58, 10%=55

Sup LM critical values, λ_0 parameters 1%=24.3, 5%=19.6, 10%=17.6

Table 2

Dumas and Solnik model - equations (4), (5) and (7)

GER,JAP,UK,US; five instruments in each set

Instrument Set	χ^2(P-value)	Deg. Freedom	Full Sup LM		Partial Sup LM	
SET A	17.67(0.61)	20	66.4	***	17.7	*
SET B	10.22(0.96)	20	80.9	***	10.2	
SET C	25.62(0.18)	20	106.8	***	25.6	***

Significance of tests is reported as: * at 10%, ** at 5%, *** at 1%.

Sup LM critical values, all parameters 1%=57, 5%=50.5, 10%=47.1

Sup LM critical values, λ_0 parameters 1%=21.9, 5%=17.9, 10%=15.6

Table 3

Dumas and Solnik model - equations (4), (5) and (7)

The US and another country; five instruments in each set

Model	Instrument set	χ^2(P-value)	Deg. Freedom	Full LM		Partial LM	
GER,US	A	9.78(0.46)	10	61.03	***	27.8	***
	B	10.99(0.36)		48.1	***	19.9	**
	C	10.01(0.44)		91.1	***	57.7	***
JAP,US	A	10.94(0.36)	10	59.3	***	19.2	**
	B	6.63(0.76)		37.9	**	24.2	***
	C	10.33(0.41)		49.2	***	19.8	**
UK,US	A	2.07(0.99)	10	25.5		12.8	
	B	8.32(0.60)		41.3	***	20.4	**
	C	7.62(0.67)		111.9	***	70.0	***

SET A contains: a constant, lag 1 of world equity excess returns, a january dummy, excess junk bond spread, excess dividend yield; SET B contains: a constant, lag 1 of world equity excess returns, a january dummy, excess junk bond spread, short term rate; SET C contains: a constant, lag 1 of world equity excess returns, a january dummy, excess dividend yield, short term rate

Significance of tests is reported as: * at 10%, ** at 5%, *** at 1%.

Sup LM critical values, all parameters 1%=40.1, 5%=34.3, 10%=31.7

Sup LM critical values, λ_0 parameters 1%=21.9, 5%=17.9, 10%=15.6

Table 4

Dumas and Solnik model - equations (4), (5) and (6)

GER,JAP,UK,US; five instruments in each set

Instrument Set	χ^2(P-value)	Degrees of Freedom	Full Sup LM Test	
SET A	16.7(0.89)	25	862.6	***
SET B	5.5(0.99)	25	304.4	***
SET C	18.2(0.83)	25	1222.6	***

For the definitions of the instrument sets, see Table 3.

Significance of tests is reported as: * at 10%, ** at 5%, *** at 1%.

Sup LM critical values, all parameters 1%=83.5, 5%=75.2, 10%=71.8

Table 5

Dumas and Solnik model - equations (4), (5) and (6)

The US and another country; five instruments in each set

Model	Instrument set	χ^2(P-value)	Deg. freedom	Full Sup LM Test	
GER, US	A	14.01(0.52)	15	119.2	***
	B	23.08(0.08)		119.2	***
	C	21.79(0.11)		179.2	***
JAP,US	A	11.34(0.73)	15	116.2	***
	B	6.41(0.97)		70.4	***
	C	13.54(0.56)		407.4	***
UK,US	A	17.63(0.28)	15	176.9	***
	B	3.24(0.58)		36.5	
	C	15.41(0.42)		104.7	***

For the definitions of the instrument sets, see Table 3.

Significance of tests is reported as: * at 10%, ** at 5%, *** at 1%.

Sup LM critical values, all parameters 1%=47.8, 5%=41.9, 10%=39.0

Table 6

Ferson and Harvey model - equations (18) and (19)

The US and another country; five instruments in each set; constant betas

Model	Inst. set	χ^2(P-value)	DF	Sup LM							
				Full		δ		γ		β	
GER,US	A	5.04(0.75)	8	128.5	***	24.9	***	121.6	***	65.4	***
	B	7.69(0.43)		40.9	**	9.4		34.4	***	8.0	
	C	2.82(0.95)		85.4	***	33.8	***	80.3	***	29.7	***
JAP,US	A	3.07(0.93)	8	30.3		3.5		15.0		15.1	**
	B	3.17(0.92)		53.0	***	15.2		42.1	***	22.2	***
	C	5.30(0.72)		32.2		10.5		17.9		13.4	**
UK,US	A	3.68(0.88)	8	39.3	**	12.4		36.4	***	7.2	*
	B	8.51(0.39)		31.3		8.3		20.5		3.4	
	C	1.73(0.99)		120.7	***	32.9	***	116.9	***	77.9	**

For the definitions of the instrument sets, see Table 3.

