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Author(s): Geert Bekaert and Michael S. Urias
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Diversification, Integration and Emerging Market Closed-End Funds

GEERT BEKAERT and MICHAEL S. URIAS*

ABSTRACT

We study a new class of unconditional and conditional mean-variance spanning tests that exploits the duality between Hansen-Jagannathan bounds (1991) and mean-standard deviation frontiers. The tests are shown to be equivalent to standard spanning tests in population, but we document substantial differences in the small sample performance of alternative tests. Our empirical application examines the diversification benefits from emerging equity markets using an extensive new data set on U.S. and U.K.-traded closed-end funds. We find significant diversification benefits for the U.K. country funds, but *not* for the U.S. funds. The difference appears to relate to differences in portfolio holdings rather than to the behavior of premiums in the United States versus the United Kingdom.

THE DIVERSIFICATION BENEFITS FROM exposure to emerging equity markets recently have attracted enormous attention from individual and institutional investors in the U.S. and abroad. This article studies these benefits using a new data set of closed-end funds and a new class of mean-variance spanning tests. Throughout the article, we maintain that a set of asset returns provides diversification benefits relative to a set of benchmark returns if adding these returns to the benchmark leads to a significant leftward shift in the mean-standard deviation frontier. The analysis makes use of an historical sample, however, so any benefits we find may have little bearing on future performance.

Our analysis contributes to three different literatures in finance: the econometric literature on mean-variance spanning tests, a rapidly growing literature on emerging financial markets, and the literature on closed-end funds. First, we provide a unified development of unconditional and conditional tests

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of mean-variance spanning in the framework of Hansen and Jagannathan (1991), extending the work of Snow (1991) and DeSantis (1995). The tests exploit the duality between Hansen-Jagannathan bounds and mean-standard deviation frontiers to derive mean-variance spanning restrictions as a function of the marginal rate of substitution of representative investors. We provide analytical expressions for the test statistics and, in the unconditional case, show their equivalence with the standard regression-based test of Huberman and Kandel (1987). Furthermore, we demonstrate that the moment restrictions of the conditional test are equivalent to restrictions on the coefficients of a system of equations that are a close analogue to the Huberman and Kandel restrictions in the unconditional case. The tractability of our expressions facilitates an analysis of the small sample properties of the alternative tests.

Second, choosing the tests with the most favorable size and power properties, we document the diversification benefits from emerging equity markets. DeSantis (1994), Divecha et al. (1992), Harvey (1995), and other authors document substantial diversification benefits from investing in emerging equity market indices. A critical drawback of these studies is their use of market indices that generally ignore the high transaction costs, low liquidity, and investment constraints associated with investments in emerging markets. In effect, the indices represent an unrealistic performance benchmark for most investors, and they violate the investability assumptions behind the tests.

In an effort to address this problem, we examine diversification benefits from exposure to emerging markets that are actually attainable to international investors: closed-end country funds. The distinguishing feature of closed-end funds is that fund share prices generally deviate from their portfolio value (known as "net asset value" or NAV), and often trade at a premium when the assets are invested in closed or restricted markets. As a result, the returns from holding the fund shares may differ from those of the (possibly unattainable) portfolio in which the fund invests. In particular, investors may forego some of the diversification benefits from emerging markets by holding the funds instead of their underlying assets. We also compare the benefits from exposure to country funds with the benefits from the International Finance Corporation (IFC) Investable indices.

Several authors, including Bailey and Lim (1992), Urias (1992), Bodurtha et al. (1995), Chang et al. (1993), and Diwan et al. (1993), cast doubt on the diversification benefits from emerging market closed-end funds. We contribute to this literature in several ways. First, we study new tests that are more robust to the nonstandard features of the data, we incorporate conditioning information in some of our tests, and we motivate the choice of statistics with an analysis of their small sample properties. Second, we use an extensive new dataset of more than 40 U.S. and U.K.-traded closed-end funds, contrasting the performance of funds in the two trading markets. Finally, in the spirit of Bonser-Neal et al. (1990), we examine the extent to which diversification benefits from the funds are affected by the relaxation of investment barriers in several heavily restricted markets.

The main empirical results of the article can be summarized as follows. We find significant diversification benefits for the U.K. country funds, but *not* for U.S. funds. The performance difference for the funds appears to relate to differences in the risk characteristics of the assets of comparable U.S. and U.K. funds, and not to the behavior of premiums in the U.S. versus the U.K. Holding the IFC Investable indices corresponding to the funds in our sample provides significant diversification gains. Using lagged premium changes as conditioning instruments, we find significant diversification benefits for both U.S. and U.K. funds. The relaxation of investment barriers in Taiwan had a significant negative impact on the benefits from holding its country funds.

The remainder of the article is organized as follows. Section II develops the conditional and unconditional mean-variance spanning tests and discusses their small sample performance. Section III describes the characteristics of the U.S. and U.K. closed-end fund markets. Section IV details results from the conditional and unconditional tests. Section V examines the robustness of the results and proposes an interpretation, and Section VI concludes.

I. Mean-Variance Spanning Tests

Our tests for mean-variance spanning rely on the framework of Hansen and Jagannathan (1991). Recall the general conditional asset pricing restriction:

$$\mathrm{E}[(R_{t+1} + \iota)m_{t+1}|\Phi_t] = \iota, \tag{1}$$

where R_{t+1} represents a vector of *net* security returns at time $t + 1$, m_{t+1} is an investor's marginal rate of substitution or discount factor, ι is the unit vector, and Φ_t is the information available at time t.

The distinguishing feature of an asset pricing model is its specification for the discount factor, m_{t+1}. Note that equation (1) assumes frictionless markets and that the Law of One Price holds. The equality becomes an inequality if transaction costs are introduced (see Luttmer (1996)).

In the following sections we describe unconditional and conditional method-of-moments tests for mean-variance spanning that take the form of restrictions on the stochastic discount factor, m_{t+1}. The tests exploit the duality between mean-standard deviation frontiers of asset returns and Hansen-Jagannathan (1991) bounds on the mean-standard deviation frontier of investor marginal rates of substitution.

Section I.*A* demonstrates that our unconditional tests are equivalent to the standard regression-based tests for spanning developed by Huberman and Kandel (1987). Section I.*B* develops the conditional spanning tests. Section I.*C* introduces the test statistics. We choose one particular test in the conditional and unconditional cases based on its superior small sample properties, which are discussed in Section I.*D*. Readers interested only in the empirical results can proceed directly to Section II.

A. Unconditional Mean-Variance Spanning

Using the law of iterated expectations and equation (1), we can write the unconditional asset pricing restriction:

$$\mathrm{E}[R_{t+1}m_{t+1}] + \mathrm{E}[m_{t+1}] - \iota = 0. \tag{2}$$

Hansen and Jagannathan (1991) show that the linear projection of m_{t+1} onto the set of asset returns being priced has minimum variance in the class of all discount factors that satisfy equation (2). Hence, the discount factor,

$$m_{t+1}^{\alpha} \equiv \alpha + [R_{t+1} - \mathrm{E}(R_{t+1})]'\beta^{(\alpha)},$$

formed from the projection of m_{t+1} onto one-period returns, satisfies equation (2).

Partition $R_{t+1} \equiv [R'_{1,t+1}, R'_{2,t+1}]'$ and $\beta^{(\alpha)} \equiv [\beta_1^{(\alpha)\prime}, \beta_2^{(\alpha)\prime}]'$. We use this framework to test whether the n_1-vector of returns, $R_{1,t+1}$, spans the $n_1 + n_2$-vector of returns, R_{t+1}. The spanning restrictions in the Hansen and Jagannathan (1991) framework are:

$$\beta_2^{(\alpha)} = 0_{n_2} \qquad \wedge \qquad \mathrm{E}[R_{t+1}m_{t+1}^{\alpha}] + \mathrm{E}[m_{t+1}^{\alpha}]\iota_n = \iota_n \qquad \forall\, \alpha, \tag{3}$$

where $n = n_1 + n_2$. In the tests to follow, we will denote the assets with returns $R_{1,t+1}$ as the benchmark assets and the assets with returns $R_{2,t+1}$ as the test assets. R_{t+1} is spanned by the benchmark assets if the test assets, $R_{2,t+1}$, need not be included in the linear pricing kernel that prices R_{t+1}.

To develop some intuition for the restrictions in expression (3) it is helpful to relate them to the restrictions of Huberman and Kandel (1987) (HK-restrictions). Denote the restrictions in expression (3) as HJ-restrictions.

PROPOSITION 1. *The HJ-restrictions are equivalent to the HK-restrictions, where the HK-restrictions are given by:*

$$R_{2,t+1} = a + B\, R_{1,t+1} + \varepsilon_{t+1} \tag{4}$$

$$\mathrm{E}[\varepsilon_{t+1}] = \mathrm{E}[\varepsilon_{t+1}R_{1,t+1}] = 0$$

$$a = 0_{n_2} \qquad B\,\iota_{n_1} = \iota_{n_2}.$$

Defining the population mean returns, μ, and population covariance matrix, Σ_R, for the set of returns $R_{t+1} = [R'_{1,t+1}, R'_{2,t+1}]'$, partitioned as:

$$\mu = \begin{bmatrix} \mu_1 \\ \mu_2 \end{bmatrix}, \qquad \sum\nolimits_R = \begin{bmatrix} \sum_R^{(11)} & \sum_R^{(12)} \\ \sum_R^{(21)} & \sum_R^{(22)} \end{bmatrix},$$

the HK-restrictions are equivalent to:

$$\begin{aligned} \mu_2 - \textstyle\sum_R^{(21)}\left(\sum_R^{(11)}\right)^{-1} \mu_1 &= 0_{n_2} \\ \textstyle\sum_R^{(21)}\left(\sum_R^{(11)}\right)^{-1} \iota_{n_1} &= \iota_{n_2}. \end{aligned} \tag{5}$$

Proof: We begin by showing that the HJ-restrictions in expression (3) imply the HK-restrictions in expression (5). First, since m^{α}_{t+1} is only a function of $R_{1,t+1}$, pricing $R_{1,t+1}$ will identify $\beta^{(\alpha)}_1$. That is, the pricing restriction, $\mathrm{E}[(R_{1,t+1}+\iota_{n_1})m^{\alpha}_{t+1}]=\iota_{n_1}$, implies, using the definition of m^{α}_{t+1}, that:

$$\beta^{(\alpha)}_1 = (\textstyle\sum^{(11)}_R)^{-1}[\iota_{n_1}(1-\alpha) - \alpha\mu_1].$$

Pricing $R_{2,t+1}$ with this kernel, it follows that, $\forall\alpha$:

$$\alpha(\mathrm{E}[R_{2,t+1}] + \iota_{n2}) + \mathrm{E}[R_{2,t+1}(R_{1,t+1} - \mathrm{E}[R_{1,t+1}])']\beta^{(\alpha)}_1 = \iota_{n2}.$$

Rewriting, we find

$$\alpha(\mu_2 - \textstyle\sum^{(21)}_R(\sum^{(11)}_R)^{-1}\mu_1) \\ + [\textstyle\sum^{(21)}_R(\sum^{(11)}_R)^{-1}\iota_{n_1} - \iota_{n2} - [\sum^{(21)}_R(\sum^{(11)}_R)^{-1}\iota_{n_1} - \iota_{n2}]\alpha = 0_{n2}.$$

Since the above expression must hold for every value of α, we obtain the HK-restrictions in expression (5).

