Economic News and Stock Market Linkages: Evidence from the U.S., U.K., and Japan*

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Abstract

This paper examines the cross-market equity return and volatility linkages for the U.S., U.K., and Japan. We investigate the extent to which these linkages can be explained by macroeconomic news announcements in the three countries, including money supply, industrial production, price inflation, unemployment rate, and trade deficit during the 1985-1996 period. Our results indicate that these news announcements account for little of the direct intermarket return spillovers. However, we find that these announcements can affect the size of intermarket return spillovers. We also find these news announcements appear to explain, at least partially, the volatility spillovers among the three markets.

1 Introduction

Considerable empirical evidence has now accumulated that return correlations

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among markets are rising, returns in one market influence returns in subsequent markets, and that market volatility spills from one market to another. Further, these spillovers appear to have pronounced asymmetries; e. g., returns in New York have a much larger impact on Tokyo returns than the reverse. The nature and source of these co-movements remain an interesting issue.

In this paper, we investigate whether economic news announcements are a potentially important source of the co-movements. Specifically, we model the impact of macroeconomic news announcements on daily close-to-open and open-to-close mean returns, volatility, and covariances among the U.S., Japan, and the U.K. Our contribution to the literature has several facets. First, relatively little research has measured the impact of explicit economic news from one country on capital markets in another nation.[1] We study the reactions of stock index returns in the U.S., U.K., and Japan to a handful of potentially significant economic announcements made in each country. Second, our extensive data set covers both intraday and overnight return series for the three countries from 1985 to 1996. And, importantly, we replace the "stale" official opening prices in the U.S. and Japanese markets with the 10:00 prices in both markets. Third, our study sheds some new light on the important question—to what extent can the observed market linkages be explained by "fundamental" economic news? If public economic news announcements play little or no role in explaining market linkages, there is room to consider alternative explanations for the market spillovers such as the contagion effect documented in King and Wadhwani (1990).

Our empirical model of returns and news allows macroeconomic news to affect returns and to affect the spillover from the preceding market return. This nonlinear approach is very much in the spirit of the Karolyi-Stulz (1996) model. Unlike their setup, however, we address only the nature of the return-generating model, not a joint hypothesis that involves a specific asset-pricing model. Karolyi and Stulz (1996) find that co-movements between the U.S. and Japan are not systematically different on days when U.S. macroeconomic news is announced. Controlling for interest rate shocks and industry effects, news also has little impact on market covariances. They show that co-movements are particularly strong when contemporaneous absolute returns on U.S. and Japanese stock indices are large.

We examine what role economic news announcements play in return spillovers between national stock markets. Our results show that each domestic market's open-to-close and close-to-open returns are significantly linked to the open-to-close returns of the two previous foreign markets, even after controlling for the effect of economic news announcements. Furthermore, the *asymmetric* return spillover patterns among the three national stock markets persist in our study. Specifically, in the open-to-close return series, the impact of the S&P500 on the

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[1] Lin (1994) investigates the effects of U.S. economic news announcements on returns and trading volume in the Japanese stock market for the period October 1982 through December 1991. She uses a larger set of U.S. news announcements than we do, but does not control for the impact of Japanese economic news or model volatility as we do in this paper.

Nikkei225 returns is *five* times the impact of the Nikkei225 on the S&P500. Likewise, the impact of the FTSE100 on the Nikkei225 is about *three* times the impact of the Nikkei225 on the FTSE100. Finally, the impact of the FTSE100 on the S&P500 is *ten* times the size of the S&P500's effect on the FTSE100 index return.

In addition, we find little evidence that economic news announcements in the three countries have systematic, *independent* effects on the return process in any of the three national stock markets. In this respect, our results are consistent with generally weak results noted in earlier studies of the effects of economic news announcements (e.g., Karolyi and Stulz 1996). Nonetheless, our results also indicate that some individual, and mostly domestic, economic news announcements significantly affect the *size* of return spillovers between markets. For example, U.S. money supply news announcements significantly increase the size of the return spillover from the Nikkei225 into S&P500, while significantly reduce the size of the return spillover from the FTSE100. Thus, our results suggest that while economic news announcements by themselves do not *directly* affect the return process in each national stock market, they play an important role in explaining the variation in the return spillovers between markets.

A second part of our work addresses the role of economic news announcements in explaining volatility spillovers from market to market. Following the logic in Ross (1989), we enter news announcements directly into the conditional variance function for each country. We investigate these effects using a well-known asymmetric conditional volatility model proposed by Glosten, Jagannathan, and Runkle (1993) to capture the asymmetry effect in volatility spillovers between markets.

We first treat all three countries' economic news announcements as dummy variables, which shifts the return volatility process of each national stock market. In this setup, almost all squared return surprises from the previous markets still exert statistically significant effects, except that previous Nikkei225 return volatility does not affect current S&P500 volatility. This result is consistent with the asymmetric volatility spillover effect between the U.S. and Japanese markets noted in earlier studies (e.g., Hamao et al. 1990). But, the evidence is different when we allow these volatility spillovers to have a different impact depending on whether the return surprise is good or bad news. In this case, we find that open-to-close Nikkei225 volatility raises open-to-close S&P500 volatility, particularly when there is bad news.

Surprisingly, return volatility spillover effects between markets are significantly affected in an empirical model where specific domestic and foreign economic news announcements are incorporated in each market's return volatility process. Moreover, some of these news announcements have significantly positive impacts on the return volatility of domestic or foreign markets. For example, using the open-to-close sample, Japanese consumer price inflation announcements raise FTSE100 return volatility; Japanese money supply announcements raise Nikkei225 return volatility; and U.S. money supply announcements raise

Nikkei225 and S&P500 return volatility, while U.S. industrial production announcements raise S&P500 return volatility. Taken together, these results suggest that economic news announcements may *directly* account for the observed volatility spillovers between markets.

In a world without economically significant investment restrictions, there are several potential explanations for market interdependence. First, if domestic stock prices are driven primarily by domestic economic factors, domestic economic news may lead to portfolio rebalancing by internationally diversified investors. Depending on the volume of trade on equity markets worldwide and on elasticities, price changes may accompany this rebalancing. Second, firms may have cash flows from operations in many countries and news about determinants of the size and timing of those cash flows (i.e., economic news from other countries) may lead to a change in the firm's share price. If there are many firms with shares listed on an exchange with significant cash flows from another country, economic news from the foreign country may have important share price effects in the home equity market.[2]

If there is a global factor in equity pricing, investors may follow price changes in other markets because these price movements reveal information about the global pricing factor. In this way, information revealed during trading in Tokyo may affect London price movements and then subsequently, New York price changes. For example, in studying the predictability of monthly excess returns in the U.S. and Japanese stock markets, Campbell and Hamao (1992) find that similar variables, such as the dividend-price ratio and interest rate variables, help to forecast excess returns in each country. Correlations among fundamentals seem to imply co-movements among equity prices in the two countries. As several authors have noted, there are difficulties with the argument for co-movement based on macroeconomic fundamentals. King, Sentana, and Wadhwani (1994) report that little of the covariance dynamics among 16 major equity markets can be explained with macroeconomic variables. An alternative hypothesis, proposed by King and Wadhwani (1990), is that market contagion can spread across markets producing evidence of changing interdependence between stock markets around the world. This hypothesis implies that macroeconomic variables will be largely unable to explain the co-movements among markets.

There is an interesting literature that examines share price volatility spillovers in major capital markets, especially since the October 1987 downturn.[3] Engle, Ito, and Lin (1990) tested two alternative explanations of volatility spillovers. The heat wave hypothesis holds that higher volatility is a local phenomenon and does not spread from market to market in any sequential fashion. By contrast, the meteor shower hypothesis posits that higher volatility spills from one market to the next over the course of the global trading day. In their study of the Tokyo-

[2] This seems particularly likely for Japanese firms which export aggressively to world markets, the most important of which is the US market.

[3] Roll (1988) discusses the transmission of share price changes across capital markets at the Crash of October 1987.

