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Threshold bipower variation and the impact of jumps on volatility forecasting[☆]

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ABSTRACT

This study reconsiders the role of jumps for volatility forecasting by showing that jumps have a positive and mostly significant impact on future volatility. This result becomes apparent once volatility is separated into its continuous and discontinuous components using estimators which are not only consistent, but also scarcely plagued by small sample bias. With the aim of achieving this, we introduce the concept of threshold bipower variation, which is based on the joint use of bipower variation and threshold estimation. We show that its generalization (threshold multipower variation) admits a feasible central limit theorem in the presence of jumps and provides less biased estimates, with respect to the standard multipower variation, of the continuous quadratic variation in finite samples. We further provide a new test for jump detection which has substantially more power than tests based on multipower variation. Empirical analysis (on the S&P500 index, individual stocks and US bond yields) shows that the proposed techniques improve significantly the accuracy of volatility forecasts especially in periods following the occurrence of a jump.

1. Introduction

The importance of jumps in financial economics is widely recognized. A partial list of recent studies on this topic includes the test specification of Aït-Sahalia (2004), Jiang and Oomen (2008), Barndorff-Nielsen and Shephard (2006), Lee and Mykland (2008) and Aït-Sahalia and Jacod (2009), as well as the empirical studies of Bollerslev et al. (2008), Maheu and McCurdy (2004), Bollerslev et al. (2009), Andersen et al. (2007) and Cartea and Karyampas (2010); non-parametric estimation in the presence of jumps, as in Bandi and Nguyen (2003), Johannes (2004) and Mancini and Renò (forthcoming); option pricing as in Duffie et al. (2000), Eraker et al. (2003) and Eraker (2004); applications to bond and stock markets, as in Das and Uppal (2004) and Wright and Zhou (2007). Interesting references for a review are Cont and Tankov (2004) and Barndorff-Nielsen and Shephard (2007). However, while jumps have been shown to be relevant in economic and financial applications, previous research has not found evidence that jumps help in forecasting volatility.

In this study we start from an apparent puzzle contained in the studies of Andersen et al. (2007) (henceforth ABD), Forsberg and Ghysels (2007), Giot and Laurent (2007) and Busch et al. (forthcoming): these works find a negative (or null) impact of jumps on future volatility. We find this result puzzling in at least two respects. First, visual inspection of realized variance time series reveals that bursts in volatility are usually initiated by a large and unexpected movement of asset prices, suggesting that jumps *should* have a forecasting power for volatility. Second, it is well known that volatility is associated with dispersion of beliefs and heterogeneous information; see e.g. Shalen (1993), Wang (1994) and Buraschi et al. (2007). If the occurrence of a jump increases the uncertainty on fundamental values, it is likely to have a positive impact on future volatility. Why, then, have jumps been found to be irrelevant? Our first contribution is to show that this puzzle is spuriously due to the small sample bias of bipower variation, a very

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popular measure of continuous quadratic variation introduced by Barndorff-Nielsen and Shephard (2004). We show by means of realistic simulations that, even if bipower variation is consistent, in finite samples it is greatly upper biased in the presence of jumps, and this implies a large underestimation of the jump component. Unfortunately, this problem cannot be accommodated by simply shrinking the observation interval, since market microstructure effects would jeopardize the estimation in an unpredictable way.[1]

The second contribution of this paper is then to introduce an alternative estimator of integrated powers of volatility in the presence of jumps which is less affected by small sample bias. We introduce (Section 2) the concept of *threshold multipower variation*, which can be viewed as a combination of multipower variation and the threshold realized variance of Mancini (2009). We show consistency and asymptotic normality in the presence of jumps of the newly proposed estimator (a central limit theorem for bipower variation has been recently proved by Vetter, 2010). Moreover, using realistic simulation of asset prices (Section 4), we show that threshold bipower variation is nearly unbiased on continuous trajectories and, importantly, also in the presence of jumps. Finally, it is robust to the choice of the threshold function, in the sense that the impact of the threshold choice on estimation is marginal. Thus, it is an ideal candidate for estimating models of volatility dynamics in which we use separately the continuous and discontinuous volatility as explanatory variables.

Our third contribution is the introduction of a novel test for jump detection in time series (Section 3). Our C-Tz test is basically a correction of the z statistics of Barndorff-Nielsen and Shephard (2006) which softens the finite sample bias in estimating the integral of the second and fourth powers of continuous volatility in the presence of jumps. We show, by using simulated data (Section 4), that the two tests are equally sized under the null, but that in the presence of jumps the C-Tz test has substantially more power than the z test, especially when jumps are consecutive, a situation which is quite frequent for high frequency data.

Empirical results (Section 5) constitute our fourth contribution. We work with stock index futures, individual stocks and Treasury Bond data. We find that lagged jumps have a positive and highly significant impact on realized variance, a result which cannot be observed when bipower variation is employed because of its inherent bias. We also document, uniformly on our data sets, that it is possible to get higher R^2 (especially in days following a jump) and generally significant lower relative root mean square error, just by using measures based on threshold multipower variation instead of measures based on multipower variation, with no other changes in the model used for volatility forecasting. Concluding remarks are in Section 6.

2. Threshold multipower variation

2.1. Introductory concepts

We work in a filtered probability space $(\Omega, (\mathcal{F}_t)_{t\in[0,T]}, \mathcal{F}, \mathcal{P})$, satisfying the usual conditions (Protter, 2004). We assume that an economic variable X_t, for example the logarithmic price of a stock or an interest rate, satisfies the following assumption:

Assumption 2.1. $(X_t)_{t\in[0,T]}$ is a real-valued process such that X_0 is measurable with respect to $\mathcal{F}_0$ and

$$\mathrm{d}X_t = \mu_t \mathrm{d}t + \sigma_t \mathrm{d}W_t + \mathrm{d}J_t \tag{2.1}$$

where μ_t is predictable, σ_t is cádlág; $\mathrm{d}J_t = c_t \mathrm{d}N_t$ where N_t is a non-explosive Poisson process whose intensity is an adapted stochastic process λ_t, the times of the corresponding jumps are $(\tau_j)_{j=1,\dots,N_T}$ and c_j are i.i.d. adapted random variables measuring the size of the jump at time τ_j and satisfying, $\forall t \in [0, T]$, $\mathcal{P}(\{c_j = 0\}) = 0$.

Typically, a time window T is fixed, e.g. one day, and we define the quantities of interest on a time span of length T. Quadratic variation of such a process over a time window is defined as

$$[X]_t^{t+T} := X_{t+T}^2 - X_t^2 - 2\int_t^{t+T} X_{s-}\mathrm{d}X_s, \tag{2.2}$$

where t indexes the day. It can be decomposed into its continuous and discontinuous component as

$$[X]_t^{t+T} = [X^c]_t^{t+T} + [X^d]_t^{t+T} \tag{2.3}$$

where $[X^c]_t^{t+T} = \int_t^{t+T}\sigma_s^2 \mathrm{d}s$ and $[X^d]_t^{t+T} = \sum_{j=N_t}^{N_{t+T}} c_j^2$. To estimate these quantities, we divide the time interval $[t, t+T]$ into n subintervals of length $\delta = T/n$. On this grid, we define the evenly sampled returns as

$$\Delta_{j,t}X = X_{j\delta+t} - X_{(j-1)\delta+t}, \quad j = 1, \dots, n. \tag{2.4}$$

For simplicity, in what follows we omit the subscript t and we simply write $\Delta_j X$. A popular estimator of $[X]_t^{t+T}$ is the realized variance, defined as

$$\mathsf{RV}_\delta(X)_t = \sum_{j=1}^{n}\left(\Delta_j X\right)^2, \tag{2.5}$$

which converges in probability to $[X]_t^{t+T}$ as $\delta \to 0$. To disentangle the continuous quadratic variation from the discontinuous one, multipower variation has been introduced by Barndorff-Nielsen and Shephard (2004). It is defined as

$$\mathsf{MPV}_\delta(X)_t^{[\gamma_1,\dots,\gamma_M]} = \delta^{1-\frac{1}{2}(\gamma_1+\dots+\gamma_M)}\sum_{j=M}^{[T/\delta]}\prod_{k=1}^{M}|\Delta_{j-k+1}X|^{\gamma_k}, \tag{2.6}$$

and it converges in probability, as $\delta \to 0$, to $\int_t^{t+T}\sigma_s^{\gamma_1+\dots+\gamma_M}\mathrm{d}s$ times a suitable constant. Asymptotic properties of multipower variation have been studied by Barndorff-Nielsen et al. (2006b) in the absence of jumps, and by Barndorff-Nielsen et al. (2006a) and Woerner (2006) in the presence of jumps. For practical applications, multipower variation is used for the estimation of $\int_t^{t+T}\sigma_s^2\mathrm{d}s$ and $\int_t^{t+T}\sigma_s^4\mathrm{d}s$. Bipower variation is defined as

$$\mathsf{BPV}_\delta(X)_t = \mu_1^{-2}\mathsf{MPV}_\delta(X)_t^{[1,1]} = \mu_1^{-2}\sum_{j=2}^{[T/\delta]}|\Delta_{j-1}X|\cdot|\Delta_j X|, \tag{2.7}$$

with $\mu_1 \simeq 0.7979$, and it converges in probability, as $\delta \to 0$, to $\int_t^{t+T}\sigma_s^2\mathrm{d}s$. For estimators of $\int_t^{t+T}\sigma_s^4\mathrm{d}s$ based on multipower variation, see Appendix A.

