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INTERNATIONAL STOCK PRICE SPILLOVERS AND MARKET LIBERALIZATION: EVIDENCE FROM KOREA, JAPAN, AND THE UNITED STATES

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Abstract

In August 1991 the Korean government announced that the stock exchange would undergo a significant liberalization in January 1992, by allowing foreigners to directly own shares in Korean stocks. This paper examines the repercussions on the relationship between the stock markets of Korea, Japan, and the United States. We estimate GARCH models to quantify the importance of "volatility spillovers" from Japan and the U.S. on the mean and variance of Korean returns. Such spillovers have increased since the announced opening, with most of the effect on the opening prices of the Korean stock market.

International Stock Price Spillovers and Market Liberalization: Evidence From Korea, Japan, and the United States

Sang W. Kim and John H. Rogers[1]

I. INTRODUCTION

In spite of the substantial recent literature on the global integration of financial markets among the four major stock markets, there has not been much discussion about the emerging markets of developing nations. Although most of these markets have been relatively closed to foreign investors, the recent trend has been to open them. Thus, while official entry barriers may have so far prevented interactions between the emerging and major markets, such interactions may increase in the future.

With the increasing integration of major financial markets around the world, the transmission of stock return movements among the major markets has become a much researched topic. Eun and Shim (1989) apply the vector autoregression (VAR) methodology to daily market index data in a study of the transmission of stock return movements among the world's nine largest stock markets. They find a substantial amount of multi-lateral interaction among the markets, with the U.S. stock market as the most influential. Jeon and von Furstenberg (1990), also using VARs, show evidence of increased international integration of the four major equity markets in the 1980s. Similarly, Cumby (1990) shows that real stock returns in the major markets are more consistent with consumption-based models of international asset pricing in a post-1980 sample than in a sample beginning in 1974.

Studying the transmission of stock market movements is a joint study of the spillover of price changes and the volatility of price changes. Ultimately, it is the perceived importance of the

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information contained in other markets' price movements that influences investors in the market to which the spillover occurs. As indicated by several studies, beginning with the seminal work of Kyle (1985), much of the information would be revealed in the volatility of stock price movements rather than the price movements themselves. King and Wadhwani (1990) develop a contagion model where volatility is transmitted across markets, and examine a data set composed of fifteen minute index returns for New York, London, and Tokyo.

Because stock return data display volatility clustering, where one period of high volatility is followed by another [see, e.g., Schwert (1989)], applying the VAR methodology to the study of the transmission of stock price movements may not allow a study of all aspects of the transmission of price movements. The Generalized Autoregressive Conditional Heteroskedasticity (GARCH) methodology, introduced by Engle (1982) and Bollerslev (1986), models conditional variance as a normally distributed Autoregressive Moving Average (ARMA) process, and thus is able to model the fat tailed distributions of stock returns that display volatility clustering. GARCH has been used extensively to model financial data, as we do below.[2]

Among such studies, Hamao, Masulis, and Ng (1990) is the closest in nature to this paper. By using the GARCH methodology, Hamao et al. are able to model the foreign market's volatility directly into both the mean and variance equations of their model. Using daily index returns from 1985 to 1987, they find significant spillover effects: from New York to Tokyo and London and from London to Tokyo. This agrees with the findings of researchers who used the VAR methodology that the transmission of stock price movements among major stock markets is quite significant.

Other researchers applying GARCH have found increasing volatility spillovers among foreign exchange markets. Engle, Ito, and Lin (1991) examine whether the volatility clustering in exchange rate data is mainly a one country phenomenon or whether there exists a transmission of volatility around the globe. Most of the volatility clustering is attributed to volatility spillovers. Baillie and

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[2]The survey by Bollerslev, Chou, and Kroner (1992) lists over 300 articles that apply GARCH.

Bollerslev (1991) study spot exchange rate volatility and find hourly patterns to be similar around the globe, with significant spillovers between major currency markets.

Given these studies, and the recent opening of several emerging capital markets, it is worth seeing if these countries display increased receptivity to the transmission of stock price movements and their volatility from major markets. Because the stock markets of interest are open at different times around the world we need intra-day data. During the regular trading hours of the New York Stock Exchange, the stock markets in, e.g., East Asia have already completed their trading day. Thus, investors in Asia will have information on the previous day's stock price movements in New York before the commencement of trading in their own market. If the information is pertinent to investors in the Asian markets, and there is market efficiency with respect to information, there will be spillover effects showing up in the opening prices of the Asian markets. Another reason to expect such spillovers comes from portfolio management/hedging strategy. Harvey (1993) suggests investing in the emerging markets as a way to enhance and optimize a portfolio. Harvey (1993) and Ferson and Harvey (1993) find that emerging market returns (i) are predictable, (ii) yield high returns, and (iii) have low correlations with the major markets. These present portfolio and fund managers with an attractive option for enhancing their investment portfolios. As such investments increase, any spillovers that would occur due to information coming from New York trading would show up in the opening price of the liberalized emerging market. However, reliable opening price data for most of these emerging markets are notoriously difficult to obtain. In this paper we focus on Korea, for which we do have good data, and the transmission of stock price movements and volatility from Japan and the United States.