Now we show that if the returns R_{t+1} satisfy the HK-restrictions, we can find pricing kernels $m^{\alpha}_{t+1}=\alpha+[R_{t+1}-\mathrm{E}(R_{t+1})]'\beta^{(\alpha)}$ with $\beta^{(\alpha)}_2=0$ such that $\mathrm{E}[R_{2,t+1}m^{\alpha}_{t+1}]+\mathrm{E}[m^{\alpha}_{t+1}]\iota_{n_2}=\iota_{n_2}\forall\alpha$. Using the HK-restrictions and letting $m^{\alpha}_{t+1}=\alpha+[R_{1,t+1}-\mathrm{E}(R_{1,t+1})]'\beta^{(\alpha)}_1$, the lefthand side of the pricing restriction equals

$$B\mu_1\alpha + B\textstyle\sum^{(11)}_R\beta^{(\alpha)}_1 + \alpha\iota_{n_2}.$$

Substituting for $\beta^{(\alpha)}_1$, we find

$$B\mu_1\alpha + B[\iota_{n_1}(1-\alpha) - \alpha\mu_1] + \alpha\iota_{n_2},$$

which equals ι_{n_2} since the HK-restrictions hold. Q.E.D.

Proposition 1 shows that the unconditional mean-variance spanning restrictions in the Hansen-Jagannathan framework are equivalent—in population—to the standard mean-variance spanning restrictions of Huberman and Kandel (1987). Both test the same moment restrictions, as Ferson (1994) also shows.

Intuitively, the set of returns $[R'_{1,t+1}, R'_{2,t+1}]'$ is spanned by the subset of returns $R_{1,t+1}$ if $R_{2,t+1}$ can be written, up to an orthogonal error term, as a weighted average of $R_{1,t+1}$ with the weights summing to one. That is, up to a mean-zero, orthogonal factor, the returns on the test assets can be perfectly mimicked by a portfolio of the benchmark assets.

B. Conditional Mean-Variance Spanning

In contrast with standard approaches, the market-subset approach enables us to easily incorporate conditioning information for tests of conditional mean-variance spanning. Following a long tradition in empirical asset pricing, we add *scaled returns* to the set of assets under study. Define $z_t \equiv [z'_{1,t}, z'_{2,t}]'$ to be a k-vector of instruments with $z_{i,t}$ having predictive power for $R_{i,t+1}$, and $k = k_1 + k_2$. Let

$$q_{i,t+1} = \begin{bmatrix} R_{i,t+1} + \iota_{n_i} \\ (R_{i,t+1} + \iota_{n_i}) \otimes z_{i,t} \end{bmatrix}.$$

The symbol $\otimes$ indicates Kronecker product. For a valid pricing kernel m_{t+1}, it follows from equation (1) and the law of iterated expectations that

$$\mathrm{E}[q_{i,t+1} m_{t+1}] = p_i \equiv \mathrm{E}\begin{bmatrix} \iota_{n_i} \\ \iota_{n_i} \otimes z_{i,t} \end{bmatrix}. \tag{6}$$

As Cochrane (1994) explains, one can view $q_{i,t+1}$ $(i = 1, 2)$ as the payoffs to managed portfolios, where the amount invested depends on the realization of the instrumental variable, $z_{i,t}$. The expected price of these payoffs is p_i.

From equation (6) it is straightforward to define the conditional mean-variance spanning restrictions in the Hansen-Jagannathan framework (conditional HJ-restrictions). For the scaled returns $q_{t+1} \equiv [q'_{1,t+1}, q'_{2,t+1}]'$, let $\mu_{q,i}$ and $\sum_q^{(ij)}(i, j = 1, 2)$ represent the population first and second moments. The conditional mean-variance spanning restrictions are then given by

$$\mathrm{E}[q_{2,t+1} m_{t+1}^{\alpha}] = p_2 \qquad \forall\ \alpha, \tag{7}$$

where

$$m_{t+1}^{\alpha} = \alpha + [q_{1,t+1} - \mu_{q,1}]' \beta_{q,1}^{(\alpha)}$$

and

$$\beta_{q,1}^{(\alpha)} = (\textstyle\sum_q^{(11)})^{-1}[p_1 - \alpha \mu_{q,1}].$$

Note here that the weight on $q_{2,t+1}$ in the expression for $\beta_{q,1}^{(\alpha)}$ is already constrained to be zero.

The following proposition demonstrates that the conditional HJ-restrictions can be written as a conditional analogue to the standard unconditional HK-restrictions, which we will call the *conditional HK-restrictions*. This terminology may seem odd since the restrictions embedded in expression (7) are unconditional. However, they apply to a much expanded space of asset payoffs and involve expected prices of the scaled returns, which account for the conditioning information (see also Cochrane (1994)).

PROPOSITION 2. *The conditional HJ-restrictions in expression (7) are equivalent to the conditional HK-restrictions:*

$$q_{2,t+1} = \mathcal{A} + \mathcal{B}\, q_{1,t+1} + v_{t+1} \tag{8}$$

$$\mathrm{E}[v_{t+1}] = \mathrm{E}[v_{t+1} q_{1,t+1}] = 0$$

$$\mathcal{A} = \mu_{q,2} - \mathcal{B}\mu_{q,1} = 0_{n_2 k_2} \qquad \mathcal{B}p_1 = p_2.$$

The proof of Proposition 2 follows the arguments in Proposition 1 closely and is given in Appendix A.

Although conditioning information is incorporated into our spanning tests, the intuition remains similar to that of the unconditional case. Whereas returns are payoffs with prices equal to one, the use of scaled returns implies that some payoffs have prices different from one. In particular, the prices of the assets with payoffs $q_{i,t+1}$ are p_i. The interpretation of the conditional mean-variance spanning restrictions is that $q_{1,t+1}$ spans $[q'_{1,t+1}, q'_{2,t+1}]'$, when the assets or managed portfolios having payoffs $q_{2,t+1}$ can be written as a portfolio of the assets having payoffs $q_{1,t+1}$ with the same expected cost.

C. Test Statistics

In order to test the Hansen-Jagannathan mean-variance spanning restrictions, we prespecify two values for α, α_1 and α_2, and test the pricing restrictions using $m_{t+1}^{\alpha_1}$ and $m_{t+1}^{\alpha_2}$.[1] We can either impose $\beta_2^{(\alpha_i)} \equiv 0_{n_2}$, a likelihood ratio-type test, or we can estimate $\beta_2^{(\alpha_i)}$ and test $\beta_2^{(\alpha_i)} = 0_{n_2}$ ($i = 1, 2$), a Wald test.

Let us first consider a Wald test. In what follows, sample moments are indicated by hats above the symbols corresponding to population moments in the preceding sections. The sample moment conditions for the Wald test are:

$$h_T(\beta^{(\alpha_1)}, \beta^{(\alpha_2)}) = 1/T \sum_{t=1}^{T} \begin{Bmatrix} \{R_{t+1}[\alpha_1 + (R_{t+1} - \hat{\mu})'\beta^{(\alpha_1)}]\} + \alpha_1 - \iota_n \\ \{R_{t+1}[\alpha_2 + (R_{t+1} - \hat{\mu})'\beta^{(\alpha_2)}]\} + \alpha_2 - \iota_n \end{Bmatrix} = 0. \tag{9}$$

Or, more succinctly,

$$h_T(\beta) = C_T\beta + d_T, \tag{10}$$

[1] Testing the pricing restrictions for a single value of α is equivalent to a test that the mean-standard deviation frontier formed from the set of assets R_{t+1} *intersects* the mean-standard deviation frontier formed from $R_{1,t+1}$. Snow (1991) first studied this test. From two-fund separation, we know that specifying two different values for α to test jointly for frontier intersection at two points along the frontier is equivalent to a test of mean-variance spanning.

with $\beta \equiv [\beta^{(\alpha_1)\prime}, \beta^{(\alpha_2)\prime}]'$, and where the dimensionality of h_T is $2n \times 1$. C_T is a square matrix:

$$C_T = \begin{bmatrix} \widehat{\sum_R} & 0 \\ 0 & \widehat{\sum_R} \end{bmatrix}, \tag{11}$$

$$d_T = \begin{bmatrix} \alpha_1(\hat{\mu} + \iota_n) - \iota_n \\ \alpha_2(\hat{\mu} + \iota_n) - \iota_n \end{bmatrix}.$$

Thus, the estimator for β is

$$b = -C_T^{-1} d_T. \tag{12}$$

Let A be a $2n_2 \times 2n$ matrix that selects the elements of the $2n \times 1$ vector β that are zero under the null of spanning. When the weighting matrix in the GMM-system, W_T, is chosen optimally (Hansen (1982)), the variance-covariance matrix of β equals $(C_T W_T C_T')^{-1}$. Denote the Wald statistic by $\mathrm{MV_W}$. Let $Q_T = C'_T W_T C_T$. Then

$$\mathrm{MV_W} = T(AC_T^{-1}d_T)'[AQ_T^{-1}A']^{-1}(AC_T^{-1}d_T). \tag{13}$$

So there is a simple, analytical expression for this test.[2]

DeSantis (1995) proposes a similar test, by restricting the coefficients $\beta_2^{(\alpha_1)}$ and $\beta_2^{(\alpha_2)}$ in equation (9) to be zero. Hence, the DeSantis test is nothing more than a likelihood ratio-type test which requires estimation under the null. It is useful to think about the DeSantis test as arising from a restricted estimation. Let $\bar{b} = [b_1^{(\alpha_1)\prime}, b_1^{(\alpha_2)\prime}]'$ denote the restricted estimator of β (i.e., with the zero restriction on the coefficients of the test assets imposed). Then,

$$\bar{b} = \text{argmin } h_T(\beta)' W_T\, h_T(\beta) \quad \text{s.t.} \quad A\,\beta = 0. \tag{14}$$

Imposing the zero restrictions $\beta_2^{(\alpha_1)} = 0$ and $\beta_2^{(\alpha_2)} = 0$, we can write

$$h_T(\overline{\beta}) = \overline{C_T}\overline{\beta} + \overline{d_T} \tag{15}$$

with:

$$\overline{C_T} = \begin{bmatrix} 1/T \sum_{t=1}^{T} R_{t+1}(R'_{1,t+1} - \hat{\mu}_1) & 0 \\ 0 & 1/T \sum_{t=1}^{T} R_{t+1}(R'_{1,t+1} - \hat{\mu}_1) \end{bmatrix},$$

$$d_T = \begin{bmatrix} \alpha_1(\hat{\mu}_1 + \iota_{n_1}) - \iota_{n_1} \\ \alpha_2(\hat{\mu} \;\; + \iota_{n_1}) - \iota_{n_1} \end{bmatrix}$$

[2] Using the fact that $Ab = 0 = -AC_T^{-1}\, dT$ and the partitioned inverse formula, the link with the HK-restrictions is immediate. Ferson (1994) provides such a derivation.

and $\bar{\beta} = [\beta_1^{(\alpha_1)\prime}, \beta_1^{(\alpha_2)\prime}]'$. Conditional on the weighting matrix, W_T, the GMM-estimator is

$$\bar{b} = -[\overline{C'_T}W_T\overline{C_T}]^{-1}[\overline{C'_T}W_T\overline{d_T}]. \tag{16}$$

Note that the orthogonality conditions in equation (14) are linear in the parameters and so is the constraint. Consequently, Proposition 4 in Newey and West (1987a) implies that the Wald test and the likelihood ratio-type test are numerically equivalent. Denote the likelihood ratio-type test statistic by MV. It is given by:

$$\text{MV} = Th_T(\bar{b})'W_Th_T(\bar{b}). \tag{17}$$

Hence, conditional on an estimate for the weighting matrix, both tests have analytical representations. Of course, the equivalence relies on the use of the same weighting matrix. In small samples, the two approaches will not yield the same test values. The likelihood ratio test requires a two stage GMM estimation, with the sample moments, h_T, evaluated at second stage estimates of $\beta_1^{(\alpha_1)}$ and $\beta_1^{(\alpha_2)}$, and W_T chosen optimally following Hansen (1982). By contrast, the Wald test is computed in one stage. Both tests have an asymptotic chi-square distribution with degrees of freedom $2n_2$.