On the basis of a threshold function $\Theta(\delta)$, Mancini (2009) provides alternative estimators of squared and fourth-power integrated volatilities. The threshold realized variance is defined as follows:

$$\mathsf{TRV}_\delta(X)_t = \sum_{j=1}^{[T/\delta]}|\Delta_j X|^2 I_{\left\{|\Delta_j X|^2 \le \Theta(\delta)\right\}} \tag{2.8}$$

[1] Attempts to study and correct bipower variation under microstructure noise can be found in Podolskij and Vetter (2009), and Christensen et al. (forthcoming). ABD and Huang and Tauchen (2005) propose a staggered version of bipower variation. Fan and Wang (2007) study jumps and microstructure noise with wavelet methods. Jacod et al. (2009) pre-average the data to get rid of microstructure noise. The impact of microstructure noise on threshold estimation of Mancini (2009) is however unknown. Directly incorporating microstructure noise can improve volatility forecasting; see e.g. Aït-Sahalia and Mancini (2008).

where $I_{\{\cdot\}}$ is the indicator function and the threshold function has to satisfy

$$\lim_{\delta\to 0}\Theta(\delta)=0,\qquad \lim_{\delta\to 0}\frac{\delta\log\frac{1}{\delta}}{\Theta(\delta)}=0, \tag{2.9}$$

that is it has to vanish more slowly than the modulus of continuity of the Brownian motion in order to have convergence in probability of $\mathsf{TRV}_\delta(X)_t$, as $\delta \to 0$, to $\int_t^{t+T}\sigma_s^2 ds$. Mancini (2009) also establishes a central limit theorem for TRV and provides a similar quarticity estimator, which is defined in Eq. (A.3) in Appendix A.

2.2. The definition and asymptotic theory

We now introduce our family of estimators. In what follows, we use a strictly positive random threshold function $\vartheta_s : [t, t+T] \to \mathbb{R}^+$. For brevity, we write $\vartheta_j := \vartheta_{j\delta+t}$.

Definition 2.2. Let $\gamma_1, \ldots, \gamma_M > 0$. We define the realized threshold multipower variation as

$$\mathsf{TMPV}_\delta(X)_t^{[\gamma_1,\ldots,\gamma_M]} = \delta^{1-\frac{1}{2}(\gamma_1+\cdots+\gamma_M)}\sum_{j=M}^{[T/\delta]}\prod_{k=1}^{M}|\Delta_{j-k+1}X|^{\gamma_k} I_{\left\{|\Delta_{j-k+1}X|^2\le\vartheta_{j-k+1}\right\}}. \tag{2.10}$$

The intuition behind the concept of threshold multipower variation is the following. Suppose $|\Delta_j X|$ contains a jump. In the case of bipower variation, it multiplies $|\Delta_{j-1}X|$ and $|\Delta_{j+1}X|$. Asymptotically, both these returns have to vanish and bipower variation converges to the integrated continuous volatility. However, for finite δ these returns will not vanish, causing a positive bias which will become larger as $|\Delta_j X|$ increases. This consideration suggests that the bias of multipower variation will be extremely large in the case of consecutive jumps. For estimators (2.10) instead, if $|\Delta_j X|$ contains a jump larger than ϑ_j, the corresponding indicator function vanishes, thus correcting for the bias. This intuition is supported by the analysis in the subsequent sections.

Importantly, threshold multipower variation admits a central limit theorem in the presence of jumps, which is stated as follows.

Theorem 2.3. *Let Assumption 2.1 hold, and let $\vartheta_t = \xi_t\Theta(\delta)$, where $\Theta(\delta)$ is a real function satisfying conditions (2.9) and ξ_t is a stochastic process on $[0, T]$ which is a.s. bounded and with a strictly positive lower bound. Then, as $\delta \to 0$,*

$$\mathsf{TMPV}_\delta(X)_t^{[\gamma_1,\ldots,\gamma_M]} \xrightarrow{p} \left(\prod_{k=1}^{M}\mu_{\gamma_k}\right)\int_t^{t+T}\sigma_s^{\gamma_1+\cdots+\gamma_M}ds \tag{2.11}$$

where the above convergence is in probability, and

$$\delta^{-\frac{1}{2}}\left(\mathsf{TMPV}_\delta(X)_t - \left(\prod_{k=1}^{M}\mu_{\gamma_k}\right)\int_t^{t+T}\sigma_s^{\gamma_1+\cdots+\gamma_M}ds\right)$$
$$\longrightarrow c_\gamma\int_t^{t+T}\sigma_s^{\gamma_1+\cdots+\gamma_M}dW_s' \tag{2.12}$$

where W' is a Brownian motion independent of W, the above convergence holds stably in law and

$$c_\gamma^2 = \prod_{k=1}^{M}\mu_{2\gamma_k} - 2(M-1)\prod_{k=1}^{M}\mu_{\gamma_k}^2 + 2\sum_{j=1}^{M-1}\left(\prod_{k=1}^{j}\mu_{\gamma_k}\prod_{k=M-j+1}^{M}\mu_{\gamma_k}\prod_{k=1}^{M-j}\mu_{\gamma_k+\gamma_{k+j}}\right). \tag{2.13}$$

Proof. Write

$X = Y + Z$

where $Y_t = \int_t^{t+T}\mu_s ds + \int_t^{t+T}\sigma_s dW_s$. If $Z = 0$, the theorem has been proved by Barndorff-Nielsen et al. (2006c) for multipower variation. Since Z is a finite activity jump process and Y is a Brownian semimartingale, by virtue of Theorem 1 in Mancini (2009) and Remark 3.4 in Mancini and Renò (forthcoming) we can write, for δ small enough,

$$\mathsf{TMPV}_\delta(X)_t^{[\gamma_1,\ldots,\gamma_M]} = \delta^{1-\frac{1}{2}(\gamma_1+\cdots+\gamma_M)}\left(\sum_{j=M}^{[T/\delta]}\prod_{k=1}^{M}|\Delta_{j-k+1}Y|^{\gamma_k} - \sum_{j=M}^{[T/\delta]}I_{j,M}^*\prod_{k=1}^{M}|\Delta_{j-k+1}Y|^{\gamma_k}\right)$$

where $I_{j,M}^*$ is 1 if there has been a single jump between $t_{j\delta}$ and $t_{j+M\delta}$, and zero otherwise. Thus, using the modulus of continuity of the Brownian motion,

$$\begin{aligned}\mathsf{TMPV}_\delta(X)_t^{[\gamma_1,\ldots,\gamma_M]} &- \mathsf{MPV}_\delta(Y)_t^{[\gamma_1,\ldots,\gamma_M]}\\ &= \delta^{1-\frac{1}{2}(\gamma_1+\cdots+\gamma_M)}O_p\left(N_T\left(\delta\log|\delta|\right)^{\frac{1}{2}(\gamma_1+\cdots+\gamma_M)}\right)\\ &= O_p\left(\delta\left(\log|\delta|\right)^{\frac{1}{2}(\gamma_1+\cdots+\gamma_M)}\right)\end{aligned}$$

which is $o_p(\delta^{\frac{1}{2}})$; thus $\mathsf{TMPV}(X)$ has the same limit as $\mathsf{MPV}(Y)$ both in probability and in distribution. □

In Theorem 2.3 we could also allow for infinite activity jumps under suitable conditions on the coefficients $\gamma_1, \ldots, \gamma_M$; see e.g. Jacod (2008) and Mancini (2009).

A simple special case of TMPV is *threshold bipower variation*, obtained with $\gamma_1 = \gamma_2 = 1$ and multiplying by a suitable constant:

$$\begin{aligned}\mathsf{TBPV}_\delta(X)_t &= \mu_1^{-2}\mathsf{TMPV}_\delta(X)_t^{[1,1]}\\ &= \mu_1^{-2}\sum_{j=2}^{[T/\delta]}|\Delta_{j-1}X|\cdot|\Delta_j X| I_{\left\{|\Delta_{j-1}X|^2\le\vartheta_{j-1}\right\}} I_{\left\{|\Delta_j X|^2\le\vartheta_j\right\}}.\end{aligned} \tag{2.14}$$

For estimators of the integrated fourth power of volatility using threshold multipower variation, see Appendix A. Alternative estimators of the integrated variance, developed with the aim of reducing the bias induced by jumps, have been recently proposed by Andersen et al. (2008), using the minimum of $|\Delta_{j-1}X|\cdot|\Delta_j X|$ or the median value of $|\Delta_{j-1}X|\cdot|\Delta_j X|\cdot|\Delta_{j+1}X|$. Their estimators also admit a central limit theorem in the presence of jumps, and have the advantage of not needing the specification of the threshold function, but the disadvantage of a larger asymptotic variance with respect to TBPV. Related work can be found in Christensen et al. (forthcoming) and Boudt et al. (2008).

For applications, it is important to note that our central limit theorem holds even with a stochastic threshold ϑ_t fulfilling the assumptions of the theorem. This can be important since it is natural to scale the threshold function with respect to the local spot variance, which is estimated with the data themselves. In this paper, we write

$$\vartheta_t = c_\vartheta^2\cdot\widehat{V}_t, \tag{2.15}$$

where $\widehat{V}_t$ is an auxiliary estimator of σ_t^2 and c_ϑ is a scale-free constant. The dimensionless parameter c_ϑ can be used to change the threshold, and by varying it we can test the robustness of proposed estimators with respect to the choice of the threshold. As we show below by means of simulations, the choice of the auxiliary estimator $\widehat{V}_t$, among unbiased estimators of the local variance, is immaterial for our purposes. The estimator of the local variance used in this paper with simulated and actual data is described in Appendix B.

In addition to providing a feasible asymptotic theory, the newly proposed threshold multipower variation is expected to perform better for finite samples. Indeed, bipower variation is greatly biased by big jumps but is less affected by small jumps. On the other hand, threshold realized variance is problematic with small jumps while much more effective with large ones. Thus, the joint combination of bipower variation and threshold is doubly effective in disentangling diffusion from jumps since each method tends to compensate for the weakness of the other. We provide evidence of the benefit of this "double-sword" technique below, using simulated data.