According to the International Monetary Fund, the Korean stock market has been among the most restrictive and tightly controlled of the emerging markets. For years, the market was open to foreign investors only through special funds, such as the closed end Korea Fund which is traded on the New York Stock Exchange. Although the Korean economy itself began a liberalization program as early as 1985 with direct foreign investment in some industries, opening the stock market was only a rumor as late as 1988, and even that was to be with severe restrictions on foreign investors. Finally, in August 1991, the Korean Ministry of Finance announced that the stock market would be open to direct investment by foreign nationals beginning in January 1992. The announced restrictions, especially on the key issue of the repatriation of funds, were significantly less than expected, although total foreign ownership in any one company is to be limited to 10%.

From the time of this announcement until early 1992, the Korea Fund's premium, the share price over the net asset value of the underlying portfolio, dropped steadily. For example, on August 12, 1991, the premium was 51.25%, while on February 17, 1992 it was 14.62%. This could mean that either the value of stocks in the Korean market rose or that investors who held shares in the Korea Fund were investing directly in the Korean market. The data indicate the latter: between August 12, 1991 and February 17, 1992, the Korea Fund share price fell from 17½ to 13¼, while the net asset value of the underlying portfolio changed only from 11.57 to 11.56. While these numbers do not necessarily indicate significant increases in foreign investment directly in the Korean Stock Market, they are quite suggestive. An important issue for Korea is whether the rising foreign investment and capital inflow will increase the transmission of stock return movements and their volatility from the major stock markets.

In this paper, we use GARCH methodology to examine whether there has been any change in the transmission of volatility from Japan and the U.S. to Korea following the liberalization announcement. We examine the full data set and a subperiod after the announcement, which we call the "post-announcement" subperiod, starting in September 1, 1991. As indicated by the fall in the Korea Fund premium, the impact of the announcement was fairly immediate. We examine whether the volatility of returns in the two major markets have any spillover effects on the Korean market, and if this spillover has increased in the post-announcement subperiod. We find that volatility spillovers have increased, particularly spillovers from Japan, and that the effect is on the volatility of returns much more than on returns themselves. We also find that such spillovers affect opening prices more than closing prices.

In section II we describe the data set used. In section III, we set up and estimate GARCH models for intra-day stock returns and perform various hypothesis tests. Section IV concludes.

II. PROPERTIES OF STOCK RETURN DATA

For this study, daily opening and closing quotes of the major stock market indices of Korea, Tokyo, and New York are collected. For Korea, 9:45 A.M. quotes of the Korea Composite Stock Price Index (KOSPI) of the Korea Stock Exchange (KSE) are used for all opening prices, and 3:20 P.M. quotes are used for closing prices on Mondays through Fridays (12:00 P.M. quotes are used for Saturday, when there is a half day trading). The KOSPI is composed of all 668 firms traded in the Korea Stock Exchange as of 1991. For Japan, we use the Nikkei 225 (NIKKEI) of the Tokyo Stock Exchange (TSE), an index of "first section" stocks of the largest 225 firms in Japan. We collected 9:01 A.M. and 3:00 P.M. quotes for opening and closing prices, respectively. Until January 1989, TSE was open on the second and fourth Saturdays of the month. The closing prices for these Saturdays are 12:00 P.M. quotes. For the U.S., we use the Standard and Poors 500 Composite Index (S&P), whose stocks are mostly listed in the New York Stock Exchange (NYSE) and the American Stock Exchange (ASE). We use 9:31 A.M. and 4:00 P.M. quotes for opening and closing prices. All reported times are local and all three data series range from October 2, 1985 to March 23, 1992. Figure 1 shows the trading hours of the three exchanges in Eastern Standard Time. We calculate two types of continuously compounded index returns: (OC) open-to-close return ($\log[close_t]-\log[open_t]$)x100 and (CO) close-to-open return ($\log[open_t]-\log[close_{t-1}]$)x100, where $open_t$ and $close_t$ stand for opening and closing index levels.

Several researchers have found that the trading mechanism and institutional features of a stock exchange can have significant effects on price volatility and correlations, especially at the market open and close. For example, Stoll and Whaley (1990) find on average that NYSE stocks do not begin trading until 4.15 to 5.46 minutes after the opening of the exchange. This could present some difficulties for us because, in constructing the opening price index, the closing quotes of the previous day would be substituted for those stocks not yet traded. This "stale prices" problem may also be present in our Nikkei opening prices, although we are not aware of similar studies for NIKKEI data. Amihud and Mendelson (1991) find that the opening auction or *Iyatose* may account for up to 11.6% of the total daily trading volume; this may reduce the magnitude of the problem in our Japanese data.

The KSE operates as a continuous auction market where an auctioneer matches best bids and offers. This is very similar to the trading mechanism found in the TSE. On Mondays through Fridays there are morning and afternoon trading sessions which begin at 9:40 A.M. and 1:30 P.M. and end at 12:00 P.M. and 3:20 P.M., respectively, while on Saturdays there is just one morning session. In addition, the opening prices of the morning and afternoon sessions are determined by an opening auction where the price is determined by referring to all bids and offers in the first five minutes of the trading session. Because we do not know the extent of trading volume in the opening auction, we cannot speculate on the extent of the "stale prices" problem for our KOSPI data.

