Volatility patterns in global financial markets

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Abstract

This paper investigates intraday patterns in global foreign exchange, equity and bond markets using recent advances in the measurement of volatility. The specific objective is to draw conclusions as to how news propagates around the global marketplace. The so-called meteor shower and heatwave hypotheses are rejected for all markets, which highlights a more complicated structure of links between them. The impulse response function analysis confirms that potential shocks have generally a positive effect on volatility, through a complex array of relationships. Patterns in variance decomposition confirm the predominance of country specific news across all three markets.

Keywords
Volatility, realized volatility, news arrival, vector autoregression, impulse response functions

JEL Classification Numbers
C58, G15

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1 Introduction

The importance of the volatility of financial assets in financial decision making and risk management has given rise to a voluminous body of research on the patterns in and transmission of volatility at both the domestic and international level. Broadly speaking, there are two of strands of research identifiable in the existing literature. The first, and more prominent of these, focusses on the time-series transmission of volatility in a single asset across international trading zones, while the second examines the transmission of volatility between different asset markets.

Engle, Ito and Lin (1990) examine international linkages in foreign exchange volatility. Using the framework of Ito (1987) and Ito and Roley (1987), Engle, Ito and Lin (1990) partition each 24 hour period (calendar day) into four discrete trading zones, Asia, Japan, Europe and the United States which have a natural ordering within each day.$^1$ Two alternative patterns in news arrival and hence volatility across these zones are then proposed. The first is the ‘heatwave’ effect in which high volatility is expected to be followed by high volatility in the same trading zone on the following calendar day. The alternative is the ‘meteor shower’ effect where high volatility is expected to be followed by high volatility in the subsequent trading zone in within the same calendar day. The major conclusion that emerges from this line of research is that volatility in the foreign exchange market is best described as a meteor shower. Given that volatility is linked to news arrival (Andersen, 1996; Clark, 1973; Ederington and Lee, 1993; Tauchen and Pitts, 1983), the implication of this result is that news is global phenomenon. In contrast to this result, using a similar research protocol Fleming and

$^1$A two hour Asia trading period is followed by Japan, Europe and finally the U.S. in non-overlapping trading periods.
Lopez (1999) and Savva, Osborn and Gill (2005) find that the heatwave hypothesis best describes the behaviour of volatility in bond and equity markets, respectively.

Despite the fact that in recent decades the globalisation of financial markets has been rapid, there is still a limited understanding of how volatility is transmitted internationally across foreign exchange, equity, and bond markets. Although there have been some significant achievements in this area (Ehrman, Fratzscher and Rigobon 2011; Hakim and McAleer, 2010), there remains scope for research in this area.

This paper uses a specially constructed data set comprising high-frequency foreign exchange, equity, and bond market data to explore the the transmission of volatility between these markets and across international trading zones. The calendar structure used by Engle, Ito and Lin (1990) is amended slightly so that three seven-hour trading zones for Japan, Europe and the United States are established and high frequency returns are used to construct realised volatility estimates for each asset class in each zone for each calendar day. The behaviour of the volatility is then examined from a number of perspectives, namely, transmission across asset classes in local markets, linkages between international trading zones for each asset class and finally the most general case of linkages between all asset classes in the global market.

The rest of the paper proceeds as follows. Section 2 discusses jump-robust measures of integrated volatility which informs the construction of the data set. Section 3 describes the construction of the global trading day and also the high-frequency data set used in the paper. Section 4 addresses the issue of the transmission of volatility between the foreign exchange, equity and bond markets of a single trading zone. This is the simplest
case to address and employs simple vector autoregressive models to explore volatility linkages. The analysis of Section 5 explores volatility patterns between trading zones but within a given market. The analysis is now complicated by the calendar structure of the global trading day which allows contemporaneous influences between zones. Structural vector autoregression models are used to account for these calendar restrictions. Section 6 the estimates a general model that allows for unrestricted interaction between markets and global trading zones. In a nutshell, the results of this research indicated that volatility linkages between different markets and across global trading zones are fairly complex. Contemporaneous influences from other global trading zones within the period of a 24 hour global day are significant and this means that volatility patterns and cannot be described in terms of the heatwave hypothesis. On the other hand, lagged volatility from the same zone is always an important explanatory variable so that the pure meteor shower hypothesis is also not appropriate.

2 Computing Realised Volatility

The central purpose of this research is to explore volatility linkages between important financial markets and also between the main financial hubs of the global market, namely, Japan, Europe, and the United States. To achieve this, it is necessary to put together a comprehensive data set capturing the volatility of these asset markets and trading zones. This begs the question as to how volatility is to be defined. Earlier papers looking at this question (Engle, Ito and Lin, 1990; Fleming and Lopez, 1999; Savva, Osborn and Gill, 2005) treat volatility as unobserved and use the GARCH modelling framework pioneered by Engle (1982) and Bollerslev (1986).
By contrast, we opt to use the realised volatility framework in which an observed proxy for volatility is constructed form high-frequency return variation (Anderson, Bollerslev, Diebold and Labys, 2001, 2003). The use of such observed proxies for volatility means that traditional vector time series techniques can be used to examine in detail the patterns in volatility within each market, and also across the respective markets. One can also obtain a clear picture of the impact of shocks to volatility emanating from the various trading zones within an impulse response framework. Such analysis would be much more difficult within a GARCH framework. For the purposes of estimating volatility and its associated components, define a jump-diffusion process for the logarithm of price,

\[ dp(t) = \mu(t)dt + \sigma(t)dW(t) + \kappa(t)dq(t) \] (1)

in which \( \mu(t) \) is a drift process, \( \sigma(t) \) is a positive stochastic volatility process, \( dW(t) \) is the increment of a Wiener process and \( q(t) \) is a counting process with intensity \( \lambda(t) \), \( t = 1, ..., T \). \( P[dq(t) = 1] = \lambda(t) \) and \( \kappa(t) \) reflects the size of discrete price jumps. It is well known that realised variation (commonly known as realised volatility) is defined as

\[ RV_{t+1}(\Delta) \equiv \frac{1}{\Delta} \sum_{j=1}^{1/\Delta} r_{t+j,\Delta,\Delta}^2, \] (2)

which is the sum of intraday squared returns and converges to the quadratic variation

\[ QV_{t+1} = \int_t^{t+1} \sigma^2(s)ds + \sum_{t<s\leq t+1} \kappa^2(s). \] (3)

The proxy for volatility in equation (3) includes contributions from both the continuous and jump components of prices. Anderson, Bollerslev and Diebold (2007), however, demonstrate that information pertaining to future
volatility is best captured by the persistent diffusive component of volatility. Using the the diffusive component realised volatility is therefore likely to provide more reliable estimates of volatility linkages in what might be loosely termed ‘normal’ market conditions. As these linkages are the primary focus of this research, a necessary prerequisite is a reliable method for decomposing total volatility into a continuous diffusive process and a discrete jump process.

