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Can the Gains from International Diversification Be Achieved without Trading Abroad?

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ABSTRACT

We examine whether portfolios of domestically traded securities can mimic foreign indices so that investment in assets that trade only abroad is not necessary to exhaust the gains from international diversification. We use monthly data from 1976 to 1993 for seven developed and nine emerging markets. Return correlations, mean-variance spanning, and Sharpe ratio test results provide strong evidence that gains beyond those attainable through home-made diversification have become statistically and economically insignificant. Finally, we show that the incremental gains from international diversification beyond home-made diversification portfolios have diminished over time in a way consistent with changes in investment barriers.

THE BENEFITS OF INTERNATIONAL DIVERSIFICATION have been emphasized over the past 40 years by financial economists, who have shown that investing in foreign indices reduces the volatility of U.S. market portfolios, with gains attributed to low return correlations between national equity indices.[1] Such investment in foreign indices requires holding securities that trade abroad, involving additional costs and potential barriers to international investment. Yet, over the past 20 years, an increasing number of country funds and depository receipts have started trading in the U.S. that, along with shares of multinational corporations, can be used to gain benefits from international diversification. In this paper, we examine whether investors can take advantage of the gains of international diversification by forming a portfolio of securities that trade in the United States, and we find that this

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[1] See, Solnik (1974), Errunza (1997), DeSantis and Gerard (1997), and Stulz (1997) for a detailed discussion of gains from international diversification.

is indeed the case. In other words, investors can mimic foreign indices by holding domestically traded assets; investing in assets that only trade abroad is no longer necessary to gain the benefits of international diversification.[2]

Specifically, we address three questions. First, is it possible to mimic foreign market indices with domestically traded securities? To answer this question, we develop diversification portfolios that measure U.S. investors' ability to obtain the benefits of international diversification using domestically traded assets. We study seven developed markets (DMs) and nine emerging markets (EMs) from 1976 to 1993. For each country, we construct diversification portfolios using U.S. market indices, 12 U.S. industry portfolios, 30 multinational corporation (MNC) stocks, closed-end country funds (CFs), and American Depositary Receipts (ADRs). Some of the portfolios involve claims on foreign assets. Traditionally, international diversification has involved foreign assets that only trade abroad; home-made international diversification includes claims on foreign assets that trade in the home market.

Second, has it become possible to exhaust the benefits from international diversification by investing in U.S. traded assets? As the mimicking portfolio is sequentially augmented with MNCs, CFs, and ADRs, it should become increasingly correlated with the foreign market portfolio, and investors should obtain most of the benefits of international diversification by investing in assets traded in their home market. At the limit, the benefits of investing in foreign indices would evaporate in spite of the low return correlation between home and foreign market indices. We study this issue within the mean-variance spanning framework of Huberman and Kandel (1987), DeSantis (1994), and Bekaert and Urias (1996).

Finally, has our ability to mimic foreign indices changed over time? Several recent studies report considerable time variation in return correlations.[3] We use the new generalized dynamic covariance (GDC) multivariate generalized autoregressive conditional heteroskedasticity (GARCH) model of Kroner and Ng (1998) to estimate time variation in the conditional correlation between foreign market indices and their respective mimicking portfolios.

The main results of the paper can be summarized as follows. During the period from 1976 to 1993, as the availability of MNCs, CFs, and ADRs rose, U.S. investors could effectively mimic foreign market returns with domestically traded securities. The mimicking portfolios, based on U.S. market indices and industry portfolios, are significantly enhanced by MNCs, CFs, and ADRs. The return correlations of home-made diversification portfolios with foreign market indices are higher than those with the S&P 500 index. Hence, the index level

[2] Due to investment barriers, regulatory restrictions on institutional portfolios and personal preferences, investors primarily hold securities traded in their home market. Indeed, home-made diversification is consistent with the observed home bias in investors' portfolios. See Cooper and Kaplanis (1994) and Kang and Stulz (1997) for a discussion of home bias.

[3] See DeSantis (1994), Longin and Solnik (1995), Karolyi and Stulz (1996), Erb, Harvey, and Viskanta (1996), Bekaert and Harvey (1997), and Bekaert, Erb et al. (1998) for evidence and explanation of time variation in return correlations.

correlations overstate the gains from investing in securities that only trade abroad. The likelihood that overseas market index returns are spanned in a mean-variance context by a representative set of benchmark assets is greater with increased augmentation of the diversification portfolios. In fact, for most markets, the gains beyond those attainable through home-made diversification are statistically and economically insignificant. Although investors should continue to be aware of their exposure to foreign risk, they no longer need to trade abroad to achieve an internationally mean-variance efficient portfolio. Finally, the substantial time variation in conditional return correlations between foreign indices and home-made diversification portfolios is consistent with changes in investment barriers, such as new listings of CFs/ADRs, changes in rules governing foreign portfolio investments, and national, political, and economic events.

The paper is organized as follows. Section I describes the construction of diversification portfolios. Section II investigates the potential of home-made international diversification to substitute for international diversification involving foreign-traded assets. In Section III, we report test results for mean-variance spanning and change in Sharpe ratios to assess the ability of domestically traded assets to exhaust diversification benefits. Section IV presents conditional correlations based on GDC multivariate GARCH model estimation and model diagnostics, and provides an explanation for time variation in diversification gains. Conclusions are presented in Section V.

I. Construction of Diversification Portfolios

A diversification portfolio (D) is defined as the portfolio of domestically traded securities that is most highly correlated with a target foreign market index. Our analysis is conducted from the perspective of U.S. investors. Although all securities that trade in the domestic (U.S.) market should be considered eligible for constructing diversification portfolios, simplifying assumptions must be made. We follow Breeden, Gibbons, and Litzenberger (1989) and disaggregate the market into 12 U.S. two-digit SIC industry portfolios. Several authors have reported the importance of industry level diversification for global investment strategy,[4] hence we use the fitted values of the following regression to obtain the portfolio ($D1$):[5]

$$R_{I,t} = \beta_1 R_{e1,t} + \cdots + \beta_{12} R_{e12,t} + \epsilon_{I,t}, \tag{1}$$

where $R_{I,t}$ is the return on the Ith foreign market index during period t and $R_{e1}, \ldots, R_{e12}$ are returns on the 12 two-digit SIC industry indices.

[4] See Griffin and Karolyi (1998) and references therein.

[5] Our formulation of the diversification portfolio is similar in spirit to Breeden et al. (1989) in which the fitted values from a regression of consumption on portfolios of securities is used as a measure of the maximally correlated consumption portfolio.

The hypothesis that multinational corporations (MNCs) indirectly provide benefits of international diversification has received much attention. On the one hand, Agmon and Lessard (1977) and Fatemi (1984) find that U.S. investors recognize the benefits of corporate international diversification in pricing MNC stocks and that the higher the degree of international involvement, the lower the beta relative to the domestic market portfolio. Errunza and Senbet (1981) and Bodnar, Tang, and Weintrop (1998) report a positive relationship between a firm's international involvement and its market value. On the other hand, Jacquillat and Solnik (1978) conclude that although MNCs provide some diversification benefits, they are a poor substitute for international diversification based on foreign-traded securities. Thus, although investments in MNCs may not be sufficient to capture the benefits from international diversification, the mimicking portfolio should improve by their inclusion. Hence, we use a sample of 30 large U.S. MNCs along with three U.S. market indices and 12 industry portfolios to obtain a more inclusive portfolio ($D2$). In order to preserve degrees of freedom, we use stepwise regressions to determine these portfolios.[6]

For many countries, closed-end country funds and/or American Depositary Receipts were introduced in the U.S. market during our sample period. Since these securities represent claims on foreign assets, they are generally viewed as international assets even though they trade on U.S. markets. Portfolios that incorporate these assets are called augmented diversification (AD) portfolios. Country funds generally trade in the U.S. markets at prices that differ from the market value of the underlying securities on the local market. As a result, the returns that U.S. investors can obtain on CFs may do a poor job of tracking the returns of the underlying assets. Further, the closed-end funds are actively managed, adding to the difficulty of mimicking the foreign index returns. Several studies suggest that although CFs provide some benefits of international diversification, their ability to substitute for (unattainable) foreign market index returns is limited.[7] Nonetheless, these CFs represent an attainable diversification opportunity for U.S. investors. We isolate the impact of country funds by estimating the following regressions:

$$R_{I,t} = \varphi_1 R_{D2,t} + \varphi_2 R_{c,t} + e_{I,t}, \tag{2}$$

where $R_{D2,t}$ is the return on portfolio ($D2$) and $R_{c,t}$ is the return on the relevant country fund. For countries with multiple CFs, we select the one with the longest history. Regressions are based on the full sample of available monthly data on market returns (i.e., January 1976 to December 1993). For the period prior to country fund inception, $R_{c,t}$ is set to zero. The coefficients φ_1, φ_2 are interpreted as the portfolio weights of an augmented diversification portfolio ($AD1$) which incorporates both the previous portfolio

[6] In order to maintain the identity of each asset in the portfolio, we do not reduce the pool of assets by means of principal components.

[7] See Errunza, Senbet, and Hogan (1998) and references therein.

($D2$) and the country funds. The significance of the components can be tested based on their respective t-statistics. The fitted values of equation (2) represent the returns associated with portfolio ($AD1$) and are incorporated in the next sections to assess the impact of country funds on diversification gains.

American Depositary Receipts represent a claim on a specific number of underlying shares that trade on the foreign market. By allowing U.S. investors to own foreign shares indirectly they offer a convenient vehicle for international diversification. At times, ADRs have traded at a premium or discount to the market value of the underlying security. They are issued primarily by well-established firms and hence it may not be possible to duplicate a well-diversified foreign market portfolio with a basket of U.S.-traded ADRs. Nonetheless, recent studies suggest that ADRs can be used to replicate local market indices.[8] Hence, the diversification portfolio is further augmented to incorporate ADRs as follows:

$$R_{I,t} = \varphi_1 R_{D2,t} + \varphi_2 R_{c,t} + \varphi_3 R_{adr,t} + e_{I,t}, \tag{3}$$

where $R_{adr,t}$ is the return on the relevant ADR. In the case of multiple ADRs, we select the one with the longest history. Again, for the period of estimation prior to the inception of the ADR, $R_{adr,t}$ is set to zero.[9] The resulting portfolio is denoted as $AD2$.