3. The jump detection test

The test that we propose for jump detection is a modification of the test proposed by Barndorff-Nielsen and Shephard (2006) based on the difference between RV and BPV, which is expected to be small in the absence of jumps (the null) and large in the presence of the jumps (the alternative): we modify it by replacing estimators based on multipower variation with estimators based on threshold multipower variation. However, when using the latter for finite δ, when $|\Delta_j X|^2 > \vartheta_j$ the corresponding return is annihilated. This is a potential issue when testing for the presence of jumps, since variations larger than the threshold exist even in the absence of jumps and annihilating them is a source of negative bias for threshold multipower variation. This issue can be attenuated if, when $|\Delta_j X|^2 > \vartheta_j$, we replace $|\Delta_j X|^\gamma$ with its expected value under the assumption that $\Delta_j X \sim \mathcal{N}(0, \sigma^2)$, which is given by

$$\mathbb{E}\left[|\Delta_j X|^\gamma \,|(\Delta_j X)^2 > \vartheta\right] = \frac{\Gamma\left(\frac{\gamma+1}{2}, \frac{\vartheta}{2\sigma^2}\right)}{2N\left(-\frac{\sqrt{\vartheta}}{\sigma}\right)\sqrt{\pi}} \left(2\sigma^2\right)^{\frac{1}{2}\gamma}, \tag{3.1}$$

where $N(x)$ is the standard normal cumulative function and $\Gamma(\alpha, x)$ is the upper incomplete gamma function.[2] Then, writing $\vartheta = c_\vartheta^2 \sigma^2$, we define the corrected realized threshold multipower estimator as

$$\text{C-TMPV}_\delta(X)_t^{[\gamma_1,\ldots,\gamma_M]} = \delta^{1-\frac{1}{2}(\gamma_1+\cdots+\gamma_M)} \sum_{j=M}^{[T/\delta]} \prod_{k=1}^{M} Z_{\gamma_k}(\Delta_{j-k+1}X, \vartheta_{j-k+1}) \tag{3.2}$$

where the function $Z_\gamma(x, y)$ is defined as

$$Z_\gamma(x, y) = \begin{cases} |x|^\gamma & \text{if } x^2 \le y \\ \dfrac{1}{2N(-c_\vartheta)\sqrt{\pi}} \left(\dfrac{2}{c_\vartheta^2} y\right)^{\frac{\gamma}{2}} \Gamma\left(\dfrac{\gamma+1}{2}, \dfrac{c_\vartheta^2}{2}\right) & \text{if } x^2 > y. \end{cases} \tag{3.3}$$

Relevant cases which will be examined in what follows are $\gamma = 1, 2, 4/3$. In these special cases we have, with $c_\vartheta = 3$ and $x^2 > y$, $Z_1(x, y) \simeq 1.094 \cdot y^{\frac{1}{2}}$, $Z_{4/3}(x, y) \simeq 1.129 \cdot y^{\frac{2}{3}}$, and $Z_2(x, y) \simeq 1.207 \cdot y$. For example, the corrected version of (2.14) is the *corrected threshold bipower variation* defined as

$$\begin{aligned} \text{C-TBPV}_\delta(X)_t &= \mu_1^{-2} \text{C-TMPV}_\delta(X)_t^{[1,1]} \\ &= \mu_1^{-2} \sum_{j=2}^{[T/\delta]} Z_1(\Delta X_j, \vartheta_j) Z_1(\Delta X_{j-1}, \vartheta_{j-1}). \end{aligned} \tag{3.4}$$

The test statistic that we propose is based on this correction and it is defined by

$$\text{C-Tz} = \delta^{-\frac{1}{2}} \frac{(\text{RV}_\delta(X)_T - \text{C-TBPV}_\delta(X)_T) \cdot \text{RV}_\delta(X)_T^{-1}}{\sqrt{\left(\frac{\pi^2}{4} + \pi - 5\right) \max\left\{1, \frac{\text{C-TTriPV}_\delta(X)_T}{(\text{C-TBPV}_\delta(X)_T)^2}\right\}}}, \tag{3.5}$$

where $\text{C-TTriPV}_\delta(X)_T$ is a quarticity estimator which is a subcase of threshold multipower variation and is precisely defined in Eq. (A.5) in Appendix A. We immediately prove the following:

Corollary 3.1. *Under the assumptions of Theorem 2.3 we have that if* $\mathrm{d}J_t = 0$ *then* C-Tz $\to \mathcal{N}(0, 1)$ *stably in law as* $\delta \to 0$.

Proof. The correction, for small δ, affects only the jumps which are finite in number. Then we use the delta method as in Barndorff-Nielsen and Shephard (2004) and Theorem 2.3, which implies that the denominator converges in probability to the integrated quarticity, thus yielding the desired result. □

Simulations in the next section show that the C-Tz test is correctly sized in finite samples under the null.[3]

4. The simulation study

To assess the small sample properties of concurrent estimators we use Monte Carlo simulations of realistic stochastic processes which have been extensively used to model stock index prices. The purpose of this section is to show that, in finite samples, bipower variation is a biased estimator of integrated volatility in the presence of jumps, while threshold-based estimators are much less sensitive to jumps and accordingly less biased. Moreover, we show that while the threshold realized variance (2.8) is particularly sensitive to the choice of the threshold, the threshold bipower variation (2.14) is instead highly robust against this choice. This latter feature is particularly important, since it suggests that the results obtained in our empirical applications using threshold multipower variation do not depend crucially on the threshold employed.

The model that we simulate is a one-factor jump–diffusion model with stochastic volatility, described by a couple of stochastic differential equations:

$$\begin{aligned} \mathrm{d}X_t &= \mu\, \mathrm{d}t + \sqrt{v_t}\mathrm{d}W_{x,t} + c_t \mathrm{d}N_t, \\ \mathrm{d}\log v_t &= (\alpha - \beta \log v_t)\, \mathrm{d}t + \eta \mathrm{d}W_{v,t}, \end{aligned} \tag{4.1}$$

where W_x and W_v are standard Brownian motions with corr$(\mathrm{d}W_x, \mathrm{d}W_v) = \rho$, v_t is a stochastic volatility factor, $c_t \mathrm{d}N_t$ is a compound Poisson process with random jump size which is normally distributed with zero mean and standard deviation σ_J. We use the model parameters estimated by Andersen et al. (2002) for *S&P*500 prices: $\mu = 0.0304$, $\alpha = -0.012$, $\beta = 0.0145$, $\eta = 0.1153$, $\rho = -0.6127$, $\sigma_J = 1.51$, where the parameters are expressed in daily units and returns are in percentages. Similar

[2] See Abramowitz and Stegun (1965). The functions $N(\cdot)$ and $\Gamma(\cdot, \cdot)$ are defined by

$$N(x) = \int_{-\infty}^{x} \frac{1}{\sqrt{2\pi}} \mathrm{e}^{-\frac{1}{2}s^2} \mathrm{d}s, \qquad \Gamma(\alpha, x) = \int_{x}^{+\infty} s^{\alpha-1} \mathrm{e}^{-s} \mathrm{d}s.$$

When $\alpha = 1$, $\Gamma(1, x) = \mathrm{e}^{-x}$. For large γ, it is useful to use the integration by parts formula $\Gamma(\alpha + 1, x) = \alpha\Gamma(\alpha, x) + x^\alpha\, \mathrm{e}^{-x}$.

[3] Under the alternative, the corrected version of threshold multipower variation is subject to a source of upper bias. Indeed, suppose that there is a jump J such that $\Delta_j X = \Delta_j X^c + J$. If $(\Delta_j X)^2 > \vartheta_j$, we replace $|\Delta_j X|^\gamma$ with $\mathbb{E}\left[|\Delta_j X^c|^\gamma |(\Delta_j X^c)^2 > \vartheta\right]$, which is much larger than $\mathbb{E}\,[|X^c|^\gamma]$, which is our estimation target.

estimates have been obtained by Bates (2000), Pan (2002) and Chernov et al. (2003). The numerical integration of the system (4.1) is performed with the Euler scheme, using a discretization step of $\Delta = 1$ second. Each day, we simulate $7 \cdot 60 \cdot 60$ steps corresponding to seven hours. We then use $\delta = 5$ minutes, that is 84 returns per day. The threshold is set as in Eq. (2.15), with the local variance estimator $\widehat{V}_t$ defined in Appendix B and $c_\vartheta = 3$.

We generate samples of 1000 "daily" replications in the following way. In the first sample (no jumps), we do not generate jumps at all. In the second sample (one jump), we generate a single jump for each day which is located randomly within the day. In the third sample (two jumps), we generate exactly two jumps per day which are located randomly within the day. In the fourth sample (two consecutive jumps), we generate two jumps per day and we force them to be consecutive (i.e., the first is located randomly and the second jump is forced to occur 300 s after the first). This procedure allows us to compute the expected value conditional on the presence of zero, one, two jumps and two consecutive jumps. For every simulated daily trajectory, we compute the estimates of BPV (and its staggered version; see ABD) and their fourth-power counterparts QPV, TriPV as well as threshold estimators TRV, TQV, TTriPV and threshold multipower estimators TBPV, TQPV, TTriPV (precise definitions of these estimators are given in Appendix A). We compute daily percentage estimation error and compute averages and standard deviations across the sample.

The results reported in Table 1 are compelling. Bipower variation is substantially biased if there is one jump in the trajectory (+48.04%) and greatly biased (+102.03%) if there are two jumps in the trajectory. If the two jumps are consecutive, the bias is huge (+595.57%) and can only be marginally softened by using staggered bipower variation (+97.07%, like for the case of two jumps). The bias of multipower variation in estimating integrated quarticity is even larger.

