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Volatility transmission in global financial markets☆

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ABSTRACT

This paper considers the transmission of volatility in global foreign exchange, equity and bond markets. Using a multivariate GARCH framework which includes measures of realised volatility as explanatory variables, significant volatility and news spillovers are found to occur on the same trading day between Japan, Europe, and the United States. All markets exhibit significant degrees of asymmetry in terms of the transmission of volatility associated with good and bad news. There are also strong links between diffusive volatilities in all three markets, whereas jump activity is only important within the equity markets. The results of this paper deepen our understanding of how news and volatility are propagated through global financial markets.

1. Introduction

The practical importance of modelling the volatility of financial assets has given rise to a voluminous body of research. Much of the modern literature stems from the seminal work of Engle (1982) and Bollerslev (1986) who treat volatility as an unobserved quantity. More recent developments treat volatility as a realised (observed) variable, which is estimated from the squared returns of high-frequency financial asset returns (Andersen et al., 2003; Barndorff-Nielsen and Shephard, 2002). Hansen et al. (2012) provide one avenue for combining these two approaches. The vast majority of this work relates to modelling and forecasting the trajectory of the volatility of an asset in one particular market.

A relatively recent but growing area of interest addresses the important question of how volatility is propagated from one region of the world to another. The series of papers by Ito (1987), Ito and Roley (1987) and Engle et al. (1990) examine how volatility is transmitted through different regions of the world during the course of a global financial trading day. Their approach is to partition each 24 hour period (calendar day) into four trading zones, namely, Asia, Japan, Europe and the United States, and examine international linkages in volatility between these regions in the context of the foreign exchange market. Engle et al. (1990) describe two particular patterns, namely, the *heat wave* in which volatility in one region is primarily a function of the previous day's volatility in the same region, and the *meteor shower* in which volatility in one region is driven by volatility in the region immediately preceding it in

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terms of calendar time. Their major conclusion is that volatility in the foreign exchange market is best described by the meteor shower pattern. Using a similar research protocol, Fleming and Lopez (1999) and Savva et al. (2005) find that the heat wave hypothesis best describes the behaviour of volatility in the bond and equity markets, respectively. Treating volatility as an observed measure, Melvin and Melvin (2003) investigate the transmission of volatility in the foreign exchange market using realised volatility. While there is some support for both meteor shower and heat wave hypotheses, evidence marginally favours the latter.[1]

This paper contributes to the literature examining volatility patterns in global foreign exchange, equity, and bond markets. In the spirit of Fleming and Lopez (1999), the trading day for each market is partitioned into three zones, Japan, Europe, and the United States. The original Engle et al. (1990) model is extended in order to entertain a more complex set of volatility interactions. Incorporating realised volatility as an explanatory variable in the regional conditional variance equations turns out to be a valuable way of investigating volatility linkages and allows a number of extensions. Using jump and continuous components of realised volatility reveals that the diffusive component has more explanatory power than the jump component. The jump component which reflects more extreme news arrival is only significant for the equity market. Furthermore, decomposing realised volatility into positive and negative semivariances allows the asymmetric transmission of volatility to be explored. It is found that the bond market is mainly driven by a negative semivariance, whereas both negative and positive semivariances have an explanatory power for the foreign exchange and equity markets.

The rest of the paper proceeds as follows. Section 2 describes the data and justifies the structure of the calendar day. Section 3 revisits the original Engle et al. (1990) framework and re-examines their results in the context of a different sample period and for a broader selection of financial markets. Section 4 extends the GARCH framework to try and separate the effects of the news and smoothed conditional variances on volatility transmission between regions. Section 5 constructs the measures of realised volatility in each of the trading zones and decomposes the basic measure into a number of constituent components which are used later in the estimation. Section 6 introduces the realised volatility as an explanatory factor to describe volatility interactions across the three regions and examines how different constituents of realised volatility contributed to our understanding of the global transmission of volatility. Section 7 is a brief conclusion.

2. Data

A high frequency (10 min) data set was gathered from Thomson Reuters Tick History for the period from 3 January 2005 to 28 February 2013 for instruments representative of the foreign exchange, bond and equity markets. Specifically, the following series were collected:

1. Euro–Dollar United States futures contracts traded on the Chicago Mercantile Exchange (foreign exchange market);
2. United States 10 year Treasury bond futures contracts (one of the most traded securities in the bond market (Fleming and Lopez, 1999));
3. S&P 500 futures contracts (equity markets).

Each of these instruments is traded continuously (23 h per calendar day). Following the standard approach in the literature (Andersen et al., 2003) days where one market is closed are eliminated, as are public holidays or other occasions when trading is significantly curtailed.[2]

This continuously-traded, high-frequency data on futures contracts is used to construct returns to the instruments in each of three trading zones. The protocol of Fleming and Lopez (1999) to delimit the global trading day is adopted, in which the Japanese trading zone is defined as 12 am to 7 am GMT, the European trading zone is taken to be 7 am to 12:30 pm GMT and the United States zone is 12:30 pm to 9 pm GMT.

The setup may be illustrated as follows:

$$\underbrace{\overbrace{12\,\text{am}\cdots 7\,\text{am}}^{\text{Japan}} \quad \overbrace{7\,\text{am}\cdots 12:30\,\text{pm}}^{\text{Europe}} \quad \overbrace{12:30\,\text{pm}\cdots 9\,\text{pm}}^{\text{U.S.}}}_{\text{One Trading Day}}$$

Note that the period denoted as Asian trading (2 h prior to Japan opening) by Engle et al. (1990) is excluded here because very little trading activity occurs during this time. This lack of volume during this period means that little reliable high-frequency data is available to construct realised volatility estimates.

A second and possibly more important point concerns the decision to start the United States zone at 12:30 pm GMT which is 7.30 am in the eastern United States. This choice is motivated by Dungey et al. (2009), who define an additional trading zone, 12:30 pm to 2:15 pm GMT, immediately prior to the market opening in the United States. Although this period overlaps with late trading in Europe, it is associated with a significant increase in trading activity immediately before pit trading in Chicago begins

[1] Recent research has also documented significant linkages in volatilities between different financial markets within a particular region. See for example, Hakim and McAleer (2010), Bubák et al. (2011), Ehrmann et al. (2011) and Engle et al. (2012).

[2] Each of these data series may be regarded as U.S. centric in the sense that they relate to U.S. based assets. However, in order to check the robustness of the volatility patterns to the choice of assets, the analysis was repeated using both Japanese (Topix equity futures and JGB bond futures) and German (DAX equity futures and BUND bond futures) data. While the results of this parallel exercise are not presented in this paper, they are broadly similar to the results reported here. An online appendix together with all the data is available for download from http://www.ncer.edu.au/data/data.jsp.

and therefore it is reasonable to assume that the increased trading activity is related to United States news. Consequently, this period is simply subsumed into the United States trading zone.

These two decisions concerning the construction of trading zones are supported by the plots in Fig. 1 which illustrate the diurnal pattern in average trading volume in the foreign exchange, equity and bond markets over the sample period 3 January 2005 to 28 February 2013. The figure reveals similar patterns for all three markets: low volume during the Japanese trading hours, an increase in trading activity after the opening in Europe (7:00 GMT) and a significant increase in volume after the beginning of trading in the United States (12:30 GMT). The shaded regions represent the 12:30 pm to 2:15 pm GMT pre-opening period showing the dramatic increase in trading activity in anticipation of the opening of the Chicago market.

Following Engle et al. (1990) the return in each zone is calculated as the difference between the last and the first transaction price within the same calendar day which is normalised by the length of the day. The returns can be expressed as

$$R_t^i = \left(\ln PC_t^i - \ln PO_t^i\right) / \sqrt{T_t^i}, \tag{1}$$

in which, PC_t^i is a closing price in zone i on day t, and PO_t^i is an opening price in zone i on day t, and T_t^i is the duration of trading in zone i on day t measured in hours. Returns in each region are computed for each of the 1976 trading days.

