On Stock Market Return Co-Movements: Macroeconomic News, Dispersion of Beliefs, and Contagion

by

Robert A. Connolly

Kenan-Flagler Business School
CB3490, McColl Building
University of North Carolina - Chapel Hill
Chapel Hill, NC 27599-3490
(919) 962-0053 (phone)
(919) 962-2068 (fax)
connollr@bschool.unc.edu
http://itr.bschool.unc.edu/faculty/connolly

F. Albert Wang

Jesse H. Jones School of Management
MS - 531
Rice University
Houston, TX 77005-1892
(713) 348-5404 (phone)
(713) 348-5251 (fax)
wangfa@rice.edu
http://www.ruf.rice.edu/~wangfa

June 2000

We would like to thank Takatoshi Ito, Oleg Bondarenko, and conference participants at the 1999 Western Finance Association Meeting and the 7th Global Finance Conference for helpful comments. We are also grateful to the Bank of Japan for providing some Japanese macroeconomic announcement data and particularly to Chris Stivers for return dispersion data and useful comments. The usual caveat applies.

On Stock Market Return Co-Movements: Macroeconomic News, Dispersion of Beliefs, and Contagion

Abstract

We document and explain the return co-movement for the U.S., U.K., and Japanese equity markets for the period of 1985-1996. Our empirical results show the importance of imperfect signal-extraction in explaining the equity market return co-movement. In such a setting, domestic investors try to extract the unobservable global factors from foreign market returns and use the extracted information in their subsequent domestic trading. In this imperfect learning environment, domestic investors respond to the foreign return signal more strongly if the signal is more precise. In addition, trading noise may also affect the return-generating process in domestic markets. We find this contagion effect is most pronounced in the extreme down markets.

JEL Classification: G15, F30.

Keywords: market co-movement, contagion, signal extraction, macroeconomic news, dispersion of beliefs.

On Stock Market Return Co-Movements:
Macroeconomic News, Dispersion of Beliefs, and Contagion

I. Introduction

With the advent of new technology and market integration around the globe, the capital allocation process in international equity markets increasingly extends beyond national boundaries. One noticeable feature of this interdependence is that these markets tend to influence each other and often appear to move together (e.g., Becker, Finnerty, and Gupta (1990), Hamao, Masulis, Ng (1990), and Lin, Engle, and Ito (1994)). The nature and origins of stock market linkages has attracted attention from academics, investment professionals, and government regulators. Drawing on international asset pricing models such as Solnik (1974a, 1974b), Stulz (1981), and Adler and Dumas (1983), the traditional view stresses the role of common fundamental factors. Nonetheless, these models seem to be unable to explain important characteristics of global market interdependence, particularly when there are extreme market movements such as the Crash in October 1987.

Macroeconomic variables may be reasonable proxies for the common factors because they reflect the economic activity affecting cash flows and/or discount rates. In this approach, the unexpected component of macroeconomic news announcements may have important market valuation effects. Early empirical evidence is mixed on this (e.g., Cornell (1983), Pearce and Roley (1985), Hardevoulis (1987), Ito and Roley (1987), McQueen and Roley (1993)). To the extent that there are global common factors in business cycles, macroeconomic news in one country may reveal information about future cash flows or discount rates in many countries, not just in the home country. This suggests that one source of market return co-movement may be macroeconomic news announcements. On the other hand, even if a macroeconomic news announcement reveals information only for the home country, foreign markets may still respond to such a local news through cross-market hedging demands from internationally diversified investors (Fleming, Kirby, and Ostdiek (1998)).

King and Wadhwani (1990) propose a signal-extraction model of cross-market return and volatility spillovers. Consider two non-concurrent markets, where the trading in market 1 opens and closes before the trading day begins in market 2. When market 2 opens, investors in this market incorporate all available information when making their trading decisions. One element of their information set is the previous intraday trading result in market 1. To the extent that a portion of trading in market 1 may reflect global factors with potential valuation implications for market 2, investors in market 2 will try to extract that valuable information from market 1 trading results and use the extracted information in their subsequent trading. Such an information extraction process can produce the return co-movement phenomenon observed between national equity markets.

Lin, Engle, and Ito (1994) provide early evidence consistent with the signal-extraction hypothesis for the U.S. and Japanese markets. They estimate the common factors from the foreign intraday returns with a signal-extraction model and find that the extracted common factors are relevant for pricing in multiple markets. Since they do not control for macroeconomic news announcements, it is difficult to interpret what the extracted common factors represent. In particular, it is unclear to what extent such common factors reflect observable public economic news or unobservable global factors.

This is important because the signal-extraction hypothesis hinges on a critical premise that there are some *unobservable* global factors significantly affecting many markets, and that information about these factors may be extracted, albeit imperfectly, from foreign market returns. However, if return co-movements between markets are driven primarily by observable global factors, such as public macroeconomic news, then the previous foreign market returns represent essentially redundant information. In this case, investors trading in market 2 can trade optimally on the basis of the observable global factors and, as a result, foreign market returns should have no obvious valuation implications.

In this paper, we test the signal-extraction hypothesis for equity market return co-movements, while controlling for the potential effect of macroeconomic news announcements. Specifically, we examine the co-movement between the domestic intraday (or overnight) returns and foreign intraday returns for the U.S., U.K, and Japanese markets, conditional on a representative set of macroeconomic news announcements from these three countries for the period of 1985-1996. If foreign intraday returns significantly influence domestic market returns even after controlling for the potential macroeconomic news effects, this renders more convincing evidence in favor of the signal-extraction hypothesis.

In our empirical analysis, several interesting patterns of market return co-movements emerge. First, foreign intraday returns significantly and positively affect domestic market returns. Second, foreign market returns affect domestic overnight returns more profoundly than domestic intraday returns. Since domestic overnight returns precede domestic intraday returns, the greater impact on the domestic overnight returns implies that the domestic market processes the information contained in foreign market trading quickly and efficiently. This result strengthens confidence in similar findings in Lin, Engle, and Ito (1994) for the U.S. and Japanese markets. Third, the nearby foreign market exerts a greater influence on domestic intraday (overnight) market than the more distant foreign market and the preceding domestic overnight (intraday) market. The only exception is the U.K. overnight market where the more distant U.S. market has a greater influence than the nearby Japanese market. Fourth, domestic intraday markets tend to reverse returns realized in the preceding domestic overnight markets, whereas domestic overnight markets tend to display momentum relative to returns in the preceding domestic intraday markets.

In addition, we find that foreign economic news announcements have larger effects on domestic returns when the announcements are accompanied by large foreign market returns. This result is consistent with McQueen and Roley (1993) and Kaminsky and Schmukler(1999) and suggests that the effect of macroeconomic announcements depend on the context in which investors interpret the announcements, not just the news itself. Nevertheless, the four distinct patterns of return co-movements persist with essentially the same magnitude, even after controlling for the effect of macroeconomic news announcements. Furthermore, in each country the two foreign intraday returns are much more important than the lagged domestic market returns and all economic news announcements taken together in explaining the domestic intraday and overnight returns.

Our empirical results suggest that the great bulk of the observed market return co-movement between national equity markets cannot be attributed to public information flow, as measured by the news about macroeconomic fundamentals. These results are consistent with recent evidence in suggesting that macroeconomic news announcements are not a primary determinant of the observed market return co-movement (e.g., King, Sentana, and Wadhwani (1994), Karolyi and Stulz (1996), and Connolly and Wang (1998)). Instead, foreign market returns may convey important information distinct from public economic news announcements. This suggests the importance of the *unobservable* global factors that domestic investors try to extract from foreign market returns.

Apart from global factors, some trading in foreign markets could be due to local factors that have no value implication for the domestic market. Domestic investors, however, cannot distinguish perfectly between the influences of global and local factors on foreign market returns. As a result, the return signal from foreign market trading is essentially "imperfect". The *imperfect* signal-extraction process has an important implication: the response of domestic investors to the return signal from foreign market trading depends on the quality of the signal. That is, holding the true signal constant, greater precision in the foreign return signal leads domestic investors to rely more heavily on foreign returns in their subsequent trading and this leads to greater market return co-movement.

We test this implication using each market's cross-sectional return dispersion as an inverse measure of the quality of the market return signal (Stivers (1999)). In this context, higher return dispersion reflects a less precise signal about the unobservable global factors. Therefore, the return co-movement between markets should decrease as the return dispersion increases, *ceteris paribus*. Our empirical evidence is consistent with this implication. Specifically, we find a smaller market return co-movement between the domestic market return and the nearby foreign market return in the presence of a higher foreign market return dispersion. Furthermore, the effect is more pronounced for domestic overnight returns than for domestic intraday returns.

The *imperfect* signal-extraction process has another important implication: trading noise in foreign markets could also spillover into the domestic market through this imperfect learning process. King and Wadhwani (1990) term this phenomenon a "contagion" effect.[1] Christie and Huang (1995) and

1. For interesting alternative models of financial market contagion see Calvo and Mendoza (1999) and Kodres and Pritsker (1999).

Bae, Karolyi, and Stulz (2000) measure the presence of contagion by extreme market returns. In this paper, we argue that extreme negative returns from one market may trigger a liquidity crunch among large financial institutions that they may be forced to liquidate some of their positions subsequently in other markets. Such a liquidity crunch can thereby cause further market co-movements for the extreme negative return days. The effect is asymmetric: there is no reason to expect the same result for the extreme positive return days. Accordingly, we expect the contagion effect to be "asymmetric" in the sense that market returns are highly correlated in bad times, but much less correlated in good times.[2]

In this paper, we test the *asymmetric* contagion effect using extreme returns as the proxy for contagion. Our empirical results confirm the existence of the asymmetric contagion effect. Specifically, we find that the return co-movement between domestic and foreign markets are significantly higher when foreign markets are experiencing extreme negative returns, but not with extreme positive returns. Furthermore, once we account for the existence of the asymmetric contagion effect, there is no reliable remaining return co-movement between Japanese intraday returns and the U.S. and U.K. intraday returns. In other words, there is evidence that the return co-movements between these market segments are driven primarily by the asymmetric contagion effect.

Overall, our empirical results lend support to the imperfect signal-extraction hypothesis (King and Wadhwani (1990)) in explaining the return co-movement between national equity markets. In particular, there is evidence that domestic investors try to extract the unobservable global factors from foreign market returns and use the extracted information in their subsequent domestic trading. In this imperfect learning environment, domestic investors respond to the foreign return signal more strongly if the signal is more precise, i.e., the return dispersion is relatively small. Furthermore, trading noise from foreign market can also spillover into domestic markets and such a contagion effect is more pronounced for the extreme negative return days than for the extreme positive return days.

The paper proceeds as follows. Section II introduces the empirical models used in our analysis, and discusses the corresponding predictions. Section III describes data and presents a preliminary

2. See Kyle and Xiong (1999) for a theoretical model of asymmetric contagion.

correlation analysis. Section IV reports our regression results and an empirical model comparison analysis. Section V concludes.

II. Modeling Market Return Co-Movements

A. General Framework and Notation

In our study of market return co-movements, we use the Nikkei 225 index for the Japanese market, the FTSE 100 index for the U.K. market, and the S&P 500 index for the U.S. market. In each market, the daily close-to-close returns are separated into overnight (i.e., close-to-open) and intraday (i.e., open-to-close) returns. To emphasize this decomposition, denote by $NKON_t$ and $NKID_t$ the Japanese overnight and intraday returns, respectively, on calendar day t. Similarly, denote by $FTON_t$ and $FTID_t$ for U.K. overnight and intraday returns, and by $SPON_t$ and $SPID_t$ for U.S. overnight and intraday returns on calendar day t.

We decompose further each of the two stock index returns into two components: one that is expected, conditional on the current information, and another which is the unexpected (residual) component. We focus initially on the information set generated by the current overnight and intraday returns and macroeconomic news announcements from both the domestic and foreign countries. Specifically, the current information set for the domestic intraday stock index returns on calendar day t, denoted by $\Omega_{i,t}^{ID}$, contains the domestic and foreign returns and announcements, if any, that occur between domestic market i's close on calendar day t-1 and the subsequent domestic market close on calendar day t. On the other hand, the domestic overnight stock index returns on calendar day t, denoted by $\Omega_{i,t}^{ON}$, consists of the domestic and foreign market returns and announcements, if any, that occur between domestic market i's open on calendar day t-1 and the subsequent domestic market open on calendar day t.

The current information set for Japanese domestic intraday stock index returns, $\Omega_{i,t}^{ID}$, contains the previous domestic overnight return ($NKON_t$), the previous U.S. intraday return ($SPID_{t-1}$), the previous U.K. intraday return ($FTID_{t-1}$), Japanese news announcements ($NEWS_{JP,t}$), U.S. news announcements

($NEWS_{US,t-1}$), and U.K. news announcements ($NEWS_{UK,t-1}$). Note that since Japanese markets precede the U.K. and U.S. markets in calendar time, the current information set relevant to Japanese intraday returns on calendar day t includes returns and news from the U.K. and U.S. markets on calendar day t-1. Thus, the decomposition of Japanese intraday returns ($NKID_t$), is

Intraday: $$NKID_t = E(NKID_t | \Omega^{ID}_{JP,t}) + e^{ID}_{JP,t}, \tag{1}$$

where *E(?)* is the expected return component, $e^{ID}_{JP,t}$ is the unexpected (residual) return component, and the information set is given by $\Omega^{ID}_{JP,t} = \{NKON_t,\ SPID_{t-1},\ FTID_{t-1},\ NEWS_{JP,t},\ NEWS_{US,t-1},\ NEWS_{UK,t-1}\}$. Similarly, the decomposition of Japanese overnight returns, $NKON_t$, is

Overnight: $$NKON_t = E(NKON_t | \Omega^{ON}_{JP,t}) + e^{ON}_{JP,t}, \tag{2}$$

where *E(?)* is the expected return component, $e^{ON}_{JP,t}$ is the unexpected (residual) return component, and the information set is given by $\Omega^{ON}_{JP,t} = \{NKID_{t-1},\ SPID_{t-1},\ FTID_{t-1},\ NEWS_{JP,t-1},\ NEWS_{US,t-1},\ NEWS_{UK,t-1}\}$.

The information set for the domestic overnight return on calendar day t contains the domestic intraday return and news on calendar day t-1. This reflects the fact that domestic overnight returns precede domestic intraday returns in our setup. Otherwise, the current intraday and overnight information sets contain the same information regarding foreign market returns and news. The decomposition of U.K. and U.S. intraday and overnight returns into expected and residual components is analogously done.

B. Conditional Mean Models

B.1. The Linear News Model

In this paper, we use a representative set of macroeconomic news announcements from all three countries to proxy for the observable global factor(s). In the following discussion of empirical models, we continue to use Japanese return models as our regular working examples. The U.K. and U.S. return models are analogously formulated, and hence are omitted to conserve space.

Our baseline model for the intraday and overnight return co-movement between markets, ignoring the news effect, is shown in equations (3) and (4):

Intraday: $$NKID_t = a + r_1 \cdot NKON_t + r_2 \cdot SPID_{t-1} + r_3 \cdot FTID_{t-1} + j \cdot MHD_{JP,t} + e^{ID}_{JP,t} \tag{3}$$

Overnight: $NKON_t = a + r_1 \cdot NKID_{t-1} + r_2 \cdot SPID_{t-1} + r_3 \cdot FTID_{t-1} + j \cdot MHD_{JP,t} + e^{ON}_{JP,t}$ (4)

Our model for intraday and overnight return co-movement between markets incorporating news effect is shown in equations (5) and (6):

$$\begin{aligned} NKID_t = a + r_1 \cdot NKON_t + r_2 \cdot SPID_{t-1} + r_3 \cdot FTID_{t-1} + b_1 \cdot NEWS_{JP,t} + b_2 \cdot NEWS_{US,t-1} \\ + b_3 \cdot NEWS_{UK,t-1} + j \cdot MHD_{JP,t} + e^{ID}_{JP,t} \end{aligned} \tag{5}$$

$$\begin{aligned} NKON_t = a + r_1 \cdot NKID_{t-1} + r_2 \cdot SPID_{t-1} + r_3 \cdot FTID_{t-1} + b_1 \cdot NEWS_{JP,t-1} + b_2 \cdot NEWS_{US,t-1} \\ + b_3 \cdot NEWS_{UK,t-1} + j \cdot MHD_{JP,t} + e^{ON}_{JP,t} \end{aligned} \tag{6}$$

It is worth noting that all news variables represent vectors of the *unexpected* components of the corresponding macroeconomic announcements.[3] We also control for the Monday and post-holiday effect (French (1980) and Gibbons and Hess (1981)) using the Monday and post-holiday dummy variable, $MHD_{JP,t}$. Using these models, we can measure the effects of various market return co-movements with the coefficients r_1, r_2 and r_3.

If proxies for observable global factors, such as public macroeconomic news, capture the primary source of return co-movement, then the return co-movement between markets should diminish substantially, once we account for the observable global factors. That is, we expect to find that estimates of coefficients r_1, r_2 and r_3 should decrease substantially in (5) and (6) (relative to (3) and (4)) if observable global factors dominate unobservable global factors in asset valuation. On the other hand, if we incorporate measures of observable global factors and the return co-movements persist with similar magnitudes, this constitutes evidence favoring the primacy of *unobservable* global factors and the signal-extraction hypothesis.

It is independently interesting to compare the impact of the previous domestic overnight (intraday) return, i.e., the coefficient r_1, on the current domestic intraday (overnight) return relative to the previous foreign intraday returns, i.e., the coefficients r_2 and r_3. Note that all foreign intraday trading considered in this paper occurs essentially during the domestic overnight trading period. Therefore, if the domestic market is efficient in processing information during the preceding intraday foreign intraday session, we expect that the foreign intraday returns significantly affect the domestic overnight returns, but have a smaller effect on domestic intraday returns. Lin, Engle, and Ito (1994) find evidence of this for the U.S. and Japanese markets. In this paper, we assess whether this result still holds after incorporating public information flows and a third country, the U.K.