We use three U.S. indices, 12 U.S. value-weighted industry portfolios, a sample of 30 multinational firms, CFs, and ADRs listed on the New York Stock Exchange as the eligible set. The composition of the industry portfolios is identical to Breeden et al. (1989). The sample of multinational firms is selected from the 50 largest U.S. multinationals in 1976 ranked by sales as reported by *Fortune* magazine.[10] The year 1976 corresponds to the beginning of our test period. All U.S.-based return data are from the Center for Research in Securities Prices (CRSP). Stocks that are no longer listed as of December 31, 1993, or for which data are missing from CRSP during the sample period are deleted from the eligible set. This leaves us with a sample of 30 MNCs. A complete list of the set of eligible securities is reported in Table I.

The total monthly returns for EMs and DMs are based on International Finance Corporation (IFC) and Morgan Stanley Capital International market indices, respectively. Indices are market value weighted and expressed in U.S. dollars. Returns from the emerging markets (Argentina, Brazil, Chile, Greece, India, Korea, Mexico, Thailand, and Zimbabwe) and developed mar-

[8] See Jorion and Miller (1997). Also, see Karolyi (1998) for a survey of the ADR literature.

[9] Note that in some instances the ADRs were traded over the counter as "Pink Sheets," for example, Telefonos de Mexico before 1991. Hence, the decline in gains from international diversification may have preceded the listing dates reported in this study. We thank an anonymous referee for this explanation.

[10] *Fortune*, May 1977, pp. 366–367.

Table I

List of Eligible Securities

The eligible set consists of 30 of the largest U.S. multinational corporations as ranked by 1976 sales reported by *Fortune* magazine, 3 broad-based U.S. market indices, 12 U.S. two-digit SIC industry portfolios, 12 closed-end country funds (CFs), and 48 American Depositary Receipts (ADRs). The information on ADR sample and their listing dates are obtained from the Bank of New York, Merrill Lynch and Company, NYSE fact book, and data tables in Gooptu (1993). All return data are from CRSP.

Panel A: Multinational Corporations (Ms)

M1	Amerada	M11	Exxon Corp.	M21	Phillips Petroleum Corp.
M2	Ashland Oil Inc	M12	Ford Motor Co.	M22	Procter and Gamble Co.
M3	Atlantic Richfield Co.	M13	General Electric Co.	M23	Rockwell International Corp.
M4	Bethlehem Steel Co.	M14	General Motors Corp.	M24	Sun Inc.
M5	Boeing Co.	M15	Goodyear Tire and Rubber Co.	M25	Tenneco Inc.
M6	Caterpillar	M16	Grace W R and Co.	M26	Texaco Inc.
M7	Chrysler Co.	M17	International Business Machines	M27	Union Carbide Corp.
M8	Dow Chemical Co.	M18	Mobil Corp.	M28	United Technologies
M9	Du Pont E I De Nemours Co.	M19	Monsanto Co.Tr	M29	Westinghouse Electric Corp
M10	Eastman Kodak Co.	M20	Occidental Petroleum Co.	M30	Xerox Corp

Panel B: Broad U.S. Market Indices (Is)

I1	Value-weighted market return, including dividends
I2	Equal-weighted market return, including dividends
I3	Standard and Poors 500 Composite index

Panel C: Two-Digit SIC Industry Portfolios (IIs)

		Codes	Two-digit SIC industry portfolios (IIs)		Codes
II1	Petroleum	13 29	II7	Capital Goods	34 35 38
II2	Finance and Real Estate	60 61 62 63 64 65 66 67 68 69	II8	Transportation	40 41 42 44 45 47
II3	Consumer Durables	25 30 36 37 50 55 57	II9	Utilities	46 48 49
II4	Basic Industries	10 12 14 24 26 28 33	II10	Textiles and Trade	22 23 31 51 53 56 59
II5	Food and Tobacco	01 20 21 54	II11	Services	72 73 75 80 82 89
II6	Construction	15 16 17 32 52	II12	Leisure	27 58 70 78 79

Panel D: Closed-End Country Funds

Emerging market CFs		Start Date	Emerging market CFs		Start Date
ARGCF	Argentina Fund Inc.	11/91	KORCF	Korea Fund	8/84
BRACF	Brazil Fund Inc.	3/88	MEXCF	Mexico Fund	6/81
CHICF	Chile Func Inc.	9/89	THACF	Thailand Fund Inc.	2/88
INDCF	India Growth Fund Inc.	8/88			

Panel E: Developed Market CFs					
FRACF	France Fund Inc.	5/86	UKCF	United Kingdom Fund	8/87
GERCF	Germany Fund Inc.	2/90	JAPCF	Japan OTC Equity Fund Inc.	3/90
ITACF	Italy Fund Inc.	2/86			
Panel F: American Depository Receipts					
Emerging Market ADRs					
MEXADR	Telefonos de Mexico	5/91			
BRAADR	Aracruz Celulose (Brazil)	6/92			
CHIADR	Compania de Telefonos de Chile	7/90			
Panel G: Developed Market ADRs					
Australia:			UK:		
$AUSADR_1$	National Australia Bank	6/88	$UKADR_1$	Nat Westminister Bank Plc	10/86
$AUSADR_2$	Orbital Engine Ltd	12/91	$UKADR_2$	BET Plc	8/87
$AUSADR_3$	Coles Myer Ltd	10/88	$UKADR_3$	Automated Security Holdings	7/92
$AUSADR_4$	FAI Insurances Ltd	9/88	$UKADR_4$	Barclays Plc	9/86
$AUSADR_5$	News Corp Ltd	5/86	$UKADR_5$	Bass Plc	2/90
$AUSADR_6$	Westpac Banking Corp Ltd	3/89	$UKADR_6$	British Airways Plc	2/87
France:			$UKADR_7$	British Petroleum Plc	10/87
$FRAADR_1$	Alcatel Alsthom Co. General	5/92	$UKADR_8$	British Steel Plc	12/88
$FRAADR_2$	Rhone Poulenc Rohrer Inc	9/63	$UKADR_9$	British Tel. Plc	12/84
$FRAADR_3$	Total SA	10/90	$UKADR_{10}$	Grand Metropolitan Plc	3/91
$FRAADR_4$	Western Mining Corp Holding	5/86	$UKADR_{11}$	Enterprice Oil Plc	10/92
Italy:			$UKADR_{12}$	Cable and Wireless Plc	9/89
$ITAADR_1$	Benetton Group Spa	6/89	$UKADR_{13}$	Hanson Plc	11/86
$ITAADR_2$	Fiat Spa	2/89	$UKADR_{14}$	Huntingdon International	2/89
$ITAADR_3$	Fila Holding Spa	5/93	$UKADR_{15}$	Imperial Chemicals Inds Plc	7/62
$ITAADR_4$	Industie Natuzzi Spa	5/93	$UKADR_{16}$	RTZ Plc	6/90
$ITAADR_5$	Luxottica Group Spa	12/90	$UKADR_{17}$	Shell Transport and Trading Co.	2/82
$ITAADR_6$	Montedison	7/87	$UKADR_{18}$	Smithklin Beetcham Plc	7/89
Japan:			$UKADR_{19}$	Zeneca Group	6/93
$JAPADR_1$	Honda Ltd	2/77	$UKADR_{20}$	Willis Corroon Group Plc	10/90
$JAPADR_2$	Kubota Corp	11/76	$UKADR_{21}$	Waste Management Intl Plc	4/92
$JAPADR_3$	Kyocera Corp	5/80	$UKADR_{22}$	Vodafone Group Plc	10/88
$JAPADR_4$	Matsushita Electrical Indl.	12/71	$UKADR_{23}$	Unilever	7/62
$JAPADR_5$	Sony Corp	9/70			
$JAPADR_6$	TDK Corp	6/82			

kets (Australia, Canada, France, Germany, Italy, Japan, and the United Kingdom) are included from January 1976 to December 1993. There are two eligible IFC indices, the IFC Global (IFCG) and its subset, the IFC Investable (IFCI). Although the IFCI indices take into account the openness of each market from the perspective of the foreign institutional investor, both the IFCG and the IFCI indices represent unattainable performance benchmarks for most investors. We use IFCG indices given their longer history and broader coverage as opposed to the IFCI indices. The gains from international diversification inferred from IFCG indices would probably be higher than those based on IFCI indices given their more unattainable nature. Thus, the IFCG indices impose a higher hurdle than the IFCI indices for demonstrating that the gains from international diversification are disappearing over time.

Table II reports descriptive statistics.[11] The behavior of developed and emerging market returns is similar to that reported in Harvey (1995) and DeSantis (1994). Briefly, EM returns on average are much higher, display greater volatility, and are more autocorrelated than their developed market counterparts. We also note the propensity for extreme outliers in the EM sample.

II. Home-Made International Diversification versus Foreign Asset Based International Diversification

In this section, we first report the composition of various diversification portfolios. We then discuss the ability of domestically traded securities to mimic foreign market indices.

A. Composition of Diversification Portfolios

Table III, Panel A, reports the resulting composition of portfolio $D2$ for EMs from the stepwise regressions.[12] The U.S. market indices, industry portfolios, and MNCs constitute $D2$ in the case of Chile, Korea, Mexico, and Zimbabwe. For Greece, the portfolio comprises industry portfolios and MNCs. In the case of Argentina, Brazil, India, and Thailand, the portfolio $D2$ is made up entirely of MNCs. Although in no case does a diversification portfolio include the petroleum industry, a number of MNCs with significant oil interests are included. The exclusion of the leisure industry is not surprising.[13] These results are consistent with past studies regarding the impor-

[11] All results are reported in U.S. dollars. Local currency results are very similar and are available from the authors.

[12] The stepwise procedure is based on a forward and backward *p-value* threshold of 0.20. This effectively lowers the dimensionality of eligible securities from 45 to a range of 2 to 16, with an average of approximately eight significant elements per diversification portfolio.

[13] Note that the diversification portfolios have been constructed ex post. That is, the construction of the portfolios is based on information that would not have been available to market participants. To investigate how sensitive the results of this paper are to the ex post construction of the diversification portfolios, we compare them to results for ex ante diversification portfolios based on 72-month rolling regressions. We found the conclusions of this paper to be the same regardless of whether portfolios are constructed ex post or ex ante. Details are available from the authors.