Significance of tests is reported as: * at 10%, ** at 5%, *** at 1%.

Sup LM critical values, all parameters $1\%= 43.3, 5\%= 37.5, 10\%= 34.5$

Sup LM critical values, δ parameters $1\%= 21.9, 5\%= 17.9, 10\%= 15.6$

Sup LM critical values, γ parameters $1\%= 32.0, 5\%= 26.4, 10\%= 24.0$

Sup LM critical values, β parameters $1\%= 15.1, 5\%= 11.3, 10\%= 9.6$

Table 7

Factor ARCH model

The US and another country; Instrument sets are Z^1_{t-1}, Z^2_{t-1}, Z^3_{t-1}

MODEL	χ^2(P-value)	Deg. Freedom	Full Sup LM Test
GER,US	0.01(0.99)	3	34.6
JAP,US	1.94(0.58)	3	31.8
UK,US	0.02(0.99)	3	20.9

Significance of tests is reported as: * at 10%, ** at 5%, *** at 1%.

Sup LM critical values, all parameters 1%= 44.1, 5%= 38.8, 10%= 36.1

Table 8

CS Tests for the Dumas and Solnik Model

Two-Country Cases; Alternative instrument set is AZ^4_{t-1}

RESIDUAL	INSTR. SET	GERMANY-US	JAPAN-US	UK-US
P-value				
u_t	A	0.999	0.000	0.998
h_{jt}		0.999	0.000	1.000
h_{USt}		0.999	0.998	0.996
h_{wt}		0.118	1.000	0.999
hcd_{jt}		0.995	0.999	0.891
u_t	B	0.991	0.999	0.999
h_{jt}		0.991	0.828	1.000
h_{USt}		0.995	1.000	0.998
h_{wt}		0.999	0.992	0.988
hcd_{jt}		0.020	0.997	0.991
u_t	C	0.999	0.977	0.989
h_{jt}		0.951	0.886	0.994
h_{USt}		0.999	0.930	0.998
h_{wt}		0.958	0.999	0.999
hcd_{jt}		0.943	0.990	0.999

For the definitions of the instrument sets, see Table 3. Residual u_t is defined in equation (7), the h error terms are from equation (5): h_{jt} is for equity returns of country j, h_{USt} is for US equity returns, h_{wt} is for world equity returns, and hcd_{jt} is for returns on a deposit in the currency of country j.

Table 9
CS Tests for the Ferson and Harvey Model
Case of the US and Germany

RESIDUAL	INSTR. SET	AZ^1_{t-1}	AZ^2_{t-1}	AZ^3_{t-1}
		P-value		
u_{1t}	A	0.51	0.74	0.72
u_{2xt}		0.17	0.00	0.07
u_{2wt}		0.26	0.06	0.77
u_{3xt}		1.00	1.00	1.00
u_{3wt}		1.00	1.00	1.00
u_{1t}	B	0.53	0.86	0.70
u_{2xt}		0.14	0.00	0.26
u_{2wt}		0.48	0.25	0.60
u_{3xt}		1.00	1.00	1.00
u_{3wt}		1.00	1.00	1.00
u_{1t}	C	0.60	0.57	0.25
u_{2xt}		0.03	0.00	0.27
u_{2wt}		0.93	0.12	0.35
u_{3xt}		1.00	1.00	1.00
u_{3wt}		1.00	1.00	1.00

For the definitions of the instrument sets, see Table 3. Alternative instrument sets are denoted $AZ^1_{t-1}, AZ^2_{t-1}, and AZ^3_{t-1}$.
The first contains lags 1,2, and 3 of excess equity returns for Germany, AZ^2_{t-1} contains lags 1,2, and 3 of excess returns to a DM deposit, and AZ^3_{t-1} contains lags 1,2, and 3 of excess US equity returns.

