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ABSTRACT

Using an extensive new data set on U.S. and U.K.-traded closed-end funds, we examine the diversification benefits from emerging equity markets and the extent of their integration with global capital markets. To measure diversification benefits, we exploit the duality between Hansen-Jagannathan bounds [1991] and mean-standard deviation frontiers. 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. To investigate global market integration, we compute the reduction in expected returns an investor would be willing to accept to avoid investment barriers in six countries. We find evidence of investment restrictions for Indonesia, Taiwan and Thailand, but not for Korea, the Philippines or Turkey.

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1 Introduction

The diversification benefits from exposure to emerging financial markets recently have attracted enormous attention among individual and institutional investors in the U.S and abroad. By some estimates, almost 90 billion dollars have flowed into the financial markets of developing countries since 1991, much of it in the form of closed and open-end mutual funds. Closed-end mutual funds, whose shares trade on an exchange different from the fund's assets, were the original vehicles for foreign investment in emerging financial markets.[1]

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 funds invest. The first objective of this paper is to formally assess the diversification benefits from holding the shares of closed-end funds investing in emerging markets and to compare them to the benefits from holding the emerging market indices. Our second objective is to use information on the differences between emerging market fund returns and net asset value returns to assess the degree of global capital market integration of emerging markets.

DeSantis [1993], Divecha et al. [1992], Harvey [1994] and other authors document substantial diversification benefits from investing in emerging equity market indices, but they ignore the high transaction costs and other constraints associated with investing in emerging equity markets. By contrast, a number of studies, including Bailey and Lim [1992], Urias [1992], Bodurtha et al. [1993], Chang et al. [1993] and Diwan et al. [1993], cast doubt on the diversification benefits from emerging market closed-end funds, but offer no formal assessment. The diversification gains from holding shares of closed-end fund portfolios can actually be achieved by foreign investors at relatively low cost when compared with holding emerging market shares directly.

To accomplish our first objective, we rely on the market-subset methodology developed by De Santis [1992] and Snow [1992]. A set of asset returns provides diversification benefits relative to a set of benchmark returns if adding these returns to the benchmark returns leads to a significant shift in the mean-standard deviation frontier. We characterize the economic

[1] For example, The Korea Fund, launched in 1984 and trading on the New York Stock Exchange, was the first financial vehicle through which foreign investors gained access to the Korean stock market.

importance of the frontier shift and we test its statistical significance using a mean-variance spanning test. The test exploits the duality between Hansen-Jagannathan [1991] bounds and mean-standard deviation frontiers of asset returns to derive the mean-variance spanning restrictions as a function of the marginal rate of substitution of representative investors. The test is robust to conditional heteroscedasticity and autocorrelation and can easily incorporate conditioning information. We propose a Wald test alternative to the likelihood ratio-type tests of DeSantis and Snow, and investigate the small sample properties of the two methods using Monte Carlo experiments.

The degree to which a particular market is integrated with global capital markets may have some bearing on the diversification potential of that market. Researchers generally conduct tests of market integration using different assets from different markets, conditioning on an asset pricing model that may or may not account for investment restrictions.[2] Bekaert and Harvey [1994] propose an alternative framework that accounts for time-variation in capital market integration using a regime-switching model. They document the difficulty in inferring the degree of capital market integration from the restrictions imposed in particular markets. Differences in the local and global pricing of equities have the potential to better describe the incidence and effectiveness of investment barriers. Hence, our second objective is to exploit the price segmentation inherent in country fund premia for information on the extent of global market integration for a number of countries.

We begin by specifying and estimating the discount factor which prices all traded assets in the U.S. and the U.K., including the emerging market funds. Using this discount factor we then estimate the expected return a global investor is willing to forego in order to avoid the investment barriers in individual countries. In effect, when investment barriers are binding except for the country funds, price segmentation may arise forcing investors to give up some of the diversification benefits from emerging markets.

Surprisingly, there have been few attempts to use country fund data to address questions of market integration. Bonser-Neal et al. [1990] use event study methods to determine the impact of specific investment liberalizations on country fund premiums, while Diwan et al. [1992] and Eun et al. [1994] analyze country fund price segmentation in the context of static mean-variance models of fund prices when direct access to the local market is restricted to

[2]Examples of models that assume global integration are Harvey [1991] and Wheatley [1988]. Errunza and Losq [1985] and Stulz[1981], among others, propose models that specifically incorporate certain market frictions.

some constant fraction of outstanding shares.

The richness of our data set makes our approach to diversification benefits and market integration feasible. Our tests are applied to a sample of over 40 U.S. and U.K. traded emerging market closed-end funds. We are able to compare the diversification benefits of U.S. and U.K. funds in light of measured differences in portfolio holdings. In our examination of market integration, we have typically several funds that target the same emerging market country or region, which should enhance the power of our tests. To control for the special features of closed-end fund pricing, we also carry out tests that use country funds investing in the equity markets of developed countries (what we call 'mature market' funds) and domestic funds (funds holding assets of the market where they trade) as benchmark assets. Hence we do not take a stand on what explains the discounts on domestic closed-end funds (see, for example, Lee et al. [1991] and Bekaert and Urias [1994a] for potential explanations).

The main results of the paper can be summarized as follows. We find significant diversification benefits for the U.K. country funds, but *not* for U.S. funds. Holding the International Finance Corporation (IFC) Investable indices corresponding to the funds in our sample provides significant diversification gains. Using price and NAV information for funds targeting six countries, we find evidence of effective barriers to investment in Indonesia, Taiwan and Thailand, but not in Korea, the Philippines or Turkey.

Of course, since our analysis is based on an historical sample, the results may have no bearing on future performance. In particular, as the emerging financial markets covered in our sample move to increase accessibility and improve efficiency, diversification benefits may decline.

The remainder of the paper is organized as follows. Section 2 describes the data and some of the different characteristics of the U.S. and U.K. closed-end fund markets. Section 3 details the tests for mean-variance spanning and results from the tests. Section 4 presents our measure of market integration and applies it to the data, and Section 5 concludes.

2 Data Description

Table 1 summarizes emerging equity market fund assets at the end of 1991 and 1993. The last two years have seen a 300 percent increase in mutual fund assets invested in emerging equity markets. 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 1 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. Closed-end funds primarily trade on organized exchanges in the U.S., London and Hong Kong, and over-the-counter in London.

Table 2 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,[3] and a smaller group of closed-end funds that 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.[4]

U.K. investment trusts are the equivalent of U.S. closed-end funds, but there are a number of institutional differences. While 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.[5] 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. See Bekaert and Urias [1994a] for more details.

For each fund, the sample contains weekly price, NAV, exchange rate, dividend, and

[3]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.

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

[5]Table 2 details which funds in our sample are included in the FT-Allshare Index.

volume information for the life of the fund. For the U.K. trusts, 'diluted' NAV and warrant prices are 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.[6] The U.S. data was obtained from primary sources including *Barrons* and the *CRSP* database. Data on investment trusts was obtained from County NatWest Securities and S.G. Warburg.

Diversification benefits are measured in relation to a set of mature market benchmark returns. Individual country and regional total return indices for the mature markets were provided by Goldman, Sachs & Co.[7] 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. 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. Additional details of the sample are contained in the Data Appendix.

Table 2 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 premiums are 0.92% for the U.S. emerging market funds and −9.51% for U.K. emerging market trusts. For the mature market funds, the numbers are −7.09% for the U.S. and −13.06% for the U.K.

[6]The five investment trusts in our sample for which only diluted NAV is available are indicated in Table 2.

[7]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.

For the U.S. funds investing in domestic securities, the average premium in our sample is -8.25%, and for U.K. domestic market trusts the average is -14.64%.

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 U.S. or the U.K. 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 control for this possibility by using a sample of mature market and domestic closed-end funds from the U.S. and U.K. as the benchmark assets.

Finally, Table 2 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.

3 Diversification Benefits

3.1 Mean-Variance Spanning

We perform tests of mean-variance spanning based on the market subset methodology studied by DeSantis [1992] and Snow [1991]. The tests exploit the duality between Hansen-Jagannathan bounds [1991] on investor marginal rates of substitution and the mean-standard deviation frontier of asset returns. They are related to the mean-variance spanning tests of Huberman and Kandel [1987] and Ferson et al. [1993] (see Appendix A).

We apply this test to the question of whether emerging market country fund returns are spanned by index returns from mature financial markets. Returns from the mature financial markets serve as the benchmark or spanning assets, and the emerging market fund returns are the test assets. The test assesses whether the leftward shift in the mean-standard deviation frontier is statistically significant.

We begin with the general conditional asset pricing restriction:

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

where R_{t+1} represent a vector of security returns, 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.

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

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

Note that (1) and (2) assume frictionless markets and that the Law of One Price holds. The equality becomes an inequality if transaction costs are introduced (see Luttmer [1991]).

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 (2). The discount factor $m^{\alpha}_{t+1} \equiv \alpha + [R_{t+1} - E(R_{t+1})]'\beta$, formed from the projection of m_{t+1} onto one-period returns, then satisfies (2) and may be substituted for m_{t+1}:

$$E[R_{t+1}m^{\alpha}_{t+1}] + E[m^{\alpha}_{t+1}] - \iota = 0. \tag{3}$$

Conditional on a value for α, m^{α} can be used to place bounds on the standard deviation of random variables which satisfy the asset pricing equation for the asset returns R since it is the minimum variance discount factor. From these bounds comes the duality with the mean-variance frontier of asset returns.