A number of methods exist to effect this decomposition and provide volatility indicators that are robust to jumps, the earliest of which is the bi-power variation (Barndorff-Nielsen and Shephard, 2004, 2002), given by

\[
BV_{t+1}(\Delta) \equiv \mu_1^{-2} \sum_{j=2}^{1/\Delta} \left| r_{t+j,\Delta,\Delta} \right| \left| r_{t+(j-1),\Delta,\Delta} \right|
\]  

(4)

in which \( \mu_1 = \sqrt{2/\pi} \). This measure converges to integrated volatility, it is possible to decompose the total volatility into the contribution from jumps,

\[
RV_{t+1}(\Delta) - BV_{t+1}(\Delta) \to \sum_{t<s \leq t+1} \kappa^2(s).
\]  

(5)

An important result that follows from equations (4) and (5) is that by construction, the bi-power variation can be used as an estimator of quadratic variance robust to jumps. Ait-Sahalia, Jacod and Li (2012) and Mancini (2009) propose two such estimators. These are truncated realised realised volatility, given by

\[
TRV_{t+1}(\Delta, u_n) \equiv \sum_{j=1}^{1/\Delta} r_{t+j,\Delta,\Delta}^2 \cdot 1_{\{||r_{t+j,\Delta,\Delta}|| \leq u_n\}}
\]  

(6)

and truncated power variation

\[
TPV_{t+1}(\Delta, u_n, p) \equiv \sum_{j=1}^{1/\Delta} \left| r_{t+j,\Delta,\Delta} \right|^p \cdot 1_{\{|r_{t+j,\Delta,\Delta}| \leq u_n\}}
\]  

(7)
in which \( u_n = \alpha \Delta^\varpi \) is a suitable sequence going to 0, \( \alpha > 0 \), \( \varpi \) is an arbitrary constant, and \( p \geq 2 \) is a positive integer.

Of course, in practice suitable choices for \( \alpha \) and \( \varpi \) must be chosen. Todorov, Tauchen and Grynkiv (2011) argue that \( \alpha = 3\sqrt{BV_{t+1}} \) and \( \varpi \in (0,1/2) \) and these conditions are intuitively reasonable. However, it is necessary to note that in choosing these parameters there is a risk of throwing away many Brownian increments, which makes it difficult to use this method in practice.

Andersen, Dobrev and Schaumburg (2009) introduce an alternative jump robust estimator known as minimum realised volatility

\[
\text{MinRV}_{t+1}(\Delta) \equiv \frac{\pi}{\pi - 2} \left( \frac{1}{1 - \Delta} \right)^{1/\Delta} \sum_{j=2}^{1/\Delta} \min(|r_{t+j-1,\Delta}|, |r_{t+(j-1),\Delta}|)^2. \tag{8}
\]

Andersen, Dobrev and Schaumburg (2009) justify that minimum realised volatility measure provides a better finite sample properties than bi-power variation. Due to this fact, and taking into account the arbitrary character of choosing the threshold \( \varpi \) in truncated power variation (even though it is more asymptotically efficient than bi-power variation and minimum realised volatility) for volatility estimation, the MinRV measure of integrated volatility robust to jumps is used.

Based on the asymptotic results of Barndorff-Nielsen and Shephard (2004), Barndorff-Nielsen and Shephard (2006) and using the fact that\(^2\)

\[
\sqrt{\frac{1}{\Delta}} \left( \text{MinRV}_{t+1} - \int_t^{t+1} \sigma^2(s)ds \right) \overset{\text{stableD}}{\rightarrow} MN \left( 0, 3.81 \int_t^{t+1} \sigma^4(s)ds \right), \tag{9}
\]

---

\(^2\)See proposition 2 and 3 in Andersen, Dobrev and Schaumburg (2009), p. 78.
statistically significant jumps are identified according to

$$Z_{t+1}(\Delta) \equiv \frac{[RV_{t+1}(\Delta) - \text{MinRV}_{t+1}(\Delta)]/RV_{t+1}(\Delta)}{[1.81\Delta \max(1, \text{MinRQ}_{t+1}(\Delta)/\text{MinRV}_{t+1}(\Delta)^2)]^{1/2}} \sim N(0, 1)$$

(10)

where MinRQ is a minimum realised quarticity

$$\text{MinRQ}_{t+1}(\Delta) \equiv \frac{\pi}{\Delta(3\pi - 8)} \left( \frac{1}{1 - \Delta} \right)^{1/\Delta} \sum_{j=2}^{1/\Delta} \min(|r_{t+j, \Delta, \Delta}|, |r_{t+(j-1), \Delta, \Delta}|)^4.$$  

(11)

Significant jumps, at an $\alpha$ level of significance are identified by $Z_{t+1}(\Delta) > \Phi_{1-\alpha}$,

$$J_{t+1}(\Delta)(Z) \equiv 1[Z_{t+1}(\Delta) > \Phi_{1-\alpha}] \cdot [RV_{t+1}(\Delta) - \text{MinRV}_{t+1}(\Delta)].$$

(12)

In constructing the series of realised volatility to be used in the empirical sections of the paper, equation (8) is used. This methodology can also be used to estimate the significant jump component of realised volatility and thus facilitate the exploration of volatility transmission in periods of ‘abnormal’ volatility. This avenue of research is being pursued in another paper.

3 Data

In order to compute the preferred proxy for volatility, namely, minimum realised volatility, a data set is collected comprising high frequency (1 minute) data for foreign exchange, equity and bond markets for each of three regions, Japan, Europe and the United States. The data was gathered from the Thomson Reuters Tick History database and covers the period from 3 January 2005 to 30 June 2011. Days where one market is closed are eliminated, as are public holidays or other occasions when trading is significant
curtailed. These high frequency data are then used to construct minimum realised volatility consisting of 1099 full trading days.

Before setting out the exact specification of the data that was collected it is necessary to define the global trading day which is integral to this research. Each calendar day is split into three trading zones, namely Japan (JP), Europe (EU) and the United States (US). The Japan trading zone is defined as 12am to 7am, the European trading zone 7am to 2pm and the United States zone 2pm to 9pm, where all times are taken to be Greenwich Mean Time (GMT). This setup may be illustrated as follows:

\[
\begin{array}{ccc}
\text{Japan} & \text{Europe} & \text{U.S.} \\
12am \cdots 7am & 7am \cdots 2pm & 2pm \cdots 9pm \\
\hline
\text{One Trading Day}
\end{array}
\]

The foreign exchange rate data in each of the three trading zones consists of closing prices for 10 minute intervals on Yen-Dollar futures contracts traded on the Chicago Mercantile Exchange. The bond market data consists of 10 minute prices for Japanese, German and United States Treasury note 10-year bond futures contracts. For equity markets, 10 minute prices were collected for TOPIX (JP), DAX (EU) and S&P500 futures contracts.

The nine series for realised volatility, calculated using equation (8) applied to each asset class in each trading zone, are plotted in Figure 1. To the naked eye it appears that the estimates of realised volatility in foreign exchange and equity markets have similar patterns across the trading zones. The volatility in the United States is perhaps a little more pronounced during the Global Financial Crises period of 2007 - 2009. The similarity across the

\[\text{9}\]

\[\text{3The period denoted as Asian trading (2 hours prior to Japan opening) by Engle, Ito and Lin (1990), is excluded here.}\]
three zones in not as pronounced, however, in the bond markets. Figure 1 indicates that while realised volatility in the Japanese and United States is very similar, realised volatility in the European zone appears to experience more volatility events (appears more spiked) that the other zones.

Figure 1: Minimum realised volatility estimates for the foreign exchange, equity and bond markets in Japan, Europe and United States, respectively. The daily estimate of realised volatility for the period 3 January 2005 to 30 June 2011 is computing using (8) and then scaled by 1000 before plotting.

In some instances, to enhance the appearance of the figures, the scales on the y-axes of Figure 1 are different for some of the trading zones, making it difficult to compare the relative sizes of realised volatility. Table 1 reports
summary statistics for the logarithm of the minimum realised volatility series. The statistics are reported for the logarithm of each series because it is these transformed data that are used in the estimation.