B.2. The Nonlinear News Model

There is evidence that the effect of macroeconomic announcements depends on the context in which investors interpret the announcements, not just the news itself (McQueen and Roley (1993) and Kaminsky and Schmukler(1999)). In this spirit, we posit that foreign macroeconomic announcements are more likely to contain relevant new information for the domestic stock market if they are accompanied by measurable foreign trading returns. As a result, such conditional foreign news announcements may have a significant impact on the domestic stock market, while the unconditional foreign news by itself may have no effect at all. Allowing for this possibility, we develop a nonlinear model to gauge the interactive effect of foreign intraday returns and news announcements on the domestic intraday and overnight returns. Specifically, we address conditional effects in the following model:

$$\begin{aligned} NKID_t &= a + (r_1 + b_1 \cdot NEWS_{JP,t}) \cdot NKON_t + (r_2 + b_2 \cdot NEWS_{US,t-1}) \cdot SPID_{t-1} \\ &\quad + (r_3 + b_3 \cdot NEWS_{UK,t-1}) \cdot FTID_{t-1} + j \cdot MHD_{JP,t} + e^{ID}_{JP,t} \end{aligned} \tag{7}$$

$$\begin{aligned} NKON_t &= a + (r_1 + b_1 \cdot NEWS_{JP,t-1}) \cdot NKID_{t-1} + (r_2 + b_2 \cdot NEWS_{US,t-1}) \cdot SPID_{t-1} \\ &\quad + (r_3 + b_3 \cdot NEWS_{UK,t-1}) \cdot FTID_{t-1} + j \cdot MHD_{JP,t} + e^{ON}_{JP,t} \end{aligned} \tag{8}$$

If our conjecture, that the news effect is greater when accompanied by measurable foreign trading returns, is right, then we should find more news announcements are significant in the nonlinear news model in comparison to the corresponding estimates from the linear news model in (5) and (6). Furthermore, one can gauge whether or not the patterns of return co-movements established under the linear news model change substantially in the nonlinear news model. This issue can be addressed by comparing the coefficients of return co-movements (i.e., r_1, r_2 and r_3) in the nonlinear model to the corresponding coefficients in the linear model.

---

3. We discuss the methods for calculating the unexpected component of news announcements in Section III.

B.3. Nonlinear News Model with Return Dispersion

While some trading in market 1 may be related to global factors, some trading may reflect local factors that have no valuation implications for market 2. It may be quite difficult for investors in market 2 to distinguish global from local factors based on the intraday trading return in market 1. That is, the signal provided by trading results from market 1 may be imperfect. These limits on the signal-extraction process have one important implication: the reaction of investors in market 2 to the trading in market 1 depends on the quality of the signal. That is, holding the information content constant, the better the quality of the signal from market 1 is, the stronger the reaction of investors in market 2.

In this paper, we test this implication using each market's cross-sectional return dispersion as an inverse measure of the quality of the market's return signal. Higher return dispersion may reflect less agreement among investors concerning the observed market return or a less precise return signal about the unobservable global factors. As a result, the investors in market 2 may discount the signal provided by foreign market returns when the corresponding return dispersion is large. That is, return co-movement should decrease as the return dispersion increases, *ceteris paribus*.

Following Stivers (1999), the market cross-sectional return dispersion (RD) at time t is given by:

$$RD_{M,t} = \sqrt{\frac{1}{n-1}\sum_{i=1}^{n}(R_{i,t} - R_{M,t})^2} \qquad (9)$$

where n is the number of firms in the market, and subscript M indicates market-wide values.[4] To implement our empirical tests, we measure RD using daily returns for the U.S. firms in the largest capitalization decile of the NYSE-AMEX, but we use the Nikkei 225 for Japan and the FTSE-100 for the U.K. This approach reduces to a minimum the chance of spurious results related to high small-firm idiosyncratic volatility or non-synchronous trading.

4. Amihud and Mendelson (1989) use return dispersion to measure the inefficiency of the price discovery process and Bessembinder, Chan, and Seguin (1996) use it to proxy for firm-level information flows.

As noted in earlier studies, it is essential to recognize the relation between the market's cross-sectional return dispersion and the market return.[5] Our relative RD measure (RRD) is intended to reflect the level of the market's cross-sectional return dispersion after controlling for variation in return dispersion attributed to the magnitude of the market return. Our relative RD measure, denoted RRD, is the residual, u_t, from a regression of RD on a function of the aggregate excess market return:

$$RD_t = f_0 + g_1 \left| R_t^e \right| + g_2 Dum1_t \left| R_t^e \right| + u_t \tag{10}$$

where R_t^e is the excess return of the market-portfolio return for period t using Euro-currency interest rates as the risk-free return, $Dum1_t = 1$ if the portfolio excess return is negative, and is 0 otherwise. This formulation admits a potentially asymmetric relation between portfolio returns and RD as noted in Lamoureux and Pannikath (1994).

We incorporate the relative return dispersion (RRD) into our nonlinear news model for intraday and overnight returns because we believe that market return co-movement may depend on the dispersion of beliefs during earlier domestic and foreign trading sessions. The specific empirical models are:

$$\begin{aligned} NKID_t = a + (r_1 + b_1 \cdot NEWS_{JP,t}) \cdot NKON_t + (r_2 + b_2 \cdot NEWS_{US,t-1} + d_2 \cdot RRD_{US,t-1}) \cdot SPID_{t-1} \\ + (r_3 + b_3 \cdot NEWS_{UK,t-1} + d_3 \cdot RRD_{UK,t-1}) \cdot FTID_{t-1} + j \cdot MHD_{JP,t} + e_{JP,t}^{ID} \end{aligned} \tag{11}$$

$$\begin{aligned} NKON_t = a + (r_1 + b_1 \cdot NEWS_{JP,t-1}) \cdot NKID_{t-1} + (r_2 + b_2 \cdot NEWS_{US,t-1} + d_2 \cdot RRD_{US,t-1}) \cdot SPID_{t-1} \\ + (r_3 + b_3 \cdot NEWS_{UK,t-1} + d_3 \cdot RRD_{UK,t-1}) \cdot FTID_{t-1} + j \cdot MHD_{JP,t} + e_{JP,t}^{ON} \end{aligned} \tag{12}$$

Using (11) and (12), we can test the first implication of the imperfect signal-extraction hypothesis: the reaction of investors in market 2 to the trading in market 1 depends on the quality of the signal. We expect the return co-movement should decrease as the return dispersion increases. This means that the coefficients of the relative return dispersions (RRD) should be negative (i.e., $d_2 < 0$ and/or $d_3 < 0$). In addition, by comparing the estimates of the return co-movements (i.e., r_1, r_2, and r_3) in (11) and (12) to the same estimates in (7) and (8), we may gauge the incremental effect of controls on signal quality on estimates of market return co-movement.

5. Stivers (1999) contains an extensive analysis of the relation between a portfolio's cross-sectional return dispersion and the market's excess return.

B.4. Nonlinear News Model with Extreme Returns

Another implication of the signal-extraction hypothesis is that trading noise in market 1 (created, say, by shifts in investor sentiment) can also spillover into market 2 through this imperfect information extraction mechanism. King and Wadhwani (1990) call this phenomenon a contagion effect. If investors could distinguish between noise and information as the source of foreign intraday trading, then they would only respond to the foreign market movement when it reflects new information. In this approach, contagion is a by-product of the imperfect signal-extraction process.

Although it is difficult to distinguish between the two possible sources of the co-movement between markets (noise or information), one may still gauge the separate influence of contagion from that of information. Imagine that on a 'normal' day, there are some positive noise and some negative noise and in the aggregate market return, the two may roughly offset each other. On such a day, the imperfect signal-extraction process would reflect mainly the influence of information and hence, investors in market 2 may rationally extract information from the return movement in market 1. On the other hand, there are days when there may exist a significant imbalance between the positive noise and the negative noise. In particular, investors in market 2 may herd on extreme (positive or negative) noise from market 1, which leads to extreme (positive or negative) return movements in market 2. Unless the extreme return movements are coupled with significant rational revaluation of the market (due, perhaps, to some discrete news release), the extreme return movements may be regarded as a manifestation of investor herding or contagion. In light of this, we posit that contagion involves an extraordinary co-movement between markets under the imperfect signal-extraction process.[6]

In this paper, we measure potential contagion effects on market return co-movement by adding negative and positive extreme return dummies for the left and right tails of the return distribution, denoted D^L and D^U, to the nonlinear news model.[7] The resulting model is given by:

---

6. On the definition of contagion, see also Forbes and Rigobon (1998, 2000).

7. Christie and Huang (1995) and Bae, Karolyi, and Stulz (2000) also use extreme return movements as proxies for investor herding or contagion.

$$NKID_t = a + (r_1 + b_1 \cdot NEWS_{JP,t}) \cdot NKON_t + (r_2 + b_2 \cdot NEWS_{US,t-1} + h_{21} \cdot D^L_{US,t-1} + h_{22} \cdot D^U_{US,t-1}) \cdot SPID_{t-1}$$
$$+ (r_3 + b_3 \cdot NEWS_{UK,t-1} + h_{31} \cdot D^L_{UK,t-1} + h_{32} \cdot D^U_{UK,t-1}) \cdot FTID_{t-1} + j \cdot MHD_{JP,t} + e^{ID}_{JP,t} \qquad (13)$$

$$NKON_t = a + (r_1 + b_1 \cdot NEWS_{JP,t-1}) \cdot NKID_{t-1} + (r_2 + b_2 \cdot NEWS_{US,t-1} + h_{21} \cdot D^L_{US,t-1} + h_{22} \cdot D^U_{US,t-1}) \cdot SPID_{t-1}$$
$$+ (r_3 + b_3 \cdot NEWS_{UK,t-1} + h_{31} \cdot D^L_{UK,t-1} + h_{32} \cdot D^U_{UK,t-1}) \cdot FTID_{t-1} + j \cdot MHD_{JP,t} + e^{ON}_{JP,t}. \qquad (14)$$

The dummy variables $D^L_{US,t-1}(D^U_{US,t-1})$ are one if the lagged U.S. intraday return is in the first (99th) percentile of its own return distribution, and zero otherwise. We define the other dummy variables in a completely analogous fashion.[8] If contagion is important in explaining return co-movement relative to the nearby foreign market, then the coefficients of the corresponding dummy variables should be significantly positive (i.e., $h_{21} > 0$ and $h_{22} > 0$). The same analysis applies to the extreme market movements involving the more distant foreign market (i.e., $h_{31} > 0$ and $h_{32} > 0$).

Large financial institutions tend to have positions in many markets. When one market is experiencing extreme market down movement, these institutions or their clients may need to liquidate their positions in the other markets in order to cover losses in the first market. Such a liquidity crunch or wealth constraint may generate further market co-movements on extreme return days.[9] Note that such a liquidity effect implies an asymmetric return co-movement between markets. That is, markets tend to *fall* together for extreme downward movements from foreign markets (i.e., $h_{21} > 0$ and $h_{31} > 0$), but not for extreme upward movements. In other words, the liquidity effect exacerbates downward market co-movements. Note, however, that this liquidity effect may be triggered either by a sudden, significant downward revaluation of the market, or simply by investor herding on pessimistic sentiment. Regardless of the source of the liquidity crunch, an asymmetric contagion effect will result.

In this expended structure, the coefficients r_2 and r_3 are measures of the normal return co-movement between the domestic market return and the two foreign market returns due to the usual signal-extraction process. And, the coefficients h_{21} and h_{31} are measures of the incremental return co-movement due to specifically the asymmetric contagion effect between these markets. If the return co-movement is

8. We also considered the fifth (95th) percentile rule, but it did not affect our results.

principally due to the asymmetric contagion effect, then the h_{21} and h_{31} estimates should be statistically and economically significant while the r_2 and r_3 estimates should be appreciably smaller.

B.5. Nonlinear News Model with Return Dispersion and Extreme Returns

To contrast the effects of macroeconomic news, difference in beliefs, and contagion, we employ a nonlinear news model with both return dispersion (RRD) and extreme returns (D^L and D^U). The encompassing (all-inclusive) model is given by:

$$\begin{aligned} NKID_t = a + (r_1 + b_1 \cdot NEWS_{JP,t}) \cdot NKON_t + (r_2 + b_2 \cdot NEWS_{US,t-1} + h_{21} \cdot D^L_{US,t-1} + h_{22} \cdot D^U_{US,t-1} \\ + d_2 \cdot RRD_{US,t-1}) \cdot SPID_{t-1} + (r_3 + b_3 \cdot NEWS_{UK,t-1} + h_{31} \cdot D^L_{UK,t-1} + h_{32} \cdot D^U_{UK,t-1} \\ + d_3 \cdot RRD_{UK,t-1}) \cdot FTID_{t-1} + j \cdot MHD_{JP,t} + e^{ID}_{JP,t} \end{aligned} \tag{15}$$

$$\begin{aligned} NKON_t = a + (r_1 + b_1 \cdot NEWS_{JP,t-1}) \cdot NKID_{t-1} + (r_2 + b_2 \cdot NEWS_{US,t-1} + h_{21} \cdot D^L_{US,t-1} + h_{22} \cdot D^U_{US,t-1} \\ + d_2 \cdot RRD_{US,t-1}) \cdot SPID_{t-1} + (r_3 + b_3 \cdot NEWS_{UK,t-1} + h_{31} \cdot D^L_{UK,t-1} + h_{32} \cdot D^U_{UK,t-1} \\ + d_3 \cdot RRD_{UK,t-1}) \cdot FTID_{t-1} + j \cdot MHD_{JP,t} + e^{ON}_{JP,t} \end{aligned} \tag{16}$$

Of course, using this model, we can estimate the unconditional market return co-movements given by r_1, r_2, and r_3. Most interestingly, if these unconditional estimates are statistically and economically significant even after controlling for news, dispersion of beliefs, and contagion, this constitutes support for the signal-extraction hypothesis. In essence, it suggests that the return co-movements be driven primarily by the unobservable global factors. Domestic investors try to extract these factors from the previous foreign market returns and, subsequently, they use this extracted information in their trading. Note that as long as the learning is imperfect, the signal-extraction process can give rise to the contagion effect. That is, noise in one market can also spillover into the next market in this process.

C. Conditional Volatility Models

Extensive empirical evidence indicates that stock return volatility exhibit clustering phenomenon, i.e., high volatility in one period tends to be followed by high volatility in the next period (Bollerslev, Chou, and Kroner (1992)). In addition, this volatility clustering effect is asymmetric, i.e., the effect is

---

9. See Kyle and Xiong (1999) for more on this.

more pronounced when the market is *falling* than rising (Nelson (1990), Engle and Ng (1993) and Glosten, Jagannathan and Runkle (1993)). In order to account for the asymmetric volatility clustering effect, we model the intraday and overnight residual terms, i.e., $e^{ID}_{JP,t}$ and $e^{ON}_{JP,t}$, respectively, using the Glosten-Jagannathan-Runkle (GJR) asymmetric GARCH volatility model as follows:

$$e^{ID}_{i,t} \sim N(0, h^{ID}_{i,t}),\ \ h^{ID}_{i,t} = w_{i,0} + w_{i,1} \cdot e^{2,ID}_{i,t-1} + w_{i,2} \cdot h^{ID}_{i,t-1} + w_{i,3} \cdot e^{2-,ID}_{i,t-1} + w_{i,4} \cdot MHD_{i,t} + w_{i,5} \cdot INT_{i,t-1} \tag{17}$$

$$e^{ON}_{i,t} \sim N(0, h^{ON}_{i,t}),\ h^{ON}_{i,t} = w_{i,0} + w_{i,1} \cdot e^{2,ON}_{i,t-1} + w_{i,2} \cdot h^{ON}_{i,t-1} + w_{i,3} \cdot e^{2-,ON}_{i,t-1} + w_{i,4} \cdot MHD_{i,t} + w_{i,5} \cdot INT_{i,t-1} \tag{18}$$

where i = JP, UK, US and $e^{2-}_{i,t-1} = e^{2}_{i,t-1}$ if $e_{i,t-1} < 0$, but is 0 otherwise.

Note that the term $e^{2-}_{i,t-1}$ accounts for the asymmetric clustering effect. In addition, our volatility model captures any Monday and post-holiday-related variations in conditional volatility by including the Monday and post-holiday dummy variable, $MHD_{i,t}$. The model also accounts for the interest rate effect (Glosten, Jagannathan, and Runkle (1993)) by incorporating a comparable domestic interest rate, $INT_{i,t-1}$, for each country. We use a nonlinear maximum-likelihood method to estimate each conditional mean model along with the corresponding conditional volatility model given by (17) and (18). We applied the volatility model specification tests (Engle and Ng (1993)) to insure that our estimates are based on well-behaved volatility functions.

D. Model Comparison

The multiplicity of models raises the issue of selecting a 'best' model for explaining equity market return co-movements. To compare the various conditional mean models delineated in subsection B, we compare the posterior odds of the different models. The following example illustrates the principal features of the posterior odds analysis.

Suppose a hypothesis, H_1, is associated with return generating model 1, and another hypothesis, H_2, is associated with return generating model 2. Denote the conditional likelihood function giving the probability of the (say) intraday return on the Nikkei 225 as $p(NKID|ß = ß_i)$. With the prior probability of hypothesis indicated by $p(H_i)$, Bayes Theorem says that the posterior probability of hypothesis i, conditional on the data is given by $p(H_i|NKID) = p(H_i)p(NKID|ß = ß)/p(NKID)$. Here, $p(NKID)$ is the weighted likelihood function where the prior probabilities in $p(H_i)$ serve as the weights. Using these results, the posterior probability of hypothesis 1 relative to hypothesis 2 is given by

$$\frac{p(H_1|NKID)}{p(H_2|NKID)} = \frac{p(H_1)}{p(H_2)} \cdot \frac{p(NKID|b = b_1)}{p(NKID|b = b_2)} \tag{19}$$

where the posterior odds are on the left-hand side of (19), the prior odds are given by $p(H_1)/ p(H_2)$, and the likelihood ratio (also known as the Bayes Factor) is given by $p(NKID|ß = ß_1)/ p(NKID|ß = ß_2)$.