London-New York yen/dollar markets, the evidence fairly convincingly favors the meteor shower hypothesis. Hamao, Masulis, and Ng (1990) examine the stock volatility spillovers for the three countries' equity markets and find significant interdependence among these markets.[4]

One of the more remarkable findings remains the documentation of significant asymmetries in the mean and volatility functions. For example, Campbell and Hamao (1992) discover an asymmetry in the return generating processes for the U.S. and Japanese equity markets. They report that "Japanese variables do not help to forecast U.S. excess stock returns, but U.S. variables do help to forecast Japanese excess stock returns in 1980's." Becker, Finnerty, and Gupta (1990) find that U.S. equity market performance strongly influences subsequent open-to-close equity returns for the Nikkei index but the price change in Tokyo has only a marginal impact on subsequent NYSE price changes. They note that transaction costs in the Tokyo market make it very unlikely that the NYSE-Nikkei predictability will result in excess returns. Neumark, Tinsley, and Tosini (1991) demonstrate that transaction costs make the foreign market response to domestic market news more likely when domestic market volatility is high. They conclude that asymmetries and temporal instability in cross-market return correlations are consistent with rational investor behavior in markets that display time-varying volatility and non-zero transaction costs. Lin, Engle, and Ito (1994) show the contemporaneous correlations of stock returns across the U.S. and Japanese stock markets increase significantly during volatility periods thus confirming the Neumark et al. (1991) analysis. They also report evidence of a structural change in the effect of U.S. returns on Tokyo stock market returns in October 1987.

Interestingly, much as Campbell and Hamao (1992) found asymmetries in the conditional mean dependence, there is evidence the conditional volatility functions display distinct asymmetries. Eun and Shim (1989) and Hamao et al. (1990) both discover asymmetries in volatility spillovers. Both papers report evidence that the innovations in the U.S. are rapidly transmitted to other markets, but not in the other direction. Despite the apparent pervasiveness of these co-movements, relatively little is understood about potential economic foundations for these observed connections among capital markets. Indeed, Engle, Ito, and Lin (1990) close their study of volatility spillovers in the world-wide yen-dollar exchange market by posing the question of whether their findings are due to "correlated fundamental news or a failure of strong-form market efficiency" (p. 541). Our findings in this paper shed some light on the understanding of the effects of macroeconomic news announcements on the return and volatility spillovers among national stock markets.

The plan of the rest of this paper is as follows. In the next section, we describe

[4] A higher degree of interdependence or integration among world capital markets in these recent studies may reflect recent elimination of some investment barriers, such as restrictions on cross-border capital flows. In fact, Gultekin, Gultekin and Penati (1989) find a higher degree of integration between the US and Japanese stock markets after the Japanese government eliminated its capital controls in late 1980.

the conditional mean, volatility, and covariance models used in our empirical investigation. Section 3 contains a description of our data and sample methods and provides details of our econometric methods. Section 4 reports our empirical results. Section 5 concludes this paper.

2 Modeling market linkages, news, and volatility

2.1 News and conditional means

Our first return-generating model isolates the impact on stock market index returns of domestic and foreign economic news and other factors:

$$R_{d,t} = \beta \cdot X_t + \gamma \cdot DN_t + \delta \cdot FN_t + \alpha \cdot R_{f,t-1} + \varepsilon_{d,t} \qquad (1)$$

where $R_{d,t}$ is the domestic stock index return at time t, $R_{f,t-1}$ is the foreign stock index return at time t-1; X_t is a vector of return predictors that includes day-of-the-week dummy variables and, potentially, other variables which might capture predictable variation in stock returns; DN_t is a vector of unexpected percentage changes in domestic economic aggregates (e.g. money stock, industrial production, etc.) at time t; and FN_t is a vector of unexpected percentage changes in foreign economic aggregates (e.g. money stock, industrial production, etc.) at time t. (We discuss specific variable definitions in Section 4 along with pertinent measurement details.) In addition, we indicate parameter vectors by β, γ, δ, and α, and the error term is $\varepsilon_{d,t}$.

To model the interdependence of returns, we add the previous market return to the return generating model for each market. The existing literature focuses on the size of α for each market, the stability of these estimates, and whether these estimates vary systematically with market conditions such as volatility. Since none of these studies account directly for domestic and international economic news, our first question focuses on whether the findings reported in earlier papers are affected by explicitly including such measures in our return models. Since volatility is related to information flow (see Ross (1989)), our model explicitly accounts for a subset of potentially important economic information flows that may raise volatility and that may have caused the appearance of instability in other models that included only the lagged foreign market return.[5]

At the same time, we can also learn more about the connections between national equity markets (i.e., the U.S., U.K., and Japan). The interdependent market hypothesis implies that each stock market reacts to both domestic and international economic news. If we find significant cross-country responses to economic news in the respective equity markets, this would constitute potentially

[5] Other researchers have proposed that the variation in cross-correlations may be due to temporal variation in the heterogeneity of trader priors or noise trading. We do not propose to distinguish between these two alternatives; rather we seek to determine the extent to which identifiable news announcements and measurable variation in market conditions can explain the observed market return cross-correlations.

important new evidence of interdependence among markets.

As we have noted in our discussion, earlier studies have found that α varies quite significantly across markets. One potential explanation for this result is an asymmetric (i.e., lead-lag) response to one or more common state variables. As a common state variable changes, the stock return response in one country leads the other country's stock market response. To the extent that economic news, for example, about the money supply or industrial production, proxies for movements in the (unobservable) underlying common state variables, asymmetric stock return responses to economic news announcements may be the explanation for observed asymmetries among world equity markets. Later, we report some evidence based on model (1) for whether the observed asymmetry between U.S., U.K., and Japanese stock markets is related to a lead-lag pattern in the stock market responses to a set of economic news announcements occurring in each country. If this hypothesis is correct, accounting for news directly should narrow the differences in estimated α across markets.

We also investigate a related hypothesis that market interdependence (measured unconditionally by α) varies systematically with economic news. Here, the idea is that market spillovers are systematically different on days when there is economic news revealed in either the domestic or foreign market. To assess this idea, we use the following return-generating model:

$$R_{d,t} = \beta \cdot X_t + (\alpha + \varphi \cdot DN_t + \theta \cdot FN_t) \cdot R_{f,t-1} + \varepsilon_{d,t} \qquad (2)$$

where all symbols are defined as before and we indicate new parameter vectors φ and θ. Karolyi and Stulz (1996) examined models of this type, but did not investigate the role of specific macroeconomic news announcements in explaining potential variation in the return spillovers between markets. Finding that φ and/or θ are significantly non-zero establishes evidence that the variation in return spillovers is conditional on economic news. Karolyi and Stulz generally report no evidence of this in their work, but they account for economic news with a single dummy variable that is one on days of U.S. economic news announcements and is zero otherwise. In our empirical work, we measure the surprise component in an array of foreign and domestic news announcements.

2.2 News and conditional volatility models

We hypothesize that the prediction error terms can be characterized as follows:

$$\varepsilon_{i,t} \big| \Omega_{t-1} \sim N(0, h_{i,t}) \qquad i = d, f \qquad (3)$$

where Ω_{t-1} denotes the information set at time t-1. The important practical issue is the specification of the conditional volatility function. It is well-known that stock returns are conditionally heteroskedastic and that modifications of Engle's (1982) Autoregressive Conditional Heteroskedastic (ARCH) model are useful in capturing the temporal dependence in the second moment of stock returns. Bollerslev's

(1986) Generalized ARCH (GARCH) model has been applied to U.S. and foreign index returns with considerable success. The conditional variances (h_t) of a GARCH(1,1) model can be represented as follows

$$h_t = \tau_0 + \tau_1 \cdot h_{t-1} + \tau_2 \cdot \varepsilon_{t-1}^2 \tag{4}$$

where the ε term is the conditional error. In this more general model, current conditional variances depend not only on last period's actual prediction errors (residuals), but also on last period's conditional variances.[6]

There is considerable evidence favoring asymmetric models of conditional volatility in which positive and negative news can generate volatility shocks of different sizes. A common finding in equity market applications (see Engle and Ng, in particular) is that these asymmetry coefficients are negative, meaning that negative news has a bigger impact on conditional volatility than good news. One particularly simple way to capture the asymmetries in conditional volatility is the model proposed by Glosten, et al.[7] Applied to our setting, their model gives the conditional volatility as

$$h_{i,t} = \omega_0 + \omega_1 \cdot h_{i,t-1} + \omega_2 \cdot \varepsilon_{i,t-1}^2 + \omega_3 \cdot \varepsilon_{i,t-1}^{2-} \qquad i = d, f \tag{5}$$

where $\varepsilon_{i,t-1}^{2-}$ are the squared unexpected returns on days when there is bad news. In this model, finding that ω_3 is nonzero provides evidence of asymmetries in the volatility consequences of bad and good news. Otherwise, this model is essentially a GARCH(1,1) model of conditional volatility.

