Stock Index Futures Trading and Volatility in International Equity Markets

HUSEYIN GULEN
STEWART MAYHEW

This article examines stock market volatility before and after the introduction of equity-index futures trading in twenty-five countries, using various models that account for asynchronous data, conditional heteroskedasticity, asymmetric volatility responses, and the joint dynamics of each country's index with the world-market portfolio. We found that futures trading is related to an increase in conditional volatility in the United States and Japan, but in nearly every other country, we found either no significant effect or a volatility-dampening effect. This result appears to be robust to model specification and is corroborated by further analysis of the relationship between volatility, trading volume, and open interest in stock futures. An increase in conditional covariance between country-specific and world returns

The authors would like to thank Mike Cooper, Jin-Chuan Duan, Robert Engle, Andrew Karolyi, Ken Kroner, Alexandra MacKay, John McConnell, Ken Ratliff, Joshua Rosenberg, and two anonymous referees. Earlier versions of this paper were presented at Virginia Tech, the Spring 1998 Chicago Board of Trade Research Symposium, the Istanbul Stock Exchange, Purdue University, the 1999 meetings of the American Finance Association and 1999 INFORMS meetings in Cincinnati. S. Mayhew would like to thank the Center for International Business Education and Research (CIBER) for funding this project.

Received October, 1999; Accepted February, 2000

- *Huseyin Gulen is a Ph.D. candidate in Finance at Krannert Graduate School of Management at Purdue University in West Lafayette, Indiana.*
- *Stewart Mayhew is an Assistant Professor in the Department of Banking and Finance at Terry College of Business at the University of Georgia in Athens, Georgia.*

The Journal of Futures Markets, Vol. 20, No. 7, 661–685 (2000)

at the time of futures listing is also documented. Jrl Fut Mark 20:661–685, 2000

INTRODUCTION

The world's first stock index futures contract was the Value Line contract, introduced by the Kansas City Board of Trade on February 24, 1982. Today, stock index futures and options trade in markets all over the world, with new contracts launched nearly every year. Table 1 reports launch dates for thirty nations that introduced stock index futures between 1982 and January, 1998. In addition, plans are underway for exchange-listed index futures in many other nations, including India, Indonesia, Czech Republic, Slovakia, Turkey, and others.

As exchange-traded stock index futures and other derivatives become more pervasive in the world's financial markets, it is increasingly important to understand the effect of derivatives trading on the underlying markets. The previous literature on the effects of stock index futures trading has focused primarily on developed markets, and it is unclear to what extent these results are applicable to less-developed markets. Moreover, the existing research has come to conflicting conclusions regarding the effect of futures trading on volatility. While some authors have found that volatility appears to increase with the introduction of futures, others have found no significant effect, and still others have found that volatility decreases.[1]

In this paper, we examine the time series properties of stock indexes in twenty-five countries in order to investigate the impact of stock index futures listing and subsequent trading activity on the volatility structure of the underlying cash market. We examined this issue in two ways. First, we tested for structural changes at the time of futures listing by comparing properties of the returns series before and after listing. Second, we tested whether volatility in the post-listing period is related to futures market volume and open interest. The results of both tests showed that futures trading is associated with increased volatility in the United States and Japan, but this was not the case in virtually every one of the other twenty-three countries. In some countries, there is no robust, significant effect, and in many others, futures trading is associated with lower volatility.

Many theories have been advanced elsewhere for how the introduction of futures might impact the volatility of the underlying market. As

[1]For a detailed summary of this literature, see surveys by Hodges (1992), Damodaran and Subrahmanyam (1992), Sutcliffe (1997), and Mayhew (1999).

pointed out by Hodges (1992), Mayhew (1999), and others, most of these theories predicted that volatility can increase or decrease with the introduction of futures depending on the underlying assumptions, or depending on the parameter values used in the models. Due to the large number of competing theoretical models with overlapping and ambiguous predictions, we are reluctant to interpret our results as favoring any particular model. Perhaps futures markets influence cash markets through multiple,

TABLE I

Launch Dates for Index Futures Contracts

Country	*Underlying Index*	*Launch Date*
United States	Value Line	24 February 1982
	S&P 500	21 April 1982
Australia	All Ordinaries	16 February 1983
United Kingdom	FT-SE 100	3 May 1984
Canada	TSE 300	16 January 1984
Brazil	BOVESPA	14 February 1986
Hong Kong	Hang Seng	6 May 1986
Japan (SIMEX)	Nikkei 225	3 September 1986
(Osaka)	OSE 50	9 June 1987
(Osaka)	Nikkei 225	3 September 1988
(Tokyo)	Topix	3 September 1988
New Zealand	Barclay Share	January 1987
Sweden	OMX	3 April 1987
Finland	FOX	2 May 1988
Netherlands	AEX	24 October 1988
France	CAC 40	9 November 1988
Denmark	KFX	7 December 1989
South Africa	All Share	30 April 1990
Switzerland	SMI	9 November 1990
Germany	DAX	23 November 1990
Chile	IPSA	December 1990
Spain	IBEX 35	14 January 1992
Austria	ATX	7 August 1992
Norway	OBX	4 September 1992
Belgium	BEL 20	29 October 1993
Italy	MIB 30	28 November 1994
Hungary	BSI	31 March 1995
Israel	Maof 25	27 October 1995
Malaysia	KLCI	15 December 1995
Korea	KOSPI 200	3 May 1996
Portugal	PSI-20	20 June 1996
Russia	RTS	March 1997
Venezuela	IBC	5 September 1997
Poland	WIG20	16 January 1998
Greece	FTSE/ASE-20	27 August 1999

Initial trading dates for various Index Futures contracts. Sources: information published by the individual exchanges, telephone conversations with exchange officials, and Futures Industry Association Fact Book. It should be noted that the trading of Japanese stock index futures initiated in Singapore.

offsetting channels. Perhaps futures play a more-important stabilizing role in markets that lack alternative stabilization mechanisms.

Our paper contributes to the existing literature in several ways. To the best of our knowledge, it is the most comprehensive examination to date on the impact of stock index futures on cash markets. We examined a much broader cross section of international futures introductions than any prior study, and this study is the first to examine the impact of futures markets in emerging economies. Examining the existing literature, as summarized by Mayhew (1999), one is left with the impression that the introduction of stock index futures is equally likely to be associated with increasing or decreasing volatility. In examining the results from twenty-five different countries, we saw a pattern emerging—in most cases, futures markets are associated with decreasing volatility.

In addition, we believe that this paper improves on the methodology used in prior studies. Like Lee and Ohk (1992) and Antoniou, Holmes, and Priestly (1998), we used a framework that allows for generalized autoregressive conditional heteroskedasticity (GARCH), but (wherever possible) we used considerably larger sample periods. This is important given the difficulty of obtaining reliable GARCH estimates in small samples.[2] Moreover, rather than selecting a GARCH specification ad hoc, we performed various specification tests to determine the appropriate model, and then tested the robustness of our results using several alternative GARCH models. In addition, we examined the properties of excess returns over the world-market index, which enabled us to avoid attributing worldwide price movements, such as the crash of October 1987, to the listing of futures in the local market.

To our knowledge, this also is the first paper to examine the relationship between volatility and futures trading volume and open interest for a large cross section of markets. We found that some, but not all, of the results reported in the literature are robust across countries. In examining this issue, we also generalized the methodology of Bessembinder and Seguin (1992) to a GARCH-based framework.

Finally, we believe this to be the first paper to examine the joint dynamics of country-specific and world-market returns for a cross section of countries using a multivariate GARCH framework. This framework allowed us to test whether the introduction of futures impacted the conditional covariance between country and world returns, a measure of the country's integration with world markets. We found that futures markets

[2]Engle and Mezrich (1995) suggested using at least eight years of daily data for proper GARCH estimation.

appeared to help countries become more integrated with the world market.

The basic approach of our analysis is as follows. First, we examined the impact of futures introduction on volatility using a modification of the Generalized Auto-Regressive Conditional Heteroskedasticity model suggested by Glosten, Jagannathan, and Rundle (GJR-GARCH; 1993). To test for the impact of futures trading, we incorporated a multiplicative dummy variable in the conditional variance equation. We checked the robustness of our results using various alternative specifications. Next, using a technique similar to the one employed by Bessembinder and Seguin (1992), we decomposed the trading volume and open-interest time series into permanent and temporary components, and then we tested how these components affected volatility by inserting them into the conditional volatility equation. Finally, we analyzed the joint dynamics of each country with world-market portfolios using the bivariate GARCH specification advanced by Engle and Kroner (1995), commonly known as the BEKK model.[3] This richer framework allowed us to more carefully control for movements in global markets. It also allowed us to test whether the conditional covariance between a country's return and the world-market return changed with futures listing.