Now we can examine whether or not a subset of the assets to be priced, say the subvector R^1_{t+1} from $R_{t+1} \equiv [R^{1\prime}_{t+1}, R^{2\prime}_{t+1}]'$, prices all of the assets in R_{t+1}. Let there be n_1 returns in R^1_{t+1}, the benchmark returns, and n_2 returns in R^2_{t+1}, the test returns. If we restrict the n_2 coefficients β_2 on R^2_{t+1} in equation (3) to be zero, we can test the overidentifying restrictions implied by (3):

$$E\{R_{t+1}[\alpha + (R^1_{t+1} - E[R^1_{t+1}])'\beta_1]\} + \alpha - \iota = 0. \tag{4}$$

Given a value for α, there are $n_1 + n_2$ moment conditions and n_1 parameters to estimate, hence there are n_2 overidentifying restrictions.

When a risk-free asset is known, the test is equivalent to a test that the mean-standard deviation frontier formed from the set of assets R intersects the mean-standard deviation frontier formed from R_1 at the point with the highest Sharpe ratio. DeSantis [1993] proposes prespecifying two different values for α to test jointly for frontier intersection at two points along the frontier. From two fund separation the test is equivalent to a test of mean-variance spanning. The sample moment conditions in the test for mean variance spanning are:

$$h_T(\beta_{11}, \beta_{12}) = 1/T \sum_{t=1}^{T} \left\{ \begin{array}{l} \{R_{t+1}[\alpha_1 + (R^1_{t+1} - E[R^1_{t+1}])'\beta_{11}]\} + \alpha_1 - \iota \\ \{R_{t+1}[\alpha_2 + (R^1_{t+1} - E[R^1_{t+1}])'\beta_{12}]\} + \alpha_2 - \iota \end{array} \right\} = 0. \qquad (5)$$

In the empirical work, we choose α_1 to be the mean of one divided by one plus the conditionally riskless rate. Using the duality between 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.[8] Hence we accompany the test results for diversification benefits with a measure of the economic importance of these benefits.

Consistent estimates for β_{11} and β_{12} can be computed analytically from (5) using a two-stage Generalized Method of Moments (GMM) estimator. Following Hansen [1982], an optimally chosen weighting matrix W_T and sample moments h_T are evaluated at second stage estimates of β_{11} and β_{12}. Define b_{11} and b_{12} to be the estimates. Then the quadratic form $Th_T(b_{11}, b_{12})'W_T h_T(b_{11}, b_{12})$ has an asymptotic χ^2 distribution with $2n_2$ degrees of freedom under the null hypothesis that the coefficients on the test assets in the discount factors are equal to zero.

There are several advantages of the market-subset approach compared with the mean-variance spanning tests proposed by Huberman and Kandel [1987]. It allows us to easily incorporate conditioning information for tests of conditional mean-variance spanning. The test does not require a distributional assumption on asset returns, and it is robust to general forms of conditional heteroscedasticity and autocorrelation.[9]

We incorporate conditioning information into our tests by scaling asset returns in equation (1) by variables in the information set at time t. Consider a vector of predetermined variables, z_t. Applying the law of iterated expectations, the moment conditions we exploit are:

$$E[[(R_{t+1} + 1) \otimes z_t]m_{t+1}] - E(\iota \otimes z_t) = 0, \qquad (6)$$

where $\otimes$ denotes Kronecker product. As with many of our tests, the conditional tests will

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

[9] Ferson et al. [1992] propose a test of mean-variance spanning using GMM as a special case of a test for latent variables.

employ indices of fund returns for benchmark and test assets instead of individual fund returns. Let z_t be the unit vector augmented by the index of individual fund premium changes (price returns minus NAV returns) corresponding to each index return in R_{t+1} (z_t has dimension $n1+n2+1$). To keep the dimensionality of our test small, we consider a subset of the vector $(R_{t+1} \otimes z_t)$, taking only each return scaled by the unit vector and its *own* index of individual fund premium changes, which doubles the number of overidentifying restrictions in the spanning test to $4n_2$. Scaled asset returns have been interpreted as managed portfolios by a number of authors. Many authors, including 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.

3.2 Test Statistics

We can rewrite the orthogonality conditions (5) as

$$h_T(b_{11}, b_{12}) = 1/T \sum_{t=1}^{T} f_t(b_1) = C_T b_1 + d_T, \tag{7}$$

with

$$C_T = \begin{bmatrix} 1/T \sum_{t=1}^{T} R_{t+1}(R_{t+1}^{1\prime} - E[R_{t+1}^{1\prime}]) & 0 \\ 0 & 1/T \sum_{t=1}^{T} R_{t+1}(R_{t+1}^{1\prime} - E[R_{t+1}^{1\prime}]) \end{bmatrix}$$

$$b_1 = \begin{bmatrix} b_{11} \\ b_{12} \end{bmatrix} \quad d_T = \begin{bmatrix} \alpha_1[1/T \sum_{t=1}^{T}(R_{t+1}+1)] - \iota \\ \alpha_2[1/T \sum_{t=1}^{T}(R_{t+1}+1)] - \iota \end{bmatrix}.$$

Conditional on the weighting matrix, W_T, the GMM-estimator is

$$b_1 = -[C_T' W_T C_T]^{-1}[C_T' W_T d_T]. \tag{8}$$

The weighting matrix is chosen to be a consistent estimate of the inverse of the spectral density at frequency zero of the orthogonal conditions, $S = \sum_{j=-\infty}^{\infty} E[f_t f_{t-j}']$. A popular estimate uses Bartlett weights (Newey and West [1987]):

$$\widehat{S} = 1/T \sum_{t=j+1}^{T} \sum_{j=1}^{\rho} (1 - \frac{j}{\rho + 1})[f_t f_{t-j}' + f_{t-j} f_t'] + 1/T \sum_{t=j}^{T} f_t f_t'.$$

Although the DeSantis 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 $\widehat{S}$. However, removing the mean of f_t and allowing for some serial correlation in $\widehat{S}$ 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 below, we examine 5 sub-statistics MV_i ($i = 1, \ldots, 5$). The properties of the different statistics can be summarized as follows:

Test Statistics

Statistic	Demeaned	ρ Correction	# 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

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 6 to 9 iterations in the GMM estimation, the weighting matrix changed very little.

Given our concern with small sample properties, we also investigate a Wald test, MV_W, as an alternative to the tests MV_i ($i = 1, \ldots, 5$). In the Wald test, the stochastic discount factor is a linear function of both the benchmark and test assets. The orthogonality conditions for the test are similar to equation (6), except that

$$C_T = \begin{bmatrix} 1/T\sum_{t=1}^{T} R_{t+1}(R'_{t+1} - E[R'_{t+1}]) & 0 \\ 0 & 1/T\sum_{t=1}^{T} R_{t+1}(R'_{t+1} - E[R'_{t+1}]) \end{bmatrix}$$
$$= \begin{bmatrix} \Sigma_R & 0 \\ 0 & \Sigma_R \end{bmatrix},$$

where Σ_R is the variance-covariance matrix of the returns.

As Appendix A shows, there is a simple, analytical expression for MV_W. Let A be a $2n_2 \times 2(n_1 + n_2)$ matrix that selects the elements of the $2(n_1 + n_2) \times 1$ vector β that are zero under the null. Let $Q_T = C_T' W_T C_T$. Then

$$MV_W = T(AC_T^{-1}d_T)'[AQ_T^{-1}A']^{-1}(AC_T^{-1}d_T). \quad (9)$$

Hence, computation of 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_{W_j} $(j = 1, 2)$. For $j = 2$, there is no serial correlation correction. For $j = 1$, we construct S using the optimal bandwidth procedure.

Appendix B describes in detail a Monte Carlo experiment that examines both the size and power of the test statistics MV_i $(i = 1, \ldots, 5)$, MV_{W_j} $(j = 1, 2)$. In our brief discussion here, denote the n_1 benchmark asset returns by R_t^1 and the n_2 test asset returns by R_t^2. The benchmark returns are assumed to follow a vectorautoregressive process with conditionally heteroscedastic errors. The test asset returns are then a linear function of the benchmark returns and an 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 for $[R_t^1, R_t^2]$, on which the mean variance spanning tests are performed.

To investigate the power of the 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 2 describes this index). The alternative is calibrated so that the variance of the test assets explained by the new benchmark is 20% higher than under the null. Details are provided in the Appendix.

The Monte Carlo results, reported in Appendix B, are striking. Employing a serial correlation correction according to the optimal bandwith scheme of Andrews [1991], but not de-meaning the orthogonality conditions, leads to dramatically superior size properties for MV_3. Although its power is weaker compared with some of the other tests, the relative loss in power is only significant for large GMM systems.

In general, both the size and power properties seem to 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. 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 statistic, with the caveat that the power may be weak when many test assets are included in the test.

3.3 Empirical Results

In this section, we test whether adding emerging market assets to a number of different benchmark portfolios significantly shifts the investment opportunity set. Specifically, we examine the hypothesis of mean-variance spanning by testing whether the frontier of benchmark and test assets intersects the frontier of benchmark assets at two distinct points. In addition, 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. The Sharpe Ratio measures the slope of the line from the riskless rate to the tangency portfolio on the efficient frontier, 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. For a particular rejection of mean-variance spanning, the change in the Sharpe Ratio measures the economic importance of the shift in the efficient frontier.

Using the monthly IFC Global Indices and Morgan Stanley Capital International (MSCI) indices for mature markets, DeSantis [1993] finds that the frontier of IFC indices along with the U.S., European or World market portfolio does not intersect the frontier of the U.S., European or World market benchmark, but that the mature markets do intersect the frontier of the benchmark. Similarly, Harvey [1994] rejects the hypothesis that the Global Indices and the mature markets intersect the frontier of all eighteen MSCI mature market indices. Both authors examine mean-variance intersection at only one point on the efficient frontier.

We compare the diversification benefits from emerging market closed-end funds with the benefits from holding the corresponding IFC Investable Indices.[10] As described in Section 2, 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.