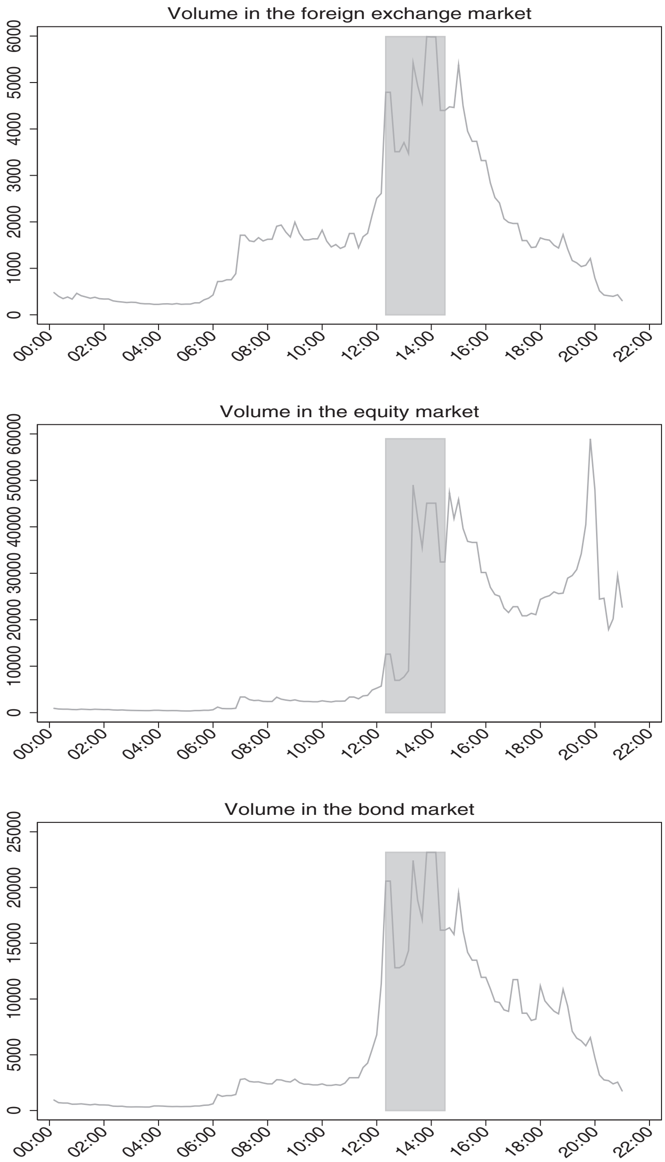

Fig. 1. Average volume in S&P 500, United States 10 year Treasury bond, and Euro–Dollar futures contracts for the period 3 January 2005 to 28 February 2013. The horizontal axis is GMT and the pre-opening trading period (12:30 to 14:15 GMT) in the United States is highlighted.

Descriptive statistics for the returns from each zone are presented in Table 1. Consistent with prior expectations, bond markets have lower average returns and volatility in terms of standard deviation than equity returns. While bond returns during Japanese trading and equity returns in the United States exhibit negative means, they are not significantly different from zero given the large degree of volatility. Turning to the higher moments, none of the series exhibit large degrees of skewness. On the other hand, all markets exhibit excess kurtosis, with equity returns being by far the most kurtotic with the exception of bond returns in the Japanese zone. Volatility clustering is a well known empirical phenomenon in financial asset returns. While the results are not reported here, the standard test for ARCH effects of Engle (1982) indicates that all the series demonstrate strong ARCH effects at the 5% level.

3. Revisiting heat waves and meteor showers

In their seminal paper, Engle et al. (1990) find evidence that the role of news from adjacent regions (the meteor shower) is to be preferred to local influences from the previous day (the heat wave) as an explanation of the transmission of volatility in the foreign exchange market. In this section the results of Engle et al. (1990) for the foreign exchange market are revisited in the context of the current data 2005 to 2013, and extended to the equity and bond markets.

The original model proposed by Engle et al. (1990) applies to $i = 1, \cdots, n$ non-overlapping trading zones and takes the form

$$R_t^i = \epsilon_t^i \quad \epsilon_t^i \sim N\left(0, h_t^i\right) \tag{2}$$

$$h_t^i = \kappa_i + \alpha_{ii} h_{t-1}^i + \sum_{j=1}^{i-1} \beta_{ij} \epsilon_{j,t}^2 + \sum_{j=i}^{n} \gamma_{ij} \epsilon_{j,t-1}^2, \tag{3}$$

in which R_t^i is a time-series of returns, h_t^i is the conditional variance of returns all defined for zone i at time t, and $\epsilon_{j,t}^2$ is the squared innovation (news) defined for zone j at time t. This specification deviates from a traditional GARCH model by recognising that the structure of the global trading day allows for news from preceding regions to influence volatility on the same trading day, a feature which is labelled the 'intra-day' effect.

The calendar structure implied by Eq. (3) is perhaps best seen by writing the equation in matrix form and tailoring it to the current modelling environment in which $n = 3$. The relevant equation is now

$$h_t = K + A h_{t-1} + B \epsilon_t^2 + G \epsilon_{t-1}^2 \tag{4}$$

where $h_t = [h_{\text{jp},t}\, h_{\text{eu},t}\, h_{\text{us},t}]'$, $\epsilon_t^2 = [\epsilon_{\text{jp},t}^2\, \epsilon_{\text{eu},t}^2\, \epsilon_{\text{us},t}^2]'$ and the subscripts are self evident. The parameter matrices of interest are

$$A = \begin{bmatrix} \alpha_{11} & 0 & 0 \\ 0 & \alpha_{22} & 0 \\ 0 & 0 & \alpha_{33} \end{bmatrix}, \quad B = \begin{bmatrix} 0 & 0 & 0 \\ \beta_{21} & 0 & 0 \\ \beta_{31} & \beta_{32} & 0 \end{bmatrix}, \quad G = \begin{bmatrix} \gamma_{11} & \gamma_{12} & \gamma_{13} \\ 0 & \gamma_{22} & \gamma_{23} \\ 0 & 0 & \gamma_{33} \end{bmatrix}.$$

The calendar structure is now apparent, particularly in the matrix B. Specifically, new developments in Japan, $\epsilon_{\text{jp},t}^2$, at the start of the trading day can potentially influence volatility in Europe and the United States via the coefficients β_{21} and β_{31}. Similarly news from Europe, $\epsilon_{\text{eu},t}^2$, can influence volatility in the United States on the same day, β_{32}. The natural calendar structure, however, implies that events in the United States will be transmitted to Japan only on the following day. The restrictions on the matrix G making it an upper diagonal matrix are not strictly necessary as all information at $t - 1$ can affect all trading zones. These restrictions are imposed by Engle et al. (1990) in their original formulation implying that information originates during United States trading times. The system (4) is characterised by an information matrix which is block diagonal with respect to the required parameters. For this reason single equation estimation of the model by the maximum likelihood can be performed on each zone.

Table 1
Descriptive statistics multiplied by 1000 for daily estimates of the returns from Eq. (1) in the foreign exchange, equity and bond markets in Japan, Europe and the United States.

		Mean	St.dev.	Min.	Max.	Skew.	Kurt.
FX	Japan	0.0319	1.0116	−4.8511	5.5888	−0.0141	5.6395
	Europe	−0.0715	1.5607	−10.6287	6.7892	−0.2929	5.8742
	U.S.	0.0466	1.5629	−7.7562	9.6225	0.0490	5.3783
Equity	Japan	0.0650	1.5851	−19.1890	18.1024	−0.0953	37.6800
	Europe	0.0605	2.0833	−13.2551	15.5146	0.0556	10.2371
	U.S.	−0.0660	3.9423	−31.4683	26.8688	−0.6466	12.7934
Bond	Japan	−0.0051	0.6633	−4.9703	4.2260	−0.5978	14.0123
	Europe	0.0182	0.7926	−4.0205	3.6861	−0.1086	5.2929
	U.S.	0.0353	1.0977	−5.5323	5.5262	−0.1329	4.8355

Table 2
Coefficient estimates of Eq. (4) with t-statistics based on QML standard errors in parentheses. (* denotes significance at the 5% level).

