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Stock Index Futures Trading and Volatility in International Equity Markets

Abstract

We examine 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 find that futures trading is related to an increase in conditional volatility in the United States and Japan, but in nearly every other country, we find either no significant effect, or 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 index futures. We also document an increase in conditional covariance between country-specific and world returns at the time of futures listing.

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 Argentina, Colombia, Costa Rica, Mexico and Peru, India, Indonesia, Czech Republic, Slovakia, Greece, and Turkey.

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, particularly in emerging economies. Previous literature on the effects of stock index futures trading has focused primarily on developed markets. Moreover, the existing research has come to conflicting conclusions regarding the effect of futures trading on volatility. Some authors have found that volatility appears to increase with the introduction of futures, and some find no significant effect, and some find that volatility decreases.[1]

This paper examines 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. Not only do we test for structural changes at the time of futures listing by comparing properties of the returns series before and after listing, but we also test whether volatility in the post-listing period is related to futures market volume and open interest. The results of both tests show that futures trading is associated with increased volatility in the United States and Japan, but in virtually every one of the other twenty-three countries, this is not the case.

Various theories have been advanced elsewhere for how the introduction of futures might impact

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

the volatility of the underlying market. As pointed out by Hodges (1992), Mayhew (1999) and others, many of these theories predict 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. One interpretation of our result is that futures influence the underlying market through multiple, offsetting channels, with the relative importance of the effects depending on the extent of the development of the market. In particular, it appears that futures markets may play an important role in stabilizing less-developed markets.

Our paper contributes to the existing literature in several ways. To the best of our knowledge, it examines a much broader cross section of international futures introductions than any prior study, and is the first to examine the impact of futures markets in a cross-section of emerging nations. Included in our sample are the markets examined by previous authors, such as Lee and Ohk (1992), and Antoniou, Holmes and Priestly (1998), but with considerably larger sample periods.[2] We examine the properties of excess returns over the world market index. This enables us to avoid attributing worldwide price movements, such as the crash of October 1987, to the listing of futures in the local market. In addition, this is the first paper, to our knowledge, that examines the relationship of volatility to futures market volume and open interest for a large cross-section of markets.

The basic approach of our anaysis is to test the the impact of futures introduction on volatility using a modification of the Generalized Auto-Regressive Conditional Heteroskedasticity model suggested by Glosten, Jagannathan and Rundle (1993) (GJR-GARCH). To test for the impact of futures trading we incorporate a multiplicative dummy variable in the conditional variance equation. We check the robustness of our results using various alternative specifications. Next, using a technique similar to then employed by Bessembinder and Seguin (1992), we decompose the trading volume and open interest time series into permanent and temporary components, and test how these components affect volatility by inserting them into the conditional volatility equation. Fi-

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

nally, we analyze the joint dynamics of each country with the world market portfolio using the bivariate GARCH specification advanced by Engle and Kroner (1995), commonly known as the BEKK model.[3] This richer framework allows us to more carefully control for movements in global markets. It also allows 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. Section I describes the data we used in our analysis. It also reports the results of simple variance ratio tests of whether the variance of raw returns is higher or lower after the introduction of futures. The results of these simple tests suggest that volatility declines more often than it increases in response to futures listing. In section II, we describe the univariate GJR-GARCH framework we use to analyze the data, we verify that conditional heteroskedasticity is present in the returns in all twenty-five countries, and we test whether volatility is higher or lower after the introduction of futures trading. We also discuss the robustness of the results to model specification. In section III we examine whether conditional volatility is related to the temporary and permanent components of open interest and trading volume. In section IV we present our analysis of the joint dynamics of country-specific and world returns using the BEKK bivariate model.

I Data

Daily stock market index data were obtained from Datastream[4] for twenty-five of the thirty nations listed in Table 1. Russia, Venezuela, and Poland, which listed futures after July 1996, were excluded because in our judgment, there was insufficient data in the post-event 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 use data on the more popular S&P 500 index instead of the Value Line index. In some cases, very little data exists

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

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 do not believe this to be a major problem. To illustrate, for Norway, we use data on the Oslo Stock Exchange (OSE) General Stock Index instead of the OBX index 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 use the Helsinki Stock Exchange General Index (HEX) instead of the FOX index, and in the U.K. and Italy, we use 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 use all the stock index data available on Datastream between 1973 and 1997. In most cases, data are 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 2.

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.

Table 3 reports the results of simple variance ratio tests, testing whether stock index volatility changes with the introduction of stock index futures trading.[6] Results based on daily data are reported for various fixed event windows ranging from six months to five years. The final column reports results using all data available through Datastream.

[5]In the case of the U.K., 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 use daily data.

[6]We conducted this test by calculating the variance of continuous daily returns using data before and after the date of futures introduction. The variance ratio has an F distribution with degrees of freedom determined by the number of observations in the pre- and post-event samples.

In this table, up and down arrows indicate volatility increases and decreases that are statistically significant at the 5% level, while plus and minus signs indicate changes that are not statistically significant. The results reported for Australia, Japan, Hong Kong and the U.S. correspond with the findings reported by Lee and Ohk (1992): no significant effect in Australia, results for Hong Kong that are sensitive to the choice of window size, and increased volatility in Japan and the United States. Despite the results for Japan and the United States, the overall impression given by Table 3 is that the introduction of futures trading is much more often associated with a volatility decrease than with an increase. Note that in nine countries, event windows may be selected to make the volatility effect of futures trading appear significantly positive or significantly negative. We do not place much confidence in these simple tests, inasmuch as they do not account for movements in the world market portfolio, autocorrelation due to infrequent trading, or conditional heteroskedasticity. These problems are addressed in the following sections.