The remainder of this paper proceeds as follows. The next section describes the data we used in our analysis. We then describe the univariate GJR-GARCH framework we used to analyze the data, verifying that conditional heteroskedasticity is present in the returns in all twenty-five countries, and we tested whether volatility is higher or lower after the introduction of futures trading. We also discuss the robustness of the results to model specification. We follow up by examining whether conditional volatility is related to the temporary and permanent components of open interest and trading volume. We then present our analysis of the joint dynamics of country-specific and world returns using the BEKK bivariate model.

DATA

Daily stock-market-index data were obtained from Datastream[4] for twenty-five of the thirty-one nations listed in Table I. Russia, Venezuela, Poland, and Greece, which listed futures after July 1996, were excluded because, in our judgment, there was insufficient data in the post-event

[3]The acronym refers to Baba, Engle, Kraft, and Kroner, the original developers of the model.
[4]Datastream International, Inc.

period to draw any meaningful conclusions. Brazil and New Zealand were excluded due to lack of data.

For twenty countries, time-series data were obtained for the stock index underlying the first equity futures contract listed in the respective country. For the United States, we used data on the more popular S&P 500 Index instead of the Value Line Index. In some cases, very little data existed for the underlying index prior to the futures listing date, often because the index was designed specifically to underlie the futures contract and didn't exist very long prior to the introduction of the futures. Given the high correlations typically observed between different indices on the same market, we did not believe this to be a major problem. To illustrate, for Norway, we used data on the Oslo Stock Exchange (OSE) General stock index instead of the Oslo Bors Index (OBX) due to the lack of data on the OBX index prior to the listing date. Over a recent subsample for which data are available on both indices, we calculated a correlation of .96 between them. Likewise, in Finland, we used the Helsinki Stock Exchange General Index (HEX) instead of the Finnish Options Index (FOX), and in the UK and Italy, we used market indexes calculated by Datastream due to insufficient daily data in the pre-event samples.[5] In Japan, we used the first introduction of Nikkei 225 futures on the Singapore International Monetary Exchange (SIMEX) as our event date.

Daily data were obtained on Datastream's World Market Index from January 2, 1973 through December 31, 1997. For each country, we used all the stock-index data available on Datastream between 1973 and 1997. In most cases, data were only available for part of this period. The time periods covered by our index data for each country, along with the number of daily observations in the pre- and post-event subsamples, are reported in Table II.

In addition, we were able to collect daily contract volume and open-interest data for seventeen of the countries in our sample. In most cases, these data were obtained from Datastream. Data from the Canadian market were provided by the Toronto Stock Exchange.

VOLATILITY EFFECTS OF FUTURES LISTING

Empirical Framework for Univariate Modeling

We begin our analysis by modeling the time series of excess country returns net of the world-market portfolio as a univariate GARCH process.

[5]In the case of the UK, weekly data are available for a large window prior to futures listing, but we felt that, in order to make the result comparable to the other countries, we should stick to daily data.

TABLE II

Data Periods

Country	*Data Period*	*Obs. Pre-*	*Obs. Post-*
Australia	2 January 1980–31 December 1997	779	3572
Austria	20 November 1987–31 December 1997	1162	1334
Belgium	2 January 1990–31 December 1997	941	1037
Canada	2 January 1973–31 December 1997	2740	3516
Chile	2 January 1987–31 December 1997	879	1741
Denmark	10 December 1979–31 December 1997	2476	2037
Finland	2 January 1987–31 December 1997	333	2424
France	9 July 1987–31 December 1997	330	2270
Germany	21 November 1977–31 December 1997	3215	1771
Japan	4 January 1980–31 December 1997	2098	2298
Hong Kong	2 January 1973–31 December 1997	3263	2888
Hungary	2 January 1991–31 December 1997	1056	674
Israel	2 January 1992–31 December 1997	928	527
Italy	2 January 1973–31 December 1997	5507	780
Korea	3 January 1990–31 December 1997	1540	398
Malaysia	2 January 1980–31 December 1997	3902	508
Netherlands	3 January 1983–31 December 1997	1402	2313
Norway	3 January 1983–31 December 1997	2418	1333
Portugal	1 January 1993–31 December 1997	853	376
South Africa	10 April 1985–31 December 1997	1136	1891
Spain	6 January 1987–31 December 1997	1238	1501
Sweden	2 January 1986–31 December 1997	311	2694
Switzerland	1 July 1988–31 December 1997	590	1792
United Kingdom	2 January 1973–31 December 1997	2871	3485
United States	2 January 1973–31 December 1997	2340	3967

Description of the data period used for each country, including the number of daily return observations before and after stock index futures listing.

This framework is parsimonious, which allowed us to capture many of the salient features of the data, and to partially account for movements in the world market in a model with relatively few parameters. Later, we would estimate a multivariate GARCH model that would allow us a richer model of the joint dynamics of country-specific and world-market returns.

Following Pagan and Schwert (1990) and Engle and Ng (1993), the first step in our univariate GARCH analysis was to remove from the time series any predictability associated with lagged returns or day-of-the-week effects. For each country, the following regression was estimated:

$$R_t - R_{Wt} = a_0 + a_1 R_{Wt-1} + \sum_{j=2}^{5} a_j DAY_j + u_t \tag{1}$$

where R_t is the daily return on the country's stock index and R_{Wt} is the daily return on the World Market Index on day t, R_{Wt-1} is the lagged return on the World Market Index, and DAY_j are day-of-the-week dummies for Tuesday through Friday.

We used the excess return relative to the World Market Index as our dependent variable and the lagged World Market Index return as an independent variable in an effort to remove the effect of worldwide price movements on volatility.[6] It should be noted that because of differences in time zones, different markets line up differently with the world-market return. This makes it difficult to compare directly the coefficients of the first-stage regression. For example, if the U.S. market is influenced by Asian markets, this will be reflected through the contemporaneous market return on the left-hand side of the regression equation. On the other hand, if the Asian markets are influenced by the US, this will be reflected through the lagged market portfolio.

Regression results are reported in Table III. It should be noted that the coefficients on the day-of-the-week dummies in this model, which is based on excess returns relative to the world market, should not be compared directly to day-of-the-week dummies based on models of raw returns.

To correct for any remaining predictability, and to correct for spurious autocorrelation induced by nonsynchronous trading,[7] we performed the usual autocorrelation adjustment:

$$u_t = b_0 + \sum_{j=1}^{5} b_j u_{t-j} + \varepsilon_t. \tag{2}$$

Table IV reports parameter estimates for this equation. Following Engle and Ng (1993), we reported Ljung–Box test statistics for twelfth-order serial correlation both in the residuals and their squares. The Ljung–Box statistics reported for the residual levels revealed that the regression model removes serial correlation in the stock-return series in most of the countries. At the 5% (1%) significance level, there is no serial autocorrelation left in 9 (5) of 25 countries, and in several other cases, the test statistic is only marginally significant. This suggests that the adjustment procedure removed the predictable part of the return series for most of the countries. The Ljung–Box test statistics for the squared residuals are highly significant in all cases, which is consistent with the existence of

[6]In separate tests not reported here, we included contemporaneous world returns on the right-hand side, allowing each country to have its own beta with respect to the world portfolio. These results, which are available on request, are similar to those reported here. We elected to use the current formulation because when contemporaneous variables are included in the first-stage regression, the GARCH volatility equation cannot be interpreted strictly as a conditional volatility.
[7]See Scholes and Williams (1977), Lo and MacKinlay (1988), Nelson (1991).