		Japan	Europe	United States
FX market	$\epsilon^2_{jp,t}$	–	0.0441* (2.50)	0.0620* (3.63)
	$\epsilon^2_{eu,t}$	–	–	0.0019 (0.30)
	$\epsilon^2_{us,t}$	–	–	–
	$\epsilon^2_{jp,t-1}$	0.0235* (2.55)	–	–
	$\epsilon^2_{eu,t-1}$	0.0018 (0.62)	0.0228* (2.75)	–
	$\epsilon^2_{us,t-1}$	0.0119* (2.77)	0.0179* (2.19)	0.0257* (2.85)
	h^i_{t-1}	0.9431* (59.4)	0.9294* (57.6)	0.9404* (63.2)
Equity market	$\epsilon^2_{jp,t}$	–	0.0784* (2.68)	0.3225* (3.25)
	$\epsilon^2_{eu,t}$	–	–	0.3102* (4.62)
	$\epsilon^2_{us,t}$	–	–	–
	$\epsilon^2_{jp,t-1}$	0.0583* (2.44)	–	–
	$\epsilon^2_{eu,t-1}$	0.0301* (3.84)	0.0754* (5.94)	–
	$\epsilon^2_{us,t-1}$	0.0081* (3.23)	0.0141* (2.60)	0.1056* (5.03)
	h^i_{t-1}	0.8271* (25.9)	0.8546* (39.9)	0.7110* (27.1)
Bond market	$\epsilon^2_{jp,t}$	–	0.0094 (1.17)	0.0027 (0.43)
	$\epsilon^2_{eu,t}$	–	–	0.0101 (1.35)
	$\epsilon^2_{us,t}$	–	–	–
	$\epsilon^2_{jp,t-1}$	0.2596* (5.41)	–	–
	$\epsilon^2_{eu,t-1}$	0.0379* (3.31)	0.0774* (5.07)	–
	$\epsilon^2_{us,t-1}$	0.0428* (4.06)	0.0143* (2.74)	0.0394* (4.63)
	h^i_{t-1}	0.6207* (17.3)	0.8890* (41.4)	0.9505* (94.5)

Table 2 reports the estimation results for Eq. (4) based on the foreign exchange, equity and bond markets.[3] To ensure robust inference, t-statistics based on quasi-maximum likelihood (QML) standard errors are reported. There are two general conclusions that emerge from inspection of these results.

1. It is immediately apparent that the pattern of volatility interaction is a combination of both heat waves and meteor showers. There is no support for the hypothesis that one of these patterns dominates, although the argument in favour of a heat wave is probably strongest in the bond market, where intra-day news has no explanatory effect on volatility in any of the trading zones.
2. There is a significant impact of the previous day's volatility in each region, with all the coefficients on the lagged conditional variance term, h^i_{t-1}, being significant. It is probably fair to say that this impact is uniformly greatest in the foreign exchange market. Moreover, the sizes of the coefficients on h^i_{t-1} in the foreign exchange market are larger than those of Engle et al. (1990). Perhaps this difference is attributable to the different periods that are used in the analysis. Engle et al. (1990) use data from 1985 to 1986, while this paper deals with data from 2005 to 2013, a period dominated by the turbulence induced by the global financial crisis.

Rather surprisingly, the intra-day impact of European news on the United States volatility, $\epsilon^2_{eu,t}$, is only significant at the 5% level in the equity market. This may perhaps be due to the fact that Japanese news has a strong intra-day effect on both Europe and the United States and this crowds out the effect of European news. In the equity market all the coefficients across all the regions are significant. This pattern of interaction suggests that world equity markets are strongly interrelated. This result in the equity market contrasts sharply with the bond market and supports the results of Fleming and Lopez (1999) and Savva et al. (2005), who find that the heat wave hypothesis best describes the behaviour of volatility in the bond market.

These results vindicate the original insight of Engle et al. (1990), made in the context of the foreign exchange market, to the effect that the transmission of news between different regions of the world on the same trading day is a potentially important explanation of

[3] In Table 2 and all subsequent results, the constant term in the variance equation is suppressed.

volatility. The result stands the test of time, from the perspectives of different sample periods and different markets and provides the motivation to pursue this avenue of inquiry beyond the confines of the original study.

4. Volatility spillovers and news

It has been shown that volatility in the current zone depends on current and lagged innovations from the preceding zones and also lagged conditional variance in the current zone. However, this specification does not allow for an intra-day volatility effect, h_t^i, from the immediately preceding zone to influence volatility in the current zone. This effect may be labelled fairly loosely as a volatility spillover. To take this into account the model must be specified as follows

$$\widetilde{A}h_t = K + Ah_{t-1} + B\epsilon_t^2 + G\epsilon_{t-1}^2, \tag{5}$$

in which the matrix $\tilde{A}$ now captures the effect of the conditional variance in preceding zones on the current zone on the same trading day. In the case of the three zones dealt with here, the two A matrices are

$$\widetilde{A} = \begin{bmatrix} 1 & 0 & 0 \\ -\widetilde{\alpha}_{21} & 1 & 0 \\ 0 & -\widetilde{\alpha}_{32} & 1 \end{bmatrix}, \quad A = \begin{bmatrix} \alpha_{11} & 0 & \alpha_{13} \\ 0 & \alpha_{22} & 0 \\ 0 & 0 & \alpha_{33} \end{bmatrix},$$

which emphasises that the current conditional variance term appropriate for Japan is in fact the lagged conditional variance from the previous days close in the United States with coefficient α_{13}. This is a system of equations that must be estimated simultaneously.

Table 3
Coefficient estimates of Eq. (5) with t-statistics based on QML standard errors in parentheses. (* denotes significance at the 5% level).

		Japan	Europe	United States
FX market	$h_{jp,t}$	–	0.5410* (88.0)	–
	$h_{eu,t}$	–	–	0.1897* (54.0)
	$h_{us,t-1}$	0.1323* (140)	–	–
	$\epsilon^2_{jp,t-1}$	0.0723* (39.0)	–	–
	$\epsilon^2_{eu,t-1}$	–	0.0264* (8.44)	–
	$\epsilon^2_{us,t-1}$	–	–	0.0535* (16.0)
	h^i_{t-1}	0.6054* (183)	0.7365* (234)	0.7579* (410)
Equity market	$h_{jp,t}$	–	0.3424* (25.0)	–
	$h_{eu,t}$	–	–	0.3273* (46.0)
	$h_{us,t-1}$	0.0577* (9.69)	–	–
	$\epsilon^2_{jp,t-1}$	0.2358* (51.0)	–	–
	$\epsilon^2_{eu,t-1}$	–	0.1343* (15.0)	–
	$\epsilon^2_{us,t-1}$	–	–	0.1398* (15.0)
	h^i_{t-1}	0.3266* (138)	0.7273* (121)	0.7466* (75.0)
Bond market	$h_{jp,t}$	–	0.0108* (21.7)	–
	$h_{eu,t}$	–	–	0.0071 (1.52)
	$h_{us,t-1}$	0.0361* (3.25)	–	–
	$\epsilon^2_{jp,t-1}$	0.3254* (40.0)	–	–
	$\epsilon^2_{eu,t-1}$	–	0.0927* (3.85)	–
	$\epsilon^2_{us,t-1}$	–	–	0.0382* (3.50)
	h^i_{t-1}	0.6438* (76.0)	0.8935* (54.0)	0.9545* (585)

As a precursor to the estimation of Eq. (5) a model is estimated in which the restriction $B = 0$ is imposed and therefore there are no intra-day effects of the news in preceding regions. The results are reported in Table 3 from which it is immediately apparent that the pattern of the estimated coefficients is very similar to that in Table 2. For the foreign exchange and equity markets, the inclusion of the conditional variance from the immediately preceding zone has the effect of dramatically reducing the impact of volatility from the preceding trading day (heat wave) as can be ascertained from the size of the estimated coefficient on lagged volatility. Once again,

Table 4
Coefficient estimates of Eq. (5) with t-statistics based on QML standard errors in parentheses. (* denotes significance at the 5% level).