Recall that our test of mean-variance spanning produces a chi-square statistic with de-

[10]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 a Morgan Stanley 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.

grees of freedom equal to two times the number of overidentifying restrictions when we impose the restriction that a set of test asset returns and benchmark returns is priced by the subset of benchmark returns. The following table summarizes the test assets we examine in Tables 3 and 5.[11] In all cases, we report the probability value for the mean-variance spanning test statistic MV_3, which incorporates a serial correlation correction but no demeaning of the sample moment conditions, and 2 stages in the GMM estimation. Where available we also report the probability value according to the empirical distribution under the null.

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

The first set of benchmark assets cousists of the FT-Actuaries World Index for the U.S., the U.K., Europe excluding the U.K., and the Pacific. The benchmark was chosen to represent the portfolio of a globally diversified investor in 1990, the start of the test period. Most combinations of individual country indices are spanned by this benchmark. The test period was chosen to maximize the number of funds included in the tests.

Table 3, Panel 1 presents results from mean-variance spanning tests where we restrict investors to hold only funds which have corresponding IFC Investable indices, and only one fund for a particular country or region if there are multiple such funds, to enhance the power

[11]Note that 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.

of our tests.[12] In our discussion, we will use 5% as the size of the test. Given that the power of our tests ranges between 13% and 73%, readers concerned about Type-2 errors may wish to employ a higher size. We fail to reject mean-variance spanning for the twelve individual U.S. emerging market funds. 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% 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.

It is possible that closed-end fund industry factors could affect the 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 which 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. This controls for the presence of closed-end fund-specific factors in the U.S. and the U.K. while providing a benchmark of mature market assets similar to the FT-Actuaries benchmark. It turns out that the closed-end fund benchmark spans the FT-Actuaries benchmark, but not vice versa.

Panel 1 of Table 3 reports test results using the control benchmark. We again fail to reject mean-variance spanning for the twelve U.S. funds and for the U.S. fund index. The rejection for the U.K. fund index and the U.S. and U.K. fund index together no longer holds. but we stil reject spanning for the seven individual U.K. emerging market funds using the chi-square and the empirical distribution under the null. Overall, the results from Panel 1 suggest that the individual U.K. emerging market funds together provide diversification benefits relative to the benchmark portfolios. Apparently, constructing an index of these funds changes the distribution of returns so that they no longer provide such benefits. Using

[12]Choosing alternate funds from the same market does not change any of the results.

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.[13] 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 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.

Table 3, Panel 2 presents results from mean-variance spanning tests for the IFC Investable Indices corresponding to the test funds. The U.S. Investables are the individual IFC Investable indices corresponding to the countries and regions covered by the U.S. funds. and the U.S. Investables Index is the equally-weighted index of those index returns. For both the FT-Actuaries and the control benchmark, we can reject strongly mean-variance spanning for the individual IFC indices corresponding to the U.S. and U.K. emerging market funds using the chi-square distibution under the null. When we construct their equally-weighted indices, we can still reject the spanning hypothesis at the five percent level. 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.

The economic impact of differences in performance for the U.S. and U.K. funds and the IFC Investable indices is represented by the change in the Sharpe Ratio that accompanies each test. For example, the rejection of mean-variance spanning for the individual U.K. funds and the control benchmark (Table 3, Panel 1) is associated with a Sharpe Ratio change of about 0.096. By contrast, the rejection for the corresponding IFC indices (Table 3, Panel

[13]The value-weighted indices are constructed using the value of individual fund assets in dollars in April. 1993 (see Table 2).

[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.

2) is associated with a Sharpe Ratio change of 0.28. This difference indicates that the slope of the line with intercept (approximately) at the mean of the conditionally riskless rate and tangent to the efficient frontier becomes three times more steep when the IFC indices are included with the benchmark than when the U.K. funds are included with the benchmark.

What explains the performance differences between U.S. and U.K. funds? First, 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. 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 for the seven asset case. 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 U.S. and the U.K.), we reject spanning for the U.K. funds but not for U.S funds. Figure 1 plots the mean-standard deviation frontiers of the U.S. and U.K. funds with common coverage (seven in each market), and the IFC Investable indices covered by the funds, with both benchmarks. While standard errors are associated with each of the frontiers, the frontier formed from the U.K. funds and the benchmark assets appears well to the left of the frontier formed from the U.S. funds.

A second possible explanation relates to differences in managerial skill. These differences could be reflected in the NAV compositions of two funds investing in the same market, or in the market timing abilities of different managers. Our framework is not suitable to investigate market timing ability, but we can address differences in fund portfolio risk exposure. Table 4 contrasts the portfolio exposure of U.S. and U.K. funds that invest in the same emerging market or region during the test period. The table presents the risk exposure of each fund's NAV to the local market index, the IFC Global index, and the IFC Investable index, as the coefficient from a multivariate regression of NAV returns on the returns of the indices.[15] The residual from a linear projection of the local market index and the IFC Global index return onto the IFC Investable index return represents the component of those returns uncorrelated with the IFC Investable index, $r^{L \cdot I}$ and $r^{G \cdot I}$, respectively. 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, while

[15]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.

for the U.K. fund JAVA, $\beta_{G \cdot I}$ is 1.226 and significant. However the sensitivity of the U.S. Indonesian funds to the IFC Investable is greater than that of JAVA, 0.454 and 0.481 versus 0.269. The index returns explain a large share of the variation in NAV returns for most funds, as indicated by $\overline{R^2}$. The results in Table 4 suggest that the superior diversification benefits from U.K. emerging market funds could arise from portfolio selection.

Table 5 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. Recall that we scale each index return in the benchmark and test assets by the unit vector and its *own* index of individual fund premium changes (price minus NAV returns), which doubles the number of overidentifying restrictions in the spanning test to $4n_2$. Interestingly, we reject spanning for the scaled U.S. fund index as well as the scaled index of U.K. funds, and for the U.S. and U.K. fund indices together, using the chi-square distribution. Apparently fund premium changes predict all fund returns in our sample and induce diversification gains for U.S. funds not present from holding only the funds themselves.

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, and they are robust to the currency in which we express returns. By contrast, the IFC investable indices corresponding to the U.K. *and* the U.S. funds provide unequivocal diversification benefits. While the U.S. funds appear to perform poorly compared with the smaller set of U.K. funds, the NAV composition of U.S. and U.K. funds investing in the same market seems to be quite different. The use of conditioning information leads to diversification benefits for the U.S. as well as the U.K. funds.

4 Market Integration

4.1 Price Segmentation

4.1.1 Framework

Despite recent capital market liberalizations, most emerging equity markets are not completely integrated in global capital markets. Investment restrictions of various forms persist and poor liquidity or political risk often serve to effectively segment emerging markets from world markets (see Bekaert [1994] for a detailed analysis). As Bekaert and Harvey [1994] stress, the degree to which a national capital market is integrated with world capital markets is very difficult to measure. Looking at investment restrictions is problematic because there are numerous types of restrictions with some being more important than others in different countries. In addition, statutory investment restrictions may not be binding.

Our data on country funds offer a unique opportunity to measure the effective degree of market integration. With closed-end funds we observe two prices for the same assets: fund price and NAV. All other things constant, binding investment restrictions will raise the price of a fund's shares relative to its net asset value by approximately the amount the marginal investor is willing to pay to avoid these restrictions. This intuition was used by Bonser-Neal et al. [1990], who tested whether announcements of changes in investment restrictions were related to changes in the price/NAV ratios of funds investing in those countries. They document significant decreases in fund premiums following announcements of investment liberalizations.

The main purpose of this section is to develop a measure of market integration that exploits the apparent price segmentation between funds and NAVs. The measure we propose can be interpreted as the reduction in expected return a representative world investor would be willing to suffer in order to obtain barriers-free exposure to emerging equity markets. We will compute this measure for six countries: Indonesia, Korea, the Philippines, Taiwan, Thailand and Turkey.

Our framework builds heavily on the rational asset pricing framework used to test for diversification benefits. As will become clear below, we are able to control for the closed-end fund puzzle in our computations. Domestic funds also trade at prices that differ from their net asset values and discounts vary both across funds and over time. Hence, country fund premiums and discounts cannot solely be due to investment restrictions. It is important to note that some researchers have claimed that the behavior of premiums on country funds defies rational explanations. Indeed, the empirical facts seem puzzling at first sight. Country fund returns show high positive correlation with the trading market (see Bailey and Lim [1992], Diwan et al. [1993] and Bekaert and Urias [1994b]), and Hardouvelis et al. [1994] find a large common component in the premiums on country funds. This is true despite the fact that the funds invest in equity markets whose returns show low correlation with U.S. returns. This has motivated Bodurtha et al. [1993] and Hardouvelis et al. [1993] to claim that country fund premiums are primarily driven by investor sentiment.

On the other hand, Diwan et al. [1992] and Eun et al. [1993] explore the effect of investment restrictions on country fund pricing in the context of static mean-variance models. When the underlying assets of the fund are not spanned by the world assets and cross-border arbitrage is not possible, country fund premiums can arise in these models. The main variables driving premiums are the different degrees of risk aversion in the local and world markets and the different portfolios against which systematic risks are measured in the world and local markets. Note that to prevent arbitrage, the local investors must also be prevented, for example, from freely short-selling U.S.-traded country funds and buying local stocks.[16]

Rather than using a fully specified model, we base our measure directly on the intertemporal relationship between stochastic discount factors and asset returns that underlies our spanning tests in Section 3. Consider a world in which arbitrage between country funds and the underlying assets is difficult and costly. If the markets are effectively segmented, world investors will price the country funds, whereas local investors will price the local assets. Both the fund and the underlying assets are claims to identical cash flows. Hence, the difference between local and global pricing can be completely described by the differences in the pricing kernels that price net asset values and country funds, respectively.