Threshold-based estimators, however, are much more robust to the presence of jumps. The bias of the threshold realized variance in estimating integrated squared volatility is around −5% in the absence of jumps and ranges from −6% to −7% in the presence of one and two jumps, consecutive or not. The same happens when estimating quarticity, the bias being around −16%. The presence of a negative bias is due to the fact that, by their proper definitions, threshold estimators remove completely observations larger than the threshold. When we correct for this as indicated in Section 3, the bias turns out to be positive since, when an observation is above the threshold, we replace it with its expected value under the assumption that the observation was actually above threshold; which is true under the null of no jumps, but need not be true under the alternative—see footnote 3.

The estimators based on threshold multipower variation, introduced in this study, yield equally good results. Threshold bipower variation has a bias of −4.15% in the case of no jumps, of −4.83% with a single jump, of −5.65% in the case of two jumps, and of −4.70% in the case of two consecutive jumps. When estimating quarticity, these biases range between −9% and −15% according to the number of jumps and the estimator used. Again, the corrected versions greatly correct the bias under the null of no jumps, but turn the negative bias into a positive one in the case of jumps. However, from our simulated experiment we can conclude that threshold-based estimators perform much better than multipower variation in the presence of jumps.

Between the two competing threshold estimators (threshold realized volatility and threshold bipower variation), our simulation experiments highlight a substantial advantage in using the second: as shown in Fig. 1 integrated volatility measures are much more robust with respect to the parameter c_ϑ. Bipower variation does not depend on the value of the threshold but it is greatly biased. Threshold estimators are less biased; however we can see that

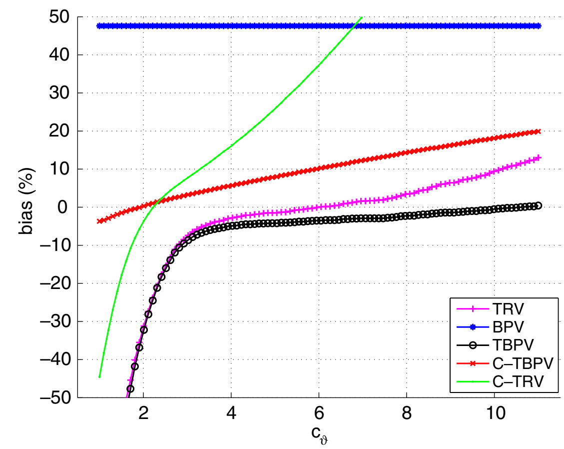

Fig. 1. Relative bias of the different estimators of $\int \sigma_s^2 ds$ in the presence of a single jump as a function of the threshold parameter c_ϑ.

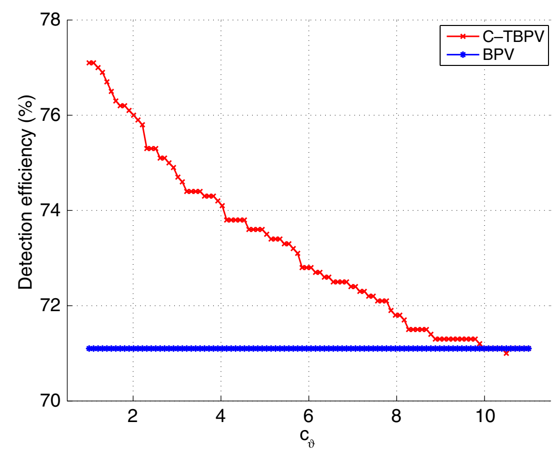

Fig. 2. Jump detection power for the model (4.1) in the presence of a single jump, as a function of the threshold parameter c_ϑ.

threshold bipower variation is less sensitive to the choice of the threshold than threshold power variation. This is basically due to the fact that, even if both TBPV and TRV converge to $[X^c]$ as $\delta \to 0$, for fixed δ we have $\mathrm{TBPV}_\delta \underset{c_\vartheta \to \infty}{\longrightarrow} \mathrm{BPV}_\delta$ while $\mathrm{TRV}_\delta \underset{c_\vartheta \to \infty}{\longrightarrow} \mathrm{RV}_\delta$.

We also use Monte Carlo experiments to evaluate the size and power of the z statistics constructed with multipower variation methods (Barndorff-Nielsen and Shephard, 2006; see Eq. (A.7) in Appendix A for its precise definition), and the C-Tz statistics (3.5). Results with different confidence levels are reported in Table 2 in the case of no jumps, a single jump, and two consecutive jumps. The C-Tz (with $c_\vartheta = 3$) has the largest power in detecting jumps, still preserving a size which is virtually identical to that of the z statistics. With higher confidence level, the advantage in using C-Tz statistics increases. Such an advantage is very large if the jumps in the simulated trajectory are consecutive. In this case, the power of the z test at the 99.99% confidence level is just 42.4%, while the corresponding power of the C-Tz test is 93.1%. We conclude that, on our simulations, there is a clear advantage in using the C-Tz statistics. Again, we can check the sensitivity of C-Tz tests with respect to the choice of the parameter c_ϑ. Fig. 2 shows a direct comparison between threshold estimators in the case of a single jump in every trajectory. We observe that the C-Tz test always has more power than the z test and it is reasonably robust to the choice of c_ϑ.

Summarizing our results on simulated experiments, we conclude that:

1. When measuring integrated volatility in the presence of jumps, bipower variation is greatly upward biased, while threshold-

Table 1
Mean percentage relative error in estimating $\int \sigma_s^2 ds$ and $\int \sigma_s^4 ds$ in the case of no jumps, one jump, two jumps and two consecutive jumps when simulating model (4.1). The estimators that we compare are: the bipower variation used by ABD (BPV) and its staggered version (stag − BPV); the threshold estimator of Mancini (2009) (TRV) and its corrected version (C-TRV); the threshold bipower variation (TBPV) proposed in this paper and defined in Eq. (2.14), and its correction (C-TBPV); and the corresponding estimators for quarticity—see Appendix A. For all estimators, the usual small sample correction $N/(N-(M-1)-k)$ is applied, where $N = [T/\delta]$ and k is the number of addends excluded because of the threshold. In parentheses, the standard error of the mean is reported.

Quantity	Estimator	Relative bias (%)			
		No jumps	One jump	Two jumps	Two consecutive jumps
$\int \sigma_s^2 ds$	BPV	−1.00 (0.53)	48.04 (1.74)	102.03 (3.36)	595.57 (21.07)
	stag − BPV	−1.20 (0.53)	47.60 (1.72)	114.77 (6.32)	97.07 (2.43)
	TRV	−5.56 (0.49)	−5.95 (0.52)	−7.00 (0.53)	−6.93 (0.52)
	C-TRV	−1.39 (0.46)	9.40 (0.55)	18.69 (0.61)	18.94 (0.61)
	TBPV	−4.15 (0.56)	−4.83 (0.60)	−5.65 (0.58)	−4.70 (0.58)
	C-TBPV	−0.58 (0.53)	7.87 (0.62)	15.26 (0.66)	24.57 (0.74)
$\int \sigma_s^4 ds$	QPV	−1.53 (1.33)	101.90 (5.41)	272.32 (22.79)	1601.81 (88.71)
	TQV	−16.32 (0.91)	−15.98(0.94)	−16.47 (1.00)	−16.31 (1.00)
	C-TQV	−4.10 (1.01)	37.75 (1.53)	75.96 (2.00)	77.10 (2.06)
	TQPV	−7.39 (1.28)	−8.92 (1.36)	−12.04 (1.32)	−9.10 (1.36)
	C-TQPV	−1.18 (1.33)	16.52 (1.71)	30.44 (1.94)	57.50 (2.88)
	TriPV	−1.66 (1.24)	210.32 (11.64)	687.56 (94.69)	7841.87 (468.15)
	TTriPV	−7.94 (1.21)	−8.47 (1.28)	−10.76 (1.25)	−8.87 (1.28)
	C-TTriPV	−1.41 (1.25)	18.12 (1.69)	34.42 (1.95)	77.61 (3.16)

Table 2
Percentage of detected jumps in the case of trajectories with no jumps (Panel A), a single jump per day (Panel B), and two consecutive jumps per day (Panel C), for different significance levels. The C-Tz statistic is computed with $c_\vartheta = 3$.

	Panel A No jumps				Panel B Single jump				Panel C Two consecutive jumps			
	50%	95%	99%	99.99%	50%	95%	99%	99.99%	50%	95%	99%	99.99%
z	53.0	5.7	1.4	0.1	93.4	81.2	77.6	68.6	98.1	79.1	64.4	42.4
C-Tz	54.0	6.0	1.6	0.1	93.7	83.6	80.6	74.6	99.2	97.3	96.3	93.1

based estimators are slightly downward biased (with the corrected versions being upward biased, though much less than bipower variation).

2. When measuring integrated variance, threshold bipower variation is nearly insensitive to the choice of the threshold for $c_\vartheta \geq 3$ while threshold realized variance is more sensitive to this choice.
3. When testing for the presence of jumps, tests based on threshold multipower variation yield a substantial advantage with respect to those based on multipower variation.

Thus, in our empirical analysis we will use threshold bipower variation as an estimator of integrated volatility, and C-Tz statistics as a jump detection test.

5. Empirical analysis

Our data set covers a long time span of almost 15 years of high frequency data for the S&P500 futures and the US Treasury Bond with maturity 30 years, and nearly five years for six individual stocks. The purpose of this section is mainly to analyze the impact of jumps on future volatility when threshold bipower variation is employed as a jump measure. We will focus not only on the impact of jumps on future realized variance, but also on the performance of models which explicitly incorporate lagged jumps in the volatility dynamics.

All of the analysis presented in this section is based on measures of threshold multipower variation with the threshold defined as in Eq. (2.15), with $c_\vartheta = 3$ and $\widehat{V}_t$ defined as in Appendix B. When not differently specified, jumps are detected using the C-Tz statistics. Our tables are built at confidence level $\alpha = 99.9\%$ but the most interesting quantities will be computed and plotted for different values of α as well. Further results with $c_\vartheta = 4, 5$ are qualitatively very similar and can be provided upon request.