		Japan	Europe	United States
FX market	$h_{jp,t}$	–	0.1352* (4.39)	–
	$h_{eu,t}$	–	–	0.1131* (13.2)
	$h_{us,t-1}$	0.1163* (4.48)	–	–
	$\epsilon^2_{jp,t}$	–	0.0452* (2.25)	–
	$\epsilon^2_{eu,t}$	–	–	0.0000 (0.00)
	$\epsilon^2_{us,t}$	0.0000 (0.00)	–	–
	$\epsilon^2_{jp,t-1}$	0.0683* (2.52)	–	–
	$\epsilon^2_{eu,t-1}$	–	0.0176* (2.56)	–
	$\epsilon^2_{us,t-1}$	–	–	0.0587* (1.97)
	h^i_{t-1}	0.6478* (12.5)	0.8983* (154)	0.8299* (31.0)
Equity market	$h_{jp,t}$	–	0.5691* (77.0)	–
	$h_{eu,t}$	–	–	0.0338 (0.93)
	$h_{us,t-1}$	0.0365* (3.91)	–	–
	$\epsilon^2_{jp,t}$	–	0.1391* (12.0)	–
	$\epsilon^2_{eu,t}$	–	–	0.3887* (24.4)
	$\epsilon^2_{us,t}$	0.0088* (2.35)	–	–
	$\epsilon^2_{jp,t-1}$	0.1331* (7.16)	–	–
	$\epsilon^2_{eu,t-1}$	–	0.0927* (4.92)	–
	$\epsilon^2_{us,t-1}$	–	–	0.0997* (4.06)
	h^i_{t-1}	0.5008* (11.9)	0.6204* (30.1)	0.7707* (85.0)
Bond market	$h_{jp,t}$	–	0.0000 (0.00)	–
	$h_{eu,t}$	–	–	0.0000 (0.00)
	$h_{us,t-1}$	0.0000 (0.00)	–	–
	$\epsilon^2_{jp,t}$	–	0.0192* (2.09)	–
	$\epsilon^2_{eu,t}$	–	–	0.0112 (1.93)
	$\epsilon^2_{us,t}$	0.0546* (5.72)	–	–
	$\epsilon^2_{jp,t-1}$	0.2819* (14.9)	–	–
	$\epsilon^2_{eu,t-1}$	–	0.0818* (3.81)	–
	$\epsilon^2_{us,t-1}$	–	–	0.0402* (10.7)
	h^i_{t-1}	0.6178* (73.0)	0.9027* (182)	0.9502* (333)

volatility in the bond market is most likely to be generated by a heat wave but the ambiguity of this claim is much increased. An important difference in comparison with the previous model is in increased intra-day volatility effects in the foreign exchange market.

There seems little doubt that a correctly specified model of volatility and news transmission between regions must make allowance for the intra-day effect of conditional variance from the preceding zones. The Engle et al. (1990) approach which does not allow for this effect therefore runs the risk of being misspecified.

The results obtained after estimation of the unrestricted version of Eq. (5) are reported in Table 4. The message here is somewhat mixed, particularly in the foreign exchange and bond markets. It appears that in the foreign exchange market making allowance for both volatility and news spillovers from the immediately preceding regions obfuscates the effect of news. Indeed only one of the three possible coefficients is significant. In the case of the bond market all coefficients representing volatility spillovers are insignificant, while news from Japan and the United States are important. The degree of ambiguity is small in the equity market where the coefficients on both volatility and news spillovers are significant except European effects captured by $h_{eu,t}$.

It seems that extending the Engle et al. (1990) framework to allow for intra-day influences from both news and smoothed conditional variances causes somewhat of a conundrum. Although there is strong evidence for each of these channels of influence individually, the current attempt to estimate a general model seems to be unsatisfactory. It may be that too much is being asked in the estimation of a model where the proliferation of coefficients requiring estimation is leading to inefficiency. Ideally what is required now is to be able to include both these effects in a parsimonious framework, a task to which attention is now turned.

5. Computing realised volatility

The ten-minute data from the Thomson Reuters Tick History for the period from 3 January 2005 to 28 February 2013 is now used to construct realised volatility series for each of the financial assets in each of the trading zones. There are a wide variety of estimators of asset price variation constructed from high-frequency data (so-called “realised measures”) available to choose from. Following the evidence presented by Patton et al. (2013) that it is difficult to beat the simple realised variance (RV) estimator, it is this estimator which is used here. To start with, the intra-daily returns at the frequency Δ for each trading zone are calculated as

$$r^i_{j,t}(\Delta) = log\ p^i_{j,t} - log\ p^i_{j-1,t}, \tag{6}$$

in which $p^i_{j,t}$ is the ten-minute price in zone i on the day t and in the time interval j. Once the intra-daily returns are available, a daily realised volatility for each zone i is easily computed as a sum of the squared intra-day returns

$$RV^i_t(\Delta) \equiv \sum_{j=1}^{1/\Delta} \left(r^i_{j,t}\right)^2. \tag{7}$$

In order to ease the notation, the superscript i is suppressed in all the subsequent discussion of realised volatility.

To extract the jump component of realised volatility at the frequency Δ, the minimum realised volatility estimator

$$\mathrm{Min}RV_t(\Delta) \equiv \frac{\pi}{\pi-2}\left(\frac{1}{1-\Delta}\right)\sum_{j=2}^{1/\Delta} min\left(|r_{j,t}|, |r_{j-1,t}|\right)^2 \tag{8}$$

of Andersen et al. (2012) is used. This jump robust estimator provides better finite sample properties than well known bi-power variation of Barndorff-Nielsen and Shephard (2002). Based on the asymptotic results of Barndorff-Nielsen and Shephard (2004,2006) and using the fact that[4]

$$\sqrt{\frac{1}{\Delta}}\left(\mathrm{Min}RV_{t+1} - \int_t^{t+1}\sigma^2(s)\mathrm{d}s\right) \overset{stableD}{\rightarrow} MN\left(0, 3.81\int_t^{t+1}\sigma^4(s)\mathrm{d}s\right),$$

statistically significant jumps are identified according to

$$Z_t(\Delta) \equiv \frac{[RV_t(\Delta) - \mathrm{Min}RV_t(\Delta)]/RV_t(\Delta)}{\left[1.81\Delta\ max\left(1, \mathrm{Min}RQ_t(\Delta)/\mathrm{Min}RV_t(\Delta)^2\right)\right]^{1/2}} \sim N(0,1)$$

where MinRQ is a minimum realised quarticity

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

[4] See propositions 2 and 3 in Andersen et al. (2012).

Significant jumps at an α level of significance are identified as

$$J_t(\Delta)(Z) \equiv 1[Z_t(\Delta) > \Phi_{1-\alpha}] \cdot [RV_t(\Delta) - \text{Min}RV_t(\Delta)]. \tag{9}$$

Following Huang and Tauchen (2005) and Evans (2011), the level of significance $\alpha = 0.999$ is chosen. The continuous component can be consistently estimated by minimum realised volatility MinRV_t, but to maintain the property that the sum of the continuous and jump components is equal to realised volatility, the continuous component is defined by

$$CC_t(\Delta) = RV_t(\Delta) - J_t(\Delta)(Z). \tag{10}$$

The volatility and jump estimates for the foreign exchange, equity, and bond markets calculated using Eqs. (7) and (9) are presented in Figs. 2, 3 and 4 respectively.

To the naked eye it appears that the estimates of realised volatility in foreign exchange market have similar patterns across the trading zones. The volatility in the United States is perhaps a little more pronounced during the Global Financial Crisis period of 2007–2009. However, the similarity across the three zones is not as pronounced in the equity and bond markets. Fig. 3 indicates that while realised volatility in the European and the United States equity markets is very similar, Japanese volatility for this market is much lower and less prone to jump activity. Fig. 4 shows that realised volatility in the bond market during Japanese trading hours appears to behave differently to the other zones.

Table 5 reports the summary statistics for the realised volatility series. On average it seems that the mean level of volatility in the equity market is greater than that in the foreign exchange market which in turn is greater than the mean volatility in the bond market. Across the three markets, the United States zone consistently experiences higher mean volatility than Europe which is in turn larger than Japan. Engle et al. (1990) find that 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 from macroeconomic announcements released during New York trading hours. The results in Table 5 support the notion that United States volatility is uniformly higher. This finding is also confirmed by the skewness and kurtosis statistics indicating a marked difference from what would be expected from a normal distribution.