The conditional HJ-restrictions can be tested in the same way. The analytical expressions for both tests remain valid, with the sample moments of $q_{i,t+1}$ replacing the moments of $R_{i,t+1}$ in the expressions for C_T, $\overline{C_T}$, d_T, and $\overline{d_T}$, and d_T and $\overline{d_T}$ reflecting the expected prices of q_{t+1}, which may differ from 1. In effect, the conditional test is an unconditional test with scaled returns added to the sets of benchmark and test asset returns, where the scaled returns have prices different from 1.

The Wald and likelihood ratio tests proposed above require the specification of the weighting matrix. Define f_t by the equation $h_T = 1/T \sum_{t=1}^{T} f_t$. The optimal weighting matrix is a consistent estimate of the inverse of the spectral density at frequency zero of the orthogonal conditions, $S = \sum_{j=-\infty}^{\infty} \mathrm{E}[f_t f'_{t-j}]$. A popular estimate uses Bartlett weights (Newey and West (1987b)):

$$\hat{S} = 1/T \sum_{t=j+1}^{T} \sum_{j=1}^{\rho} \left(1 - \frac{j}{\rho+1}\right)[f_t f'_{t-j} + f_{t-j} f'_t] + 1/T \sum_{t=1}^{T} f_t f'_t.$$

Although the MV test is easy to compute, many choices remain that could affect its small sample properties. Under the null hypothesis, f_t should have zero mean and be serially uncorrelated. Hence ρ could be set equal to 0 in the expression for $\hat{S}$. However, removing the mean of f_t and allowing for some serial correlation in $\hat{S}$ can improve the reliability of the estimators (see Cochrane (1994) for a general discussion). Furthermore, iterating on the weighting matrix might also improve the small sample properties of GMM-estimators (see Ferson and Foerster (1994)). In a Monte Carlo experiment described

Table I
Test Statistics

Statistic	Demeaned	ρ Correction	No. stages
MV_1	no	no	2
MV_2	yes	no	2
MV_3	no	yes	2
MV_4	yes	yes	2
MV_5	yes	yes	7
MV_{W1}	—	yes	1
MV_{W2}	—	no	1

below, we examine 5 sub-statistics MV_i ($i = 1, \ldots, 5$). The properties of the different statistics are summarized in Table I.

To choose ρ in the serial correlation correction, we use the "optimal bandwidth" procedure described by Andrews (1991). In the applications below, we found that after six to nine iterations in the GMM estimation, the weighting matrix changed very little.

Computation of the Wald test, MV_W, requires an estimate for S, but no multiple stage estimation is necessary. Moreover, since the GMM system is exactly identified, the mean of the orthogonality conditions is zero. We compute two Wald statistics, MV_{Wj} ($j = 1, 2$). For $j = 2$, there is no serial correlation correction. For $j = 1$, we construct S using the optimal bandwidth procedure.

By contrast, the standard Huberman and Kandel (1987) mean-variance spanning test based on the restrictions in expression (4) assumes homoscedastic error terms and multivariate normal-distributed returns. It has an exact F-distribution with $2n_2$ and $2(T - n)$ degrees of freedom. Denote the Huberman-Kandel test by HK. It can be written as

$$HK = \left(\frac{T - n}{n_2}\right)\left[\frac{1}{\ell} - 1\right], \tag{18}$$

where ℓ represents the ratio of the determinants of the maximum likelihood estimators of the error covariance matrix for the unrestricted model (no spanning) and the restricted model (spanning), respectively. We accompany our unconditional mean-variance spanning results with results from the HK test.

D. Small Sample Properties

We conduct a Monte Carlo experiment to examine both the size and power of the test statistics MV_i ($i = 1, \ldots, 5$), MV_{Wj} ($j = 1, 2$) and HK.

Under the null of unconditional mean-variance spanning, we assume that the benchmark returns follow a vectorautoregressive process with conditionally heteroskedastic errors (following Bollerslev (1990)). The test asset returns are then a linear function of the benchmark returns and a normal, homoske-

Table II

Empirical Size and Power of Test Statistics and Sharpe Ratio

	Closed-End Fund Benchmark Assets			
	$\chi^2(2)$		$\chi^2(24)$	
	Size	Power	Size	Power
MV_1	6.30	69.54	15.76	21.60
MV_2	6.88	69.86	35.28	28.66
MV_3	6.02	67.16	6.52	13.98
MV_4	8.08	67.38	52.10	21.58
MV_5	7.82	66.76	47.30	21.72
MV_{W1}	7.32	66.42	40.56	13.70
MV_{W2}	6.50	68.20	24.56	15.88
HK	9.20	78.68	11.58	40.52
ΔSR	0.057	10.90	0.254	6.22

The empirical size (in percent) is for a 5 percent-size test. The power is the percentage of experiments yielding a value larger than the 5 percent critical value of the empirical distribution under the null. In the column heading "$\chi^2(n)$," $n/2$ is the number of test assets. In the case of $n = 2$, an equally-weighted U.K. fund index is used, and in the case of $n = 24$, twelve U.S.-traded funds constitute the test assets. ΔSR refers to the change in the Sharpe Ratio; the size column reports the change in the Sharpe Ratio at the 95th percentile of the empirical distribution under the null, whereas the power column records the percentage of experiments yielding a larger change than this value under the alternative. A total of 5,000 Monte Carlo replications are employed throughout.

dastic error term such that their mean and variance dynamics are completely driven by the benchmark returns and the mean-variance spanning restrictions are imposed. The model is estimated from the data and then used to generate samples of length 152 (our sample size in the empirical work) for $[R'_{1,t+1}, R'_{2,t+1}]$, on which the mean variance spanning tests are performed.

To investigate the power of the unconditional tests, we assume that an additional factor generates the test assets besides the benchmark assets. For the additional factor we use the IFC Investable Composite index return as a proxy for emerging market-specific risk (Section II describes this index). The alternative is calibrated so that the variance of the test assets explained by the new benchmark is 20 percent higher than under the null.

Details on the data generation processes and complete results are provided in an appendix available from the authors. Table II contains partial results that illustrate the main findings for the unconditional tests. Employing a serial correlation correction according to the optimal bandwidth scheme of Andrews (1991), but not demeaning the orthogonality conditions, leads to dramatically superior size properties for MV_3. This is evident when the number of test assets is twelve or as few as one. Although MV_3's power is weaker compared with some of the other tests, the relative loss in power is only significant for large GMM systems. Interestingly, the power of the HK statistic is superior to that of all the moment-based tests. However, the size distortion is worse than for the MV_3 statistic.

In general, both the size and power properties deteriorate substantially as the number of assets is increased. We conjecture that the saturation ratio (see Gallant and Tauchen (1991)) of the GMM system is the driving factor in the results. The saturation ratio is the total number of observations divided by the number of parameters to be estimated (including the parameters of the weighting matrix). When $n_2 \geq 7$, the saturation ratio drops below 10 for the unconditional tests. Unfortunately, GMM-systems with such low saturation ratios are common in the empirical finance literature. In the empirical work to follow, we use the MV_3 and HK statistics, with the caveat that the power may be weak when many test assets are included in the test.

For the conditional tests, we report MV_3. We suspect, however, that the properties of the moment-based tests may be poor in the conditional case, not only because of the dimensionality of the system, but also because of the effect of an additional dynamic variable in the orthogonality conditions, which can render the weighting matrix ill-behaved (Section I.*C* describes the construction of the conditional estimators). Preliminary Monte Carlo results confirm this suspicion, but we defer formal analysis to future work.

II. Data Description

Table III illustrates a 300 percent increase in mutual fund assets invested in emerging equity markets from the end of 1991 through 1993. Recall that mutual funds can be characterized by their open or closed format. Unlike open-end mutual funds, closed-end funds issue a fixed number of shares that trade on an exchange, freeing portfolio managers from trading to meet net redemptions or net purchases of fund shares. The advantages of a closed format are especially relevant for funds investing in illiquid or volatile markets, and often local market regulations will favor the closed format early in the liberalization process. As Table III demonstrates, open-end fund assets have increased sharply in certain regions since 1991, but closed-end funds remain an important component of foreign capital invested in emerging financial markets.[3] Closed-end funds primarily trade on organized exchanges in the United States, London, and Hong Kong, and over-the-counter in London.

Table IV contains some details of our sample of U.S. and U.K. closed-end funds. The sample consists of forty-three U.S. funds and thirty-seven U.K. investment trusts for the period beginning January 1986 through August 1993. Twenty-three of the U.S. funds and nineteen of the U.K. trusts are classified as emerging market funds, corresponding to the emerging markets defined by the IFC. The remainder of the funds in each group includes country funds investing in the securities of "mature" markets as defined by inclusion in the FT-Actuaries World Indices,[4] and a smaller group of closed-end funds that

[3] Our sample amounts to nearly 20 percent of outstanding emerging market closed-end fund assets in 1993.

[4] There are two exceptions. Mexico and Malaysia are part of the FT-Actuaries World Indices and the IFC emerging markets. We classify them as emerging markets.

Table III

Emerging Market Funds: Total Net Assets

Country or Region	Closed-End	Open-End	Total
	Panel A: End of 1991		
Southeast Asia	10367.5	7886.3	18253.8
Latin America	3272.3	367.4	3639.7
Global	2730.4	85.7	2816.1
Other	3234.0	1.2	3235.2
Total	15815.1	7129.0	22944.1
	Panel B: End of 1993		
Southeast Asia	12340.8	34121.7	46462.5
Latin America	5569.4	9266.4	14835.8
Global	9379.7	18271.2	27650.9
Other	3655.2	1496.0	5151.2
Total	30945.1	63155.3	94100.4

Figures are in millions of U.S. dollars. There is no overlap between country and regional totals. Includes all funds worldwide that invest in equity or fixed-income markets. Less than 20 percent of total fund assets in 1993 are fixed-income, with the majority of those invested in Latin American issues. There is a group of funds which invests in emerging Asia and Japan, with total assets of about $21 billion in 1993, that is excluded from the totals. Source is Lipper Emerging Markets Fund Service.

do not invest internationally. The sample attempts to include all emerging market U.S. funds and U.K. investment trusts with initial offerings prior to 1992. In all but two cases, the entire trading history of the emerging market funds is included in the sample.[5]

U.K. investment trusts are the equivalent of U.S. closed-end funds, but there are a number of institutional differences. Whereas U.S. closed-end funds are required to distribute 90 percent of realized gains to shareholders in a given year to qualify for exclusion from corporation tax, investment trusts must retain capital gains for reinvestment. In addition, fund expenses are deductible from taxable income for U.K. trusts, but not for U.S. funds. British institutions are the primary holders of investment trusts, in part because of their inclusion in the FT-Allshare Index, which is commonly indexed by large shareholders in Britain.[6] U.S. closed-end funds, by contrast, are largely the realm of individual investors, although institutions make up a larger share of the ownership of emerging market funds than of country funds investing in mature markets.

For each fund, the sample contains weekly price, NAV, exchange rate, and dividend information for the life of the fund. The database also includes information on capital changes, the composition of dividend payments, and

[5] The Mexico Fund and The Korea Fund, both listed on the NYSE, began trading prior to 1986.

[6] Table IV details which funds in our sample are included in the FT-Allshare Index.