II Volatility Effects of Futures Listing

A 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. This framework is parsimonious, allowing us to capture many of the salient features of the data, and to at least partially account for movements in the world market, in a model with relatively few parameters. Later, we will estimate a multivariate GARCH model that allows 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 is to remove from the time series any predictability associated with lagged returns or day-of-the-week effects. For each country, the following regression is estimated:

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

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 use 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.[7] Regression results are reported in Table 4.

To correct for any remaining predictability, and to correct for spurious autocorrelation induced by non-synchronous trading,[8] we perform the usual autocorrelation adjustment:

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

Table 5 reports parameter estimates for this equation. Following Engle and Ng (1993), we report 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 tell us that the regression model removes serial correlation in the stock return series in most of 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 time varying volatility of index returns in all countries. We take 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 proceed to test for the effect of futures introduction on the conditional volatility of the spot market, using various GARCH specifications.

[7] 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 strictly be interpreted as a conditional volatility.

[8] See Scholes and Williams (1977), Lo and MacKinlay (1988), Nelson (1991).

B 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 from the autoregression equation 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 seems to model the returns data. The results of these specification tests are not reported here, but are available on request. 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 where positive and negative shocks can have different effects. We tested the symmetric model and three alternative asymmetric models including the asymmetric GARCH (GJR-GARCH) model of Glosten, Jagannathan and Runkle (1993), the nonlinear GARCH (NGARCH) model of Engle and Ng (1993), and the exponential GARCH (EGARCH) model of Nelson (1991). Specification tests indicate that these asymmetric models fit the data better then the symmetric GARCH model, with the GRJ-GARCH performing marginally better than the others. Therefore, we base 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 interact 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 zero prior to futures introduction and a value of one after futures introduction. A significant negative parameter estimate for α_M would indicate an decrease in the volatility associated with futures introduction.

Results are reported in Table 6. Defining statistical significance at the five percent level, we find that out of twenty-five 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 sixteen 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.

C 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 6. The results of all these tests are available on 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 GARCH (NGARCH) model of Engle and Ng (1993), and the exponential GARCH (EGARCH) model of Nelson (1991). In some specifications, the volatility increase in Japan loses 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 yield results supporting the same conclusion: outside of the United States and Japan, volatility has tended to decrease with futures listing, or at least to remain unchanged.

In our analysis, we account 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 beta of each country's return with respect to the world is one. In other results, reported in an earlier version of this paper, we also estimated a model where each country had its own beta. The results of this specification are similar to those reported here.

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.

III The Effect of Futures Trading Activity

If stock index futures markets truly have a stabilizing effect on cash markets, as our previous results suggest, then presumably the amount of stabilization would be related to the level of futures market activity. A naive approach to testing this would be to see if prices are more or less volatile in periods when futures trading is more active. The problem with this, of course, is that causality may go both ways: high spot market volatility may induce more people to trade futures.

Bessembinder and Seguin (1992) suggest a way to address this problem. Using an ARIMA model, they decompose 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 find that market volatility is positively related to the unexpected components of volume and open interest, reflecting the positive effect of volatility on volume, but that market volatility is negatively related to the expected component, suggesting an underlying stabilizing influence.

We follow a similar procedure using futures market trading volume and open interest data from seventeen of our twenty-five countries, for which data were available. First, we analyze the volume and open interest time series from each country to select an ARIMA model that appears 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 select, on the basis of the autocorrelation structure, a different model for each time series. The models we selected, along with corresponding Ljung-Box(12) test statistics for model specification, are reported in Table 7.

We then use these models to decompose each time series into expected and unexpected components, and then insert them as additional explanatory variables in the GJR-GARCH conditional volatility equation:

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

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 8. Out of the seventeen countries analyzed, the coefficient α_7 on the unexpected component of open interest is negative in all seventeen, 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 fifteen countries, significantly so in seven. Note that these results very closely correspond to those reported in our earlier analysis. We interpret 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 find, like Bessembinder and Seguin (1992), that it has a positive 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, has 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 twelve countries.

IV Modeling the Joint Dynamics of Country and World Volatility

The univariate models we have 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 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 wish to capture the dynamic interaction between world market volatility, country-specific volatility and conditional covariance, we use 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 estimate the following bivariate process:

$$\begin{aligned} R_{i,t} &= a_0 + \sum_{j=1}^{5} a_j \, R_{i,t-j} + \sum_{k=2}^{5} b_k \, DAY_k + \varepsilon_{i,t} \\ R_{w,t} &= w_0 + \sum_{j=1}^{5} w_j \, R_{w,t-j} + \sum_{k=2}^{5} d_k \, DAY_k + \varepsilon_{w,t} \end{aligned}$$

where the error terms are multivariate normal:

$$\varepsilon_t \mid \mathcal{F}_{t-1} \sim N(\mathbf{0}, \mathbf{H}_t)$$

[9]This model has also 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).

with conditional covariance matrix

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

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 $\mathbf{C}$ represents a matrix of constants, $\mathbf{A}$ represents a matrix of ARCH coefficients, $\mathbf{G}$ represents a matrix of GARCH coefficients and Φ represents a matrix of dummy coefficients. Matrices $\mathbf{H}$, $\mathbf{C}$, $\mathbf{G}$, and $\mathbf{\Phi}$ are symmetric. Our main purpose in using the multivariate GARCH model is 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, and the element Φ_{22} is zero.