The KSE also limits the amount of daily change in a stock price by employing a schedule of "bands" within which the price of any given stock may fluctuate. These bands are determined according to the market price of the individual stock and on average the width of the band ranges from approximately $\pm 4\%$ to $\pm 6\%$ for stocks priced at 50,000 Won or less, to $\pm 2\%$ to $\pm 3\%$ for stocks priced higher. If these bands are met frequently during trading, this could potentially cause significant problems, the most obvious of which is artificially low volatility. As prices are forced to fluctuate only within the band on any given day, computed return volatility will reflect true volatility plus some unknown negative bias. Furthermore, the volatility may be spread out as investors take several days to make their trades using information available to them. Similarly, computed returns will not reflect the "true" return but the sum of the true return and some bias. Also, as stock prices take days rather than hours to reach the "true" price, this may cause some short term autocorrelation in daily returns. This autocorrelation should occur in addition to that caused by non-synchronous trading and minimum-sized price changes [see Scholes and William (1977)]. Our autocorrelation analysis, which is available on request, indicates that the KOSPI returns display a short term autocorrelation of only one lag, which is comparable to that of NIKKEI and S&P. This leads us to believe that the autocorrelation problem caused by the "bands" may not be very large. In general, although there exist different ways to empirically account for these problems, we do not know just how significant they are, and so will proceed on the assumption that they are small enough not to affect the main results of the paper.

Finally, our KOSPI and NIKKEI data contain observations for the half-day Saturday trading. In a study of "day-of-the-week (DOW) effects" in this data, Kim (1994) finds that the variance of Saturday KOSPI returns are significantly smaller than other days. For the NIKKEI data, Kim found that alternate Saturday half day trading did not cause a similar low variance; he does find a negative Monday DOW effect in variance and positive Tuesday DOW effect in mean. For the S&P data, Kim finds no DOW effects in the returns although he reports a negative Wednesday DOW effect in variance. This is contrary to earlier findings of Monday DOW effects [*e.g.*, French (1980)].

The sample means and variances for all three data sets are reported in Table 1. In general, most mean returns are either zero or close to zero and the OC return variances are larger than the CO return variances. We also report the coefficient of skewness, S, and coefficient of kurtosis, K. As these statistics show, our return series are not normally distributed and are highly leptokurtic. Also, the differences in our full and post-announcement subperiod statistics suggest a time varying volatility.

Cross-correlations for the return levels are reported in Table 2. These cross-correlations are all fairly small compared to those between major markets, a result also reported by Ferson and Harvey (1993). In general, the cross-correlations between the KOSPI and NIKKEI returns are higher than those between KOSPI and S&P. They are also higher for CO KOSPI returns than OC. We should note that, as shown in Figure 1, at the opening of the KSE, traders in Korea have, along with all domestic public and private information, the previous day's New York and Tokyo opening and closing prices, and the same day's opening prices in Tokyo since the KSE opens thirty minutes after the TSE. The efficient market hypothesis would predict that Korean investors will incorporate their knowledge of the previous day's market movements in New York and Tokyo as well as the same day

Tokyo opening prices into their own opening bids. Any resulting spillovers would affect KOSPI prices much more at the opening than closing, and thus should have stronger effects on CO returns. Also, at KSE's market close, there is no new information from New York (NYSE has not opened yet for the new trading day) while there is new information from Tokyo about the TSE's trading activities and market closing data. Thus, KOSPI closing prices will only reflect the spillovers from Tokyo. This suggests that spillovers from New York will only occur to KOSPI CO returns while spillovers from Tokyo will occur to both OC and CO returns.[3] Finally, note that the cross-correlations are generally higher in the post-announcement subperiod, although the post-announcement correlations are still not as large as those between the major markets [*cf.* Hamao et al. (1991)].

Because our main interest lies with volatility spillovers, it is worth examining whether similar cross-correlation patterns exist in the squared returns. Table 2 indicates that the correlation of the squares is generally higher than the levels. More importantly, the increase in correlation in the post-announcement subperiod is quite large, much larger than those observed for the levels. It is worth noting that the full sample and post-announcement subperiod correlations for the *squares* are relatively much higher between KOSPI and NIKKEI than between KOSPI and S&P compared to those for the *return levels*. Thus, if volatility spillovers are an indication of the transmission of information, there seems to be much more information passed from Japan than the United States. Furthermore, this seems to have increased with the liberalization.

III. GARCH MODELS OF KOSPI RETURNS

The review of recent research in section I and the statistical and time series properties of our data set described in section II indicate the appropriateness of using the GARCH methodology

[3]It is also possible that geographic proximity and high volume of trade between Korea and Japan leads to a much stronger relationship between the TSE and KSE than between the NYSE and KSE.

pioneered by Engle (1982) and Bollerslev (1986).[4] In this section we set up and estimate our GARCH model of OC and CO KOSPI returns.

3.1. A Basic GARCH(1,1)-M-*t* Model of OC KOSPI Returns

We first set up and estimate a GARCH model for KOSPI, NIKKEI, and S&P returns to examine volatility spillovers in Korea's recently-liberalized stock market. Asset pricing models in finance all have a variable that accounts for the mostly positive relation between risk and return. Of these, the most widely used model of stock returns is the Capital Asset Pricing Model (CAPM), which expresses stock returns as a linear function of risk. Empirical tests to determine the most appropriate form of the risk terms in the CAPM equation have been carried out by several researchers with varying results. French et al. (1987) and Baillie and DeGennaro (1990) are among the numerous GARCH studies of the relation between risk and stock returns.