Table IV

Country Fund Sample General Characteristics

Fund	Symbol	Start	Assets	Premium	$\%\sigma(r_t^P)$	VR1	VR2
Panel 1: U.S. Emerging Market Funds							
Argentina	AF	911021	55.9	12.833	4.780	1.121	0.308
Asia Pacific	APB	870424	184.1	−4.426	5.554	0.774	0.414
Brazil	BZF	880331	191.4	−11.143	6.193	1.532	1.806
Chile	CH	890926	165.0	−5.455	5.470	0.822	0.268
Templeton EM	EMF	870226	192.5	3.775	4.961	0.782	0.270
First Philippine	FPF	891108	137.8	−17.282	4.407	0.803	0.155
Indonesia	IF	900301	40.3	7.563	5.609	0.886	0.152
India Growth	IGF	880812	56.2	−3.634	5.447	1.180	0.584
Jakarta Growth	JGF	900410	35.6	1.712	5.290	0.943	0.157
Korea	KF	860102	254.7	54.816	6.109	0.916	0.247
Latin Am. Invt.	LAM	900725	73.9	−6.233	5.178	1.049	0.314
Latin Am. Equity	LAQ	911022	83.9	−5.773	3.956	0.967	0.437
Emerging Mexico	MEF	901002	103.9	−8.153	5.395	0.654	0.345
Malaysia	MF	870508	131.2	−2.185	6.785	0.833	0.210
Morgan Stanley EM	MSF	911025	184.4	4.998	2.998	1.099	0.273
Mexico	MXF	810608	697.7	−17.154	7.074	0.767	0.553
Portugal	PGF	891101	51.1	−5.278	5.352	0.866	0.187
ROC Taiwan	ROC	890512	251.1	−1.714	5.585	0.759	0.451
Scudder New Asia	SAF	870618	123.1	−10.701	4.841	0.808	0.423
Thai Capital	TC	900523	75.6	−7.992	4.954	0.866	0.508
Thai	THF	880217	210.4	14.310	5.578	0.940	0.400
Turkish Investment	TKF	891215	50.0	1.336	5.895	1.208	1.145
Taiwan	TWN	861216	161.7	26.897	6.992	1.534	0.915
Panel 2: U.K. Emerging Market Trusts							
Abtrust New Thai	ABTHAI	891222	28.2	−16.876	5.259	0.612	0.745
Abtrust New Dawn†	ABTST	890512	59.0	−12.315	3.963	0.799	0.622
Beta Global EM†	BGEM	900306	62.8	−9.961	2.962	1.296	0.832
Brazil*	BRAIT	920504	65.9	−11.643	4.881	0.762	0.730
Drayton Korea	DRKT	911206	43.0	−13.358	3.745	0.897	0.948
EFM Dragon†	EFMDR	880311	240.1	−9.243	5.702	0.986	0.306
Fleming EM†	FEMR	910726	120.8	−2.925	2.473	1.328	0.900
First Philippine†	FPT	891215	51.8	−30.048	4.635	1.330	1.030
Gartmore Pacific†	GEP	900119	112.9	−14.665	3.822	0.850	0.657
EFM Java	JAVA	900518	15.7	−5.364	4.241	1.328	0.624
Korea Liberalisation	KLF	900615	34.5	−20.012	5.710	1.424	0.679
Korea Europe*	KOREUR	890623	153.8	24.985	5.941	0.788	0.348
Latin America*	LATIT	900703	153.8	−6.227	5.207	1.149	0.555
Pacific Horizon*	PHIT	890829	25.1	−14.816	4.404	1.081	0.411
Schroder Korea*	SCHRKT	911213	47.1	−2.818	5.600	0.718	0.391
Siam Select	SIAM	900406	43.9	−7.948	2.718	2.124	1.831
Thornton Asian†	TAEMIT	890728	164.8	−11.809	3.740	0.927	0.759
Templeton EM†	TMPLT	890929	375.0	−1.097	3.086	1.228	0.782
Turkey	TURK	900803	39.8	−14.566	3.050	2.557	2.401

Table IV—*continued*

Country Fund Sample General Characteristics

Fund	Symbol	Start	Assets	Premium	$\%\sigma(r_t^P)$	VR1	VR2
Panel 3: U.S. Mature Market Funds							
Europe	EF	900427	95.2	−7.411	3.783	1.032	0.269
France Growth	FRF	900511	183.5	−13.328	4.415	0.923	0.256
Emerging Germany	FRG	900329	114.6	−12.504	4.502	0.954	0.416
Germany	GER	860718	142.6	3.805	5.933	0.912	0.200
New Germany	GF	900125	331.2	−13.272	4.712	1.004	0.480
Growth Fd Spain	GSP	900215	172.3	−13.602	3.684	0.721	0.461
First Australia	IAF	860102	98.2	−12.687	4.819	1.310	0.605
First Iberia	IBF	880419	51.4	−4.926	5.302	0.929	0.220
Irish Investment	IRL	900330	45.4	−17.186	3.496	1.090	0.441
Italy	ITA	860226	58.0	−9.906	5.205	0.932	0.297
Japan OTC	JOF	900314	79.0	6.267	5.376	1.016	0.597
Austria	OST	890922	64.8	−4.985	6.320	1.062	0.399
Singapore	SGF	900725	66.3	−7.599	4.058	0.932	0.205
Spain	SNF	880621	90.0	13.093	6.066	0.970	0.339
Swiss Helvetia	SWZ	870819	160.1	−5.202	3.945	0.911	0.331
United Kingdom	UK	870807	45.0	−14.047	4.470	0.811	0.392
Panel 4: U.K. Mature Market Trusts							
Fleming American†	FAMN	870206	378.2	−14.978	3.469	0.603	0.598
F & C Eurotrust†	FCEU	870102	191.4	−2.730	3.249	0.717	0.522
Fleming Far East†	FFET	870102	811.2	−17.096	3.792	0.556	0.502
First Ireland	FIC	900316	NA	−17.825	2.867	1.389	1.193
Fleming Japanese†	FLJA	870206	426.8	−13.317	4.002	0.615	0.579
First Spanish	FSPAN	880805	40.8	−10.740	3.508	0.799	0.670
German†	GRMIT	900302	54.9	−14.734	3.332	0.792	0.582
German Smaller Cos.†	GRMS	870102	62.8	−18.375	3.044	0.888	0.727
GT Japan†	GTJA	870102	222.8	−12.332	3.996	0.666	0.645
Govett Oriental†	LVIT	870102	855.2	−16.834	3.953	0.563	0.661
Paribas French†	PRBS	870515	58.1	−9.845	3.332	0.784	0.497
TR Pacific†	TRIPT	871127	152.2	−7.963	4.194	0.887	0.535
Panel 5: U.S. Domestic Market Funds							
Adams Express	ADX	860102	832.2	−7.197	2.409	0.957	0.678
Baker Fentress	BKF	871127	427.3	−17.127	2.433	0.753	0.453
General Am. Invest.	GAM	860102	514.0	−11.301	3.259	0.558	0.541
Source Capital	SOR	860102	277.3	3.547	2.052	0.828	0.420
Tricontinental Corp.	TY	860102	2047.8	−9.177	2.298	0.805	0.956
Panel 6: U.K. Domestic Market Trusts							
Alliance†	ALNCT	870102	1399.6	−15.952	2.195	0.431	0.857
Edinburgh†	EDIN	870102	1696.2	−17.575	2.801	0.333	0.535
Fleming Claverhouse†	FCLVR	870102	147.5	−9.226	2.949	0.543	0.829
Foreign & Colonial†	FRCL	870102	2168.5	−14.444	3.057	0.439	0.521

Table IV—*continued*

Country Fund Sample General Characteristics

Fund	Symbol	Start	Assets	Premium	$\%\sigma(r_t^P)$	VR1	VR2
Panel 6: U.K. Domestic Market Trusts							
Scottish Mortgage†	SMT	870102	1589.5	−17.690	2.880	0.344	0.634
TR City of London†	TRCL	870102	459.7	−7.211	3.063	0.289	0.633
Witan†	WTAN	870102	1404.3	−21.383	2.755	0.369	0.763

Notes: Figures cover the trading life of the funds between January 1986 and August 1993. All U.S. funds trade on the New York Stock Exchange (NYSE), except for IAF and IBF, which trade on the American Stock Exchange (AMEX). All investment trusts trade on the London Stock Exchange. Fund net assets are in millions of U.S. dollars, computed in April 1993. The source for investment trust assets is the *Association of Investment Trust Companies Monthly Information Service.* "NA" indicates not available. Other statistics are computed in the currency in which the fund trades. The investment trusts in the sample are quoted in pence except for BRAIT, KOREUR, LATIT, and SCHRKT, which are quoted in dollars.

Start is the date when fund coverage begins. It corresponds to the IPO date for U.S. funds offered after 1985 and for U.K. funds offered after 1986.

PRM_t is the premium, defined by: $P_t \equiv \mathrm{NAV}_t\mathrm{PRM}_t$. Premium is 100 times the average of $(\mathrm{PRM}_t - 1)$. $\sigma(r_t^P)$ is the percentage weekly standard deviation of log price returns.

The log premium change is then the difference in log price (r_t^P) and log NAV (r_t^N) returns: $\Delta\log(\mathrm{PRM}_t) = r_t^P - r_t^N$. (Dividends are included in price and NAV returns.)

VR1 is the variance ratio $\sigma^2(\Delta\log(\mathrm{PRM}_t))/\sigma^2(r_t^P)$, and VR2 is the variance ratio $\sigma^2(r_t^N)/\sigma^2(r_t^P)$.

'*' indicates investment trusts for which diluted NAV is used. A superscript † indicates U.K. funds in our sample that are included in the FT-Allshare Index as of January, 1994. See Section II for details.

missing values for each fund.[7] For the U.K. trusts, "diluted" NAV is also included. It is customary in Britain to quote diluted NAV, which is the net asset value per share diluted to the level that would result if all outstanding warrants that are in the money were exercised immediately. Warrants trade for U.S. funds as well, but are much less common. Our analysis of NAV characteristics uses "undiluted" NAV in all but five cases.[8]

The price, dividend and volume data was obtained daily from the CRSP database for U.S. funds; NAV information was obtained weekly from *Barrons* and the *Wall Street Journal*, and daily for several funds from Scudder, Stevens, and Clark. County Natwest Securities supplied the weekly price, NAV and diluted NAV numbers for the U.K. funds, and S.G. Warburg provided daily prices and diluted NAV. Seven-day deposit LIBOR rates in dollars and pounds were obtained from Bloomberg and the *Financial Times of London*.

In constructing weekly fund data, the correct U.S. price was paired with the weekly NAV data, which normally is computed on Friday, but also on Wednes-

[7] Missing values are only a problem for the U.S. NAV series for a few funds. In cases where we were unable to complete the series using alternate sources, we used the previous week's figure.

[8] The five investment trusts in our sample for which only diluted NAV is available are indicated in Table IV.

day or Thursday in several cases. Where possible the timing of indices was matched to the price and NAV information. A similar procedure was followed to construct weekly U.K. files from daily series.

Diversification benefits are measured in relation to a set of mature market benchmark returns. Individual country and regional total return indices and exchange rates for the mature markets were provided by Goldman, Sachs & Co. beginning in January 1986.[9] To assess the exposure of emerging market fund portfolios to the local market, we use three weekly total return dollar indices available from the Emerging Markets Data Base (EMDB) maintained by the IFC since December, 1988. These include local market indices representing a broad cross-section of shares in each country, the IFC Global Indices representing a set of the largest capitalization and most liquid shares in a given country or region, and the IFC Investable Indices. The IFC Investable Indices represent a value-weighted subset of the equities in the IFC Global Indices for which ownership restrictions are not legally binding. The Investable Indices are better than the Global Indices as proxies for local market exposure that is actually *attainable*, and hence they are more relevant in comparisons with closed-end funds. Nevertheless, they only include assets satisfying certain liquidity requirements, and high transaction costs and other market frictions limit the effective availability of investable shares. As there are some exceptions in coverage, Appendix B details the date when the series for each of the three IFC indices begins. The table also indicates the name of the local market index. For the emerging markets, we obtained weekly exchange rate data from the IFC, and daily rates from Tradeline International.

Table IV also includes information about the average departure of prices from NAV (known as the premium) for each fund. This aspect of the "closed-end fund puzzle" is well known for U.S. funds (see, for example, Lee et al. (1991) and Hardouvelis et al. (1993)). Ammer (1990) documents premium characteristics for U.K. investment trusts. The average weekly dollar premiums are 0.92 percent for the U.S. emerging market funds and -9.51 percent for U.K. emerging market trusts. For the mature market funds, the numbers are -7.09 percent for the United States and -13.06 percent for the United Kingdom. For the U.S. funds investing in domestic securities, the average premium in our sample is -8.25 percent, and for U.K. domestic market trusts the average is -14.64 percent.