<table>
<thead>
<tr>
<th></th>
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</thead>
<tbody>
<tr>
<td><strong>FX</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Japan</td>
<td>-13.838</td>
<td>0.834</td>
<td>-16.054</td>
<td>-10.176</td>
<td>0.531</td>
<td>3.783</td>
</tr>
<tr>
<td>Europe</td>
<td>-13.243</td>
<td>0.760</td>
<td>-15.421</td>
<td>-9.194</td>
<td>0.378</td>
<td>3.855</td>
</tr>
<tr>
<td>U.S.</td>
<td>-13.424</td>
<td>0.881</td>
<td>-15.942</td>
<td>-10.083</td>
<td>0.511</td>
<td>3.691</td>
</tr>
<tr>
<td><strong>Equity</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Japan</td>
<td>-11.565</td>
<td>0.919</td>
<td>-14.037</td>
<td>-7.423</td>
<td>0.586</td>
<td>4.283</td>
</tr>
<tr>
<td>Europe</td>
<td>-12.026</td>
<td>1.100</td>
<td>-14.717</td>
<td>-7.535</td>
<td>0.494</td>
<td>3.551</td>
</tr>
<tr>
<td>U.S.</td>
<td>-11.915</td>
<td>1.171</td>
<td>-14.640</td>
<td>-7.078</td>
<td>0.852</td>
<td>3.833</td>
</tr>
<tr>
<td><strong>Bond</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Japan</td>
<td>-14.938</td>
<td>0.889</td>
<td>-17.184</td>
<td>-11.083</td>
<td>0.595</td>
<td>3.719</td>
</tr>
<tr>
<td>Europe</td>
<td>-14.259</td>
<td>0.744</td>
<td>-16.507</td>
<td>-12.022</td>
<td>0.164</td>
<td>2.794</td>
</tr>
<tr>
<td>U.S.</td>
<td>-14.055</td>
<td>0.887</td>
<td>-16.270</td>
<td>-10.598</td>
<td>0.300</td>
<td>2.826</td>
</tr>
</tbody>
</table>

Table 1: Descriptive statistics for daily estimates of the logarithm of realised volatility in the foreign exchange, equity and bond markets in Japan, Europe and United States for the period 3 January 2005 to 30 June 2011.

As can be seen from the Table 1 the level of volatility in each market is similar irrespective of global trading zone. Interestingly enough mean volatility is highest in the Japanese bond market. Surprisingly, the mean volatilities in all three markets in the United States are not uniformly larger mean volatilities in all the trading zones (although it is true that the variability of the logarithm of realised volatility is generally higher in the United States). This appears to contradict the original view of Engle, Ito and Lin (1990), who comment that Treasury market volatility is substantially higher during the New York trading hours than during Tokyo or London trading hours. Their view is that much of this volatility seems to originate with macroeconomic announcements released during New York trading hours. On the basis of the summary statistics presented in Table 1, however, there is little
evidence to support the conjecture that if volatility spillovers do occur, they probably flow from New York to the overseas trading centres. Neither is there any reason to expect, a priori that the main result reported by Engle, Ito and Lin (1990) that the meteor shower form of volatility spillover is more likely to be found for Tokyo and London than for New York.

Attempting to model the transmission of volatility across markets and between trading zones necessarily requires that there be some structure in the volatility series being modelled. To explore whether or not there is a prima facie case for continuing the investigation, the sample autocorrelations out to ten lags for each of the nine series are reported in Table 2.

<table>
<thead>
<tr>
<th>Lag</th>
<th>Foreign exchange</th>
<th>Equity</th>
<th>Bond</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>0.540</td>
<td>0.515</td>
<td>0.558</td>
</tr>
<tr>
<td>2</td>
<td>0.491</td>
<td>0.459</td>
<td>0.515</td>
</tr>
<tr>
<td>3</td>
<td>0.452</td>
<td>0.389</td>
<td>0.467</td>
</tr>
<tr>
<td>4</td>
<td>0.432</td>
<td>0.416</td>
<td>0.476</td>
</tr>
<tr>
<td>5</td>
<td>0.382</td>
<td>0.364</td>
<td>0.458</td>
</tr>
<tr>
<td>6</td>
<td>0.399</td>
<td>0.390</td>
<td>0.462</td>
</tr>
<tr>
<td>7</td>
<td>0.408</td>
<td>0.363</td>
<td>0.471</td>
</tr>
<tr>
<td>8</td>
<td>0.398</td>
<td>0.344</td>
<td>0.429</td>
</tr>
<tr>
<td>9</td>
<td>0.397</td>
<td>0.344</td>
<td>0.447</td>
</tr>
<tr>
<td>10</td>
<td>0.360</td>
<td>0.320</td>
<td>0.389</td>
</tr>
</tbody>
</table>

Table 2: Sample autocorrelations of the logarithm of realised volatility in the foreign exchange, equity, and bond markets of Japan, Europe, and United States, respectively.

The sample auto-correlations of the realised volatility for the global foreign exchange, equity, and bond markets are all statistically significant and indicate a fair amount of persistence. This is to be expected given that
autorcorrelation in the squares of financial returns is a well-known and well-
documented phenomenon (Pagan, 1996). An interesting result is that the
autocorrelation coefficients appear to be quite smaller in the Japanese bond
and equity markets. This may be evidence that the ‘heatwave’ hypothesis
is weakest in Japan which would compete with the broad conclusions of

4 Volatility Transmission Between Markets

The section analyses volatility interaction between the foreign exchange,
equity, and bond markets in each of the three trading zones. For the moment
the assumption is that each of the trading blocks is unaffected by the others.
This is a rather strict assumption which will be relaxed subsequently, after
these simple benchmark models have been examined.

The econometric model to be estimated here is a simple Vector AutoRegras-
sion (VAR) in each of the three global trading zones. Let $y_t$ be the vector
of logarithms of minimum realised volatilities for the foreign exchange, eq-
uiity and bond markets for a given trading zone. The VAR model for this
particular trading zone is therefore

$$y_t = \sum_{p=1}^{P} \Phi_p y_{t-p} + \nu_t, \quad \nu_t \sim iid N(0, V), \quad (13)$$

in which $\nu_t$ is the vector of reduced form disturbances with covariance matrix
$V$, the $\Phi_p$ are matrices of parameters.\(^4\)

The first practical issue at hand is the correct choice of optimal lag length,
$P$. Given the rapid dissemination of news in financial markets, intuition

\(^4\)The intercept term has been omitted in this equation for notational convenience. Intercept terms were included in the estimation.
would suggest that one week would be enough to capture all the relevant information in lagged values of realised volatility. This conjecture is only partly supported by the results reported in Table 3 which reports the optimal choice of $P$ in equation (13) in terms of a number of well known information criterion. As expected, the HIC and SIC favour a more parsimonious lag structure than the FPE and the AIC. The latter indicate that a lag structure of about a week is sufficient while the former suggest that a two-week period is perhaps more appropriate.

<table>
<thead>
<tr>
<th>Trading Zone</th>
<th>FPE</th>
<th>AIC</th>
<th>HIC</th>
<th>SIC</th>
</tr>
</thead>
<tbody>
<tr>
<td>Japan</td>
<td>8</td>
<td>8</td>
<td>4</td>
<td>4</td>
</tr>
<tr>
<td>Europe</td>
<td>7</td>
<td>7</td>
<td>4</td>
<td>3</td>
</tr>
<tr>
<td>U.S.</td>
<td>9</td>
<td>9</td>
<td>4</td>
<td>4</td>
</tr>
</tbody>
</table>

Table 3: Optimal choice of lag length for equation (13) for Japan, Europe, and the United States as determined by the FPE, AIC, HIC and SIC information criteria.