Time-variation in these discount factors can explain some of the salient features of country fund premiums described above. For example, a rise in U.S. interest rates may decrease discount factors in world capital markets, without much effect on discount factors in segmented markets. This in turn leads to lower fund prices for all country funds, without much of an impact on NAVs. Hence, one potential measure of market integration could be based on the difference between the stochastic discount factor that prices country funds in the trading market and the discount factors that price the assets of the funds, used by local investors.[17] Recall that the fundamental asset pricing equation (1) places no restrictions on the form of the discount factor. We assume that the stochastic discount factors are *linear* functions of a limited number of benchmark returns.[18] Unfortunately, we cannot identify

[16]Of course, since such arbitrage will likely involve costs and since there is no guarantee that country fund prices will actually return to their net asset value, it may not be truly risk free (see Tuckman and Vila [1993] for a formal model).

[17]This is in fact the basic idea behind the non-parametric measure of market integration developed by Chen and Knez [1994].

[18]Bansal et al. [1993] study the properties of non-linear pricing kernels in an international setting.

the discount factor that prices local assets. The net asset values of the country funds in our sample are not spanned by local indices. Portfolio managers clearly do not index the local indices in our dataset (see also Table 4). However, we can identify the discount factor (pricing kernel) that prices world assets, including country funds.

Define the discount factor pricing country fund returns to be

$$m_{t+1}^{T} = \alpha^{T} + (R_{t+1}^{BT} - E[R_{t+1}^{BT}])'\beta^{T}, \tag{10}$$

where R_{BT} denotes a benchmark of investable assets. We identify α^T as the mean of one divided by one plus the conditionally riskless rate. The conditionally risk free rate is measured as the one week dollar LIBOR rate. As benchmark assets, we use the closed-end fund control benchmark and a mimicking portfolio for the IFC Investable Composite index constructed from global funds trading in the U.S. and the U.K.[19] Hence, we control for the presence of possible closed-end fund-specific pricing factors by including closed-end fund returns in the benchmark assets for fund returns. The mimicking portfolio for the IFC Investable Composite is constructed by forming the weighted sum of returns on the global closed-end funds that spans the IFC index. These weights are computed using the fact that when mean-variance spanning holds, the test returns can be written as a linear function of the benchmark returns with the coefficients summing to one (see also Appendix A).

Ideally, this benchmark should span all individual country fund returns. Since there are a total of 16 funds investing in the six countries of interest, it is impossible to test this hypothesis in a meaningful way. Instead, we impose the weaker requirement that the benchmark prices six indices of the individual country funds for our six countries. We cannot reject that our benchmark spans the six NAV value-weighted indices of fund returns representing each country. The benchmark also spans the set of local index-mimicking portfolios of fund returns. These mimicking portfolios are constructed from the spanning weights for the IFC Investable index in each country using the individual NAV returns as benchmark assets. The resulting portfolio can be interpreted as that portfolio of individual country funds with exposure closest to the IFC Investable index.[20]

[19]The global emerging market funds used to span the Investable Composite are EMF, ABTST, BGEM and TMPLT. These funds may invest in any emerging market.

[20]This interpretation remains valid when mean-variance spanning is rejected, as it is for a number of countries. Note that we do reject mean-variance spanning of the equally-weighted index of fund returns for the six countries. Using the FT-Actuaries benchmark with the Investable Composite-mimicking portfolio

If the discount factor m^T_{t+1} correctly prices all traded assets in the U.K. and the U.S., all returns available to U.S. and U.K. investors should satisfy equation (1). That is, viewing the returns as payoffs, their prices should be one. Our integration measure arises from using the definition for covariance to rewrite equation (3) and solving for the expected return[21]:

$$E[R_{t+1}] = \frac{1}{\alpha^T} - \iota - \frac{Cov(m^T_{t+1}, R_{t+1})}{\alpha^T}. \tag{11}$$

When returns covary positively with the discount factor, expected returns are lower. We first compute the expected return for country fund returns, $E[R^P_{t+1}]$. Although our pricing kernel does not price NAV-returns, we also use (11) to compute the expected return for net asset values, $E[R^N_{t+1}]$. We denote the difference by $\Delta \equiv E[R^P_{t+1}] - E[R^N_{t+1}]$.

We expect Δ to be positive when investment restrictions are binding. If the market was fully open, net asset values and country fund returns would be priced similarly, and the difference in expected returns would be small. If investment restrictions are effectively binding and investors cannot freely access the local market except through the country fund, price segmentation arises and investors "give up" part of the diversification benefits of emerging market exposure. The statistic Δ measures the reduction in expected return investors are willing to "pay" to receive the original exposure.

The measure is not perfect for several reasons. First, identifying m^T_{t+1} as a function of benchmark returns is difficult. For example, when an emerging market is opened up through a country fund and its stock returns are not spanned by world assets, the investment opportunity set of world investors changes and so will their pricing kernels. Second, the measure is potentially affected by factors such as local market inefficiencies, foreign exchange risk, and political risk, that may affect the local and global pricing of the underlying assets so that investment barriers have to be interpreted very broadly. In fact, our framework of stochastic discount factors which price returns in equation (1) is strictly valid only in complete, frictionless markets (see, for example, Hansen and Jagannathan [1991] and Luttmer [1991]). Third, the measure is affected by the extent of diversification benefits the underlying assets offer. For example, it is conceivable that a fully segmented market only trades stocks that have perfectly correlated cash flows with companies in developed countries. In this case, Δ will probably be low or negative, despite the market being fully segmented. Implicitly, we

provides similar results.

[21]Cochrane [1994] provides a thorough development.

assume that local emerging market returns are not spanned by developed market returns. Of course, this is exactly the claim made by DeSantis [1993] and Harvey [1994]. In what follows we also show that net asset value returns for our six countries are not spanned by the benchmarks we used in Section 3.

4.1.2 Empirical Results

We begin by reporting simple mean-variance spanning tests for the six countries of interest, using the country fund benchmarks from Section 3. Table 6, Panel 1 reports tests for both net asset value (NAV) returns and country fund returns. For the NAV and the country fund returns, we form a value-weighted index of funds targeting each country. Therefore, each test has two degrees of freedom. Note first that only the net asset value returns for Turkey are spanned (marginally) by the benchmark. When the test employs the NAV returns for the individual Turkish funds rather than an index, we reject spanning at the 5% but not at the 1% level. Hence, the returns in the six countries do offer diversification benefits, conditional on their markets being open. More surprisingly, for five of the six countries, the country fund price returns do not offer diversification benefits.[22] The exception is the value-weighted portfolio of Korea funds. When individual funds are used, the Turkish and Thai funds also offer diversification benefits. Hence, for three countries, the differential pricing between country funds and the underlying assets has led to the disappearance of all diversification benefits. In our pricing framework, it is natural to expect some reduction in the diversification benefits of funds relative to net asset values because of time-variation in the stochastic discount factor, m_{t+1}^T, which is common to all funds. It is not clear why they should completely disappear for some countries but remain for others, although this may depend on the nature of investment restrictions.

We also report the weights on a mimicking portfolio of the benchmark assets for the price returns. As above, these weights are computed by exploiting the fact that when mean-variance spanning holds, the test returns can be written as a linear function of the benchmark returns with the coefficients summing to one (see also Appendix A). The mimicking portfolios for fund returns place large weights on mature market closed-end funds.[23]

Results on expected returns and the market integration measure, Δ, are reported in Table

[22] These results are similar using the FT-Actuaries benchmark.

[23] For the FT-Actuaries benchmark, the Pacific Index receives a large weight for Thailand, Taiwan and Turkey.

6, Panel 2 for each country. We identify m_{t+1}^T as the discount factor which prices the six value-weighted indices of fund returns, using the closed-end fund benchmark described in Section 4.1 (recall equation (10) and the ensuing discussion). We also computed the market integration measure for each country using the m_{t+1}^T that prices the six IFC Investable-mimicking fund portfolios, and the results are very similar. The Δ measure is computed for value-weighted indices of fund returns targeting each country.

The segmentation measure Δ is relatively high (greater than 5%) for three countries (Indonesia, Thailand, and Taiwan), and low or negative for the three remaining countries in the analysis (Korea, Philippines and Turkey). As an aid in interpreting the results. Table 7 lists some characteristics of the six stock markets. Table 7 illustrates the difficulty in gauging the degree of market integration from general market characteristics and the regulatory framework of different countries. For example, Korea, Taiwan and Thailand are relatively large, liquid markets with low political risk, stable exchange rate policies and numerous possibilities for indirect access by foreigners using country funds and American Depositary Receipts (ADRs). However, there are foreign ownership restrictions in all three countries. Thailand maintains two separate listings for common stocks which have reached foreign ownership limits, one for locals (the "Main Board") and one for foreigners (the "Alien Board"). Prices on the Alien Board are typically higher than prices on the Main Board (see Bailey and Jagtiani [1994]). Our finding of $\Delta = 5.75\%$ for Thailand confirms the findings in Bailey and Jagtiani who document the effectiveness of foreign ownership restrictions in the Thai capital markets.

Taiwan also has strict limits on foreign ownership. In January 1991, Qualified Foreign Institutional Investors (QFIIs) were allowed to invest directly in Taiwan's stock market. In the aggregate, foreign investment in each of the listed companies is limited to 10% of the shares outstanding, and QFIIs can buy no more than 5% of the shares. In addition, there is a ceiling on total foreign investment which was recently raised to $10 billion (in March 1993). The value of 6.60% for Δ confirms the effectiveness of these restrictions. The surprising result is for Korea, where Δ is only 1.41%. Regulations on foreign participation prohibited direct access to the Korean market until January 1992. Foreign ownership is limited to 10% in so-called unlimited industries and 8% in limited industries (which includes communications and defense). Recently, the 10% ceiling was raised to 25% for 45 firms which hit the 10% cap. However, the low number may reflect the many country funds available to investors. Bekaert and Harvey [1994] also find that the Korean market appears to be integrated with global capital markets. The fact that both Korea and Taiwan underwent capital market liberalizations during the sample may affect the results and we address this issue in Section 4.2.