In summary,

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

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 are attempting 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 9. Examining the coefficients on the futures introduction dummy variable in the country-specific conditional volatility equation, we find a significant volatility increase in four countries (Germany, Japan, Hungary and Spain) and a significant decrease in twelve countries. Note that under this specification, for the

United States the volatility effect is still positive but is no longer statistically significant. Although these results are not as one-sided as those from the GJR-GARCH model, we still observe a clear propensity for volatility to decrease after futures introduction.

Examining the dummy coefficients in the conditional covariance equations, we find that conditional covariance with the world market increases in twenty-one out of twenty-five countries, with statistical significance in thirteen cases. We may interpret this as evidence that futures markets contribute to an increase in the level of world market integration. On the other hand, 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.

V Conclusion

In this paper, we have examined the time series properties of returns in twenty-five markets around the world before and after the introduction of stock index futures.

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 indicate 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 find the opposite result 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 which allows for the possibility of volatility spillover and time-varying conditional covariance between country-specific and world returns. In this framework, the basic result of our previous analysis is preserved—country-specific conditional variance is likely to decline with the introduction of stock index futures. We also document 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 be 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 really interpret this as we would a traditional event study.[11] Different countries have different contract designs, trading mechanisms, and regulatory environments. Some countries 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 1980's, a group of Western European and other developed markets listing in the late 1980's, and emerging markets listing in the 1990's.

Third, it should be noted that a relatively long time series is required to obtain reliable GARCH parameter estimates. In some cases, particularly for the most recent listings in our sample, our window length may be too short. This may explain the unusual parameter estimates reported for

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

Hungary and Portugal.

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 less-developed markets, the introduction of stock index futures contributes to a decrease in conditional variance.

References

[1] Antoniou, Antonios, Phil Holmes, and Richard Priestley, 1998, The Effects of Stock Index Futures Trading on Stock Index Volatility: An Anslysis of the Asymmetric Response of Volatility to News, Journal of Futures Markets 18, 151-166.

[2] Bessembinder, Hendrik and Paul J. Seguin, 1992, Futures-Trading Activity and Stock Price Volatility, Journal of Finance 47, 2015-2034.

[3] Bollerslev, T. 1986, Generalized Autoregressive Conditional Heteroscedasticity, Journal of Econometrics 31, 307-327

[4] Chan, K.C. and G. Andrew Karolyi, 1991, The Volatility of the Japanese Stock Market: Evidence from 1977 to 1990, in Japanese Financial Market Research, W.T.Ziemba, W. Bailey and Y. Hamao, eds., Elsevier.

[5] Damodaran, Aswath and Marti G. Subrahmanyam, 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.

[6] Engle, Robert F. and Kenneth F. Kroner, 1995, Multivariate Simultaneous Generalized ARCH, Econometric Theory 11, 122-150.

[7] Engle, Robert and Joseph Mezrich, 1995, Grappling with GARCH, Risk, 8, 112-117.

[8] Engle, Robert and Victor Ng, 1993, Measuring and Testing the Impact of News on Volatility, Journal of Finance 48, 1749-1778.

[9] Glosten, Lawrence R., Ravi Jagannathan and David E. Rundle, 1993, On the Relation between the Expected Value and the Volatility of the Nominal Excess Return on Stocks, Journal of Finance 48, 1779-1801.

[10] Hodges, Stewart, 1992, Do Derivative Instruments Increase Market Volatility?, 1992, Options: Recent Advances in Theory and Practice vII (chapter 12), Stewart Hodges, ed., Manchester University Press.

[11] Karolyi, G. Andrew, 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.

[12] Kroner, Kenneth F. and Victor K. Ng, 1998, Modeling Asymmetric Comovements of Asset Returns, Review of Financial Studies 11, 817-844.

[13] Lee, Sang Bin and Ki Yool Ohk, 1992, Stock Index Futures Listing and Structural Change in Time-Varying Volatility, Journal of Futures Markets 12, 493-509.

[14] Lo, Andrew and Craig MacKinlay, 1988, Stock Market Prices Do Not Follow Random Walks: Evidence From Simple Specification Test, Review of Financial Studies 1, 41-66.

[15] Mayhew, Stewart, 1999, The Impact of Derivatives on Cash Markets: What Have We Learned?, Working paper, Purdue University.

[16] Nelson, D. 1991, Conditional Heteroscedasticity in Asset Returns: A New Approach, Econometrica 59, 347-370.

[17] Pagan, A. and G. W. Schwert, 1990, Alternative Models For Conditional Stock Volatility, Journal of Econometrics 45, 267-290.

[18] Rhode, William, 1997, Risk in an Era of Reform, AsiaRisk (Supplement to Risk), October 19-23.

[19] Scholes, Myron and J. Williams, 1977, Estimating Betas from Nonsynchronous Data, Journal of Financial Economics 5, 309-327.

[20] Stewart, Jules and Andrew Priest, 1997, Exchange of Fortunes, LatinRisk (Latin America Supplement to Risk), October 21-23.

[21] Sutcliffe, Charles, 1997, Stock Index Futures: Theories and International Evidence, 2nd ed., International Thomson Business Press.