TABLE III

Coefficients from the First-Stage Regression

Country	*Intercept*	R_{Wt-1}	*Tuesday*	*Wednesday*	*Thursday*	*Friday*
Australia	0.0371	0.3758*	−0.1204*	−0.0890	−0.0508	−0.0243
	(0.03)	(0.02)	(0.05)	(0.05)	(0.05)	(0.05)
Austria	0.0398	0.1876*	−0.0824	−0.0563	−0.0206	−0.0156
	(0.05)	(0.04)	(0.08)	(0.08)	(0.08)	(0.08)
Belgium	−0.0324	−0.0988*	0.0277	0.0515	0.0822	0.0423
	(0.04)	(0.03)	(0.06)	(0.06)	(0.06)	(0.06)
Canada	−0.0328	−0.0252*	0.0242	0.0190	0.0407	0.0617*
	(0.02)	(0.01)	(0.03)	(0.03)	(0.03)	(0.03)
Chile	−0.0204	−0.0920*	0.0618	0.1528	0.1366	0.3528*
	(0.07)	(0.04)	(0.09)	(0.09)	(0.09)	(0.09)
Denmark	0.0774*	0.0316	−0.0854	−0.0986	−0.0349	−0.0680
	(0.04)	(0.02)	(0.05)	(0.05)	(0.05)	(0.05)
Finland	0.0171	0.0571	−0.0725	−0.0306	0.0480	0.0167
	(0.05)	(0.03)	(0.07)	(0.07)	(0.07)	(0.07)
France	−0.1388*	−0.0669*	0.1907*	0.1727*	0.1973*	0.1479*
	(0.05)	(0.03)	(0.07)	(0.07)	(0.07)	(0.07)
Germany	0.0018	0.0070	−0.0074	−0.0106	0.0078	−0.0341
	(0.03)	(0.02)	(0.05)	(0.05)	(0.05)	(0.05)
Hong Kong	−0.1182*	0.3306*	0.0893	0.2151*	0.0638	0.2121*
	(0.05)	(0.03)	(0.08)	(0.08)	(0.08)	(0.08)
Hungary	0.0940	0.3063*	−0.0519	0.0740	−0.1148	−0.0072
	(0.08)	(0.06)	(0.11)	(0.11)	(0.11)	(0.11)
Israel	−0.0261	0.0977	0.1360	0.0891	0.1029	−0.0052
	(0.10)	(0.08)	(0.14)	(0.14)	(0.14)	(0.14)
Italy	−0.0345	−0.0395	−0.0255	0.0126	0.1258*	0.1427*
	(0.04)	(0.02)	(0.05)	(0.05)	(0.05)	(0.05)
Japan	−0.0232	0.1249*	−0.0352	0.1518	0.0221	−0.0263
	(0.03)	(0.02)	(0.05)	(0.05)	(0.05)	(0.05)
Korea	−0.1100	0.0689	−0.0134	0.1928	−0.0294	0.0696
	(0.09)	(0.06)	(1.12)	(1.12)	(1.12)	(1.12)
Malaysia	−0.1248*	0.1102*	−0.0055	0.1473*	0.1684*	0.1975*
	(0.05)	(0.03)	(0.07)	(0.07)	(0.07)	(0.07)
Netherlands	−0.0325	−0.0676*	0.0886	0.0894	−0.0007	0.0701
	(0.04)	(0.02)	(0.06)	(0.06)	(0.06)	(0.06)
Norway	−0.0054	0.1130*	−0.0183	−0.0103	0.0422	0.1054
	(0.04)	(0.03)	(0.06)	(0.06)	(0.06)	(0.06)
Portugal	0.0026	0.0822	0.0076	0.0784	0.0665	0.0306
	(0.05)	(0.04)	(0.08)	(0.08)	(0.08)	(0.08)
South Africa	−0.0348	0.0238	0.0083	0.1459*	0.0765	−0.013
	(0.05)	(0.03)	(0.07)	(0.07)	(0.07)	(0.07)
Spain	0.1297*	0.1270*	−0.1283*	−0.2286*	−0.1434*	−0.1060
	(0.05)	(0.03)	(0.07)	(0.07)	(0.07)	(0.07)
Sweden	−0.0159	0.0663*	0.0228	0.0245	0.0667	0.0637
	(0.05)	(0.03)	(0.07)	(0.07)	(0.07)	(0.07)
Switzerland	−0.0123	−0.0359	0.0078	0.0992	0.0705	0.0484
	(0.04)	(0.03)	(0.06)	(0.06)	(0.06)	(0.06)

TABLE III (Continued)

Coefficients from the First-Stage Regression

Country	*Intercept*	R_{Wt-1}	*Tuesday*	*Wednesday*	*Thursday*	*Friday*
United Kingdom	−0.0709*	0.0228	0.1513*	0.0589	0.0620	0.1269*
	(0.03)	(0.02)	(0.04)	(0.04)	(0.04)	(0.04)
United States	0.0319	−0.1647*	−0.0027	−0.0430	−0.0387	−0.0350
	(0.02)	(0.01)	(0.03)	(0.03)	(0.03)	(0.03)

Results from the first-stage regression of country-specific returns on lagged world-market index and day-of-the-week dummies. The model is

$$R_t - R_{Wt} = a_0 + a_1 R_{Wt-1} + \sum_{j=2}^{5} a_j DAY_j + u_t,$$

where R_t is the daily return on the country's stock index, R_{Wt} is the daily return on the World Market Index on day t, R_{Wt-1} is the lagged return on the World Market Index, and DAY_j are day-of-the-week dummies for Tuesday through Friday. Standard errors are shown in parentheses. An asterisk indicates statistical significance at the 5% level.

time-varying volatility of index returns in all countries. We took this as evidence that some type of GARCH specification is necessary to properly model index returns in all countries.

Using $\{\varepsilon_t\}$ as our new return series, we proceeded to test for the effect of futures introduction on the conditional volatility of the spot market using various GARCH specifications.

Volatility Effect of Futures Introduction

Having demonstrated the need to account for conditional heteroskedasticity in returns, we now address the issue of futures listing using a GARCH model. In GARCH modeling, the residuals ε_t arising from estimating the autoregression equation [eq. (2)] are assumed to be distributed $N(0, h_t)$, or alternatively $\varepsilon_t = \epsilon_t \sqrt{h_t}$, where ϵ_t has a conditional distribution that is $N(0, 1)$, and the conditional volatility h_t depends on the GARCH specification.

In order to determine which GARCH specification we should use in our analysis, we conducted extensive tests to see which form of the conditional volatility equation best seemed to model the returns data. The results of these tests are not reported here, but nearly are identical to those reported in Gulen and Mayhew (1999), where we used a slightly different formulation for the first-stage regression. The main focus of this analysis was to determine whether we should use the symmetric GARCH model of Bollerslev (1986), in which positive and negative shocks of equal magnitude have the same effect on subsequent volatility, or a model

TABLE IV

Coefficients from the Residual Autoregression

Country	*Constant*	b_1	b_2	b_3	b_4	b_5	*LBQ(12) (Levels)*	*LBQ(12) (Squares)*
Australia	0.0001*	0.1012	−0.0134	0.0253	−0.0181	0.0232	8.4	630.27**
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)**		
Austria	−0.0001	0.2276*	−0.0210	−0.0246	−0.0125	0.0310	20.6**	438.65**
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)**		
Belgium	0.0000	0.0274	0.066*	0.0399	−0.0026	−0.0358	6.9	271.55**
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)**		
Canada	0.0000	0.0808*	−0.0023	0.0100	0.0272	0.0104	7.9	639.94**
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)**		
Chile	−0.0001	0.1981*	−0.0534*	−0.0394*	0.0598*	−0.0050	7.1	176.466**
	(0.03)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)**		
Denmark	−0.0000	0.0947*	−0.0037	−0.0017	−0.0402*	0.0051	7.4	41.96**
	(0.02)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)**		
Finland	0.0001	0.1764*	0.0001	0.0052	0.0153	0.0220	10.2	340.722**
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)**		
France	0.0000	−0.0797*	0.0533*	−0.0059	0.0016	0.0140	15.5*	917.54**
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)**		
Germany	0.0000	−0.0307*	−0.0001	0.0200	−0.0164	0.0141	4.8	438.591**
	(0.02)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)**		
Hkong	−0.0000	0.0432*	−0.0237	0.0601*	−0.0093	−0.0183	18.9**	1083.33**
	(0.02)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)**		
Hungary	0.0003	0.1301*	0.0531*	0.0070	−0.0113	0.0141	33.8**	727.973**
	(0.04)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)**		
Israel	−0.0001	−0.0174	−0.0365	−0.0207	0.0117	−0.0581*	16.3*	223.239**
	(0.04)	(0.03)	(0.03)	(0.03)	(0.03)	(0.03)**		
Italy	0.0000	0.1592*	−0.0717*	0.0392*	0.0013	−0.0010	12.9	2050.29**
	(0.02)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)**		
Japan	−0.0000	−0.0792*	−0.0668*	0.0229	−0.0021	−0.0061	8.1	801.76**
	(0.01)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)**		
Korea	0.0004	−0.0021	−0.0422	−0.0219	−0.0372	−0.0528*	11.9	1009.59**
	(0.04)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)**		
Malaysia	0.0000	0.1209*	0.0070	−0.0114	0.0065	0.0105	10.9	1352.07**
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)**		
Netherlands	0.0001	−0.0971*	0.0274	0.0257	−0.0166	0.0453*	21.4**	780.3**
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)**		
Norway	0.0000	0.1182*	−0.0068	−0.0129	−0.0466*	0.0373*	14.9*	501.91**
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)**		
Portugal	−0.0000	0.1516*	0.0490	−0.0522	−0.0074	−0.0127	13.8	154.161**
	(0.02)	(0.03)	(0.03)	(0.03)	(0.03)	(0.03)**		
South Africa	−0.0000	0.0834*	−0.0252	0.0018	−0.0218	−0.0308	9.6	38.1064**
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)**		
Spain	−0.0001	0.0915*	0.0149	−0.0540*	0.0364	−0.0186	18.4*	782.843**
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)**		
Sweden	0.0002	0.0555*	0.0252	−0.0135	0.0106	0.0178	11.6	715.95**
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)**		
Switzerland	0.0000	−0.0200	0.0089	−0.0404	0.0052	−0.0248	7.2	56.3926**
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)**		