A mean-variance efficient CAPM model of stock returns can be expressed as a GARCH in mean (GARCH-M) model, which simply adds a conditional variance as a term in the mean equation along with other explanatory variables and with the conditional variance specified as GARCH(1,1) with some independent variables. Our first model, which will be used to check for the significance of the liberalization announcement, is shown in equation (3.1) below. This model, which will be referred to as Model 1, was estimated using an iterative Maximum Likelihood procedure applying the Berndt, Hall, Hall, and Hausman algorithm using numerical derivatives:

$$\begin{aligned} R_t^i &= \gamma_0 + \gamma_1 h_t^{1/2} + \gamma_2 DL_t + \theta\,\varepsilon_{t-1} + \varepsilon_t\,, \\ h_t &= \alpha_0 + \alpha_1 \varepsilon_{t-1}^2 + \beta_1 h_{t-1} + \delta_1 DL_t + \delta_2 DS_t\,, \end{aligned} \tag{3.1}$$

where, R_t^i is the KOSPI return (i=OC,CO) and h_t is the conditional variance. Estimation results are reported in Panel A of Table 3.

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[4]Although not reported, we also carried out Engle's ARCH test to check for the appropriateness of applying GARCH. Our results are similar to those reported by other researchers [see Engle and Susmel (1993)]. Like others using high frequency asset returns, we also find leptokurtosis and time-varying volatility. This, along with Q tests and the higher cross-correlations in the post-announcement subperiod (especially in the squares), provides the basis for the following.

In the mean equation, Model 1 includes a risk term, $h_t^{1/2}$, which is expressed as a square root of the contemporaneous conditional variance. The coefficient on $h_t^{1/2}$ is large and significant, meaning that the risk-return tradeoff in the KOSPI OC returns is significant. The mean equation also includes a first order moving average or MA(1) term. Inclusion of the MA(1) term is motivated by the slight first order serial correlation found in the return data, which may have been induced by non-synchronous trading of individual stocks, minimum sized price changes, bid-ask spreads [see Scholes and Williams (1977)], and, if binding, the price change "bands" described in section II. However, as shown by Blume and Stambaugh (1983), the bid-ask spread effect is less significant in aggregated stock portfolio returns such as those used in this study, so most of the serial correlation in our data is likely to be the result of non-synchronous trading and minimum sized price changes. For completeness, an MA(1) term is included in all of our models.

The mean equation also includes a "post-announcement" dummy, DL_t, which is 1 in the post-announcement subperiod and 0 elsewhere. The coefficient on the post-announcement dummy is large in absolute value and at t=-3.40 is significant at all reasonable significance levels. This implies that the liberalization announcement did have a strong negative effect on the stock returns in Korea. It should be noted, however, that this may be the result of the bear market in Korea in 1991 and early 1992 and not entirely due to market liberalization. In fact, the average OC return in the post-announcement subperiod is -0.29.

The variance equation is specified as a GARCH(1,1) model with a post-announcement (DL_t) and Saturday DOW (DS_t) dummy variables. The second and third coefficients are ARCH and GARCH parameters, respectively, and are significant at any reasonable levels. The size and statistical significance of the Saturday DOW variable is interesting. The negative sign indicates that the variance of the half day trading is about two-thirds less than a normal trading day, a result that is supportive of the hypothesis that trading activity causes volatility [see French (1985)].

Most interesting for our purposes is the coefficient on DL_t. It is small and statistically not different from zero. The liberalization announcement seems to have caused no structural change in market volatility. This implies that the increase in volatility in our post-announcement subperiod must come from other sources, such as spillovers from the two major markets. In our next model, where we include the spillovers, we drop this term from the estimated model.

We carried out various specification and hypothesis tests on Model 1. These Likelihood Ratio (LR) statistics are also reported in Table 3. Our test for GARCH(1,1) specification with the null hypothesis of no GARCH effect is rejected at all reasonable levels at LR(3) of 250.2. The test for constant mean and constant variance is also rejected at LR(7) of 266.0. Most importantly, the LR(2) test with the null hypothesis that the post-liberalization dummy variables are zero is easily rejected. Since our DL_t coefficient in the variance equation was found to be insignificant using the t statistic, this rejection must be due to the DL_t coefficient in the mean equation.

In addition to the specification tests reported above, goodness of fit is also checked through an analysis of the standardized residuals ($\hat{\varepsilon}_t/\hat{h}_t^{1/2}$) and squared standardized residuals ($\hat{\varepsilon}_t/\hat{h}_t$) for higher order serial correlation. If the GARCH effect was modelled properly, these residuals should lack any higher order serial correlation. As shown in Table 3, both residuals and squared residuals lack serial correlation as indicated by the Q statistics which are all significant at 5% levels. Also, a high t statistic for the $1/\nu$ term indicates the appropriateness of using the conditional t distribution. This is further backed up by our calculation of the implied kurtosis, $3(\hat{\nu}-2)(\hat{\nu}-4)^{-1}=4.38$, which is very close to its sample analogue for $\hat{\varepsilon}_t^4/\hat{h}_t^2$, $K=4.30$.

3.2. A GARCH Model of KOSPI OC Returns with Volatility Spillovers

We now set up and estimate a GARCH model that can be used to test the increase in the volatility spillovers from New York and Tokyo to Korea following the market liberalization.