Table II

Summary Statistics

Descriptive monthly statistics for emerging market and developed market dollar returns for the sample period from January 1976 to December 1993. ρ_i denotes autocorrelation of lag i, Q(12) refers to Ljung–Box statistics for serial correlation based on 12 lags and $Q^2(12)$ are Ljung–Box statistics for squared returns.

	Mean	Standard Deviation	Minimum	Maximum	ρ_1	ρ_2	ρ_3	Q(12)	$Q^2(12)$
Emerging markets									
Argentina	0.053	0.286	−0.650	1.781	0.052	0.065	0.119	12.033	15.413
Brazil	0.023	0.171	−0.569	0.575	0.025	−0.032	−0.031	9.211	42.106
Chile	0.031	0.111	−0.280	0.629	0.165	0.249	−0.010	30.614	27.418
Greece	0.006	0.101	−0.308	0.586	0.130	0.169	0.019	26.19	44.718
India	0.017	0.079	−0.244	0.353	0.104	−0.077	0.000	20.109	53.605
Korea	0.018	0.090	−0.192	0.448	0.006	0.074	−0.003	7.366	11.852
Mexico	0.024	0.124	−0.593	0.396	0.244	−0.074	−0.040	24.613	57.254
Thailand	0.020	0.078	−0.338	0.322	0.103	0.176	0.022	21.860	18.056
Zimbabwe	0.011	0.100	−0.279	0.460	0.169	0.187	0.260	50.465	13.714
Developed markets									
Australia	0.013	0.074	−0.445	0.208	0.013	−0.111	−0.029	11.493	3.297
Canada	0.010	0.056	−0.220	0.179	−0.030	−0.052	0.052	14.073	20.260
France	0.014	0.073	−0.232	0.265	−0.005	−0.028	0.081	9.083	11.739
Germany	0.012	0.061	−0.176	0.202	−0.031	0.001	0.111	16.67	23.843
Italy	0.012	0.079	−0.204	0.309	0.106	−0.030	0.080	18.597	11.789
Japan	0.015	0.068	−0.194	0.242	0.029	−0.048	0.058	6.787	34.3152
U.K.	0.015	0.064	−0.215	0.227	−0.002	−0.090	−0.075	11.281	9.563

Table III

Composition of Diversification Portfolios for the Emerging and Developed Markets

Diversification portfolios ($D2$) are based on stepwise regression procedures over three broad U.S. indices (Is), 12 U.S. industry indices (IIs), and 30 multinational corporations (MNCs). The numbers in each column correspond to the identification in Table I. Augmented diversification portfolios AD1 and AD2 are obtained by augmenting portfolio D2 with each country's country fund and ADRs respectively. AD1s are the fitted values from the regression $R_{I,t} = \varphi_1 R_{D2,t} + \varphi_2 R_{c,t} + e_{I,t}$ and AD2s are the fitted values from the regression $R_{I,t} = \varphi_1 R_{D2,t} + \varphi_2 R_{c,t} + \varphi_3 R_{adr,t} + e_{I,t}$. R_{It} are Ith market returns, R_{D2} are $D2$ returns, R_c are country fund returns and R_{adr} are ADR returns.

Columns two, three, and four report the parameter estimates (t-statistics) for these regressions. Columns five, six, and seven report the composition of portfolio $D2$. Diversification portfolios ($D2$) are based on stepwise regression procedures over three broad U.S. indices (Is), 12 U.S. industry indices (IIs), and 30 multinational corporations. Column eight reports significant ADRs (* denotes the ADR used in the regression to estimate ϕ_3).

Country	ϕ_1	ϕ_2	ϕ_3	Is	IIs	MNCs	ADRs
Panel A: Emerging Markets							
Argentina:							
AD1	0.9803	0.7842				3, 11, 17, 24	
	(3.536)	(1.257)					
Brazil:							
AD1	0.5816	1.0224				4, 10, 17, 19, 29, 30	
	(2.583)	(7.341)					
AD2	0.6063	1.0204	0.2694				
	(2.676)	(7.325)	(0.970)				
Chile:							
AD1	0.9436	0.3119		3	2, 5, 6, 8	1, 5, 6, 16–18, 22	
	(5.913)	(2.748)					
AD2	0.9287	0.2841	0.1299				
	(5.782)	(2.402)	(0.845)				
Greece:					2, 4, 5, 10	5, 13, 19, 24, 25, 27	
India:							
AD1	0.7759	0.3847				2, 6	
	(2.685)	(5.123)					
Korea:							
AD1	0.8580	0.2745		1, 3	3, 4, 6, 9	1–3, 5, 6, 7, 12, 24, 26, 29	
	(8.488)	(5.338)					

Mexico:							
AD1	0.4640	0.5702		2	7, 10, 11	5, 17, 18	
	(3.531)	(9.381)					
AD2	0.4632	0.5689	0.0149				
	(3.510)	(9.046)	(0.086)				
Thailand:							
AD1	0.7535	0.4034				2, 10, 13, 14, 16, 17, 20, 22, 23, 26, 30	
	(5.320)	(5.501)					
Zimbabwe:				2	4, 5, 6, 11	7, 22,	
Panel B: Developed Markets							
Australia:				2	4, 5, 9, 10, 12	2, 6, 8, 11, 14, 16, 19, 20, 24, 30	1*, 2, 5
AD2	0.3793		0.2508				
	(4.664)		(2.802)				
Canada:				2	3–5, 8, 10	7, 11, 14, 15, 19, 20, 21, 26	
France:							
AD1	0.9721	0.0978			4–6, 10	2, 8, 9, 14–16, 19, 28	1, 2*
	(9.467)	(0.904)					
AD2	0.9793	0.0955	0.8617				
	(9.686)	(0.897)	(2.797)				
Germany:							
AD1	0.8179	0.3031			5	1, 2, 15–17, 19, 23, 26	
	(7.296)	(6.110)					
Italy:							
AD1	0.8770	0.2624		3	2, 4, 6, 9, 11	6, 9–12, 15, 22, 27, 28	2*, 3
	(7.338)	(4.414)					
AD2	0.8388	0.1604	0.3839				
	(7.262)	(2.577)	(4.187)				
Japan:							1, 2, 3, 5*
AD1	0.9222	0.3604			2, 4, 9, 10, 12	1, 9, 11, 15, 22, 27, 28	
	(8.311)	(5.107)					
AD2	0.6814	0.2770	0.3839				
	(7.262)	(4.387)	(7.873)				
UK:							2, 5, 6, 7, 21, 23*
AD1	0.8708	0.3577		2	4, 5, 9, 10, 12	2, 6, 8, 11, 14, 16, 19, 20, 24, 30	
	(10.199)	(4.230)					
AD2	0.8198	0.1206	0.3514				
	(9.890)	(1.230)	(4.328)				

tance of industry level diversification. Inclusion of MNCs in the case of all EMs provides new evidence in support of the role of corporate diversification in providing international diversification benefits.

Parameter estimates for augmented portfolio *AD*1 illustrate the extent to which home-made diversification is further enhanced by the introduction of country funds. For situations in which more than one country fund is traded, only the oldest fund enters the portfolio significantly. This is similar to the seasoning effect suggested by Diwan, Errunza, and Senbet (1993) for CFs. With the exception of Argentina, the portfolio weight associated with the country fund (β_2) is statistically significant. Furthermore, in the cases of Brazil and Mexico, the portfolio weight associated with the country fund exceeds that of all other assets combined. During our sample period there are only three ADRs from emerging markets, each with a short history of U.S. trading. The parameter estimates for augmented portfolio *AD*2 suggest insignificant improvement in home-made diversification from inclusion of these ADRs.

Table III, Panel B, reports the results for DMs. These results are quite similar to those of EMs with respect to the contribution of industry portfolios and MNCs to home-made diversification. The DM country funds are of relatively recent origin compared to ADRs. Nonetheless, CFs play an important role in the cases of Germany (which has no ADRs at this time), Italy, and Japan. In the case of France, the contribution of the country fund is not significant, irrespective of whether ADRs are included in the eligible set. For the U.K., the contribution of the country fund is subsumed by ADRs. This is not surprising given that 23 U.K. ADRs trade on the NYSE, two of which traded throughout our sample period. To preserve degrees of freedom, we use stepwise procedures to select only those ADRs that enhance home-made diversification in a statistically significant way. In all five DMs with ADR listings, multiple ADRs are found to enhance home-made diversification benefits. This is likely attributable to the fact that DM ADRs are more prevalent and, in many instances, have traded throughout our sample period.

B. Benefits of Home-Made Diversification

Table IV reports the unconditional return correlations of foreign markets indices with three U.S. market indices; the portfolio *D*1 based on 12 industry portfolios; the portfolio *D*2 selected from three U.S. market indices, 12 industry portfolios, and 30 MNCs; and the augmented portfolios including the initial country fund (*AD*1) and ADRs (*AD*2). The return correlation between the U.S. market index and a target foreign market index is a traditional measure of the benefits of international diversification: The lower the correlation, the greater the potential benefits. The return correlation between a diversification portfolio constructed from domestically traded assets and a target foreign market index is a measure of home-made diversification benefits. The higher the correlation, the greater the opportunity to realize international diversification benefits through domestically traded assets.

Table IV

Unconditional Correlations

I1, I2, and I3 are value-weighted, equal-weighted, and Standard and Poors 500 return indices. D1 denotes diversification portfolios based on 12 industrial indices. D2 denotes the diversification portfolio selected from three broad-based U.S. indices, 12 industrial portfolios, and 30 multinational corporations by the stepwise procedure. AD1 are augmented diversification portfolios in which D2 is augmented using each country's country fund. AD2 are augmented portfolios in which AD1 is augmented by the country's representative ADRs.