Table 10
CS Tests for the Ferson and Harvey Model
Case of the US and Japan

RESIDUAL	INSTR. SET	AZ^1_{t-1}	AZ^2_{t-1}	AZ^3_{t-1}
P-value				
u_{1t}	A	0.56	0.44	0.40
u_{2xt}		0.14	0.00	0.16
u_{2wt}		0.75	0.55	0.63
u_{3xt}		1.00	1.00	1.00
u_{3wt}		1.00	1.00	1.00
u_{1t}	B	0.35	0.32	0.03
u_{2xt}		0.00	0.00	0.04
u_{2wt}		0.45	0.41	0.05
u_{3xt}		1.00	1.00	1.00
u_{3wt}		1.00	1.00	1.00
u_{1t}	C	0.67	0.42	0.25
u_{2xt}		0.84	0.00	0.05
u_{2wt}		0.65	0.55	0.35
u_{3xt}		1.00	1.00	1.00
u_{3wt}		1.00	1.00	0.00

For the definitions of the instrument sets, see Table 3. Alternative instrument sets are denoted $AZ^1_{t-1}, AZ^2_{t-1}, and AZ^3_{t-1}$.
The first contains lags 1,2, and 3 of excess equity returns for Japan, AZ^2_{t-1} *contains lags 1,2, and 3 of excess returns to a Yen deposit, and* AZ^3_{t-1} *contains lags 1,2, and 3 of excess US equity returns.*

Table 11

CS Tests for the Ferson and Harvey Model

Case of the US and UK

RESIDUAL	INSTR. SET	AZ^1_{t-1}	AZ^2_{t-1}	AZ^3_{t-1}
		P-value		
u_{1t}	A	0.04	0.05	0.91
u_{2xt}		0.47	0.00	0.02
u_{2wt}		0.08	0.73	0.92
u_{3xt}		1.00	1.00	1.00
u_{3wt}		1.00	1.00	1.00
u_{1t}	B	0.11	0.06	0.70
u_{2xt}		0.90	0.00	0.16
u_{2wt}		0.19	0.30	0.39
u_{3xt}		1.00	1.00	1.00
u_{3wt}		1.00	1.00	1.00
u_{1t}	C	0.32	0.10	0.78
u_{2xt}		0.01	0.00	0.00
u_{2wt}		0.78	0.36	0.24
u_{3xt}		1.00	1.00	1.00
u_{3wt}		1.00	1.00	1.00

For the definitions of the instrument sets, see Table 3. Alternative instrument sets are denoted $AZ^1_{t-1}, AZ^2_{t-1}, and AZ^3_{t-1}$.
The first contains lags 1,2, and 3 of excess equity returns for the UK, AZ^2_{t-1} *contains lags 1,2, and 3 of excess returns to a £deposit, and* AZ^3_{t-1} *contains lags 1,2, and 3 of excess US equity returns.*

Table 12

CS Tests for Purely Autoregressive Model

Case of the US and Germany

RESIDUAL	XSJUNK	XSDIV	STRATE
CS Statistic χ^2(P-value)			
Returns			
u_{xt}	1.06(0.787)	0.74(0.863)	0.80(0.849)
u_{mt}	0.37(0.947)	0.38(0.944)	0.34(0.952)
u_{jt}	4.05(0.256)	2.24(0.524)	2.474(0.480)
Variances			
η_{xt}	4.03(0.258)	4.30(0.231)	3.59(0.309)
η_{mt}	1.64(0.650)	1.71(0.635)	2.04(0.565)
Covariances			
η_{xjt}	5.68(0.129)	4.79(0.188)	6.37(0.095)
η_{mjt}	151.92(0.000)	101.07(0.000)	192.0(0.000)

XSJUNK means lags 1,2,3 of excess junk bond spread,
XSDIV means lags 1,2,3 of excess dividend yield,
STRATE means lags 1,2,3 of Eurodollar rate.

Table 13
CS Tests for Purely Autoregressive Model
Case of the US and Japan

RESIDUAL	XSJUNK	XSDIV	STRATE
CS Statistic χ^2(P-value)			
Returns			
u_{xt}	0.055(0.996)	0.031(0.998)	0.093(0.993)
u_{mt}	0.360(0.948)	0.211(0.976)	0.057(0.996)
u_{jt}	10.404(0.015)	8.133(0.043)	3.618(0.306)
Variances			
η_{xt}	3.161(0.367)	3.334(0.342)	3.100(0.376)
η_{mt}	0.878(0.831)	0.960(0.811)	1.170(0.760)
Covariances			
η_{xjt}	62.981(0.000)	57.351(0.000)	31.123(0.000)
η_{mjt}	347.728(0.000)	313.341(0.000)	212.299(0.000)

For the definitions of XSJUNK,XSDIV & STRATE see Table 12.