From Table 7, Indonesia, the Philippines and Turkey appear less developed, less liquid, and more prone to political and currency risk than the preceding three markets. In addition, they have fewer country funds and ADR programs. On the other hand, none of these countries appears to have binding foreign ownership restrictions. Our segmentation measure implies that Indonesia is the most segmented of the six markets. It is also the only country for which Δ is significantly different from zero. Since 1988, foreigners have been able to buy up to 49% of shares in listed companies. However, certain blue chip shares frequently reach that limit and foreign investors are forced to pay a premium for these shares. Investors are only willing to forego about 1% in expected return to obtain barriers-free exposure to Turkey. For the Philippines, Δ is actually negative. The results for Turkey and the Philippines may reflect the fact that these markets are effectively integrated with global markets[24], or it could reflect idiosyncratic features of the funds. As an example of the latter, Diwan et al. [1993] find that the U.S.-traded Turkish Investment Fund (TKF) held a large share of its portfolio in U.S. Treasury bills early after inception, a period included in our sample.

4.2 Investment Liberalizations and Mean-Variance Spanning

It has sometimes been argued that the movement toward market integration will reduce the diversification benefits of emerging markets. While our sample is too short to allow a fully dynamic analysis, for four heavily restricted markets a significant investment liberalization takes place in the middle of the test period. An investment liberalization can be viewed as a move toward market integration. 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 they may be after the liberalization. To test this conjecture, we modify the specification of the discount factor

[24] Although foreign investors were given complete access to shares in the Philippines with the Foreign Investment Act of November 1991, some companies have subsequently been given foreign ownership limits. The restricted shares now trade at substantial premiums to local shares.

as follows. Recall that

$$m^{\alpha_i}_{t+1} = \alpha_i + \beta_{1i}\left(R^1_{t+1} - E[R^1_{t+1}]\right) + \beta_{2i}\left(R^2_{t+1} - E[R^2_{t+1}]\right), \tag{12}$$

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 $\beta_{2i} = 0$ ($i = 1, 2$), we redefine $\beta_{2i} \equiv \widetilde{\beta_{2i}}\, D_{t+1}$, where D_{t+1} is equal to 1 before the liberalization and 0 afterwards. If the hypothesis $\widetilde{\beta_{2i}} = 0$ ($i = 1, 2$) cannot be rejected, the benchmark spans the test assets before *and* after the liberalization. On the other hand, if $\widetilde{\beta_{2i}} \neq 0$ ($i = 1, 2$), then the test assets are spanned by the benchmark assets after the liberalization but not before.

The results for the investment liberalizations are contained in Table 8. The table considers funds investing in four countries: Brazil, India, Korea and Taiwan. In the first column, the table reports a Wald statistic for the test that $\widetilde{\beta_{2i}} = 0$ ($i = 1, 2$), estimated from the GMM system. In all four cases we cannot reject the hypothesis that $\widetilde{\beta_{2i}} = 0$ ($i = 1, 2$) 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 6 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.

5 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 study we examine the diversification benefits from holding closed-end country funds which 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, while the IFC Investables are attainable in theory but ignore all effective investment costs or restrictions with the exception of foreign ownership restrictions.

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 tests have fairly low power against one interesting alternative for large GMM systems, but the power is about 70% 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.

Our analysis of price segmentation and diversification benefits suggests that the differential pricing of fund shares and NAVs has contributed to the disappearance of diversification benefits for half of the countries we examine. Despite the existence of substantial premiums on Korean funds, they continue to provide strongly significant benefits. We propose a measure of market segmentation which captures the notion that investors may be willing to forego expected return for direct access to a fund's NAV instead of the fund's shares. The measure suggests that barriers to investment, broadly interpreted, are especially effective in Indonesia, Taiwan, and Thailand, but not in Korea, Turkey and the Philippines. Finally, the opening of Taiwan's stock market to foreign investment in January, 1991, had a significant negative impact on the diversification benefits from holding its country funds.

The power and small sample properties of our mean-variance spanning tests deserve further study. In particular, the mean-variance spanning hypothesis is closely related to the mean-variance efficiency hypothesis used in tests of single or multifactor asset pricing models. Our results suggest that tests using large GMM systems may have considerable size distortions and poor power properties.

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Table 1

Emerging Market Funds: Total Net Assets

Country or Region	Closed-End	Open-End	Total
	End 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
	End 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

Notes:

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**.

Table 2

Country Fund Sample General Characteristics:

January 1986-August 1993

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†	FEMK	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†	GEF	900119	112.9	-14.665	3.822	0.850	0.657
EFM Java	JAVA	900518	15.7	-5.361	4.241	1.328	0.624
Korea Liberalisation	KLF	900615	34.5	-20.012	5.710	1.424	0.679
Korea Europe*	KORFUR	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 2

Country Fund Sample General Characteristics:

January 1986-August 1993

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
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 (Table 2):

All U.S. funds trade on the NYSE, except for IAF and IBF, which trade on the 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: $P_t \equiv NAV_t PRM_t$. **Premium** is 100 times the average of $(PRM_t - 1)$.

$\sigma(r_t^P)$ 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: $\Delta \log(PRM_t) = r_t^P - r_t^N$.

VR1 is the variance ratio $\sigma^2(\Delta \log(PRM_t))/\sigma^2(r_t^P)$, 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 2 for details.

Table 3

Unconditional Mean-Variance Spanning Tests

September 1990 - August 1993

Panel 1: U.S. and U.K. Emerging Market Funds						
	Test Assets: p-value MV_3 (p-value Empirical Distribution) [Change in 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
U.S., U.K., Europe, Pacific	0.6743 (0.7910) [0.1618]	0.0109 (0.0280) [0.1294]	0.4178 [0.2405]	0.1215 (0.1280) [0.0680]	0.0325 (0.0395) [0.0564]	0.0213 (0.0255) [0.0875]
U.S. Domestic Index, U.K. Domestic Index, U.S. Mature Market Index, U.K. Mature Market Index	0.6015 (0.7280) [0.1311]	0.0002 (0.0010) [0.0958]	0.5382 [0.1988]	0.1001 (0.1150) [0.0473]	0.2721 (0.3090) [0.0246]	0.1032 (0.1150) [0.0578]

Panel 2: IFC Investable Indices				
Benchmark Assets	U.S. Investables	U.K. Investables	U.S Investables Index	U.K. Investables Index
U.S., U.K., Europe, Pacific	0.0006 [0.4126]	0.0030 [0.3051]	0.0007 [0.0940]	0.0016 [0.0239]
U.S. Domestic Index, U.K. Domestic Index, U.S. Mature Market Index, U.K. Mature Market Index	0.0080 [0.3732]	0.0014 [0.2801]	0.0078 [0.0584]	0.0011 [0.0117]

Notes:

The p-value is the probability value of the chi-square statistic MV_3. Numbers in parentheses are approximate probability values from the empirical distribution of MV_3 under the null. See Section 3.2 or the Appendix for details. Numbers in square 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.

Notes (Table 3):

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.

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 4
Country Fund Portfolio Risk Characteristics: Common U.S. and U.K. Emerging Market Coverage

October 1990-August 1993

Country/Region	Fund	Domicile	α	$\beta_{L \cdot I}$	$\beta_{G \cdot I}$	β_I	$\overline{R}^2$
Asia-Pacific	APB	U.S.	0.002		-0.019	0.622	0.385
			(0.001)		(0.068)	(0.055)	
	EFMDR	U.K.	0.001		-0.126	0.626	0.354
			(0.001)		(0.085)	(0.078)	
	GEP	U.K.	0.002		-0.004	0.639	0.353
			(0.001)		(0.074)	(0.073)	
	PHIT	U.K.	0.001		0.129	0.559	0.252
			(0.001)		(0.072)	(0.073)	
	SAF	U.S.	0.000		0.455	0.617	0.394
			(0.002)		(0.166)	(0.097)	
	TAEMIT	U.K.	0.000		0.042	0.610	0.174
			(0.002)		(0.097)	(0.080)	
Global	ABTST	U.K.	0.002		0.270	0.435	0.151
			(0.002)		(0.111)	(0.092)	
	BGEM	U.K.	0.001		0.223	0.454	0.201
			(0.001)		(0.093)	(0.059)	
	EMF	U.S.	0.004		0.039	0.708	0.509
			(0.001)		(0.059)	(0.047)	
	TMPLT	U.K.	0.004		-0.026	0.388	0.092
			(0.002)		(0.141)	(0.078)	
Indonesia	IF	U.S.	-0.001	0.331	0.009	0.454	0.586
			(0.001)	(0.131)	(0.172)	(0.030)	
	JAVA	U.K.	-0.000	-0.046	1.226	0.269	0.168
			(0.002)	(0.419)	(0.412)	(0.083)	
	JGF	U.S.	-0.000	0.532	0.114	0.481	0.694
			(0.001)	(0.163)	(0.172)	(0.029)	
Latin America	LAM	U.S.	0.003		0.048	0.536	0.245
			(0.001)		(0.064)	(0.088)	
	LATIT	U.K.	-0.001		0.065	0.798	0.404
			(0.001)		(0.050)	(0.076)	
Philippines	FPF	U.S.	0.003	-0.009	0.408	0.241	0.417
			(0.001)	(0.058)	(0.115)	(0.040)	

Table 4

Country Fund Portfolio Risk Characteristics:

Common U.S. and U.K. Emerging Market Coverage

October 1990-August 1993

Country/Region	Fund	Domicile	α	$\beta_{L\cdot I}$	$\beta_{G\cdot I}$	β_I	$\overline{R}^2$
Philippines	FPT	U.K.	0.002	0.162	0.895	0.414	0.330
			(0.003)	(0.184)	(0.318)	(0.089)	
South Korea	KF	U.S.	0.001	0.796		0.439	0.595
			(0.001)	(0.111)		(0.057)	
	KLF	U.K.	-0.003	0.189		0.327	0.067
			(0.003)	(0.307)		(0.152)	
	KOREUR	U.K.	0.001	0.922		0.476	0.463
			(0.002)	(0.143)		(0.050)	
Thailand	ABTHAI	U.K.	0.000	0.331	0.552	0.755	0.609
			(0.002)	(0.173)	(0.362)	(0.043)	
	SIAM	U.S.	0.001	0.064	-0.668	0.463	0.268
			(0.002)	(0.306)	(0.647)	(0.091)	
	TC	U.S.	0.000	0.282	1.150	0.761	0.790
			(0.001)	(0.134)	(0.508)	(0.033)	
	THF	U.S.	0.000	0.170	0.908	0.815	0.909
			(0.001)	(0.087)	(0.259)	(0.022)	
Turkey	TKF	U.S.	0.002	0.523	0.138	0.799	0.806
			(0.002)	(0.125)	(0.360)	(0.035)	
	TURK	U.K.	0.003	-0.026	0.719	0.258	0.109
			(0.003)	(0.125)	(0.375)	(0.080)	

Notes:

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

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

The dependent variable is the return on NAV for the fund. $r^{L\cdot I}$ 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 spaces 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}. $\overline{R}^2$ is the R^2 adjusted for degrees of freedom.