A Information About Stock Index Futures Markets

For information on plans for derivatives trading in Latin markets, see Stewart and Priest (1997). For a discussion of plans for derivatives in India, see Rhode (1997). More information may be available on the following web sites:

Country	Source of Information	URL
Czech Republic	Prague Stock Exchange	www.pse.cz/defaulten.htm
India	National Stock Exchange	www.nseindia.com
Indonesia	Surabaya Stock Exchange	www.bes.co.id
Mexico	Mexican Stock Exchange	www.bmv.com.mx
Slovakia	Federation of Euro Asian Stock Exchanges	www.feas.org/newsltr.htm
Turkey	Istanbul Stock Exchange	www.ise.org/

Table 1:

Launch Dates for Index Futures Contracts

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

Country	Underlying Index	Launch Date
United States	Value Line	24 Feb 1982
	S&P 500	21 Apr 1982
Australia	All Ordinaries	16 Feb 1983
UK	FT-SE 100	03 May 1984
Canada	TSE 300	16 Jan 1984
Brazil	BOVESPA	14 Feb 1986
Hong Kong	Hang Seng	06 May 1986
Japan (SIMEX)	Nikkei 225	03 Sep 1986
(Osaka)	OSE 50	09 Jun 1987
(Osaka)	Nikkei 225	03 Sep 1988
(Tokyo)	Topix	03 Sep 1988
New Zealand	Barclay Share	Jan 1987
Sweden	OMX	03 Apr 1987
Finland	FOX	02 May 1988
Netherlands	AEX	24 Oct 1988
France	CAC 40	09 Nov 1988
Denmark	KFX	07 Dec 1989
South Africa	All Share	30 Apr 1990
Switzerland	SMI	09 Nov 1990
Germany	DAX	23 Nov 1990
Chile	IPSA	Dec 1990
Spain	IBEX 35	14 Jan 1992
Austria	ATX	07 Aug 1992
Norway	OBX	04 Sep 1992
Belgium	BEL 20	29 Oct 1993
Italy	MIB 30	28 Nov 1994
Hungary	BSI	31 Mar 1995
Israel	Maof 25	27 Oct 1995
Malaysia	KLCI	15 Dec 1995
Korea	KOSPI 200	03 May 1996
Portugal	PSI-20	20 Jun 1996
Russia	RTS	Mar 1997
Venezuela	IBC	05 Sep 1997
Poland	WIG20	16 Jan 1998

Table 2:

Data Periods

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

Country	Data Period	Obs. Pre-	Obs. Post-
Australia	02 Jan 1980 - 31 Dec 1997	779	3572
Austria	20 Nov 1987 - 31 Dec 1997	1162	1334
Belgium	02 Jan 1990 - 31 Dec 1997	941	1037
Canada	02 Jan 1973 - 31 Dec 1997	2740	3516
Chile	02 Jan 1987 - 31 Dec 1997	879	1741
Denmark	10 Dec 1979 - 31 Dec 1997	2476	2037
Finland	02 Jan 1987 - 31 Dec 1997	333	2424
France	09 Jul 1987 - 31 Dec 1997	330	2270
Germany	21 Nov 1977 - 31 Dec 1997	3215	1771
Japan	04 Jan 1980 - 31 Dec 1997	2098	2298
Hong Kong	02 Jan 1973 - 31 Dec 1997	3263	2888
Hungary	02 Jan 1991 - 31 Dec 1997	1056	674
Israel	02 Jan 1992 - 31 Dec 1997	928	527
Italy	02 Jan 1973 - 31 Dec 1997	5507	780
Korea	03 Jan 1990 - 31 Dec 1997	1540	398
Malaysia	02 Jan 1980 - 31 Dec 1997	3902	508
Netherlands	03 Jan 1983 - 31 Dec 1997	1402	2313
Norway	03 Jan 1983 - 31 Dec 1997	2418	1333
Portugal	01 Jan 1993 - 31 Dec 1997	853	376
South Africa	10 Apr 1985 - 31 Dec 1997	1136	1891
Spain	06 Jan 1987- 31 Dec 1997	1238	1501
Sweden	02 Jan 1986 - 31 Dec 1997	311	2694
Switzerland	01 Jul 1988 - 31 Dec 1997	590	1792
United Kingdom	02 Jan 1973 - 31 Dec 1997	2871	3485
United States	02 Jan 1973 - 31 Dec 1997	2340	3967

Table 3:

Variance Ratio Test Results

Results of variance ratio tests of whether stock index volatility changes with the introduction of stock index futures trading. Results are reported for pre- and post-event windows ranging from six months to sixty months, and for the entire sample. An up (down) arrow indicates a statistically significant increase (decrease) at the .05 level. Plusses and minuses indicate changes that were not statistically significant. A blank indicates that either in the pre- or post-event period not enough data are available to fill out the designated event window. In Japan, we used the first listing date for stock index futures on the Singapore Exchange.

Country	Index	6M	12M	24M	60M	All Data
Australia	All Ordinaries	+	-	↓		+
Austria	ATX	↑	-	↓		↓
Belgium	BEL 20	-	+	↓		↓
Canada	TSE 300	-	-	↓	↓	↓
Chile	IPSA	↑	↑	↑		↓
Denmark	KFX	↓	↓	↓	↓	↓
Finland	FOX/HEX	↓	↓			-
France	CAC-40	↓	↓			↓
Germany	DAX 30	↓	↓	↓		↓
Hong Kong	Hang Seng	-	-	↑	-	↓
Hungary	BUX	↓	↑	-		↑
Israel	Maof 25	+	↓	↓		↓
Italy	MIB 30	-	↓	↓		↓
Japan (SIMEX)	Nikkei 225	↑	↑	↑	↑	↑
Korea	KOSPI 200	+	↑			↑
Malaysia	KLCI	+	↓	↑		↑
Netherlands	AEX	↓	↓	↓	↓	↓
Norway	OBX/OSE	+	↓	↓	↓	↓
Portugal	PSI-20	↓	↑			↑
South Africa	All Share	↓	↓	↓	↓	↓
Spain	IBEX 35	↓	↑	↓	↓	↓
Sweden	OMX	↓	↑			↓
Switzerland	SMI	↓	↓	↓		↓
United Kingdom	FT-SE 100	↑	↑	-	↑	↓
United States	S&P 500	↑	↑	+	↑	↑