5.1. The forecasting model

Empirical evidence on strong temporal dependence of realized volatility has been already found for instance in Andersen et al. (2001a,b). This evidence, together with our empirical results reported below, suggests that realized variance series should be described using models allowing for slowly decaying autocorrelation and possibly long memory; see Andersen et al. (2003), Banerjee and Urga (2005) and McAleer and Medeiros (2008). These features are displayed by the HAR model of Corsi (2009), which has been extended by ABD to allow for the separation of quadratic variation into its continuous part and jumps. We then borrow the model of ABD to forecast realized variance. However, when implementing for a finite sample, we use the newly proposed measures based on threshold multipower variation.

Let $t = 1, \ldots$ be the day index and write $\mathrm{RV}_t = \mathrm{RV}_\delta(X)_t$. For two days t_1 and $t_2 \geq t_1$, define

$$\mathrm{RV}_{t_1:t_2} = \frac{1}{t_2 - t_1 + 1} \sum_{t=t_1}^{t_2} \mathrm{RV}_t. \tag{5.1}$$

The HAR model reads

$$\begin{aligned}\mathrm{RV}_{t:t+h-1} &= \beta_0 + \beta_d \mathrm{RV}_{t-1} + \beta_w \mathrm{RV}_{t-5:t-1} \\ &\quad + \beta_m \mathrm{RV}_{t-22:t-1} + \varepsilon_t. \end{aligned} \tag{5.2}$$

ABD proposed the HAR-CJ model[4]:

$$\mathrm{RV}_{t:t+h-1} = \beta_0 + \beta_d \widehat{C}_{t-1} + \beta_w \widehat{C}_{t-5:t-1} + \beta_m \widehat{C}_{t-22:t-1} + \beta_j \widehat{J}_{t-1} + \varepsilon_t. \tag{5.3}$$

The daily jump is defined as

$$\widehat{J}_t = I_{\{z_t > \Phi_\alpha\}} \cdot (\mathrm{RV}_t - \mathrm{BPV}_t)^+, \tag{5.4}$$

with z_t given by Eq. (A.7) in Appendix A and Φ_α being the cumulative distribution function of the normal distribution at confidence level α; $x^+ = \max(x, 0)$. The continuous part of the quadratic variation is defined by

$$\widehat{C}_t = \mathrm{RV}_t - \widehat{J}_t. \tag{5.5}$$

Finally, $\widehat{C}_{t-5:t-1}$ and $\widehat{C}_{t-22:t-1}$ are the weekly and monthly aggregation of $\widehat{C}_t$ as in Eq. (5.1). Using model (5.3), ABD estimate β_j to be not significant and negative, and a very similar conclusion has been reached by Forsberg and Ghysels (2007), Giot and Laurent (2007) and Busch et al. (forthcoming); see also the analysis of Ghysels et al. (2006).

In the light of the above sections, it is natural to define the HAR-TCJ model as

$$\mathrm{RV}_{t:t+h-1} = \beta_0 + \beta_d \widehat{TC}_{t-1} + \beta_w \widehat{TC}_{t-5:t-1} + \beta_m \widehat{TC}_{t-22:t-1} + \beta_j \widehat{TJ}_{t-1} + \varepsilon_t \tag{5.6}$$

where we employ the threshold bipower variation measure to estimate the jump component

$$\widehat{TJ}_t = I_{\{\text{C-Tz} > \Phi_\alpha\}} \cdot (\mathrm{RV}_t - \mathrm{TBPV}_t)^+ \tag{5.7}$$

and the corresponding continuous part $\widehat{TC}_t = \mathrm{RV}_t - \widehat{TJ}_t$.

The square root and logarithmic counterparts of the model read

$$\mathrm{RV}^{\frac{1}{2}}_{t:t+h-1} = \beta_0 + \beta_d \widehat{TC}^{\frac{1}{2}}_t + \beta_w \widehat{TC}^{\frac{1}{2}}_{t-5:t-1} + \beta_m \widehat{TC}^{\frac{1}{2}}_{t-22:t-1} + \beta_j \widehat{TJ}^{\frac{1}{2}}_{t-1} + \varepsilon_t \tag{5.8}$$

and

$$\log \mathrm{RV}_{t:t+h-1} = \beta_0 + \beta_d \log \widehat{TC}_{t-1} + \beta_w \log \widehat{TC}_{t-5:t-1} + \beta_m \log \widehat{TC}_{t-22:t-1} + \beta_j \log \left(\widehat{TJ}_{t-1} + 1\right) + \varepsilon_t \tag{5.9}$$

and the same transformations will be estimated for model (5.2) and (5.3). Everywhere we use annualized volatility measures (one year = 252 days).

To evaluate the forecasting performance of the different models, we use: (a) the R^2 of Mincer–Zarnowitz forecasting regressions; (b) the heteroskedasticity adjusted root mean square error suggested in Bollerslev and Ghysels (1996):

$$\mathrm{HRMSE} = \sqrt{\frac{1}{T}\sum_{t=1}^{T}\left(\frac{\mathrm{RV}_t - \widehat{\mathrm{RV}}_t}{\mathrm{RV}_t}\right)^2} \tag{5.10}$$

where RV_t is the ex post value of realized variance and $\widehat{\mathrm{RV}}_t$ is the forecast; (c) since the HRMSE is not a robust loss function (see Patton, forthcoming), we also employ the QLIKE loss function, defined as

$$\mathrm{QLIKE} = \frac{1}{T}\sum_{t=1}^{T}\left(\log \mathrm{RV}_t + \frac{\widehat{\mathrm{RV}}_t}{\mathrm{RV}_t}\right). \tag{5.11}$$

The QLIKE loss function is robust, in the sense defined by Patton (forthcoming). Significant differences among competing forecasting models are evaluated through the Diebold and Mariano (1995) test at the 5% confidence level.

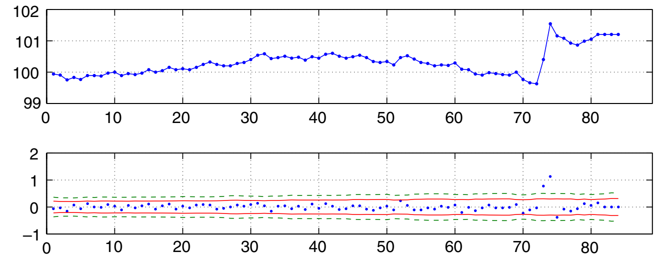

Fig. 3. Rescaled time series (top) and five-minute logarithmic returns (bottom) for the S&P500 on 4th December 1990. The solid and dashed lines show our estimated thresholds with $c_\vartheta = 3$ and $c_\vartheta = 5$ respectively. The jump statistics are $z = -0.2545$, C-Tz $= 4.5055$ with $c_\vartheta = 3$, C-Tz $= 4.4745$ with $c_\vartheta = 5$.

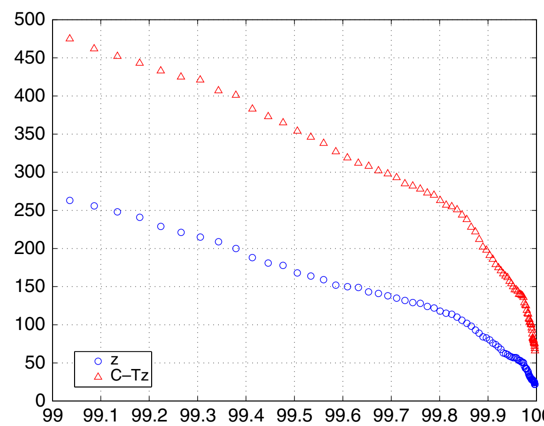

Fig. 4. Number of days which contain jumps in the S&P500 sample obtained with the C-Tz statistics (3.5), as a function of the confidence level α. The total number of days is 3736.

5.2. Stock index futures S&P500 data

The first and most important data set that we analyze is the S&P500 futures time series. We have all high frequency transactions from January 1990 to December 2004 (3736 days). In order to mitigate the impact of microstructure effects on our estimates, we choose, following ABD, a sampling frequency of $\delta = 5$ min, corresponding to 84 returns per day.

Fig. 3 is an example in which using the C-Tz statistics is effective. It displays the S&P500 time series on one specific day, in which there is an abrupt price change. However, on this day, the z statistic, which is based on multipower variation, is negative and does not reveal a jump at any reasonable significance level; while the C-Tz statistic, which is based on threshold multipower variation, does reveal a very significant jump. Our interpretation for this day is that, since the jump appears in the form of two consecutive and very large returns, this creates a huge bias (especially in the quarticity estimates) which makes the z statistic very noisy. This bias is instead completely removed by the corrected threshold estimators.

Fig. 4 shows the number of jumps detected in the sample with the conditions C-Tz $> \Phi_\alpha$ and $z > \Phi_\alpha$ as a function of α. We see that with the statistics based on threshold multipower variation, we detect a higher number of jumps, reflecting the greater power of the C-Tz test.

We then estimate and compare the HAR-TCJ model (5.6) and the HAR-CJ model (5.3). We also estimate the standard HAR model (5.2) as a benchmark. Results are reported in Tables 3–5 for daily, weekly and monthly volatility forecasts respectively.

The results are unambiguous. When the jump component is measured by means of bipower variation, as in the HAR-CJ model, its coefficient is significantly negative for the square

[4] ABD also consider weekly and monthly aggregation of the jump component. In this paper, we focus on the daily component only, to better evaluate the impact of a single jump on future volatility. Weekly and monthly components are considered in Corsi and Renò (2010), together with a heterogeneous leverage component.