To investigate potential volatility spillovers across markets, we adapted the volatility spillover model of Engle, Ito, and Lin (1990). In their model, volatility in a domestic market segment depends on squared return surprises in preceding

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[6] We do not investigate GARCH-M models here. Since at least the work of Sharpe, Lintner, and Mossin, financial economics has posited that expected returns depend on some measure of risk. The GARCH-M, or GARCH in Mean model, is a natural extension of the simple GARCH model in which the current conditional variance also appears in the return equation. See Engle, Lilien, and Robins (1987) for the development of the model. In particular, a GARCH-M model may be written

$$R_t = \mathbf{X} \cdot \beta + \delta \cdot h_t + \varepsilon_t$$

$$h_t = \tau_0 + \tau_1 \cdot h_{t-1} + \tau_2 \cdot \varepsilon_{t-1}^2$$

where $\mathbf{X}$ is a vector of return determinants, β is a coefficient vector, and all other variables are as defined earlier. French, Schwert, and Stambaugh (1987) and Baillie and DeGennaro (1989) have reported some success in modeling stock returns with a structure similar to this. Hamao, Masulis, and Ng (1990) applied this model in their analysis. We experimented with this model extensively and found no evidence the GARCH-in-Mean term contributed any explanatory power to the model. Hamao, et al. also found little evidence favoring the GARCH-in-Mean hypothesis. See also the discussion in Glosten, Jagannathan, and Runkle (1993).

[7] Another approach to modelling these volatility asymmetries is Nelson's (1990) Exponential GARCH (EGARCH) model which, in our application, is given by

$$h_{d,t} = \exp[\varphi_0 + \varphi_1 \cdot |\varepsilon_{d,t-1}|/\sqrt{h_{d,t-1}} + \varphi_2 \cdot \varepsilon_{d,t-1}/\sqrt{h_{d,t-1}} + \varphi_3 \cdot \log(h_{d,t-1})]$$

A thorough discussion and analysis of this approach to modeling conditional volatility is provided by Pagan and Schwert (1990), Bollerslev, Chou, and Kroner (1992), Engle and Ng (1993), and Glosten, Jagannathan, and Runkle (1993). The asymmetries arise when φ_2 or κ_2 are non-zero, since this is equivalent to finding that the impact of bad news on volatility differs from the impact of good news on volatility. See Campbell and Hentschel (1992) for another asymmetric conditional volatility model.

foreign and domestic market segments. Including terms that account for volatility asymmetries, we have

$$h_{d,t} = \omega_0 + \omega_1 \cdot h_{d,t\text{-}1} + \omega_2 \cdot \varepsilon^2_{d,t\text{-}1} + \omega_3 \cdot \varepsilon^2_{d,t\text{-}1} + \omega_4 \cdot \varepsilon^2_{f,t\text{-}1} + \omega_5 \cdot \varepsilon^{2-}_{f,t-1} \quad (6)$$

where all terms are as defined earlier. In the Engle, Ito, and Lin model, ω_3 and ω_5 are both zero, so (6) is a potentially interesting extension of their model.

Another feature of our work is that we estimate the direct impact of news announcements on market volatility. Given Bailey (1988) and Ross (1989), it seems especially reasonable to model the volatility effects of information flows contained in these economic news announcements. Accordingly, we modified (6) to accommodate potential volatility effects of news announcements and to include the effects of interest rate levels ($i_{d,t}$) on conditional volatility:[8]

$$h_{d,t} = \omega_0 + \omega_1 \cdot h_{d,t-1} + \omega_2 \cdot \varepsilon^2_{d\,t-1} + \omega_3 \cdot \varepsilon^{2-}_{d\,t-1} + \omega_4 \cdot \varepsilon^{2-}_{f\,t-1} + \omega_5 \cdot \varepsilon^{2-}_{f\,t-1} + \omega_6 \cdot i_{d,t} + \nu \cdot DN_t + \psi \cdot FN_t + \Theta_d \cdot DOW_t \quad (7)$$

Since Baillie and Bollerslev (1990) and others have reported evidence that volatility differs by day of the week, we also include the DOW dummy variable vector in each of the conditional volatility models.[9]

[8] See Glosten, Jagannathan, and Runkle (1993) for a recent analysis of interest rate effects on conditional volatility.

[9] We worked extensively, but with limited success, on the issue of whether economic news had systematic effects on conditional covariances. The most basic conditional correlation model is given by

$$h_{df,t} = \rho_0 + \rho_1 \cdot h_{df,t\text{-}1} + \rho_2 \cdot \varepsilon^2_{d,t} \cdot \varepsilon^2_{f,t}$$

where $h_{df,t}$ is the conditional covariance and the ρ_i are parameters. In this model, news announcements affect covariances only through the product of the individual volatility terms. There is empirical evidence to suggest that conditional covariances among international equity returns vary systematically through time (see Karolyi and Stulz (1996) for references and discussion). More complicated multivariate versions of the GARCH volatility model provide alternative ways to explore the behavior and economic determinants of conditional covariances. Following the example in Karolyi and Stulz (1996), we might specify such a model as (2), (7), and

$$h'_{dft} = (\rho_3 + \rho_4 \cdot i_{d,t} + \lambda \cdot DN_{d,t} + \chi \cdot FN_{f,t} + \kappa_d \cdot DOW_t)(h_{d,t} \cdot h_{f,t})^{1/2}$$

where h'_{dft} is the conditional correlation between the domestic and foreign market returns at time t, all other terms are as defined earlier, and λ, χ, and κ are the new parameter vectors. This structure for correlation dynamics supports testing for direct economic news effects on correlation movements, aside from the indirect effects of economic news on conditional variances, and hence, on conditional covariances. Specifically, we expect λ and/or χ to be nonzero if economic news affects the conditional correlation between domestic and foreign equity markets.

3 Data and empirical methods

3.1 Data sample

We estimate our empirical models with daily data from the U.S., U.K., and Japan for the sample period January 1, 1985 through December 31, 1996. We formed returns to estimate these models, in part due to the differences in hours of operation for these equity markets. Consistent with the timing of the Engle, Ito, and Lin (1990), we first calculated a sequence of non-overlapping open-to-close returns on the Nikkei225, open-to-2:00 p.m. returns on the FTSE100, and open-to-close returns on the S&P500 index. We also computed close-to-open returns using the same dating conventions. Most of this data is from Datastream, except where noted.

Since there are potential stale price problems with opening values of the U.S. and Japan stock indices, we collected intraday values for these indices and recomputed close-to-open returns with post-opening values of the respective indices. For Japan, we collected 9:15 and 10:00 Nikkei quotes from *Nihon Keizai Shinbun* and computed two sets of close-to-open returns. We found little apparent difference in the return series using the 9:15 quotes, so we relied on the 10:00 data for the Japanese market. For our U.S. return series, we obtained half-hourly quotes for the S&P500 index from Prudential Securities for the 1985 - 1993 period and hourly quotes from Standard and Poor's for the last three years. We calculated close-to-open returns using 10:00 quotes for the S&P500.

We found significant differences in the variances of returns using the later quotes of the respective indices, and in the estimates we computed. (These results are not reported here, but a summary is available upon request.) We take this to be consistent with the warnings from earlier researchers that find stale price problems to be an important issue in empirical work. Following what appears to be common practice, we used opening quotes, not the 10:00 data in computing the open-to-close return series for these markets. Figure 1 shows the timing of the three markets and the corresponding open-to-close and close-to-open returns in each market.

Our U.S. economic news variables include money supply, industrial production, consumer price inflation, wholesale price inflation, and the merchandise trade balance. We acquired the actual announcement values from MMS (Money Market Services) International. In addition to the announced values, we have the median estimates of the announcement values for both variables. These median values are calculated from the MMS International surveys and serve as proxies for the market's expectations. Our U.S. economic news variables are calculated as the difference between the actual value of the economic aggregate less the median expected value. Money supply announcements are made weekly, but the other announcements are made on a monthly basis.[10]

[10] Ederington and Lee (1993) provide a more detailed discussion of economic news announcements and their impacts on the U.S. market.