As a quick consistency check on the accuracy of the procedure for isolating the jump component of realised volatility, days which exhibited the greatest volatility are presented in Table 6. It is useful to try and ascertain whether the largest jumps actually correspond to important events in the relevant markets and in this way establish the internal consistency of the method. The largest levels of volatility occurred in the equity market in the last quarter of 2008, which can be related to the bankruptcy of Lehman Brothers and the bridging loan from the Federal Reserve to the world largest insurance company A.I.G. The highest volatility in the crisis period was experienced in the equity market with the major effect in the United States. Another interesting event that significantly affects the equity market occurred in the middle of August 2011.

During August 2011 markets fell on fears of contagion of the European sovereign debt crisis to Spain and Italy leading to a fall in the S&P 500 of 79.92 points (6.7%). As a result, European and Japanese equity markets experienced their third and fourth largest volatility

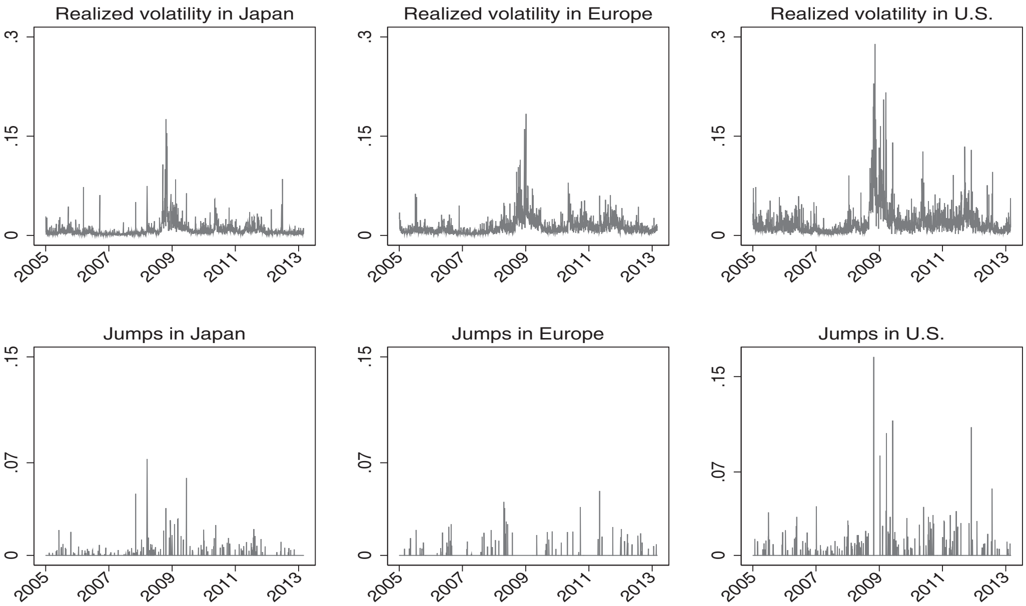

Fig. 2. Realised volatility and jump estimates (scaled by 1000) for the foreign exchange market during Japanese, European, and United States trading hours, respectively, for the period 3 January 2005 to 28 February 2013.

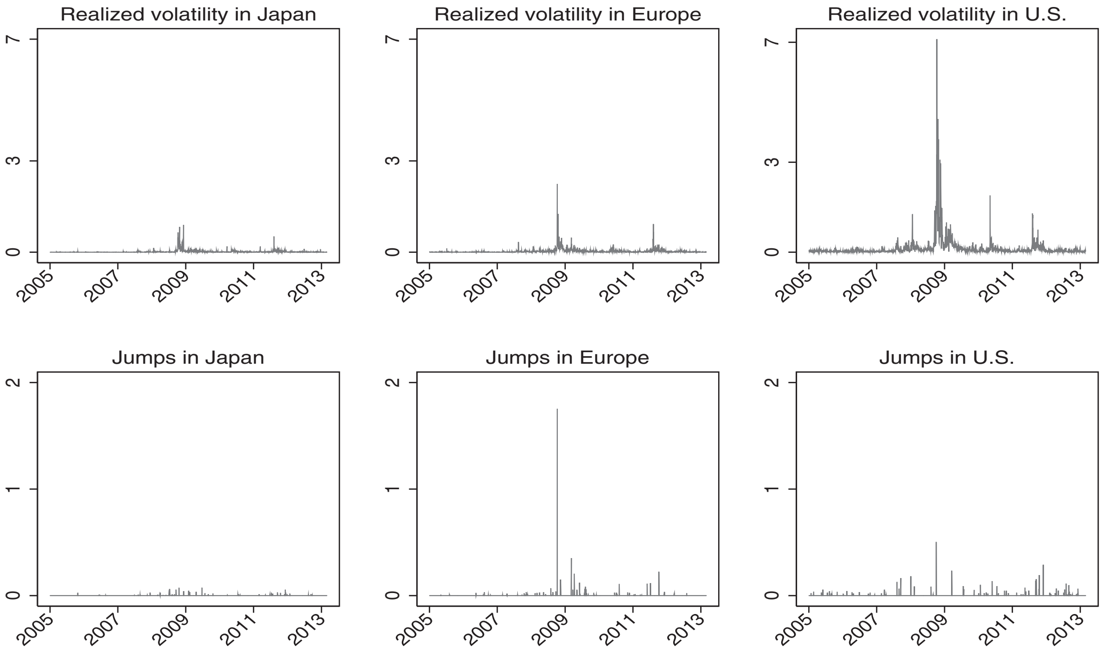

Fig. 3. Realised volatility and jump estimates (scaled by 1000) for the equity market during Japanese, European and United States trading hours, respectively, for the period 3 January 2005 to 28 February 2013.

episodes on 9 August 2011. In the bond market, by far the second largest crisis appears to have been the decision by the United States Treasury borrowed over $1 trillion in September 2008.

There is one final manipulation of the realised volatility series which proves useful and this relates to the asymmetric transmission of volatility relating to good and bad news. Note that this effect should not be confused with the well established leverage effect in equity markets. The leverage effect allows for asymmetric impacts on volatility due to bad and good news of an identical size. In the multivariate context, the asymmetric BEKK model of Kroner and Ng (1998) and the matrix exponential GARCH of Kawakatsu (2006) are the most popular models to capture such asymmetric effects.

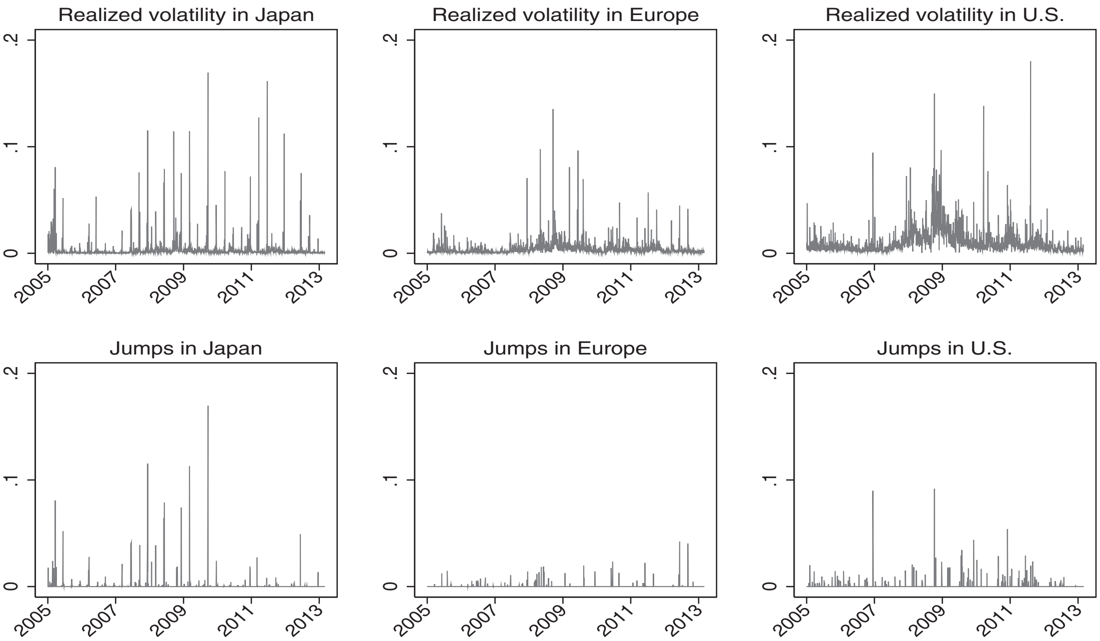

Fig. 4. Realised volatility and jump estimates (scaled by 1000) for the bond market during Japanese, European and United States trading hours, respectively, for the period 3 January 2005 to 28 February 2013.