The premiums on closed-end funds, while potentially informative about market integration, could bias our tests for diversification benefits. If closed-end funds in the United States or the United Kingdom contain a fund-specific risk factor not spanned by the benchmark assets, our inferences could reflect only the gains from exposure to this factor and not the enhancement from emerging market assets added to the benchmark portfolios. We will return to this issue in Section III.*A*.

[9] The FT-Actuaries World Indices (TM/SM) are jointly compiled by the Financial Times Limited, Goldman, Sachs & Co., and County NatWest Securities in conjunction with the Institute of Actuaries and the Faculty of Actuaries.

Finally, Table IV lists the value of fund assets in April of 1993, the volatility of fund price returns, and a variance decomposition of fund returns showing the proportion of fund return variance due to NAV returns and premium change variance. The variance decomposition indicates that in the majority of cases, and in each fund category, the change in the log fund premium explains a larger share of price return variance than NAV returns explain. Hence the premium anomaly is clearly evident in the second moment properties of fund returns.

III. Diversification Benefits from Emerging Markets

In this section, we test whether adding emerging market assets to a number of different benchmark portfolios from mature financial markets leads to a statistically significant shift in the mean-standard deviation frontier. Hence, we examine whether a set of benchmark asset returns spans the returns on both benchmark and test assets.

Recall that our test statistics follow an asymptotic chi-square distribution with degrees of freedom equal to two times the number of test assets. The degrees of freedom arise from the fact that we test jointly for frontier intersection at two distinct points along the frontier. From two fund separation, this is equivalent to a test of mean-variance spanning. Our conditional and unconditional test results are based on the MV_3 statistic, which had superior small sample properties in the Monte Carlo experiment (see Section I.*D*). We also report the standard Huberman and Kandel (HK) test with the unconditional results.

A. The Choice of Benchmark

The choice of benchmark assets is an important consideration for both implementing and evaluating the results of our mean-variance spanning tests. We prioritize several issues when making this choice. First, we wish to approximate the benchmark of a globally diversified dollar investor in 1990, the start of the test period. Alternatively, we might consider the benefits to an investor exclusively in U.S. assets. However, we find that the FT-Actuaries World Indices for the U.K., Europe excluding the U.K. and the Pacific are *not* spanned by the U.S. index. Hence, the benchmark of the four country and regional indices together provides a more stringent test of spanning. In addition, most combinations of individual mature country indices are spanned by this four-index benchmark, suggesting that it is a reasonable benchmark for the global investor.

A second consideration when choosing benchmarks is the possibility that closed-end fund industry factors could affect the test results. If the test asset returns contain a closed-end fund-specific risk factor not spanned by the benchmark FT-Actuaries returns, our tests could produce rejections of mean-variance spanning that do not reflect international diversification benefits. Since we are interested in whether emerging market funds provide diversification benefits relative to a worldwide mature market portfolio, we wish to control for closed-end fund effects. To do so, we construct a set of benchmark equally-weighted index returns from domestic closed-end funds and mature market country funds for both the United States and the United Kingdom.[10] This controls for the presence of closed-end fund-specific factors while providing a benchmark of mature market assets similar to the FT-Actuaries benchmark.[11]

An important final consideration in choosing the benchmark is the small sample properties of the tests when a large number of benchmark assets is included. Our Monte Carlo results suggest that we need to keep the number of assets in the tests small, hence we proceed using the two benchmarks of four assets.

B. Unconditional Spanning Results

In previous work, DeSantis (1994) finds that a U.S., European or World market portfolio spans the frontier of mature markets but not the frontier of emerging markets. DeSantis (1994) uses the Morgan Stanley Capital International (MSCI) indices for mature markets and the IFC Global Indices for the emerging markets. Similarly, Harvey (1995) rejects the hypothesis that the Global Indices and the eighteen MSCI mature market indices intersect the frontier of the mature market indices at a single point on the frontier.

In this section, we compare the diversification benefits from emerging market closed-end funds with the benefits from holding the corresponding IFC Investable Indices.[12] As described in Section II, the IFC Investable Indices correct for foreign ownership restrictions and exclude illiquid stocks to some extent, but they do not account for other barriers to investment such as high transaction costs or poor accounting standards. Table V summarizes the test assets we examine in Tables VI and VII. For example, we compare a set of twelve individual U.S. emerging country and regional funds with the set of twelve IFC Investable indices corresponding to the countries and regions the funds represent. We restrict investors to hold only funds that have corresponding IFC Investable indices, and only one fund for a particular country or region

[10] Most combinations of individual mature and domestic country funds are spanned by this index-of-funds benchmark.

[11] In fact, it turns out that the closed-end fund benchmark spans the FT-Actuaries benchmark, but not vice versa.

[12] There are at least three index funds which attempt to match the performance of the IFC Investable indices or some comparable. There is the open-end International Equity Index Fund Emerging Markets Portfolio managed by Vanguard Group, which tracks an MSCI index, a "semi-open" fund managed by the IFC tracking the IFC Investable Composite Index, and a closed-end fund, Baring Securities Emerging Market Index Tracker Fund, which trades over-the-counter in London and tracks an index published by Baring Securities. Of course, an investor could buy the stocks in the index directly at some nontrivial cost. A comparison between holding portfolios of country funds and holding investable indices presumes that the marginal investor in country funds can obtain the theoretically investable portfolios.

Table V

Mean-Variance Spanning Tests

Test Assets	n_2
U.S. Emerging Market Funds	12
Corresponding IFC Investables	12
U.K. Emerging Market Trusts	7
Corresponding IFC Investables	7
U.S. & U.K. Funds Emerging Market Funds	19
U.S. Emerging Market Fund Index	1
Corresponding IFC Investables Index	1
U.K. Emerging Market Fund Index	1
Corresponding IFC Investables Index	1
U.S. & U.K. Emerging Market Fund Indices	2

if there are multiple such funds, to enhance the power of our tests.[13] We also examine the equally-weighted and value-weighted index of the twelve funds and the corresponding twelve Investable indices. All fund and benchmark returns are measured at a weekly frequency in dollars.

For all tests, we report the p-value for the mean-variance spanning test statistic MV_3, estimated in two stages, which incorporates a serial correlation correction but does not demean the orthogonality conditions. Where available we also report the p-value according to its empirical distribution under the null. In our discussion, we will use 5 percent as the size of the test. Given that the power of our MV_3 tests ranges between 13 percent and 73 percent (see the Monte Carlo Appendix), readers concerned about Type-2 errors may wish to employ a higher size. For comparison, we report the p-value for the HK test as well.

Table VI, Panel 1, presents results from mean-variance spanning tests using the benchmark of four FT-Actuaries market and regional indices. We fail to reject mean-variance spanning for the twelve individual U.S. emerging market funds using MV_3. The test statistic falls within the third decile of the empirical distribution under the null. We also fail to reject spanning at the five percent level using all nineteen funds. However, we do reject strongly mean-variance spanning for the seven individual U.K. emerging market funds, according to both the chi-square distribution and the null empirical distribution at the 5 percent level.

The same tests were performed using equally-weighted indices of country fund returns, reducing the dimensionality of the problem considerably. We fail to reject spanning for the U.S. fund index at the five percent level, but we reject spanning for the U.K. fund index and the two indices together using the chi-square and empirical distributions for MV_3. The same pattern emerges for the HK statistic using the null empirical distribution (note that HK rejects

[13] Choosing alternate funds covering the same market does not change any of the results.

Table VI

Unconditional Mean-Variance Spanning Tests

Benchmark Assets	U.S. Funds	U.K. Funds	U.S. & U.K. Funds	U.S. Fund Index	U.K. Fund Index	U.S. & U.K. Fund Indices
Panel 1: U.S. and U.K. Emerging Market Funds						
U.S., U.K.,	0.6743	0.0109	0.4178	0.1215	0.0325	0.0213
Europe,	(0.7848)	(0.0304)		(0.1308)	(0.0370)	(0.0210)
Pacific	[0.3693]	[0.0000]	[0.0000]	[0.8615]	[0.0000]	[0.0001]
	[(0.3926)]	[(0.0000)]		[(0.7176)]	[(0.0010)]	[(0.0028)]
U.S. Domestic Index,	0.6015	0.0002	0.5382	0.1001	0.2721	0.1032
U.K. Domestic Index,	(0.7270)	(0.0010)		(0.1188)	(0.3052)	(0.1064)
U.S. Mature Market Index,	[0.7578]	[0.0000]	[0.0000]	[0.7254]	[0.0745]	[0.1415]
U.K. Mature Market Index	[(0.6970)]	[(0.0000)]		[(0.5902)]	[(0.1168)]	[(0.1714)]

Benchmark Assets	U.S. Investables	U.K. Investables	U.S. Investables Index	U.K. Investables Index
Panel 2: IFC Investable Indices				
U.S., U.K.,	0.0006	0.0030	0.0007	0.0016
Europe,	[0.0000]	[0.0000]	[0.0000]	[0.0000]
Pacific				
U.S. Domestic Index,	0.0080	0.0014	0.0078	0.0011
U.K. Domestic Index,	[0.0000]	[0.0000]	[0.0000]	[0.0000]
U.S. Mature Market Index,				
U.K. Mature Market Index				

We report p-values for the chi-square statistic MV_3 on the first line. Numbers in parentheses are approximate p-values from the empirical distribution of MV_3 under the null. See Section I.*C* or the appendix on the Monte Carlo available from the authors. Numbers in square brackets are the p-values of the F-statistic HK; bracketed parentheses indicate the p-values from the empirical distribution of HK under the null.

All returns are computed in dollars, and all indices are equally-weighted. The tests employ 152 weekly observations from September 1990 through August 1993.

In Panel 1, the U.S. funds included as tests assets are APB, BZF, CH, EMF, FPF, IF, LAM, MF, MXF, PGF, THF, and TKF. The U.K. funds are ABTHAI, ABTST, EFMDR, FPT, JAVA, LATIT, and TURK. These funds make up the fund indices as well. In cases where more than one fund covers a particular country or region, the longer-lived fund was chosen. The same rule applies to the mature market fund benchmarks.

FT-Actuaries World Indices make up the first set of benchmark assets. The Europe index does not include the United Kingdom. The second set of benchmark assets contains indices of closed-end funds.

In Panel 2, the IFC test assets were chosen to cover the same markets as the corresponding country funds assets. Both indices of IFC Investable returns are equally-weighted.

spanning for all nineteen funds using the F distribution at the five percent level).

Panel 1 of Table VI reports test results using the control benchmark of closed-end fund indices described in Section III.*A*. We again fail to reject mean-variance spanning for the twelve U.S. funds and for the U.S. fund index using MV_3. The rejection for the U.K. fund index and the U.S. and U.K. fund index together no longer holds, but we still reject spanning for the seven

Table VII

Conditional Mean-Variance Spanning Tests

Benchmark Assets	U.S. Fund Index	U.K. Fund Index	U.S. & U.K. Fund Indices
U.S. Domestic Index U.K. Domestic Index U.S. Mature Market Index U.K. Mature Market Index	0.0000	0.0000	0.0002

Each country fund index in the benchmark and test assets was scaled by a vector of ones and an index of its own lagged premium changes (the difference between price and NAV returns), doubling the degrees of freedom in the test to $4n_2$.

We report the p-value for the chi-square statistic MV_3.

All returns are computed in dollars, and the indices are equally-weighted. The tests employ 152 weekly observations from September 1990 through August 1993.

individual U.K. emerging market funds using the chi-square and the empirical distribution under the null. Again, the same rejections hold using the null empirical distribution of the HK statistic. Overall, the results from Panel 1 suggest that the individual U.K. emerging market funds provide diversification benefits relative to the benchmark portfolios. Constructing an index of these funds changes the distribution of returns so that they no longer provide such benefits.