In application, the trick is to find explicit forms for the prior probability densities and the likelihood functions that deliver tractable and economically interesting posterior odds computations. Using diffuse priors, Leamer (1978) demonstrates that (19) may be written as

$$p(H_1|NKID)/p(H_2|NKID) = T^{-(d/2)}(SSE_2/SSE_1)^{T/2} \tag{20}$$

where T is the sample size, d is the difference in the number of parameters under the two hypotheses, and SSE_i is the sum of squared errors for the i^{th} hypothesized model. In the event that the two hypotheses imply nested empirical models, the posterior odds may be rewritten in terms of the usual F statistic:

$$p(H_1|NKID)/p(H_2|NKID) = T^{-(d/2)}[(dF/(T\text{-}k))+1]^{T/2} \tag{21}$$

where k is the number of parameters in the larger model and F is the usual F statistic.

Note that with a fixed number of parameters, any difference in the sum of squared errors (or model R^2) can be rendered statistically significant by increasing the sample size. That is, with large sample sizes (for us, $T \tilde{} 2500$), there is a strong chance that the p-value of standard test statistics will be a misleading indicator of the probability associated with a particular hypothesis.[10] Klein and Brown (1984) note that (19) may be interpreted “in terms of an adjusted likelihood ratio test in which for large samples the significance level is a specified function of the sample size.” While our empirical applications of posterior odds rely on diffuse priors, extending this analysis to informative priors is relatively direct.[11]

10. See Connolly (1989, 1991) for an extensive analysis of these issues in the context of the weekend effect anomaly literature. Shanken (1987) applies posterior odds analysis to asset pricing model tests.

11. See Connolly (1991, Section 5.2) for an application of informative priors in the context of posterior odds analysis.

III. Data and Correlation Analysis

A. Stock Returns

We estimate our empirical models with daily stock index returns from the U.S., U.K., and Japan for the sample period January 1, 1985 through December 31, 1996. We measure the stock return as the change in the logarithm of the stock price index. Most of the stock price index data are from Datastream. Figure 1 shows the timing of the three markets. The Japanese market opens at 9:00am and closes at 3:00pm, although there is a two-hour break between 11:00am to 1:00pm (Tokyo local time). The U.K. market opens at 9:00am and closes at 4:30pm (London local time). The U.S. market opens at 9:30am and closes at 4:00pm (New York local time).

In the study, we consider close-to-open "overnight" returns and open-to-close "intraday" returns. Since there are potential stale-price problems associated with opening prices of the U.S. (Stoll and Whaley (1990)) and Japanese (Hamao, Masulis, Ng (1990)) stock indices, we collected intraday values for these indices and computed intraday and overnight returns with post-opening prices of the respective indices. For Japan, we collected 9:15 and 10:00 Nikkei 225 quotes by hand from *Nihon Keizai Shinbun*.[12] We found important differences in the distribution moments for returns calculated using the official opening vs. the later quotes. Generally, overnight returns calculated with the opening quotes displayed a much smaller mean return and lower volatility than when overnight returns use the 10:00 quote. After consulting with experts at the Bank of Japan, we choose the 10:00 price as the opening price in our analysis to minimize the stale-price problems in the Japanese market. The Japanese overnight return on day t ($NKON_t$) is calculated as ln(Nikkei 225 at 10 a.m. on day t) – ln(Nikkei 225 at close on day t-1). The Japanese intraday return on day t ($NKID_t$) is calculated as ln(Nikkei 225 at close on day t) – ln(Nikkei 225 at 10:00 a.m. on day t).

For our U.S. return series, we used the S&P500 index as our measure of the stock market. We obtained daily opening and closing quotes on the S&P500 index from Datastream. We also obtained half-hourly quotes for the S&P500 index from Prudential Securities for the 1985 - 1993 period, and hourly index values from Standard and Poor's for the last three years of our sample. As with the Japanese return data, we found important differences in the distribution moments of returns using the later quotes.[13] Following Becker, Finnerty, and Tucker (1992) and Craig, Dravid and Richardson (1995), we choose the 10:00 price as the opening price in our analysis to minimize the stale-price problems in the U.S. market. The U.S. overnight return on day t ($SPON_t$) is calculated as ln(S&P500 at 10 a.m. on day t) – ln(S&P500 at close on day t-1). The U.S. intraday return on day t ($SPID_t$) is the percentage change in the value of the S&P500 from 10:00 to the close of trading on day t, calculated as ln(S&P500 at close on day t) – ln(S&P500 at 10:00 a.m. on day t).

For the U.K., we use FTSE 100 index with the 9:00 a.m. opening prices and 4:30 p.m. closing prices. The stale price problem is much less important for this market, because firm quotes are used instead of transaction prices and there is a half-hour pre-market period each day. Hence, the U.K. overnight return on day t ($FTON_t$) is calculated as ln(FTSE 100 at 9:00 a.m. on day t) – ln(FTSE 100 at close on day t-1). The U.K. intraday return on day t ($FTID_t$) is the percentage change in the value of the FTSE 100 from 9:00 to the close of trading, calculated as ln(FTSE 100 at close on day t) – ln(FTSE 100at 9:00 a.m. on day t).

Figure 1 shows the corresponding intraday and overnight returns in each market. Note that our Japanese intraday return does not overlap with intraday returns in the U.S. and U.K. Japanese overnight returns completely overlap with the previous day's intraday trading period in the U.S. and the U.K. There is a short overlap of one and one-half hours between our U.S. intraday and U.K. intraday measures. This creates some greater cross-market lead-lag correlation between the U.K and U.S. returns. Overnight returns in the U.S. overlap with intraday returns in Japan and for all but one and one-half hours of the U.K. intraday return period. Overnight returns in the U.K. overlap with all but the first one and one-half hours of U.S. intraday trading, and the U.K. overnight period completely encompasses the intraday trading period for Japan.

---

12. In earlier work, we used the broader TOPIX index, which includes 500 stocks in the Japanese stock market. We were unable to find a source of intraday data for the TOPIX index, so we shifted our focus to the Nikkei 225 index.
13. These results are not reported here, but a summary is available upon request.

Table 1, Panel A gives the mean, standard deviation, minimum, maximum, and autocorrelation with various lags for our overnight and intraday returns. The summary statistics are consistent with findings in earlier studies. The volatility of intraday returns is larger than the volatility of overnight returns. Similarly, the average return is often larger for intraday returns than overnight returns for the U.S. and U.K. data. The volatility finding may reflect greater information flow during the intraday period and/or it may reflect the conclusions reached in some recent papers that stress the importance of trading behavior in generating higher volatility.

B. Macroeconomic News Announcements

Our choice of the macroeconomic news announcements is based on a trade-off between a fair representation of economic fundamentals and the need for parsimony. Our U.S. economic news variables include money supply (MS), industrial production (IP), consumer price inflation (CPI), producer price inflation (PPI), the unemployment rate (UR), and the merchandise trade balance (TD). Earlier work has investigated the valuation impact of all of these announcements. Among a large number of papers, Cornell (1983) provides an early assessment of money supply announcements. Hardevoulis (1987) analyzes the stock price effects of the trade deficit announcement, Pearce and Roley (1985) and Hardevoulis study inflation announcements, and recently McQueen and Roley (1993) examine industrial production and unemployment rate among other major announcements. These announcements form a fair and parsimonious representation of the economic fundamentals for the U.S. markets (McQueen and Roley (1993), Pearce and Roley (1993), and Ederington and Lee (1993)).

We provide a quick sketch of the U.S. economic news announcement schedule (see McQueen and Roley (1993) and Ederington and Lee (1993) for more detail). U.S. industrial production announcements are released monthly by the Federal Reserve Board before the U.S. market opens. Unemployment data are released by the Bureau of Labor Statistics monthly before the U.S. market opens. The merchandise trade deficit data are announced monthly by the Foreign Trade Division of the Department of Commerce, also before the U.S. market begins trading. Consumer price index and producer price index data are announced monthly by the Bureau of Labor Statistics at 8:30 a.m. Finally, weekly U.S. money stock

(M1) announcements are made by the Federal Reserve on Thursday afternoons, after the U.S. market closes, but before trading begins in Japan. We obtain the U.S. actual announcement data and the median survey estimates from MMS International (Money Market Services). The median values are calculated from the MMS surveys of market participants and observers, and they serve as a measure of the market's expected value of the particular announcement.

The Japanese economic news announcements that we study include money supply (MS), industrial production (IP), consumer price index (CPI), and wholesale price index (PPI). The money supply announcement reports the monthly money stock (M2 plus certificates of deposit) estimate. The Japanese money supply announcements are usually made on Friday of the first or second week of the month. Until November 28, 1996, the Bank of Japan announced the money supply and wholesale price index in the afternoon while the market was still open. From that point forward, the announcements have been made at 8:50 a.m. Tokyo time. Japanese industrial production data are the monthly index of industrial production computed by the Ministry of International Trade and Industry (MITI). This announcement does not have a fixed release date, but is usually made toward the end of the month. The industrial production announcement is made by the MITI, generally at 1:30 p.m. (while the market in Tokyo is open). The consumer price index is released monthly at slightly irregular times by the Management and Coordination Agency after it is reported to the Japanese cabinet. In our regressions, we treat this announcement as if it is made during the trading day. We collected the Japan monthly *preliminary* announcement data by hand from the *Nihon Keizai Shinbun* for the 1985-1994 period. The Bank of Japan provided some additional announcement data for 1995 and 1996.

Our macroeconomic news announcements for the U.K. include M3 money supply (MS), retail price inflation (CPI), industrial production (IP), and the unemployment rate (UR). These announcements are made monthly at 11:30 a.m. while the market in London is open (King and Wadhwani (1990)). The U.K. announcement data are collected from the *Financial Times*.

Since stock prices might be expected to respond only to the new information contained in each announcement, we use the unexpected component of each announcement in our study, not the raw announcement figures. The unexpected components of the U.S. announcements are calculated as the percentage difference between the actual announcement values and the median expected values, using the actual values as the base. Since we do not have expected values on the macroeconomic announcements for Japan and the U.K., we built a separate ARIMA model for each actual news announcement series. Specifically, we use residual at time t from the final ARIMA model for a specific announcement to measure the unexpected value of the news announcement at time t. These residuals can be interpreted as the unexpected percentage change in the news announcement series, and in this way, they are directly comparable to the measurement approach we use for U.S. announcements. We do not report the details here, but generally, the models are often fairly simple ARMA(1,1) models with an AR(12) term added.

Figure 1 shows the timing of trading in the three markets and the corresponding intraday and overnight returns in each market. Some examples may help to clarify the considerations that went into the determination of the timing of macroeconomic announcements (i.e. specifying the timing on particular announcements) that underlie the regression models (5) – (16).

Example 1: Five of the six U.S. announcements are made before the market opens in the U.S.; the U.S. money supply data are released after the U.S. market closes. Accordingly, the five U.S. announcements are included in both the overnight U.S. and the intraday U.K. models at time t.

Example 2: The U.S. money supply announcement enters both the intraday and overnight U.S., U.K., and Japanese return models at time t-1.

Example 3: Japanese economic news announcements enter U.K. and U.S. return models at time t, but they only enter the intraday Japanese return model at time t. The Japanese news announcements enter the overnight return models at time t-1, because they are made after the 10:00 Nikkei index value is recorded.

Example 4: U.K. announcements enter the U.K. intraday return model and the U.S. overnight and intraday return models at time t. U.K. news announcements appear at time t-1 in the U.K. overnight return model.

Table 1, Panel B reports summary statistics on our macroeconomic news announcement data.

The final, non-news variable that we need to describe is the interest rate term that is included in our conditional volatility model. There are some potential difficulties in finding comparable domestic interest rates for each country. To address this issue, we use the one-month Eurodollar interest rate for the U.S., the one-month Euro-yen interest rate for Japan, and the one-month Euro-sterling interest rate for the U.K. We gathered this data from Datastream, and it consists of afternoon observations taken 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 intraday and overnight return data, we adjust for timing variations so that interest rate observations precede the last equity index value used in calculating equity returns. This renders the interest rate observations pre-determined variables in our empirical work.

C. Cross-market Contemporaneous Correlations

C.1. Unconditional Correlations

Documenting and explaining equity market return co-movements is the principal purpose of this paper. Table 1, Panel C reports cross-market contemporaneous return correlations between foreign intraday returns and domestic overnight returns, while Panel D reports cross-market lead-lag return correlations between foreign intraday returns and subsequent domestic intraday returns.

The analysis of the contemporaneous correlations provides a preliminary, univariate summary of the data that our overnight regression models are designed to explain. The contemporaneous correlations have several interesting features. First, foreign intraday returns are positively and substantially related to domestic overnight returns. Second, U.S. intraday returns exert the greatest influence on both the U.K. and Japanese overnight returns, while the U.K. intraday returns have a greater impact on U.S. overnight returns than Japanese intraday returns. Third, there are distinct asymmetries in the spillover from foreign intraday returns to domestic overnight returns. Specifically, the contemporaneous correlation between the U.S. intraday return at t-1 and the Japanese (U.K.) overnight return at time t is .223 (.556). In contrast, the contemporaneous correlation between the Japanese (U.K.) intraday return at t and the subsequent U.S.

overnight return at time t is .090 (.338). This asymmetry has been noted in earlier studies (e.g., Hamao, Masulis, and Ng (1990)).

C.2. Conditional Correlations

We also stratified the data to provide preliminary evidence on how cross-market correlations vary with news and return dispersion, and on the likely importance of contagion effects. To focus this discussion, we identified three related questions about market co-movements that the correlation analysis might address. First, is there systematic evidence that market return co-movements are affected by public information flows (i.e., economic news)? Second, do cross-market correlations change systematically when there are extremely large negative (or positive) market movements? Finally, does the dispersion of beliefs in a foreign market systematically affect the market return co-movements? We address each question in turn.

Two results emerge from the news-related variations in the correlation of Japanese market returns with foreign market returns. First, the correlations between Japanese overnight returns and foreign intraday returns are highest on days when there is news in all markets, followed by the days with news from any one market. This correlation is at its nadir on no-news days. Second, the correlation between Japanese intraday returns and U.S. overnight returns, after conditioning on news, is much higher than in the whole sample. These results suggest that the co-movement between Japanese returns and foreign market returns is quite sensitive to the news effect. Nonetheless, the unconditional result that Japanese intraday returns have the least impact on foreign overnight returns in the whole sample is still robust in every news-related subsample. For example, the correlation of U.K. overnight returns with lagged U.S. intraday returns is greater than with Japanese intraday returns in every news-related subsample. A similar result holds for U.S. overnight returns.

The second question focuses on return-related variations in market return co-movements. With one exception, market return co-movements are larger and *positive* when there is bad news compared to the whole sample. For example, the correlation of Japanese overnight returns with lagged U.S. intraday returns is .223 in the whole sample, but .799 in the extreme negative returns (1% fractile) subsample. The only exception is in the U.S. overnight market.

We also find that the contemporaneous correlations are generally smaller when we form samples based on positive return extremes compared with negative return extremes. The tests for differences in correlations confirm that five out of six correlations are higher in the extreme negative return sample than the extreme positive return sample. We also note that the correlations based on extreme returns are dramatically and statistically significantly different from correlations based on samples where returns fall in the interquartile range. These results are consistent with the contagion effect that unusually large noise in foreign intraday returns may also spillover into domestic overnight returns on days with extreme foreign intrady returns. Moreover, the effect is more pronounced with extreme negative returns in foreign markets. The only exception is in the U.S. overnight market, where the Japanese extreme positive returns have a greater influence than its extreme negative returns do.

The final set of return correlations that we analyze is based on variations in the relative return dispersion (RRD). For Japanese overnight returns, the U.S. RRD-sorted correlations (5% vs. 95% fractiles) show partial evidence that the contemporaneous correlations are negatively related to the dispersion of beliefs in U.S. intraday returns. In other words, investors whose trading determines returns in the Japanese overnight market seem to discount the influence of U.S. intraday returns if they find the U.S. return signal is accompanied by unusually high return dispersion. For U.S. overnight returns, the U.K. RRD-sorted correlations display a similar pattern, thus providing further evidence consistent with the signal extraction hypothesis.

Overall, the general patterns in the contemporaneous correlations prevail under a variety of sample stratifications - news, extreme returns, and return dispersion. That is, foreign intraday returns are positively and substantially related to subsequent domestic overnight returns. U.S. intraday returns exert the greatest influence on both U.K. and Japanese overnight returns, while U.K. intraday returns have a greater impact on U.S. overnight returns than Japanese intraday returns. Finally, the asymmetric correlation patterns among the U.S., U.K., and Japanese market returns noted by Hamao, Masulis, and Ng (1990) persists.

D. Cross-market Lead-lag Correlations

D.1. Unconditional Correlations

Table 1, Panel D reports the cross-market lead-lag correlations of intraday returns. The analysis of these correlations provides a preliminary, univariate summary of the data that our intraday regression models are designed to explain.

In the full sample, the lead-lag correlations indicate that foreign intraday returns continue to exert positive impact on domestic intraday returns in the Japanese and U.S. markets, even after their first and substantial impact on the domestic overnight returns. However, the lead-lag correlation is generally less than half of the corresponding contemporaneous correlation. For example, the lead-lag correlation between the U.S. intraday returns and Japanese intraday returns is .096, whereas the contemporaneous correlation between the U.S. intraday returns and Japanese overnight returns is .223. It appears that the signal-extraction process is not completed in the domestic overnight market and the learning continues, albeit to a lesser extent, in the subsequent domestic intraday market.

The largest entry in the full sample portion of the table is the lead-lag cross-correlation between intraday returns in the U.K. and U.S. (.191). In part, this correlation reflects the one and one-half hour overlap of the intraday trading periods for these two markets. Aside from this relatively large value, the rest of the full sample correlations are relatively small.

D.2. Conditional Correlations

Here, stratifying the samples by news, extreme return, and return dispersion produces some interesting changes in the correlations. The news stratification shows that the lead-lag correlations between domestic overnight returns and subsequent domestic intraday returns are generally much higher when there is news from any of the markets. For Japan and the U.K., when we sort the data on domestic news, we find the correlation between domestic overnight and domestic intraday news is negative, relatively sizeable, and statistically different from the comparable correlation in the no-news sample.

The stratification by extreme return indicates that the lead-lag correlations between the foreign intraday returns and subsequent domestic intraday returns are substantially higher when the foreign intraday returns fall into the extreme negative categories. Moreover, the test for differences in correlation for extreme returns indicates that the lead-lag correlations are significantly and negatively related to the extreme returns. This result is consistent with a similar finding in contemporaneous correlations in Panel C. Both results suggest an asymmetric contagion effect: markets tend to move down together in bad times, but may not move up together in good times.