In the light of the results of the lag-length selection tests the VAR in equation (13) is estimated using both 5 and 10 lags. As the focus of the research is on overall patterns in volatility, individual parameter estimates for the VAR are suppressed. Suffice to say that, as indicated by the results reported in Table 2, the own lags are highly significant in each of the zone-specific VAR models estimated. In the light of the conflicting evidence on the optimal choice of lag order presented in Table 3, it comes as no surprise to find that the first three own lags are highly significant followed by a number of insignificant lags and then a fairly significant lag at order 8 or 9.

The primary focus of the current analysis is on the volatility linkages between the markets for the different assets in each of the trading zones. Formally,
the interaction between these markets is explored in the context of Granger causality and the results are reported in Table 4 for the VAR(5) and Table 5 for the VAR(10).

<table>
<thead>
<tr>
<th></th>
<th>Japan</th>
<th>Europe</th>
<th>U.S.</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$\chi^2$</td>
<td>p-value</td>
<td>$\chi^2$</td>
</tr>
<tr>
<td>Equity</td>
<td>28.293</td>
<td>0.000</td>
<td>25.748</td>
</tr>
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<td>FX</td>
<td>3.645</td>
<td>0.602</td>
<td>19.164</td>
</tr>
<tr>
<td>Bond</td>
<td>40.317</td>
<td>0.000</td>
<td>72.815</td>
</tr>
<tr>
<td>All</td>
<td>16.388</td>
<td>0.006</td>
<td>6.890</td>
</tr>
<tr>
<td>Equity</td>
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<td>0.051</td>
<td>7.829</td>
</tr>
<tr>
<td>Bond</td>
<td>26.170</td>
<td>0.004</td>
<td>16.996</td>
</tr>
<tr>
<td>All</td>
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<td>0.468</td>
<td>8.532</td>
</tr>
<tr>
<td>Equity</td>
<td>3.615</td>
<td>0.606</td>
<td>26.400</td>
</tr>
<tr>
<td>Bond</td>
<td>10.354</td>
<td>0.410</td>
<td>40.647</td>
</tr>
</tbody>
</table>

Table 4: Granger causality tests for the VAR(5) models estimated for each of the trading zones. The Wald statistics for the null hypothesis that there is no Granger causality and associated p-values are reported.

A careful analysis of the patterns in the Granger causality results in Table 4 reveal a number of interesting results. Granger causality from the equity and bond markets to the foreign exchange market is significant in all three trading zones. The p-value on all of the Wald statistics is 0.000 which indicates the strength of the statistical result. This result makes intuitive sense as volatility in the key domestic markets is likely to influence the foreign exchange market.

The volatility linkages in the equity market appear to be weak. In Europe the equity market which appears not to be Granger caused by either the foreign exchange or the bond market. The Wald statistic for Granger causality from both foreign exchange and bond markets to the equity market
is $\chi^2 = 16.996$ with a p-value of 0.074 indicating that the null hypothesis that there is no Granger causality cannot be rejected at 5%. This result is completely unexpected as the equity market is usually regarded as fairly sensitive to developments elsewhere. Interestingly enough this pattern of an unresponsive equity market is mirrored by the results for the United States with the relevant Wald statistic being 16.485 with a p-value of 0.087. In Japan, however, realised volatility in both foreign exchange and bond markets helps to predict volatility in the equity market.

There is one particularly striking result for the bond market and that is the apparent decoupling of the Japanese bond market from the foreign exchange and equity markets, at least in terms of Granger causality. The Wald statistic for Granger causality from both foreign exchange and equity markets to the bond market is $\chi^2 = 10.354$ with a p-value of 0.410 indicating that the null hypothesis that there is no Granger causality cannot be rejected. The bond market’s influence on both the foreign exchange market and the equity market, when causality in the opposite direction is tested, is marginal at the 5% significance level. This evidence taken together with statistically significant own lags in the VAR(5), which although the estimates are not reported can easily be deduced from Table 2 establishes a fairly strong prima facie case for the ‘heatwave’ hypothesis in the Japanese bond market. This conclusion would concur with the view of Savva, Osborn and Gill (2005), but as will become apparent in the subsequent analysis, such a conclusion is premature at this stage.

The Granger causality tests for the VAR(10) are reported in Table 5. The patterns identified by the VAR(5) are broadly similar for the VAR(10) although evidence of causality is slightly weaker in this case. It now appears that realised volatility in the Japanese equity market is not Granger caused
5 Volatility Transmission Between Zones

This section considers volatility patterns between the three global trading zones, Japan, Europe, and U.S. in each of the three financial markets. Define $y_t$ as the vector of logarithms of realised volatilities to a particular asset in

<table>
<thead>
<tr>
<th></th>
<th>Japan</th>
<th>Europe</th>
<th>U.S.</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$\chi^2$</td>
<td>p-value</td>
<td>$\chi^2$</td>
</tr>
<tr>
<td>Equity</td>
<td>21.039</td>
<td>0.021</td>
<td>23.942</td>
</tr>
<tr>
<td>FX</td>
<td>11.694</td>
<td>0.306</td>
<td>26.785</td>
</tr>
<tr>
<td>Bond</td>
<td>38.969</td>
<td>0.007</td>
<td>69.811</td>
</tr>
<tr>
<td>All</td>
<td>15.876</td>
<td>0.103</td>
<td>10.688</td>
</tr>
<tr>
<td>Equity</td>
<td>12.634</td>
<td>0.245</td>
<td>16.027</td>
</tr>
<tr>
<td>Bond</td>
<td>26.870</td>
<td>0.139</td>
<td>30.100</td>
</tr>
<tr>
<td>All</td>
<td>8.896</td>
<td>0.542</td>
<td>14.925</td>
</tr>
<tr>
<td>Equity</td>
<td>4.532</td>
<td>0.920</td>
<td>22.460</td>
</tr>
<tr>
<td>Bond</td>
<td>14.736</td>
<td>0.791</td>
<td>38.853</td>
</tr>
</tbody>
</table>

Table 5: Granger causality tests for the VAR(10) models estimated for each of the trading zones. The Wald statistics for the null hypothesis that there is no Granger causality and associated p-values are reported.

by realised volatility in either the foreign exchange or bond markets. Furthermore, the United States bond market also now appears to be decoupled from foreign exchange and equity markets at the 5% level. While there is statistical evidence in favour of a longer lag structure, our belief is that the more parsimonious VAR(5) is probably defensible, based on the idea that the dissemination of news in global markets is likely to be complete within 5 working days. Consequently, the choice of 5 lags in the modelling will be adopted.
each of the trading zones and the question of interest is whether or not the volatilities of returns to this particular asset class are linked across trading zones.

5.1 Estimation

The first problem to overcome is that, unlike the analysis of Section 4, there is scope for contemporaneous interaction between the trading zones. For example, events in the foreign exchange market in Japan can influence both Europe and the United States on the same trading day. In fact there is a natural ordering in each calendar day in which imposes the structure $y_{t}^{JP} \rightarrow y_{t}^{EU} \rightarrow y_{t}^{US}$. Consequently the VAR methodology must be augmented slightly and a structural VAR (SVAR) must be estimated in which the calendar structure of the trading day imposes a recursive set of short-run restrictions on the contemporaneous interactions of the variables.