Table 3
OLS estimate for daily ($h = 1$) HAR, HAR-CJ and HAR-TCJ volatility forecast regressions for S&P500 futures from January 1990 to December 2004 (3736 observations). The significant daily jumps are computed using a critical value of $\alpha = 99.9\%$ and the C-Tz statistics computed with $c_\vartheta = 3$. Reported in parentheses are the t-statistics based on Newey–West correction with order 5. The performance measures are the Mincer–Zarnowitz R^2, the *HRMSE* as in Eq. (5.10) and the *QLIKE* as in (5.11), computed unconditionally, conditionally on having had a jump at time $t-1$ (J-R^2, *J-HRMSE*, *J-QLIKE*) and conditionally on no jump at time $t-1$ (C-R^2, *C-HRMSE*, *C-QLIKE*), using the Diebold–Mariano test at the 5% confidence level.

Daily ($h = 1$) S&P500 regressions

HAR: $\mathrm{RV}_{t:t+h-1} = \beta_0 + \beta_d \mathrm{RV}_{t-1} + \beta_w \mathrm{RV}_{t-5:t-1} + \beta_m \mathrm{RV}_{t-22:t-1} + \varepsilon_t$
HAR-CJ: $\mathrm{RV}_{t:t+h-1} = \beta_0 + \beta_d \widehat{C}_{t-1} + \beta_w \widehat{C}_{t-5:t-1} + \beta_m \widehat{C}_{t-22:t-1} + \beta_j \widehat{J}_{t-1} + \varepsilon_t$
HAR-TCJ: $\mathrm{RV}_{t:t+h-1} = \beta_0 + \beta_d \widehat{TC}_{t-1} + \beta_w \widehat{TC}_{t-5:t-1} + \beta_m \widehat{TC}_{t-22:t-1} + \beta_j \widehat{TJ}_{t-1} + \varepsilon_t$

	RV_t			$\mathrm{RV}_t^{1/2}$			$\log \mathrm{RV}_t$		
	HAR	HAR-CJ	HAR-TCJ	HAR	HAR-CJ	HAR-TCJ	HAR	HAR-CJ	HAR-TCJ
β_0	34.266	26.823	23.312	0.986	0.906	0.771	0.254	0.274	0.286
	(3.779)	(3.373)	(2.760)	(3.973)	(3.817)	(3.318)	(4.498)	(4.906)	(5.206)
β_d	0.220	0.378	0.420	0.323	0.371	0.421	0.335	0.336	0.356
	(2.329)	(5.736)	(6.171)	(6.348)	(8.961)	(11.958)	(13.143)	(13.146)	(14.838)
β_w	0.321	0.263	0.298	0.336	0.317	0.308	0.359	0.356	0.342
	(3.820)	(3.156)	(2.608)	(6.077)	(5.812)	(5.427)	(9.489)	(9.420)	(9.441)
β_m	0.313	0.288	0.253	0.268	0.260	0.237	0.256	0.256	0.252
	(4.814)	(4.783)	(3.850)	(6.820)	(6.704)	(6.081)	(8.829)	(8.916)	(9.028)
β_j		−0.581	0.045		−0.101	0.096		0.007	0.055
		(−2.968)	(0.925)		(−1.627)	(2.653)		(0.680)	(6.386)
R^2	0.339	0.374	0.387	0.583	0.592	0.604	0.679	0.681	0.684
HRMSE	0.867	0.794[a]	0.749[a,b]	0.652	0.630	0.605[a,b]	0.555	0.552[a]	0.537[a,b]
QLIKE	6.429	6.390[a]	6.371[a,b]	6.228	6.223[a]	6.216[a,b]	6.139	6.139	6.136
J-R^2	0.320	0.363	0.386	0.573	0.596	0.626	0.685	0.691	0.707
J-HRMSE	1.137	0.955[a]	0.778[a,b]	0.799	0.729[a]	0.624[a,b]	0.644	0.611[a]	0.530[a,b]
J-QLIKE	6.513	6.405[a]	6.317[a,b]	6.265	6.214[a]	6.149[a,b]	6.148	6.114[a]	6.054[a,b]
C-R^2	0.341	0.375	0.388	0.584	0.592	0.603	0.679	0.680	0.682
C-HRMSE	0.846	0.782[a]	0.747[a,b]	0.640	0.623	0.604	0.548	0.547	0.537
C-QLIKE	6.423	6.389[a]	6.375[a,b]	6.226	6.224	6.220	6.139	6.140	6.142

[a] Denotes significant improvement in the forecasting performance with respect to the HAR model.
[b] Denotes significant improvement in the forecasting performance with respect to the HAR-CJ model.

Table 4
OLS estimate for weekly ($h = 5$) HAR, HAR-CJ and HAR-TCJ volatility forecast regressions for S&P500 futures from January 1990 to December 2004 (3736 observations). The significant daily jumps are computed using a critical value of $\alpha = 99.9\%$. Reported in parentheses are the t-statistics based on Newey–West correction with order 10. The performance measures are the Mincer–Zarnowitz R^2, the *HRMSE* as in Eq. (5.10) and the *QLIKE* as in (5.11), computed unconditionally, conditionally on having had at jump a time $t-1$(J-R^2, *J-HRMSE*, *J-QLIKE*) and conditionally on no jump at time $t-1$(C-R^2, *C-HRMSE*, *C-QLIKE*), using the Diebold–Mariano test at the 5% confidence level.

Weekly ($h = 5$) S&P500 regressions

HAR: $\mathrm{RV}_{t:t+h-1} = \beta_0 + \beta_d \mathrm{RV}_{t-1} + \beta_w \mathrm{RV}_{t-5:t-1} + \beta_m \mathrm{RV}_{t-22:t-1} + \varepsilon_t$
HAR-CJ: $\mathrm{RV}_{t:t+h-1} = \beta_0 + \beta_d \widehat{C}_{t-1} + \beta_w \widehat{C}_{t-5:t-1} + \beta_m \widehat{C}_{t-22:t-1} + \beta_j \widehat{J}_{t-1} + \varepsilon_t$
HAR-TCJ: $\mathrm{RV}_{t:t+h-1} = \beta_0 + \beta_d \widehat{TC}_{t-1} + \beta_w \widehat{TC}_{t-5:t-1} + \beta_m \widehat{TC}_{t-22:t-1} + \beta_j \widehat{TJ}_{t-1} + \varepsilon_t$

	$\mathrm{RV}_{t:t+4}$			$\mathrm{RV}_{t:t+4}^{1/2}$			$\log \mathrm{RV}_{t:t+4}$		
	HAR	HAR-CJ	HAR-TCJ	HAR	HAR-CJ	HAR-TCJ	HAR	HAR-CJ	HAR-TCJ
β_0	47.320	41.284	37.873	1.541	1.465	1.348	0.407	0.426	0.438
	(4.335)	(3.987)	(3.384)	(4.334)	(4.252)	(3.966)	(4.843)	(5.118)	(5.384)
β_d	0.097	0.190	0.210	0.176	0.213	0.244	0.205	0.211	0.229
	(1.892)	(4.858)	(4.403)	(5.360)	(7.927)	(9.395)	(11.744)	(12.132)	(13.821)
β_w	0.367	0.351	0.393	0.369	0.356	0.353	0.360	0.353	0.333
	(4.677)	(4.299)	(3.564)	(6.157)	(6.006)	(5.869)	(8.245)	(8.148)	(8.070)
β_m	0.335	0.320	0.304	0.342	0.337	0.329	0.353	0.355	0.359
	(4.897)	(4.371)	(3.702)	(6.361)	(6.174)	(6.028)	(8.548)	(8.728)	(9.159)
β_j		−0.394	0.007		−0.105	0.040		−0.005	0.031
		(−2.571)	(0.379)		(−2.427)	(2.466)		(−0.693)	(6.193)
R^2	0.499	0.534	0.554	0.690	0.700	0.709	0.767	0.770	0.772
HRMSE	0.634	0.589[a]	0.566[a,b]	0.439	0.425[a]	0.417[a,b]	0.377	0.372[a]	0.369[a]
QLIKE	6.401	6.376[a]	6.366[a,b]	6.194	6.190[a]	6.187	6.098	6.097	6.097
J-R^2	0.489	0.539	0.567	0.656	0.684	0.712	0.742	0.751	0.766
J-HRMSE	0.761	0.652[a]	0.545[a,b]	0.526	0.465[a]	0.397[a,b]	0.415	0.391[a]	0.355[a,b]
J-QLIKE	6.456	6.377[a]	6.300[a,b]	6.220	6.173[a]	6.114[a,b]	6.101	6.069[a]	6.019[a,b]
C-R^2	0.503	0.536	0.553	0.693	0.701	0.709	0.769	0.771	0.773
C-HRMSE	0.625	0.584[a]	0.568[a,b]	0.432	0.423[a]	0.418[a]	0.373	0.371[a]	0.369[a]
C-QLIKE	6.397	6.376[a]	6.371[a]	6.193	6.191	6.192	6.098	6.099	6.102

[a] Denotes significant improvement in the forecasting performance with respect to the HAR model.
[b] Denotes significant improvement in the forecasting performance with respect to the HAR-CJ model.

model and insignificant for the log and square root models in explaining future volatility. This result, although consistent with the literature, is rather surprising in our opinion, being at odds with the economic intuition which would suggest an increase in volatility after a jump in the price process. Moreover, this result is even more puzzling, given that the unconditional correlations between realized variation and jumps lagged by one day is significantly positive and around 20% for the variances, 30% for the volatilities and 25% for the log volatilities. However, the simulated experiments in Section 4 suggest that when the jump component is

Table 5
OLS estimate for monthly ($h = 22$) HAR, HAR-CJ and HAR-TCJ volatility forecast regressions for S&P500 futures from January 1990 to December 2004 (3736 observations). The significant daily jumps are computed using a critical value of $\alpha = 99.9\%$. Reported in parentheses are the t-statistics based on Newey–West correction with order 44. The performance measures are the Mincer–Zarnowitz R^2, the *HRMSE* as in Eq. (5.10) and the *QLIKE* as in (5.11), computed unconditionally, conditionally on having had a jump at time $t-1$ (J-R^2, J-*HRMSE*, J-*QLIKE*) and conditionally on no jump at time $t-1$ (C-R^2, C-*HRMSE*, C-*QLIKE*), using the Diebold–Mariano test at the 5% confidence level.