	Correlations of Market Index Return with						
	I1	I2	I3	D1	D2	AD1	AD2
Emerging markets							
Argentina	0.026	0.050	0.018	0.242	0.209	0.209	na
Brazil	0.058	0.079	0.060	0.162	0.258	0.502	0.505
Chile	0.021	0.031	0.025	0.214	0.312	0.355	0.358
Greece	0.088	0.069	0.100	0.289	0.402	na	na
India	−0.016	0.016	−0.017	0.197	0.202	0.371	na
Korea	0.168	0.197	0.167	0.343	0.525	0.601	na
Mexico	0.283	0.281	0.283	0.358	0.414	0.639	0.639
Thailand	0.112	0.122	0.129	0.186	0.401	0.514	na
Zimbabwe	0.033	−0.019	0.030	0.323	0.364	na	na
Developed markets							
Australia	0.404	0.427	0.406	0.544	0.636	na	0.674
Canada	0.714	0.706	0.699	0.787	0.830	na	na
France	0.411	0.362	0.414	0.482	0.548	0.550	0.553
Germany	0.315	0.271	0.325	0.379	0.477	0.585	na
Italy	0.218	0.241	0.209	0.431	0.473	0.537	0.539
Japan	0.227	0.213	0.231	0.366	0.467	0.550	0.678
UK	0.499	0.467	0.504	0.578	0.610	0.649	0.653

Our results strongly suggest that it is possible to mimic the foreign market index returns with portfolios of domestically traded assets. Indeed, as we sequentially augment the mimicking portfolios, our ability to substitute home-made international diversification for foreign asset based international diversification dramatically increases. For example, the correlation between the U.S. index and the Mexico index is 0.28, compared with 0.64 between the most augmented portfolio ($AD2$) and the Mexico index. Such differences are even more extreme for other EMs in our sample. In the case of India, the correlation with the U.S. broad index is -0.02 whereas the correlation with respect to $AD1$ is 0.37. The inclusion of ADRs has virtually no effect on the correlation between the mimicking portfolios and EM returns.

With respect to DMs, the message from return correlations at the index level is also dramatically different from that based on the diversification portfolios. For example, Japan would be viewed as one of the most diversification-enhancing countries based on index correlation but would be next to the lowest based on $AD2$ correlation. In general, correlations with respect to the U.S. index overstate the gains from investing in securities that only trade abroad, but not nearly to the extent that they do for the EMs. To illustrate, the average correlation coefficient for DMs and the U.S. index is 0.4, whereas the average correlation coefficient for the most augmented portfolio is 0.64. For EMs, these average correlations are 0.08 and 0.43, respectively. Interestingly, these averages also reveal that the difference between EMs and DMs in terms of foreign asset based international diversification gains is not nearly as great as one might infer by looking only at simple index correlations. In fact, Mexico and Korea do not provide much more diversification benefit through their respective market portfolios beyond those attainable through home-made diversification than France, Germany, or Italy.[14] Indeed, these results caution against the use of correlations of market-wide index returns as a measure of international diversification gains involving foreign-traded assets.[15] Such gains must be measured beyond those attainable through home-made international diversification.

III. Mean-Variance Spanning Tests

In the previous section, we show that as the U.S. index returns are sequentially augmented with industry portfolios, MNCs, CFs, and ADRs, the diversification portfolios become increasingly correlated with the target foreign market return. The key issue addressed in this section is whether it is possible to exhaust international diversification benefits through home-made diversification. Huberman and Kandel (1987) show that for any partition of assets into a set of test assets and benchmark assets, the inclusion of additional test assets into the set of benchmark assets shifts the efficient frontier to the left if and only if the test assets are not mean-variance spanned by the benchmark assets. Mean variance spanning tests have been developed by Huberman and Kandel, DeSantis (1994), and Bekaert and Urias (1996). Both DeSantis and Bekaert and Urias apply spanning tests to study benefits of international diversification and define benchmark assets as a set of developed market assets. In contrast, we use the set of assets which comprise the various diversification portfolios developed in the previous section as benchmark assets. Our benchmark assets are summarized below with the underlying U.S. assets in each set in

[14] Canada is an interesting benchmark case because a large number of Toronto Stock Exchange (TSE) listed stocks are cross-listed in the United States. Almost 50 percent of TSE 300 market capitalization is traded on the NYSE, AMEX, or Nasdaq. Indeed, the home-made diversification including cross-listings should be virtually perfect. We thank an anonymous referee for pointing this out to us.

[15] Interestingly, the correlations based on ex ante diversification portfolios show even stronger evidence of diminished global diversification opportunities than do the ex post diversification portfolios.

parentheses: Set I (II1 to II12); Set II (II1 to II12 and M1 to M30); Set III (II1 to II12, M1 to M30, CFs and ADRs). While they are in the same spirit as the home-made diversification portfolios $D1$, $D2$, $AD1$, and $AD2$, the benchmark sets differ in two respects. First, we limit the number of assets in each set to four, based on simulation results of Bekaert and Urias who show that the power of the test is extremely sensitive to the number of benchmark assets. Second, we use a slightly different selection criterion. For each benchmark set, we choose the four assets that maximize the probability of not rejecting spanning, as measured by the p-value associated with Huberman and Kandel (HK) tests. The HK test involves estimating the following equation:

$$R_{I,t} = \alpha_i + \beta_1 R_{e1,t} + \beta_2 R_{e2,t} + \beta_2 R_{e2,t} + \beta_4 R_{e4,t} + \epsilon_{I,t}, \tag{4}$$

where $R_{I,t}$ is the return on the Ith foreign market and $R_{e1,t},\ldots,R_{e4,t}$ are the returns on the benchmark assets. Huberman and Kandel show that $R_{I,t}$ is spanned by $R_{e1,t},\ldots,R_{e4,t}$ if and only if the following two conditions hold:

$$\alpha_i = 0, \tag{5}$$

$$\sum_{i=1}^{4} \beta_i = 1. \tag{6}$$

Huberman and Kandel test these restrictions based on OLS estimates of equation (4). Bekaert and Urias (BU) use GMM estimators to form a likelihood ratio-type test in which corrections for serial correlation are made.[16]

Panel A of Table V reports the p-values associated with HK mean-variance spanning test statistics. We interpret these values as a measure of the degree to which one can reject mean-variance spanning. The higher the p-value, the more confident we are that a given market index is mean-variance spanned and hence does not enhance diversification benefits. For the first set of benchmark assets, we find our results differ dramatically from those of DeSantis (1994) and Bekaert and Urias (1996). Whereas, DeSantis and Bekaert and Urias report significant diversification gains from investing in EM indices, we reject spanning in only five of the nine cases. Specifically, we reject spanning for Chile, Greece, India, Thailand, and Zimbabwe. More surprisingly, we reject spanning for three of the seven developed markets, a result dramatically different from that of DeSantis, who finds no diversification gains from investing in developed markets. As expected, the p-values increase as

[16] The exact test statistic used in this study is their MV_3 statistic. See Bekaert and Urias (1996) for further details.

Table V

Tests of Mean-Variance Spanning and Change in Sharpe Ratio

Panels A and B report p-values associated with spanning tests of Huberman and Kandel (1987) and Bekaert and Urias (1996). Panel C reports the change in Sharpe ratio. Benchmark Set I includes four of the 12 industrial portfolios that maximize the degree of spanning. Set II includes four assets from the list of 12 industry indices and the 30 multinational corporations. Set III includes four assets chosen from 12 industry indices, 30 multinational corporations, country funds, and American Depositary Receipts. The tests employ monthly observations from January 1976 to December 1993. Note that based on Monte Carlo experiment, Bekaert and Urias report change in the Sharpe ratio of 0.057 at the 95th percentile of the empirical distribution under the null of spanning.

	Panel A: Huberman–Kandel (OLS) Benchmark Assets			Panel B: Bekaert–Urias (GMM) Benchmark Assets			Panel C: Change in Sharpe Ratio Benchmark Assets		
	Set I	Set II	Set III	Set I	Set II	Set III	Set I	Set II	Set III
Emerging markets									
Argentina	0.0678	0.1144	na	0.2312	0.1431	na	0.0533	0.0435	na
Brazil	0.1475	0.299	0.6250	0.0096	0.2018	0.5141	0.0076	0.0102	0.0026
Chile	0.0000	0.0007	0.0044	0.0295	0.0098	0.1178	0.0641	0.0926	0.0637
Greece	0.0000	0.0001	na	0.1483	0.0744	na	0.0000	0.0022	na
India	0.0000	0.0000	0.0000	0.0071	0.0055	0.0074	0.0411	0.0372	0.0424
Korea	0.2765	0.4057	0.6435	0.5830	0.6739	0.6783	0.0147	0.0102	0.0093
Mexico	0.3672	0.4223	0.5004	0.6255	0.6898	0.6503	0.0185	0.0132	0.0132
Thailand	0.0000	0.0000	0.0041	0.0465	0.0826	0.1267	0.0558	0.0550	0.0713
Zimbabwe	0.0000	0.0000	na	0.0275	0.0217	na	0.0004	0.0054	na
Developed markets									
Australia	0.1023	0.2708	0.9556	0.2889	0.4788	0.6337	0.0020	0.0016	0.0001
Canada	0.3874	0.9169	na	0.4206	0.4650	na	0.0117	0.0017	na
France	0.5898	0.7309	0.9541	0.5976	0.6702	0.9593	0.0031	0.0015	na
Germany	0.0088	0.0684	0.1795	0.2580	0.2335	0.3762	0.0016	0.0038	0.0220
Italy	0.0303	0.0303	0.8661	0.1101	0.1101	0.7715	0.0024	0.0024	0.0021
Japan	0.0137	0.0153	0.6946	0.1625	0.0892	0.7482	0.0129	0.0069	0.0190
U.K.	0.3223	0.4607	0.8481	0.4389	0.5460	0.8830	0.0087	0.0109	na

we move sequentially from set I to set III benchmark assets. Although the inclusion of MNCs (set II) and cross-listings (set III) leads to a reduction in the likelihood of rejecting spanning, the overall conclusion remains unaltered for EMs; that is, we reject spanning for Chile, Greece, India, Thailand, and Zimbabwe at the five percent critical level. This is not true for the DMs. Once cross-listed securities are considered in the benchmark (i.e., set III) we fail to reject spanning in all cases. Hence, cross-listed securities play an important role in enhancing home-made diversification in the case of DMs.