Table 4:

Coefficients from the first-stage regression

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

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

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

Country	Intercept	R_{Wt-1}	Tuesday	Wednesday	Thursday	Friday
Australia	0.03713 (1.12)	0.37582 (18.47)	-0.12036 (-2.60)	-0.08899 (-1.92)	-0.05075 (-1.10)	-0.02427 (-0.52)
Austria	0.03977 (0.74)	0.18755 (5.17)	-0.08241 (-1.09)	-0.05627 (-0.75)	-0.0206 (-0.27)	-0.01558 (-0.20)
Belgium	-0.03235 (-0.81)	-0.09882 (-3.73)	0.02773 (0.50)	0.0515 (0.92)	0.08215 (1.46)	0.04227 (0.75)
Canada	-0.03279 (-1.74)	-0.02523 (-2.16)	0.02417 (0.93)	0.01898 (0.73)	0.04067 (1.56)	0.0617 (2.36)
Chile	-0.02036 (-0.31)	-0.09196 (-2.35)	0.06182 (0.66)	0.15283 (1.62)	0.13664 (1.46)	0.35277 (3.73)
Denmark	0.07744 (2.15)	0.0316 (1.42)	-0.08543 (-1.69)	-0.09859 (-1.95)	-0.0349 (0.68)	-0.06804 (-1.34)
Finland	0.01714 (0.035)	0.05714 (1.92)	-0.07254 (-1.05)	-0.03055 (-0.44)	0.04796 (0.69)	0.01671 (0.24)
France	-0.13878 (-2.72)	-0.06692 (-2.23)	0.19068 (2.69)	0.17274 (2.44)	0.19733 (2.78)	0.14794 (2.08)
Germany	0.00181 (0.05)	0.00702 (0.32)	-0.00736 (-0.15)	-0.01063 (-0.22)	0.00782 (0.16)	-0.03405 (-0.70)
Hong Kong	-0.11824 (-2.09)	0.33057 (9.41)	0.08932 (1.14)	0.21506 (2.74)	0.06375 (0.81)	0.21209 (2.69)
Hungary	0.09398 (1.17)	0.30634 (5.25)	-0.0519 (-0.46)	0.074 (0.65)	-0.1148 (-1.02)	-0.0072 (-0.06)
Israel	-0.0261 (-0.27)	0.09774 (1.29)	0.136 (0.99)	0.0891 (0.64)	0.1029 (0.75)	-0.0052 (-0.04)
Italy	-0.03453 (-0.89)	-0.0395 (-1.60)	-0.02546 (-0.47)	0.01258 (0.23)	0.12575 (2.30)	0.14268 (2.61)

Table 4: Continued

Coefficients from the first-stage regression

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

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

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

Country	Intercept	R_{Wt-1}	Tuesday	Wednesday	Thursday	Friday
Japan	-0.02319 (-0.68)	0.12491 (6.10)	-0.03524 (-0.73)	0.1518 (0.32)	0.02208 (0.46)	-0.02629 (-0.55)
Korea	-0.11001 (-1.25)	0.06893 (1.19)	-0.0134 (-0.11)	0.1928 (1.55)	-0.0294 (-0.24)	0.0696 (0.56)
Malaysia	-0.12483 (-2.57)	0.11017 (3.69)	-0.00552 (-0.08)	0.1473 (2.16)	0.16837 (2.47)	0.19752 (2.89)
Netherlands	-0.03254 (-0.81)	-0.06764 (-2.72)	0.08861 (1.58)	0.08935 (1.60)	-0.0007 (-0.01)	0.07009 (1.24)
Norway	-0.00539 (-0.13)	0.11298 (4.33)	-0.0183 (-0.31)	-0.01034 (-0.18)	0.04217 (0.72)	0.10545 (1.80)
Portugal	0.0026 (0.05)	0.08223 (1.84)	0.0076 (0.10)	0.07838 (1.02)	0.06654 (0.86)	0.03059 (0.39)
South Africa	-0.03475 (-0.68)	0.0238 (0.84)	0.00832 (0.12)	0.14594 (2.04)	0.07647 (1.06)	-0.013 (-0.18)
Spain	0.12966 (2.67)	0.12702 (4.32)	-0.12834 (-1.87)	-0.22861 (-3.33)	-0.14339 (-2.08)	-0.10601 (-1.54)
Sweden	-0.01588 (-0.32)	0.06626 (2.21)	0.02283 (0.33)	0.02446 (0.35)	0.06669 (0.95)	0.06367 (0.91)
Switzerland	-0.01227 (-0.29)	-0.03589 (-1.29)	0.00782 (0.13)	0.09924 (1.66)	0.07047 (1.17)	0.04842 (0.80)
United Kingdom	-0.07092 (-2.55)	0.02278 (1.30)	0.15129 (3.91)	0.05893 (1.53)	0.06196 (1.61)	0.1269 (3.28)
United States	0.03189 (1.67)	-0.16465 (-13.60)	-0.00269 (-0.10)	-0.04298 (-1.61)	-0.03865 (-1.44)	-0.03503 (-1.31)

Table 5:

Coefficients from the residual autoregression

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. t-statistics are shown in parentheses. Ljung-Box statistics testing for 12th order serial autocorrelation in ε and ε^2 are also reported.