To capture the possible direct effect of foreign market volatility on the own market returns, we change our stock return specification for the mean equation by separating the mean equation time varying risk term into two parts: a risk measure that is derived from the own market's volatility, $h_t^{1/2}$, and four risk measures that capture the direct spillover effects from influential foreign markets. We denote volatility spillover from New York by XS_t and the spillover from Tokyo by XN_t. The two post-announcement spillovers are denoted $XS_t \cdot DL_t$ and $XN_t \cdot DL_t$. This approach differs from that taken by Hamao et al. (1990), where the foreign market's returns rather than volatility are modeled into the mean equation. Although the levels of foreign stock market returns may be important determinants of own market returns, actual information is revealed in the movement of the returns, as reflected in the volatility of the returns.[5] As such, the positive correlation between the two market's returns may be a result of the markets reacting similarly to some new information that would be revealed and transmitted through return volatility, rather than the level of returns in one market having a causal effect on a neighboring market.

To form the spillover variables, we use once lagged squared residuals from the GARCH model, similar to Model 1, estimated for NIKKEI and S&P returns. Thus, to generate XS_t, we use the once lagged squared residuals from GARCH estimation of Model 1 for S&P returns (without the liberalization dummy variables but with a Wednesday DOW dummy in the variance equation). To obtain XN_t, we use once lagged squared residuals from GARCH estimation of Model 1 for NIKKEI returns (with volatility spillovers from New York in the mean and variance equations - an effect also found by Hamao et al. (1990) - as well as a Monday DOW term in the mean and a Tuesday DOW term in the variance equations).[6] Our estimates of NIKKEI and S&P returns are similar to those of other researchers, so we omit the results here. In deriving these spillover variables, the presence of non-synchronous holidays and the half day Saturday trading in KSE and in TSE up to 1987 presents data problems in forming the spillover variables. In this regard, we adopt the "volatility surprise" method used by Hamao et al. (1990) where the most recent available volatility spillover observation is

[5]See Kyle (1985) and King and Wadhwani (1990).

[6]We also tried using lagged estimated conditional variances for volatility spillover variables. As in Hamao et al. (1990), the results were similar to those obtained using squared residuals.

used. Finally, we impose the small country assumption for Korea. This allows us to restrict attention to cases in which spillovers come only from New York and Tokyo to Korea. This assumption also implies the absence of the international arbitrage effects discussed by Hamao et al. (1990) that would allow us to disregard the overlapping trading hours between the Tokyo and Korean markets.

Volatility spillover terms are also included in the variance equation as independent variables. This allows estimation of a direct relationship between spillovers from the two major markets and the volatility of the Korean market. Thus, the separation of the risk term in the mean equation, and the inclusion of them as explanatory variables in the variance equation, allows us to fully examine the effects of volatility spillovers on the mean and variance of KOSPI returns.

The use of post-announcement spillover variables, $XS_t \cdot DL_t$ and $XN_t \cdot DL_t$ allows a direct estimation of the *increase* in volatility spillovers after the liberalization announcement. This method of conducting our event study has two advantages. First, it allows estimation of one comprehensive model to compare the two subperiods, helping to circumvent the seemingly constant conditional variances in short time horizons, which may be a result of high persistence of GARCH effects. Second, the use of dummy variables allows application of simpler and easier hypothesis tests.

Thus, our GARCH model of volatility spillovers, Model 2, is set up as follows:

$$\begin{aligned} R_t^i &= \gamma_0 + \gamma_1 h_t^{1/2} + \gamma_2 DL_t + \gamma_3 XN_t + \gamma_4 XN_t \cdot DL_t + \gamma_5 XS_t + \gamma_6 XS_t \cdot DL_t + \theta\varepsilon_{t-1} + \varepsilon_t \,, \\ h_t &= \alpha_0 + \alpha_1 \varepsilon_{t-1}^2 + \beta_1 h_{t-1} + \delta_1 DS_t + \delta_2 XN_t + \delta_3 XN_t \cdot DL_t + \delta_4 XS_t + \delta_5 XS_t \cdot DL_t \,. \end{aligned} \tag{3.2}$$

The estimation results are reported in Panel B of Table 3.[7] It is interesting to note that the estimates of the mean equation spillovers for the full sample period are negative and similar in size for spillovers from New York and Tokyo. The positive coefficient estimates for the post-announcement subperiod supports our hypothesis of increased volatility spillovers. However, the t statistics are very small and, as reported in Table 3, an LR(2) test (LR=4.48) of the null hypothesis

[7]We also estimated all models with the interaction terms between NIKKEI and S&P spillovers, *ie.*, with $XS_t \cdot XN_t$ and $XS_t \cdot XN_t \cdot DL_t$. The coefficients were not significant and their inclusion did not improve the likelihood function. We exclude these terms from our reported results.

of zero *increase* in the post-announcement mean equation spillovers cannot be rejected even at 10%. Although this may be a result of our small post-announcement sample size, it seems that volatility spillovers onto the mean OC KOSPI returns did not increase with the liberalization announcement.[8] However, the highly significant DL_t coefficient indicates that the announcement did affect a strong structural change in the mean KOSPI returns.

The results for volatility spillovers in the variance equation is similarly inconclusive. The coefficient estimates for $XS_t \cdot DL_t$ and $XN_t \cdot DL_t$ are positive, indicating a rise in volatility spillovers since the liberalization announcement. It is also clear that the volatility spillovers from Tokyo are stronger than the spillovers from New York. But again, the t ratios are both small and the LR(2) test (LR=3.65) shows that the hypothesis of no increase in volatility spillovers following the liberalization announcement cannot be rejected. However, the LR(8) test (LR=20.68) for no volatility spillovers at all indicates that the spillovers are jointly significant, implying that while spillover effects from Tokyo and New York are important in the determination OC KOSPI returns, spillover effects have not increased since liberalization.