For Japan, we collected the monthly announced value of Japanese industrial production, money supply, consumer price inflation, and wholesale price inflation by hand from issues of *Nihon Keizai Shinbun*. In addition, the Bank of Japan provided some announcement data for 1995 and 1996. The money supply announcement gives the monthly money stock (M2+CD) estimates. The Japanese money supply announcements are usually made on Friday of the first or second week of the month. Japanese industrial production announcement data are the monthly index of industrial production as reported in *Nihon Keizai Shinbun*. The announcements do not have a fixed release date but are usually made toward the end of the month. The consumer and wholesaler inflation data are monthly announcements made at irregular intervals. Since we do not have survey data to form expectations for these economic aggregates, we obtain proxies for these expectations as one-step ahead forecast values from an appropriate ARIMA model.

Our macroeconomic news announcements for the U.K. are money supply, retail price inflation, industrial production, and the unemployment rate. We collected the announced values for these variables from the *Financial Times*; all are announced at roughly monthly intervals. We used an ARIMA model for each series to calculate expected values of these economic aggregates as one-step ahead forecast values.

Inspection of (11) and (12) shows that our volatility models also incorporate an interest rate term. There are some potential difficulties in finding comparable domestic interest rates for both countries. To address this issue, we use the one-month Eurodollar interest rate for the U.S., the one-month Euroyen interest rate for Japan, and the one-month Eurosterling interest rate for the U.K. This data was collected from Datastream, and consists of observations taken at noon in London. The advantage of the Euro-interest rates is the comparability of the underlying asset, the relative lack of regulation in the market, and ease with which the data can be collected. When matching interest rate data to open-to-close and close-to-open return data, we adjust for timing variations so that interest rate observations never precede the last equity index value used in calculating equity returns.[11]

[11] To implement the multivariate models mentioned in fn. 9, we calculated a set of open-to-close returns on Nikkei225 and S&P500 stock index futures contracts traded on the International Monetary Market (IMM) in Chicago. We acquired the return data in part from Datastream, in part by hand collection, and in part from a private data vendor. The sample extends from August 1990 (when the Nikkei contract began to trade) through the end of December 1996. We also calculated another set of returns to support multivariate modeling of stock index returns. To mitigate some of the difficulties with the nonsynchronous trading problem, we calculate open-to-close and close-to-open returns for all three markets. We focused on six sets of overlapping series: 1) U.S. open-to-close and Japanese close-to-open, 2) U.S. close-to-open and Japanese open-to-close, 3) U.K. open-to-close and Japanese close-to-open, 4) U.K. close-to-open and Japanese open-to-close, 5) U.S. open-to-close and U.K. close-to-open, and 6) U.S. close-to-open and U.K. open-to-close. Given the sparse news announcements in our sample, we find that our data are no match for such highly parameterized multivariate models and the results are inconclusive. These results are available upon request. Chan, Karolyi, and Stulz (1992) use a bivariate GARCH-in-Mean model to mitigate the nonsynchronous trading problem in daily return data. In a novel way, Karoyi and Stulz (1996) use a portfolio of Japanese ADR's trading in the U.S. to avoid the nonsynchronous trading problem completely.

3.2 Econometric methods

We estimated the parameters of our conditional return functions, (1) or (2) from open-to-close and close-to-open data using the least squares estimator. We computed standard errors and test statistics for these models using the Newey-West covariance matrix estimator. Estimating the parameters of the Engle-Ito-Lin volatility model is greatly simplified because of the recursive nature of the model. To implement the model, we retrieve residuals from the conditional mean function model and then estimate the volatility model using the squared residuals from this initial step.

To implement the Engle-Ito-Lin maximum likelihood model, we use the non-linear optimization routine in RATS version 4.10 to estimate the conditional volatility function. The package produces maximum-likelihood-based estimates using the Berndt-Hall-Hall-Hausman numerical method.[12] We used a simplex algorithm to refine starting values before beginning the maximum-likelihood procedure. We encountered some convergence difficulties in our attempts to estimate the parameters of some of the volatility functions. In particular, we found the FTSE100 volatility model and volatility models with the full set of news announcement variables required several dozen sets of starting values before the program would converge. Our attempts to estimate the U.K. volatility model with the full set of news announcement variables were particularly complex; before final convergence, we came quite close to implementing a (manual) grid search method across half a dozen particularly critical parameter estimates.

For each model that we estimated, we also computed the battery of volatility model specification tests proposed by Engle and Ng (1993). Each test has a chi-square distribution with one degree of freedom. The six specification tests are for (the Engle-Ng name is indicated in parentheses):

1) omitted interest rate influences on conditional volatility,
2) an intercept shift in the conditional volatility function for positive errors (positive sign bias test),
3) a linear differential impact of positive errors on conditional volatility (positive size bias test),
4) a nonlinear differential impact of positive errors on conditional volatility,
5) a linear differential impact of negative errors on conditional volatility (negative size bias test), and
6) a nonlinear differential impact of negative errors on conditional volatility.

Tests 4) and 6) are constructed by replacing the linear term in each test with a

[12] A discussion of the econometric properties of estimates from GARCH and related models can be found in Bollerslev, et al. (1992).

square of the term used in the positive size bias test and negative size bias test. This battery of tests provides a way to evaluate the usefulness of the various volatility models in capturing positive and negative asymmetries in the conditional volatility as well as the effectiveness of our interest rate proxy variables. We turn to our estimates and the specification test results next.

4 Empirical results and discussion

4.1 News and conditional mean returns

We turn first to the empirical evidence on the size and extent of return interdependencies among the three markets. Estimates reported in Table 1, Panel A, clearly show that each domestic stock market's open-to-close and close-to-open returns are significantly linked to the open-to-close returns of the two previous foreign markets (with the only exception of Japanese close-to-open returns to previous U.K. market returns.) The table also indicates that the largest cross-market return spillover involves the most recent market segment. For instance, the lagged U.K. return has five times the impact of the lagged Nikkei return in the U.S. stock return equation.

In the open-to-close return series, the estimates reveal a distinct pattern of asymmetric interdependencies between national markets. The impact of the S&P500 on Nikkei returns is *five* times the impact of the Nikkei on the S&P500. Likewise, the impact of the FTSE on the Nikkei is about *three* times the impact of the Nikkei on the FTSE. Finally, the impact of the FTSE on the S&P500 is *ten* times the size of the S&P500's effect on the FTSE index return. Similar results are found in Panel B of this table, where continuous, rather than dummy, variables are used as regressors.

We find little evidence that economic news announcements in the three countries have systematic, independent effects on the return process in any of the countries. This holds both for dummy news variables and continuous measures in Panels A and B, respectively. In this respect, our results are consistent with the results in Karolyi and Stulz (1996). Our results also seem to be consistent with generally weak results noted in earlier studies of the effects of macroeconomic news announcements.

When we analyze whether economic news announcements affect the size of the spillovers using dummy news variables, Table 2's Panel A shows that the spillovers are generally not systematically different on days when news announcements are made. That is, conditioning on whether there are economic news announcements does not appear generally to alter the size of the spillover in a statistically significant manner.

Table 2's Panel B reports estimates of the role of specific news announcements in return spillovers. The most significant finding in Panel B is that domestic news announcements are far more important than foreign news announcements in return spillovers among the three national stock markets. For example, more than two thirds of coefficients on news variables that are significant at the one per cent level are domestic news announcements, which affect foreign market return spillovers into the domestic market. Specifically, Japanese consumer price inflation, producer price inflation, and money supply are important in the Japanese market; UK consumer price inflation is important in the UK market; and US consumer price inflation, money supply, and industrial production are important in the US market.

4.2 News and conditional return volatility

Since information flow is formally and informally connected to volatility in financial theory (e.g. Ross 1989), we also investigated the role of domestic and foreign economic news announcements in conditional volatility modeling. Table 3, Panels A and B, report results of estimating the Engle-Ito-Lin volatility spillover model.[13] First, we address whether there are volatility spillovers among the markets, in addition to the return spillovers just discussed. Panel A shows that for open-to-close returns, all squared return surprises from the previous markets exert statistically significant effects, except that return surprises from Japan exert only a small, statistically insignificant effect on U.S. return volatility. We also note that these volatility spillovers have a different impact depending on whether the return surprise is good or bad news. Seven of nine asymmetry terms are significant. We find that Japanese return volatility now raises U.S. return volatility particularly when there is bad news. The positive volatility spillover effect from S&P500 to Nikkei occurs mainly when there is bad news in the U.S. There is also a significant negative volatility spillover effect from FTSE to Nikkei when there is bad news in the U.K.