The results based on BU mean-variance spanning test statistics reported in Panel B of Table V are similar to those of HK. However, there are some important differences. For example, using set III benchmark assets, we reject spanning for India and Zimbabwe at the five percent critical level and for Chile, Greece, and Thailand at about the 10 percent critical level. Set I benchmark assets fail to reject spanning for all DMs, albeit marginally for Italy. As in the case of the HK results, although both the MNCs and cross-listings contribute to a lower likelihood of rejecting spanning, it is the ability to invest in country funds and ADRs that allows U.S. investors to obtain international diversification benefits through home-made diversification.

Mean-variance spanning test statistics assess whether the shift in the efficient frontier is statistically significant. Bekaert and Urias suggest that economic significance can be assessed by evaluating the change in the Sharpe ratio. We measure the change in the slope of the line from the risk-free rate (30 day Eurodollar rates) to the tangency portfolio on the mean-variance frontier. The change in the slope corresponds to the change in the Sharpe ratio associated with the addition of a target foreign market index to our various benchmarks. These results are reported in Panel C of Table V. As expected, the change in the Sharpe ratio is inversely related to the p-values associated with the mean-variance spanning tests. The change in the Sharpe ratio is lower for nine of 16 markets as we move from set I to set II benchmark assets and lower still for five of 10 markets as we move from set II to set III benchmark assets.

Formally testing whether the shifts in the Sharpe ratios are statistically significant is difficult due to its unknown distribution. However, Bekaert and Urias, using Monte Carlo techniques, find that changes in Sharpe ratios of less than 0.057 are not statistically significant.[17] Based on this finding, we conclude that only in the case of Chile and Thailand are the changes in the Sharpe ratios significantly different from zero. Note that Chile and Thailand indices contribute as stand-alone international diversification; they may not do so in a broadly diversified portfolio including CFs and ADRs from other EMs and DMs. Based on these results, we conclude that the economic gains from international diversification that cannot be obtained with domes-

[17] The Bekaert and Urias study is based on 152 observations. Hence, applying their reported simulation results to our sample of 204 observations is approximation at best.

Table VI

Unconditional Correlations over Subperiods

To investigate time-series variation in the gains from diversification, we report monthly return correlation between each foreign market index and its most augmented diversification portfolio over three nonoverlapping subperiods of equal length. The most augmented diversification portfolio is either *AD*1 or *AD*2 obtained by augmenting portfolio *D*2 with each country's country fund or ADRs respectively. Diversification portfolios (*D*2) are based on stepwise regression procedures over three broad U.S. indices (Is), 12 U.S. industry indices (IIs), and 30 multinational corporations.

	Subperiods		
	Jan.'76–Nov.'81	Dec.'81–Nov.'87	Dec.'87–Dec.'93
Emerging markets			
Argentina	0.2214	0.2212	0.2306
Brazil	0.0860	0.1701	0.6729
Chile	0.4116	0.2330	0.5170
Greece	0.2355	0.4236	0.4539
India	0.2075	0.1881	0.4470
Korea	0.5998	0.5833	0.6930
Mexico	0.3331	0.6766	0.7178
Thailand	0.1951	0.2445	0.7254
Zimbabwe	0.3931	0.4094	0.3292
EM Average	0.2981	0.3499	0.5318
Developed markets			
Australia	0.6112	0.7199	0.5223
Canada	0.8210	0.8754	0.7676
France	0.5423	0.5886	0.5439
Germany	0.5562	0.5268	0.6676
Italy	0.4754	0.6300	0.4178
Japan	0.7419	0.6372	0.7956
U.K.	0.5821	0.7719	0.8325
DM Average	0.6185	0.6785	0.6461

tically traded securities are, with few exceptions, minimal. The gains appear to be greater for EMs than DMs; however, this may simply be due to the longer history of cross-listed securities in DMs.

IV. Evolution of Diversification Benefits

In the previous sections, gains from diversification are measured with the unconditional correlation and spanning relationship between diversification portfolios and market indices. We now turn to the issue of how gains from diversification have evolved through time. As a preliminary step, we compute the unconditional correlations between the most augmented diversification portfolios and their respective foreign market indices over three nonoverlapping subperiods of equal length. These results are reported in Table VI.

Some clear patterns emerge: There is a general tendency for the EM correlations to increase through time. The cross-sectional average of the correlations increases from 0.2981 (in the first period) to 0.3499 (in the second period) to 0.5318 (in the last period). The most noticeable increases occur in the cases of Brazil, India, Mexico, and Thailand. Not surprisingly, a country fund was introduced for these countries during the second or third subperiod. In contrast, the DM correlations have generally not increased through time. Only in the case of the U.K. do we observe an obvious increase in the correlation, albeit from an initially high level. The cross-sectional average correlations for DMs are 0.6185, 0.6781, and 0.6461 for the first, second, and the last subperiod, respectively. We now model the time variation in these correlations and relate it to important events such as the introduction of CFs and ADRs.

There are numerous ways in which time variation in correlation structures has been modeled. Bekaert and Harvey (1997) use the quadratic form multivariate GARCH model to assess international asset market integration. Karolyi and Stulz (1996) model the ex post cross-product residuals from regressions which function as reasonable proxies for conditional covariances. Longin and Solnik (1995) and Karolyi and Stulz augment the multivariate constant correlation model with various explanatory variables to investigate how index level correlation structures have changed over time. Although the estimates of time-varying correlations are dependent on the model, until recently there was no clear way to determine which model specification best represented a given process. The development of the generalized dynamic covariance (GDC) structure of Kroner and Ng (1995) enables one to test whether restrictions associated with the quadratic form or the constant correlation models are warranted. What follows is a brief overview of these models and the GDC model.

The bivariate GARCH model is of the following specification:

$$R_{I,t}|\Omega_{t-1} = \alpha_I + \epsilon_{It}, \tag{7}$$

$$R_{D,t}|\Omega_{t-1} = \alpha_D + \epsilon_{Dt}, \tag{8}$$

where $\alpha_I(\alpha_D)$ are constants and

$$\begin{bmatrix} \epsilon_{I,t} \\ \epsilon_{D,t} \end{bmatrix} |\Omega_{t-1} \sim N(0,H_t). \tag{9}$$

Ω_{t-1} is the information set that includes past observations of R_I and R_D, that is, $\Omega_{t-1} = (R_{I,t-1},R_{I,t-2},\ldots R_{D,t-1},R_{D,t-2},\ldots)$, and H_t is the conditional covariance matrix, that is,

$$H_t = \begin{bmatrix} h_{II,t} & h_{ID,t} \\ \cdot & h_{DD,t} \end{bmatrix}. \tag{10}$$

The quadratic form model of Engle and Kroner (1995) is of the following form:

$$H_t = C + A'\epsilon_{t-1}\epsilon'_{t-1}A + B \circ H_{t-1}, \tag{11}$$

where C, A, and B are 2×2 parameter matrices and $\circ$ denotes element-by-element multiplication. This specification is sufficiently general to allow time-varying variances, covariances, and correlations, and to impose cross-equation restrictions which help to ensure that H_t is positive definite. Note that equation (11) does not impose a quadratic specification for the parameters associated with the lagged conditional variance. We find that this specification aids model convergence greatly, but at a cost that the estimated parameters may not result in a positive definite H_t for all possible ϵ_t.[18] Given estimates of h_{II}, h_{ID}, and h_{DD}, the conditional correlation between the diversification portfolio and its corresponding market can be constructed as $\rho_t = h_{IDt}/\sqrt{h_{IIt}}\sqrt{h_{DDt}}$.

The constant correlation (CCOR) specification of H_t is of the following form:

$$H_t = V'_t G V_t, \tag{12}$$

where G is a symmetric correlation matrix and V_t is a diagonal matrix of conditional variances modeled as follows:

$$h_{II,t} = c_{11} + a_{11}\epsilon^2_{I,t-1} + b_{11}h^2_{I,t-1}, \tag{13}$$

$$h_{DD,t} = c_{22} + a_{22}\epsilon^2_{D,t-1} + b_{22}h^2_{D,t-1}. \tag{14}$$

To the extent that a CCOR structure is justified, time variation in the correlation is not significant. Comparisons between this specification and various alternatives can be made by estimating the following generalized dynamic covariance (GDC) model of Kroner and Ng (1998):

$$H_t = D'_t R D_t + C'_t \Phi C_t, \tag{15}$$

where

$$D_t = \begin{bmatrix} \sqrt{\theta_{11,t}} & 0 \\ 0 & \sqrt{\theta_{22,t}} \end{bmatrix}, \qquad R = \begin{bmatrix} 1 & \rho \\ \rho & 1 \end{bmatrix}, \tag{16}$$

$$C_t = \begin{bmatrix} 0 & \sqrt{\theta_{12,t}} \\ \sqrt{\theta_{12,t}} & 0 \end{bmatrix}, \qquad \Phi = \begin{bmatrix} 0 & \varphi \\ \varphi & 0 \end{bmatrix}, \tag{17}$$

[18] Sufficient conditions for H_t, as specified in equation (11) to be positive definite, are for H_o and B to be positive definite. Proof is available from the authors on request.

and $\theta_{ij,t} = h_{ij,t}$ are the i,j elements of H_t as stated in equation (11). Kroner and Ng show that the GDC specification allows for a greater degree of flexibility in modeling the dynamic behavior of higher moments. The extent to which the restrictions imposed by either CCOR or the quadratic specification are overly restrictive can be tested within the GDC framework. Nested within the GDC specification are the quadratic and constant correlation models. This is readily seen by expressing the variance equations of the GDC model as follows:

$$h_{iit} = \theta_{iit} \quad \text{for all } i. \tag{18}$$

$$h_{ijt} = \rho_{ij}\sqrt{\theta_{iit}}\sqrt{\theta_{jjt}} + \phi_{ij}\theta_{ijt} \quad \text{for all } i \neq j. \tag{19}$$

$$\theta_{ijt} = \omega_{ij} + a_i'\epsilon_{t-1}\epsilon_{t-1}'a_j + b_{ij} \circ H_{t-1}. \tag{20}$$

Thus, the GDC model reduces to the quadratic specification when $\rho = 0$ and $\phi = 1$, and to a CCOR specification when $\phi = 0$.[19]

A. Bivariate GARCH Diagnostics and Inferences

GDC bivariate GARCH(1,1) models between each country's return and its most augmented diversification portfolio are estimated. As reported in Table II, significant serial correlation exists in the return series for Chile, Greece, and Zimbabwe. Hence, an autoregressive term is added to the mean equations for these countries. In this section, we discuss various diagnostics results that serve to support our model specification. This is followed by two alternative tests for time varying correlation: the nested approach of Kroner and Ng (1998) and the nonnested approach of Vuong (1989).