Table 5

Conditional Mean-Variance Spanning Tests

September 1990 - August 1993

	Test Assets: p-value MV_3 [Change in Sharpe Ratio]		
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 [0.0380]	0.0000 [0.0270]	0.0002 [0.0514]

Notes:

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 $4n_2$.

The p-value is the probability value of the chi-square statistic MV_3. See Section 3.2 or the Appendix for details. Numbers in square 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 the indices are equally-weighted. The tests employ 152 weekly observations

Table 6

Market Integration

Panel 1: Mean-Variance Spanning						
	Indonesia	Korea	Philippines	Taiwan	Thailand	Turkey
Return	*p*-value MV_3 [Change in Sharpe Ratio]					
Price	0.9847	0.0078	0.1141	0.9224	0.8877	0.4441
	[0.0000]	[0.0014]	[0.0409]	[0.0003]	[0.0034]	[0.0014]
NAV	0.0000	0.0045	0.0004	0.0100	0.0070	0.0724
	[0.0056]	[0.0003]	[0.0331]	[0.0000]	[0.0141]	[0.0031]
Benchmark	**Weights for Price Returns**					
U.S. Domestic Index	0.3973	0.2912	0.3755	0.2623	0.4625	-0.0530
U.K. Domestic Index	0.0196	0.1539	0.1857	-0.1821	-0.7293	-0.2784
U.S. Mature Market Index	0.2529	0.3660	0.1887	0.4840	0.4779	0.7119
U.K. Mature Market Index	0.3302	0.1888	0.2501	0.4357	0.7888	0.6195
Panel 2: Cost of Barriers to Investment						
	Indonesia	Korea	Philippines	Taiwan	Thailand	Turkey
	(Standard Error)					
$E[R^P_{t+1}]$	18.3935	11.8572	12.0525	19.9329	15.9425	11.8455
	(8.4980)	(7.8532)	(7.7457)	(10.2335)	(9.5493)	(10.3070)
$E[R^N_{t+1}]$	1.0870	10.4492	13.0712	13.3198	10.1680	10.7808
	(2.9520)	(5.5802)	(4.5110)	(5.2301)	(6.9726)	(8.1578)
Δ	17.3065	1.4080	-1.0187	6.6131	5.7745	1.0646
	(7.3065)	(5.6596)	(5.3308)	(8.0604)	(5.3666)	(7.9335)
$Mean(R^P_{t+1})$	18.4450	14.7772	37.3945	21.7580	22.7877	9.1692
$Mean(R^N_{t+1})$	-1.9817	5.8931	23.7340	10.2454	21.0070	0.6624

Notes:

Panel 1

The *p*-value is the probability value of the chi-square statistic MV_3. See Section 3.2 or the Appendix for details.

The test assets are a value-weighted price or value-weighted NAV index return, and the control benchmark is used. Numbers in square brackets indicate the approximate change in the Sharpe Ratio from adding the country index to the benchmark assets. Spanning weights (that sum to one) are given for the tests using the index of fund price returns.

Notes (Table 6):

Panel 2

$E[R^P_{t+1}]$ and $E[R^N_{t+1}]$ are computed according to equation (11) in the text using the discount factor that prices the six value-weighted country fund indices. Δ is the difference between expected price and NAV returns. It represents the cost, in terms of a reduction in expected return, that investors would be willing to accept in exchange for direct access to the assets of the funds.

To compute standard errors for Δ and the expected returns, we first derived the GMM variance-covariance matrices for b_{11} (see Section 3.2) and the covariance between benchmark returns and price and NAV returns for the six countries. Denote the latter moment estimates as COV. Since the expected returns and Δ are a function of b_{11} and COV, the mean value theorem was applied to derive their standard errors. We assumed a block-diagonal structure for the variance-covariance matrix of b_{11} and COV.

$Mean(\cdot)$ denotes the time series average of returns.

All expected return measures are multiplied by 5200.

Table 7

Stock Market Characteristics and Investment Restrictions

Measure	Indonesia	Korea	Philippines	Taiwan	Thailand	Turkey
Market Cap	14385.0	94289.7	14077.1	124818.0	48252.4	11291.1
Market Cap/ GDP	0.09	0.25	0.17	0.32	0.35	0.12
Turnover	N.A.	123.19	22.19	192.80	149.25	72.68
Percent Investable	47.6	9.6	47.3	3.0	27.0	97.3
Exchange Rate Regime	Managed Float	Pegged to Dollar	Free Float	Managed Float	Pegged to Basket	Free Float
# Country Funds	12	16	5	10	17	2
# ADR Programs	1	6	7	5	5	3
Credit Rating	50.5	67.6	25.2	77.5	61.3	43.9
Inflation Variability	N.A.	6.3	9.0	N.A.	5.8	28.0

Notes:

Market Cap is the average dollar market capitalization of the local market index for each country, in millions (see the Data Appendix for specific indices). **Turnover** is the annual value of traded local stocks divided by **Market Cap**. **GDP** data are taken from the *IMF International Financial Statistics*, except for Taiwan. For Taiwan, GNP obtained from its Central Bank Annual Report is substituted for **GDP**. The **Market Cap**, **Turnover** and **GDP** data are measured in 1992, approximately the middle of our sample, except for Indonesia's **Market Cap**, which is measured in March 1993.

Percent Investable is the ratio of the IFC Investable index, which controls for foreign ownership restrictions, and the IFC Global index. Figures are taken from the *IFC Index Methodology* in March, 1993. Section 2 describes the indices.

The **Exchange Rate Regime** is taken from the *1994 IMF Annual Report on Exchange Agreements and Exchange Restrictions*.

 # Country Funds and **# ADR Programs** is taken from Table 6 in Bekaert, Garcia and Harvey [1994]. Only publically-traded country funds focusing on the target market are included.

Notes (Table 7):

The **Credit Rating** is taken from *Institutional Investor*, March 1992. A value of 100 represents a perfect credit rating.

Inflation Variability is the standard deviation of monthly rates, taken from the *IMF International Financial Statistics* for the period 1986-1992, multiplied by 1200.

N.A. indicates the number is not available.

Table 8

Mean-Variance Spanning and Investment Liberalizations

December 1989 - August 1993

		p-value [Change in Sharpe Ratio]	
Test Assets	$[\widetilde{\beta_{21}}, \widetilde{\beta_{22}}] = 0$	Before Liberalized	After Liberalized
BZF	0.6512	0.6631	0.7563
		[0.0032]	[0.0021]
IGF	0.2719	0.2509	0.9731
		[0.0009]	[0.0008]
KF	0.3229	0.2722	0.0176
		[0.0242]	[0.0002]
TWN & ROC Index	0.1116	0.0082	0.5085
		[0.0040]	[0.0003]

Notes:

Information on restricted markets was obtained from Bekaert [1994] 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. 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. India opened to foreign investors in November, 1992, with a 25 percent cap on holdings by registered foreign institutions. A variety of overriding restrictions apply in all cases. The test period begins in December 1989.

The hypothesis, $[\widetilde{\beta_{21}}, \widetilde{\beta_{22}}] = 0$, 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 the probability value. See Section 3.2 or the Appendix for details. Numbers in square 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 benchmark assets are the closed-end fund mature market and domestic indices.

Appendix

A A GMM-Based Wald Test of Mean-Variance Spanning

Huberman and Kandel [1987] define mean-variance spanning in a linear regression model. Let R_t^1 denote a vector of benchmark asset returns. To test whether these returns span the vector of returns $R_t = [R_t^{1\prime}, R_t^{2\prime}]'$, consider the linear regression:

$$R_t^2 = \alpha + BR_t^1 + u_t. \tag{1}$$

Huberman and Kandel show that testing for spanning of R_t by R_t^1 is equivalent to testing:

$$\alpha = 0 \qquad B\iota = \iota, \tag{2}$$

with i a vector of ones.

In this appendix, we will first write the De Santis test in an alternative way to prepare for the derivation of the Wald test. Since these GMM-based tests seem very different from the standard Huberman and Kandel [1987] tests, we then demonstrate how the Wald test and the Huberman and Kandel test are tests of particular restrictions on the first and second moments of $R_t = [R_t^{1\prime}, R_t^{2\prime}]'$. A similar derivation can be found in Ferson [1993].