TABLE IV (Continued)

Coefficients from the Residual Autoregression

Country	*Constant*	b_1	b_2	b_3	b_4	b_5	*LBQ(12) (Levels)*	*LBQ(12) (Squares)*
United Kingdom	−0.0000	0.0761*	0.0107	0.0101	0.004	−0.0099	13.5	4791.4**
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)**		
United States	−0.0000	−0.1303*	−0.0076	−0.0187	−0.0513*	−0.0097	20.2**	1252.86**
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)**		

Estimated parameters of the residual autoregression

$$u_t = b_0 + \sum_{j=1}^{5} b_j u_{t-j} + \varepsilon_t.$$

where u_t is the residual from regression 1. Ljung-Box statistics testing for 12th order serial autocorrelation in ε and ε^2 are also reported. Standard errors are shown in parentheses. * indicates statistical significance at the 5% level. ** indicates statistical significance at the 1% level.

where positive and negative shocks can have different effects. We tested the symmetric model and three alternative asymmetric models, including the asymmetric GARCH model (GJR-GARCH; Glosten, Jagannathan, & Runkle, 1993), the nonlinear GARCH model (NGARCH; Engle & Ng, 1993), and the exponential GARCH model (EGARCH; Nelson, 1991). Specification tests indicated that these asymmetric models fit the data better then the symmetric GARCH model, with the GJR-GARCH performing marginally better than the others. Therefore, we based our main analysis on the GJR-GARCH model.

In this model, the conditional volatility equation takes the form:

$$h_t = \alpha_0 + \alpha_1 h_{t-1} + \alpha_2 \varepsilon_{t-1}^2 + \alpha_3 \max(0, -\varepsilon_{t-1})^2.$$

In order to estimate the impact of futures introduction, we interacted the GJR-GARCH conditional-volatility equation with a multiplicative dummy, as follows:

$$h_t = (1 + \alpha_M D_t)[\alpha_0 + \alpha_1 h_{t-1} + \alpha_2 \varepsilon_{t-1}^2 + \alpha_3 \max(0, -\varepsilon_{t-1})^2],$$

where D_t takes on a value of 0 prior to futures introduction and a value of 1 after futures introduction. A significant negative-parameter estimate for α_M would indicate an decrease in the volatility associated with futures introduction.[8]

[8]In estimating this model, proper inequality constraints were placed on the parameters of the conditional variance equations to ensure the stationarity of the GARCH models and the nonnegativity of the conditional variances. Constrained maximum-likelihood estimates of these parameters were computed using the BHHH algorithm.

Results are reported in Table V. Defining statistical significance at the 5% level, we found that out of 25 countries, the coefficient α_M is positive and significant only for the United States and Japan, indicating an increase in conditional volatility associated with futures introduction in these countries. On the other hand, α_M is significantly negative for Australia, Austria, Belgium, Chile, Denmark, France, Germany, Hong Kong, Israel, Italy, Malaysia, Netherlands, Norway, South Africa, Switzerland, and the United Kingdom, a total of 16 countries. There is no significant effect in the remaining seven countries. The experience in the United States and Japan appears to be the exception, not the rule.

At this point, we would like to interject an important caveat. Since we have not analyzed the data for countries with rapidly maturing financial institutions that did not list futures, it is possible that this decreasing volatility reflects other developments that tend to coincide with the introduction of stock index futures.

Robustness Checks and Additional Tests

This section summarizes various other specifications we tested, but did not report here because the results are substantively similar to those in Table V. The results of all these tests are available upon request.

Another approach to analyzing the effect of futures introduction on volatility is to put an additive dummy variable into the GARCH equation:

$$h_t = \alpha_0 + \alpha_1 h_{t-1} + \alpha_2 \varepsilon_{t-1}^2 + \alpha_3 \max(0, -\varepsilon_{t-1})^2 + \alpha_A D_t.$$

We repeated our analysis using this additive dummy specification for the GJR-GARCH model. In addition, we examined the standard GARCH(1,1) model of Bollerslev (1986), the nonlinear NGARCH, and the exponential EGARCH. In some specifications, the volatility increase in Japan lost its significance, and in some specifications, other countries, including Canada, Hungary, and Korea, exhibit significant increases in conditional volatility. By and large, however, all of these specifications yielded results supporting the same conclusion: outside of the United States and Japan, volatility tends to decrease with futures listing, or at least has remained unchanged.

In our analysis, we accounted for movements in the world index simply by estimating the dynamics of excess returns of the country index relative to the world market. Implicitly, this assumes that the β of each country's return with respect to the world is 1. In other results, reported in an earlier version of this paper, we also estimated a model where each

TABLE V

Effect of Futures Introduction on GJR-GARCH Volatility

Country	α_0	α_1	α_2	α_3	α_M
Australia	0.1311*	0.7463*	0.1910*	−0.1028*	−0.0443*
	(0.016)	(0.022)	(0.009)	(0.015)	(0.013)
Austria	0.0969*	0.7511*	0.1991*	−0.0041	−0.0264*
	(0.012)	(0.018)	(0.013)	(0.021)	(0.011)
Belgium	0.0450*	0.8577*	0.0864*	−0.0112	−0.0308*
	(0.012)	(0.027)	(0.017)	(0.020)	(0.011)
Canada	0.0153*	0.8423*	0.1405*	−0.0261*	0.0012
	(0.002)	(0.007)	(0.008)	(0.009)	(0.004)
Chile	0.2407*	0.7074*	0.2625*	−0.0337	−0.0837*
	(0.032)	(0.021)	(0.021)	(0.020)	(0.013)
Denmark	0.0512*	0.9242*	0.0586*	−0.0284*	−0.0362*
	(0.007)	(0.007)	(0.006)	(0.006)	(0.006)
Finland	0.0744*	0.8446*	0.0812*	0.0145	0.0112
	(0.011)	(0.018)	(0.011)	(0.012)	(0.010)
France	0.0579*	0.8856*	0.0629*	0.0544*	−0.0272*
	(0.014)	(0.016)	(0.012)	(0.018)	(0.009)
Germany	0.0307*	0.8989*	0.0668*	0.0253*	−0.0075*
	(0.004)	(0.007)	(0.008)	(0.009)	(0.003)
Hong Kong	0.0870*	0.8231*	0.1189*	0.1142*	−0.0206*
	(0.007)	(0.005)	(0.009)	(0.010)	(0.004)
Hungary	0.3701*	0.5299*	0.3692*	−0.1810*	0.0427
	(0.032)	(0.028)	(0.024)	(0.029)	(0.023)
Israel	0.1748*	0.7909*	0.1315*	0.0617*	−0.0432*
	(0.034)	(0.023)	(0.021)	(0.025)	(0.015)
Italy	0.0529*	0.9020*	0.0749*	−0.0130	−0.0094*
	(0.007)	(0.008)	(0.007)	(0.007)	(0.004)
Japan	0.0309*	0.8725*	0.0664*	0.0450*	0.0127*
	(0.004)	(0.010)	(0.005)	(0.009)	(0.004)
Korea	0.1154*	0.8351*	0.0799*	0.0900*	0.0149
	(0.025)	(0.021)	(0.014)	(0.022)	(0.012)
Malaysia	0.0817*	0.8225*	0.1191*	0.0437*	−0.0211*
	(0.009)	(0.009)	(0.007)	(0.010)	(0.009)
Netherlands	0.1216*	0.8071*	0.0771*	0.0931*	−0.0836*
	(0.013)	(0.015)	(0.012)	(0.013)	(0.010)
Norway	0.0524*	0.8674*	0.0975*	0.0120	−0.0371*
	(0.005)	(0.005)	(0.006)	(0.011)	(0.005)
Portugal	0.3137*	0.2647*	0.2987*	0.0129	−0.0115
	(0.050)	(0.084)	(0.031)	(0.048)	(0.048)
South Africa	0.2772*	0.7237*	0.3579*	−0.2170*	−0.1817*
	(0.028)	(0.019)	(0.028)	(0.035)	(0.013)
Spain	0.0702*	0.8157*	0.1448*	−0.0366*	−0.0012
	(0.010)	(0.016)	(0.013)	(0.015)	(0.007)
Sweden	0.0495*	0.8988*	0.0499*	0.0492*	−0.0113
	(0.009)	(0.011)	(0.008)	(0.011)	(0.008)
Switzerland	0.2491*	0.6765*	0.1098*	0.0390	−0.1255*
	(0.058)	(0.060)	(0.021)	(0.023)	(0.029)