Finally, conditional on extremely small RRD (1% fractile) in the U.K. market, we find the correlation of the U.K. intraday returns with subsequent U.S. intraday returns is much higher than the whole sample result (i.e., .501 vs. .191). Similarly, conditional on extremely small RRD (1% fractile) in the U.S. market, the correlation of the U.S. intraday returns with subsequent U.K. intraday returns is much higher than the whole sample result (i.e., .144 vs. -.004). On the other hand, conditional on unusually high RRD (99% fractile) in the U.S. market, we find the correlation of the U.S. intraday returns with subsequent Japanese intraday returns becomes significantly negative and much less than the whole sample result (i.e., -.523 vs. .096). These results are consistent with the signal-extraction hypothesis, which predicts that the lead-lag correlation between market returns will be negatively related to the dispersion of beliefs in the leading foreign market. We find further evidence for this view in the test for differences in correlation for the two extreme return dispersions (1% vs. 99% fractiles) for all three countries. Note, however, the results based on the less extreme return dispersions (5% vs. 95% fractiles) provide similar evidence only for the lead-lag correlation between the U.K. and U.S. intraday returns. It appears that return dispersion may have a nonlinear effect on the lead-lag correlations.

IV. Empirical Results

A. Linear Models

In this section, we report and discuss the results from estimating various return-generating models described in Section II. Table 2 reports the estimates for our baseline model. These estimates establish four distinct patterns of return co-movements across the three national equity markets studied here. First, foreign intraday returns significantly and positively affect subsequent domestic market returns (i.e., $r_2 > 0$ and/or $r_3 > 0$). Second, the spillover effect is greater for domestic overnight returns than for domestic intraday returns, a result consistent with the findings in Lin, Engle, and Ito (1994). Third, the nearby foreign market returns have a larger impact than the more distant foreign market on the subsequent domestic returns (i.e., $r_2 > r_3$). The only exception involves the U.K. overnight market, where U.S. market returns have a larger impact than Japanese market returns. Fourth, domestic intraday markets tend to reverse the direction of the return movement in the previous domestic overnight markets (i.e., $r_1 < 0$), whereas domestic overnight markets tend to follow the course of the return movement in the previous domestic intraday markets (i.e., $r_1 > 0$).

Table 3 shows that the four distinct patterns about return co-movements persist and have essentially the same magnitudes, even after controlling for the effects of the macroeconomic news announcements. This invariance result suggests that the return co-movements between markets are more likely to be driven by unobservable global factors than by observable macroeconomic news. This result tends to confirm a similar conclusion in King, Sentana, and Wadhwani (1994) about monthly return co-movements between national stock markets. Table 3 also show that both Japanese and U.K. markets respond to domestic as well as foreign macroeconomic news. In contrast, U.S. markets tend to respond only to the domestic macroeconomic news.

B. Nonlinear Models

B.1. Basic Results

In Table 4, we report estimates of the nonlinear news model parameters. The four distinct patterns about return co-movements across the three national equity markets remain and the corresponding parameter estimates have similar magnitudes in this model. However, we find that across all three countries there are more statistically significant macroeconomic news announcement coefficients in the nonlinear model compared to the linear model. In other words, macroeconomic news effects seem to be more important in the nonlinear model than in the linear model. This finding confirms our conjecture that macroeconomic announcements from a country are more likely to contain relevant new information for the domestic stock market if they are accompanied by measurable stock return movements. This result further strengthens the recent findings that the effect of macroeconomic announcements depends on the context in which investors interpret the announcements, not just the news itself (McQueen and Roley (1993) and Kaminsky and Schmukler (1999)).

From taking a closer look at the effect of each individual news announcement in Table 4, the following results seem clear. First, for domestic news announcements, consumer price index (CPI) and money supply (MS) are two important domestic news announcements for Japan and U.K. Industrial production (IP) is also important in Japan, while unemployment rate (UR) is the only important domestic news in U.S. Second, concerning the effect of U.S. news on foreign markets, U.S. consumer price index (CPI) and trade deficit (TD) both have measurable impacts on Japanese and U.K. markets. In contrast, U.S. industrial production (IP) and money supply (MS) affect Japanese markets, whereas U.S. producer price index (PPI) and unemployment rate (UR) affect U.K. markets. Third, concerning the effect of U.K. news on foreign markets, U.K. consumer price index (CPI) and money supply (MS) have measurable impacts on both Japanese and U.S. markets. In addition, U.K. unemployment rate (UR) affects Japanese markets. Fourth, concerning the effect of Japanese news on foreign markets, Japanese industrial production (IP) has measurable impacts on both U.K. and U.S. markets. In addition, Japanese producer price index (PPI) only affects U.K. markets, whereas Japanese consumer price index (CPI) only affects U.S. markets.

B.2. Nonlinear Return Model with Return Dispersion

Table 5 reports estimates of nonlinear news model parameters where the model incorporates foreign return dispersion (RRD) as an additional conditional determinant of equity market return co-movement. One particularly important finding about the role of the return dispersion emerges. Aside from the Japanese intraday market, the greater the return dispersion in the nearby foreign market, the smaller the return co-movement between that foreign market and the domestic market (i.e., $d_2 < 0$). This result is particularly evident for the domestic overnight market. For example, the d_2 coefficients for the

Japanese, U.K., and U.S. overnight markets are all statistically significant with values of -9.15, -0.03, and -0.03, respectively. This result is consistent with the signal extraction hypothesis and suggests that the response of the domestic traders to the return signal from the nearby foreign market decrease with the dispersion of beliefs. Interestingly, the return co-movement between Japanese intraday returns and the U.S. (U.K.) intraday returns, i.e., r_2 (r_3), in Table 4 decreases from 0.21 (0.08) to 0.13 (0.06) in Table 5 after including the two return dispersion variables—about a 30% drop. Apart from these new findings, the results in Table 5 are similar to those in Table 4. Again, the four distinct patterns about return co-movements across the three national equity markets remain, even after controlling for both the effects of public news and dispersion of beliefs.

B.3. Nonlinear Return Model with Contagion

Table 6 reports estimates of the parameters of the nonlinear news model that incorporates extreme foreign return dummies (i.e., 1% and 99% tails) to capture the effects of potential contagion from foreign markets. The results are generally consistent with our conjecture about asymmetric contagion effects in down markets. Specifically, domestic markets tend to move down with foreign markets when the foreign markets are experiencing extreme negative returns (i.e., $h_{21} > 0$ and/or $h_{31} > 0$). On the other hand, domestic markets generally do not move up with foreign markets when the foreign markets are experiencing extreme positive returns. This asymmetric contagion result is consistent with the prediction of Kyle and Xiong (1999). For example, in Japanese intraday market the coefficients of the 1% dummies of the U.S. and U.K. market returns (i.e., h_{21} and h_{31}) are statistically significant with values of 0.38 and 0.15, respectively. On the other hand, the coefficients on the 99% dummies of the U.S. and U.K. market returns in the Japanese market (i.e., h_{22} and h_{32}) are negative or statistically insignificant. More interestingly, the usual co-movement between Japanese intraday returns and the U.S. intraday returns (measured by r_2) disappears, after the extreme foreign returns are incorporated. It seems that the bulk of the return co-movement related to the Japanese intraday market may reflect an asymmetric contagion effect. However, in all other markets the general return co-movement patterns persist and have similar magnitudes even after controlling for both effects of public economic news and contagion. To check for robustness, we repeated the analysis in Table 6 with 5% and 95% return dummies and find essentially the same results.

B.4. All-Inclusive Nonlinear Return Model

In Table 7, we report parameter estimates for the all-inclusive non-linear news model. This specification renders possible simultaneous comparison of macroeconomic news, dispersion of beliefs, and contagion effects on equity market return co-movements. The most important new finding emerges from the Japanese intraday market. Unlike in Table 5, U.S. return dispersion now has a significant, negative impact (d_2 = -6.55) on the return co-movement between the U.S. and Japanese intraday returns. This result is further confirmation of the signal extraction hypothesis, but we find this only after controlling for extreme U.S. and U.K. returns. Like in Table 6, the usual co-movements between Japanese intraday returns and the U.S. and the U.K. intraday returns (i.e., r_2 and r_3) disappear, after both the dispersion of beliefs and the extreme foreign returns are incorporated. This result strengthens the finding in Table 6 and suggests that the bulk of the return co-movement related to the Japanese intraday market may be due to the asymmetric contagion effect. In general, Table 7 confirms both the effect of dispersion of beliefs found in Table 5 and the asymmetric contagion effect found in Table 6.

Above all, the estimates of return co-movements reveal that the distinct patterns among the three equity markets by and large persist with similar magnitudes. First, positive foreign intraday returns tend to raise domestic market returns, especially for domestic overnight markets. Second, the nearby foreign markets tend to exert a greater influence on the domestic markets. The only exception remains the U.K. overnight market, which is much more sensitive to the distant U.S. market than to the nearby Japanese market. Finally, domestic intraday returns tend react negatively to the previous domestic overnight returns, whereas domestic overnight returns tend react positively to the previous domestic intraday returns.

These patterns are established first in our baseline model and are robust in the all-inclusive, non-linear news model. These robust results suggest that the return co-movements, to a large extent, are driven by unobservable global factors. Domestic investors try to extract these factors from the previous foreign market returns and they use this extracted information in subsequent trading. This signal-extraction process may also give rise to a contagion effect in which noise in one market spills over into the next market through this imperfect learning process.

C. Model Comparison

Finally, in Table 8 we compare various conditional mean models in equations (3) – (16). Panel A gives each model's R^2 value and Panel B provides a posterior odds-based comparison of the various models. It is not surprising that R^2 values increase as more variables are added to the models. Posterior odds analysis takes this into account, penalizes models with more parameters, and compares different models based on the posterior probability of each model given the data.

We compare first the baseline model (eqs. 3 and 4) to all other models. Two remarkable findings are in order. First, the baseline model dominates the linear news model (eqs. 5 and 6) and the nonlinear news model (eqs. 7 and 8) for all six intraday and overnight markets. This result strengthens one of our basic conclusions that macroeconomic news announcements have little independent role in explaining the return co-movements between national equity markets. To the extent that our new announcements proxy for observable global factors, the relative orthogonality of news effect and return co-movement strongly suggests that the return co-movements may be driven primarily by unobservable global factors. This implication is consistent with the conclusion of King, Sentana, and Wadhwani (1994), who emphasize the importance of unobservable global factors.

Second, for the U.K. and U.S. overnight markets and the Japanese intraday market the baseline model can be improved by incorporating return dispersion and/or extreme returns. This result suggests that the dispersion of beliefs and market contagion be important factors affecting the return co-movements in these markets. On the other hand, this result also implies that the parsimonious, baseline model is sufficient to capture the return co-movements in the U.K. and U.S. intraday markets and the Japanese overnight market.

The comparison between the linear news model (eqs. 5 and 6) and other models shows that the linear news model may best explain the U.K. intraday market. But, the linear news model is inferior to all other models essentially for all other five markets. For example, posterior odds favor the nonlinear news model (eqs. 7 and 8) over the linear news model for all markets except for the U.K intraday market. This result basically confirms our previous finding that foreign macroeconomic announcements are more likely to contain relevant new information for the domestic stock market if they are accompanied by measurable stock return movements. This result is also consistent with the recent findings that the effect of macroeconomic announcements depends on the context in which investors interpret them, not just the news itself (McQueen and Roley (1993) and Kaminsky and Schmukler (1999)).

Relative to the basic nonlinear news model (eqs. 7 and 8), the posterior odds favor the more complex nonlinear news models with return dispersion and/or extreme returns (eqs. 11 – 16) for all three overnight markets and the Japanese intraday market. This result highlights again the importance of incorporating the potential effects of the dispersion of beliefs and contagion as one examines the return co-movements in these markets.

Finally, posterior odds favor the most complex (all-inclusive) model (eqs. 15 and 16) for explaining the return co-movements for the U.K. and U.S overnight markets. By comparison, posterior odds favor the baseline model for the Japanese overnight market and the U.K. and U.S intraday markets, and the nonlinear news models with extreme returns (eqs. 13 and 14) for the Japanese intraday market.

V. Summary and Conclusions

In this paper, we document and explore the sources of return co-movement for the U.S., U.K., and Japanese equity markets for the 1985-1996 period. Several distinct patterns of market return co-movements emerge from our empirical analysis. First, foreign intraday returns significantly influence subsequent domestic market returns, and the effect is stronger for domestic overnight returns than for domestic intraday returns. Second, the U.S. market exerts the greatest influence on both the U.K. and Japanese markets, while the U.K. market has more influence on the U.S. market than the Japanese market has. Third, domestic intraday returns tend to have a negative correlation with the preceding domestic overnight returns, whereas domestic overnight returns tend to have a positive correlation with the preceding domestic intraday returns.

We establish these distinct patterns first in our baseline model excluding macroeconomic news, and find these results robust essentially in all other models where we incorporate macroeconomic news, dispersion in beliefs, and contagion. In addition, we find that foreign macroeconomic news announcements have larger effects on domestic market returns when the announcements are accompanied by large foreign market returns. Still, the news effect on returns is small in comparison to the effects of foreign market returns. In particular, for all three countries the two foreign intraday returns are much more important in explaining the domestic intraday and overnight returns than the lagged domestic market returns and all macroeconomic news announcements taken together. This result confirms a similar finding in Connolly and Wang (1998).

These empirical results are consistent with the imperfect signal-extraction explanation (King and Wadhwani (1990)) of equity market return co-movements. Specifically, our results are consistent with a model in which domestic investors try to extract the unobservable global factors from foreign market returns and use the extracted information in their subsequent domestic trading. In this imperfect learning environment, domestic investors discount the foreign return signal more strongly as the signal is less precise. Finally, we find evidence that trading noise may also affect the return-generating process in domestic markets and that this contagion effect is more pronounced for the extreme down markets than for the extreme up markets. The Crash in October 1987 appears to provide a leading example of this asymmetric contagion effect.

REFERENCES

Adler, M. and B. Dumas, 1983, "The International Portfolio Choice and Corporate Finance," *Journal of Finance*, 38, 925-984.

Amihud, Y. and H. Mendelson, 1989, "Market Microstructure and Price Discovery on the Tokyo Stock Exchange," *Japan and the World Economy* I, 341-370.

Bae, K-H.., Karolyi, A., and Stulz, R., 2000, "A New Approach to Measuring Financial Contagion," Ohio State University Working Paper.

Baig, T. and Goldfajn, I. (1998), "Financial Market Contagion in the Asian Crisis," International Monetary Fund Working Paper.

Becker, Kent G., Finnerty, Joseph E., and Gupta, Manoj, 1990, "The Intertemporal Relation Between the U.S. and Japanese Stock Markets," *Journal of Finance*, 45 (December 1990), 1297-1306.

Becker, Kent G., Finnerty, Joseph E., and A.L. Tucker, 1992, "The Intraday Dependence Structure Between U.S. and Japanese Equity Markets," *Journal of Financial Research*, 45, 1297-1306.

Bessembinder, H., K. Chan, and P. Seguin, 1996, "An Empirical Examination of Information, Differences of Opinion, and Trading Activity," *Journal of Financial Economics* 40, 105-134.

Bollerslev, Tim, Chou, Ray U., and Kroner, Kenneth F., 1992, "ARCH Modeling in Finance: A Review of the Theory and Empirical Evidence," *Journal of Econometrics*, 15, 27-37.

Calvo, G. A., and Mendoza, E. G. (1999), "Rational Contagion and the Globalization of Securities Markets," National Bureau of Economic Research Working Paper.

Christie, W. and Huang, R., 1995, "Following the Pied Piper: Do Individual Returns Herd around the Market?," *Financial Analysts Journal*, 31-37.

Connolly, R., 1991, "A Posterior Odds Analysis of the Weekend Effect," *Journal of Econometrics*, 49, 51-103.

Connolly, R., 1989, "An Examination of the Robustness of the Weekend Effect," *Journal of Financial and Quantitative Analysis*, 24, 133-171.

Connolly, R. and F.A. Wang, 1998, "International Equity Market Co-Movement: Economic Fundamentals or Contagion?," University of North Carolina – Chapel Hill and Rice University Working Paper.

Cornell, B., 1983, "The Money Supply Announcements Puzzle," *American Economic Review*, 83, 644-657.

Craig, A., Dravid A., and Richardson, M., 1995, "Market Efficiency Around the Clock: Some Supporting Evidence Using Foreign-based Derivatives," *Journal of Financial Economics*, 39, 161-180.

Ederington, Louis H. and Lee, Jae Ha, 1993 "How Markets Process Information: News Releases and Volatility," *Journal of Finance*, 48, 1161-1192.

Engle, Robert F. and Ng, Victor, 1993, "Measuring and Testing the Impact of News on Volatility," *Journal of Finance*, 48, 1749-1778.

Fleming, J, Kirby C., and Ostdiek B., 1998, "Information and Volatility Linkage in the Stock, Bond, and Money Markets," *Journal of Financial Economics*, 49, 111-137.

Forbes, K. and Rigobon, R. (2000), “No Contagion, Only Interdependence: Measuring Stock Market Co-Movements,” MIT Working Paper.

Forbes, K., and Rigobon, R. (1998), “Measuring Contagion: Conceptual and Empirical Issues,” MIT Working Paper.

French, K., 1980, "Stock Returns and the Weekend Effect," *Journal of Financial Economics*, 8, 55-69.

Gibbon, M.R. and P. Hess, 1981, "Day of the Week Effect and Asset Returns" *Journal of Business*, 54, 579-596.

Glosten, L.R., Jagannathan, R., and Runkle, D.E., 1993"On the Relation Between the Expected Value and the Volatility of the Nominal Excess Return on Stocks," *Journal of Finance*, 48, 1779-1802.

Hamao, Y., Masulis, Ronald W., and Ng, V., 1990, "Correlations in Price Changes and Volatility across International Stock Markets," *Review of Financial Studies*, 3, 281-307.

Hardouvelis, G.A., 1987, "Macroeconomic Information and Stock Prices," *Journal of Economics and Business*, 39, 131-140.