The orthogonality conditions for the De Santis test are given in equation (7) in the text. As above, let $R_t = [R_t^{1\prime}, R_t^{2\prime}]'$ and let the dimension of $R_t^1(R_t^2)$ be $n_1(n_2) \times 1$. Denote $n = n_1 + n_2$. Let β be a $2n \times 1$ vector. The orthogonality conditions for the Wald test can be written as:

$$h_T(\beta) = C_T\beta + d_T \tag{3}$$

where the dimensionality of h_T is $2n \times 1$ and C_T is a square matrix. Hence,

$$\beta = -C_T^{-1}d_T. \tag{4}$$

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 weigthing matrix in the GMM-system, W_T, is chosen optimally (Hansen [1982]), the variance-covariance matrix of β equals $(C_T W_T C_T')^{-1}$. The expression for the Wald test in equation (9) in the text follows straightforwardly.

The De Santis test is nothing more than a likelihood ratio-type test which requires estimation under the null. It is useful to think about the De Santis test as arising from a restricted estimation. Let $\tilde{\beta}$ denote the restricted estimator of β (i.e., with the zero restriction on the coefficients of the test assets imposed). Then,

$$\begin{aligned} \tilde{\beta} &= \operatorname{argmin} h_T(\beta)' W_T\, h_T(\beta) \\ & \qquad s.t. \qquad A\,\beta = 0 \\ h_T(\beta) &= C_T\,\beta + d_T. \end{aligned} \tag{5}$$

Note that the orthogonality conditions are linear in the parameters and so is the constraint. Consequently, Proposition 4 in Newey and West [1986] implies that the Wald test and the likelihood ratio-type test are numerically equivalent. Denote the De Santis test statistic by L. It is given by:

$$L = T\, h_T(\tilde{\beta})'\, W_T\, h_T(\tilde{\beta}). \tag{6}$$

Of course, the equivalence relies on the use of the same weighting matrix. As our Monte Carlo results show, the choice of the weighting matrix is one important determinant of the small sample properties of the test statistics.

Finally, let's consider the moment restrictions implied by mean-variance spanning tests. We will first rewrite the spanning restrictions derived by Huberman and Kandel [1987], reproduced in equation (2). Let μ_1 denote $E[R_t^1]$ and let μ_2 denote $E[R_t^2]$. Using least squares algebra, the restrictions in equation (2) are equivalent to:

$$\begin{aligned} \mu_2 - \Sigma_R^{12}\left(\Sigma_R^{11}\right)^{-1}\mu_1 &= 0 \\ \Sigma_R^{12}\left(\Sigma_R^{11}\right)^{-1}\iota &= \iota, \end{aligned} \tag{7}$$

where $\sum_R$ is the covariance matrix of the set of returns $R_t = [R_t^{1\prime}, R_t^{2\prime}]'$, partitioned as:

$$\Sigma_R = \begin{bmatrix} \Sigma_R^{11} & \Sigma_R^{12} \\ \Sigma_R^{21} & \Sigma_R^{22} \end{bmatrix}. \tag{8}$$

The Wald test, on the other hand, implies:

$$A\,\hat{b} = 0 = -A\,C_T^{-1}\,d_T. \tag{9}$$

Recall the expressions for C_T and d_T (see equation (7) in the text):

$$\begin{aligned} C_T &= \begin{bmatrix} \Sigma_R & 0 \\ 0 & \Sigma_R \end{bmatrix} \\ d_T &= \begin{bmatrix} \alpha_1(\mu + \iota) - \iota \\ \alpha_2(\mu + \iota) - \iota \end{bmatrix}, \end{aligned} \tag{10}$$

where $\mu = [\mu_1', \mu_2']'$.

Using (10) and the partitioned inverse formula in (9), we find:

$$\Sigma_R^{21}\left(\Sigma_R^{11}\right)^{-1}[(\mu_1 + \iota)\alpha_j - \iota] = (\mu_2 + \iota)\alpha_j - \iota$$

for $j = 1, 2$.

This can only be true for both $j = 1$ and $j = 2$ if the Huberman-Kandel restrictions proposed in equation (7) hold.

B Monte Carlo Experiments

The Date Generating Process (DGP) for the returns $[R_t^1, R_t^2]$ under the null of mean variance spanning is assumed to be:

$$\begin{aligned} R_t^1 &= \mu + A\,R_{t-1}^1 + \epsilon_t \\ R_t^2 &= B\,R_t^1 + u_t \\ &\qquad s.t. \quad B\iota = \iota. \end{aligned} \tag{1}$$

The restriction $B\iota = \iota$ imposes mean variance spanning (see Appendix A). Let

$$\begin{aligned} \epsilon_t &\sim N(0, \Sigma_t) \\ u_t &\sim N(0, \Sigma_u), \end{aligned} \tag{2}$$

where $\sum_t$ follows a constrained correlation GARCH-model (Bollerslev [1991]), or

$$\begin{aligned} \Sigma_t^{ii} &= w_i + \alpha_i(\epsilon_{t-1}^i)^2 + \beta_i \Sigma_{t-1}^{ii} \\ &\qquad i = 1 \ldots, n_1 \\ \Sigma_t^{ij} &= \rho_{ij} \Sigma_t^{ii} \Sigma_t^{jj}, \quad i \neq j, \end{aligned} \tag{3}$$

where the superscripts indicate the position in the matrix or vector.

The parameter vector $[\mu', vec(A)', vec(B)', vech(\Sigma_u)']'$ is estimated from our data set using GMM. We use the residuals ϵ_t to estimate the multivariate GARCH model in equation (3). In the estimation of (3), we constrain w_i to match the unconditional sample variance of ϵ_t^i. i.e., we let:

$$w_i = \frac{\frac{1}{T}\Sigma(\epsilon_t^i)^2}{1 - \alpha_i - \beta_i}. \tag{4}$$

We consider two sets of four benchmark returns (see Section 3.3). Hence, the number of parameters to be estimated in the multivariate GARCH process becomes 14 under the null. When no ARCH effects were found for residual i, the estimation was redone constraining $\alpha_i = \beta_i = 0$. So the returns $[R_t^1, R_t^2]$ have time-varying conditional means and variances. but the mean and variance dynamics for R_t^2 are completely driven by R_t^1. Estimation was performed using the MAXLIK package in GAUSS.

Under the null of mean variance spanning, all test statistics have a χ^2 distribution with $2n_2$ degrees of freedom. The choice of n_2 is guided by our empirical work. We set $n_2 = 1$, $n_2 = 2$, $n_2 = 7$ and $n_2 = 12$, to correspond to the dimension of the test assets described in Section 3.3 of the text.[1]

Table A1 reports the percentage of the 2,000 experiments constructed under the null hypothesis for which the values of the test statistics are greater than the five percent critical values of the null distribution. A standard error for the empirical size can be computed as

$$\left[\frac{(0.05)(0.95)}{2,000}\right]^{1/2} = 0.0049.$$

The table shows that the size properties of the statistics are very different. First of all, de-meaning the orthogonality conditions (compare MV_2 to MV_1 and MV_4 to MV_3) invariably worsens the size distortion. Second, a serial correlation correction leads to an empirical size close to 5%, only if the orthogonality conditions are not de-meaned. That is, MV_3 performs better than MV_1 but MV_2 performs better than MV_4. Note that by using the optimal bandwidth scheme of Andrews [1991], the serial correlation correction may not always be applied. The use of a fixed number of Newey-West [1987] lags would probably lead to inferior size properties. Iterating on the weighting matrix marginally improves the size properties of the test. The Wald statistic with no serial correlation correction, MV_{W2}, performs better than MV_{W1}, but the size properties are poor when compared with MV_3.

For $n_2 = 1$ and $n_2 = 2$, all test statistics perform reasonably well, and the empirical sizes of MV_1, MV_3 and MV_{W2} are within two standard error bands around the nominal size. For $n_2 = 7$, the size distortions become problematic for most test statistics and the superior small sample behavior of MV_3 stands out. We conclude that MV_3, the test statistic with a serial correlation correction and the orthogonality conditions not de-meaned, gives the least small sample distortion.

We also examine the power of the tests. In order to do so, we need to formulate a DGP under an alternative hypothesis. To make the alternative easy to interpret, we let the benchmark assets follow the same process as under the null. The new factor is assumed independent of the old factors and follows a univariate AR(1) process. Its error terms are

[1] We did not perform the Monte Carlo analysis for $n_2 = 19$. As will become clear, the test statistics are likely to be ill-behaved for such a large number of test assets.

normally distributed with the conditional variance following a GARCH (1,1) process. To estimate parameters for this process, we use the IFC Investable Composite index return as a proxy for emerging market-specific risk (Section 2 describes this index).

Finally, we have to determine the new matrix B. For each return in R_t^2, the corresponding row in the B matrix can be interpreted as the portfolios weights on the different benchmark assets that span it. Indeed, the weights sum to one but they need not be positive. For each test asset, we choose the new weights starting from the matrix B under the null, but redistribute some of each weight (say y) from the original benchmark assets to the new benchmark asset. Hence, the weight on the new benchmark asset becomes $4y$, since there are four benchmark assets under the null. The redistribution is such that the proportion of test asset variance explained by the new benchmark is 20% higher than the proportion explained by the benchmark under the null. Note that this is not a particularly strong alternative. If the R^2 were 30% under the null, our alternative implies an increase in R^2 to 36%. In the actual data, including the IFC Investable Composite index in a regression of country funds indices onto the null benchmark assets increases the R^2 by 40% rather than 20%.

In order to prevent the implied weights on the new factor from becoming too large, we had to scale up its variance by a factor of 3. The resulting variance is of the same magnitude as that of the individual country fund with the lowest sample variance in our sample. The implied weights on the new benchmark asset vary between 0.26 and 0.66.