A.1. Model Diagnostics

Table VII reports various model misspecification tests for EMs and DMs, respectively. The extent to which the GARCH(1,1) adequately captures time dependence can be inferred from Ljung–Box statistics for the residuals and the cross-product matrix of the residuals. Based on these standard measures, the GARCH(1,1) specification appears to be quite reasonable.

We also report conditional moment tests for the presence of deterministic trends in second moments as suggested by Kroner and Ng (1998).[20] The test results for EMs are reported in Panel B of Table VII. We document significant misspecification attributable to a deterministic trend in the conditional variance of market returns for Brazil, Chile, and India, and the

[19] Technically, in order for conditional variance equations of the GDC model to reduce to the CCOR specifications as stated in equations (13) and (14) the additional condition of $a_{12} = a_{21} = b_{12} = b_{21} = 0$ is needed. However, this is not required for constant correlation models with alternative conditional variance specifications.

[20] See Appendix A for additional details on conditional moment tests.

Table VII

Bivariate GARCH (1,1) Misspecification Tests for Emerging Markets

Misspecification diagnostics for bivariate GARCH(1,1) models with the GDC variance-covariance specification between each market and its most augmented diversification portfolio are reported below. $Q(\cdot)$ refers to Ljung–Box statistics based on 12 lags. Conditional moment tests for deterministic trends are based on Wooldridge (1990) and Kroner and Ng (1998). h_{11} refers to the conditional variance of the market return, h_{22} is the conditional variance of the most augmented diversification portfolio, and h_{12} is the conditional covariance between a market and its most augmented diversification portfolio. Market index returns and their most augmented diversification portfolio returns are denoted as r_t and $r_{dp,t}$ respectively. The conditional moment test statistics are distributed chi-squared with one degree of freedom.

Panel A: Ljung–Box Tests for Serial Correlation

Emerging Markets	Argentina	Brazil	Chile	Greece	India	Korea	Mexico	Thailand	Zimbabwe
$Q(r_t)$	12.034	9.211	18.837	22.362	20.109	4.528	17.490	21.860	26.98
$Q(r_{dp,t})$	8.246	18.056	18.756	15.286	10.621	8.577	11.206	18.683	11.00
$Q(h_{11,t})$	15.413	20.429	7.301	14.701	12.598	4.557	9.003	10.359	8.694
$Q(h_{12,t})$	25.327	11.454	18.156	7.433	25.425	6.058	14.998	7.261	5.250
$Q(h_{22,t})$	12.660	12.874	10.892	6.153	6.432	8.337	6.988	9.572	6.417
Developed Markets	Australia	Canada	France	Germany	Italy	Japan	U.K.		
$Q(r_t)$	11.494	14.074	9.084	16.679	18.597	6.788	11.282		
$Q(r_{dp,t})$	13.202	7.081	12.558	10.531	8.917	13.222	9.297		
$Q(h_{11,t})$	4.317	9.146	7.930	5.414	6.194	13.750	8.150		
$Q(h_{12,t})$	4.358	12.425	4.785	12.243	11.222	14.805	6.992		
$Q(h_{22,t})$	9.187	13.120	21.459	11.933	5.807	17.690	7.366		

Panel B: Robust Conditional Moment Test for Trend

Emerging Markets	Argentina	Brazil	Chile	Greece	India	Korea	Mexico	Thailand	Zimbabwe
h_{11}	0.049	8.573*	5.760*	2.644	7.351*	0.857	2.835	0.454	0.530
h_{12}	0.068	2.667	1.233	1.115	3.814	1.215	0.051	2.900	0.549
h_{22}	1.444	2.809	2.465	6.313*	4.001*	4.186*	1.465	3.847*	0.548
Developed Markets	Australia	Canada	France	Germany	Italy	Japan	U.K.		
h_{11}	0.184	2.569	0.334	3.133	0.178	10.567*	3.236		
h_{12}	3.389	3.883*	0.081	1.432	0.000	2.347	3.361		
h_{22}	1.652	1.510	0.019	0.078	0.050	0.004	5.552*		

*Denotes significance at α = 0.05.

conditional variance of most augmented diversification portfolio returns for Greece, India, Korea, and Thailand. Based on these results, we include trend terms in the appropriate variance-covariance equations for these particular cases.

The DM conditional moment tests for deterministic trends are reported in Panel B of Table VII. We find unaccountable trends in the conditional variance of market returns for Japan, the conditional covariance for Canada, and the conditional variance of the most augmented diversification portfolio returns for Japan. Trend terms are again included to account for these findings.

A.2. Statistical Significance of Time Variation in Correlations

To assess whether the time-variation in correlation implied from the quadratic form model estimates is statistically significant, we estimate generalized dynamic covariance structure GARCH(1,1) models as stated in matrix form in equation (15). The relevant parameters for our purpose are ρ and φ, which determine whether the GDC model reduces to either the constant correlation specification or the quadratic form specification. Specifically, $\phi = 0$ is consistent with constant correlation, whereas $\rho = 0$ and $\phi = 1$ are consistent with the quadratic specification.[21]

In Table VIII we report parameter estimates and 95 percent confidence intervals for ρ and ϕ corresponding to the most augmented diversification portfolio for a given country. In most instances, we are unable to justify reducing the GDC specification to either the CCOR or the quadratic specification. This is most apparent for Argentina, Chile, Greece, Mexico, and Zimbabwe where we fail to reject both relevant hypotheses with respect to ϕ (i.e., we fail to reject $\phi = 0$, which would be consistent with CCOR, and we fail to reject the hypothesis that $\phi = 1$, which would be consistent with the quadratic correlation specification). For Brazil and Korea, the results are consistent with quadratic specification and not the CCOR specification. For India and Thailand, the results are consistent with the CCOR specification. Overall, the inconclusiveness of these results suggests that the GDC specification cannot be reduced to either the CCOR or quadratic specification without significantly limiting the dynamic behavior of the second moments.

For DMs the evidence is again mixed. Only in the cases of Australia and Italy can we rule out constant correlation and only for Australia are we able to claim the quadratic form is appropriate. For Germany and Canada we fail to reject constant correlation. And for Japan, the U.K., and France the results are inconclusive. Overall, these results do not form a strong basis for reducing the GDC specification to either the CCOR or the quadratic specification.

The Kroner and Ng (1998) approach to testing nonnested models follows the lines of Atkinson (1969, 1970) in which competing models are combined into a larger model that nests both alternatives. Alternatively, Cox (1961,

[21] All reported t-statistics are based on the robust standard errors suggested by Bollerslev and Wooldridge (1992).

Table VIII

Selected Generalized Dynamic Covariance Estimates

Selected parameter estimates from the following generalized dynamic covariance (GDC) bivariate generalized autoregressive conditional heteroskedasticity (GARCH) model:

$$h_{ii} = \theta_{iit} \quad \text{for } i = 1,2$$

$$h_{12t} = \rho\sqrt{\theta_{11t}}\sqrt{\theta_{22t}} + \phi_{12}\theta_{12t}$$

$$\theta_{ijt} = \omega_{ij} + B_{ij} \circ h_{ijt-1} + a_i'\epsilon_{t-1}\epsilon_{t-1}'a_j \quad \text{for all } i, j,$$

where h_{ii} are the conditional second moments, ρ is the unconditional correlation, ϵ_{t-1} are the residuals from the mean equations, and ϕ_{12}, B_{ij}, a_i are parameters. In general, $\varphi \neq 0$ implies that the conditional correlation is time varying—that is, the GDC model reduces to a constant correlation model. $\rho = 0$ and $\phi = 1$ imply that the GDC model reduces to the quadratic form model of Engle and Kroner (1995). L.B and U.B. are 95 percent confidence interval lower and upper bounds respectively.

Country	L.B.	ρ	U.B	L.B.	ϕ	U.B.
Emerging markets						
Argentina	−1.7341	−0.7364	0.2616	−1.1766	1.8039	4.7844
Brazil	−0.7567	0.2378	2.2032	0.1964	0.8251	1.0676
Chile	−0.3322	0.2019	0.7360	−0.5521	0.4901	1.5321
Greece	−0.8126	0.5061	1.82664	−0.8567	0.4055	1.6681
India	−0.1582	0.2076	0.5733	−0.0674	0.3502	0.7678
Korea	−2.6462	0.1500	3.7238	0.044	0.6981	1.3521
Mexico	−2.1210	0.3523	2.8256	−2.072	0.4318	2.935
Thailand	0.3408	0.7918	1.2427	−0.0971	0.0614	0.2199
Zimbabwe	−0.1284	0.3700	0.8684	−0.4898	0.2949	1.0796
Developed markets						
Australia	−0.5694	0.2874	1.1441	0.2550	1.5638	2.8726
Canada	0.7177	0.9384	1.1591	−1.1200	−0.4737	0.2527
France	−1.1772	2.9572	7.0915	−13.6713	−5.8564	1.9585
Germany	−.2411	0.3182	0.8776	−0.0512	0.0129	0.0769
Italy	−3.9447	−2.2964	−0.6481	1.1199	3.1470	5.1742
Japan	−5.2986	−1.1207	3.0571	−2.4664	1.9752	6.4167
U.K.	−0.3285	0.7907	1.9099	−1.0338	0.4271	1.8881

1962) and White (1982) show that the standard likelihood ratio (LR) statistic for nonnested models, once properly normalized and centered, has a limiting normal distribution if one of the models is correctly specified. On the other hand, Vuong (1989) derives the asymptotic distribution of the LR statistic under more general conditions in which both models are correctly specified, only one model is correctly specified, or neither model is correctly specified. Hence, this approach is best viewed as a means of determining which specification is closer to the true model[22].

[22] See Appendix B for additional details on the Vuong (1989) normalized likelihood statistic.

Table IX

Normalized Likelihood Ratio Statistics

Normalized likelihood ratio statistic of Vuong (1989) for model selection tests between quadratic form and constant correlation specifications. LL are log likelihood values, and LR* are normalized likelihood ratio statistics that are standard Normal. LR* greater (less) than zero implies the time varying (fixed) correlation model outperforms the fixed (time invariant) correlation model.