Table 5
Descriptive statistics multiplied by 1000 for daily estimates of the realised volatility in the foreign exchange, equity and bond markets in Japan, Europe and the United States.

		Mean	St.dev.	Min.	Max.	Skew.	Kurt.
FX	Japan	0.0083	0.0120	0.0002	0.1756	6.1819	59.5118
	Europe	0.0128	0.0127	0.0008	0.1837	4.1659	36.5686
	U.S.	0.0202	0.0226	0.0008	0.2894	4.0994	29.8966
Equity	Japan	0.0160	0.0466	0.0004	0.8970	10.5115	152.8709
	Europe	0.0306	0.0824	0.0005	2.2443	14.0351	307.4340
	U.S.	0.1263	0.3269	0.0015	7.1036	9.8323	148.8653
Bond	Japan	0.0032	0.0104	0.0001	0.1696	9.1152	107.1831
	Europe	0.0041	0.0068	0.0002	0.1354	8.9967	125.1152
	U.S.	0.0103	0.0121	0.0004	0.1801	4.7502	43.8053

A potentially interesting avenue of research is one which allows for the transmission of volatility to be different depending on whether the volatility is due to good or bad news. To capture this asymmetry, realised volatility is decomposed into realised volatility related to positive and negative returns (Barndorff-Nielsen et al., 2008) as follows

$$RV_t = RS_t^- + RS_t^+, \tag{11}$$

in which

$$RS_t^- = \sum_{\Omega=\{r_{j,t}<0\}} \left(r_{j,t}\right)^2 1_{r_{j,t}\in\Omega} \tag{12}$$

is the downside realised semivariance and

$$RS_t^+ = \sum_{\mho=\{r_{j,t}\geq 0\}} \left(r_{j,t}\right)^2 1_{r_{j,t}\in\mho} \tag{13}$$

is the upside realised semivariance, in which 1_a is the indicator function taking the value 1 if the argument a is true, and $\Omega \cup \mho = \{1, 2,..., 1\backslash\Delta\}$. Note that zero returns are treated as an indication of good news which is different from the formulation of Barndorff-Nielsen et al. (2008).

Table 6
Ten largest changes for the foreign exchange, equity, and bond prices in Japan, Europe, and the United States. The dates of the events are in the bottom cells. The values of realised volatility on this day in the basis points (multiplied by 1000) are in the upper cell.

Order	Foreign exchange			Equity			Bond		
	Jp	Eu	U.S.	Jp	Eu	U.S.	Jp	Eu	U.S.
1	0.1757	0.1838	0.2894	0.8970	2.2444	7.1036	0.1696	0.1354	0.1802
	22/10/08	05/01/09	13/11/08	08/12/08	08/10/08	10/10/08	22/09/09	17/09/08	09/08/11
2	0.1547	0.1608	0.2297	0.8311	1.2596	4.4425	0.1615	0.0977	0.1499
	31/10/08	19/12/08	29/10/08	27/10/08	16/10/08	23/10/08	22/06/11	02/05/08	08/10/08
3	0.1365	0.1143	0.2162	0.6536	0.9237	3.7652	0.1274	0.0964	0.1383
	30/10/08	30/10/08	18/03/09	10/10/08	09/08/11	29/10/08	23/03/11	12/06/09	22/03/10
4	0.1346	0.1066	0.2055	0.5181	0.8268	3.2545	0.1153	0.0808	0.0969
	05/11/08	24/10/08	20/02/09	09/08/11	10/10/08	08/10/08	12/12/07	12/03/09	19/12/08
5	0.1073	0.1032	0.1946	0.4938	0.6663	3.0842	0.1146	0.0705	0.0944
	16/09/08	08/10/08	24/10/08	08/10/08	11/08/11	13/11/08	06/03/09	12/12/07	15/12/06
6	0.1044	0.0961	0.1754	0.3819	0.5932	2.9657	0.1144	0.0696	0.0806
	24/10/08	18/09/08	19/11/08	13/10/08	17/10/08	20/11/08	18/09/08	07/08/09	23/01/08
7	0.1042	0.0804	0.1652	0.3436	0.5730	2.6830	0.1123	0.0571	0.0799
	29/10/08	06/01/09	15/01/09	24/11/08	07/10/08	16/10/08	21/12/11	08/07/11	29/09/08
8	0.1008	0.0796	0.1604	0.3338	0.5115	2.4697	0.0808	0.0476	0.0785
	28/10/08	07/05/10	05/11/08	16/10/08	27/10/08	24/10/08	22/03/05	03/09/10	29/10/08
9	0.1000	0.0754	0.1453	0.3271	0.4791	2.2121	0.0791	0.0447	0.0771
	13/10/08	23/01/09	25/11/08	24/10/08	09/03/09	21/11/08	09/06/08	11/06/12	07/05/10
10	0.0978	0.0743	0.1445	0.3130	0.4774	2.1020	0.0772	0.0417	0.0739
	27/10/08	27/01/09	25/03/09	21/11/08	29/10/08	14/10/08	23/03/10	07/09/12	01/12/08

6. Volatility linkages: asymmetry and jumps

At the end of Section 4, it emerged that using a traditional GARCH framework to investigate the intra-day effects of both the smoothed conditional variance and the news yielded ambiguous results. Armed with estimates of realised volatility from all the regions, however, this questions may be addressed using the following econometric model

$$h_t = K + Ah_{t-1} + BRV_t + G\epsilon^2_{t-1}, \tag{14}$$

in which $RV_t = [RV_{\mathrm{jp},t}\, RV_{\mathrm{eu},t}\, RV_{\mathrm{us},t-1}]'$, parameter matrices K, A and G are diagonal and B now has form

$$B = \begin{bmatrix} 0 & 0 & \beta_{13} \\ \beta_{21} & 0 & 0 \\ 0 & \beta_{32} & 0 \end{bmatrix}.$$

It is important to note that (somewhat paradoxically) the intra-day volatility effect on Japan comes from realised volatility at the close of the previous trading day in the United States.

In this specification the vector of realised volatilities will contain observable elements of both smoothed conditional variances and news from the preceding region and therefore Eq. (14) provides a parsimonious way of incorporating both elements into a comprehensive explanation of volatility transmission. This model may be regarded as a variant of the multiplicative error model of Engle and Gallo (2006), modified slightly in order to deal with the calendar structure imposed by the global trading day. The relevant asymptotic theory for the model with realised variance in the dynamic equation for conditional variance is given in Shephard and Sheppard

Table 7
Coefficient estimates for Eq. (14) with t-statistics based on QML standard errors in parentheses. (* denotes significance at the 5% level).

		Japan	Europe	United States
FX market	$RV_{\mathrm{jp},t}$	–	0.0082* (5.89)	–
	$RV_{\mathrm{eu},t}$	–	–	0.0157* (7.25)
	$RV_{\mathrm{us},t-1}$	0.0038* (6.35)	–	–
	$\epsilon^2_{\mathrm{jp},t-1}$	0.0303* (3.07)	–	–
	$\epsilon^2_{\mathrm{eu},t-1}$	–	0.0156* (3.69)	–
	$\epsilon^2_{\mathrm{us},t-1}$	–	–	0.0148 (1.55)
	h^i_{t-1}	0.8938* (89.0)	0.9492* (190)	0.8976* (109)
Equity market	$RV_{\mathrm{jp},t}$	–	0.0934* (9.78)	–
	$RV_{\mathrm{eu},t}$	–	–	0.2228* (17.2)
	$RV_{\mathrm{us},t-1}$	0.0057* (7.86)	–	–
	$\epsilon^2_{\mathrm{jp},t-1}$	0.0823* (2.15)	–	–
	$\epsilon^2_{\mathrm{eu},t-1}$	–	0.0535* (2.21)	–
	$\epsilon^2_{\mathrm{us},t-1}$	–	–	0.0268 (1.80)
	h^i_{t-1}	0.5687* (16.2)	0.6482* (20.2)	0.4445* (13.8)
Bond market	$RV_{\mathrm{jp},t}$	–	0.0027* (2.07)	–
	$RV_{\mathrm{eu},t}$	–	–	0.0018 (1.33)
	$RV_{\mathrm{us},t-1}$	0.0070* (4.88)	–	–
	$\epsilon^2_{\mathrm{jp},t-1}$	0.2448* (5.01)	–	–
	$\epsilon^2_{\mathrm{eu},t-1}$	–	0.0782* (7.16)	–
	$\epsilon^2_{\mathrm{us},t-1}$	–	–	0.0382* (6.79)
	h^i_{t-1}	0.6213* (20.3)	0.9064* (100)	0.9521* (203)

(2010) and in a recent paper by Han and Kristensen (2014), who show that standard inference based on quasi-maximum likelihood standard errors is applicable.