The close-to-open return volatility spillovers measured in Panel A look fairly different. There are no spillovers for Japan and the U.S., and only modestly-sized volatility spillovers from the U.S. and Japan into the U.K. In addition, only four of nine asymmetry terms are significant in the close-to-open volatility models, and two of them are significant only at the 10 per cent level. In general, the estimated asymmetry terms are much smaller in the close-to-open return volatility models.

Table 3 also contains estimates of the volatility effects of economic news announcements. Estimates of volatility changes on news announcement days for the open-to-close models (Panel A) show that Nikkei return volatility is lower when there is domestic news, while it is higher when there is news from the U.K. FTSE return volatility is lower when there is news from either U.K. or U.S. markets. U.S. return volatility is lower when there is economic news from either U.K. or Japanese markets. For the close-to-open sample, eight of nine estimates

[13] Some of the volatility model intercept terms are negative indicating the possibility of a negative conditional variance estimate. We computed the sequence of conditional volatilities implied by our model estimates for all markets and found no instance when the conditional volatility was non-positive.

of volatility shifts on news announcement days are significant, and all of them are positive.

Table 3, Panel B, reports estimates of the volatility consequences of individual economic news announcements. Interestingly, eight of 18 significant estimates are positive for the open-to-close models, and eighty percent of the significant estimates in the close-to-open sample are positive. In addition, we find that in the open-to-close sample, 11 of 18 significant estimates are on foreign news announcements. The balance is domestic news. Of the fifteen significant estimates in the close-to-open sample, nine are on foreign news announcements. These results suggest that domestic new announcements may be somewhat less important than foreign news in explaining domestic return volatility. This result is moderately at odds with our finding in Table 2, Panel B, that domestic news is more important than foreign news in explaining domestic returns.

Perhaps more interestingly, return volatility spillover effects between markets that we document in Panel A are often dramatically different in Panel B, where the dummy news variables are replaced with specific economic news announcements. Six asymmetry terms from Panel A are insignificant in Panel B, and seven volatility spillover terms from Panel A are insignificant in Panel B. This result suggests that economic news announcements may have significant *direct* effects on the observed volatility spillovers between markets. It also suggests that researchers need to be very careful about using dummy variables to measure economic news effects. We found a similar disparity of results from dummy and continuous news variables in our return spillover analysis.

Finally, we found our volatility models invariably fared well when the Engle-Ng specification tests were applied.[14] We found only a very small handful of cases where test statistics were significantly different from zero, indicating model misspecification. In these cases, we found that a new set of starting values generally raised the value of the likelihood function and produced insignificant specification test statistics.

5 Conclusion

In this paper, we have studied the role of macroeconomic news announcements in explaining return and volatility spillovers between three major national stock markets. Specifically, we examine the extent to which observed market linkages can be explained by “fundamental” economic news. We have used carefully sequenced data on market returns in Tokyo, London, and New York in conjunction with a series of important macroeconomic news announcements made in each country for the 1985 - 1996 period. Our work extends earlier studies by using a longer sample period, explicit measurement of the news surprises, and separate intraday and overnight return series. We have attempted to minimize the stale price problem in the U.S. and Japanese markets by using 10:00 quotes instead of

[14] Further details are available from the authors upon request.

opening index quotes. We also generalize the standard GARCH conditional volatility model to admit asymmetries in the volatility impact of good and bad news from both domestic and foreign markets.

Our results show that while domestic macroeconomic news announcements are far more important than foreign news announcements in explaining domestic returns, the reverse appears to be true for volatility. We find little evidence that economic news announcements in the three countries have systematic *independent* effects on the return process in any of the three national stock markets. Furthermore, the return spillover estimates do not appear to vary systematically from days when there are announcements to days when there are no announcements. These results seem to be consistent with generally weak results noted in earlier studies of the effects of macroeconomic news announcements. However, there is evidence that some specific, mostly *domestic*, news announcements affect the return spillover from foreign market to domestic market.

There is evidence that volatility spillover effects differ from days with good news to days with bad news. We find that return volatility spillover effects between markets are affected with many becoming insignificant when specific economic news announcements are incorporated in our conditional volatility model. This result provides evidence that economic news announcements may *directly* account for the observed volatility spillovers between markets.

Overall, our results suggest that macroeconomic news announcements play a more important role in explaining volatility linkage between stock markets than in explaining their return linkage. Economic news announcements form a subset of investors' public information; our finding is therefore consistent with Ross' (1989) intuition that information flow is formally and informally connected to volatility in financial markets.

Table 1: Estimates of news announcement effects on the conditional mean function

$R_{i,t} = \beta \cdot \mathbf{X}_t + \gamma \cdot \mathbf{DN}_t + \delta \cdot \mathbf{FN}_t + \alpha \cdot R_{f,t-1} + \varepsilon_{i,t}$ i = JP, UK, US

	Open to Close Sample			Close to Open Sample		
Panel A	$\mathbf{R}_{JP}$	$\mathbf{R}_{UK}$	$\mathbf{R}_{US}$	$\mathbf{R}_{JP}$	$\mathbf{R}_{UK}$	$\mathbf{R}_{US}$
$R_{US, t-1}$	.388*	.037***		.231*	.326*	
	(.076)	(.020)		(.052)	(.031)	
$R_{UK, t-1}$	.138*		.383*	.257		.262*
	(.039)		(.125)	(.183)		(.054)
$R_{JP, t-1}$		.051*	.069*		.102*	.055**
		(.014)	(.017)		(.022)	(.028)
ADV_{US}	.001	-.0001	-.001*	.0001	.0001	-.001***
	(.001)	(.0004)	(.0003)	(.002)	(.0003)	(.0003)
ADV_{UK}	-.002**	.0004	-.0001	.001	-.001**	-.0002
	(.001)	(.0004)	(.0007)	(.001)	(.0003)	(.0004)
ADV_{JP}	-.0003	.0005***	.0004	.002***	-.0003	.0000
	(.001)	(.0003)	(.0004)	(.001)	(.0003)	(.0007)

Note: The X vector includes a constant term and a Monday-holiday dummy variable. The DN variable is a dummy variable for news announcements in the domestic market, and FN is a vector of dummy variables for news announcements in foreign markets. The specific variables are ADV_{US}, ADV_{UK}, and ADV_{JP}. Standard errors are reported in parentheses beneath parameter estimates. The standard errors are estimated using the Newey-West consistent covariance matrix estimator. We indicate statistical significance at the one, five and ten per cent levels by *, **, and ***.

continued on next page

Table 1, continued

$R_{i,t} = \beta \cdot \mathbf{X}_t + \gamma \cdot \mathbf{DN}_t + \delta \cdot \mathbf{FN}_t + \alpha \cdot R_{f,t-1} + \varepsilon_{i,t}$ i = JP, UK, US

	Open to Close Sample			Close to Open Sample		
Panel B	$\mathbf{R_{JP}}$	$\mathbf{R_{UK}}$	$\mathbf{R_{US}}$	$\mathbf{R_{JP}}$	$\mathbf{R_{UK}}$	$\mathbf{R_{US}}$
$R_{US,\ t-1}$	.339*	-.036***		.231*	.327*	
	(.076)	(.020)		(.052)	(.031)	
$R_{UK,\ t-1}$	.136*		.389*	.255		.262*
	(.039)		(.125)	(.180)		(.055)
$R_{JP,\ t-1}$		.050*	.070*		.102*	.055**
		(.014)	(.017)		(.022)	(.028)
CPI_{US}	.002	-.001	-.001	.002	.001	-.0005
	(.002)	(.001)	(.001)	(.002)	(.001)	(.0006)
CPI_{UK}	.003	.002	-.006*	-.003	-.0006	-.004*
	(.006)	(.002)	(.002)	(.003)	(.003)	(.001)
CPI_{JP}	.0006	-.0002	.0008	.0003	-.0003	-.001**
	(.0005)	(.0002)	(.0009)	(.001)	(.0002)	(.0004)
PPI_{US}	-.0001	-.0004	.0003	.0008	.0003	.0004
	(.001)	(.001)	(.001)	(.001)	(.003)	(.0003)
PPI_{JP}	.003	.001	-.001	.001	.0007	.001
	(.002)	(.001)	(.002)	(.002)	(.002)	(.001)
MS_{US}	-.0000***	-.00001**	.0000	-.000	.0000	.0000
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
MS_{UK}	-.009	.002	.001	-.0000	-.001	.0001
	(.006)	(.001)	(.002)	(.002)	(.002)	(.001)
MS_{JP}	-.115	-.016	-.130	-.291**	.020	-.050
	(.143)	(.075)	(.110)	(.136)	(.061)	(.077)
IP_{US}	.003***	.001**	-.002*	.001	-.001**	-.0002
	(.002)	(.0005)	(.001)	(.001)	(.0006)	(.0004)
IP_{UK}	-.0003	.0003	-.0006	.0003	.001	-.0003
	(.001)	(.001)	(.0007)	(.001)	(.001)	(.0005)