TABLE V (Continued)

Effect of Futures Introduction on GJR-GARCH Volatility

Country	α_0	α_1	α_2	α_3	α_M
United Kingdom	0.0110*	0.9370*	0.0474*	0.0130	−0.0096*
	(0.002)	(0.005)	(0.005)	(0.007)	(0.003)
United States	0.0046*	0.9204*	0.0446*	0.0309*	0.0133*
	(0.001)	(0.005)	(0.005)	(0.005)	(0.002)

Constrained Maximum Likelihood parameter estimates are reported for the GJR-GARCH model with a multiplicative dummy:

$$h_t = (1 + \alpha_M D_t)[\alpha_0 + \alpha_1 h_{t-1} + \alpha_2 \varepsilon_{t-1}^2 + \alpha_3 \max(0, -\varepsilon_{t-1})^2],$$

where ε_t is the residual from autoregression 2, and α_M is a dummy variable equal to zero before and one after the futures introduction. Standard errors are shown in parentheses. An asterisk indicates statistical significance at the 5% level.

country had its own β. The results of this specification are similar to those reported here.

In order to investigate whether the mismatching between time zones materially affects our results, we reestimated the model for the Asian countries in our sample using only lagged world-market returns, and this had no impact on the signs or significance of any of our tests. Also, we estimated our model using weekly instead of daily data. Here, despite the difficulty in estimating good GARCH parameters on weekly data, we still found results in line with those reported above.

Some authors, such as Chan and Karolyi (1991) and Lee and Ohk (1992), have tested for more-general structural changes in the GARCH equation at the time of futures listing by interacting a dummy variable separately for each term in the conditional volatility equation. By examining these coefficients, one can measure whether there is a change in the speed with which volatility shocks dissipate. We also estimated such a model for each country in our sample. Although some of the coefficients on the individual dummy variables were statistically significant, no clear pattern emerged across countries.

THE EFFECT OF FUTURES TRADING ACTIVITY

In the previous section, we tested whether there appears to be any structural change in the underlying market at the time of futures introduction. In this section, we will test whether there appears to be a relationship, *after the futures have been listed,* between the level of future trading activity and the volatility of the underlying index.

Our approach is based on that of Bessembinder and Seguin (1992). Using an autoregressive integrated moving average (ARIMA) model, they decomposed the time series of futures trading volume and open interest into expected and unexpected components. Bursts of trading activity stimulated by unexpected price changes should be picked up in the unexpected component, while the expected component should reflect the "background" level of futures trading. They found that market volatility was positively related to the unexpected components of volume and open interest, reflecting the positive effect of volatility on volume, but that market volatility was negatively related to the expected component, suggesting an underlying stabilizing influence.

We followed a similar procedure using futures market trading volume and open-interest data from 17 of our 25 countries for which volume and open interest data were available. First, we analyzed the volume and open-interest time series from each country to select an ARIMA model that appeared to fit the data reasonably well. Restricting our attention to models with five or less autoregressive lags and five or less moving average lags, we selected, on the basis of the autocorrelation structure, a different model for each time series. We then used these models to decompose each time series into expected and unexpected components, and then inserted them as additional explanatory variables in the GJR-GARCH conditional volatility equation:

$$h_t = \alpha_0 + \alpha_1 h_{t-1} + \alpha_2 \varepsilon_{t-1}^2 + \alpha_3 \max(0, -\varepsilon_{t-1})^2 + \alpha_4\, ExpVol + \alpha_5\, UnexpVol + \alpha_6\, ExpOI + \alpha_7\, UnexpOI,$$

where *ExpVol* and *UnexpVol* are the expected and unexpected components of volume, and *ExpOI* and *UnexpOI* are the expected and unexpected components of open interest.

Estimation results are reported in Table VI. Of the 17 countries analyzed, the coefficient α_7 on the unexpected component of open interest is negative in all 17, and statistically significant in eight. The coefficient α_6 on the expected component of open interest is positive and significant only in Japan, and it is positive but not significant in the United States. The coefficient is negative in the remaining 15 countries, significantly so in seven. Note that these results very closely correspond to those reported in our earlier analysis. We interpreted this as additional evidence that it is, in fact, futures trading, not spuriously correlated factors, that drives the results.

With respect to the unexpected component of futures trading volume, we found, like Bessembinder and Seguin (1992), that it has a posi-

TABLE VI

Effect of Futures Trading Activity on Volatility

Country *Vol|OI*	α_0	α_1	α_2	α_3	α_4	α_5	α_6	α_7
Australia	0.1342	0.7087*	0.1745*	−0.0941*	0.0346	0.4671*	−0.0508	−0.06379
$10^4\|10^5$	(0.166)	(0.026)	(0.009)	(0.016)	(0.046)	(0.059)	(0.042)	(0.163)
Austria	0.2032*	0.8200*	0.0963*	−0.0100	−0.6353*	2.0146*	−0.0812	−0.7615*
$10^4\|10^5$	(0.027)	(0.033)	(0.022)	(0.026)	(0.182)	(0.272)	(0.069)	(0.204)
Belgium	0.0547*	0.8425*	0.0575*	0.0064	−0.0077	0.3832	−0.0133	−0.4163
$10^4\|10^4$	(0.022)	(0.048)	(0.019)	(0.028)	(0.124)	(0.266)	(0.014)	(0.217)
Canada	0.0263*	0.8455*	0.1432*	−0.0767*	−0.0093	0.0491*	−0.0009	−0.2244*
$10^3\|10^4$	(0.004)	(0.015)	(0.015)	(0.014)	(0.008)	(0.015)	(0.005)	(0.057)
Denmark	0.2727*	0.3764*	0.2096*	−0.0904*	1.9812*	3.8229*	−0.2113*	−0.3571
$10^4\|10^4$	(0.047)	(0.064)	(0.036)	(0.044)	(0.525)	(0.327)	(0.076)	(0.829)
France	0.2950*	0.5485*	0.0781*	0.0479	0.7452	3.0331*	−0.3960	−2.6735
$10^5\|10^6$	(0.053)	(0.070)	(0.025)	(0.035)	(0.391)	(0.334)	(0.289)	(2.838)
Germany	0.3404*	0.4334*	0.0330	0.0899	1.0981*	4.2413*	−0.1279*	−0.2444
$10^5\|10^5$	(0.050)	(0.072)	(0.029)	(0.046)	(0.289)	(0.364)	(0.040)	(0.390)
Hong Kong	0.4442*	0.4131*	0.1507*	0.1310*	2.6692*	11.2339*	−0.5390	−5.5596*
$10^5\|10^5$	(0.055)	(0.056)	(0.033)	(0.049)	(1.202)	(0.817)	(0.449)	(1.484)
Italy	0.5237*	0.3059*	0.2601*	−0.1106	1.8269	8.5268*	−0.6310	−5.899*
$10^5\|10^5$	(0.151)	(0.130)	(0.081)	(0.099)	(1.542)	(1.126)	(0.9690)	(2.053)
Japan	0.0121*	0.9111*	0.0217*	0.0979*	−0.0030	0.4131*	0.0757*	−0.4698
$10^4\|10^6$	(0.003)	(0.010)	(0.007)	(0.013)	(0.010)	(0.098)	(0.025)	(1.201)
Netherlands	0.0488*	0.9156*	0.0474*	0.0022	−0.0122	0.2918*	−0.0972*	−0.7806*
$10^4\|10^5$	(0.010)	(0.016)	(0.013)	(0.013)	(0.026)	(0.077)	(0.044)	(0.388)
Norway	0.0403*	0.8644*	0.0965*	−0.0026	0.0650	0.1827*	−0.1504*	−0.5672
$10^3\|10^4$	(0.010)	(0.019)	(0.015)	(0.025)	(0.062)	(0.085)	(0.068)	(0.515)
Spain	0.2958*	0.6943*	0.0484	0.0993*	0.3125	4.8436*	−0.3706*	−0.9556
$10^5\|10^5$	(0.051)	(0.071)	(0.025)	(0.037)	(0.506)	(0.502)	(0.160)	(0.764)
Sweden	0.0009	0.8883*	0.0032	0.0535	0.1223	0.2158	−0.0172	−0.1134
$10^4\|10^5$	(0.072)	(0.059)	(0.027)	(0.038)	(0.140)	(0.167)	(0.127)	(0.366)
Switzerland	0.2207*	0.7317*	0.0357	0.0523	−0.0024	0.8582*	−0.4029*	−2.3471*
$10^4\|10^5$	(0.036)	(0.082)	(0.021)	(0.030)	(0.097)	(0.085)	(0.175)	(0.132)
United Kingdom	0.0393*	0.9036*	0.0450*	0.0164	−0.0103*	0.1071*	−0.0081	−0.3577*
$10^4\|10^5$	(0.007)	(0.016)	(0.010)	(0.011)	(0.005)	(0.018)	(0.005)	(0.115)
United States	0.0052*	0.9193*	0.0554*	0.0231*	0.0007	0.2241*	0.0083	−0.7909*
$10^5\|10^6$	(0.002)	(0.005)	(0.007)	(0.007)	(0.003)	(0.017)	(0.005)	(0.168)