Ito, T. and V.V. Roley, 1987, "News from the U.S. and Japan: Which Moves the Yen/Dollar Exchange Rate?" *Journal of Monetary Economics*, 19, 255-277.

Kaminsky, G. L. and Schmukler, S. L. (1999), “What Triggers Market Jitters? A Chronicle of the Asian Crisis,” *Journal of International Money and Finance*, 18, 537 – 560.

Karolyi, G. Andrew and Stulz, René M., 1996, “Why do Markets Move Together? An Investigation of U.S.-Japan Stock Return Co-movements,” *Journal of Finance*, 51, 951-986.

King, Mervyn, Sentana, Enrique, and Wadhwani, Sushil, 1994, “Volatility and Links Between National Stock Markets,” *Econometrica*, 62, 901-934.

King, Mervyn, and Wadhwani, Sushil, 1990, "Transmission of Volatility Between Stock Markets," *The Review of Financial Studies*, 3, 5-33.

Klein, R., and Brown, S., 1984, “Model Selection When There is ‘Minimal’ Prior Information,” *Econometrica*, 52, 1291-1312.

Kodres, L. E. and Pritsker, M. (1999), “A Rational Expectations Model of Financial Contagion,” International Monetary Fund and Federal Reserve Board Working Paper.

Kyle, A. and Xiong W., 1999, “Contagion as a Wealth Effect of Financial Intermediaries,” Duke University Working Paper.

Lamoureux, C. and S. Panikkath, 1994, “Variation in Stock Returns: Asymmetries and Other Patterns,” Working Paper, Washington University in St. Louis.

Leamer, E., 1978, *Specification Searches: Ad Hoc Inference with Non-Experimental Data*, New York: Wiley.

Lin, W., Engle, R.F., and Ito, T., 1994, “Do Bulls and Bears Move Across Borders? International Transmission of Stock Returns and Volatility,” *Review of Financial Studies*, 7, 507-538.

McQueen G. and V.V. Roley, 1993, "Stock Prices, News, and Business Condition," *Review of Financial Studies*, 6, 683-707.

Nelson, Daniel B., 1991, "Conditional Heteroskedasticity in Asset Returns: A New Approach," *Econometrica*, 59, 347-370.

Pearce, D.K. and V.V. Roley, 1985, "Stock Prices and Economic News," *Journal of Business*, 58, 49-67.

Shanken, J., 1987, “A Bayesian Approach to Testing Portfolio Efficiency,” *Journal of Financial Economics*, 18, 195-215.

Solnik, B., 1974a, "The International Pricing of Risk: An Empirical Investigation of the World Capital Market Structure," *Journal of Finance*, 29, 48-54.

Solnik, B., 1974b, "An Equilibrium Model of the International Capital Market," *Journal of Economic Theory*, 8, 500-524.

Stivers, C., 1999, “Differential Information, Cross-sectional Return Dispersion, and Time-Variation in Stock Market Volatility,” Working Paper, University of Georgia.

Stoll, H. and R. Whaley, 1988, "Stock Market Structure and Volatility," *Review of Financial Studies*, 3, 37-71.

Stulz, R.M., 1981, "A Model of International Asset Pricing," *Journal of Financial Economics*, 11, 383-406.

Table 1

Summary Statistics and Correlation Analysis For U.S., U.K., and Japan

Panel A: Summary Statistics for Overnight and Intraday Returns

Market	Return	Mean	Std. Dev.	Min.	Max.	?(1)	?(2)	?(3)	?(12)
U.S.	***SPID***	.0006	.0077	-.163	.045	.012	.006	-.012	-.034
	SPON	.00004	.0054	-.066	.087	-.015	-.204	.045	.026
U.K.	***FTID***	.0005	.0073	-.054	.047	-.047	.059	.022	.002
	FTON	.0001	.0058	-.095	.061	-.059	.076	.054	-.028
Japan	***NKID***	-.0002	.0103	-.130	.083	-.060	.026	.021	.003
	NKON	.0005	.0086	-.045	.070	.042	-.048	.014	.025

Panel B: Summary Statistics for Economic News Announcement

Market	News	Mean	Std. Dev.	Min.	Max.
U.S.	$\mathbf{CPI_{US}}$	-.0057	.139	-2.00	3.00
	$\mathbf{IP_{US}}$	.014	.177	-2.00	3.00
	$\mathbf{PPI_{US}}$	.024	.483	-3.00	4.00
	$\mathbf{UR_{US}}$	-.0005	.0062	-.067	.053
	$\mathbf{TD_{US}}$	-.0003	.041	-.833	.303
	$\mathbf{MS_{US}}$	.078	1.72	-49.0	26.0
U.K.	$\mathbf{CPI_{UK}}$	.0007	.070	-.957	1.10
	$\mathbf{IP_{UK}}$	-.0043	.253	-4.01	3.03
	$\mathbf{UR_{UK}}$	.0005	.038	-1.534	.259
	$\mathbf{MS_{UK}}$	.0005	.099	-1.32	1.75
Japan	$\mathbf{CPI_{JP}}$	.0037	.179	-1.323	4.133
	$\mathbf{PPI_{JP}}$	.001	.093	-1.248	2.47
	$\mathbf{IP_{JP}}$	-.005	.475	-5.827	6.779
	$\mathbf{MS_{JP}}$	.0033	1.454	-.016	.021

Table 1 (cont.)

Summary Statistics and Correlation Analysis For U.S., U.K., and Japan

Panel C: Cross-Market Contemporaneous Return Correlations

Stratification		Correlation of $NKON_t$ with		Correlation of $FTON_t$ with		Correlation of $SPON_t$ with	
		$SPID_{t-1}$	$FTID_{t-1}$	$NKID_t$	$SPID_{t-1}$	$FTID_t$	$NKID_t$
Whole Sample		0.223	0.193	0.240	0.556	0.338	0.090
News	**All News**	0.522	0.352	0.122	0.669	0.400	0.290
	US News	0.262	0.235	0.141	0.470	0.388	0.117
	UK News	0.263	0.315	0.045	0.631	0.390	0.156
	Japan News	0.209	0.296	0.402	0.701	-0.023	-0.267
	No News	0.208	0.148	0.202	0.478	0.421	0.220
Return							
	1% Fractile	0.799	0.265	0.583	0.929	0.294	-0.632
	99% Fractile	0.184	0.196	0.487	0.824	-0.18	0.358
	5% Fractile	0.459	0.263	0.504	0.889	0.294	-0.362
	95% Fractile	0.16	0.133	0.394	0.444	-0.044	0.351
	Interquartile Range	0.193	0.165	-0.009	0.482	0.147	0.024
Return Dispersion							
	1% Fractile	0.15	-0.115	-0.154	0.387	0.655	0.225
	99% Fractile	-0.183	0.255	0.149	-0.332	0.555	0.329
	5% Fractile	0.429	0.13	-0.068	0.566	0.655	0.055
	95% Fractile	0.267	0.389	0.406	0.612	0.162	0.389
	Interquartile Range	0.171	0.185	0.162	0.546	0.377	0.168

Stratification		Tests for Differences in Correlation of $NKON_t$ with		Tests for Differences in Correlation of $FTON_t$ with		Tests for Differences in Correlation of $SPON_t$ with	
		$SPID_{t-1}$	$FTID_{t-1}$	$NKID_t$	$SPID_{t-1}$	$FTID_t$	$NKID_t$
No News vs.	**US News**	-1.26	-1.99**	1.39	0.23	0.87	2.34*
	UK News	-0.97	-2.96*	2.67*	-3.72*	0.62	1.11
	Japan News	-0.02	-2.76*	-3.91*	-6.16*	8.34*	8.79*
	All News	-2.21**	-1.31	0.49	-1.73**	0.15	-0.45
All News vs.	**US News**	1.85**	0.76	-0.11	1.77**	0.08	1.08
	UK News	1.79**	0.24	0.45	0.38	0.07	0.82
	Japan News	2.14**	0.36	-1.77**	-0.35	2.60*	3.33*
Return							
	1% vs. 99%	14.44*	1.74**	2.24*	7.97*	39.60*	-18.56*
	5% vs. 95%	27.55*	11.88*	11.81*	77.39*	30.42*	-63.79*
	5% vs. Inter. Range	34.26*	12.63*	68.70*	101.63*	19.03*	-49.14*
	95% vs. Inter. Range	-4.02*	-4.05*	50.61*	-5.71*	-23.79*	40.74*
Return Dispersion							
	1% vs. 99%	6.78*	-10.52*	-7.03*	17.34*	5.90*	-2.49*
	5% vs. 95%	15.70*	-28.13*	-43.04*	-5.98*	62.36*	-30.55*
	5% vs. Inter. Range	34.30*	-8.32*	-27.99*	3.47*	57.11*	-13.85*
	95% vs. Inter. Range	12.01*	30.31*	32.59*	11.83*	-31.62*	29.14*

Table 1 (cont.)

Summary Statistics and Correlation Analysis For U.S., U.K., and Japan

Panel D: Cross-Market Lead-Lag Return Correlations

Stratification		Correlation of $NKID_t$ with		Correlation of $FTID_t$ with		Correlation of $SPID_t$ with	
		$NKON_t$	$SPID_{t-1}$	$FTON_t$	$SPID_{t-1}$	$SPON_t$	$FTID_t$
Whole Sample		-0.039	0.096	-0.043	-.004	0.086	0.191
News	**All News**	0.284	0.018	-0.138	-0.152	0.238	0.183
	US News	-0.072	-0.008	-0.039	-0.047	0.037	0.197
	UK News	-0.203	0.016	-0.194	-0.196	0.235	0.228
	Japan News	-0.324	0.366	0.074	0.162	-0.058	0.209
	No News	0.051	-0.020	-0.046	-0.021	0.176	0.179
Return							
	1% Fractile	-0.036	0.920	0.188	0.567	0.534	0.388
	99% Fractile	-0.641	-0.050	-0.207	-0.505	-0.224	0.095
	5% Fractile	0.120	0.725	0.074	0.373	0.417	0.381
	95% Fractile	-0.504	-0.087	-0.208	-0.329	-0.209	0.064
	Interquartile Range	0.009	-0.009	0.042	-0.07	-0.015	0.011
Return Dispersion							
	1% Fractile	-0.979	0.016	-0.202	0.144	0.229	0.501
	99% Fractile	0.114	-0.523	0.187	-0.153	-0.147	0.095
	5% Fractile	-0.825	-0.005	-0.128	0.019	0.071	0.410
	95% Fractile	0.339	0.430	0.104	0.241	-0.238	0.192
	Interquartile Range	0.142	0.012	-0.051	-0.044	0.243	0.105

Stratification		Tests for Differences in Correlation of $NKID_t$ with		Tests for Differences in Correlation of $FTID_t$ with		Tests for Differences in Correlation of $SPID_t$ with	
		$NKON_t$	$SPID_{t-1}$	$FTON_t$	$SPID_{t-1}$	$SPON_t$	$FTID_t$
No News vs.	**US News**	2.72*	-0.26	-0.15	0.57	3.11*	-0.41
	UK News	4.30*	-0.60	2.52*	2.97*	-1.03	-0.85
	Japan News	6.84*	-7.13*	-2.12**	-3.26*	4.17*	-0.55
	All News	-1.45	-0.23	0.56	0.79	-0.39	-0.02
All News vs.	**US News**	2.16**	0.15	-0.59	-0.63	1.22	-0.09
	UK News	2.88*	0.01	0.33	0.26	0.02	-0.27
	Japan News	3.66*	-2.13**	-1.24	-1.84**	1.75**	-0.16
Return							
	1% vs. 99%	12.25*	27.11*	6.62*	19.83*	13.66*	7.81*
	5% vs. 95%	58.00*	82.77*	24.80*	60.40*	56.60*	29.56*
	5% vs. Inter. Range	13.38*	105.68*	3.95*	52.67*	55.95*	47.95*
	95% vs. Inter. Range	-68.70*	-9.23*	-30.85*	-32.03*	-23.83*	6.58*
Return Dispersion							
	1% vs. 99%	-52.70*	13.73*	-11.01*	6.89*	8.78*	12.73*
	5% vs. 95%	-131.03*	-39.44*	-23.42*	-19.25*	26.62*	24.24*
	5% vs. Inter. Range	-159.01*	-2.04**	-11.45*	7.56*	-21.21*	48.67*
	95% vs. Inter. Range	25.39*	53.27*	21.08*	34.48*	-58.35*	12.08*

Table 1 (cont.)

Summary Statistics and Correlation Analysis For U.S., U.K., and Japan

Notes: The return and news announcement data in Panels A and B, and the sample stratification and the tests for differences in correlation in Panels C and D are explained as follows.

Returns: We calculate the overnight return in the Japanese market (*NKON*) on day t as ln(Nikkei 225 at 10 a.m. on day t) - ln(Nikkei 225 at close on day t-1). Our intraday return in the Japanese market (*NKID*) on day t is calculated as ln(Nikkei 225 at close on day t) – ln(Nikkei 225 at 10:00 a.m. on day t). The U.S. overnight return (*SPON*) on day t is given by ln(S&P500 at 10 a.m. on day t) - ln(S&P500 at close on day t-1). The U.S. intraday return (*SPID*) on day t is calculated as ln(S&P500 at close on day t) – ln(S&P500 at 10:00 a.m. on day t). We calculate the U.K. overnight return (*FTON*) on day t as ln(FTSE 100 at 9:00 a.m. on day t) - ln(FTSE 100 at close on day t-1). Our U.K. intraday return (*FTID*) on day t is calculated as ln(FTSE 100 at close on day t) – ln(FTSE 100at 9:00 a.m. on day t).

News: Our U.S. economic news variables include money supply (MS), industrial production (IP), consumer price inflation (CPI), producer price inflation (PPI), the unemployment rate (UR), and the merchandise trade balance (TD). The Japanese economic news announcements that we study include money supply (MS), industrial production (IP), consumer price index (CPI), and wholesale price index (PPI). Our macroeconomic news announcements for the U.K. include M3 money supply (MS), retail price inflation (CPI), industrial production (IP), and the unemployment rate (UR). We employ the unexpected component of each announcement in our study, not the raw announcement figures. The unexpected components of the U.S. announcements are calculated as the percentage difference between the actual announcement values and the median expected values, using the actual values as the base. Since we do not have expected values on the macroeconomic announcements for Japan and the U.K., we built a separate ARIMA model for each actual news announcement series. We use residual at time t from an ARIMA model for a specific announcement to measure the surprise in the news announcement at time t.

Correlation: The correlations in the row marked whole sample are based on the entire sample of data for the 1985 – 1996 period. The sample stratifications based on news were made as follows. The all news sample consists of observations for days when there is news in any of the markets. The U.S. (U.K., Japan) news sample includes only observations from days when there are economic news announcements in the U.S. (U.K., Japan). The no-news sample consists of observations on days when there are no announcements in any of the three markets.

The return and return dispersion samples are formed by including observations where the retum in the foreign market involved in the correlation computation (not the home market) satisfies the return or return dispersion criterion. For the Japanese (U.K., U.S.) overnight return correlations, the sample is formed based on the U.S. (Japan, U.K.) intraday returns from the previous day. For example, the 1% return fractile sample used to compute the correlation of $NKON_t$ with $SPID_{t-1}$ consists of observations where the value of $SPID_{t-1}$ fell into the 1% return fractile. The return dispersion samples are formed in a completely analogous fashion.

The tests for differences in the correlation coefficients are a heteroskedasticity-robust t-test studied by Forbes and Rigobon (1998) and implemented by Baig and Goldfajn (1998) from whom this explanation is taken. The null and alternative hypotheses are given by H_0: $r^0_{i,j} \geq r^1_{i,j}$ and H_1: $r^0_{i,j} < r^1_{i,j}$ where the 0 and 1 superscripts indicate the first and second samples, the i and j subscripts indicate the ith and jth countries, and r is the correlation coefficient. The test uses a Fisher transformation of the sample correlation given by $m_t = .5 \cdot \ln[(1+r^t_{i,j})/(1-r^t_{i,j})]$ to compute a test statistic (distributed as a t with m degrees of freedom) given by $U = (m_0 - m_1)/\sqrt{(s_0^2/n_0)+(s_1^2/n_1)}$ where $s_t^2 = (n_t - 3)^{-1}$ is the sample variance after the Fisher transformation. The number of degrees of freedom is given by the integer part of $[(s_0^2/n_0)+(s_1^2/n_1)]^2/[(s_0^2/n_0)/(n_0-1)+(s_1^2/n_1)/(n_1-1)]$. We indicate statistically significant differences in correlations with * (1%), ** (5%), or *** (10%).