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Table 1, continued

$R_{i,t} = \beta \cdot \mathbf{X}_t + \gamma \cdot \mathbf{DN}_t + \delta \cdot \mathbf{FN}_t + \alpha \cdot R_{f,t-1} + \varepsilon_{i,t}$ i = JP, UK, US

	Open to Close Sample			Close to Open Sample		
Panel B	**R_{JP}**	**R_{UK}**	**R_{US}**	**R_{JP}**	**R_{UK}**	**R_{US}**
IP_{JP}	.0002	.0000	-.0005***	.001	.0004	.0001
	(.0005)	(.0003)	(.0003)	(.001)	(.0002)	(.0002)
UR_{US}	-.017	-.030	.055	-.110	-.040***	.002
	(.050)	(.025)	(.036)	(.090)	(.023)	(.025)
UR_{UK}	.010	.009*	-.003	.005	.001	.0004
	(.014)	(.002)	(.003)	(.012)	(.002)	(.002)
TD_{US}	-.002	.005***	-.0002	-.002	.003	-.001
	(.006)	(.003)	(.004)	(.005)	(.004)	(.003)

Note: The X vector includes a constant term and a Monday-holiday dummy variable. The news variables give the unexpected value of the specific economic news announcements for the domestic and foreign markets. The specific news announcements are as follows: CPI_j is the consumer price inflation, PPI_j is the producer price inflation, MS_j is the money supply, IP_j is industrial production, UR_j is the unemployment rate, and TD_j is the trade deficit, all in the j^{th} country. Standard errors are reported in parentheses beneath parameter estimates. The standard errors are estimated using the Newey-West consistent covariance matrix estimator. We indicate statistical significance at the one, five and ten per cent levels by *, **, and ***.

Table 2: Estimates of news announcement effects on conditional mean function

$$R_{i,t} = \beta \cdot \mathbf{X}_t + (\alpha + \varphi \cdot \mathbf{DN}_t + \theta \cdot \mathbf{FN}_t) \cdot R_{f,t-1} + \varepsilon_{d,t} \qquad i = \text{JP, UK, US}$$

	Open to Close Sample			Close to Open Sample		
Panel A	**R_{JP}**	**R_{UK}**	**R_{US}**	**R_{JP}**	**R_{UK}**	**R_{US}**
$R_{US,t-1}$	.313* (.059)	-.059** (.026)		.326* (.035)	.319* (.045)	
$R_{UK,t-1}$	.127* (.045)		.420** (.184)	.328 (.222)		.234* (.053)
$R_{JP,t-1}$		.048* (.014)	.110* (.040)		.102* (.023)	.087* (.021)
$ADV_{US} \bullet R_{US}$	-.077 (.075)	.077 (.050)		-.106 (.117)	.011 (.052)	
$ADV_{US} \bullet R_{UK}$			-.064 (.157)			.082 (.063)
$ADV_{US} \bullet R_{J}$			-.034 (.046)			.026 (.026)
$ADV_{UK} \bullet R_{US}$		-.039 (.064)			.017 (.088)	
$ADV_{UK} \bullet R_{UK}$	-.191 (.145)		-.129 (.142)	-.292 (.233)		-.004 (.078)
$ADV_{UK} \bullet R_{JP}$		-.044 (.035)			-.027 (.048)	
$ADV_{JP} \bullet R_{US}$	.100 (.101)			-.152*** (.086)		
$ADV_{JP} \bullet R_{UK}$	.126 (.120)			-.286 (.264)		
$ADV_{JP} \bullet R_{JP}$		.041*** (.025)	-.169 (.117)		.018 (.041)	-.219 (.155)

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Table 2, continued

$R_{i,t} = \beta \cdot \mathbf{X}_t + (\alpha + \varphi \cdot \mathbf{DN}_t + \theta \cdot \mathbf{FN}_t) \cdot R_{f,t-1} + \varepsilon_{d,t}$ i = JP, UK, US

Panel B	Open to Close Sample R_{JP}	Open to Close Sample R_{UK}	Open to Close Sample R_{US}	Close to Open Sample R_{JP}	Close to Open Sample R_{UK}	Close to Open Sample R_{US}
$R_{US,t-1}$	.294*	-.036***		.277*	.326*	
	(.042)	(.020)		(.044)	(.034)	
$R_{UK,t-1}$	.129*		.383*	.265		.258*
	(.039)		(.131)	(.183)		(.056)
$R_{JP,t-1}$		.049*	.073*		.100*	.061*
		(.014)	(.020)		(.022)	(.017)
$CPI_{US} \bullet R_{US}$	.340	.262*		.079	.214**	
	(.223)	(.101)		(.144)	(.105)	
$CPI_{UK} \bullet R_{US}$		.219			.094	
		(.331)			(.696)	
$CPI_{JP} \bullet R_{US}$	.029			-.034		
	(.077)			(.043)		
$CPI_{US} \bullet R_{UK}$			-.089			-1.147*
			(.228)			(.214)
$CPI_{UK} \bullet R_{UK}$	-1.990***		.948	-1.390***		-.230*
	(1.14)		(.676)	(.867)		(.091)
$CPI_{JP} \bullet R_{UK}$	-.025			-.353*		
	(.120)			(.118)		
$CPI_{US} \bullet R_{JP}$			.132			.062
			(.091)			(.060)
$CPI_{UK} \bullet R_{JP}$		-.742*			.537	
		(.297)			(.350)	
$CPI_{JP} \bullet R_{JP}$		.010	-.048**		-.021	.065**
		(.011)	(.021)		(.020)	(.029)
$PPI_{US} \bullet R_{US}$	.458**	.015		.040	.091	
	(.200)	(.069)		(.227)	(.071)	
$PPI_{US} \bullet R_{UK}$			.128			.009
			(.165)			(.074)

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Table 2, continued

$R_{i,t} = \beta \cdot \mathbf{X}_t + (\alpha + \varphi \cdot \mathbf{DN}_t + \theta \cdot \mathbf{FN}_t) \cdot R_{f,t-1} + \varepsilon_{d,t}$ i = JP, UK, US

	Open to Close Sample			Close to Open Sample		
Panel B	$\mathbf{R_{JP}}$	$\mathbf{R_{UK}}$	$\mathbf{R_{US}}$	$\mathbf{R_{JP}}$	$\mathbf{R_{UK}}$	$\mathbf{R_{US}}$
$PPI_{US} \bullet R_{JP}$			-.031 (.057)			-.021 (.029)
$PPI_{JP} \bullet R_{US}$	-.792* (.310)			-.184 (.170)		
$PPI_{JP} \bullet R_{UK}$	-.120 (.489)			-.407 (.441)		
$PPI_{JP} \bullet R_{JP}$		-.123 (.183)	-.221 (.266)		-.041 (.200)	-.079 (.174)
$MS_{US} \bullet R_{US}$	.002 (.001)	-.001 (.001)		-.002 (.002)	.0000 (.0004)	
$MS_{US} \bullet R_{UK}$			-.010* (.003)			-.007* (.002)
$MS_{US} \bullet R_{JP}$			.007* (.002)			.005* (.001)
$MS_{UK} \bullet R_{US}$		-.255 (.262)			.272 (.359)	
$MS_{UK} \bullet R_{UK}$	-.493 (.670)		-.106 (.361)	-.164 (.448)		-.010 (.214)
$MS_{UK} \bullet R_{JP}$		-.052 (.186)			.166 (.147)	
$MS_{JP} \bullet R_{US}$	91.58* (23.97)			-58.64* (19.59)		
$MS_{JP} \bullet R_{UK}$	17.77 (27.03)			-8.78 (20.08)		
$MS_{JP} \bullet R_{JP}$		3.666 (8.09)	-39.15 (38.89)		5.542 (8.11)	-68.0 (50.6)
$IP_{US} \bullet R_{US}$	.270 (.216)	.123 (.098)		-.055 (.143)	.222*** (.137)	