3.3. Estimating Models of CO KOSPI Returns

Given our empirical finding that volatility spillover effects have not increased with the liberalization announcement for the OC returns, it is important to see if such spillovers have increased for the CO returns. If so, this would imply that all new and pertinent information would reveal itself in opening KOSPI prices, thereby having a stronger effect on the CO returns, particularly following the liberalization announcement. To check this hypothesis, we estimated Model 1 for CO returns. In the results reported in Panel A of Table 4, we see that the MA(1) coefficient is large and highly significant, unlike the estimates for the OC returns. More importantly, the DL_t coefficient is small and not different from zero. In the variance equation, on the other hand, the DL_t coefficient is highly

[8]Our spillover variables may suffer from the "generated regressor bias" described by Pagan and Ullah (1988), which would impart bias in our LR statistics. The complexity of the model make correction of this problem difficult. But since most of our LR statistics are very large, our results are still suggestive even with this problem.

significant. This is the exact opposite of that found for OC returns. The GARCH and ARCH coefficients are also different from those for the OC returns, reflecting the different basic statistics shown for the two returns. In addition, $1/\nu$ is larger, indicating that the more leptokurtic CO returns are being modelled properly. The implied kurtosis is 27.99, close to its sample analogue of 29.15. The GARCH specification is appropriate, as the LR(3) statistic for no GARCH effects is strongly rejected (*cf.* Table 4). Finally, the LR(2) statistic for the significance of the two DL_t coefficients is significant at 5%. Thus, we conclude that the liberalization announcement caused a structural change in the KOSPI CO returns, particularly in its variance.

To model the volatility spillovers on KOSPI CO returns, we estimate Model 2 and report the results in Panel B of Table 4.[9] Our GARCH estimation achieves a good fit of the more leptokurtic CO return variances as shown by the sample and implied kurtosis and large $1/\nu$ estimates.

These results are quite unlike those of OC returns in that the spillovers from New York in the variance is now as large as that from Tokyo and highly significant. This confirms our view that all New York information will show up in the KOSPI opening prices. Also, the change in spillovers following the liberalization announcement is quite significant. LR test results show that, unlike results for the OC returns, the spillover terms in CO returns are all highly significant, except for the post-liberalization spillovers in the mean equation. For example, the LR(2) test statistic with the null hypothesis of zero post-announcement increase in spillovers in the variance equation is 17.94, significant at all reasonable significance levels. However, the LR(2) test statistic for the null of zero increase in the mean equation is only 1.98, which cannot reject. The LR(4) test statistic for joint significance of mean and variance equation spillovers, however, is 16.21, which is significant at 1%. As was true of the OC returns, there are significant volatility spillovers to the levels and the volatility of CO returns. This is indicated by the large LR(8) statistics (=59.50) for a test of the null hypothesis of no volatility spillovers at all. These results indicate that spillovers from the two major

[9]Estimation of Model 2 with DL_t in the variance equation did not yield significantly different results, as the t ratio on DL_t is only 0.65. For comparison sake we drop DL_t from the variance equation and report Model 2 as estimated for OC returns.

markets affect CO returns much more strongly than OC returns. They also imply that such spillovers have increased significantly since the liberalization announcement. Finally, the results support our hypothesis that volatility spillovers affect opening prices much more than closing prices, as predicted by the efficient market hypothesis.

IV. CONCLUSION

Our results confirm the importance of volatility spillovers from the major markets in the determination of stock returns and volatilities in our emerging market, Korea. We have also shown that these spillovers increased after the market liberalization for the CO return measure. For OC returns the increases seem negligible. This indicates that information about the behavior of stock returns in the major foreign markets has become more important for opening stock prices since the liberalization. In other words, what used to be information that was relatively unimportant has become important since the market liberalization. This result further backs up the efficient use of information in the market through arbitrage and hedging by both Korean and foreign investors.

These results have also raised some unresolved questions. How large and how fast are the spillovers absorbed into stock prices? Are there any economic reasons why the Tokyo market is more influential than the New York market in the determination of Korean stock return volatility besides the geographic proximity and perhaps closer economic ties between Korea and Japan? Just when and how does information that was previously unimportant become important? Finally, exactly when do the effects of market liberalization appear in the emerging market? These unresolved issues are the subject of ongoing research.

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Table 1

Data Summary Statistics[a]

KOREAN DATA (KOSPI)

	Open-to-Close Return			Close-to-Open Return		
	Full Data	Pre Announcement	Post Announcement	Full Data	Pre Announcement	Post Announc't
Mean	0.051[b]	0.079[b]	-0.290[b]	0.029[b]	0.018[b]	0.202[b]
Variance	1.279	1.252	1.409	0.249	0.212	0.478
Skewness	0.167	0.177	0.308	0.811	1.150	0.844
Kurtosis	1.240	1.324	0.310	15.010	18.002	2.143

NEW YORK DATA (S&P 500)

	Open-to-Close Return			Close-to-Open Return		
	Full Data	Pre Announcement	Post Announcement	Full Data	Pre Announcement	Post Announc't
Mean	0.058[b]	0.061	0.040	-0.009	-0.011	0.006
Variance	1.405	1.498	0.572	0.053	0.057	0.011
Skewness	-4.631	-4.675	-0.420	-17.499	-17.058	6.690
Kurtosis	90.291	88.151	5.103	479.666	446.019	61.719

TOKYO DATA (NIKKEI 225)

	Open-to-Close Return			Close-to-Open Return		
	Full Data	Pre Announcement	Post Announcement	Full Data	Pre Announcement	Post Announc't
Mean	-0.009	0.003	-0.144	0.036[b]	0.039[b]	0.000
Variance	1.543	1.497	1.982	0.031	0.030	0.039
Skewness	-0.757	-0.850	0.462	0.182	0.216	0.111
Kurtosis	26.464	30.651	1.015	6.084	7.136	-0.620

[a] The returns for days when markets were closed were deleted. Thus the return for the day following a holiday is a return for an additional day. Full and Post-liberalization Announcement periods for this and other tables ranges from October 2, 1985 to March 23, 1992 and September 1, 1991 to March 23, 1992, respectively.