Country	Quadratic LL	CCOR LL	LR*
Emerging markets			
Argentina	−1751.55	−1752.23	1.092
Brazil	−1528.72	−1565.50	4.483
Chile	−1412.12	−1414.28	1.022
Greece	−1372.56	−1377.48	1.433
India	−1211.60	−1222.72	2.388
Korea	−1400.62	−1404.57	1.411
Mexico	−1483.72	−1502.23	2.231
Thailand	−1275.80	−1284.55	2.113
Zimbabwe	−1351.11	−1355.51	1.321
Developed markets			
Australia	−1317.59	−1327.81	2.287
Canada	−1178.00	−1183.60	1.684
France	−1298.01	−1301.58	1.340
Germany	−1208.74	−1213.74	1.553
Italy	−1330.69	−1334.85	1.449
Japan	−1280.63	−1282.87	1.061
UK	−1258.77	−1267.84	2.152

The Vuong normalized likelihood ratio statistics for quadratic form versus constant correlation specifications for both EMs and DMs are reported in Table IX. The LR statistic is positive in all cases, suggesting that the time-varying correlation model outperforms the fixed correlation model. Statistical significance can be assessed by noting that the Vuong normalized likelihood statistic is asymptotically distributed standard normal with the following interpretation: For a given critical value c from the standard normal distribution, if the reported LR statistic is greater (less) than c ($-c$), we conclude that the time-varying model is (not) significantly better than the constant correlation model. With c equal to 1.96 (i.e., with 95 percent confidence), we conclude that the time varying model is significantly better than the fixed correlation model for four EMs (Brazil, India, Mexico, and Thailand), and two DMs (Australia and the U.K.). For all other cases, we cannot statistically discriminate between constant and time-varying correlation measures, although the time varying model outperforms in all cases. Hence, we find a higher incidence of time-variation in gains from diversification for EMs than DMs.

B. Explaining Time Variation in Gains from Diversification

The results of the previous section suggest significant time variation in the correlation between some (but not all) of the target market returns and their most augmented diversification portfolios. A more interesting issue is whether the time variation manifests itself in terms of trending behavior; that is, are the international diversification gains involving foreign- traded assets disappearing? To some extent the evidence reported in Table VI suggests that this is the case for EMs which have experienced an increase in home traded assets that are good substitutes for foreign-traded assets during the sample period. However, these subperiod correlations are only suggestive. To further investigate this issue, Figure 1 plots the time series of the conditional correlations between each target market return and the various diversification portfolios. Based on the time variation in correlations between foreign market returns and the S&P 500 index, these plots do not suggest that gains from foreign asset based diversification are disappearing over time. This result is consistent with a number of existing studies that report no clear trend in index level correlations. For example, Bekaert and Harvey (1997) find that the change in index correlations following market liberalizations are statistically significant but economically small. However, consistent with the subperiod unconditional correlations reported in Table VI, we find some evidence of positive trend for EMs where substitute assets are added during the sample period. There are, however, large and sustained departures from these trends. Though we may ascribe these fluctuations to estimation error, the more interesting possibility is that other factors besides substitute assets affect the diversification benefits to investing abroad. Hence, while the average gains from foreign asset based diversification may be negligible, there are periods (which are as likely to happen today as they were 10 years ago) when international markets provided meaningful diversification that could not be replicated at home. What follows is a brief discussion of events that explain some of the observed time variation in the conditional correlations.

B.1. Emerging Markets

A general decline in gains from investing abroad for emerging markets is evident in the cases of Brazil, Chile, India, Mexico, and Thailand. For Brazil, we note a large increase in conditional correlation during the first quarter of 1988 in anticipation of the introduction of the Brazil Fund. The average conditional correlation of 0.08 (for *AD*1 vis-à-vis IFCG) prior to the introduction of the CF rises to 0.66 in the following period. This increase reflects the events of May 1991 when Brazil officially removed all restrictions on foreign investments in preferred stock and allowed up to 49 percent ownership of voting common stock. It also reflects the events of June 1992 when the first Brazilian ADR was issued on the NYSE. (Note that the IFC considered approximately 60 percent of the IFCG investable at the end of 1993.)

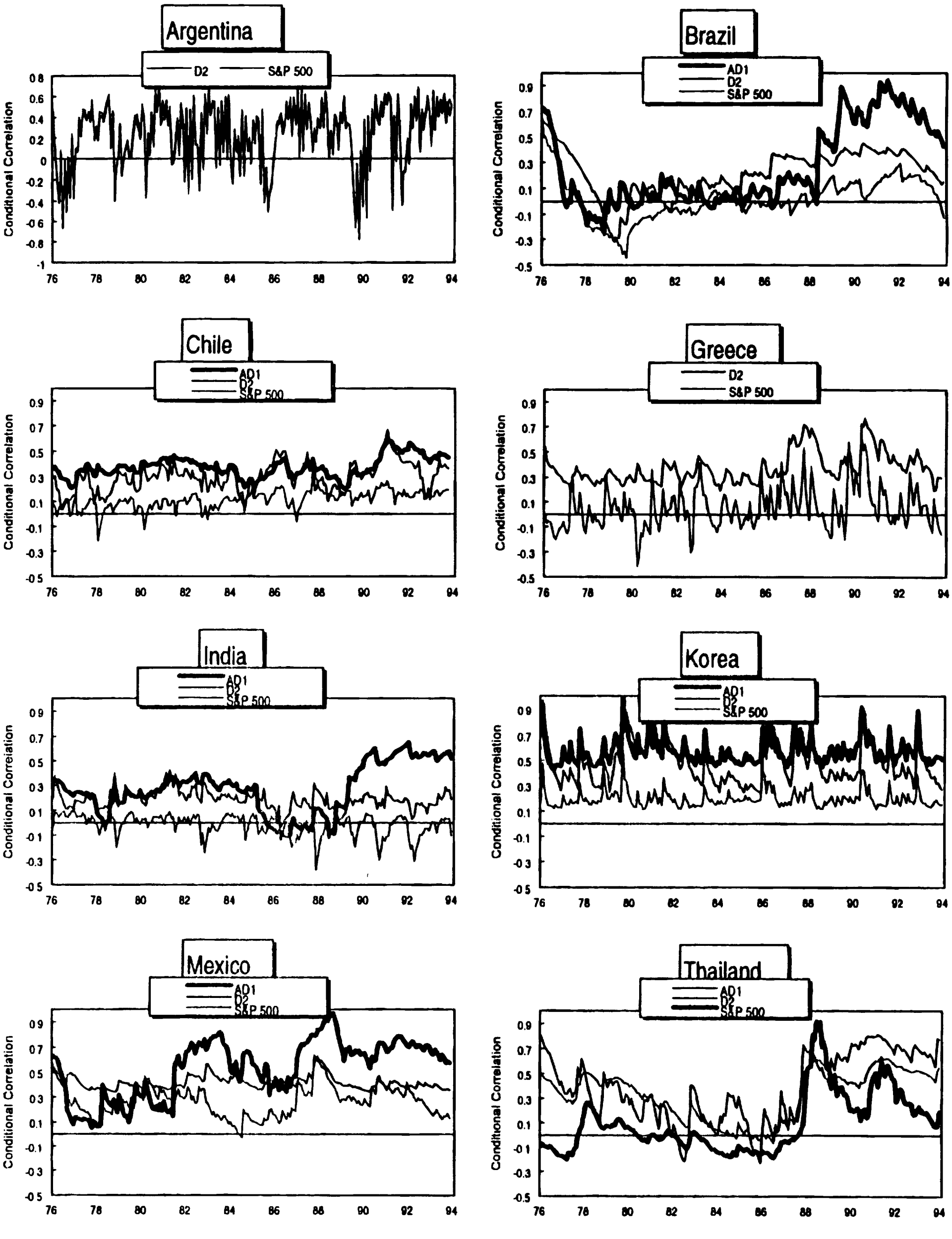

Figure 1. Conditional correlations. The conditional return correlations between foreign market indices and various diversification portfolios based on U.S. traded assets are illustrated. The conditional correlations are estimated using the generalized dynamic covariance model of Kroner and Ng (1998). D2 denotes the diversification portfolio selected from three broad-based U.S. indices, 12 industrial portfolios, and 30 multinational corporations by the stepwise procedure. AD1 are augmented diversification portfolios in which D2 is augmented using each country's country fund. AD2 are augmented portfolios in which AD1 is augmented by the country's representative ADRs.