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Table 2, continued

$R_{i,t} = \beta \cdot \mathbf{X}_t + (\alpha + \varphi \cdot \mathbf{DN}_t + \theta \cdot \mathbf{FN}_t) \cdot R_{f,t-1} + \varepsilon_{d,t}$ i = JP, UK, US

	Open to Close Sample			Close to Open Sample		
Panel B	$\mathbf{R_{JP}}$	$\mathbf{R_{UK}}$	$\mathbf{R_{US}}$	$\mathbf{R_{JP}}$	$\mathbf{R_{UK}}$	$\mathbf{R_{US}}$
$IP_{US} \bullet R_{UK}$			.084			-.102**
			(.117)			(.044)
$IP_{US} \bullet R_{JP}$			-.052			-.049*
			(.042)			(.020)
$IP_{UK} \bullet R_{US}$		-.090			.234***	
		(.078)			(.135)	
$IP_{UK} \bullet R_{UK}$	-.120		-.059	-.114		-.061
	(.115)		(.161)	(.119)		(.080)
$IP_{UK} \bullet R_{JP}$		.011			.076	
		(.031)			(.062)	
$IP_{JP} \bullet R_{US}$	-.126			-.077		
	(.122)			(.094)		
$IP_{JP} \bullet R_{UK}$	-.145			-.092		
	(.099)			(.089)		
$IP_{JP} \bullet R_{JP}$		-.035	.016		.042	.029
		(.030)	(.040)		(.058)	(.039)
$UR_{US} \bullet R_{US}$	6.665*	1.260		5.716**	.456	
	(2.52)	(4.08)		(2.69)	(2.07)	
$UR_{US} \bullet R_{UK}$			-7.470			.058
			(7.44)			(3.87)
$UR_{US} \bullet R_{JP}$			4.727***			3.89***
			(2.63)			(2.21)
$UR_{UK} \bullet R_{US}$		-.091			-.959	
		(.645)			(.858)	
$UR_{UK} \bullet R_{UK}$	-1.465		.292**	2.472		-.256*
	(3.41)		(.142)	(2.59)		(.073)
$UR_{UK} \bullet R_{JP}$		-.045			-.025	
		(.490)			(.303)	

continued on next page

Table 2, continued

$R_{i,t} = \beta \cdot \mathbf{X}_t + (\alpha + \varphi \cdot \mathbf{DN}_t + \theta \cdot \mathbf{FN}_t) \cdot R_{f,t-1} + \varepsilon_{d,t}$ i = JP, UK, US

	Open to Close Sample			Close to Open Sample		
Panel B	$\mathbf{R_{JP}}$	$\mathbf{R_{UK}}$	$\mathbf{R_{US}}$	$\mathbf{R_{JP}}$	$\mathbf{R_{UK}}$	$\mathbf{R_{US}}$
$TD_{US} \bullet R_{US}$	-.198	-.301		-.100	-.117	
	(.366)	(.272)		(.271)	(.401)	
$TD_{US} \bullet R_{UK}$			.019			-.376***
			(.471)			(.224)
$TD_{US} \bullet R_{JP}$			-.374			-.474***
			(.569)			(.302)

Notes to Panel A: The X vector includes a constant term and a Monday-holiday dummy variable. The DN variable is a dummy variable for news announcements in the domestic market, and FN is a vector of dummy variables for news announcements in foreign markets. The ADV_{US}, ADV_{UK}, and ADV_{JP} variables are dummy variables for news announcements in the U.S., U.K., and Japanese markets. Standard errors are reported in parentheses beneath parameter estimates. The standard errors are estimated using the Newey-West consistent covariance matrix estimator. We indicate statistical significance at the one, five and ten per cent levels by *, **, and ***.

Notes to Panel B: The X vector includes a constant term and a Monday-holiday dummy variable. The DN vector contains values of the surprise component in domestic economic news announcements while FN contains a vector of variables giving the surprise component in foreign economic news announcements. The specific news announcements are as follows: CPI_j is the consumer price inflation, PPI_j is the producer price inflation, MS_j is the money supply, IP_j is industrial production, UR_j is the unemployment rate, and TD_j is the trade deficit, all in the j^{th} country. Standard errors are reported in parentheses beneath parameter estimates. The standard errors are estimated using the Newey-West consistent covariance matrix estimator. We indicate statistical significance at the one, five and ten per cent levels by *, **, and ***.

Table 3: Estimates of news announcement effects on conditional volatility function

$$h_{i,t} = \omega_0 + \omega_1 \cdot h_{i,t-1} + \omega_2 \cdot \varepsilon^2_{d\,t-1} + \omega_3 \cdot \varepsilon^{2-}_{d\,t-1} + \omega_4 \cdot \varepsilon^2_{f\,t-1} + \omega_5 \cdot \varepsilon^{2-}_{f\,t-1} + \omega_6 \cdot i_{i,t} + \nu \cdot DN_t + \psi \cdot FN_t$$

i = JP, UK, US

	Open to Close Sample			Close to Open Sample		
Panel A	**R_{JP}**	**R_{UK}**	**R_{US}**	**R_{JP}**	**R_{UK}**	**R_{US}**
Intercept	.041* (.005)[a]	.018* (.002)[a]	-.009* (.003)[a]	-.007* (.001)[a]	.011* (.002)[a]	.007* (.003)[a]
ε^2_{US}	-.039* (.006)	.035* (.011)	.015*** (.009)	.002 (.026)	.024* (.008)	.092* (.023)
ε^2_{UK}	.108* (.044)	.129* (.022)	.022*** (.011)	.007 (.023)	.096* (.025)	.001 (.012)
ε^2_{JP}	.120* (.015)	.019* (.005)	-.003*** (.001)	.279* (.013)	.021* (.002)	.003 (.003)
h_{t-1}	.433* (.026)	.467* (.027)	.857* (.012)	.520* (.015)	-.003* (.001)	.001* (.0003)
$\varepsilon^2_{US} \bullet DN$	.221* (.008)	.075* (.013)	.081* (.015)	-.001 (.026)	.000 (.001)	070*** (.039)
$\varepsilon^2_{UK} \bullet DN$	-.147* (.048)	-.158* (.021)	.004 (.003)	.407* (.045)	.048*** (.026)	.017 (.021)
$\varepsilon^2_{JP} \bullet DN$	.215* (.028)	.010 (.008)	.073* (.065)	.346* (.039)	.001 (.007)	.017 (.015)
D_{US}	-.019* (.006)[a]	-.013* (.001)[a]	.003 (.002)	.009* (.002)[a]	.003** (.001)[a]	.0004 (.002)[a]
D_{UK}	.008 (.008)[a]	-.005* (.002)[a]	-.003*** (.002)	.006* (.002)[a]	.002* (.0003)	.003* (.0005)
D_{JP}	-.011* (.005)[a]	.005* (.002)[a]	-.005** (.002)	.033* (.002)	.003* (.001)	.003* (.0004)
i	.0003 (.0005)[a]	.001* (.0001)[a]	.001* (.0001)[a]	-.0003*** (.0002)	.001* (.0003)	.001** (.0005)

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Table 3, continued

$$h_{i,t} = \omega_0 + \omega_1 \cdot h_{i,t-1} + \omega_2 \cdot \varepsilon^2_{d\,t-1} + \omega_3 \cdot \varepsilon^{2-}_{d\,t-1} + \omega_4 \cdot \varepsilon^2_{f\,t-1} + \omega_5 \cdot \varepsilon^{2-}_{f\,t-1} + \omega_6 \cdot i_{i,t} + \nu \cdot DN_t + \psi \cdot FN_t$$