[b] These means are significantly different from 0.00 at 5% significance level.

Table 2

Cross Correlations of KOSPI Stock Returns

	Returns in Levels				Squares of Returns			
	Full Sample		Post-Announcement		Full Sample		Post-Announcement	
	OC	CO	OC	CO	OC	CO	OC	CO
Nikkei OC	0.067	-	0.037	-	0.082	-	0.135	-
Nikkei OC(-1)[a]	-0.023	0.100	-0.202	0.093	0.033	0.051	0.245	0.139
Nikkei CO	0.011	0.082	0.122	-0.147	-0.004	-0.012	0.004	-0.067
S&P OC(-1)	0.033	0.105	-0.043	0.109	0.056	-0.001	-0.047	0.104

[a] (-1) stands for the lag of one for the series.

Table 3

GARCH Estimation of OC KOSPI Returns and Volatility Spillovers to Korea

PANEL A: Model 1 (Log-Likelihood = -2768.37)

$$R_t^{oc} = \gamma_0 + \gamma_1 h_t^{1/2} + \gamma_2 DL_t + \theta\,\varepsilon_{t-1} + \varepsilon_t\,,$$
$$h_t = \alpha_0 + \alpha_1 \varepsilon_{t-1}^2 + \beta_1 h_{t-1} + \delta_1 DL_t + \delta_2 DS_t\,.$$

γ_0	γ_1	γ_2	θ		
-0.160 (-1.833)	0.201 (2.307)	-0.349 (-3.403)	0.026 (1.053)		
α_0	α_1	β_1	δ_1	δ_2	$1/\nu$
0.153 (5.264)	0.158 (6.299)	0.787 (26.787)	0.021 (0.577)	-0.421 (-3.909)	0.120 (199.656)

Goodness of Fit Statistics	Value	Null Hypothesis	Value
Sample Kurtosis	4.151	LR(3) $\gamma_2=\alpha_1=\beta_1=0$	250.25*
Implied Kurtosis	4.387	LR(2) $\gamma_2=\delta_1=0$	13.82*
Q(20)	24.707	LR(7) $\gamma_1=\gamma_2=\theta=\alpha_1=\beta_1=\delta_1=\delta_2=0$	266.00*
$Q^2(20)$	18.491		

PANEL B : Model 2 (Log-Likelihood = -2758.03)

$$R_t^{oc} = \gamma_0 + \gamma_1 h_t^{1/2} + \gamma_2 DL_t + \gamma_3 XN_t + \gamma_4 XN_t\cdot DL_t + \gamma_5 XS_t + \gamma_6 XS_t\cdot DL_t + \theta\,\varepsilon_{t-1} + \varepsilon_t\,,$$
$$h_t = \alpha_0 + \alpha_1 \varepsilon_{t-1}^2 + \beta_1 h_{t-1} + \delta_1 DS_t + \delta_2 XN_t + \delta_3 XN_t\cdot DL_t + \delta_4 XS_t + \delta_5 XS_t\cdot DL_t\,.$$

γ_0	γ_1	γ_2	γ_3	γ_4	γ_5	γ_6	θ	
-0.112 (-1.294)	0.202 (1.900)	-0.464 (-3.691)	-0.005 (-0.579)	0.055 (1.550)	-0.005 (-0.685)	0.031 (0.349)	0.031 (1.245)	
α_0	α_1	β_1	δ_1	δ_2	δ_3	δ_4	δ_5	$1/\nu$
0.135 (4.795)	0.151 (5.897)	0.786 (25.558)	-0.382 (-3.620)	0.021 (2.462)	0.008 (0.287)	0.000 (0.002)	0.008 (0.305)	0.110 (15.099)

Goodness of Fit Statistics	Value	Null Hypothesis	Value
Sample Kurtosis	4.056	LR(2) $\gamma_4=\gamma_6=0$	4.48
Implied Kurtosis	4.181	LR(2) $\delta_3=\delta_5=0$	3.65
Q(20)	25.701	LR(4) $\gamma_4=\gamma_6=\delta_3=\delta_5=0$	5.74
$Q^2(20)$	17.001	LR(8) $\gamma_3=\gamma_4=\gamma_5=\gamma_6=\delta_2=\delta_3=\delta_4=\delta_5=0$	20.68*

* Significant at 1% significance level.
Note: figures in parenthesis are t statistics.