The coefficient estimates for Eq. (14) estimated on data from the foreign exchange, equity and bond markets are reported in Table 7. The most striking result is that eight out of nine coefficients on the intra-day effects are significant and therefore there is little doubt that the meteor shower effect of news being transmitted from region to region is an important component of volatility. It is also true that all the own lagged conditional variance terms are significant indicating that the heat wave effect is also present, although this

Table 8
Coefficient estimates of Eq. (15) with t-statistics based on QML standard errors in parentheses. (* denotes significance at the 5% level).

		Japan	Europe	United States
FX market	$CC_{jp,t}$	–	0.0071 (0.39)	–
	$CC_{eu,t}$	–	–	0.0160* (5.97)
	$CC_{us,t-1}$	0.0039* (4.45)	–	–
	$J_{jp,t}$	–	0.0227 (0.13)	–
	$J_{eu,t}$	–	–	0.0044 (0.48)
	$J_{us,t-1}$	0.0028 (0.90)	–	–
	$\epsilon^2_{jp,t-1}$	0.0311* (3.01)	–	–
	$\epsilon^2_{eu,t-1}$	–	0.0155* (2.10)	–
	$\epsilon^2_{us,t-1}$	–	–	0.0135 (1.53)
	h^i_{t-1}	0.8935* (57.0)	0.9490* (145)	0.8995* (71.0)
Equity market	$CC_{jp,t}$	–	0.0931* (5.09)	–
	$CC_{eu,t}$	–	–	0.2269* (12.8)
	$CC_{us,t-1}$	0.0060* (6.26)	–	–
	$J_{jp,t}$	–	0.0997* (1.99)	–
	$J_{eu,t}$	–	–	0.1942* (1.96)
	$J_{us,t-1}$	0.0000 (0.00)	–	–
	$\epsilon^2_{jp,t-1}$	0.0826* (1.96)	–	–
	$\epsilon^2_{eu,t-1}$	–	0.0535* (2.12)	–
	$\epsilon^2_{us,t-1}$	–	–	0.0254 (1.52)
	h^i_{t-1}	0.5600* (12.3)	0.6484* (19.7)	0.4394* (9.34)
Bond market	$CC_{jp,t}$	–	0.0035* (1.97)	–
	$CC_{eu,t}$	–	–	0.0038* (2.48)
	$CC_{us,t-1}$	0.0074* (4.23)	–	–
	$J_{jp,t}$	–	0.0014 (0.13)	–
	$J_{eu,t}$	–	–	0.0000 (0.00)
	$J_{us,t-1}$	0.0016 (0.34)	–	–
	$\epsilon^2_{jp,t-1}$	0.2508* (5.13)	–	–
	$\epsilon^2_{eu,t-1}$	–	0.0776* (3.30)	–
	$\epsilon^2_{us,t-1}$	–	–	0.0356* (6.12)
	h^i_{t-1}	0.6157* (16.0)	0.9059* (360)	0.9486* (126)

effect is fairly muted in the equity market. Indeed, the most remarkable results appear to be those obtained in the equity market where it is apparent that the intra-day effect of realised volatility from Japan on Europe and Europe on the United States is significant. These intra-day effects on the volatility in the United States appear much larger in absolute size than the effects obtained from the United States and Japanese realised variances.

Another really interesting result is the fact that now the heat wave hypothesis seems more appropriate in the bond market in comparison with the results of Table 3. While the intra-day effects from Japan to Europe and the United States to Japan are significant again, the values of all intra-day coefficients are closer to zero than coefficients on $h_{jp,t}$, $h_{eu,t}$, and $h_{us,t-1}$ from Table 3.

At this stage is seems appropriate to make use of Eqs. (9) and (10) to decompose the realised volatility series into its constituent diffusive and jump components in each of the three trading zones. The obvious extension to Eq. (14) is therefore

$$h_t = K + Ah_{t-1} + BCC_t + \widetilde{B}J_t + G\epsilon^2_{t-1}, \tag{15}$$

in which $CC_t = [CC_{jp,t}\, CC_{eu,t}\, CC_{us,t-1}]'$ represents the continuous components of realised volatility and $J_t = [J_{jp,t}\, J_{eu,t}\, J_{us,t-1}]'$ represents the jump component. The matrices B and $\widetilde{B}$ have the same structure as matrix B in Eq. (14).

The coefficient estimates for Eq. (15) are reported in Table 8. The major result is fairly unambiguous. Jumps are only really important in the transmission of volatility in the equity market and play no significant role in the intra-day transmission of volatility across regions in the foreign exchange and bond markets. On the other hand, the continuous component of realised volatility exhibits strong explanatory power across all three markets in all trading zones.

Given the importance of the intra-day effects of realised volatility in explaining the conditional variance and also the relative importance of the diffusive component of volatility relative to the jump component, the estimates of the realised semivariances can be used to build an asymmetric volatility model for each of the three markets. The relevant equation for the conditional variance is

$$h_t = K + Ah_{t-1} + BRS^+_t + \widetilde{B}RS^-_t + G\epsilon^2_{t-1}, \tag{16}$$

where the vector RV_t as defined after Eq. (14) is decomposed into its positive and negative semi-variances.[5]

The coefficient estimates from Eq. (16) are presented in Table 9. The results reported in Table 9 allow several general observations to be drawn about the asymmetric transmission of volatility relating to good and bad news.

1. There is strong evidence to support the hypothesis that volatility responds asymmetrically to the realised volatility from preceding zones. Remembering that the overall intra-day effects are seen in Table 7 to be significant in all cases, there is only one instance in which there are significant coefficients of comparable size recorded in Table 9. This is the case of the European influence on the United States in the foreign exchange market.
2. Volatility in the bond market responds mainly to the intra-day impact of the negative semivariances, as all coefficients on RS^+ are insignificant.
3. The largest intra-day impact on the foreign exchange and equity markets has the European negative semivariance. It is notable that the coefficient on $RS^-_{eu,t}$ in the equity market is even bigger than the effect from the lagged smoothed volatility h^i_{t-1}, which is a strong support of the meteor shower. Such a strong influence can be related to the crisis facing the euro area.
4. Equity markets in Europe and the United States appear particularly susceptible to volatility arising from bad news in the immediately preceding zones (Japan and Europe, respectively). This provides limited support for a claim that bad news travels fast in equity markets. The fact that Japan seems immune to this effect is perhaps due to the long bear market in the Japanese equity market.

These results indicate fairly strongly the need for further research in this area.

7. Conclusion

The paper investigates volatility transmission in global financial markets. In so doing, the seminal analysis of Engle et al. (1990) is extended in a number of different directions using a high frequency data set drawn from continuously-traded futures contracts from the foreign exchange, equity and bond markets. Returns to the futures contracts from different trading zones form a hypothetical global trading day in which developments in Japan can influence Europe and the United States on the same calendar day. Similarly events in Europe can influence the United States. The influence of the United States on Japan occurs at the beginning of the next trading day. The potential volatility and news spillovers in the transmission of volatility across trading zones are referred to as intra-day effects.

On the evidence presented here, it is clear that the meteor shower effect in the transmission of volatility from one region to another on the same trading day is significant. This effect is at least as important as the heat wave effect, which is the normal effect estimated in traditional volatility models. Moreover, the result of previous research in the bond market which suggests that the intra-day effect of volatility from preceding zones on the same trading day is not important is clearly refuted when realised volatility from previous zones is used as an explanatory factor. There can therefore be no doubt that there exist significant volatility linkages between financial markets.