The SVAR model for the realised volatility of a particular asset in each of the trading zones is now represented by the system of equations

$$B_0 y_t = \sum_{p=1}^{P} B_p y_{t-p} + u_t, \quad u_t \sim iid N(0, D), \quad (14)$$

in which $B_0$ is a $(3 \times 3)$ matrix representing the contemporaneous interaction between the variables, $B_i, i = 1, \ldots, P$ are $(3 \times 3)$ parameter matrices and $u_t$ is a vector of disturbances, with covariance matrix $D$, representing the structural shocks.\(^5\)

As already mentioned, the model has a natural recursive structure with the

\(^5\) As in Section 4, the intercept term has been suppressed for simplicity without loss of generality.
contemporaneous matrix, $B_0$, restricted to be the lower triangular matrix

$$
B_0 = \begin{bmatrix}
1 & 0 & 0 \\
-\alpha_{21} & 1 & 0 \\
-\alpha_{31} & -\alpha_{32} & 1
\end{bmatrix},
$$

(15)
in which $\alpha_{21}$ captures the influence of the Japanese zone on the European zone, and $\alpha_{31}$ and $\alpha_{32}$ model, respectively, the effects of Japan and Europe on the United States. It is important to emphasise at this point that the ‘contemporaneous’ effects of Europe and the United States on Japan come from the previous day prior to Japan opening trading at the beginning of the next global trading day. The coefficients of interest are in the matrix $B_1$ and the elements are $\beta_{12}$ for the European influence and $\beta_{13}$ for the effect from the United States.

In this framework, the heatwave hypothesis of Engle, Ito and Lin (1990) requires that the off-diagonal elements of both $B_0$ and $B_1$ are zero. If these zero restrictions are satisfied, realised volatility for the asset class in each trading zone is only a function of the realised volatility from the same trading zone on the previous day.

On the other hand, the pure meteor shower hypothesis requires that the $\alpha_{31} = 0$ and all the elements of $B_1$ are zero, apart from $\beta_{13}$ which captures the effect of the United States on the previous trading day on Japan as the market opens. If these restrictions are satisfied volatility in each zone depends only the volatility in the zone immediately preceding it in the global trading day.

Moreover, the meteor shower effect may take one of two forms, namely world-wide and country-specific news flows. In the former case the impact of volatility is independent of the trading zone with one process describing the evolution of volatility in all zones, which is equivalent to restriction $\alpha_{21} =$
\( \alpha_{32} = \beta_{13} \). In the case of country specific news, volatility in each trading zone has potentially different impacts on subsequent volatility, which means that \( \alpha_{21} \neq \alpha_{32} \neq \beta_{13} \). These hypotheses provide insights into volatility patterns across the markets and how shocks in one trading zone propagate to the other zones.

Taking into account that the structural shocks, \( u_t \), in equation (14) can be expressed in terms of standardized residuals as \( z_t = (D)^{-1/2} u_t \), the dynamics of the structural model can be summarized by its VAR representation

\[
y_t = \sum_{p=1}^{P} B_0^{-1} B_p y_{t-p} + B_0^{-1} (D^{1/2} z_t) = \sum_{p=1}^{P} \Phi_p y_{t-p} + \nu_t, \quad (16)
\]

in which \( \nu_t \sim N(0, V) \), \( \Phi_p = B_0^{-1} B_p \) and \( \nu_t = B_0^{-1} D^{1/2} z_t = S z_t \). Note that covariance matrix \( V \) is not necessarily a diagonal matrix.

Now the structural VAR in equation (14) can be estimated using the following two step procedure (see details of the procedure in Martin, Hurn and Harris (2012, p.558). On the first step, each equation of the VAR from (16) is estimated by OLS to yield \( \hat{\Phi}_1, ..., \hat{\Phi}_1 \). The VAR residuals are given by

\[
\hat{\nu}_t = y_t - \sum_{p=1}^{P} \hat{\Phi}_p y_{t-p}, \quad (17)
\]

which are used to compute the covariance matrix

\[
\hat{V} = \frac{1}{T-P} \sum_{P+1}^{T} \hat{\nu}_t \hat{\nu}_t'.
\]

On the second step, the full information maximum likelihood (FIML) estimates of \( B_0 \) and \( D \) are obtained by maximizing

\[
\ln L_T = -\frac{N}{2} \ln 2\pi - \frac{1}{2} \ln |V| - \frac{1}{2(T-P)} \sum_{t=P+1}^{T} \hat{\nu}_t V^{-1} \hat{\nu}_t, \quad (18)
\]
in which \( N \) is the number of variables (in our case \( N = 3 \)), \( p = 5 \) is a number of lags, and an estimate of covariance matrix \( \hat{V} \) is defined from the first step.

The results of estimating the SVAR for the foreign exchange market is given in Table 6 with the estimates of the elements of \( B_0 \) shown in the shaded panel. The foreign exchange market appears to have fairly complex dynamics with most of the coefficients of the SVAR significant at the 5% level. The contemporaneous linkages, represented by the coefficients \( jpfx_t \) for Europe and the United States, and \( eufx_t \) for Europe, are significant as is the coefficient \( usfx_{t-1} \), which captures the ‘contemporaneous’ effect of the United States on Japan. The unambiguous conclusion seems to be that contemporaneous effects matter in the foreign exchange market. The size of these contemporaneous effects does differ however, so it is unlikely that there is one world-wide news process describing the evolution of volatility.

Another striking result volatility at time \( t - 1 \) is statistically significant in all equations. This suggests that there are also strong lagged volatility linkages, both from the same trading zone and also from both other trading zones. The only other set of coefficients that are statistically significant in all the equations are those pertaining to the United States at \( t - 2 \). In other words, lagged volatility from the United States is important in explaining volatility in all the zones, suggesting that the foreign exchange market is dominated by developments in the United States. This result does support the conjecture of Engle, Ito and Lin (1990) that if volatility spillovers do occur, they probably flow from New York to the overseas trading centres.

Results for the same analysis of the equity market are presented in Table 7. Again, all the coefficients of the contemporaneous matrix are significant at the level 5%, but their magnitudes are smaller in this market. This observation includes the significance of the coefficient on \( useq_{t-1} \) which represents
Table 6: Coefficient estimates of the SVAR in (14) estimated using realised volatility in the foreign exchange market for each of the three trading zones. The shaded panel contains estimates of the elements of the contemporaneous matrix $B_0$. Coefficients that are significant at the 5% level are marked (*)

<table>
<thead>
<tr>
<th></th>
<th>Japan</th>
<th>Europe</th>
<th>U.S.</th>
</tr>
</thead>
<tbody>
<tr>
<td>$jpf_{t-1}$</td>
<td>0.153*</td>
<td>0.073*</td>
<td>0.086*</td>
</tr>
<tr>
<td>$jpf_{t-2}$</td>
<td>0.095*</td>
<td>0.021</td>
<td>0.014</td>
</tr>
<tr>
<td>$jpf_{t-3}$</td>
<td>0.074*</td>
<td>0.050</td>
<td>0.021</td>
</tr>
<tr>
<td>$jpf_{t-4}$</td>
<td>0.064*</td>
<td>0.022</td>
<td>-0.045</td>
</tr>
<tr>
<td>$jpf_{t-5}$</td>
<td>0.021</td>
<td>-0.010</td>
<td>0.051</td>
</tr>
</tbody>
</table>

The contemporaneous effect just prior to market opening in Japan. Apart from this contemporaneous effect from the United States, Japanese volatility is entirely affected by country specific news. The European zone is driven by domestic news with an external effect from the American volatility on the previous day. The United States volatility pattern for the equity market is

22
Table 7: Coefficient estimates of the SVAR in (14) estimated using realised volatility in the equity market for each of the three trading zones. The shaded panel contains estimates of the elements of the contemporaneous matrix $B_0$. Coefficients that are significant at the 5% level are marked (*).