Country	Constant	b_1	b_2	b_3	b_4	b_5	LBQ(12) (Levels)	LBQ(12) (Squares)
Australia	0.0232 (1.56)	0.00006 (0.00)	0.1012 (6.81)	-0.0134 (-0.90)	0.0253 (-1.21)	-0.0181 (1.69)	8.4	630.27
Austria	0.031 (1.55)	-0.00009 (-0.00)	0.2276 (11.36)	-0.021 (-1.02)	-0.0246 (-1.20)	-0.0125 (-0.61)	20.6**	438.65
Belgium	-0.0358 (-1.59)	0 (0.0)	0.0274 (1.22)	0.066 (2.93)	0.0399 (1.77)	-0.0026 (-0.12)	6.9	271.55
Canada	0.0104 (0.82)	0 (-0.0)	0.0808 (6.39)	-0.0023 (-0.18)	0.01 (0.79)	0.0272 (2.14)	7.9	639.94
Chile	-0.00011 (-0.00)	0.1981 (10.12)	-0.0534 (-2.68)	-0.0394 (-1.98)	0.0598 (3.00)	-0.005 (-0.26)	7.1	176.466
Denmark	-0.00001 (-0.00)	0.0947 (6.36)	-0.0037 (-0.25)	-0.0017 (-0.11)	-0.0402 (-2.69)	0.0051 (0.34)	7.4	41.96
Finland	0.00005 (0.00)	0.1764 (9.25)	0.0001 (0.00)	0.0052 (0.27)	0.0153 (0.79)	0.022 (1.15)	10.2	340.722
France	0.00002 (0.00)	-0.0797 (-4.06)	0.0533 (2.70)	-0.0059 (-0.30)	0.0016 (0.08)	0.014 (0.71)	15.5*	917.54
Germany	0.00001 (0.00)	-0.0307 (-2.16)	-0.0001 (-0.01)	0.02 (1.41)	-0.0164 (-1.16)	0.0141 (1.00)	4.8	438.591
Hkong	-0.00003 (-0.00)	0.0432 (3.39)	-0.0237 (-1.85)	0.0601 (4.71)	-0.0093 (-0.73)	-0.0183 (-1.43)	18.9**	1083.33
Hungary	0.0003 (0.01)	0.1301 (5.40)	0.0531 (2.19)	0.007 (0.29)	-0.0113 (-0.46)	0.0141 (0.59)	33.8**	727.973
Israel	-0.00008 (-0.00)	-0.0174 (-0.66)	-0.0365 (-1.39)	-0.0207 (-0.79)	0.0117 (0.45)	-0.0581 (-2.22)	16.3*	223.239
Italy	0.00001 (0.00)	0.1592 (12.61)	-0.0717 (-5.61)	0.0392 (3.06)	0.0013 (0.10)	-0.001 (-0.08)	12.9	2050.29

Table 5: Continued

Coefficients from the residual autoregression

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. t-statistics are shown in parentheses. Ljung-Box statistics testing for 12th order serial autocorrelation in ε and ε^2 are also reported.

Country	Constant	b_1	b_2	b_3	b_4	b_5	LBQ(12) (Levels)	LBQ(12) (Squares)
Japan	-0.00004 (0.00)	-0.0792 (-5.24)	-0.0668 (-4.41)	0.0229 (1.51)	-0.0021 (-0.14)	-0.0061 (-0.40)	8.1	801.76
Korea	0.00038 (0.01)	-0.0021 (-0.09)	-0.0422 (-1.85)	-0.0219 (-0.96)	-0.0372 (-1.62)	-0.0528 (-2.30)	11.9	1009.59
Malaysia	0.00003 (0.00)	0.1209 (8.02)	0.007 (0.46)	-0.0114 (-0.75)	0.0065 (0.43)	0.0105 (0.70)	10.9	1352.07
Netherlands	0.00006 (0.00)	-0.0971 (-5.92)	0.0274 (1.66)	0.0257 (1.56)	-0.0166 (-1.01)	0.0453 (2.76)	21.4**	780.3
Norway	0.00001 (0.00)	0.1182 (7.24)	-0.0068 (-0.41)	-0.0129 (-0.78)	-0.0466 (-2.84)	0.0373 (2.28)	14.9*	501.91
Portugal	-0.00004 (-0.00)	0.1516 (5.30)	0.049 (1.69)	-0.0522 (-1.80)	-0.0074 (-0.26)	-0.0127 (-0.44)	13.8	154.161
South Africa	-0.00002 (-0.00)	0.0834 (4.58)	-0.0252 (-1.38)	0.0018 (0.10)	-0.0218 (-1.20)	-0.0308 (-1.69)	9.6	38.1064
Spain	-0.00008 (-0.00)	0.0915 (4.78)	0.0149 (0.78)	-0.054 (-2.82)	0.0364 (1.89)	-0.0186 (-0.97)	18.4*	782.843
Sweden	0.00018 (0.01)	0.0555 (3.04)	0.0252 (1.38)	-0.0135 (-0.74)	0.0106 (0.58)	0.0178 (0.97)	11.6	715.95
Switzerland	0 (0.00)	-0.02 (-0.98)	0.0089 (0.43)	-0.0404 (-1.97)	0.0052 (0.25)	-0.0248 (-1.21)	7.2	56.3926
UK	-0.00001 (-0.00)	0.0761 (6.06)	0.0107 (0.85)	0.0101 (0.81)	0.004 (0.32)	-0.0099 (-0.79)	13.5	4791.4
US	-0.000025 (-0.00)	-0.1303 (-10.34)	-0.0076 (-0.60)	-0.0187 (-1.48)	-0.0513 (-4.04)	-0.0097 (-0.77)	20.2**	1252.86

Table 6:

Effect of Futures Introduction on GJR-GARCH Volatility

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 by the α_3 and α_M coefficients indicates statistical significance at the 5% level.