[5] This model may be considered as a special case of the volatility model with a persistent leverage effect introduced by Corsi and Reno (2012).

Table 9
Coefficient estimates of Eq. (16) with t-statistics based on QML standard errors in parentheses. (* denotes significance at the 5% level).

		Japan	Europe	United States
FX market	$RS^{+}_{jp,t}$	–	0.0103 (0.20)	–
	$RS^{+}_{eu,t}$	–	–	0.0109* (1.97)
	$RS^{+}_{us,t-1}$	0.0053* (2.39)	–	–
	$RS^{-}_{jp,t}$	–	0.0065 (0.07)	–
	$RS^{-}_{eu,t}$	–	–	0.0180* (3.46)
	$RS^{-}_{us,t-1}$	0.0026* (1.98)	–	–
	$\epsilon^{2}_{jp,t-1}$	0.0315* (3.10)	–	–
	$\epsilon^{2}_{eu,t-1}$	–	0.0155* (1.96)	–
	$\epsilon^{2}_{us,t-1}$	–	–	0.0148 (0.09)
	h^{i}_{t-1}	0.8903* (58.0)	0.9537* (119)	0.9043* (75.0)
Equity market	$RS^{+}_{jp,t}$	–	0.0130 (0.09)	–
	$RS^{+}_{eu,t}$	–	–	0.0821* (2.93)
	$RS^{+}_{us,t-1}$	0.0139* (4.19)	–	–
	$RS^{-}_{jp,t}$	–	0.1771* (1.99)	–
	$RS^{-}_{eu,t}$	–	–	0.4181* (10.2)
	$RS^{-}_{us,t-1}$	0.0000 (0.00)	–	–
	$\epsilon^{2}_{jp,t-1}$	0.1023* (2.29)	–	–
	$\epsilon^{2}_{eu,t-1}$	–	0.0616* (2.42)	–
	$\epsilon^{2}_{us,t-1}$	–	–	0.0293 (1.55)
	h^{i}_{t-1}	0.4795* (10.0)	0.6371* (18.0)	0.3896* (8.28)
Bond market	$RS^{+}_{jp,t}$	–	0.0000 (0.00)	–
	$RS^{+}_{eu,t}$	–	–	0.0038 (1.38)
	$RS^{+}_{us,t-1}$	0.0000 (0.00)	–	–
	$RS^{-}_{jp,t}$	–	0.0037* (2.50)	–
	$RS^{-}_{eu,t}$	–	–	0.0000 (0.00)
	$RS^{-}_{us,t-1}$	0.0135* (2.80)	–	–
	$\epsilon^{2}_{jp,t-1}$	0.2495* (5.83)	–	–
	$\epsilon^{2}_{eu,t-1}$	–	0.0788* (3.17)	–
	$\epsilon^{2}_{us,t-1}$	–	–	0.0383* (6.80)
	h^{i}_{t-1}	0.6155* (16.0)	0.9083* (341)	0.9516* (138)

Incorporating realised volatility measures into the GARCH framework in the context of the problem of volatility linkages is a promising avenue of research and has yielded some interesting insights. The first is that volatility transmission appears to be asymmetric, particularly in the equity market where volatility related to negative news appears to be transmitted more quickly that volatility linked to good news. Furthermore the decomposition of realised volatility into its continuous and jump components yields the unexpected result that jumps in volatility are not as readily transmitted as might be expected a priori.

References

Andersen, T.G., Bollerslev, T., Diebold, F.X., Labys, P., 2003. Modeling and forecasting realized volatility. Econometrica 71, 579–625.
Andersen, T.G., Dobrev, D., Schaumburg, E., 2012. Jump-robust volatility estimation using nearest neighbor truncation. J. Econ. 169, 75–93.
Barndorff-Nielsen, O.E., Shephard, N., 2002. Econometric analysis of realized volatility and its use in estimating stochastic volatility models. J. R. Stat. Soc. Ser. B 64, 253–280.
Barndorff-Nielsen, O.E., Shephard, N., 2004. Power and bi-power variation with stochastic volatility and jumps. J. Financ. Econ. 2, 1–37.
Barndorff-Nielsen, O.E., Shephard, N., 2006. Econometrics of testing for jumps in financial economics using bipower variation. J. Financ. Econ. 4, 1–30.
Barndorff-Nielsen, O.E., Kinnebrock, S., and Shephard, N. 2008. Measuring downside risk-realised semivariance. Unpublished Manuscript.
Bollerslev, T., 1986. Generalized autoregressive conditional heteroskedasticity. J. Econ. 31, 307–327.
Bubák, V., Kočenda, E., Žikeš, F., 2011. Volatility transmission in emerging European foreign exchange markets. J. Bank. Financ. 35, 2829–2841.
Corsi, F., Reno, R., 2012. Discrete-time volatility forecasting with persistent leverage effect and the link with continuous-time volatility modelling. J. Bus. Econ. Stat. 30, 368–380.
Dungey, M., Fakhrutdinova, L., Goodhart, C.A.E., 2009. After-hours trading in equity futures markets. J. Futur. Mark. 29, 114–136.
Ehrmann, M., Fratzscher, M., Rigobon, R., 2011. Stocks, bonds, money markets and exchange rates: measuring international financial transmission. J. Appl. Econ. 26, 948–974.
Engle, R.F., 1982. Autoregressive conditional heteroskedasticity with estimates of the variance of United Kingdom inflation. Econometrica 50, 987–1008.
Engle, R.F., Gallo, G.M., 2006. A multiple indicators model for volatility using intra-daily data. J. Econ. 131, 3–27.
Engle, R.F., Ito, T., Lin, W.L., 1990. Meteor showers or heat waves? Heteroskedastic intra-daily volatility in the foreign exchange market. Econometrica 58, 525–542.
Engle, R.F., Gallo, G.M., Velucchi, M., 2012. Volatility spillovers in East Asian financial markets: a MEM-based approach. Rev. Econ. Stat. 94, 222–223.
Evans, K.P., 2011. Intraday jumps and US macroeconomic news announcements. J. Bank. Financ. 35, 2511–2527.
Fleming, M.J., Lopez, J.A., 1999. Heat waves, meteor showers, and trading volume: an analysis of volatility spillovers in the U.S. Treasury market. Working Paper. Federal Reserve Bank of New York, p. 82.
Hakim, A., McAleer, M., 2010. Modelling the interactions across international stock, bond and foreign exchange markets. Appl. Econ. 42, 825–850.
Han, H., Kristensen, D., 2014. Asymptotic theory for the QMLE in GARCH-X models with stationary and non-stationary covariates. J. Bus. Econ. Stat. 32, 416–429.
Hansen, P.R., Huang, Z., Shek, H.H., 2012. Realized GARCH: a joint model for returns and realized measures of volatility. J. Appl. Econ. 27, 877–906.
Huang, X., Tauchen, G., 2005. The relative contribution of jumps to total price variation. J. Financ. Econ. 3, 456–499.
Ito, T., 1987. The intra-daily exchange rate dynamics and monetary policy after the G5 agreement. J. Jpn. Int. Econ. 1, 275–298.
Ito, T., Roley, V.V., 1987. News from the U.S. and Japan which moves the Yen/Dollar exchange rate? J. Monet. Econ. 19, 255–278.
Kawakatsu, H., 2006. Matrix exponential GARCH. J. Econ. 134, 95–128.
Kroner, K.F., Ng, V.K., 1998. Modeling asymmetric comovements of asset returns. Rev. Financ. Stud. 11, 817–844.
Melvin, M., Melvin, B.P., 2003. The global transmission of volatility in the foreign exchange market. Rev. Econ. Stat. 85, 670–679.
Patton, A.J., Liu, L., Sheppard, K., 2013. Does anything beat 5-minute RV? A comparison of realized measures across multiple asset classes. Working Paper. University of Oxford, Department of Economics.
Savva, C., Osborn, D.R., and Gill, L. 2005. Volatility, spillover effects and correlations in US and major european markets. Unpublished Manuscript, University of Manchester.
Shephard, N., Sheppard, K., 2010. Realising the future: forecasting with high-frequency-based volatility (HEAVY) models. J. Appl. Econ. 25, 197–231.