Monthly ($h = 22$) S&P500 regressions

HAR: $\mathrm{RV}_{t:t+h-1} = \beta_0 + \beta_d \mathrm{RV}_{t-1} + \beta_w \mathrm{RV}_{t-5:t-1} + \beta_m \mathrm{RV}_{t-22:t-1} + \varepsilon_t$
HAR-CJ: $\mathrm{RV}_{t:t+h-1} = \beta_0 + \beta_d \widehat{C}_{t-1} + \beta_w \widehat{C}_{t-5:t-1} + \beta_m \widehat{C}_{t-22:t-1} + \beta_j \widehat{J}_{t-1} + \varepsilon_t$
HAR-TCJ: $\mathrm{RV}_{t:t+h-1} = \beta_0 + \beta_d \widehat{TC}_{t-1} + \beta_w \widehat{TC}_{t-5:t-1} + \beta_m \widehat{TC}_{t-22:t-1} + \beta_j \widehat{TJ}_{t-1} + \varepsilon_t$

	$\mathrm{RV}_{t:t+21}$			$\mathrm{RV}^{1/2}_{t:t+21}$			$\log \mathrm{RV}_{t:t+21}$		
	HAR	HAR-CJ	HAR-TCJ	HAR	HAR-CJ	HAR-TCJ	HAR	HAR-CJ	HAR-TCJ
β_0	78.538	73.578	70.785	2.883	2.808	2.699	0.758	0.771	0.778
	(5.909)	(5.638)	(4.962)	(5.892)	(5.967)	(5.883)	(5.292)	(5.492)	(5.702)
β_d	0.061	0.124	0.129	0.109	0.135	0.149	0.126	0.130	0.138
	(2.555)	(5.292)	(4.876)	(5.915)	(8.410)	(9.183)	(10.055)	(10.519)	(11.717)
β_w	0.219	0.206	0.243	0.281	0.273	0.281	0.269	0.265	0.253
	(4.279)	(3.746)	(4.350)	(5.579)	(5.385)	(5.405)	(6.171)	(6.093)	(5.927)
β_m	0.385	0.385	0.385	0.399	0.397	0.395	0.454	0.456	0.462
	(4.595)	(4.310)	(4.022)	(6.358)	(6.291)	(6.373)	(9.144)	(9.246)	(9.728)
β_j		−0.284	−0.002		−0.092	0.018		−0.010	0.016
		(−2.892)	(−0.114)		(−3.278)	(1.444)		(−1.746)	(3.658)
R^2	0.471	0.500	0.517	0.649	0.659	0.667	0.739	0.742	0.746
HRMSE	0.680	0.649[a]	0.639[a,b]	0.438	0.429[a]	0.425[a]	0.376	0.373[a]	0.370[a]
QLIKE	6.483	6.467[a]	6.462[a]	6.227	6.224	6.222	6.106	6.105	6.105
J-R^2	0.438	0.476	0.513	0.651	0.672	0.692	0.746	0.755	0.761
J-*HRMSE*	0.753	0.690[a]	0.653[a,b]	0.475	0.443[a]	0.420[a,b]	0.389	0.375[a]	0.367[a]
J-*QLIKE*	6.520	6.465[a]	6.416[a,b]	6.237	6.200[a]	6.156[a,b]	6.096	6.069[a]	6.033[a,b]
C-R^2	0.474	0.502	0.517	0.650	0.658	0.666	0.738	0.742	0.745
C-*HRMSE*	0.675	0.647[a]	0.638[a]	0.435	0.428[a]	0.425	0.374	0.372[a]	0.370[a]
C-*QLIKE*	6.481	6.467[a]	6.465[a]	6.227	6.226	6.227	6.107	6.108	6.110

[a] Denotes significant improvement in the forecasting performance with respect to the HAR model.
[b] Denotes significant improvement in the forecasting performance with respect to the HAR-CJ model.

estimated via Eq. (5.4) it is greatly downward biased because of the large upward bias of bipower variation in the presence of jumps. As a consequence, the continuous component $\widehat{C}_t$ estimated using bipower variation is still contaminated by the jump component and hence the impact of jumps is also passing through the positive coefficients of the other regressors.

When instead the jump component is measured by means of threshold bipower variation, the estimated coefficients $\widehat{\beta}_j$ are positive and highly significant for the square root and log models. Most importantly, the HAR-TCJ model yields a higher R^2 and a significantly lower HRMSE and QLIKE (as witnessed by the Diebold–Mariano test at the 5% confidence level), which implies a better forecasting power. To better understand this point, we divide the sample into days immediately following the occurrence of a jump and the remainder. For these two samples, we compute the R^2, HRMSE and QLIKE statistics separately, denoting them by J-R^2, J-HRMSE, J-QLIKE and C-R^2, C-HRMSE, C-QLIKE respectively. The results in Table 3 show that the TCJ models greatly and significantly improve the forecasting power for realized variance in days immediately following a jump, and are still slightly better performing in days that do not follow a jump. Our interpretation of this result is that, since we are measuring the jump component better, we are also removing noise from the continuous component in the explanatory variables; and thus, we also get slightly better results on days in which there were no jumps before.

Fig. 5 displays the most important quantities as a function of the confidence interval α. It shows that the jump component of the HAR-TCJ model, as measured using the t-statistics of the coefficient β_j, is always positive for all models and highly significant for square root and log transformations; while the jump component of the HAR-CJ model is mainly significantly negative or not significant. Importantly, it shows that the HAR-TCJ model provides superior forecasts when measured in terms of R^2 and the HRMSE, irrespective of the confidence level used and model employed.

When forecasting weekly and monthly volatility, the β_j of the HAR-CJ model tends to be negative, sometimes significantly. However, for the HAR-TCJ model, the β_j are greatly positive and significant in the square root and log models, and insignificant in the square model. Again, the R^2 and the HRMSE confirm, in days following a jump, the better forecasting ability of the HAR-TCJ model, which is not worse than HAR-CJ in days not following a jump.[5]

5.3. Individual stocks

We analyze a sample of six individual stocks, chosen among the most liquid stocks of S&P500. The stocks are Alcoa (aa), Citigroup (c), Intel (intc), Microsoft (msft), Pfeizer (pfe) and Exxon-Mobil (xom). Our sample starts on 2 January 2001 and ends on 30 December 2005, containing 1256 days per stock. Since these stocks are traded very actively, we still use a sampling frequency of $\delta = 5$ min, corresponding to 78 returns per day. To save space, we focus on the most important quantities (the significance of the jump and the R^2, HRMSE and QLIKE of the forecasting model conditional on days after the occurrence of a jump), which are reported in Table 6 for the logarithmic model. We report results for $\alpha = 99\%$ and $\alpha = 99.9\%$.

Despite the smaller sample size and increased idiosyncratic noise, jumps still have a substantial impact in determining future volatility which can be revealed by using threshold bipower

[5] The analysis with higher values of c_ϑ reveals that the β_j coefficient of the TCJ specification is mildly significant for $c_\vartheta = 4$ and not significant for $c_\vartheta = 5$. This is not surprising, since as we increase c_ϑ we get closer to the results obtained with bipower variation. We also estimated the HAR-CJ model using the jumps detected via the z statistics (A.7) and compare it with the HAR-TCJ model estimated with the jumps detected via the C-Tz statistics (3.5). The results indicate that in this case the difference between the two models is even more pronouncedly in favor of the HAR-TCJ.