Country	α_0	α_1	α_2	α_3	α_M
Australia	0.1311	0.7463	0.191	-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	0.045	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	0.087	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	-0.181*	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	0.902	0.0749	-0.013	-0.0094*
	(0.007)	(0.008)	(0.007)	(0.007)	(0.004)

Table 6: Continued

Effect of Futures Introduction on GJR-GARCH Volatility: Multiplicative Dummy

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 by the α_3 and α_M coefficients indicates statistical significance at the 5% level.

Country	α_0	α_1	α_2	α_3	α_M
Japan	0.0309 (0.004)	0.8725 (0.010)	0.0664 (0.005)	0.045* (0.009)	0.0127* (0.004)
Korea	0.1154 (0.025)	0.8351 (0.021)	0.0799 (0.014)	0.09* (0.022)	0.0149 (0.012)
Malaysia	0.0817 (0.009)	0.8225 (0.009)	0.1191 (0.007)	0.0437* (0.010)	-0.0211* (0.009)
Netherlands	0.1216 (0.013)	0.8071 (0.015)	0.0771 (0.012)	0.0931* (0.013)	-0.0836* (0.010)
Norway	0.0524 (0.005)	0.8674 (0.005)	0.0975 (0.006)	0.012 (0.011)	-0.0371* (0.005)
Portugal	0.3137 (0.050)	0.2647 (0.084)	0.2987 (0.031)	0.0129 (0.048)	-0.0115 (0.048)
South Africa	0.2772 (0.028)	0.7237 (0.019)	0.3579 (0.028)	-0.217* (0.035)	-0.1817* (0.013)
Spain	0.0702 (0.010)	0.8157 (0.016)	0.1448 (0.013)	-0.0366* (0.015)	-0.0012 (0.007)
Sweden	0.0495 (0.009)	0.8988 (0.011)	0.0499 (0.008)	0.0492* (0.011)	-0.0113 (0.008)
Switzerland	0.2491 (0.058)	0.6765 (0.060)	0.1098 (0.021)	0.039 (0.023)	-0.1255* (0.029)
United Kingdom	0.011 (0.002)	0.937 (0.005)	0.0474 (0.005)	0.013 (0.007)	-0.0096* (0.003)
United States	0.0046 (0.001)	0.9204 (0.005)	0.0446 (0.005)	0.0309* (0.005)	0.0133* (0.002)

Table 7:

ARIMA Specifications for Volume and Open Interest

This table reports the ARIMA models used to decompose futures volume and open interest series into expected and unexpected components. The models were selected using the standard Box-Jenkins approach. The corresponding Ljung-Box test statistics are reported. An asterisk indicates statistical significance at the one percent level, meaning that some significant twelfth-order serial correlation remains in the residuals.

Country	Volume	LBQ(12)	Open Interest	LBQ(12)
Australia	ARIMA(4,1,3)	34.4*	ARIMA(5,1,5)	57.5*
Austria	ARIMA(2,0,5)	10.8	ARIMA(1,1,5)	12.6
Belgium	ARIMA(5,0,0)	8.4	ARIMA(2,1,1)	14.4
Canada	ARIMA(5,1,5)	1.3	ARIMA(0,1,5)	26.3*
Denmark	ARIMA(2,0,5)	10.5	ARIMA(5,1,0)	6.3
France	ARIMA(4,1,4)	48.5*	ARIMA(1,1,1)	10.8
Germany	ARIMA(4,1,2)	9.5	ARIMA(4,1,2)	9.8
Hong Kong	ARIMA(2,1,5)	34.8*	ARIMA(1,1,3)	11.3
Italy	ARIMA(4,1,3)	18.8*	ARIMA(4,1,1)	3.1
Japan	ARIMA(5,1,5)	10.9*	ARIMA(0,1,5)	10.6
Korea	ARIMA(4,1,5)	14.0*	ARIMA(2,1,4)	2.1
Netherlands	ARIMA(5,0,1)	35.2*	ARIMA(1,1,5)	3.9
Norway	ARIMA(5,0,0)	20.8*	ARIMA(1,1,5)	10.1
Portugal	ARIMA(3,1,1)	11.9	ARIMA(3,1,5)	1.3
Spain	ARIMA(5,0,0)	17.3	ARIMA(3,1,3)	8.4
Sweden	ARIMA(4,0,3)	8.8	ARIMA(2,1,5)	11.2
Switzerland	ARIMA(5,0,0)	7.7	ARIMA(1,1,5)	10.3
United Kingdom	ARIMA(5,0,0)	47.8*	ARIMA(1,1,1)	22.2
United States	ARIMA(4,1,1)	102.5*	ARIMA(5,1,5)	5.9

Table 8: **Effect of Futures Trading Activity on Volatility**

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

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

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.