Table 2

Estimates of the Baseline Model for U.S., U.K., and Japan, 1985 – 1996

Parameter Estimates	Japan Intraday	Japan Overnight	U.K. Intraday	U.K. Overnight	U.S. Intraday	U.S. Overnight
ρ_1	-0.209 (0.015)*	0.071 (0.010)*	-0.086 (0.020)*	0.119 (0.009)*	-0.065 (0.020)*	0.092 (0.009)*
ρ_2	0.223 (0.011)*	0.223 (0.017)*	0.063 (0.011)*	-0.011 (0.005)**	0.104 (0.017)*	0.202 (0.008)*
ρ_3	0.076 (0.025)*	0.121 (0.018)*	-0.017 (0.020)	0.380 (0.005)*	0.013 (0.013)	0.044 (0.006)*
α	0.000 (0.000)	0.000 (0.000)**	0.001 (0.000)*	0.000 (0.000)	0.000 (0.000)**	0.000 (0.000)
φ	-0.002 (0.001)*	0.001 (0.000)*	0.000 (0.000)	-0.001 (0.000)*	0.001 (0.000)*	0.000 (0.000)
ω_0	0.000 (0.000)*	0.000 (0.000)*	0.000 (0.000)*	0.000 (0.000)*	0.000 (0.000)	0.000 (0.000)*
ω_1	0.076 (0.012)*	0.153 (0.013)*	0.034 (0.012)*	0.193 (0.012)*	0.037 (0.010)*	0.035 (0.014)**
ω_2	0.782 (0.013)*	0.727 (0.016)*	0.789 (0.017)*	0.711 (0.011)*	0.740 (0.018)*	0.735 (0.017)*
ω_3	0.156 (0.022)*	0.210 (0.023)*	0.101 (0.021)*	-0.036 (0.015)**	0.143 (0.012)*	0.194 (0.018)*
ω_4	0.000 (0.000)*	0.000 (0.000)*	0.000 (0.000)**	0.000 (0.000)*	0.000 (0.000)	0.000 (0.000)
ω_5	0.000 (0.000)**	0.000 (0.000)*	0.000 (0.000)*	0.000 (0.000)*	0.000 (0.000)*	0.000 (0.000)*

Notes: In this table, we report estimates of the parameters of the conditional mean and volatility models for Japan, the U.K., and the U.S., respectively, using daily intraday and overnight returns for the 1985 – 1996 period. The conditional mean models are as follows

Japan Intraday: $NKID_t = a + r_1 \cdot NKON_t + r_2 \cdot SPID_{t-1} + r_3 \cdot FTID_{t-1} + j \cdot MHD_{JP,t} + e^{ID}_{JP,t}$,

Japan Overnight: $NKON_t = a + r_1 \cdot NKID_{t-1} + r_2 \cdot SPID_{t-1} + r_3 \cdot FTID_{t-1} + j \cdot MHD_{JP,t} + e^{ON}_{JP,t}$,

UK Intraday: $FTID_t = a + r_1 \cdot FTON_t + r_2 \cdot NKID_t + r_3 \cdot SPID_{t-1} + j \cdot MHD_{UK,t} + e^{ID}_{UK,t}$

UK Overnight: $FTON_t = a + r_1 \cdot FTID_{t-1} + r_2 \cdot NKID_t + r_3 \cdot SPID_{t-1} + j \cdot MHD_{UK,t} + e^{ON}_{UK,t}$

US Intraday: $SPID_t = a + r_1 \cdot SPON_t + r_2 \cdot FTID_t + r_3 \cdot NKID_t + j \cdot MHD_{US,t} + e^{ID}_{US,t}$,

US Overnight: $SPON_t = a + r_1 \cdot SPID_{t-1} + r_2 \cdot FTID_t + r_3 \cdot NKID_t + j \cdot MHD_{US,t} + e^{ON}_{US,t}$,

Table 2 (cont.)

Estimates of the Baseline Model for U.S., U.K., and Japan, 1985 – 1996

Standard errors, indicated in parentheses on the line beside the point estimates, are computed from a covariance matrix that accounts for conditional volatility using the Glosten-Jagannathan-Runkle (GJR) asymmetric GARCH model. Specifically, the conditional volatility models for intraday and overnight returns are given by

Intraday Volatility: $e_{i,t}^{ID} \sim N(0, h_{i,t}^{ID}),\ \ h_{i,t}^{ID} = w_0 + w_1 \cdot e_{i,t-1}^{2(ID)} + w_2 \cdot h_{i,t-1}^{ID} + w_3 \cdot e_{i,t-1}^{2-(ID)} + w_4 \cdot MHD_{i,t} + w_5 \cdot INT_{i,t}$

Overnight Volatility: $e_{i,t}^{ON} \sim N(0, h_{i,t}^{ON}),\ \ h_{i,t}^{ON} = w_0 + w_1 \cdot e_{i,t-1}^{2(ON)} + w_2 \cdot h_{i,t-1}^{ON} + w_3 \cdot e_{i,t-1}^{2-(ON)} + w_4 \cdot MHD_{i,t} + w_5 \cdot INT_{i,t}$ (i = JP, UK, US)

We measure stock returns as follows. We calculate the overnight return in the Japanese market (*NKON*) on day t as ln(Nikkei 225 at 10 a.m. on day t) - ln(Nikkei 225 at close on day t-1). Our intraday return in the Japanese market (*NKID*) on day t is calculated as ln(Nikkei 225 at close on day t) – ln(Nikkei 225 at 10:00 a.m. on day t). The U.S. overnight return (*SPON*) on day t is given by ln(S&P500 at 10 a.m. on day t) - ln(S&P500 at close on day t-1). The U.S. intraday return (*SPID*) on day t is calculated as ln(S&P500 at close on day t) – ln(S&P500 at 10:00 a.m. on day t). We calculate the U.K. overnight return (*FTON*) on day t as ln(FTSE 100 at 9:00 a.m. on day t) - ln(FTSE 100 at close on day t-1). Our U.K. intraday return (*FTID*) on day t is calculated as ln(FTSE 100 at close on day t) – ln(FTSE 100at 9:00 a.m. on day t).

For the interest rate term (INT) that is included in our conditional volatility model, we use the one-month Eurodollar interest rate for the U.S., the one-month Euro-yen interest rate for Japan, and the one-month Euro-sterling interest rate for the U.K. The underlying data are afternoon observations taken in London. We adjust for timing variations so that interest rate observations precede the last equity index value used in calculating equity returns. MHD is a dummy variable taking a value of one when an observation falls on Monday or follows a holiday.

See Sections II and III for more detail about the model and data. We indicate statistical significance of parameter estimates using the following notation: *: 1 per cent significance; **: 5 per cent significance; ***: 10 per cent significance.

Table 3

Estimates of Linear News Model for U.S., U.K., and Japan, 1985 - 1996

Parameter Estimates	Japan		U.K.		U.S.	
	Intraday	**Overnight**	**Intraday**	**Overnight**	**Intraday**	**Overnight**
ρ_1	-.148 (.015)*	.062 (.011)*	-.071 (.025)*	.122 (.008)*	-.058 (.023)*	-.016 (.010)
ρ_2	.236 (.012)*	.232 (.017)*	.066 (.012)*	.060 (.004)*	.105 (.018)*	.198 (.009)*
ρ_3	.065 (.024)	.121 (.018)*	.002 (.021)	.365 (.006)*	.015 (.014)	.030 (.006)*
β_{11}	.0002 (.001)	-.0004 (.0006)	.0002 (.0007)	.0001 (.0002)	-.0003 (.001)	-.0004 (.0004)
β_{12}	.002 (.002)	-.001 (.001)	-.0003 (.006)	.002 (.002)	-.0003 (.0006)	-.0001 (.0003)
β_{13}	-.001 (.0004)	.001 (.0002)*	.0004 (.003)	-.001 (.001)	-.0008 (.001)	-.0001 (.001)
β_{14}	-.007 (.112)	.040 (.037)	.006 (.004)***	-.0005 (.002)	.039 (.014)*	-.002 (.007)
β_{15}					.006 (.003)*	-.0006 (.001)
β_{16}					-.050 (.095)	-.001 (.009)
β_{21}	.001 (.001)***	-.003 (.0004)*	.0002 (.0004)	.00003 (.0001)	.0002 (.0007)	-.0002 (.0003)
β_{22}	.001 (.001)	.0002 (.0007)	.002 (.002)	.0001 (.001)	-.001 (.005)	-.003 (.003)
β_{23}	.001 (.001)	.001 (.0005)*	.0003 (.0003)	.0002 (.0001)	-.00003 (.003)	.001 (.002)
β_{24}	.001 (.030)	.023 (.024)	-.146 (.100)	-.070 (.038)***	-.001 (.003)	.001 (.002)
β_{25}	.004 (.004)	-.003 (.003)				
β_{26}	-.00001 (.00001)	-.0001 (.0001)				
β_{31}	-.001 (.001)	.0003 (.0006)	-.000 (.001)	-.001 (.001)	.0005 (.0001)*	.036 (.097)
β_{32}	.006 (.003)**	-.005 (.002)*	.000 (.001)	-.0004 (.0004)	-.002 (.002)	.001 (.001)
β_{33}	-.008 (.001)*	.008 (.0005)*	-.0002 (.001)	.00005 (.0003)	.004 (.304)	.0002 (.0002)
β_{34}	.005 (.008)	.003 (.004)	-.053 (.022)*	-.036 (.008)*	-.092 (.110)	-.041 (.050)
β_{35}			-.001 (.002)	-.0005 (.001)		
β_{36}			-.0260 (.050)	.00001 (.00003)		
α	-.0001 (.0002)	.0004 (.0001)*	.0005 (.0001)*	.023 (.071)	.0004 (.0001)*	.028 (.085)
φ	-.002 (.000)*	-.001 (.0003)*	-.000 (.000)	-.001 (.000)*	.001 (.000)**	-.000 (.000)

Notes: We report estimates of the parameters of the conditional mean models for Japan, the U.K., and the U.S., respectively, using daily intraday and overnight returns for the 1985 – 1996 period. To save space, we show the conditional mean models only for intraday returns as follows:

JP Intraday: $$NKID_t = a + r_1 \cdot NKON_t + r_2 \cdot SPID_{t-1} + r_3 \cdot FTID_{t-1} + b_1 \cdot NEWS_{JP,t} + b_2 \cdot NEWS_{US,t-1} + b_3 \cdot NEWS_{UK,t-1} + j \cdot MHD_{JP,t} + e^{ID}_{JP,t},$$

Table 3 (cont.)

Estimates of Linear News Model for U.S., U.K., and Japan, 1985 - 1996

UK Intraday: $FTID_t = a + r_1 \cdot FTON_t + r_2 \cdot NKID_t + r_3 \cdot SPID_{t-1} + b_1 \cdot NEWS_{UK,t} + b_2 \cdot NEWS_{JP,t} + b_3 \cdot NEWS_{US,t} + j \cdot MHD_{UK,t} + e^{ID}_{UK,t}$

US Intraday: $SPID_t = a + r_1 \cdot SPON_t + r_2 \cdot FTID_t + r_3 \cdot NKID_t + b_1 \cdot NEWS_{US,t} + b_2 \cdot NEWS_{UK,t} + b_3 \cdot NEWS_{JP,t} + j \cdot MHD_{US,t} + e^{ID}_{US,t}$,

Standard errors, indicated in parentheses on the line beside the point estimates, are computed from a covariance matrix that accounts for conditional volatility using the Glosten-Jagannathan-Runkle (GJR) asymmetric GARCH model. Specifically, the conditional volatility models for intraday returns is given by

Intraday Volatility: $e^{ID}_{i,t} \sim N(0, h^{ID}_{i,t})$, $h^{ID}_{i,t} = w_0 + w_1 \cdot e^{2(ID)}_{i,t-1} + w_2 \cdot h^{ID}_{i,t-1} + w_3 \cdot e^{2-(ID)}_{i,t-1} + w_4 \cdot MHD_{i,t} + w_5 \cdot INT_{i,t}$ (i = JP, UK, US).

We have suppressed reporting the estimates of the coefficients of the volatility models (i.e., w_0, w_1, w_2, w_3, w_4 and w_5) to save space.

We calculate the overnight return in the Japanese market (*NKON*) on day t as ln(Nikkei 225 at 10 a.m. on day t) - ln(Nikkei 225 at close on day t-1). Our intraday return in the Japanese market (*NKID*) on day t is calculated as ln(Nikkei 225 at close on day t) – ln(Nikkei 225 at 10:00 a.m. on day t). The U.S. overnight return (*SPON*) on day t is given by ln(S&P500 at 10 a.m. on day t) - ln(S&P500 at close on day t-1). The U.S. intraday return (*SPID*) on day t is calculated as ln(S&P500 at close on day t) – ln(S&P500 at 10:00 a.m. on day t). We calculate the U.K. overnight return (*FTON*) on day t as ln(FTSE 100 at 9:00 a.m. on day t) - ln(FTSE 100 at close on day t-1). Our U.K. intraday return (*FTID*) on day t is calculated as ln(FTSE 100 at close on day t) – ln(FTSE 100at 9:00 a.m. on day t).

We calculate news announcements as follows. We obtain the U.S. actual announcement data and the median survey estimates from MMS International (Money Market Services). We collected the Japan monthly *preliminary* announcement data by hand from the *Nihon Keizai Shinbun* for the 1985-1994 period. The Bank of Japan provided some additional announcement data for 1995 and 1996. The U.K. announcement data are collected from the *Financial Times*. To capture the news in each announcement, we use the unexpected component of the announcement figures. For the U.S., this is the percentage difference between the actual announcement values and the median expected values, using the actual values as the base. For Japan and the U.K., we use the residual at time t from an ARIMA model for a specific announcement to measure the unexpected value of the news announcement at time t.

The β_{ij} news announcement parameter estimates may be matched against specific announcements in the following way. The i subscript refers the country where the announcement occurred and the j subscript refers to the specific announcement. For Japan, the β_{1j} are coefficients on the ordered values of the Japanese CPI, PPI, IP, and MS announcements (j = 1, 2, 3, 4). For example, β_{13} is the coefficient on the Japanese industrial production announcement. The β_{2j} are coefficients on U.S. news announcements using the order CPI, PPI, IP, UR, TD, and MS (j = 1,…,6). The β_{3j} are coefficients on U.K. news announcements using the order IP, CPI, MS, and UR (j = 1,…, 4). Using the news announcement order given below, the same procedure is used to identify individual coefficients for the U.K. and U.S. regressions. For example, in the U.S. intraday regression, β_{15} is the coefficient on the U.S. trade deficit announcement.

For the interest rate term (INT), we use the one-month Eurodollar interest rate for the U.S., the one-month Euro-yen interest rate for Japan, and the one-month Euro-sterling interest rate for the U.K. MHD is a dummy variable taking a value of one when an observation falls on Monday or follows a holiday.

See Sections II and III for more detail about the model and data. We indicate statistical significance of parameter estimates using the following notation: *: 1 per cent significance; **: 5 per cent significance; ***: 10 per cent significance.

Table 4

Estimates of Nonlinear News Model for U.S., U.K., and Japan, 1985 - 1996

Parameter Estimates	Japan Intraday	Japan Overnight	U.K. Intraday	U.K. Overnight	U.S. Intraday	U.S. Overnight
ρ_1	-0.115 (0.016)*	0.061 (0.012)*	-0.064 (0.021)*	0.109 (0.009)*	-0.035 (0.028)	0.093 (0.010)*
β_{11}	0.327 (0.074)*	-0.016 (0.164)	-0.169 (0.036)*	-0.006 (0.043)	0.007 (0.255)	-0.150 (0.136)
β_{12}	-0.440 (0.310)	-0.011 (0.192)	-0.235 (1.034)	-0.465 (0.184)**	-0.044 (0.175)	-0.060 (0.070)
β_{13}	-0.241 (0.013)*	-0.007 (0.031)	0.251 (0.790)	0.308 (0.095)*	0.327 (0.249)	-0.009 (0.078)
β_{14}	107.025 (5.679)*	-37.815 (6.495)*	1.220 (1.089)	0.053 (0.524)	-19.510 (1.536)*	-1.684 (1.474)
β_{15}					-0.172 (0.397)	-0.173 (0.258)
β_{16}					-0.002 (0.014)	-0.009 (0.007)
ρ_2	0.209 (0.012)*	0.237 (0.016)*	0.068 (0.013)*	-0.015 (0.006)**	0.094 (0.018)*	0.204 (0.009)*
β_{21}	0.385 (0.343)	-0.717 (0.071)*	-0.018 (0.077)	-0.037 (0.038)	-0.084 (0.093)	-0.040 (0.030)
β_{22}	0.115 (0.129)	0.141 (0.106)	0.024 (0.289)	-0.147 (0.077)***	-0.118 (0.425)	-0.348 (0.206)***
β_{23}	0.219 (0.129)***	-0.085 (0.085)	-0.056 (0.029)***	-0.025 (0.013)***	0.230 (0.362)	0.289 (0.160)***
β_{24}	4.655 (4.603)	1.483 (2.959)	-7.517 (11.246)	2.758 (4.667)	0.339 (0.934)	0.237 (0.395)
β_{25}	-1.649 (0.743)**	-0.312 (0.493)				
β_{26}	0.008 (0.033)	-0.051 (0.013)*				
ρ_3	0.076 (0.025)*	0.118 (0.017)*	-0.023 (0.020)	0.370 (0.007)*	0.014 (0.013)	0.044 (0.006)*
β_{31}	0.039 (0.057)	-0.042 (0.064)	-0.033 (0.271)	0.532 (0.056)*	0.012 (0.075)	-0.040 (0.032)
β_{32}	0.338 (0.538)	0.844 (0.258)*	0.214 (0.046)*	0.154 (0.044)*	-0.058 (0.200)	-0.136 (0.060)**
β_{33}	-0.863 (0.088)*	-1.006 (0.074)*	0.092 (0.072)	-0.069 (0.080)	0.000 (0.051)	0.001 (0.017)
β_{34}	1.053 (0.400)*	0.071 (0.380)	6.140 (3.516)***	-2.318 (0.484)*	7.442 (7.312)	0.281 (5.680)
β_{35}			0.676 (0.259)*	-0.333 (0.200)***		
β_{36}			0.014 (0.015)	-0.010 (0.010)		

Notes: We report estimates of the parameters of the conditional mean models for Japan, the U.K., and the U.S., respectively, using daily intraday and overnight returns for the 1985 – 1996 period. To save space, we show the conditional mean models only for intraday returns as follows:

Japan Intraday: $NKID_t = a + (r_1 + b_1 \cdot NEWS_{JP,t}) \cdot NKON_t + (r_2 + b_2 \cdot NEWS_{US,t-1}) \cdot SPID_{t-1} + (r_3 + b_3 \cdot NEWS_{UK,t-1}) \cdot FTID_{t-1} + j \cdot MHD_{JP,t} + e^{ID}_{JP,t}$,

Table 4 (cont.)

Estimates of Nonlinear News Model for U.S., U.K., and Japan, 1985 - 1996

UK Intraday: $FTID_t = a + (r_1 + b_1 \cdot NEWS_{UK,t}) \cdot FTON_t + (r_2 + b_2 \cdot NEWS_{JP,t}) \cdot NKID_t + (r_3 + b_3 \cdot NEWS_{US,t}) \cdot SPID_{t-1} + j \cdot MHD_{UK,t} + e^{ID}_{UK,t}$

US Intraday: $SPID_t = a + (r_1 + b_1 \cdot NEWS_{US,t}) \cdot SPON_t + (r_2 + b_2 \cdot NEWS_{UK,t}) \cdot FTID_t + (r_3 + b_3 \cdot NEWS_{JP,t}) \cdot NKID_t + j \cdot MHD_{US,t} + e^{ID}_{US,t}$,

Standard errors, indicated in parentheses on the line beside the point estimates, are computed from a covariance matrix that accounts for conditional volatility using the Glosten-Jagannathan-Runkle (GJR) asymmetric GARCH model. Specifically, the conditional volatility models for intraday returns is given by

Intraday Volatility: $e^{ID}_{i,t} \sim N(0, h^{ID}_{i,t}),\ \ h^{ID}_{i,t} = w_0 + w_1 \cdot e^{2(ID)}_{i,t-1} + w_2 \cdot h^{ID}_{i,t-1} + w_3 \cdot e^{2-(ID)}_{i,t-1} + w_4 \cdot MHD_{i,t} + w_5 \cdot INT_{i,t}$ (i = JP, UK, US).