Table VI, Panel 2, presents results from mean-variance spanning tests for the IFC Investable Indices corresponding to the test funds. For both the FT-Actuaries and the control benchmark, we strongly reject mean-variance spanning for the individual IFC indices corresponding to U.S. and U.K. emerging market funds. When we construct their equally-weighted indices, we can still reject the spanning hypothesis at the one percent level using MV_3. HK provides the same rejections. These rejections remain valid using the empirical distribution generated for Panel 1.[14] It is notable that we obtain a rejection of the spanning hypothesis for the indices corresponding to the U.S. funds, while we were unable to reject the hypothesis using the funds themselves. It appears that the Investable indices offer superior diversification benefits compared with the U.S. emerging market funds. Of course, there are considerable costs associated with obtaining the index performance relative to the costs of holding the funds.

In sum, results from unconditional mean-variance spanning tests for U.S. and U.K. emerging market funds indicate that there are diversification benefits from holding the individual U.K. emerging market funds. Constructing an index of emerging market funds appears to weaken the benefits from the U.K. funds. The results do not change when we control for closed-end fund-specific factors. By contrast, the IFC investable indices corresponding to the U.K. *and* the U.S. funds provide unequivocal diversification benefits.

[14] There is no reason to expect the empirical distribution to be substantially different for the statistics generated with the IFC indices as test assets.

C. Conditional Spanning Results

Consistent with most of the literature on diversification benefits from emerging markets, our focus up to this point has been on unconditional tests. A large literature on the predictability of international equity returns, including emerging markets (see, for example, Bekaert and Harvey (1995) and Harvey (1995)), suggests that dynamic strategies may yield improvements in risk-reward tradeoffs. There are several reasons to be skeptical about the practical viability of dynamic strategies in our context. Our sample size is relatively short, complicating inference on the value of many potential instruments. Moreover, the transaction costs of trading some of the closed-end funds in our sample, particularly the U.K. funds, may be substantial. There is also no reason to expect that conditioning information would improve the risk-reward characteristics of the benchmark and test assets differently, in a way that changes the spanning results. In light of these concerns, we focus on only one instrument: the closed-end fund discount. Thompson (1978) and Brauer (1988) have argued that U.S. closed-end fund discounts predict fund returns and can be used to construct abnormal portfolio returns.

We choose for the instruments $z_{i,t}$ the index of individual fund premium changes (price returns minus NAV returns) corresponding to each index return in R_{t+1}. To keep the dimensionality of our test small, we consider a subset of the vector q_{t+1} in equation (6), taking only each return and its *own* index of individual fund premium changes, which doubles the number of overidentifying restrictions in the spanning test to $4n_2$.

Table VII reports results from conditional mean-variance spanning tests using the control benchmark of domestic and mature market funds and emerging market fund indices as the test assets. Interestingly, we reject spanning for the scaled U.S. fund index as well as the scaled index of U.K. funds, and for U.S. and U.K. fund indices together, using the chi-square distribution for MV_3. One interpretation of the result is that the predictive power of premium changes can be exploited in a dynamic strategy that leads to diversification gains for U.S. funds not present when holding them passively. However, as Section I.*D* describes, preliminary Monte Carlo experiments indicate that the nature of the conditioning tests can lead to ill-behaved weighting matrices, hence the results must be interpreted with caution.

IV. Robustness and Interpretation of Results

A. Robustness

For the tests reported in Section III.*B*, we use equally weighted indices of funds to represent benchmark and (sometimes) test assets. Using an equally-weighted index could bias the results against finding diversification benefits if the index is not efficient in the space of test asset returns. In fact, the results are not materially different when we use value-weighted indices, except that we can reject mean-variance spanning for the index of U.S. funds using the control benchmark.[15] This result appears to be due to large weights on MXF and KF during the test period.

To assess whether our results are robust to the currency in which returns are expressed, we performed MV_3 tests with the U.S. and U.K. funds expressed in British pounds. Using the closed-end fund benchmark, the rejection for the seven U.K. funds remains using the chi-square distribution and the empirical distribution computed in dollars. As before, there are no other rejections for fund indices or individual U.S. funds, and the result holds for the FT-Actuaries benchmark.

Since the U.S. sample covers emerging markets not included in the U.K. sample, the results could reflect in part the different composition of the samples.[16] In fact, using the control benchmark and seven U.S. funds corresponding to the U.K. coverage, we still fail to reject spanning using the chi-square distribution and the empirical distribution. Again, this result is robust to the funds we choose when there is duplicate coverage, and to the benchmark used. When we exclude the global and regional emerging market funds from each sample (leaving four single-country funds in the United States and the United Kingdom), we reject spanning for the U.K. funds but not for U.S. funds.

Hence the U.S.-U.K. emerging market fund performance difference is robust to the method used to construct the indices, the currency in which returns are expressed, the different market coverage of the funds, and the presence of regional funds.

It has sometimes been argued that the movement toward market integration will reduce the diversification benefits of emerging markets. Since investment liberalizations can be viewed as moves toward market integration, their presence during the sample may affect our results. For four heavily restricted markets, a significant investment liberalization takes place in the middle of the test period. Closed-end country funds often receive special rights to invest in closed markets before restrictions are lifted for other foreign investors, so we might expect that the test assets are not spanned by the benchmark before but that they may be after the liberalization. To test this conjecture, we modify the specification of the discount factor as follows. Recall that

$$m_{t+1}^{\alpha_i} = \alpha_i + \beta_1^{(\alpha_i)}[R_{1,t+1} - \mathrm{E}(R_{1,t+1})] + \beta_2^{(\alpha_i)}[R_{2,t+1} - \mathrm{E}(R_{2,t+1})], \tag{19}$$

for $i = 1, 2$, where α_i is prespecified and does not affect the test statistic for mean-variance spanning. Instead of testing the hypothesis $\beta_2^{(\alpha_i)} = 0$ ($i = 1, 2$), we redefine $\beta_2^{(\alpha_1)} \equiv \widetilde{\beta_2^{(\alpha_i)}}\,D_{t+1}$, where D_{t+1} is equal to 1 before the liberalization and 0 afterwards. If the hypothesis $\widetilde{\beta_2^{(\alpha_i)}} = 0$ ($i = 1, 2$) cannot be rejected, the benchmark spans the test assets before *and* after the liberalization. On the

[15] The value-weighted indices are constructed using the value of individual fund assets in dollars in April, 1993 (see Table IV).

[16] The countries and regions covered by the set of U.K. emerging market trusts is a subset of those covered by the U.S. funds.

Table VIII

Mean-Variance Spanning and Investment Liberalizations

Test Assets	$[\beta_2^{(\alpha 1)}, \beta_2^{(\alpha 2)}]=0$	Before Liberalized	After Liberalized
BZF	0.6512	0.6631	0.7563
IGF	0.2719	0.2509	0.9731
KF & KOREUR Index	0.3085	0.1084	0.0078
ROC & TWN Index	0.1116	0.0082	0.5085

Information on restricted markets was obtained from Bekaert (1995) and the *IFC Index Methodology*, 1993.

Country funds normally receive special permission to trade in the securities of restricted markets before the markets are opened to all foreign investors. Foreign institutions in addition to country funds were granted access to 49 percent of voting common stock in Brazil in May 1991. India opened to foreign investors in November 1992, with a 25 percent cap on holdings by registered foreign institutions. In January 1992, foreign investors were allowed to acquire up to 10 percent of Korean company securities. In January 1991, foreign investors were granted permission to hold 10 percent of Taiwanese company securities. A variety of overriding restrictions apply in all cases. The test period begins in December 1989 and ends in August 1993.

The hypothesis, $[\widehat{\beta_2^{(\alpha 1)}}, \widehat{\beta_2^{(\alpha 2)}}]=0$, is tested using a GMM Wald statistic. The mean-variance spanning hypothesis before and after the liberalization is tested using the statistic MV_3; the p-value is reported. See Section I.*C* for details.

All returns are computed in dollars, and all indices are equally-weighted. The benchmark assets are the closed-end fund mature market and domestic indices.

other hand, if $\widehat{\beta_2^{(\alpha_i)}} \neq 0$ $(i = 1, 2)$, then the test assets are spanned by the benchmark assets after the liberalization, but not before.

We consider liberalizations for funds investing in four countries: Brazil, India, Korea, and Taiwan. The results along with details about the liberalizations in each country are contained in Table VIII. In the first column, the table reports a Wald statistic for the test that $\widehat{\beta_2^{(\alpha_i)}} = 0$ $(i = 1, 2)$, estimated from the GMM system. For all four countries we cannot reject the hypothesis that $\widehat{\beta_2^{(\alpha_i)}} = 0$ $(i\ 1, 2)$ at the five percent level. Failure to reject this hypothesis is not surprising, since in general the β coefficients in the discount factor are not estimated precisely.

Table VIII also reports results from mean-variance spanning tests before and after the liberalization. If the investment restriction was a binding constraint before the liberalization, we would expect to reject spanning before but not afterwards. In fact, we are unable to reject spanning before or after the change for two of the four countries we consider. We are able to reject the liberalization hypothesis for Taiwan: the Taiwan funds are not spanned before but are spanned after the liberalization. For Korea, however, we reject spanning after the liberalization but not before. In sum, examining specific investment liberalizations for heavily restricted markets reveals that only the opening of Taiwan's market had a significant negative impact on the diversification benefits from holding closed-end funds.

B. Economic Significance

In the empirical work, we chose α_1 to be the mean of one divided by one plus the conditional riskless rate, where the conditional riskless rate is the seven-day deposit dollar LIBOR rate. Using the duality between (unconditional) mean-standard deviation frontiers of discount factors and asset returns, this choice for α_1 enables us to approximate the change in the Sharpe Ratio for a given set of test assets at the risk free rate.[17] The Sharpe Ratio measures the slope of the line from the riskless rate to the tangency portfolio on the efficient frontier, and is commonly known as the "reward to risk" ratio (see Sharpe (1994)). It gives the largest mean return per unit of standard-deviation risk attainable for the assets in question.

We document the change in the approximate Sharpe Ratio corresponding to the shift in the mean-standard deviation frontier of asset returns when the test assets are added to the benchmark assets. For a particular test of mean-variance spanning in Table VI, the change in the Sharpe Ratio in Table IX measures the economic importance of the shift in the efficient frontier. For example, the rejection of mean-variance spanning for the individual U.K. funds and the control benchmark (Table VI, Panel 1) is associated with a Sharpe Ratio change of about 0.096 in Table IX. By contrast, the rejection for the corresponding IFC indices (Table VI, Panel 2) is associated with a Sharpe Ratio change of 0.28.

As an aid for interpreting these numbers, Table IX reports the empirical p-value of the change in the Sharpe Ratio under the null of mean-variance spanning (Section I.*D* describes the Monte Carlo experiments; Table II gives partial results on the small sample distribution of the Sharpe Ratio change). Although the change in the Sharpe Ratio should be zero under the null of spanning, mean-variance mathematics implies that this number will be upwardly biased. Our Monte Carlo results demonstrate that this upward bias is severe. For example, the mean of the small sample distribution of the change for the individual U.K. funds is 0.092. Hence, despite the fact that we reject spanning for these assets, the Sharpe Ratio change of 0.096 is not significant, falling in the sixth decile of the empirical distribution. However, the Sharpe Ratio change for the corresponding IFC Investables of 0.28, indicating an increase in expected return per unit of risk three times that of the U.K. funds, is highly significant using the distribution generated for the U.K. funds.

The interpretation of the reward to risk change in the conditional setting no longer corresponds to the change in the Sharpe Ratio since we include return payoffs with prices different from one. However, we can still examine improvements in the reward to risk tradeoff for the conditional frontier relative to the point α_1. The quantity reported in Table IX for the conditional tests is given by the square root of the following expression:

$$\{[p - \alpha_1\mu_q]'(\textstyle\sum_q)^{-1}[p - \alpha_1\mu_q] - [p_1 - \alpha_1\mu_{q,1}]'(\textstyle\sum_q^{(11)})^{-1}[p_1 - \alpha_1\mu_{q,1}]\}.$$

[17] The approximation comes from the fact that we do not observe an unconditionally risk free rate. Hence because of Jensen's inequality, we do not obtain the exact Sharpe Ratio using α_1.