Estimated coefficients from a GJR-GARCH model with expected and unexpected components of futures trading activity variables:

$$h_t = \alpha_0 + \alpha_1 h_{t-1} + \alpha_2 \varepsilon_{t-1}^2 + \alpha_3 \max(0, -\varepsilon_{t-1})^2 + \alpha_4\, ExpVol + \alpha_5\, UnexpVol + \alpha_6\, ExpOI + \alpha_7\, UnexpOI$$

For computational reasons, the volume and open interest series are standardized to have a mean between zero and one. Scaling units are reported below the country name. An asterisk indicates statistical significance at the 5% level.

tive effect on volatility. This is what we would expect to see if exogenous volatility events cause high trading volume. The expected component of futures volume, on the other hand, had no robust significant effect on volatility—there is a significant positive effect in Denmark, Germany, and Hong Kong, a significant negative effect in Austria and the UK, and no significant effect in the other 12 countries.

It should be noted that we made no effort to model the seasonality in volume and open-interest data that was associated with the rolling over of contracts prior to monthly expiration dates. Thus, in our ARIMA decomposition, we may be erroneously classifying some predictable trading volume as unpredictable. An interesting topic for future research would be to investigate whether accounting for this seasonality has a significant impact on the decomposition of permanent and temporary components of trading activity.

MODELING THE JOINT DYNAMICS OF COUNTRY AND WORLD VOLATILITY

The univariate models we employed above do not allow for time-varying conditional covariance between the country and world returns. If the conditional covariance changes systematically with the introduction of stock index futures, then our previous results may be biased.

In this section, we will address this problem by estimating the joint dynamics of each country's return with the world-market return in a multivariate GARCH framework that allows for time-varying conditional covariance. Because we wished to capture the dynamic interaction between world-market volatility, country-specific volatility, and conditional covariance, we used the BEKK specification of Engle and Kroner (1995).[9] Unlike certain other well-known multivariate GARCH models, the BEKK model allows conditional variances and covariances to influence each other.[10]

For each country i, we estimated the following bivariate process:

$$R_{i,t} = a_0 + \sum_{j=1}^{5} a_j R_{i,t-j} + \sum_{k=2}^{5} b_k DAY_k + \varepsilon_{\mathbf{i,t}}$$

$$R_{w,t} = w_0 + \sum_{j=1}^{5} w_j R_{w,t-j} + \sum_{k=2}^{5} d_k DAY_k + \varepsilon_{\mathbf{w,t}}$$

where the error terms are multivariate normal:

[9]This model also has been used by Karolyi (1995) to model the joint dynamics of stock returns in Canada and the United States.

[10]For a comparison of BEKK and other multivariate GARCH models, see Kroner and Ng (1998).

$$\varepsilon_t | F_{t-1} \sim N(\mathbf{0},\ \mathbf{H}_{\mathrm{t}})$$

with conditional covariance matrix

$$H_t = \mathbf{C}'\mathbf{C} + \mathbf{A}'\varepsilon_{t-1}\ \varepsilon'_{t-1}\ \mathbf{A} + \mathbf{G}'\mathbf{H}_{\mathrm{t-1}}\mathbf{G} + \Phi D_t.$$

In the mean equations, $R_{i,t}$ represents the log country-index return, $R_{w,t}$ is the contemporaneous log world-index return, and the variables DAY_k are day-of-the-week dummies for Tuesday through Friday. In the conditional variance equations, the coefficient matrix **C** represents a matrix of constants, **A** represents a matrix of ARCH coefficients, **G** represents a matrix of GARCH coefficients, and Φ represents a matrix of dummy coefficients. Matrices **H**, **C**, **G**, and Φ are symmetric. Our main purpose in using the multivariate GARCH model was to better correct for the effect of world-market movements, not to test whether futures listing in individual countries influenced world-market volatility. Thus, we did not include a dummy variable for futures listing in the conditional-variance equation for world returns. That is, d_{22} of Φ is set to 0.

In summary,

$$\mathbf{H}_t = \mathbf{C}'\mathbf{C} + \begin{bmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \end{bmatrix}' \begin{bmatrix} \varepsilon^2_{1,t-1} & \varepsilon_{1,t-1}\varepsilon_{2,t-1} \\ \varepsilon_{2,t-1}\varepsilon_{1,t-1} & \varepsilon^2_{2,t-1} \end{bmatrix} \begin{bmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \end{bmatrix}$$
$$+ \begin{bmatrix} g_{11} & g_{12} \\ g_{12} & g_{22} \end{bmatrix}' \mathbf{H}_{\mathrm{t-1}} \begin{bmatrix} g_{11} & g_{12} \\ g_{12} & g_{22} \end{bmatrix} + \begin{bmatrix} d_{11} & d_{12} \\ d_{12} & 0 \end{bmatrix} D_t$$

The interpretation of the dummy coefficient in a country's conditional-variance equation is analogous to the dummy in the univariate GARCH equations analyzed above—a negative coefficient indicates that the introduction of futures corresponds to a volatility decrease. By including the dummy variable in the equation governing the conditional covariance of a country's return with the world-market return, we attempted to test whether futures introduction has any impact on the extent to which the country's stock market is integrated into the world market.

Maximum-likelihood estimates of these parameters are reported in Table VII. In the interest of space, we have reported only those coefficients that we feel are most relevant to the issue at hand. Examining the coefficients on the futures introduction dummy variable in the country-specific-conditional-volatility equation, we found a significant volatility increase in four countries (Germany, Japan, Hungary, and Spain) and a significant decrease in 12 countries. Note that under this specification, the volatility effect for the United States is still positive but is no longer