The results for the bond market are presented in Table 8. One can see this market has a very similar volatility patterns to the equity market, however also dominated by domestic news. The overall pattern in the equity market can then be summarised as one of significant contemporaneous interactions between zones, but little effect from lagged volatility in other trading zones.
Table 8: Coefficient estimates of the SVAR in (14) estimated using realised volatility in the bond market for each of the three trading zones. The shaded panel contains estimates of the elements of the contemporaneous matrix $B_0$. Coefficients that are significant at the 5% level are marked (*)

<table>
<thead>
<tr>
<th></th>
<th>Japan</th>
<th>Europe</th>
<th>U.S.</th>
</tr>
</thead>
<tbody>
<tr>
<td>jpx(_t)</td>
<td>1.000</td>
<td>-0.047</td>
<td>-0.063*</td>
</tr>
<tr>
<td>eufx(_t)</td>
<td>...</td>
<td>1.000</td>
<td>-0.162*</td>
</tr>
<tr>
<td>usfx(_t)</td>
<td>...</td>
<td>...</td>
<td>1.000</td>
</tr>
<tr>
<td>jpb(_{t-1})</td>
<td>0.325*</td>
<td>0.018</td>
<td>-0.029</td>
</tr>
<tr>
<td>jpb(_{t-2})</td>
<td>0.142*</td>
<td>-0.020</td>
<td>0.002</td>
</tr>
<tr>
<td>jpb(_{t-3})</td>
<td>0.092*</td>
<td>0.039</td>
<td>0.022</td>
</tr>
<tr>
<td>jpb(_{t-4})</td>
<td>0.131*</td>
<td>-0.015</td>
<td>0.008</td>
</tr>
<tr>
<td>jpb(_{t-5})</td>
<td>0.124*</td>
<td>-0.013</td>
<td>0.030</td>
</tr>
<tr>
<td>eub(_{t-1})</td>
<td>0.024</td>
<td>0.238*</td>
<td>0.118*</td>
</tr>
<tr>
<td>eub(_{t-2})</td>
<td>-0.040</td>
<td>0.108*</td>
<td>0.028</td>
</tr>
<tr>
<td>eub(_{t-3})</td>
<td>-0.013</td>
<td>0.112*</td>
<td>0.063</td>
</tr>
<tr>
<td>eub(_{t-4})</td>
<td>0.081*</td>
<td>0.080*</td>
<td>0.001</td>
</tr>
<tr>
<td>eub(_{t-5})</td>
<td>0.001</td>
<td>0.066*</td>
<td>0.012</td>
</tr>
<tr>
<td>usb(_{t-1})</td>
<td>0.005</td>
<td>0.161*</td>
<td>0.252*</td>
</tr>
<tr>
<td>usb(_{t-2})</td>
<td>0.024</td>
<td>0.041</td>
<td>0.093*</td>
</tr>
<tr>
<td>usb(_{t-3})</td>
<td>0.036</td>
<td>-0.011</td>
<td>0.134*</td>
</tr>
<tr>
<td>usb(_{t-4})</td>
<td>0.003</td>
<td>0.002</td>
<td>0.155*</td>
</tr>
<tr>
<td>usb(_{t-5})</td>
<td>-0.045</td>
<td>0.033</td>
<td>0.090*</td>
</tr>
<tr>
<td>constant</td>
<td>-1.611*</td>
<td>-2.312*</td>
<td>-0.138</td>
</tr>
</tbody>
</table>

some interesting distinctions should be discussed. First of all, surprisingly the contemporaneous effect from Japan to Europe is not significant. Moreover, the lagged inter-zonal volatility in all countries has a weak impact on the global market. In this regard, only eub\(_{t-4}\)\(_p\), usb\(_{t-1}\)\(_m\), and eub\(_{t-1}\)\(_s\) have an inter-zonal impact on the global bond market.
It is now possible to provide a formal test of the heatwave hypothesis as formulated by Engle, Ito and Lin (1990). Recall that this hypothesis requires that volatility in each of zones evolves independently. In effect the heatwave hypothesis implies a complete decoupling of inter-zonal realised volatility. Formally, in this model the heatwave hypothesis requires the restrictions
\[ \alpha_{21} = \alpha_{31} = \alpha_{32} = \beta_{12} = \beta_{13} = \beta_{23} = \beta_{21} = \beta_{31} = \beta_{32} = 0. \]

A simple glance at the significance of the coefficients reported in Tables 6-8 suggests that the heatwave restrictions will be rejected for each of the markets considered here. This casual empiricism is supported by a formal likelihood ratio test which indicates that the heatwave hypothesis is rejected with p-values $p = 0.000$ (foreign exchange market), $p = 0.000$ (equity market), and $p = 0.000$ (bond market). The results reported here both confirm the original result reported by Engle, Ito and Lin (1990) for the foreign exchange market and extend their result to the equity and bond markets.

These results suggest that it is not possible to regard each zone as being completely independent, thus the form of this interaction must now be explored in more detail. The second hypothesis of interest, namely the meteor shower hypothesis, claims that volatility in each of the zones depends only upon volatility in the other zones on the same day, subject to the calendar ordering that is imposed (that is, Japan precedes both Europe and the U.S.). Essentially the meteor shower hypothesis implies strong volatility interactions between the zones and therefore contrasts sharply with the independence implied by the heatwave hypothesis. Formally, the restrictions to be tested are
\[ \alpha_{31} = \beta_{11} = \beta_{12} = \beta_{21} = \beta_{22} = \beta_{23} = \beta_{31} = \beta_{32} = \beta_{33} = 0. \]

Once again, however, Wald tests indicate that the meteor shower hypothesis
is rejected with p-values $p = 0.000$ (foreign exchange market), $p = 0.000$ (equity market), and $p = 0.000$ (bond market).

The general conclusion to emerge from this analysis is that inter-zonal patterns of volatility in the foreign exchange, equity and bond markets are neither a pure heatwave effect nor a pure meteor shower. Instead, it appears that there are strong linkages between realised volatility in all of the three trading zones in all of the markets considered which includes elements of both these effects.

5.2 Impulse Responses and Variance Decomposition

Taking into account that equation (16) may be represented in vector moving average form as\(^6\)

$$y_t = \sum_{q=0}^{\infty} \Psi_q \nu_{t-q},$$

(19)

in which $\Psi_q$ are moving average parameter matrices. An effect of the shocks in $\nu_t$ on the future time path of $y_t, ..., y_{t+h}$ is given by matrices $\Psi_q$ and analyzed in terms of impulse response functions (IRF) \(^7\)

$$IRF_h = \frac{\partial y_{t+h-1}}{\partial z_t} = \Psi_{h-1} S, \ h = 1, 2, ...$$

(20)

Note that short run effects of orthogonalized shocks $z_t$ on output parameters $y_t$ at horizon $h = 1$ are represented by the elements of the matrix $S = B_0^{-1}D^{1/2}$, while long-run effects are captured by the cumulative sum of the elements of matrices $\Psi_q S, q = 0, ..., \hat{h}$.

Having estimated by means of impulse response functions the conditional mean of the distribution of $y_t$ the conditional variance of impulse responses

\(^6\)It is implied that the process $y_t$ is strictly indeterministic, namely deterministic component is zero \(\forall t\)

\(^7\)For a discussion of the properties of IRF see in Koop, Pesaran and Potter (1996)
in terms of the variance decomposition can be calculated as
\[ VD_h = \sum_{q=1}^{h} IRF_q \odot IRF_q, \]
(21)
in which \( \odot \) is the Hadamard product. The total variance of each variable \( y_t \) can be found as row sums of each \( VD_h \) with the elements representing the contribution of each of the zones to the total variance.