	Open to Close Sample			Close to Open Sample		
Panel B	**R_{JP}**	**R_{UK}**	**R_{US}**	**R_{JP}**	**R_{UK}**	**R_{US}**
Intercept	.058* (.010)[a]	.035* (.004)[a]	.0002 (.0002)	.037 (.118)a	.003 (.005)[a]	-.002 (.002)
ε^2_{US}	.997* (.295)	.220* (.032)	.032 (.105)	-.577 (2.94)	-.003 (.022)	.126* (.048)
ε^2_{UK}	.383** (.191)	-.038* (.010)	-.091 (.447)	-.589 (.569)	.045 (.076)	.252* (.048)
ε^2_{JP}	.158* (.014)	.001 (.001)	-.001 (.001)	.133 (.273)	.008 (.009)	-.0009* (.0002)
h_{t-1}	.086* (.004)	-.004*** (.002)	.471* (.056)	.082* (.007)	.219* (.035)	.022* (.003)
ε^2_{US} • DN	-1.049* (.310)	-.001 (.004)	.024 (.246)	-.613 (2.93)	.056 (.063)	-.098*** (.051)
ε^2_{UK} • DN	-.306*** (.180)	.116* (.026)	.393 (.681)	.370 (1.46)	.124 (.099)	.189* (.085)
ε^2_{JP} • DN	-.113* (.002)	.008 (.007)	.048 (.060)	.210 (.280)	.029 (.021)	.001 (.001)
i	.0002 (.002)[a]	.001*** (.0003)[a]	.002 (.014)[a]	.007 (.024)	.001 (.022)[a]	.002* (.0004)[a]
CPI_{US}	.067* (.011)[a]	.005 (.006)[a]	.0002 (.0003)	.006* (.0004)	.005 (.010)[a]	.010* (.001)[a]
CPI_{UK}	-.001* (.000)	.065*** (.038)[a]	-.0001 (.001)	.001 (.003)	.039 (.049)[a]	.017 (.016)[a]
CPI_{JP}	-.013 (.036)[a]	.002 (.003)[a]	-.0002* (.00004)	.003* (.0006)	.007* (.002)[a]	-.049* (.004)[a]
IP_{US}	.025 (.039)[a]	-.013** (.006)[a]	-.0002 (.0001)	.002*** (.001)	.002 (.008)[a]	.008* (.001)[a]
IP_{UK}	-.049** (.023)[a]	.004 (.009)	.0002 (.0001)	-.001 (.0005)	.002 (.013)	-.003* (.001)[a]

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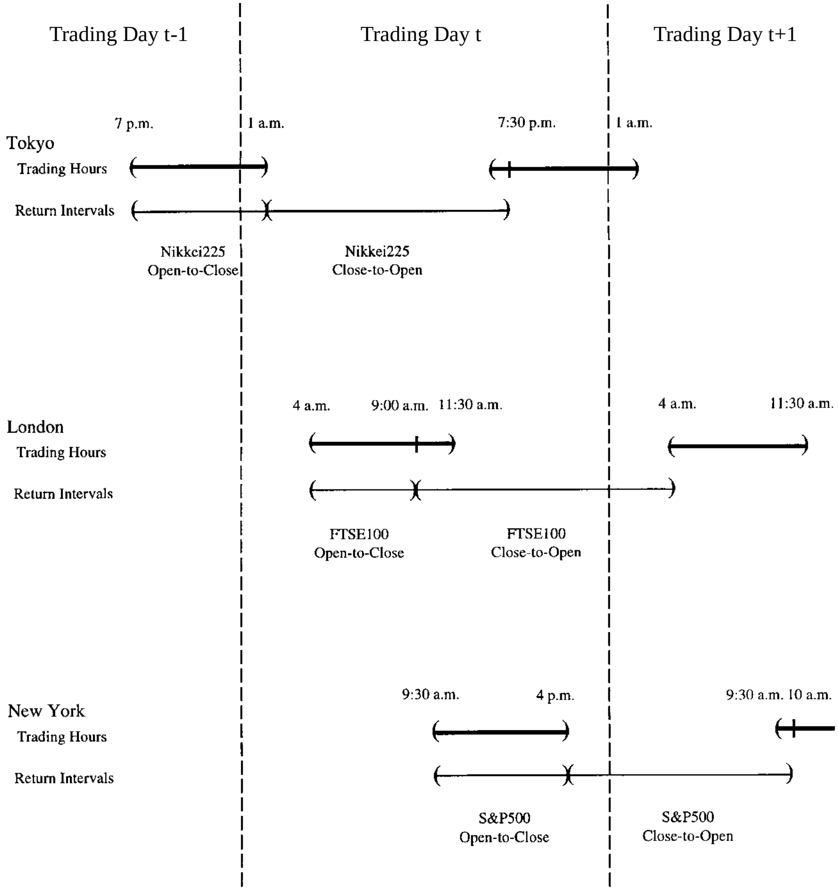
Table 3, continued

$$h_{i,t} = \omega_0 + \omega_1 \cdot h_{i,t-1} + \omega_2 \cdot \varepsilon^2_{d\,t-1} + \omega_3 \cdot \varepsilon^{2-}_{d\,t-1} + \omega_4 \cdot \varepsilon^2_{f\,t-1} + \omega_5 \cdot \varepsilon^{2-}_{f\,t-1} + \omega_6 \cdot i_{i,t} + \nu \cdot DN_t + \psi \cdot FN_t$$

	Open to Close Sample			Close to Open Sample		
Panel B	**R_{JP}**	**R_{UK}**	**R_{US}**	**R_{JP}**	**R_{UK}**	**R_{US}**
IP_{JP}	.046*	.006**	.0002*	-.0001	.002	.000
	(.005)[a]	(.002)[a]	(.0001)	(.0001)	(.004)[a]	(.000)
MS_{JP}	.011*	-.866	-.016	.017	.081	-.057
	(.001)	(.920)[a]	(.031)	(.045)	(.353)[a]	(.476)[a]
MS_{UK}	.095	-.017	.0001	.007*	-.021	-.017*
	(.059)	(.030)	(.001)	(.001)	(.030)[a]	(.006)[a]
MS_{US}	-.001	-.0002***	-.003	.020	.0001	-.000
	(.002)[a]	(.0001)[a]	(.004)[a]	(.041)	(.004)[a]	(.000)
PPI_{JP}	-.159*	.008	-.0005	.001*	-.221	.024*
	(.041)[a]	(.021)[a]	(.0005)	(.0002)	(.026)	(.003)[a]
PPI_{US}	.166*	.001	-.0003*	-.001	-.005	-.002
	(.010)[a]	(.006)[a]	(.0001)	(.001)	(.009)	(.003)
UR_{UK}	-.003*	-.082***	-.0002	.004	-.0002*	-.005
	(.0002)	(.043)	(.001)	(.005)	(.000)	(.004)[a]
UR_{US}	-.011	-.024	.002	-.001	.001*	.341*
	(.010)	(.214)[a]	(.009)	(.006)	(.0002)	(.032)[a]
TD_{US}	-.272*	.022	-.0003	-.001	.032	.035
	(.062)[a]	(.032)[a]	(.001)	(.004)	(.088)[a]	(.041)[a]

Notes to Panels A, B, and C: The $\varepsilon^2_{f\,t-1}$ vector includes lagged variance shocks (squared return residuals) from previous market segments. These are computed using the conditional mean function whose estimates are reported in Table 1, Panel B. The ADV_{US}, ADV_{UK}, and ADV_{JP} variables are dummy variables for news announcements in the U.S., U.K. and Japanese markets. The continuous news variables give the unexpected value of the specific economic news announcements for the domestic and foreign markets. The specific continuous news announcements are as follows: CPI_j is the consumer price inflation, PPI_j is the producer price inflation, MS_j is the money supply, IP_j is industrial production, UR_j is the unemployment rate, and TD_j is the trade deficit, all in the j^{th} country. Estimates are computed using the method outlined in Engle, Ito, and Lin (1990). Estimates are computed using the maximum likelihood method outlined in Engle, Ito, and Lin (1990). Standard errors are reported in parentheses beneath parameter estimates, and are estimated using the Newey-West consistent covariance matrix estimator. We indicate statistical significance at the one, five and ten per cent levels by *, **, and ***. The a superscript indicates an estimate and its standard error has been multiplied by 1000.

Figure 1: Timing conventions for intraday and overnight returns for the U.S., U.K., and Japanese markets



*All times indicated are based on New York time.

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