Table 14

CS Tests for Purely Autoregressive Model

Case of the US and the UK

RESIDUAL	XSJUNK	XSDIV	STRATE
CS Statistic χ^2(P-value)			
Returns			
u_{xt}	0.959(0.811)	1.186(0.756)	0.877(0.831)
u_{mt}	0.018(0.999)	0.056(0.997)	0.008(0.999)
u_{jt}	6.731(0.081)	9.150(0.027)	9.292(0.026)
Variances			
η_{xt}	1.338(0.720)	1.532(0.675)	1.566(0.667)
η_{mt}	2.412(0.491)	3.360(0.339)	2.750(0.432)
Covariances			
η_{xjt}	12.608(0.006)	10.050(0.018)	14.222(0.003)
η_{mjt}	603.313(0.000)	283.365(0.000)	770.338(0.000)

For the definitions of XSJUNK,XSDIV & STRATE see Table 12.

Table 15

Pricing Errors Statistics

COUNTRIES	MODEL	ABS MEAN	RMS ERROR
GER,US	D&S - Inst. A	5.4424	8.1760
	D&S - Inst. B	5.6043	9.0030
	D&S - Inst. C	5.3804	8.1594
	F&H - Inst. A	4.9365	6.5350
	F&H - Inst. B	4.8731	6.4839
	F&H - Inst. C	4.9130	6.5321
	Factor-ARCH	4.8718	6.4808
JAP,US	D&S - Inst. A	6.1112	9.4168
	D&S - Inst. B	6.2691	10.1762
	D&S - Inst. C	6.6673	11.2405
	F&H - Inst. A	5.7550	7.3630
	F&H - Inst. B	5.6307	7.3434
	F&H - Inst. C	5.6935	7.3105
	Factor-ARCH	5.5671	7.2881
UK,US	D&S - Inst. A	8.7422	12.7205
	D&S - Inst. B	6.9369	10.4199
	D&S - Inst. C	6.6939	9.6600
	F&H - Inst. A	4.9233	6.2219
	F&H - Inst. B	5.0887	6.2303
	F&H - Inst. C	5.1957	6.4217
	Factor-ARCH	5.3778	7.6048

ABS MEAN is absolute mean error, RMS ERROR is root mean square error.

Table 16

Autocorrelations and White Noise Test

COUNTRIES	MODEL	ρ_1	ρ_2	ρ_3	ρ_6	ρ_{12}	P-value LB
GER,US	D&S - Inst. A	-0.0088	-0.0814	0.1570	0.0507	-0.0965	0.036
	D&S - Inst. B	0.0349	-0.0332	0.1473	0.0966	-0.0839	0.298
	D&S - Inst. C	-0.0999	-0.1277	0.1622	0.0421	-0.0779	0.001
	F&H - Inst. A	-0.0577	0.0002	0.0972	0.0063	-0.0902	0.328
	F&H - Inst. B	-0.0400	-0.0028	0.0961	0.0208	-0.1085	0.283
	F&H - Inst. C	-0.0428	-0.0001	0.0969	0.0132	-0.0846	0.355
	Factor-ARCH	-0.0064	0.0116	0.1179	0.0104	-0.0720	0.251
JAP,US	D&S - Inst. A	-0.0713	-0.0469	0.0464	0.0007	0.0138	0.959
	D&S - Inst. B	-0.0917	-0.0404	0.0219	0.0210	0.0100	0.797
	D&S - Inst. C	-0.1081	-0.0297	0.0264	0.0011	-0.0199	0.885
	F&H - Inst. A	0.0689	-0.0442	0.0540	-0.0149	0.0712	0.553
	F&H - Inst. B	0.0837	-0.0582	0.0407	-0.0152	0.0762	0.496
	F&H - Inst. C	0.0735	-0.0599	0.0433	-0.0218	0.0653	0.634
	Factor-ARCH	0.0660	-0.0694	0.0433	-0.0280	0.0704	0.636
UK,US	D&S - Inst. A	0.0617	0.0238	-0.1594	-0.0610	-0.0454	0.374
	D&S - Inst. B	0.0022	-0.0922	-0.1206	-0.0738	-0.0303	0.392
	D&S - Inst. C	0.0143	-0.0044	-0.1555	-0.0882	-0.0348	0.346
	F&H - Inst. A	-0.0965	-0.1155	-0.0717	-0.0804	-0.1273	0.328
	F&H - Inst. B	-0.0629	-0.1180	-0.0724	-0.0791	-0.1270	0.435
	F&H - Inst. C	-0.0327	-0.0911	-0.0498	-0.0586	-0.1348	0.587
	Factor-ARCH	-0.0581	-0.0702	-0.0359	0.0303	-0.0505	0.987

Autocorrelations 1,2,3,6 and 12 are in the designated ρ's; LB test refers to the LJUNG-BOX white noise test.

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