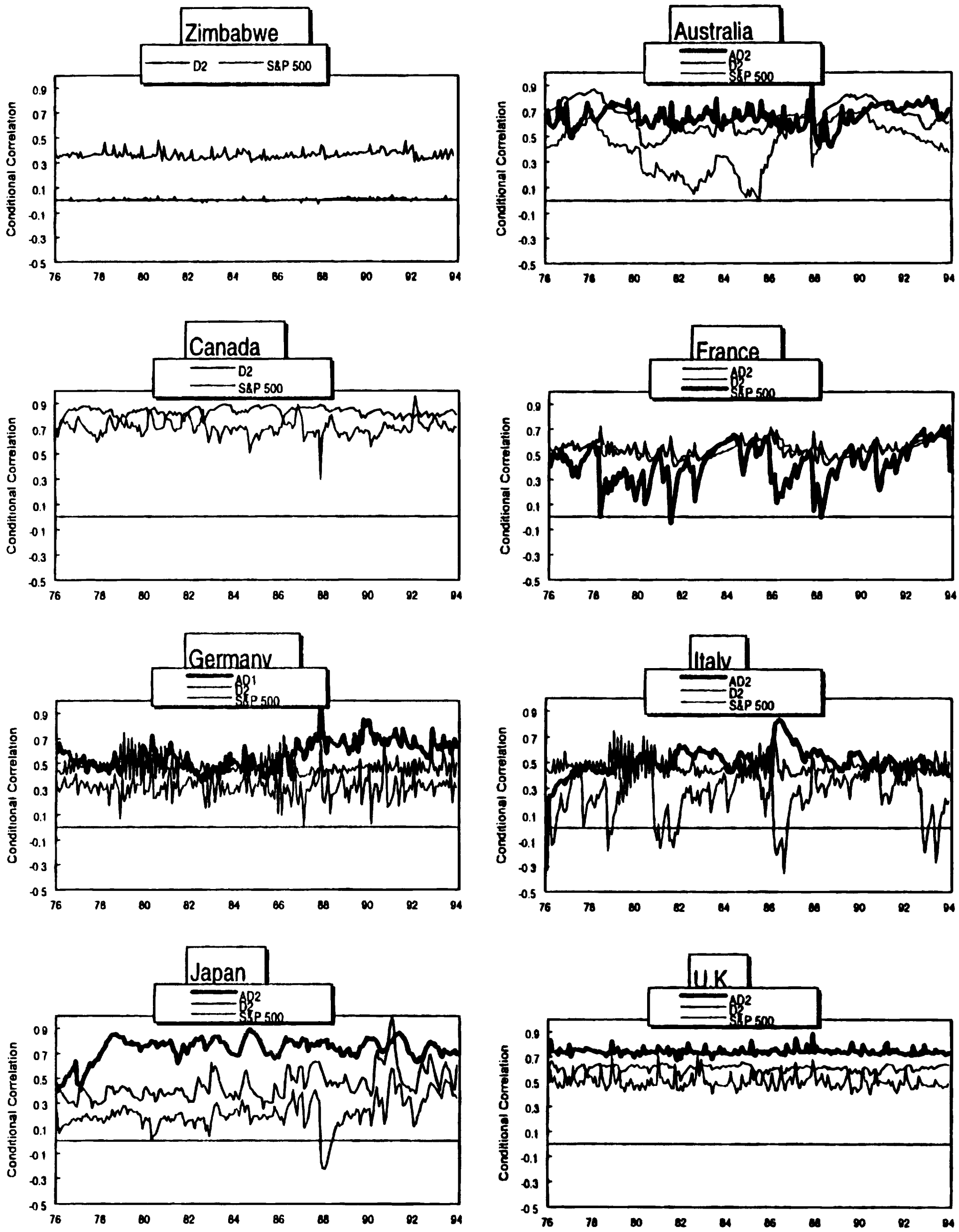

Figure 1. Continued.

In the case of Chile, the upward trend is not as apparent as in Brazil. The average conditional correlation associated with most augmented portfolio rises from 0.35 to 0.47 after the introduction of the country fund. Although the ability of U.S. investors to achieve home-made international diversification was further augmented by the listing of Chilean ADRs, the benefits of international diversification involving assets that actually trade in Chile have continued. This may be attributed to the largely uninvestable nature of the Chilean market due to entry and exit barriers such as currency controls, high taxes, and the holding period requirement. Even at the end of 1993, the IFC considered less than 25 percent of the Chilean market investable from the foreign investor perspective.

The average conditional correlation associated with the most augmented diversification portfolio for India increased from 0.19 to 0.48 following the listing of the India growth fund on the NYSE, suggesting that the benefits of diversification based on assets traded in India have declined over time. However, the absence of ADRs, the rather recent liberalization of foreign portfolio investments (1992), and low investability (approximately 20 percent at the end of 1993 according to the IFC) may suggest a continuation of benefits from investing in securities that trade in India. (Note that the precipitous decline toward the end of 1984 may reflect heightened political uncertainty, a unique national factor resulting from the assassination of Prime Minister Indira Gandhi.)

In the case of Mexico, the average conditional correlation associated with the most augmented diversification portfolio dramatically increases from 0.25 to 0.66 after the listing of the Mexico Fund in May 1981. An inexplicable downward drift from 1984 to 1987 is followed by a sharp increase reflecting the spillover effects of the market crash. The need to invest in securities that only trade in Mexico was reduced by the May 1989 removal of all restrictions on foreign direct purchases of nonbank stocks (the IFC considered approximately 90 percent of the Mexican market investable at the end of 1993), large foreign direct investments by U.S. MNCs, and the ability of U.S. investors to participate in Mexican stocks through ADRs (since 1991).

The Thai market shows a dramatic increase in average conditional correlation since the end of 1987. Indeed, it rises from 0.25 to 0.68 after the introduction of the Thai fund in February 1988. Although market access is restricted for foreign investors, the IFC considered approximately 30 percent of the Thai market investable at the end of 1993. This fact, together with U.S. MNC investments and the availability of a large number of open- and closed-end country funds worldwide, have contributed to declining benefits of investing in assets traded on the Thai market.

B.2. Developed Markets

The decline in diversification benefits of investing in securities that trade in DMs based on augmented diversification portfolios is attributable to the availability of multiple ADRs, large cross-border investments by MNCs based in the U.S, the availability of CFs, and the investability of MSCI indices. Furthermore, gains from international diversification based on the S&P 500 index seem more volatile than gains from home-made international diversification based on $D2$, $AD1$, and $AD2$ portfolios. Indeed, the discussions of time variation in broad market-wide index correlations in the case of DMs seems largely unwarranted.

V. Conclusions

This article investigates the ability of investors to mimic returns on foreign market indices with domestically traded securities, so that investing in assets that trade only abroad would not be necessary to obtain the benefits from international diversification. We construct diversification portfolios based on industry indices, MNCs, CFs, and ADRs. From 1976 to 1993, the monthly return correlations of these home-made diversification portfolios with foreign market indices are higher than those with the S&P 500 index. Indeed, correlations at the index level do not properly take into account the ability of U.S. investors to gain international diversification benefits through home-made international diversification.

The increasing availability of assets that represent claims on foreign-traded assets makes it possible to exhaust most of the diversification benefits by holding domestically traded assets. Mean-variance spanning test results provide strong evidence that gains beyond those attainable through home-made diversification are statistically and economically insignificant for 11 of the 16 markets (all seven of the DMs plus four of the EMs: Argentina, Brazil, Korea, and Mexico). Of the remaining five EMs, only Chile and Thailand deliver economically significant gains from investing overseas. Investors should continue to be aware of the foreign risks to which they are exposed, but they no longer need to trade abroad to obtain an internationally mean variance efficient portfolio. Substantial time variation in conditional correlations between overseas market returns and home-made diversification portfolios is apparent, especially in the case of EMs. These correlations vary over time differently from those computed with the S&P 500 and in ways consistent with changes in investment barriers, including new listings of CFs/ADRs, changes in rules governing foreign portfolio investments, and national events.

In summary, the use of return correlations at the market index level to infer gains from international diversification involving foreign-traded assets as it is commonly practiced in academia and the investment industry overstates the potential benefits. The gains must be measured beyond those attainable through home-made diversification. Indeed, the need to hold assets that trade only abroad to obtain international diversification benefits is fast disappearing.

Appendix A. Conditional Moment Tests

The conditional moment tests, suggested by Kroner and Ng (1998), are based on the comparison between the ex post cross-product matrix of the vector of residuals $\epsilon_{i,t}\epsilon_{j,t}$ and their estimates $h_{ij,t}$. Since the latter is the conditional expectation of the former, it should be the case that their differences are orthogonal to the available information set. Formally, we define the generalized residual to be this difference; that is,

$$u_{ij,t} \equiv \varepsilon_{i,t}\varepsilon_{j,t} - h_{ij,t}. \tag{A1}$$

If the model is properly specified, then

$$E_t(u_{ij,t}|\Omega_{t-1}) = 0, \tag{A2}$$

where Ω_{t-1} is the $t - 1$ information set. Define elements of Ω_{t-1} to be potential misspecification indicators. Denote $\lambda^t_{ij,t-1}$ as the misspecification indicator for trend and define it as follows:

$$\lambda^t_{ij,t-1} = t. \tag{A3}$$

The actual test statistic is based on the conditional moment test of Wooldridge (1990). Conditional moment tests have the attractive feature of being robust to model distribution assumptions while maintaining asymptotic efficiency. The test statistic is

$$C_{rcm} = \left[\left(\frac{1}{T}\right)\sum_{t=1}^{T} u_{ij,t}\Psi_{g,t-1}\right]^2\left[\left(\frac{1}{T}\right)\sum_{t=1}^{T} u^2_{ij,t}\Psi^2_{g,t-1}\right]^{-1}, \tag{A4}$$

where $\Psi_{g,t-1}$ is the residual from a regression of the misspecification indicator on the derivatives of $h_{ij,t}$ with respect to the parameters of the model under the null. Given certain regularity conditions, C_{rcm} is distributed chi-square with one degree of freedom.

This test statistic can be computed from two auxiliary regressions. The first regression involves regressing the generalized residual on the derivatives of $h_{ij,t-1}$ with respect to all the parameters of the model. The second regression involves regressing a vector of ones on the product of the residual from the first regression and the misspecification indicator variable. The test statistic is then formed as the product of the uncentered R^2 from the second regression multiplied by the number of observations.

Appendix B. Vuong LM Statistic model misspecification statistic

Formally, the Vuong LM statistic can be expressed as follows: Let $f(R_t{:}\theta)$ and $g(R_t{:}\gamma)$ denote the conditional densities of the quadratic form time-varying correlation model and the constant correlation model, respectively. The likelihood ratio statistic for time-varying correlation model against the constant correlation model is

$$LR_n(\theta,\gamma) = L^f(\theta) - L^g(\gamma) = \sum_{t=1}^{n}\frac{f(R_t:\theta)}{g(R_t:\gamma)}. \tag{B1}$$

Let $\hat{\omega} = \mathrm{Var}[LR_n(\theta,\gamma)]$ and estimated by

$$\hat{\omega} = \frac{1}{n}\sum_{t=1}^{n}\left[\frac{f(R_t:\theta)}{g(R_t:\gamma)}\right]^2 - \left[\frac{1}{n}\sum_{t=1}^{n}\frac{f(R_t:\theta)}{g(R_t:\gamma)}\right]^2. \tag{B2}$$

If $f(R_t:\theta) \neq g(R_t:\gamma)$, under the null hypothesis that the two models are equivalently close to the true model,

$$\frac{n^{-1/2}LR_n(\theta,\gamma)}{\hat{\omega}} \xrightarrow{d} N(0,1), \tag{B3}$$

where equivalence is defined in terms of the Kullback–Leibler Information Criterion (KLIC) in which the distance between the true distribution $H^o(R_t)$ and a specified model is measured as

$$KLIC(H^o;f) \equiv E[\log H(R_t)] - E[\log f(R_t;\theta)]. \tag{B4}$$

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