We have suppressed reporting the estimates of the coefficients $a,\ j,\ w_0,\ w_1,\ w_2,\ w_3,\ w_4$ and w_5 to save space.

We calculate the overnight return in the Japanese market (*NKON*) on day t as ln(Nikkei 225 at 10 a.m. on day t) - ln(Nikkei 225 at close on day t-1). Our intraday return in the Japanese market (*NKID*) on day t is calculated as ln(Nikkei 225 at close on day t) – ln(Nikkei 225 at 10:00 a.m. on day t). The U.S. overnight return (*SPON*) on day t is given by ln(S&P500 at 10 a.m. on day t) - ln(S&P500 at close on day t-1). The U.S. intraday return (*SPID*) on day t is calculated as ln(S&P500 at close on day t) – ln(S&P500 at 10:00 a.m. on day t). We calculate the U.K. overnight return (*FTON*) on day t as ln(FTSE 100 at 9:00 a.m. on day t) - ln(FTSE 100 at close on day t-1). Our U.K. intraday return (*FTID*) on day t is calculated as ln(FTSE 100 at close on day t) – ln(FTSE 100at 9:00 a.m. on day t).

We calculate news announcements as follows. We obtain the U.S. actual announcement data and the median survey estimates from MMS International (Money Market Services). We collected the Japan monthly *preliminary* announcement data by hand from the *Nihon Keizai Shinbun* for the 1985-1994 period. The Bank of Japan provided some additional announcement data for 1995 and 1996. The U.K. announcement data are collected from the *Financial Times*. To capture the news in each announcement, we use the unexpected component of the announcement figures. For the U.S., this is the percentage difference between the actual announcement values and the median expected values, using the actual values as the base. For Japan and the U.K., we use the residual at time t from the final ARIMA model for a specific announcement to measure the unexpected value of the news announcement at time t.

The β_{ij} news announcement parameter estimates may be matched against specific announcements in the following way. The i subscript refers the country where the announcement occurred and the j subscript refers to the specific announcement. For Japan, the β_{1j} are coefficients on the ordered values of the Japanese CPI, PPI, IP, and MS announcements (j = 1, 2, 3, 4). For example, β_{13} is the coefficient on the Japanese industrial production announcement. The β_{2j} are coefficients on U.S. news announcements using the order CPI, PPI, IP, UR, TD, and MS (j = 1,…,6). The β_{3j} are coefficients on U.K. news announcements using the order IP, CPI, MS, and UR (j = 1,…, 4). Using the news announcement order given below, the same procedure is used to identify individual coefficients for the U.K. and U.S. regressions. For example, in the U.S. intraday regression, β_{15} is the coefficient on the U.S. trade deficit announcement.

For the interest rate term (INT), we use the one-month Eurodollar interest rate for the U.S., the one-month Euro-yen interest rate for Japan, and the one-month Euro-sterling interest rate for the U.K. MHD is a dummy variable taking a value of one when an observation falls on Monday or follows a holiday.

See Sections II and III for more detail about the model and data. We indicate statistical significance of parameter estimates using the following notation: *: 1 per cent significance; **: 5 per cent significance; ***: 10 per cent significance.

Table 5

Estimates of Nonlinear News Model with Dispersion in Beliefs for U.S., U.K., and Japan, 1985 – 1996

Parameter Estimates	Japan		U.K.		U.S.	
	Intraday	Overnight	Intraday	Overnight	Intraday	Overnight
ρ_1	-0.157 (0.016)*	0.068 (0.014)*	-0.057 (0.022)*	0.128 (0.010)*	-0.028 (0.028)	0.091 (0.010)*
β_{11}	0.330 (0.082)*	-0.045 (0.175)	-0.174 (0.037)*	0.194 (0.017)*	0.019 (0.258)	-0.163 (0.120)
β_{12}	-0.343 (0.291)	0.219 (0.193)	-0.241 (1.050)	-0.322 (0.220)	-0.038 (0.171)	-0.057 (0.069)
β_{13}	-0.229 (0.014)*	0.021 (0.028)	0.294 (0.798)	0.355 (0.113)*	0.334 (0.251)	-0.005 (0.079)
β_{14}	93.404 (5.535)*	-23.084(10.819)**	1.184 (1.088)	-0.471 (0.184)**	-19.851 (1.535)*	-1.938 (1.442)
β_{15}					-0.170 (0.403)	-0.175 (0.253)
β_{16}					-0.001 (0.014)	-0.005 (0.007)
ρ_2	0.130 (0.018)*	0.219 (0.016)*	0.066 (0.013)*	-0.017 (0.007)**	0.093 (0.018)*	0.197 (0.009)*
β_{21}	0.345 (0.324)	-0.710 (0.073)*	-0.024 (0.077)	-0.050 (0.029)***	-0.087 (0.093)	-0.029 (0.030)
β_{22}	0.218 (0.116)***	0.134 (0.113)	0.033 (0.285)	-0.154 (0.079)***	-0.130 (0.421)	-0.278 (0.205)
β_{23}	0.224 (0.133)***	-0.063 (0.089)	-0.061 (0.025)**	-0.004 (0.012)	0.224 (0.360)	0.247 (0.165)
β_{24}	2.247 (5.474)	1.530 (2.665)	-6.711(10.289)	6.768 (4.642)	0.289 (0.953)	0.379 (0.388)
β_{25}	-1.186 (0.641)***	-0.325 (0.523)				
β_{26}	0.011 (0.036)	-0.036 (0.017)**				
δ_2	29.596 (2.022)*	-9.147 (2.160)*	-0.022 (0.020)	-0.031 (0.007)*	-0.023 (0.029)	-0.028 (0.014)**
ρ_3	0.058 (0.024)**	0.119 (0.019)*	-0.025 (0.020)	0.370 (0.006)*	0.008 (0.013)	0.054 (0.007)*
β_{31}	-0.064 (0.044)	-0.018 (0.058)	-0.034 (0.270)	0.097 (0.198)	0.005 (0.074)	-0.005 (0.041)
β_{32}	0.411 (0.500)	0.955 (0.275)*	0.209 (0.046)*	0.159 (0.041)*	-0.061 (0.199)	-0.155 (0.064)**
β_{33}	-0.909 (0.079)*	-1.103 (0.090)*	0.091 (0.076)	-0.010 (0.079)	-0.012 (0.046)	0.026 (0.014)***
β_{34}	0.792 (0.462)***	0.075 (0.392)	6.598 (3.583)***	-1.651 (1.129)	8.438 (8.846)	-3.771 (4.829)
β_{35}			0.690 (0.261)*	-0.634 (0.225)*		
β_{36}			0.007 (0.014)	-0.011 (0.015)		
δ_3	0.006 (0.045)	0.017 (0.036)	0.759 (3.953)	8.744 (1.006)*	-0.037 (0.020)***	0.056 (0.008)*

Notes: We report estimates of the parameters of the conditional mean models for Japan, the U.K., and the U.S., respectively, using daily intraday and overnight returns for the 1985 – 1996 period. To save space, we show the conditional mean models only for intraday returns as follows:

Table 5 (Cont.)

Estimates of Nonlinear News Model with Dispersion in Beliefs for U.S., U.K., and Japan, 1985 – 1996

Japan: $NKID_t = a + (r_1 + b_1 \cdot NEWS_{JP,t}) \cdot NKON_t + (r_2 + b_2 \cdot NEWS_{US,t-1} + d_2 \cdot RRD_{US,t-1}) \cdot SPID_{t-1} + (r_3 + b_3 \cdot NEWS_{UK,t-1} + d_3 \cdot RRD_{UK,t-1}) \cdot FTID_{t-1} + j \cdot MHD_{JP,t} + e^{ID}_{JP,t}$,

U.K.: $FTID_t = a + (r_1 + b_1 \cdot NEWS_{UK,t}) \cdot FTON_t + (r_2 + b_2 \cdot NEWS_{JP,t} + d_2 \cdot RRD_{JP,t}) \cdot NKID_t + (r_3 + b_3 \cdot NEWS_{US,t} + d_3 \cdot RRD_{US,t-1}) \cdot SPID_{t-1} + j \cdot MHD_{UK,t} + e^{ID}_{UK,t}$

US: $SPID_t = a + (r_1 + b_1 \cdot NEWS_{US,t}) \cdot SPON_t + (r_2 + b_2 \cdot NEWS_{UK,t} + d_2 \cdot RRD_{UK,t}) \cdot FTID_t + (r_3 + b_3 \cdot NEWS_{JP,t} + d_2 \cdot RRD_{JP,t}) \cdot NKID_t + j \cdot MHD_{US,t} + e^{ID}_{US,t}$,

Standard errors, indicated in parentheses on the line beside the point estimates, are computed from a covariance matrix that accounts for conditional volatility using the Glosten-Jagannathan-Runkle (GJR) asymmetric GARCH model. Specifically, the conditional volatility models for intraday returns is given by

Intraday Volatility: $e^{ID}_{i,t} \sim N(0, h^{ID}_{i,t}),\ h^{ID}_{i,t} = w_0 + w_1 \cdot e^{2(ID)}_{i,t-1} + w_2 \cdot h^{ID}_{i,t-1} + w_3 \cdot e^{2-(ID)}_{i,t-1} + w_4 \cdot MHD_{i,t} + w_5 \cdot INT_{i,t}$ (i = JP, UK, US).

We have suppressed reporting the estimates of the coefficients a, j, w_0, w_1, w_2, w_3, w_4 and w_5 to save space.

We define the stock market's cross-sectional RD in period t as: $RD_{M,t} = \sqrt{\frac{1}{n-1}\sum_{i=1}^{n}(R_{i,t} - R_{M,t})^2}$,

where n is the number of firms in the market, and subscript M indicates market-wide values. To implement our empirical tests, we measure RD using daily returns for the U.S. firms in the largest capitalization decile of the NYSE-AMEX, but we use the Nikkei 225 for Japan and the FTSE-100 for the U.K. Our relative RD measure, denoted RRD, is the residual, u_t, from a regression of RD on a function of the aggregate excess market return:

$$RD_t = f_0 + g_1 \left|R^e_t\right| + g_2 Dum1_t \left|R^e_t\right| + u_t$$

where R^e_t is the excess return of the market-portfolio return for period t using Euro-currency interest rates as the risk-free return, $Dum1_t = 1$ if the portfolio excess return is negative, and is 0 otherwise.

The β_{ij} news announcement parameter estimates may be matched against specific announcements in the following way. The i subscript refers the country where the announcement occurred and the j subscript refers to the specific announcement. For Japan, the β_{1j} are coefficients on the ordered values of the Japanese CPI, PPI, IP, and MS announcements (j = 1, 2, 3, 4). For example, β_{13} is the coefficient on the Japanese industrial production announcement. The β_{2j} are coefficients on U.S. news announcements using the order CPI, PPI, IP, UR, TD, and MS (j = 1,…,6). The β_{3j} are coefficients on U.K. news announcements using the order IP, CPI, MS, and UR (j = 1,…, 4). Using the news announcement order given below, the same procedure is used to identify individual coefficients for the U.K. and U.S. regressions. For example, in the U.S. intraday regression, β_{15} is the coefficient on the U.S. trade deficit announcement.

See Sections II and III for more detail about the model and data. We indicate statistical significance of parameter estimates using the following notation: *: 1 per cent significance; **: 5 per cent significance; ***: 10 per cent significance.

Table 6

Estimates of Nonlinear News Models with Extreme Returns for U.S., U.K., and Japan, 1985 – 1996

Parameter	Japan		U.K.		U.S.	
Estimates	Intraday	Overnight	Intraday	Overnight	Intraday	Overnight
ρ_1	-0.154 (0.015)*	0.062 (0.013)*	-0.066 (0.023)*	0.128 (0.010)*	-0.028 (0.028)	0.079 (0.010)*
β_{11}	0.333 (0.081)*	-0.012 (0.179)	-0.198 (0.033)*	0.053 (0.045)	0.011 (0.255)	0.217 (0.118)***
β_{12}	-0.411 (0.277)	0.212 (0.188)	-0.317 (1.033)	-0.232 (0.256)	-0.026 (0.169)	-0.077 (0.062)
β_{13}	-0.233 (0.014)*	0.021 (0.028)	0.321 (0.784)	0.384 (0.137)*	0.305 (0.249)	0.018 (0.089)
β_{14}	89.060 (5.755)*	-33.927 (7.071)*	0.905 (1.046)	-0.108 (0.341)	-20.337 (1.731)*	-0.468 (1.602)
β_{15}					-0.178 (0.393)	-0.018 (0.271)
β_{16}					0.001 (0.020)	0.006 (0.009)
ρ_2	0.027 (0.029)	0.227 (0.022)*	0.070 (0.017)*	-0.022 (0.008)*	0.094 (0.018)*	0.208 (0.009)*
β_{21}	0.527 (0.257)**	-0.686 (0.073)*	-0.041 (0.080)	0.115 (0.014)*	-0.079 (0.088)	-0.087 (0.023)*
β_{22}	0.096 (0.088)	0.127 (0.111)	0.035 (0.277)	-0.148 (0.090)***	-0.157 (0.419)	-0.224 (0.234)
β_{23}	0.250 (0.148)***	-0.026 (0.094)	-0.034 (0.041)	-0.014 (0.014)	0.243 (0.358)	0.289 (0.172)***
β_{24}	7.759 (3.397)**	1.952 (3.441)	-11.378(12.263)	4.595 (4.480)	0.357 (0.923)	-0.005 (0.467)
β_{25}	-1.770 (0.523)*	-0.373 (0.578)				
β_{26}	0.011 (0.034)	-0.038 (0.017)**				
η_{21}	0.381 (0.032)*	0.065 (0.027)**	-0.065 (0.041)	0.197 (0.017)*	0.837 (0.323)*	-0.623 (0.086)*
η_{22}	-0.146 (0.067)**	-0.058 (0.054)	0.063 (0.034)***	-0.116 (0.016)*	-0.093 (0.212)	0.113 (0.112)
ρ_3	0.054 (0.024)**	0.113 (0.021)*	-0.025 (0.025)	0.365 (0.011)*	0.025 (0.016)	0.057 (0.008)*
β_{31}	0.126 (0.070)***	-0.052 (0.071)	-0.031 (0.290)	0.135 (0.170)	-0.009 (0.081)	-0.021 (0.038)
β_{32}	0.683 (0.500)	0.963 (0.269)*	0.185 (0.054)*	0.164 (0.047)*	-0.061 (0.198)	-0.366 (0.056)*
β_{33}	-0.930 (0.080)*	-1.072 (0.082)*	0.061 (0.075)	-0.033 (0.084)	0.016 (0.045)	0.041 (0.007)*
β_{34}	0.895 (0.448)**	0.077 (0.381)	6.939 (3.700)***	-0.035 (0.946)	3.770 (7.398)	3.725 (8.173)
β_{35}			0.710 (0.268)*	-0.330 (0.191)***		
β_{36}			0.005 (0.014)	-0.007 (0.014)		
η_{31}	0.146 (0.063)**	-0.018 (0.048)	0.091 (0.037)**	0.118 (0.016)*	-0.022 (0.032)	0.001 (0.020)
η_{32}	-0.072 (0.138)	0.078 (0.082)	-0.085 (0.055)	-0.167 (0.022)*	-0.057 (0.040)	-0.024 (0.014)***

Table 6 (Cont.)

Estimates of Nonlinear News Models with Extreme Returns for U.S., U.K., and Japan, 1985 – 1996

Notes: We report estimates of the parameters of the conditional mean models for Japan, the U.K., and the U.S., respectively, using daily intraday and overnight returns for the 1985 – 1996 period. To save space, we show the conditional mean models only for intraday returns as follows:

Japan:
$$NKID_t = a + (r_1 + b_1 \cdot NEWS_{JP,t}) \cdot NKON_t + (r_2 + b_2 \cdot NEWS_{US,t-1} + h_{21} \cdot D^L_{US,t-1} + h_{22} \cdot D^U_{US,t-1}) \cdot SPID_{t-1}$$
$$+ (r_3 + b_3 \cdot NEWS_{UK,t-1} + h_{31} \cdot D^L_{UK,t-1} + h_{32} \cdot D^U_{UK,t-1}) \cdot FTID_{t-1} + j \cdot MHD_{JP,t} + e^{ID}_{JP,t}$$

UK:
$$FTID_t = a + (r_1 + b_1 \cdot NEWS_{UK,t}) \cdot FTON_t + (r_2 + b_2 \cdot NEWS_{JP,t} + h_{21} \cdot D^L_{JP,t} + h_{22} \cdot D^U_{JP,t}) \cdot NKID_t$$
$$+ (r_3 + b_3 \cdot NEWS_{US,t} + h_{31} \cdot D^L_{US,t-1} + h_{32} \cdot D^U_{US,t-1}) \cdot SPID_{t-1} + j \cdot MHD_{UK,t} + e^{ID}_{UK,t}$$

US:
$$SPID_t = a + (r_1 + b_1 \cdot NEWS_{US,t}) \cdot SPON_t + (r_2 + b_2 \cdot NEWS_{UK,t} + h_{21} \cdot D^L_{UK,t} + h_{22} \cdot D^U_{UK,t}) \cdot FTID_t$$
$$+ (r_3 + b_3 \cdot NEWS_{JP,t} + h_{31} \cdot D^L_{JP,t} + h_{32} \cdot D^U_{JP,t}) \cdot NKID_t + j \cdot MHD_{US,t} + e^{ID}_{US,t}$$

Standard errors, indicated in parentheses on the line beside the point estimates, are computed from a covariance matrix that accounts for conditional volatility using the Glosten-Jagannathan-Runkle (GJR) asymmetric GARCH model. Specifically, the conditional volatility models for intraday returns is given by

Intraday Volatility: $e^{ID}_{i,t} \sim N(0, h^{ID}_{i,t}),\ h^{ID}_{i,t} = w_0 + w_1 \cdot e^{2(ID)}_{i,t-1} + w_2 \cdot h^{ID}_{i,t-1} + w_3 \cdot e^{2-(ID)}_{i,t-1} + w_4 \cdot MHD_{i,t} + w_5 \cdot INT_{i,t}$ (i = JP, UK, US).