Volatility reactions to shocks in the foreign exchange, equity, and bond markets are presented in Figures 2, 3 and 4 respectively. The short run shock effects described in these plots are driven the the matrices
\[
\hat{S}^{fx} = \begin{bmatrix}
0.615 & 0 & 0 \\
0.169 & 0.557 & 0 \\
0.131 & 0.175 & 0.629
\end{bmatrix},
\hat{S}^{eq} = \begin{bmatrix}
0.581 & 0 & 0 \\
0.093 & 0.604 & 0 \\
0.104 & 0.184 & 0.565
\end{bmatrix},
\hat{S}^{bd} = \begin{bmatrix}
0.638 & 0 & 0 \\
0.030 & 0.528 & 0 \\
0.046 & 0.085 & 0.571
\end{bmatrix},
\]
(22)

Figure 2: Impulse response functions for the foreign exchange market.
Figure 3: Impulse response functions for the equity market.

Figure 4: Impulse response functions for the bond market.
The impulse responses are not particularly informative, but a number of general conclusions do emerge.

1. The standardised shock always has a pronounced instantaneous positive effect in the zone of origin.

2. The effect of a shock in non-origin zones is much smaller and also consistent with the calendar structure of the trading day. This is particularly noticeable in the effect of the United States zone on Japanese and European zones.

3. The persistence of the shocks is minimal with the major effect dying out in a matter of days.

4. The minimal impact of Japanese shocks on the other zones is apparent in all of the diagrams and the decoupling of the Japanese bond market from external volatility influences is dramatically illustrated in the first column of Figure 4.

Now consider the results of the variance decomposition. The structure of variance decomposition in the short run is presented in Table 9 for all three markets. The main factor of the variance for all three markets is domestic volatility, meaning country specific news as the main source of market volatility in the zone. The main driving factor of volatility in the foreign exchange market is United States volatility and for the bond market is Japanese volatility. The patterns for variance decomposition in the long run are similar to the short run case and for this reason they are not discussed.

The results of the structural VAR analysis aimed at exploring the transmission of volatility between global trading zones can be summarised succinctly as follows. The pattern of volatility transmission is neither a simple meteor
<table>
<thead>
<tr>
<th></th>
<th>Japan</th>
<th>Europe</th>
<th>U.S.</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Foreign exchange</strong></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Japan</td>
<td>0.387</td>
<td>0</td>
<td>0</td>
</tr>
<tr>
<td>Europe</td>
<td>0.031</td>
<td>0.316</td>
<td>0</td>
</tr>
<tr>
<td>U.S.</td>
<td>0.018</td>
<td>0.034</td>
<td>0.415</td>
</tr>
<tr>
<td><strong>Equity</strong></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Japan</td>
<td>0.356</td>
<td>0</td>
<td>0</td>
</tr>
<tr>
<td>Europe</td>
<td>0.008</td>
<td>0.379</td>
<td>0</td>
</tr>
<tr>
<td>U.S.</td>
<td>0.011</td>
<td>0.033</td>
<td>0.335</td>
</tr>
<tr>
<td><strong>Bond</strong></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Japan</td>
<td>0.445</td>
<td>0</td>
<td>0</td>
</tr>
<tr>
<td>Europe</td>
<td>0.001</td>
<td>0.293</td>
<td>0</td>
</tr>
<tr>
<td>U.S.</td>
<td>0.003</td>
<td>0.012</td>
<td>0.359</td>
</tr>
</tbody>
</table>

Table 9: Short-run (one-period) variance decomposition of the realised volatility in each of the foreign exchange, equity, and bond markets.

...shower nor a heatwave, but a mixture of these processes. The effect dominant (positive) effect on domestic volatility comes from a domestic shock and the impulse response analysis reveals that the fluctuations recede fairly quickly. The effect of the standardised shock causes an positive increase of about 0.7% in the country of origin. This is consistent with the results of Engle, Ito and Lin (1990) who find an immediate impact of less than less than 1% in response to domestic volatility shocks.

6 A General Model of Volatility Interaction

A model capable of analysing volatility patterns across both international trading zones and between financial markets simultaneously is now proposed. Essentially the VAR of Section 4 and the structural VAR of Section 5 are
combined in an unrestricted model. This general model is given by

$$B_0Y_t = \sum_{p=1}^{P} B_p Y_{t-p} + u_t, \quad u_t \sim iid \ N(0, D)$$ (23)

in which

$$B_0 = \begin{bmatrix}
1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\
-\alpha_{21} & 1 & 0 & -\alpha_{24} & 0 & 0 & -\alpha_{27} & 0 & 0 \\
-\alpha_{31} & -\alpha_{32} & 1 & -\alpha_{34} & -\alpha_{35} & 0 & -\alpha_{37} & -\alpha_{38} & 0 \\
0 & 0 & 0 & 1 & 0 & 0 & 0 & 0 & 0 \\
-\alpha_{51} & 0 & 0 & -\alpha_{54} & 1 & 0 & -\alpha_{57} & 0 & 0 \\
-\alpha_{61} & -\alpha_{62} & 0 & -\alpha_{64} & -\alpha_{65} & 1 & -\alpha_{67} & -\alpha_{68} & 0 \\
0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\
-\alpha_{81} & 0 & 0 & -\alpha_{84} & 0 & 0 & -\alpha_{87} & 1 & 0 \\
-\alpha_{91} & -\alpha_{92} & 0 & -\alpha_{94} & -\alpha_{95} & 0 & -\alpha_{97} & -\alpha_{98} & 1
\end{bmatrix}, \quad Y_t = \begin{bmatrix}
f_{x,\text{jp},t} \\
f_{x,\text{eu},t} \\
f_{x,\text{us},t} \\
f_{e,\text{jp},t} \\
f_{e,\text{eu},t} \\
f_{e,\text{us},t} \\
f_{b,\text{jp},t} \\
f_{b,\text{eu},t} \\
f_{b,\text{us},t}
\end{bmatrix},$$

the matrices $B_p, \ p \geq 1$ are parameter matrices for lag $p$ and $u_t$ is a vector of non-correlated disturbances with covariance matrix $D$.

The upper left, middle, and lower right shaded blocks of the matrix $B_0$ highlight coefficients describing the behaviour of contemporaneous volatility interaction in the foreign exchange, equity, and bond markets. As in Section 5, the structure of these matrices incorporate the calendar restrictions imposed by the definition of the global trading day. Each of these matrices corresponds the matrix of contemporaneous effects estimated in Section 5 as separate entities for each market. The main innovation in this general model is in the off-diagonal coefficient blocks which now describe the contemporaneous effects from all of the other asset markets in all the trading zones which the single-market analysis of Section 5 ignored. For example, the coefficient $\alpha_{51}$ measures the contemporaneous influence of the Japanese foreign exchange market on the European equity market. Similarly, $\alpha_{62}$ measures the contemporaneous effect from the European foreign
exchange market on the United States equity market. It is important to remember that the ‘contemporaneous’ effect from the United States (to a lesser extent Europe) to Japan will be captured by the relevant elements of the matrix $B_1$. That is, events in the United States and Europe can only effect Japan at the opening of the following global trading day.

The parameters $\theta = \{B_0, \ldots, B_5\}$ of the system of equations (23) is estimated by maximum likelihood for $P = 5$ lags using the same procedure as outlined in Section 5. The results are reported for $B_0$ and $B_1$ only as this is where the major interest lies.