To summarize the DGP under the alternative, let R_t^e denote the new benchmark return, and define $\overline{R_t^1} = [R_t^{1\prime}, R_t^e]'$. Then,

$$\begin{aligned} \overline{R_t^1} &= \overline{\mu} + \overline{A}\,\overline{R_{t-1}^1} + \overline{\epsilon_t} \\ \overline{R_t^2} &= \overline{B}\,\overline{R_t^1} + \overline{u_t} \\ & \qquad s.t. \quad \overline{B}\iota = \iota, \end{aligned} \tag{5}$$

with

$$\overline{\mu} = \begin{bmatrix} \mu \\ \mu^e \end{bmatrix} \quad \overline{A} = \begin{bmatrix} A & 0 \\ 0 & a^e \end{bmatrix} \quad \overline{\Sigma_\epsilon} = \begin{bmatrix} \Sigma_\epsilon & 0 \\ 0 & (\sigma^e)^2 \end{bmatrix}.$$

The way we construct $\overline{B}$ implies that $\overline{B} = [B - b/4, b]$, where b is a $n_2 \times 1$ vector containing the weights on the new factor. It turns out that in order to satisfy the explained variance criterion, each element in b must satisfy a quadratic equation with two real roots. We take $b_i, \forall i$ to be the positive root of this equation.[2]

The power of the test statistics hovers around 70% for $n_2 = 1$ and $n_2 = 2$, and drops below 50% for most test statistics when $n_2 \geq 7$. For $n_2 = 12$, the null and alternative distributions are quite close to each other. The MV_3 statistic – the test with the best small size properties – has the worst power for $n_2 = 12$, but not for smaller n_2. The Wald tests seem to have poor power properties in general compared with the likelihood ratio-type tests. To maximize power, orthogonality conditions should be de-meaned but serial correlation corrections should be avoided as is demonstrated by the superior performance of MV_2.

In general, both the size and power properties seem to 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. Unfortunately, GMM-systems with such low saturation ratios are common in the empirical finance literature. To conclude, MV_3 has significant superior size properties relative to the other test statistics. Although it has poor power for $n_2 = 12$, its power does not differ substantially from that of the other statistics for smaller systems.[3] Therefore, we use MV_3 for all our tests with the caveat that all the tests perform poorly when the saturation ratio is low.

[2] Another potentially interesting alternative would let $\overline{u_t}$ be serially correlated but keep the number of benchmark assets intact. We intend to explore the power of our tests against such an alternative in the future. We also plan to investigate the small sample properties of alternative mean variance spanning tests (Huberman and Kandel [1987], Jobson and Korkie [1989], and Ferson et al. [1993] are examples).

[3] In weighing the relative advantages and disadvantages of the various test statistics, power should be the more important consideration, especially since our alternative (exposure to an 'emerging market factor') is an economically interesting one. See McCloskey [1985] for a critique on the misuse of "significance tests" and their lack of power, and other considerations in hypothesis testing by economists.

C Data Appendix

For each country fund, the sample contains weekly price, NAV, exchange rate, dividend, and volume information for the life of the fund. For the U.K. trusts, 'diluted' NAV and warrant prices are also included. 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 price, diluted NAV, warrant prices, and volume. Dates for individual fund coverage are given in Table 2.

For the emerging market funds we use three weekly total return dollar indices compiled for the Emerging Markets Data Base (EMDB) maintained by the IFC. They are a local market index, the IFC Global index, and the IFC Investable index. Weekly coverage for the IFC indices begins in December 1988. As there are some exceptions in coverage, the adjoining table details the date when the series begins for each of the three IFC indices. 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. Local market and regional total return indices and exchange rates for the mature markets were provided by Goldman, Sachs & Co.; for the mature markets represented in our sample, daily coverage begins in January, 1986.

The database also includes information on capital changes, warrant exercise, the composition of dividend payments, and missing values for each fund. 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.

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 Wednesday 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.

Seven-day deposit LIBOR rates in dollars and pounds were obtained from Bloomberg and *The Financial Times of London*.

Market Indices for Emerging Equity Markets

Through August 1993

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:

All indices are taken from the Emerging Markets Data Base. Supplemental dates for the KSE Composite were obtained from the *KSE Fact Book* and for the BOVESPA from the Sao Paulo Stock Exchange.

Table A1
Empirical Size of Test Statistics

Panel 1: Closed-End Fund Benchmark Assets

	$\chi^2(2)$	$\chi^2(4)$	$\chi^2(14)$	$\chi^2(24)$
MV_1	5.55	5.80	10.45	16.80
MV_2	6.00	7.15	17.30	36.65
MV_3	5.60	4.80	7.55	6.60
MV_4	8.00	9.40	26.30	53.15
MV_5	7.65	8.90	22.70	48.05
MV_{W1}	7.00	7.90	19.40	40.90
MV_{W2}	5.95	5.95	12.30	26.05
$CV(\chi^2)$	5.99	9.49	23.69	36.42
$CV(MV_3)$	6.25	9.42	25.20	37.57

Panel 2: FT-Actuaries World Indices Benchmark Assets

	$\chi^2(2)$	$\chi^2(4)$	$\chi^2(14)$	$\chi^2(24)$
MV_1	5.85	7.00	10.45	13.05
MV_2	6.10	8.00	17.35	34.15
MV_3	5.85	6.50	9.25	7.30
MV_4	6.50	8.90	21.80	45.40
MV_5	6.45	8.50	19.30	41.60
MV_{W1}	6.30	7.80	16.15	33.55
MV_{W2}	6.15	7.00	12.85	23.85
$CV(\chi^2)$	5.99	9.49	23.69	36.42
$CV(MV_3)$	6.27	10.01	26.06	38.57

Notes:
The empirical size (in percent) is for a 5%-size test. The test statistics and Monte Carlo design are described in the text. The benchmark assets are further described in Section 3.3. In the column heading "$\chi^2(n)$," $n/2$ is the number of test assets used. The experiment for $n = 2$, uses a U. S. index of country funds. The results for U. K. indices are similar. CV stands for critical value. We report the 5% critical value for each χ^2-distribution and for the empirical distributions generated for the MV_3 test statistic.

Table A2
Power of Test Statistics

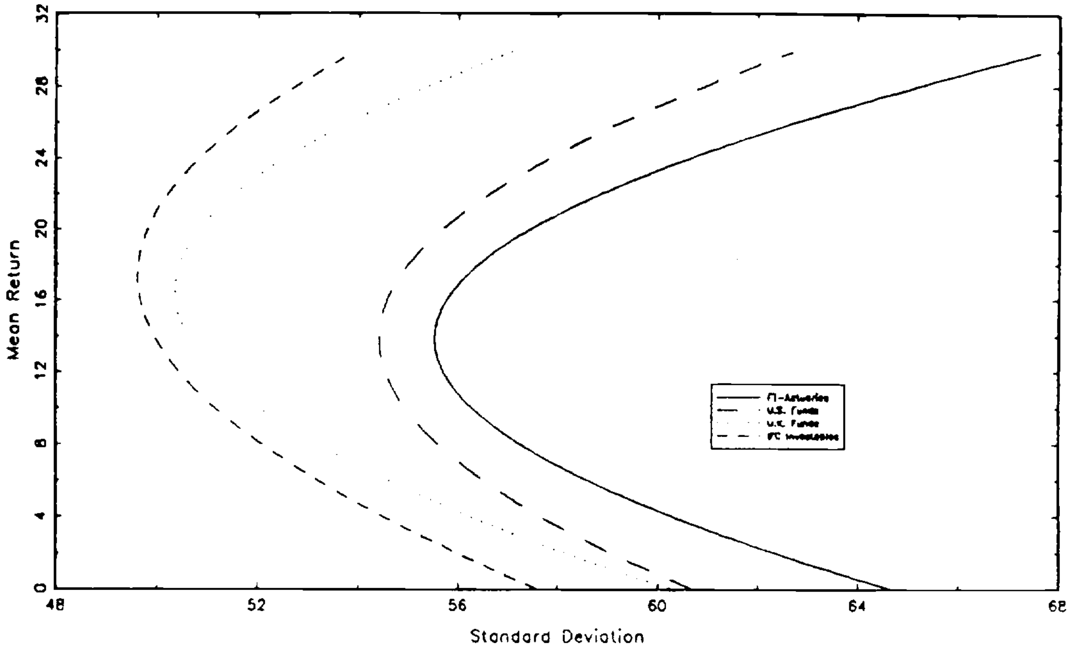
Panel 1: Closed-End Fund Benchmark Assets				
	$\chi^2(2)$	$\chi^2(4)$	$\chi^2(14)$	$\chi^2(24)$
MV_1	71.50	74.80	49.85	19.85
MV_2	71.70	75.15	52.45	27.95
MV_3	68.45	71.90	46.05	13.10
MV_4	68.35	71.50	44.65	21.00
MV_5	67.95	71.05	41.30	20.60
MV_{W1}	68.80	70.10	36.75	14.80
MV_{W2}	71.35	72.30	40.10	17.95

Panel 2: FT-Actuaries World Indices Benchmark Assets				
	$\chi^2(2)$	$\chi^2(4)$	$\chi^2(14)$	$\chi^2(24)$
MV_1	73.15	70.00	51.75	19.45
MV_2	73.15	70.85	54.00	27.85
MV_3	72.25	69.50	46.75	13.30
MV_4	73.10	70.00	49.10	26.85
MV_5	72.50	69.30	47.15	25.00
MV_{W1}	70.75	67.15	39.35	17.60
MV_{W2}	71.40	42.75	18.85	13.55

Notes:
The test statistics and the DGP under the alternative are described in the text. The benchmark assets are further described in Section 3.3. In the column heading "$\chi^2(n)$," $n/2$ is the number of test assets used. The experiment for $n = 2$, uses a U. K. index of country funds. The power using the U.S. index is slightly lower. The table reports the percent of 2,000 Monte Carlo replications that yield a higher test statistic than the 5% critical value from the empirical distribution under the null.

FIGURE 1

Unconditional Mean-Variance Spanning: FT-Actuaries Benchmark



Unconditional Mean-Variance Spanning: Control Benchmark