Country *Vol* | *OI*	α_0	α_1	α_2	α_3	α_4	α_5	α_6	α_7
Australia 10^4 | 10^5	0.1342 (0.166)	0.7087 (0.026)	0.1745 (0.009)	-0.0941 (0.016)	0.0346 (0.046)	0.4671* (0.059)	-0.0508 (0.042)	-0.06379 (0.163)
Austria 10^4 | 10^5	0.2032 (0.027)	0.82 (0.033)	0.0963 (0.022)	-0.01 (0.026)	-0.6353* (0.182)	2.0146* (0.272)	-0.0812 (0.069)	-0.7615* (0.204)
Belgium 10^4 | 10^4	0.0547 (0.022)	0.8425 (0.048)	0.0575 (0.019)	0.0064 (0.028)	-0.0077 (0.124)	0.3832 (0.266)	-0.0133 (0.014)	-0.4163 (0.217)
Canada 10^3 | 10^4	0.0263 (0.004)	0.8455 (0.015)	0.1432 (0.015)	-0.0767 (0.014)	-0.0093 (0.008)	0.0491* (0.015)	-0.0009 (0.005)	-0.2244* (0.057)
Denmark 10^4 | 10^4	0.2727 (0.047)	0.3764 (0.064)	0.2096 (0.036)	-0.0904 (0.044)	1.9812* (0.525)	3.8229* (0.327)	-0.2113* (0.076)	-0.3571 (0.829)
France 10^5 | 10^6	0.295 (0.053)	0.5485 (0.070)	0.0781 (0.025)	0.0479 (0.035)	0.7452 (0.391)	3.0331* (0.334)	-0.396 (0.289)	-2.6735 (2.838)
Germany 10^5 | 10^5	0.3404 (0.050)	0.4334 (0.072)	0.033 (0.029)	0.0899 (0.046)	1.0981* (0.289)	4.2413* (0.364)	-0.1279* (0.040)	-0.2444 (0.390)
Hong Kong 10^5 | 10^5	0.4442 (0.055)	0.4131 (0.056)	0.1507 (0.033)	0.131 (0.049)	2.6692* (1.202)	11.2339* (0.817)	-0.539 (0.449)	-5.5596* (1.484)
Italy 10^5 | 10^5	0.5237 (0.151)	0.3059 (0.130)	0.2601 (0.081)	-0.1106 (0.099)	1.8269 (1.542)	8.5268* (1.126)	-0.631 (0.969)	-5.899* (2.053)
Japan 10^4 | 10^6	0.0121 (0.003)	0.9111 (0.010)	0.0217 (0.007)	0.0979 (0.013)	-0.003 (0.010)	0.4131* (0.098)	0.0757* (0.025)	-0.4698 (1.201)
Netherlands 10^4 | 10^5	0.0488 (0.010)	0.9156 (0.016)	0.0474 (0.013)	0.0022 (0.013)	-0.0122 (0.026)	0.2918* (0.077)	-0.0972* (0.044)	-0.7806* (0.388)
Norway 10^3 | 10^4	0.0403 (0.010)	0.8644 (0.019)	0.0965 (0.015)	-0.0026 (0.025)	0.065 (0.062)	0.1827* (0.085)	-0.1504* (0.068)	-0.5672 (0.515)
Spain 10^5 | 10^5	0.2958 (0.051)	0.6943 (0.071)	0.0484 (0.025)	0.0993 (0.037)	0.3125 (0.506)	4.8436* (0.502)	-0.3706* (0.160)	-0.9556 (0.764)
Sweden 10^4 | 10^5	0.0009 (0.072)	0.8883 (0.059)	0.0032 (0.027)	0.0535 (0.038)	0.1223 (0.140)	0.2158 (0.167)	-0.0172 (0.127)	-0.1134 (0.366)
Switzerland 10^4 | 10^5	0.2207 (0.036)	0.7317 (0.082)	0.0357 (0.021)	0.0523 (0.030)	-0.0024 (0.097)	0.8582* (0.085)	-0.4029* (0.175)	-2.3471* (1.132)
United Kingdom 10^4 | 10^5	0.0393 (0.007)	0.9036 (0.016)	0.045 (0.010)	0.0164 (0.011)	-0.0103* (0.005)	0.1071* (0.018)	-0.0081 (0.005)	-0.3577* (0.115)
United States 10^5 | 10^6	0.0052 (0.002)	0.9193 (0.005)	0.0554 (0.007)	0.0231 (0.007)	0.0007 (0.003)	0.2241* (0.017)	0.0083 (0.005)	-0.7909* (0.168)

Table 9:

Joint Dynamics of Country and World Returns

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:

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

For brevity, coefficients specific to the world conditional variance equation are not reported. An asterisk by the d_{11} and d_{12} coefficients indicates statistical significance at the 5% level.

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	0.347	0.0192	0.0054	0.842	0.8412	-0.004*	-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	0.027	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	0.4576	-0.0112	0.324	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	0.403	0.0007	0.089	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	0.093	-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	0.146	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	0.165	-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	0.013	-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)

Table 9: Continued

Joint Dynamics of Country and World Returns

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:

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

For brevity, coefficients specific to the world conditional variance equation are not reported. An asterisk by the d_{11} and d_{12} coefficients indicates statistical significance at the 5% level.

Country	c_{11}	c_{12}	a_{11}	a_{12}	a_{21}	g_{11}	g_{12}	d_{11}	d_{12}
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	0.018	0.0874	0.8629	0.7726	-0.03*	0.0314*
	(0.0177)	(0.0095)	(0.0137)	(0.0068)	(0.0176)	(0.0113)	(0.0127)	(0.0145)	(0.0098)
Switzerland	0.2279	-0.01	0.2863	0.02	0.3394	0.652	0.6596	-0.0667*	0.0239*
	(0.0240)	0.0090	(0.0229)	(0.0105)	(0.0176)	(0.0288)	(0.0172)	(0.0117)	(0.0093)
UK	0.0279	0.0056	0.2607	0.0073	0.0678	0.8986	0.863	-0.0129*	0.001
	(0.0040)	(0.0019)	(0.0093)	(0.0050)	(0.0071)	(0.0057)	(0.0075)	(0.0030)	(0.0016)
US	0.0106	0.0061	0.2203	-0.0194	0.0074	0.938	0.9419	0.0005	-0.0006
	(0.0013)	(0.0007)	(0.0035)	(0.0026)	(0.0062)	(0.0025)	(0.0021)	(0.0009)	(0.0004)

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