Estimates of matrix $B_0$, with stars indicating the significance of individual coefficients at the 5% level, are as follows:

$$B_0 = \begin{bmatrix}
1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\
-0.26^* & 1 & 0 & -0.03 & 0 & 0 & -0.03 & 0 & 0 \\
-0.09^* & -0.28^* & 1 & -0.08^* & -0.02 & 0 & -0.05 & 0.01 & 0 \\
0 & 0 & 0 & 1 & 0 & 0 & 0 & 0 & 0 \\
-0.07^* & 0 & 0 & -0.16^* & 1 & 0 & -0.00 & 0 & 0 \\
0.02 & -0.04 & 0 & -0.13^* & -0.29^* & 1 & 0.01 & -0.07^* & 0 \\
0 & 0 & 0 & 0 & 0 & 0 & 1 & 0 & 0 \\
-0.04 & 0 & 0 & 0.01 & 0 & 0 & -0.04 & 1 & 0 \\
-0.02 & -0.02 & 0 & -0.03 & -0.11^* & 0 & -0.04 & -0.11^* & 1 \\
\end{bmatrix},$$

The first thing to note is that the contemporaneous volatility patterns between the zones for a particular market for the general model (shaded areas) are similar to the results presented in Section 5 (shaded panels of Tables 6, 7, 8 for the foreign exchange, equity, and bond markets respectively). What is apparent, however, is that coefficient values reported here are slightly smaller than the corresponding values reported in Tables 6, 7, 8. This accords with intuition: adding additional linkages in the general model reduces the size of existing coefficient values. The two insignificant coefficients in the
shaded regions represent the contemporaneous influence from the Japanese bond market to the European and United States bond markets. This is not surprising given the minimal impact of Japanese bond market shocks in Figure 4.

Most of the coefficients in the non-shaded panels of the matrix $B_0$ are insignificant, which means that contemporaneous effects from other asset markets is not strong. However, there are four significant coefficients, which are:

1. $\alpha_{34}$, the effect of the Japanese equity market on the United States foreign exchange market;
2. $\alpha_{51}$, the effect of the Japanese foreign exchange market on the European equity market;
3. $\alpha_{68}$, the effect of the European bond market on the United States equity market; and
4. $\alpha_{95}$, the effect of the European equity market on the United States bond market.

Taken together with the previous results for the shaded blocks which show a strong and significant pattern of influence from Europe to the United States in each of the foreign exchange, equity and bond markets ($\alpha_{32}$, $\alpha_{65}$ and $\alpha_{98}$, respectively), the overall pattern that seems to emerge is one in which the developments in European markets have a significant influence on what happens in the United States markets later on the same day.

The impact on current volatility from developments on the previous global trading day, is given by the coefficients of the matrix $B_1$. Parameter estimates for $B_1$, with stars indicating significance at the 5% level, are as
follows:

\[
\begin{bmatrix}
0.13^* & 0.14^* & 0.20^* & -0.03 & 0.06 & 0.06 & -0.02 & -0.01 & -0.04 \\
0.06 & 0.15^* & 0.17^* & 0.02 & -0.01 & 0.01 & 0.01 & 0.06 & 0.02 \\
0.08^* & 0.06 & 0.20^* & -0.05 & 0.00 & 0.02 & -0.02 & 0.03 & 0.02 \\
0.03 & -0.03 & 0.04 & 0.25^* & 0.02 & 0.13^* & 0.06^* & 0.04 & -0.05 \\
-0.00 & -0.01 & -0.02 & 0.01 & 0.28^* & 0.26^* & -0.06^* & 0.04 & -0.00 \\
0.08^* & -0.05 & 0.02 & 0.07^* & 0.06 & 0.44^* & -0.06^* & 0.03 & -0.01 \\
0.01 & 0.04 & -0.06 & 0.01 & 0.03 & 0.10^* & 0.32^* & 0.00 & -0.03 \\
0.05 & 0.03 & 0.04 & -0.03 & 0.03 & 0.04 & 0.01 & 0.22^* & 0.12^* \\
0.00 & -0.02 & 0.02 & 0.01 & 0.00 & 0.08^* & -0.04 & 0.11^* & 0.21^*
\end{bmatrix}
\]

The diagonal blocks shaded in light tray indicate the effect of lagged volatility within and between trading zones, but limited to a single market. The top left block is lagged volatility the foreign exchange market, the middle block is lagged volatility in the equity market and the bottom right block is lagged volatility the bond market.

The individual cells that are shaded a slightly darker grey indicate the ‘contemporaneous’ effect of the United States markets on Japanese markets. Three of these shaded cells are statistically significant

1. \( \beta_{13} \), the influence of the United States foreign exchange market on the Japanese foreign exchange market;

2. \( \beta_{46} \), the influence of the United States equity market on the Japanese equity market; and

3. \( \beta_{76} \), the influence of the United States equity market on the Japanese bond market.

The latter effect is particularly interesting as it is not captured by the analysis in Section 5 and appears to be a strong and significant effect. Taken
together, these results provide significant evidence that news in the United States has a pervasive influence on the opening of Japanese markets on the subsequent trading day.

To sum up the case for contemporaneous interaction between markets and trading zones, there are two broad conclusions. First, there is compelling evidence for a meteor shower pattern in which volatility in one trading zone is driven by events in the zone that immediately precedes it. This is particularly significant in terms of the transition from Europe to the United States to Japan. The one proviso to this is that the Japanese bond market does not play any role in influencing events in Europe and the United States. Second, this effect is not merely a market based phenomenon. There are enough significant coefficients outside of the diagonal blocks to suggest that the meteor shower pattern also occurs between markets for different assets.

Another important pattern evident in the parameters of $B_1$ is that most of the significant coefficients are to be found in the shaded diagonal blocks, a heatwave pattern. This pattern, however, just as in the case of the meteor shower, is not confined to a single market. In particular, the significance of the coefficient $\beta_{12}$ confirms the findings of Hong (2001) concerning the linkage between lagged volatility in the European and Japanese foreign exchange markets. Overall, the results are very similar to those reported in Section 5 in terms of the numbers of significant coefficients. The fact that most of the significant coefficients are to be found in these shaded blocks also supports the very weak evidence for Granger causality found in Section 4, particularly in Table 5, which finds no Granger causality between the different markets in the individual trading zones. Perhaps the patterns of causality found in Table 4 are due to the restricted nature of the VAR model and the fact that no international linkages are allowed.
Lagged volatility linkages between different markets (coefficients outside of the diagonal blocks) are not particularly strong. Once interesting observation concerns the parameters $\beta_{47}$, $\beta_{57}$ and $\beta_{67}$ which represent the influence of lagged volatility in the Japanese bond market on all the equity markets. It is true that this effect is small in magnitude but does appear to be statistically significant. Once again this emphasises that volatility linkages are particularly complex and not simple explanation is available.

7 Conclusion

An enormous amount of research has focused on the issue of volatility transmission through time, either within a country or a specific asset market. This paper considers patterns in realised volatility between the global foreign exchange, equity, and bond markets. Realised volatility estimates were constructed using high frequency data for each asset market and trading zone and the global trading day was divided into distinct trading zones. This marks a significant departure in the literature on volatility transmission because the use of observed estimates of volatility allow traditional time-series techniques, such as VARs and structural VARs, to be used to test hypotheses about volatility linkages.

The major conclusion to emerge from this work is that a series of significant and complex relationships link the different asset markets and trading zones. Furthermore, this interaction defies categorisation in terms of a simple meteor shower or heatwave. There are both significant contemporaneous effects from markets in the zone immediately preceding any given zone (meteor shower) but also significant effects from lagged volatility (heatwave). Moreover, the interaction is not confined to particular markets. Every asset
market is influenced by events in other markets as well as other zones.

If pushed, a tentative conclusion may be that the influence of Japan on Europe and the United States, apart from the foreign exchange market, is muted by comparison with all the other effects that are identified. This suggests that news from the opening in European markets is propagated through that the United States and then very strongly into the Japanese markets at the opening of the new global trading day. This is particularly true of the Japanese bond market, which appears to react to news from the United States but plays no role in propagating these influences any further.

References


Mancini, Cecilia. 2009. Non-parametric Threshold Estimation for Models