Table IX

The Change in the Sharpe Ratio

Benchmark Assets	U.S. Funds	U.K. Funds	U.S. & U.K. Funds	U.S. Fund Index	U.K. Fund Index	U.S. & U.K. Fund Indices
			Panel 1: U.S. and U.K. Emerging Market Funds			
U.S., U.K., Europe, Pacific	0.1618 (0.4172)	0.1294 (0.2318)	0.2405	0.0680 (0.0366)	0.0564 (0.0572)	0.0897 (0.0458)
U.S. Domestic Index, U.K. Domestic Index, U.S. Mature Market Index, U.K. Mature Market Index	0.1311 (0.5840)	0.0958 (0.4092)	0.1988	0.0473 (0.0724)	0.0246 (0.1988)	0.0578 (0.1368)

Benchmark Assets	U.S. Investables	U.K. Investables	U.S. Investables Index	U.K. Investables Index
		Panel 2: IFC Investable Indices		
U.S., U.K., Europe, Pacific	0.4126	0.3051	0.0940	0.0239
U.S. Domestic Index, U.K. Domestic Index, U.S. Mature Market Index, U.K. Mature Market Index	0.3732	0.2801	0.0584	0.0117

Benchmark Assets	U.S. Fund Index	U.K. Fund Index	U.S. & U.K. Fund Indices
	Panel 3: Conditional Tests: U.S. and U.K. Emerging Market Funds		
U.S. Domestic Index, U.K. Domestic Index, U.S. Mature Market Index, U.K. Mature Market Index	0.0488	0.0270	0.0514

Notes: On the first line for each benchmark, the table reports the approximate change in the Sharpe Ratio from adding the test assets to the benchmark assets. In Panel 3, this number is not strictly a Sharpe Ratio, but can be interpreted as the change in the conditional reward to risk ratio (See Section IV.*B*, for details). Numbers in parentheses are the approximate p-values from the empirical distribution of the change under the null hypothesis of spanning.

All returns are computed in dollars, and all indices are equally-weighted. The tests employ 152 weekly observations from September 1990 through August 1993.

In Panels 1 and 3, the U.S. funds included as tests assets are APB, BZF, CH, EMF, FPF, IF, LAM, MF, MXF, PGF, THF, and TKF. The U.K. funds are ABTHAI, ABTST, EFMDR, FPT, JAVA, LATIT, and TURK. In cases where more than one fund covers a particular country or region, the longer-lived fund was chosen. The same rule applies to the mature market fund benchmarks.

FT-Actuaries World Indices make up the first set of benchmark assets. The Europe index does not include the United Kingdom. The second set of benchmark assets contains indices of closed-end funds.

In Panel 2, the IFC test assets were chosen to cover the same markets as the corresponding country funds assets. Both indices of IFC Investable returns are equally-weighted.

Table X

Unconditional Mean-Variance Spanning Tests: U.S. and U.K. Emerging Market Fund NAVs

Benchmark Assets	U.S. NAVs	U.K. NAVs	U.S. & U.K. NAVs	U.S. NAV Index	U.K. NAV Index	U.S. & U.K. NAV Indices
U.S., U.K.,	0.0001	0.0000	0.0065	0.0000	0.0000	0.0000
Europe,	[0.0000]	[0.0000]	[0.0000]	[0.0000]	[0.0000]	[0.0000]
Pacific	{0.3084}	{0.0904}	{0.3189}	{0.1031}	{0.0491}	{0.1041}
U.S. Domestic Index,	0.0921	0.0170	0.5944	0.0006	0.0000	0.0004
U.K. Domestic Index,	[0.0000]	[0.0000]	[0.0000]	[0.0000]	[0.0000]	[0.0000]
U.S. Mature Market Index,	{0.2769}	{0.0637}	{0.2933}	{0.0599}	{0.0267}	{0.0603}

Notes: The table reports the p-value for the chi-square statistic MV_3 on the first line. The probability value of the F-statistic HK is reported in square brackets. Numbers in curly brackets indicate the approximate change in the Sharpe Ratio from adding the test assets to the benchmark assets. All returns are computed in dollars, and all indices are equally-weighted.

The tests employ 152 weekly observations from September 1990 through August 1993.

The U.S. funds whose NAVs are included as tests assets are APB, BZF, CH, EMF, FPF, IF, LAM, MF, MXF, PGF, THF, and TKF. The U.K. funds are ABTHAI, ABTST, EFMDR, FPT, JAVA, LATIT, and TURK. These funds make up the NAV indices as well. In cases where more than one fund covers a particular country or region, the longer-lived fund was chosen. The same rule applies to the mature market fund benchmarks.

FT-Actuaries World Indices make up the first set of benchmark assets. The Europe index does not include the United Kingdom. The second set of benchmark assets contains indices of closed-end fund returns.

The rejection of the hypothesis that the conditional benchmark spans the conditional test assets for the United States and the United Kingdom corresponds to a change in the reward to risk ratio of 0.049 and 0.027, respectively.

C. Interpretation

There are two interpretations for the performance difference of U.S. and U.K. funds that we have some power to distinguish in our setting. The first relates to managerial skill as reflected in the properties of the NAV for comparable funds, and the second relates to the behavior of premiums.

If there were no price segmentation between fund prices and NAV, investors would enjoy the full benefits from holding the assets of the funds. In such case, the performance differences observed would reflect the composition (perhaps changing over time) of the different funds. Table X reports results of mean-variance spanning tests that add fund NAV returns to both benchmarks. The results indicate that both U.S. and U.K. fund NAV returns provide significant diversification gains using the FT-Actuaries benchmark. However, we cannot reject the possibility that the individual U.S. fund NAVs are spanned by the closed-end fund control benchmark using MV_3, although the U.K. fund NAVs, the equally-weighted index of U.K. funds NAVs and the index of U.S. fund NAVs are not. Given our Monte Carlo results, failure to reject spanning for the

Table XI

Country Fund NAV and Premium Characteristics: Common U.S. and U.K. Emerging Market Coverage

Country or Region	Fund	Domicile	NAV α_N	NAV $\beta_{L\cdot I}$	NAV $\beta_{G\cdot I}$	NAV β_I	NAV $\bar{R}^2_N$	Premium γ_T	Premium $\bar{R}^2_{PR}$
Asia-Pacific	APB	U.S.	0.002		−0.020	0.621	0.380	0.825	0.079
			(0.001)		(0.068)	(0.055)		(0.224)	
	SAF	U.S.	0.000		0.455	0.617	0.390	0.780	0.066
			(0.002)		(0.166)	(0.097)		(0.263)	
	EFMDR	U.K.	0.001		−0.141	0.591	0.365	0.214	0.061
			(0.001)		(0.085)	(0.081)		(0.147)	
	GEP	U.K.	0.002		−0.014	0.616	0.353	0.473	0.091
			(0.001)		(0.072)	(0.078)		(0.120)	
	PHIT	U.K.	−0.001		0.117	0.530	0.256	0.208	−0.005
			(0.001)		(0.073)	(0.077)		(0.264)	
	TAEMIT	U.K.	−0.000		0.001	0.516	0.229	0.267	0.027
			(0.001)		(0.096)	(0.080)		(0.154)	
Global	EMF	U.S.	0.004		0.037	0.704	0.506	0.427	0.002
			(0.001)		(0.060)	(0.047)		(0.225)	
	ABTST	U.K.	0.002		0.275	0.448	0.146	0.585	0.193
			(0.002)		(0.112)	(0.092)		(0.119)	
	BGEM	U.K.	0.001		0.194	0.380	0.228	0.464	0.100
			(0.001)		(0.094)	(0.062)		(0.097)	
	TMPLT	U.K.	0.004		−0.042	0.347	0.093	0.477	0.089
			(0.002)		(0.142)	(0.081)		(0.082)	
Latin America	LAM	U.S.	0.002		0.054	0.524	0.242	0.845	0.040
			(0.001)		(0.064)	(0.077)		(0.316)	
	LATIT	U.K.	−0.001		0.061	0.776	0.405	−0.361	0.096
			(0.001)		(0.050)	(0.078)		(0.196)	
Indonesia	IF	U.S.	−0.001	0.331	0.009	0.454	0.583	0.465	0.001
			(0.001)	(0.132)	(0.173)	(0.031)		(0.298)	
	JGF	U.S.	−0.000	0.526	0.100	0.479	0.693	0.816	0.076
			(0.001)	(0.158)	(0.172)	(0.031)		(0.223)	
	JAVA	U.K.	−0.001	0.064	1.117	0.266	0.197	0.183	0.047
			(0.002)	(0.435)	(0.406)	(0.082)		(0.158)	
Test of Equal Coefficients (p-value):				0.3084	0.0396	0.0094	0.0007†	0.0462	
Philippines	FPF	U.S.	0.002	−0.008	0.408	0.235	0.426	0.114	0.035
			(0.001)	(0.051)	(0.102)	(0.039)		(0.196)	
	FPT	U.K.	0.002	0.170	0.849	0.400	0.333	0.362	0.060
			(0.003)	(0.184)	(0.316)	(0.090)		(0.182)	
Test of Equal Coefficients (p-value):				0.2045	0.1321	0.0760	0.0009†	0.3154	
South Korea	KF	U.S.	0.001	0.800		0.436	0.595	0.300	0.001
			(0.001)	(0.111)		(0.057)		(0.240)	
	KLF	U.K.	−0.003	0.154		0.316	0.086	−0.256	−0.010
			(0.003)	(0.304)		(0.152)		(0.194)	
	KOREUR	U.K.	0.001	0.904		0.471	0.471	−0.141	−0.013
			(0.002)	(0.148)		(0.050)		(0.157)	

Table XI—*continued*

Country Fund NAV and Premium Characteristics: Common U.S. and U.K. Emerging Market Coverage

Country or Region	Fund	Domicile	α_N	NAV $\beta_{L\cdot I}$	$\beta_{G\cdot I}$	β_I	$\bar{R}^2_N$	Premium γ_T	$\bar{R}^2_{PR}$
Test of Equal Coefficients (p-value):				0.0615		0.4545	0.1147†	0.1327	
Thailand	TC	U.S.	0.000	0.283	1.152	0.761	0.788	0.824	0.120
			(0.001)	(0.134)	(0.504)	(0.033)		(0.314)	
	THF	U.S.	0.000	0.179	0.919	0.811	0.910	0.828	0.146
			(0.001)	(0.086)	(0.240)	(0.021)		(0.208)	
	ABTHAI	U.K.	0.000	0.323	0.554	0.750	0.608	0.584	0.105
			(0.002)	(0.172)	(0.355)	(0.043)		(0.114)	
	SIAM	U.K.	0.001	0.081	−0.671	0.474	0.275	0.614	0.242
			(0.002)	(0.299)	(0.635)	(0.088)		(0.117)	
Test of Equal Coefficients (p-value):				0.4174	0.0067	0.0000	0.0000†	0.7378	
Turkey	TKF	U.S.	0.002	0.516	0.176	0.798	0.806	0.519	0.241
			(0.002)	(0.127)	(0.378)	(0.034)		(0.339)	
	TURK	U.K.	0.003	−0.017	0.717	0.254	0.106	0.403	0.028
			(0.003)	(0.128)	(0.375)	(0.083)		(0.133)	
Test of Equal Coefficients (p-value):				0.0045	0.3854	0.0000	0.0000†	0.7415	

Notes: To compare the risk exposure of fund portfolios, we estimate the following multivariate linear time series regression using OLS:

$$r_t^N = \alpha_N + \beta_T r_t^T + \beta_{L\cdot I} r_t^{L\cdot I} + \beta_{G\cdot I} r_t^{G\cdot I} + \beta_I r_t^I + \epsilon_t.$$

The dependent variable is the log return on NAV for the fund. r_t^T is the trading market return (the FT-Actuaries U.K. or U.S. index return), $r^{L\cdot I}$ and $r^{G\cdot I}$ are the residuals from a linear projection of IFC Local and Global index returns on the Investable Index, and r^I is the return on the Investable index. Numbers in parentheses are Newey-West (1987b) standard errors computed using three lags. All returns include dividends and are measured in dollars. Blank columns indicate that the corresponding index was unavailable or of insufficient length. The South Korean Investable index is available only late in the sample, so $r^{L\cdot I}$ is replaced with $r^{L\cdot G}$ and r^I with r^G. $\overline{R_N^2}$ is the R^2 adjusted for degrees of freedom.