Table 4

GARCH Estimation of CO KOSPI Returns and Volatility Spillovers to Korea

PANEL A: Model 1 (Log-Likelihood = -250.25)

$$R_t^{co} = \gamma_0 + \gamma_1 h_t^{1/2} + \gamma_2 DL_t + \theta\varepsilon_{t-1} + \varepsilon_t ,$$
$$h_t = \alpha_0 + \alpha_1 \varepsilon_{t-1}^2 + \beta_1 h_{t-1} + \delta_1 DL_t + \delta_2 DS_t .$$

γ_0	γ_1	γ_2	θ		
-0.029	0.104	0.030	0.155		
(-0.957)	(1.376)	(0.491)	(4.118)		
α_0	α_1	β_1	δ_1	δ_2	$1/\nu$
0.058	0.582	0.417	0.079	-0.054	0.234
(7.562)	(5.871)	(7.440)	(2.124)	(-4.654)	(9.538)

Goodness of Fit Statistics	Value	Null Hypothesis	Value
Sample Kurtosis	29.155	LR(3) $\gamma_2=\alpha_1=\beta_1=0$	293.85*
Implied Kurtosis	27.987	LR(2) $\gamma_2=\delta_1=0$	50.87*
Q(20)	78.210	LR(7) $\gamma_1=\gamma_2=\theta=\alpha_1=\beta_1=\delta_1=\delta_2=0$	582.17*
$Q^2(20)$	6.579		

PANEL B : Model 2 (Log-Likelihood = -190.74)

$$R_t^{co} = \gamma_0 + \gamma_1 h_t^{1/2} + \gamma_2 DL_t + \gamma_3 XN_t + \gamma_4 XN_t \cdot DL_t + \gamma_5 XS_t + \gamma_6 XS_t \cdot DL_t + \theta\varepsilon_{t-1} + \varepsilon_t ,$$
$$h_t = \alpha_0 + \alpha_1 \varepsilon_{t-1}^2 + \beta_1 h_{t-1} + \delta_1 DS_t + \delta_2 XN_t + \delta_3 XN_t \cdot DL_t + \delta_4 XS_t + \delta_5 XS_t \cdot DL_t .$$

γ_0	γ_1	γ_2	γ_3	γ_4	γ_5	γ_6	θ	
-0.018	0.085	0.039	0.003	0.002	0.007	0.009	0.164	
(-0.775)	(1.244)	(0.727)	(0.567)	(0.098)	(3.318)	(0.152)	(4.513)	
α_0	α_1	β_1	δ_1	δ_2	δ_3	δ_4	δ_5	$1/\nu$
0.034	0.578	0.421	-0.043	0.009	0.021	0.008	0.033	0.238
(6.647)	(6.383)	(8.483)	(-5.053)	(2.797)	(0.982)	(2.796)	(0.504)	(10.070)

Goodness of Fit Statistics	Value	Null Hypothesis	Value
Sample Kurtosis	34.377	LR(2) $\gamma_4=\gamma_6=0$	1.98
Implied Kurtosis	34.259	LR(2) $\delta_3=\delta_5=0$	17.94*
Q(20)	73.910*	LR(4) $\gamma_4=\gamma_6=\delta_3=\delta_5=0$	16.21*
$Q^2(20)$	2.229	LR(8) $\gamma_3=\gamma_4=\gamma_5=\gamma_6=\delta_2=\delta_3=\delta_4=\delta_5=0$	59.50*

* Significant at 1% significance level.
Note: figures in parenthesis are t statistics.

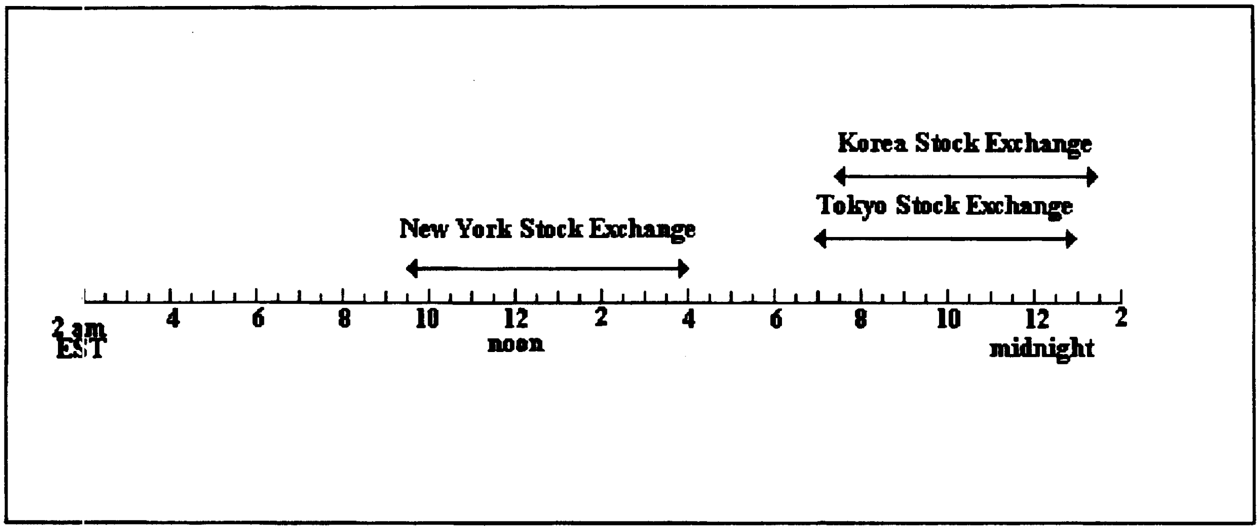

Figure 1
Stock Exchange Operating Hours

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