We have suppressed reporting the estimates of the coefficients $a,\ j,\ w_0,\ w_1,\ w_2,\ w_3,\ w_4$ and w_5 to save space.

The D^L and D^U terms that appear in each regression are dummy variables that take a value of one if the return lies in the bottom (top) one per cent of the return distribution. These are defined separately for each market and within each market, they are defined separately for overnight and intraday returns.

The β_{ij} news announcement parameter estimates may be matched against specific announcements in the following way. The i subscript refers the country where the announcement occurred and the j subscript refers to the specific announcement. For Japan, the β_{1j} are coefficients on the ordered values of the Japanese CPI, PPI, IP, and MS announcements (j = 1, 2, 3, 4). For example, β_{13} is the coefficient on the Japanese industrial production announcement. The β_{2j} are coefficients on U.S. news announcements using the order CPI, PPI, IP, UR, TD, and MS (j = 1,…,6). The β_{3j} are coefficients on U.K. news announcements using the order IP, CPI, MS, and UR (j = 1,…, 4). Using the news announcement order given below, the same procedure is used to identify individual coefficients for the U.K. and U.S. regressions. For example, in the U.S. intraday regression, β_{15} is the coefficient on the U.S. trade deficit announcement.

See Sections II and III for more detail about the model and data. We indicate statistical significance of parameter estimates using the following notation: *: 1 per cent significance; **: 5 per cent significance; ***: 10 per cent significance.

Table 7

Estimates of All-Inclusive Nonlinear News Models for U.S., U.K., and Japan, 1985 – 1996

Parameter Estimates	Japan		U.K.		U.S.	
	Intraday	**Overnight**	**Intraday**	**Overnight**	**Intraday**	**Overnight**
ρ_1	-0.030 (0.017)***	0.066 (0.014)*	-0.065 (0.023)*	0.124 (0.009)*	0.039 (0.027)	0.090 (0.010)*
β_{11}	0.302 (0.065)*	-0.032 (0.177)	-0.198 (0.033)*	-0.048 (0.037)	-0.013 (0.245)	-0.169 (0.117)
β_{12}	-0.758 (0.232)*	0.225 (0.193)	-0.306 (1.052)	-0.174 (0.242)	-0.010 (0.160)	-0.057 (0.069)
β_{13}	-0.298 (0.017)*	0.021 (0.028)	0.336 (0.786)	0.324 (0.116)*	0.419 (0.222)***	-0.006 (0.078)
β_{14}	70.734 (8.176)*	-28.447(10.211)*	0.912 (1.035)	-0.025 (0.426)	-21.370 (1.685)*	-1.914 (1.450)
β_{15}					0.717 (0.391)***	-0.177 (0.253)
β_{16}					0.001 (0.019)	-0.004 (0.007)
ρ_2	-0.001 (0.029)	0.218 (0.022)*	0.068 (0.017)*	-0.015 (0.007)**	0.087 (0.018)*	0.196 (0.009)*
β_{21}	0.349 (0.226)	-0.697 (0.073)*	-0.047 (0.079)	0.111 (0.013)*	0.072 (0.081)	-0.026 (0.030)
β_{22}	0.207 (0.121)***	0.123 (0.113)	0.047 (0.277)	-0.156 (0.084)***	0.217 (0.426)	-0.274 (0.216)
β_{23}	0.302 (0.171)***	-0.036 (0.097)	-0.040 (0.035)	-0.021 (0.012)***	0.114 (0.378)	0.246 (0.170)
β_{24}	6.445 (3.454)***	2.145 (2.767)	-10.392(11.786)	4.661 (4.928)	0.155 (0.978)	0.391 (0.391)
β_{25}	-1.375 (0.643)**	-0.391 (0.571)				
β_{26}	0.004 (0.035)	-0.041 (0.018)**				
δ_2	-6.553 (3.185)**	-8.620 (2.645)*	-0.024 (0.022)	-0.005 (0.009)	0.006 (0.030)	-0.031 (0.015)**
η_{21}	0.332 (0.037)*	0.057 (0.031)***	-0.069 (0.042)	0.195 (0.017)*	0.626 (0.262)**	0.043 (0.103)
η_{22}	-0.015 (0.080)	-0.023 (0.065)	0.061 (0.035)***	-0.062 (0.022)*	-0.203 (0.258)	0.125 (0.090)
ρ_3	0.034 (0.026)	0.118 (0.021)*	-0.024 (0.025)	0.376 (0.011)*	-0.014 (0.017)	0.052 (0.008)*
β_{31}	0.071 (0.072)	-0.035 (0.067)	-0.023 (0.286)	0.180 (0.116)	-0.021 (0.087)	-0.004 (0.042)
β_{32}	-0.199 (0.571)	0.957 (0.275)*	0.116 (0.056)**	0.153 (0.051)*	-0.056 (0.201)	-0.156 (0.066)**
β_{33}	0.094 (0.229)	-1.090 (0.090)*	0.055 (0.080)	0.003 (0.079)	-0.006 (0.042)	0.026 (0.014)***
β_{34}	0.335 (0.676)	0.082 (0.392)	6.902 (3.677)***	1.652 (0.904)***	11.446(10.129)	-3.171 (5.360)
β_{35}			0.661 (0.263)**	-0.360 (0.237)		
β_{36}			0.003 (0.014)	-0.011 (0.015)		
δ_3	0.008 (0.045)	0.011 (0.036)	0.209 (3.978)	19.004 (1.451)*	-0.068 (0.020)*	0.054 (0.008)*
η_{31}	0.162 (0.070)**	-0.020 (0.050)	0.088 (0.037)**	0.229 (0.017)*	0.002 (0.032)	0.018 (0.019)
η_{32}	-0.078 (0.134)	0.078 (0.079)	-0.057 (0.057)	-0.070 (0.024)*	-0.026 (0.041)	0.001 (0.014)

Table 7 (cont.)

Estimates of All-Inclusive Nonlinear News Models for U.S., U.K., and Japan, 1985 – 1996

Notes: We report estimates of the parameters of the conditional mean models for Japan, the U.K., and the U.S., respectively, using daily intraday and overnight returns for the 1985 – 1996 period. To save space, we show the conditional mean models only for intraday returns as follows:

Japan:
$$NKID_t = a + (r_1 + b_1 \cdot NEWS_{JP,t}) \cdot NKON_t + (r_2 + b_2 \cdot NEWS_{US,t-1} + d_2 \cdot RRD_{US,t-1} + h_{21} \cdot D^L_{US,t-1} + h_{22} \cdot D^U_{US,t-1}) \cdot SPID_{t-1}$$
$$+ (r_3 + b_3 \cdot NEWS_{UK,t-1} + d_3 \cdot RRD_{UK,t-1} + h_{31} \cdot D^L_{UK,t-1} + h_{32} \cdot D^U_{UK,t-1}) \cdot FTID_{t-1} + j \cdot MHD_{JP,t} + e^{ID}_{JP,t}$$

UK:
$$FTID_t = a + (r_1 + b_1 \cdot NEWS_{UK,t}) \cdot FTON_t + (r_2 + b_2 \cdot NEWS_{JP,t} + d_2 \cdot RRD_{JP,t} + h_{21} \cdot D^L_{JP,t} + h_{22} \cdot D^U_{JP,t}) \cdot NKID_t$$
$$+ (r_3 + b_3 \cdot NEWS_{US,t} + d_3 \cdot RRD_{US,t-1} + h_{31} \cdot D^L_{US,t-1} + h_{32} \cdot D^U_{US,t-1}) \cdot SPID_{t-1} + j \cdot MHD_{UK,t} + e^{ID}_{UK,t}$$

US:
$$SPID_t = a + (r_1 + b_1 \cdot NEWS_{US,t}) \cdot SPON_t + (r_2 + b_2 \cdot NEWS_{UK,t} + d_2 \cdot RRD_{UK,t} + h_{21} \cdot D^L_{UK,t} + h_{22} \cdot D^U_{UK,t}) \cdot FTID_t$$
$$+ (r_3 + b_3 \cdot NEWS_{JP,t} + d_3 \cdot RRD_{JP,t} + h_{31} \cdot D^L_{JP,t} + h_{32} \cdot D^U_{JP,t}) \cdot NKID_t + j \cdot MHD_{US,t} + e^{ID}_{US,t}$$

Standard errors, indicated in parentheses on the line beside the point estimates, are computed from a covariance matrix that accounts for conditional volatility using the Glosten-Jagannathan-Runkle (GJR) asymmetric GARCH model. Specifically, the conditional volatility models for intraday returns is given by

Intraday Volatility: $e^{ID}_{i,t} \sim N(0, h^{ID}_{i,t}),\ \ h^{ID}_{i,t} = w_0 + w_1 \cdot e^{2(ID)}_{i,t-1} + w_2 \cdot h^{ID}_{i,t-1} + w_3 \cdot e^{2-(ID)}_{i,t-1} + w_4 \cdot MHD_{i,t} + w_5 \cdot INT_{i,t}$ (i = JP, UK, US).

We have suppressed reporting the estimates of the coefficients $a,\ j,\ w_0,\ w_1,\ w_2,\ w_3,\ w_4$ and w_5 to save space.

The β_{ij} news announcement parameter estimates may be matched against specific announcements in the following way. The i subscript refers the country where the announcement occurred and the j subscript refers to the specific announcement. For Japan, the β_{1j} are coefficients on the ordered values of the Japanese CPI, PPI, IP, and MS announcements (j = 1, 2, 3, 4). For example, β_{13} is the coefficient on the Japanese industrial production announcement. The β_{2j} are coefficients on U.S. news announcements using the order CPI, PPI, IP, UR, TD, and MS (j = 1,…,6). The β_{3j} are coefficients on U.K. news announcements using the order IP, CPI, MS, and UR (j = 1,…, 4). Using the news announcement order given below, the same procedure is used to identify individual coefficients for the U.K. and U.S. regressions. For example, in the U.S. intraday regression, β_{15} is the coefficient on the U.S. trade deficit announcement.

See Sections II and III for more detail about the model and data. We indicate statistical significance of parameter estimates using the following notation: *: 1 per cent significance; **: 5 per cent significance; ***: 10 per cent significance.

Table 8

R^2 and Posterior Odds Comparison of Empirical Equity Market Return Co-Movement Models

Panel A: R^2 Values for Conditional Mean Models

Model	Equation	Japan Intraday	Japan Overnight	U.K. Intraday	U.K. Overnight	U.S. Intraday	U.S. Overnight
Baseline	3 and 4	0.019	0.080	0.014	0.351	0.036	0.121
Linear News	5 and 6	0.023	0.084	0.018	0.326	0.041	0.124
Nonlinear News	7 and 8	0.055	0.091	0.025	0.368	0.056	0.156
Nonlinear News + Return Dispersion	11 and 12	0.056	0.097	0.026	0.393	0.058	0.197
Nonlinear News + Extreme Returns	13 and 14	0.072	0.094	0.031	0.449	0.065	0.200
All-Inclusive	15 and 16	0.074	0.098	0.033	0.453	0.070	0.236

Panel B: Posterior Odds

Model Comparison	Japan Intraday	Japan Favored	Japan Overnight	Japan Favored	U.K. Intraday	U.K. Favored	U.K. Overnight	U.K. Favored	U.S. Intraday	U.S. Favored	U.S. Overnight	U.S. Favored
3 vs. 5	a	3	a	3	a	3	a	3	a	3	a	3
3 vs. 7	1208	3	a	3	a	3	a	3	a	3	119	3
3 vs. 11	7906	3	a	3	a	3	b	11	a	3	b	11
3 vs. 13	0.01	13	a	3	a	3	b	13	a	3	b	13
3 vs. 15	0.85	15	a	3	a	3	b	15	a	3	b	15
5 vs. 7	b	7	b	7	545	5	b	7	0.11	7	b	7
5 vs. 11	b	11	b	11	915	5	b	11	1.29	5	b	11
5 vs. 13	b	13	b	13	1243	5	b	13	0.20	13	b	13
5 vs. 15	b	15	b	15	a	5	b	15	0.93	15	b	15
7 vs. 11	6.54	7	b	11	1.68	7	b	11	11	7	b	11
7 vs. 13	b	13	0.22	13	2.28	7	b	13	1.74	7	b	13
7 vs. 15	0.001	15	1.96	7	1021	7	b	15	8.08	7	b	15
11 vs. 13	b	13	a	11	1.36	11	b	13	0.15	13	30	11
11 vs. 15	0.16	15	1.59	11	461	11	b	15	6.88	11	b	15
13 vs. 15	155	13	8.81	13	448	13	0.31	15	4.64	13	b	15

Table 8 (cont.)

R^2 and Posterior Odds Comparison of Empirical Equity Market Return Co-Movement Models

Note: An "a" represents a posterior odds value greater than 10,000 while a "b" represents a posterior odds value less than 0.00001. Posterior odds above one indicate the data favor the model in the numerator. Posterior odds below one indicate the data favor the model in the denominator. For example, the posterior odds of (3) vs. (11) for Japanese intraday returns are 7906. This means that model 3 is 7906 times more likely than model 11 given the data. The same model comparison for U.K. overnight returns indicates that the odds are less than .00001. In this case, the data strongly favor model (11) against model (3).

These calculated posterior odds employ diffuse priors and even prior odds. Under these circumstances, Leamer (1978) shows posterior odds may be calculated as $p(H_1|NKID)/p(H_2|NKID) = T^{-(d/2)}(SSE_2/SSE_1)^{T/2}$ for Japanese intraday return (*NKID*), where T is the sum of the sample size, d is the difference in the number of parameters under the two hypotheses, and SSE_i is the sum of squared errors for the i^{th} hypothesized model. This is equation (20) in the text. The SSE_i entries are retrieved from a least squares regression corresponding to the indicated model where we estimated the covariance matrix using the ROBUSTERRORS option in RATS.

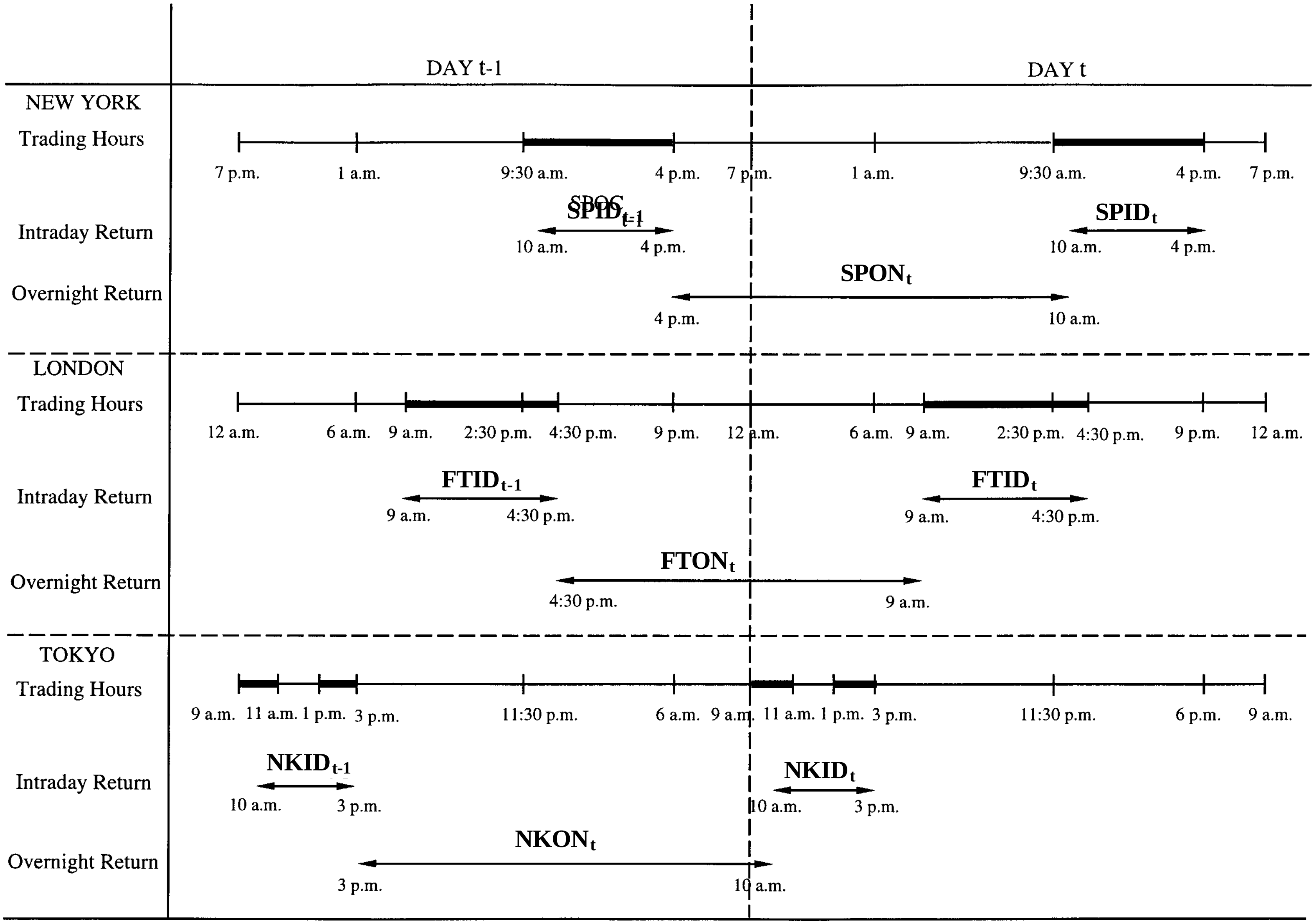

Figure 1. Timing Conventions for Intraday and Overnight Index Returns for the U.S., U.K., and Japanese Equity Markets