Recall that the fund premium at time t, PRM_t, is defined by $P_t \equiv \text{NAV}_t\text{PRM}_t$. To document fund premium characteristics, we regress the premium change for each fund, $\Delta\log(\text{PRM}_t)=r_t^P-r_t^N$, on the same set of independent variables:

$$r_t^P - r_t^N = \alpha_{PR} + \gamma_T r_t^T + \gamma_{L\cdot I} r_t^{L\cdot I} + \gamma_{G\cdot I} r_t^{G\cdot I} + \gamma_I r_t^I + v_t.$$

$\bar{R}^2_{PR}$ is the R^2 in the premium regression adjusted for degrees of freedom.

For individual countries, the table reports the p-value from a Wald χ^2 test that the regression coefficients are the same across U.S. and U.K. funds. The Andrews (1991) optimal bandwidth is employed for the GMM weighting matrix. A superscript † indicates the p-value from the test that all three (two for South Korea) coefficients in the NAV regression are jointly the same across funds.

twelve U.S. fund NAVs using the control benchmark may reflect lack of power. The HK statistic does, however, reject spanning in this case.

Of course, these tests are not strictly valid, since the NAVs are not investable in many cases. Nevertheless, the exercise demonstrates the intuition that fund investors effectively forego some of the diversification benefits from the underlying assets in order to circumvent the barriers from investing in the assets directly. It is striking that the Sharpe Ratio improvements here are similar to those obtained for the IFC Investable indices in Section IV. Hence, the fact that the funds constitute managed portfolios does not fully explain their weaker diversification potential relative to the IFC Investables. Yet the spanning tests are not very revealing about differences between U.S. and U.K. managers. Differences in managerial skill could be reflected in the NAV compositions of two funds investing in the same market, or in the market timing abilities of different managers.

Table XI contrasts the portfolio exposure of U.S. and U.K. funds that invest in the same emerging market or region during the test period. We run a multivariate regression of NAV returns on the IFC Investable index return, on the residuals from a linear projection of the local market index return and the IFC Global index return onto the IFC Investable return ($r^{L\cdot I}$ and $r^{G\cdot I}$, respectively), and on the trading market return (r^T).[18] The results clearly suggest that the risk exposure of U.K. and U.S. funds investing in the same emerging market differs. For example, $\beta_{G\cdot I}$ for the U.S. funds investing in Indonesia (IF and JGF) is less than 0.2, whereas for the U.K. fund JAVA, $\beta_{G\cdot I}$ is 1.117 and significant. However, the sensitivity of the U.S. Indonesian funds to the IFC Investable is greater than that of JAVA, 0.454 and 0.476 versus 0.266. A Wald test rejects that $\beta_{G\cdot I}$ and β_I are jointly the same for U.S. and U.K. Indonesia funds at the five percent level. Wald tests reject equal U.S. and U.K. portfolio betas for Thailand and Turkey as well.[19] A Wald test for the equality of all three coefficients together rejects for the Philippines. The index returns explain a large share of the variation in NAV returns for most funds, as indicated by $\overline{R_N^2}$.

The majority of the funds do not exhibit "abnormal" returns as measured by α_N, the intercept term in the regression. α_N is significantly different from zero (but small) primarily for the regional funds: APB, GEP, EMF, TMPLT, and LAM; it is also significant for FPF. It is well known, however, that α_N may indicate poor performance even when the fund manager is a superior market timer.[20] The differences in Table XI suggest that the superior diversification benefits from U.K. emerging market funds could arise from portfolio selection, but make no meaningful statement about the market timing ability of the managers.

Since price segmentation is prevalent in the closed-end fund market (recall Table III), the performance difference may in part reflect differences in the premium characteristics of U.S. and U.K. funds targeting the same markets.

[18] Note that nontrading problems could affect the consistency of our estimates of risk exposure to the indices. In addition, dynamic trading strategies by fund managers might reduce the explanatory power of the indices.

[19] For Thailand, only TC is used for the U.S. coverage because of high correlation between THF and TC. The result is the same using THF.

[20] Cumby and Glen (1990) study the alpha measure and alternatives that account for this problem in the context of international open-end mutual funds.

In the extreme case where a U.S. and U.K. fund hold the same portfolio, the premium behavior must be the culprit. There is no obvious reason, however, for why premiums would behave differently in the two markets. Section II highlighted the fact that the clienteles differ for U.S. and U.K. funds, with British institutions being the primary holders of U.K. funds and retail investors the primary holders of U.S. funds. This could lead to differences in the liquidity of shares, or perhaps differences in shareholder rationality (see Lee et al. (1991) for a story of U.S. shareholder sentiment).[21] It is also true that the tax treatment of U.S. and U.K. funds differs.

Consequently, although we control for closed-end fund factors in the tests by using a benchmark of domestic and mature market closed-end funds, the time-variation in premiums might still differ in a way that favors the U.K. funds. Table XI reports results from a regression of fund premium changes (defined as the difference between log price and log NAV returns) on the fund's trading market return, controlling for the local market in which the fund invests.[22] One can imagine that if U.S. fund premiums were more correlated with their trading market than comparable U.K. funds, the diversification benefits of U.S. funds would be less. In fact, such a pattern is not demonstrated by Table XI. Of the 15 U.K. funds with common U.S. coverage, eight have a statistically significant, positive coefficient (γ_T) with respect to the trading market. Similarly, six of the 11 U.S. funds with common U.K. coverage have a statistically significant and positive coefficient with respect to the trading market, controlling for the local market. In most cases, however, γ_T is larger in magnitude for the U.S. funds. Yet for funds targeting individual countries, we cannot reject that γ_T is jointly the same across the trading markets except marginally for Indonesia. Note that in general, the adjusted R^2 for the regressions is quite low.

VI. Conclusion

The fanfare surrounding emerging equity markets continues to attract international investors. Most studies that address the portfolio benefits from emerging markets gloss over the transaction costs and other barriers to investment associated with these markets. In the present article we examine the diversification benefits from holding closed-end country funds that invest in emerging markets, and compare them to the diversification benefits associated with the IFC Investable indices. Emerging market closed-end funds represent exposure to emerging markets that is actually attainable by foreign investors at relatively low cost, whereas the IFC Investables are attainable in theory but ignore all effective investment costs or restrictions with the exception of foreign ownership restrictions.

[21] It may also explain differences in portfolio composition, for example, if the optimal portfolios of retail and institutional investors differ.

[22] The results are no different using an equally-weighted index of U.K. or U.S. domestic closed-end funds as the trading market return.

We examine the issue of emerging market diversification benefits using a new class of mean-variance spanning tests based on the framework of Hansen and Jagannathan (1991). We demonstrate their equivalence with the standard tests of Huberman and Kandel (1987) in the unconditional setting, and show that our conditional tests evaluate restrictions on the coefficients of a system of regressions that are analagous to the Huberman and Kandel restrictions in the unconditional case. With this analytical framework we examine the small sample properties of our tests.

We find that U.K. emerging market funds provide statistically significant diversification gains in unconditional tests, while comparable U.S. funds do not. The IFC Investable indices corresponding to the funds in our sample yield unequivocal diversification benefits. A Monte Carlo experiment demonstrates that our unconditional tests have fairly low power against one interesting alternative for large GMM systems, but the power is about 70 percent for smaller systems. Nevertheless, the performance result is robust to a number of extensions, and appears to derive from differences in portfolio selection by fund managers. Using lagged fund premiums as conditioning information produces significant gains for both U.S. and U.K. funds.

Two unresolved issues from the analysis warrant further study. A more detailed analysis of the small sample properties of our mean-variance spanning tests would be useful. For the conditional tests, a treatment of size and power is still missing. Moreover, a number of additional tests could be explored (see, for example, Ferson et al. (1993) and Jobson and Korkie (1989)). A second unresolved issue is the performance difference between U.S. and U.K. funds. Plausible explanations include differences in the taxes, capitalization, and liquidity of shares, as well as different investment clienteles.

Appendix

A. Proof of Proposition 2

We first show that the conditional HJ-restrictions imply the conditional HK-restrictions. Using the expression for $\beta_{q,1}^{(\alpha)}$, the conditional HJ-restrictions

$$\mathrm{E}[q_{2,t+1}m_{t+1}^{\alpha}] = p_2 \quad \forall \alpha$$

imply

$$\alpha\mu_{q,2}\sum{}^{(21)}(\textstyle\sum_q^{(11)})^{-1}[p_1-\alpha\mu_{q,1}]=p_2 \quad \forall\alpha \Rightarrow \alpha[\mu_{q,2}-\mathcal{B}\mu_{q,1}]+\mathcal{B}p_1-p_2=0_{n_2k_2}.$$

Hence, the conditional HK-restrictions hold.

It remains to be shown that if the conditional HK-restrictions hold, we can find pricing kernels that satisfy the conditional HJ-restrictions in expression (7). Letting $m_{t+1}^{\alpha} = \alpha + [q_{1,t+1} - \mu_{q,1}]'\ \beta_{q,1}^{(\alpha)}$, it follows that

$$\mathrm{E}[(\mathcal{B}q_{1,t+1} + v_{t+1})m_{t+1}^{\alpha}] = \mathcal{B}\mu_{q,1}\alpha + \mathcal{B}\mathrm{E}[q_{1,t+1}(q_{1,t+1} - \mu_{q,1})]'\beta_{q,1}^{(\alpha)},$$

using the fact that v_{t+1} is orthogonal to m_{t+1}^{α}. Substituting for $\beta_{q,1}^{(\alpha)}$, we obtain

$$\mathcal{B}\mu_{q,1}\alpha + \mathcal{B}\mathrm{E}[q_{1,t+1}(q_{1,t+1} - m(q_{1,t+1} - \mu_{q,1})]'(\textstyle\sum_q^{(11)})^{-1}[p_1 - \alpha\mu_{q,1}]$$
$$= \mathcal{B}\mu_{q,1}\alpha + \mathcal{B}[p_1 - \alpha\mu_{q,1}],$$

which equals p_2 since the conditional HK-restrictions hold. Q.E.D.

B. Index Coverage

Index Coverage for Emerging Equity Markets

Country	Local Index		IFC Global	IFC Investable
		Latin America		
Argentina	Bolsa Indice	881230	881230	881230
Brazil	BOVESPA	880331	881230	881230
Chile	IGPA	881230	881230	881230
Mexico	BMV General	881230	881230	881230
		Asia		
India	FE Bombay	881230	881230	921106
Indonesia	JSE Composite	881230	900928	900928
Korea	KSE Composite	860102	881230	920103
Malaysia	KLSE Composite	881230	881230	881230
Philippines	Manila Com/Ind	881230	881230	881230
Taiwan, China	TSE Average	881230	881230	910104
Thailand	SET	881230	881230	881230
		Europe/Mideast		
Portugal	BTA	881230	881230	881230
Turkey	ISE	881230	881230	890804

Notes: The table details the start of coverage for the emerging equity market indices used in the article. The name of the local market index is noted along with its coverage. All indices are taken from the Emerging Markets Data Base maintained by the IFC, where coverage generally begins in December 1988 (note that there are several important exceptions). Supplemental dates for the KSE Composite were obtained from the *KSE Fact Book* and for the BOVESPA from the Sao Paulo Stock Exchange.

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