TABLE VII

Joint Dynamics of Country and World Returns

Country	c_{11}	c_{12}	a_{11}	a_{12}	a_{21}	g_{11}	g_{12}	d_{11}	d_{12}
Australia	0.3173*	0.0339	0.0101*	−0.0132	0.7291*	0.2779*	0.1461*	−0.0937*	0.0009
	(0.0284)	(0.0183)	(0.0083)	(0.0077)	(0.0151)	(0.0251)	(0.0543)	(0.0194)	(0.0179)
Austria	0.0041*	0.0014	0.2367*	0.0113*	0.0142	0.9446*	0.9140*	0.0017	0.0010
	(0.0013)	(0.0017)	(0.0060)	(0.0045)	(0.0156)	(0.0026)	(0.0061)	(0.0015)	(0.0014)
Belgium	0.0574*	−0.0014	0.2852*	0.0253*	0.2637*	0.7452*	0.6992*	0.0025	0.0132*
	(0.0079)	(0.0038)	(0.0204)	(0.0101)	(0.0110)	(0.0221)	(0.0192)	(0.0058)	(0.0046)
Canada	0.0235*	0.0162	0.3470*	0.0192*	0.0054	0.8420*	0.8412*	−0.0040*	−0.0052*
	(0.0021)	(0.0015)	(0.0067)	(0.0054)	(0.0063)	(0.0056)	(0.0059)	(0.0014)	(0.0010)
Chile	0.1718*	−0.0084	0.4538*	0.0270*	0.0179	0.3724*	0.0449	−0.0719*	0.0688*
	(0.0205)	(0.0268)	(0.0127)	(0.0052)	(0.0299)	(0.0123)	(0.0472)	(0.0160)	(0.0290)
Denmark	0.45768	−0.0112	0.3240*	0.0055	0.2547*	0.4579*	0.2963*	−0.2076*	0.0618*
	(0.0334)	(0.0086)	(0.0157)	(0.0073)	(0.0177)	(0.0379)	(0.0543)	(0.0153)	(0.0115)
Finland	0.0305*	0.0069	0.4030*	0.0007	0.0890*	0.7907*	0.6112*	0.0283*	0.0089
	(0.0051)	(0.0081)	(0.0140)	(0.0070)	(0.0110)	(0.0104)	(0.0306)	(0.0046)	(0.0084)
France	0.0930*	−0.0009	0.2979*	0.0315*	0.0044	0.8623*	0.7881*	−0.0247	0.0246*
	(0.0233)	(0.0081)	(0.0173)	(0.0083)	(0.0203)	(0.0145)	(0.0146)	(0.0177)	(0.0080)
Germany	0.0358*	0.0143*	0.3297*	0.0408*	0.0177	0.8585*	0.6734*	0.0198*	0.0298*
	(0.0036)	(0.0036)	(0.0102)	(0.0059)	(0.0114)	(0.0083)	(0.0217)	(0.0030)	(0.0043)
Hong Kong	0.0706*	0.0021	0.4153*	0.0104*	0.1460*	0.8149*	0.7488*	−0.0046	0.0201*
	(0.0072)	(0.0031)	(0.0082)	(0.0019)	(0.0119)	(0.0050)	(0.0190)	(0.0071)	(0.0040)
Hungary	0.1650*	−0.0099*	0.5453*	0.0109*	0.3808*	0.5829*	0.8104*	0.0553*	0.0181*
	(0.0161)	(0.0038)	(0.0193)	(0.0029)	(0.0221)	(0.0183)	(0.0257)	(0.0234)	(0.0059)
Israel	0.1294*	0.0799*	0.3963*	−0.0213*	−0.1335*	0.8059*	0.0130	−0.0207	0.0387
	(0.0312)	(0.0315)	(0.0231)	(0.0049)	(0.0593)	(0.0222)	(0.3260)	(0.0228)	(0.0362)
Italy	0.0486*	0.0058*	0.2558*	−0.0024	0.0127	0.9021*	0.8668*	0.0023	0.0058*
	(0.0043)	(0.0015)	(0.0080)	(0.0030)	(0.0100)	(0.0060)	(0.0116)	(0.0055)	(0.0029)
Japan	0.0181*	−0.0047*	0.3014*	0.0052	0.1747*	0.8525*	0.8792*	0.0317*	0.0145*
	(0.0022)	(0.0011)	(0.0096)	(0.0029)	(0.0069)	(0.0063)	(0.0047)	(0.0035)	(0.0012)
Korea	0.1169*	0.0567*	0.3488*	−0.0603*	0.1249*	0.8262*	−0.0284	0.0482	−0.0062
	(0.0193)	(0.0200)	(0.0186)	(0.0050)	(0.0416)	(0.0163)	(0.1988)	(0.0307)	(0.0437)
Malaysia	0.1102*	0.0112*	0.3904*	0.0191*	0.0989*	0.7853*	0.6658*	−0.0388*	0.0135
	(0.0071)	(0.0045)	(0.0108)	(0.0041)	(0.0122)	(0.0104)	(0.0284)	(0.0075)	(0.0075)
Netherlands	0.0944*	0.0018	0.2619*	0.0262*	0.0977*	0.8674*	0.8139*	−0.0538*	0.0066*
	(0.0081)	(0.0035)	(0.0094)	(0.0050)	(0.0139)	(0.0077)	(0.0097)	(0.0053)	(0.0029)
Norway	0.1132*	0.0473*	0.4149*	0.0112	0.1566*	0.7381*	0.3568*	−0.0486*	0.0125
	(0.0077)	(0.0027)	(0.0115)	(0.0074)	(0.0095)	(0.0115)	(0.0366)	(0.0060)	(0.0084)
Portugal	0.0247*	−0.0040	0.4738*	0.0531	0.1618*	0.7403*	0.3218*	−0.0033	0.0425*
	(0.0043)	(0.0061)	(0.0190)	(0.0162)	(0.0233)	(0.0161)	(0.1013)	(0.0059)	(0.0135)
South Africa	0.9573*	0.1467*	0.3228*	−0.0585*	0.2147*	0.6498*	0.8169*	−0.8027*	−0.1239*
	(0.0946)	(0.0212)	(0.0204)	(0.0019)	(0.0212)	(0.0306)	(0.0262)	(0.0786)	(0.0176)
Spain	0.0407*	0.0073*	0.3049*	0.0164*	−0.0292*	0.8744*	0.8396*	0.0089*	0.0042
	(0.0045)	(0.0024)	(0.0129)	(0.0065)	(0.0060)	(0.0109)	(0.0111)	(0.0037)	(0.0028)
Sweden	0.0953*	−0.0174	0.2828*	0.0180*	0.0874*	0.8629*	0.7726*	−0.0300*	0.0314*
	(0.0177)	(0.0095)	(0.0137)	(0.0068)	(0.0176)	(0.0113)	(0.0127)	(0.0145)	(0.0098)
Switzerland	0.2279*	−0.0100	0.2863*	0.0200	0.3394*	0.6520*	0.6596*	−0.0667*	0.0239*
	(0.0240)	(0.0090)	(0.0229)	(0.0105)	(0.0176)	(0.0288)	(0.0172)	(0.0117)	(0.0093)

TABLE VII (Continued)

Joint Dynamics of Country and World Returns

Country	c_{11}	c_{12}	a_{11}	a_{12}	a_{21}	g_{11}	g_{12}	d_{11}	d_{12}
United Kingdom	0.0279*	0.0056*	0.2607*	0.0073	0.0678*	0.8986*	0.8630*	−0.0129*	0.0010
	(0.0040)	(0.0019)	(0.0093)	(0.0050)	(0.0071)	(0.0057)	(0.0075)	(0.0030)	(0.0016)
United States	0.0106*	0.0061*	0.2203*	−0.0194*	0.0074	0.9380*	0.9419*	0.0005	−0.0006
	(0.0013)	(0.0007)	(0.0035)	(0.0026)	(0.0062)	(0.0025)	(0.0021)	(0.0009)	(0.0004)

Constrained Maximum Likelihood parameter estimates and standard errors are reported for the BEKK model with dummy variables for futures listing in each country's conditional variance equation and conditional covariance equation:

$$\begin{bmatrix} h_{11,t} & h_{12,t} \\ h_{12,t} & h_{22,t} \end{bmatrix} = \mathbf{C}'\mathbf{C} + \begin{bmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \end{bmatrix}' \begin{bmatrix} \varepsilon_{1,t-1}^2 & \varepsilon_{1,t-1}\varepsilon_{2,t-1} \\ \varepsilon_{2,t-1}\varepsilon_{1,t-1} & \varepsilon_{2,t-1}^2 \end{bmatrix} \begin{bmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \end{bmatrix}$$

$$+ \begin{bmatrix} g_{11} & g_{12} \\ g_{12} & g_{22} \end{bmatrix}' \begin{bmatrix} h_{11,t-1} & h_{12,t-1} \\ h_{12,t-1} & h_{22,t-1} \end{bmatrix} \begin{bmatrix} g_{11} & g_{12} \\ g_{12} & g_{22} \end{bmatrix} + \begin{bmatrix} d_{11} & d_{12} \\ d_{12} & 0 \end{bmatrix} D_t$$

For brevity, coefficients specific to the world conditional variance equation are not reported. An asterisk indicates statistical significance at the 5% level.

statistically significant. Although these results are not as consistent as those from the GJR-GARCH model, we still observed a propensity for volatility to decrease after futures introduction.

Examining the dummy coefficients in the conditional covariance equations, we found that conditional covariance with the world market increases in 21 out of 25 countries, with statistical significance in 13 cases. We may interpret this as evidence that futures markets contribute to an increase in the level of world-market integration. Again, we should interpret these results with caution as, over time, we would expect countries to become more integrated with the world, with or without futures markets.

CONCLUSION

In this article, we have examined the time-series properties of returns in 25 markets around the world before and after the introduction of stock index futures. In 17 of these countries, we also have examined the impact of trading activity in the post-listing samples.

First, in each country, we examined the time series of excess returns over the world-market index using various GARCH models to account for asynchronous trading, conditional heteroskedasticity in returns, and an asymmetric response to positive and negative news. Our results indicated that in the largest two markets, the United States and Japan, volatility may have increased after the listing of stock index futures. On the other hand, volatility decreased or stayed roughly the same in most of the other countries in our sample, with statistically significant decreases in many cases. This result appears to be robust to model specification, holding for different specifications of the dummy variable and for GARCH specifications, including models that allow for asymmetric responses to good and bad news.

Next, using a procedure inspired by Bessembinder and Seguin (1992), we found that in most countries, volatility tends to be lower in periods when open interest in stock index futures is high. The only two cases where we found the opposite results are the United States and Japan, reinforcing our previous results. In some cases, volatility is higher in periods when futures volume is high, but this is driven by the unexpected component of volume, not the expected component.

Finally, we extended our analysis to a multivariate framework that allows for the possibility of volatility spillover and time-varying conditional covariance between country-specific and world returns. In this framework, we found that the volatility increase in the United States at the time of futures listing is no longer significant, and we still found that country-specific conditional variance is more likely to decline than to increase with the introduction of stock index futures. We also documented that the markets in most countries are significantly more integrated with the world market after the introduction of stock index futures.

We do not deny that these results may be influenced by other factors, and, as always, advocate caution in interpreting empirical results. In particular, several points should be considered that may confound the interpretation of our results, and those of all the previous papers in this literature.

First, the listing of index futures is not an entirely exogenous event. The listing process involves many decisions made by exchange officials and regulators, who may have been influenced by recent or anticipated market conditions. For example, the reluctance of regulators to approve the introduction of index futures during periods of political uncertainty may introduce a selection bias.

Second, because the events in our sample are not independent draws from an homogeneous population, we cannot interpret this as we would a traditional event study.[11] Different countries have different contract designs, trading mechanisms, and regulatory environments. There have been important political and economic developments in various countries that are not captured by our model. For example, although we found a

[11]We thank Andrew Karolyi for useful comments on this issue.

reliably positive coefficient for futures listing in Japan, we are not convinced that futures listing caused this increase since we can point to other developments in the Japanese market that have led to high volatility there. Undoubtedly, one could tell a similar story for other countries in our sample. Some countries in our sample have listed index options in addition to index futures, and others have not. Some countries have competing offshore contracts, and others do not. Moreover, the events in our sample are clustered in time, with a group of English-speaking developed countries listing in the early 1980s, a group of Western European and other developed markets listing in the late 1980s, and emerging markets listing in the 1990s.

Third, it should be noted that a relatively long time series is required to obtain reliable GARCH parameter estimates. In some cases, our window length may be too short. For the most recent listings in our sample, the post-listing window may be too short to make reliable inferences. This may explain the unusual parameter estimates reported for Hungary and Portugal. In other countries, such as Finland, France, and Sweden, we have relatively short pre-event windows.

Despite these inherent difficulties, the results we have reported here do present a relatively consistent picture, which appears to be robust to model specification: in many markets, stock index futures appear to have contributed to a decrease in conditional variance. In the United States and Japan, however, the opposite appears to have been the case.

We do not know why the results appear to be different for the two largest markets in our sample. Many theoretical models have been proposed to explain why futures markets might impact the underlying market, and in most cases, these models have ambiguous predictions depending on the parameters. Perhaps the most highly developed markets correspond to parameter values in these models that would lead us to predict increased volatility.

For example, Stein (1987) develops an information-based model in which prices are determined by the interaction between hedgers and informed speculators. In this model, opening a futures market has two opposing effects. On one hand, the futures market improves risk sharing and reduces price volatility. But also, if the speculators observe a noisy but informative signal, the hedgers react to the noise in the speculative trades, producing an offsetting (volatility-increasing) effect. Perhaps in countries with highly developed financial institutions, futures introduction offers only modest risk-sharing benefits while introducing a large number of speculators with noisy information. In other countries, the introduction of futures may have greater risk-sharing benefits that outweigh the "informational externality" associated with partially informed speculators. This is only one of many examples of theoretical models that can predict a volatility increase/decrease with futures introduction, depending on the parameters.

In summary, after examining the evidence, we found little reason to expect that the introduction of new stock index futures contracts in emerging nations will destabilize stock markets. On the contrary, in nearly every country outside of the US and Japan, it appears that futures markets have had either no significant effect or else have been associated with a significant decline in volatility and increased integration with the world market.

BIBLIOGRAPHY

Antoniou, A., Holmes, P., & Priestley, R. (1998). The effects of stock index futures trading on stock index volatility: An analysis of the asymmetric response of volatility to news. Journal of Futures Markets, 18, 151–166.

Bessembinder, H., & Seguin, P. J. (1992). Futures-trading activity and stock price volatility. Journal of Finance, 47, 2015–2034.

Bollerslev, T. (1986). Generalized autoregressive conditional heteroscedasticity. Journal of Econometrics, 31, 307–327.

Chan, K. C., & Karolyi, G. A. (1991). The volatility of the Japanese stock market: Evidence from 1977 to 1990. In W. T. Ziemba, W. Bailey, & Y. Hamao (eds.), Japanese financial market research. New York: Elsevier, 121–143.

Damodaran, A., & Subrahmanyam, M. G. (1992). The effects of derivative securities on the markets for the underlying assets in the United States: A survey. Financial Markets, Institutions & Instruments, 1, 1–22.

Engle, R. F., & Kroner, K. F. (1995). Multivariate simultaneous generalized ARCH. Econometric Theory, 11, 122–150.

Engle, R., & Mezrich, J. (1995). Grappling with GARCH. Risk, 8, 112–117.

Engle, R., & Ng, V. (1993). Measuring and testing the impact of news on volatility. Journal of Finance, 48, 1749–1778.

Glosten, L. R., Jagannathan, R., & Rundle, D. E. (1993). On the relation between the expected value and the volatility of the nominal excess return on stocks. Journal of Finance, 48, 1779–1801.

Gulen, H., & Mayhew, S. (1999). The dynamics of international stock index returns. Working paper, University of Georgia.

Hodges, S. (1992). Do derivative instruments increase market volatility? In S. Hodges (ed.), Options: recent advances in theory and practice (Vol. II, Chapter 12). Manchester, UK: Manchester University Press.

Karolyi, G. A. (1995). A multivariate GARCH model of international transmissions of stock returns and volatility: The case of the United States and Canada. Journal of Business & Economic Statistics, 13, 11–25.

Kroner, K. F., & Ng, V. K. (1998). Modeling asymmetric comovements of asset returns. Review of Financial Studies, 11, 817–844.

Lee, S. B., & Ohk, K. Y. (1992). Stock index futures listing and structural change in time-varying volatility. Journal of Futures Markets, 12, 493–509.

Lo, A., & MacKinlay, C. (1988). Stock market prices do not follow random walks: Evidence from simple specification test. Review of Financial Studies, 1, 41–66.

Mayhew, S. (1999). The impact of derivatives on cash markets: What have we learned? Working paper, University of Georgia.

Nelson, D. (1991). Conditional heteroscedasticity in asset returns: A new approach. Econometrica, 59, 347–370.

Pagan, A., & Schwert, G. W. (1990). Alternative models for conditional stock volatility. Journal of Econometrics, 45, 267–290.

Rhode, W. (1997, October). Risk in an era of reform. AsiaRisk (Supplement to Risk), 19–23.

Scholes, M., & Williams, J. (1977). Estimating betas from nonsynchronous data. Journal of Financial Economics, 5, 309–327.

Stein, J. C. (1987). Informational externalities and welfare-reducing speculation. Journal of Political Economy, 95, 1123–1145.

Stewart, J., & Priest, A. (1997, October). Exchange of fortunes. LatinRisk (Latin-America Supplement to Risk), 21–23.

Sutcliffe, C. (1997). Stock index futures: Theories and international evidence (2nd ed.). London: International Thomson Business Press.