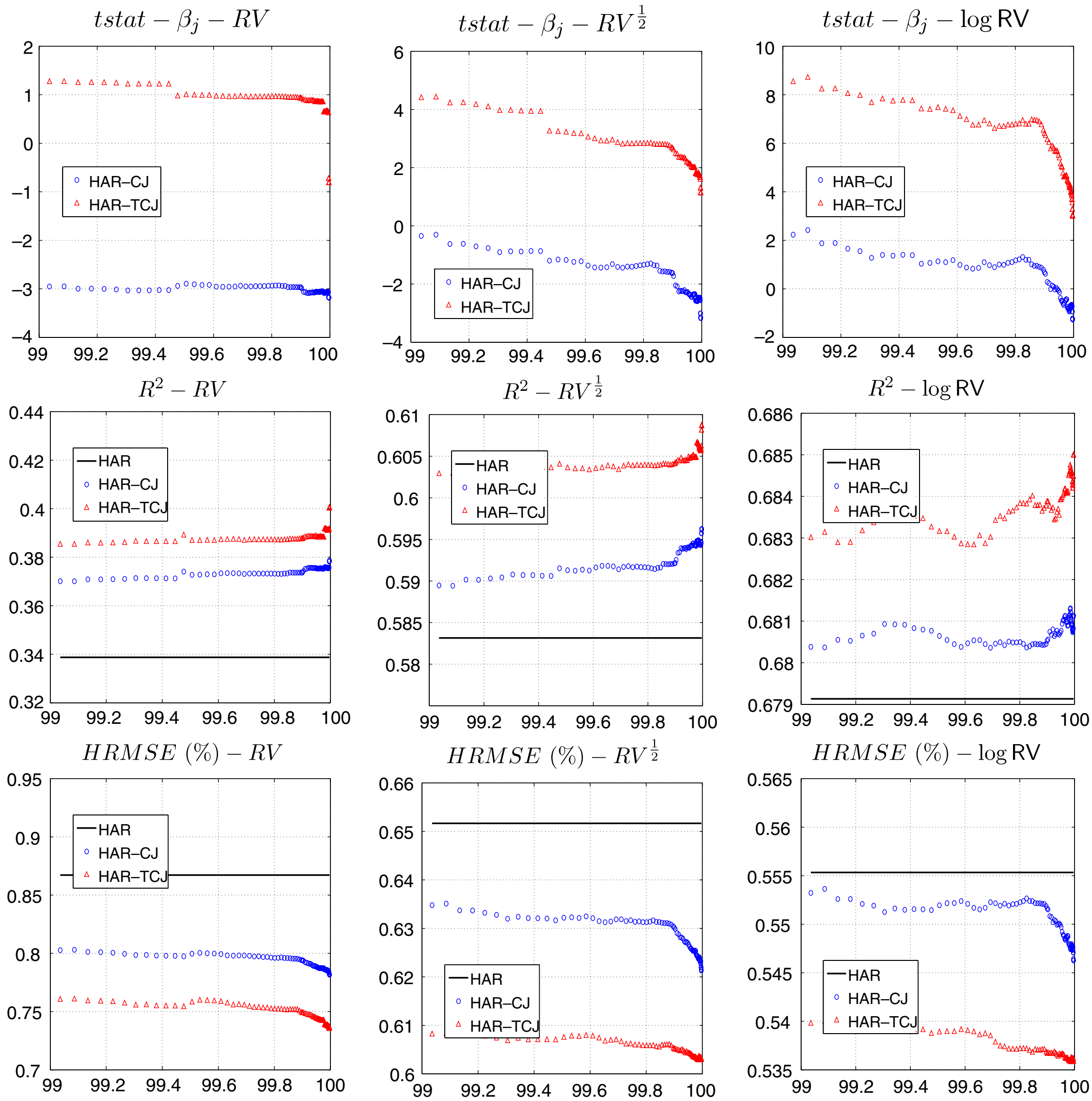

Fig. 5. The t statistics of the coefficient β_j measuring the impact of jumps on future volatility, the R^2 and the HRMSE for the three models estimated on S&P500 data for daily forecasting, for both the HAR-CJ and HAR-TCJ versions, as a function of the confidence level used for detecting jumps with the C-Tz statistics.

Table 6

Number of jumps, and t-statistics on β_j, J-R^2, J-*HRMSE* and J-*QLIKE* for the daily ($h = 1$) *logarithmic* version of HAR, HAR-CJ and HAR-TCJ volatility forecast regressions on six individual stocks. The significant daily jumps are computed using a critical value of $\alpha = 0.99$ and $\alpha = 0.999$ as reported, with the C-Tz statistics, using the Diebold–Mariano test at the 5% confidence level.

Stock	α (%)	Jumps	β_j t-statistic		J-R^2			J-*HRMSE*			J-*QLIKE*		
			HAR-CJ	HAR-TCJ	HAR	HAR-CJ	HAR-TCJ	HAR	HAR-CJ	HAR-TCJ	HAR	HAR-CJ	HAR-TCJ
aa	99.0	242	0.829	3.378	0.551	0.560	0.546	0.495	0.466[a]	0.450[a]	−1.589	−1.623[a]	−1.665[a,b]
	99.9	121	0.373	2.671	0.607	0.628	0.614	0.455	0.416[a]	0.399[a]	−1.582	−1.627[a]	−1.685[a,b]
c	99.0	196	0.746	3.677	0.827	0.830	0.835	0.538	0.520	0.486	−2.045	−2.073[a]	−2.104[a,b]
	99.9	105	−0.410	3.967	0.842	0.847	0.851	0.457	0.435[a]	0.418[a,b]	−2.091	−2.127[a]	−2.164[a,b]
intc	99.0	155	3.471	5.572	0.784	0.786	0.784	0.512	0.485	0.463[a]	−1.472	−1.506[a]	−1.535[a,b]
	99.9	78	3.112	4.935	0.831	0.827	0.821	0.474	0.460[a]	0.430	−1.401	−1.443[a]	−1.485[a,b]
msft	99.0	202	−0.305	2.899	0.807	0.813	0.815	0.604	0.547	0.506	−2.030	−2.060[a]	−2.087[a,b]
	99.9	92	0.252	2.190	0.799	0.805	0.810	0.755	0.673	0.602	−2.020	−2.058[a]	−2.102[a,b]
pfe	99.0	239	0.524	2.469	0.600	0.617	0.609	0.655	0.570	0.560	−1.956	−1.988[a]	−2.013[a,b]
	99.9	131	0.391	2.979	0.591	0.628	0.612	0.715	0.564	0.551	−1.953	−2.001	−2.033[a,b]
xom	99.0	193	1.424	3.224	0.732	0.735	0.721	0.500	0.475[a]	0.464[a]	−2.296	−2.324[a]	−2.360[a,b]
	99.9	98	0.472	2.656	0.695	0.702	0.695	0.523	0.479[a]	0.454[a,b]	−2.305	−2.348[a]	−2.394[a,b]

[a] Denotes significant improvement in the forecasting performance with respect to the HAR model.
[b] Denotes significant improvement in the forecasting performance with respect to the HAR-CJ model.

variation. The t statistic of the β_j coefficient is always larger for the HAR-TCJ model than for the HAR-CJ model, and always significant. On the whole sample, the performances of the two models are virtually the same but, conditioned on the occurrence of a jump, there is a significant advantage in using the HAR-TCJ model, especially when significance is evaluated by the robust

QLIKE. Thus, the results obtained for the S&P500 portfolio are qualitatively replicated for its most liquid constituents, indicating that the impact of jumps on future volatility is not peculiar to the S&P500 returns considered in the previous section, and suggesting that it may simply come from aggregation, at the portfolio level, of the same effect at the individual stock level.

5.4. Bond data and the no-trade bias

Finally, we use a sample of 30-year US Treasury Bond futures from January 1990 to October 2003 for a total of 3231 daily data points. All the relevant volatility and jump measures are computed again with five-minute returns, corresponding to 84 returns per day.

The first thing we note for bond data is a surprisingly high number of jumps. At the 99.9% confidence level, the C-Tz statistics detects 570 jumps, corresponding to the 17.6% of our sample. Visual inspection of time series data reveals that on most of these days there are many intervals with zero return instead.

The problem hinges on what we could call *the no-trade bias* of bipower variation. This can be explained as follows. Suppose that data are not recorded continuously but, more realistically, are recorded discretely. Denote by $\bar{\delta}$ the minimum distance between two subsequent observations. By construction, if $\delta < \bar{\delta}$, then $\mathsf{MPV}_\delta = 0$ identically! Clearly, also $\mathsf{TMPV}_\delta = 0$. This simple reasoning also explains why the presence of null returns caused by the absence of trading (stale price) in that interval induces a downward bias in multipower variation measures. Note that realized variance is immune from this bias instead. Moreover, this bias has a larger impact on the jump detecting statistics, pointing toward rejection of the null. For example, when considering the z statistics, this bias lowers both the BPV and TriPV measure, with the joint effect of increasing the numerator and decreasing the denominator, thus increasing the statistics considerably.

In our paper, this problem does not affect the S&P500 index, or the stocks considered in our empirical analysis. However, it may affect US bond data, which are greatly less liquid. Indeed, the percentage of zero five-minute return in bond data is very high, nearly 30%. We accommodate this problem as follows: for bond data, we compute the C-Tz statistics using only returns different from zero. Clearly, this biases the test toward the null, meaning that the detected jumps are those which have a larger impact for the statistics. With this correction the number of significant jumps with $\alpha = 99.9\%$ reduced to 112, representing 3.4% of the sample. Further analysis and different possible corrections for this problem are discussed in Schulz (2010).

Relevant estimation results for bond data when forecasting daily, weekly and monthly volatility are shown in Table 7 for $\alpha = 99.9\%$. We find that the HAR-TCJ model outperforms the HAR-CJ model (significantly after a jump), while both significantly outperform the HAR model. Also the impact of the jump on future volatility in the HAR-TCJ model is generally insignificant, but nevertheless not negatively significant as for the HAR-CJ estimates. An explanation for this finding might be that jumps in the bond market are less "surprising" than those in the stock markets, since most of them are related to scheduled macroeconomic announcements. Indeed, related studies, for example Bollerslev et al. (2000), find two intraday spikes at hours in which announcements are released.

Summarizing, our empirical findings further corroborate the theoretical and simulation results in the previous sections on the superior performance of the threshold method in separating and estimating the continuous and jump components of the price process. Moreover, they show that, once the two components are correctly measured and separated, the impact of past jumps on future realized variance is positive and significant, and the ability to forecast volatility increases significantly.

6. Conclusions

This paper shows that dividing volatility into jumps and continuous variation yields a substantial improvement in volatility forecasting, because of the positive impact of past jumps on future volatility. This important result has been obscured in the literature since, for finite samples, measures based on multipower variation are greatly biased in the presence of jumps. We uncover this effect by modifying bipower variation with the help of threshold estimation techniques. We show that the newly defined estimator is robust to the presence of jumps and quite inelastic with respect to the choice of the threshold. The class of TMPV estimators introduced in this paper admits a central limit theorem in the presence of jumps.

Our empirical results, obtained on US stock index, individual stocks and Treasury Bond data, also show that jumps can be effectively detected using the newly proposed C-Tz statistics, which is based on TMPV estimators. The models that we propose provide a significantly superior forecasting ability, especially for days which follow the occurrence of a jump. Clearly, these findings can be of great importance for risk management and other financial applications involving volatility estimation. Moreover, the correlation between jumps and future volatility can be helpful not only for practical applications, but also for a deeper comprehension of the asset price dynamics. We consider this line of research as an interesting direction for future studies.

Appendix A. Quarticity estimators and jump detection statistics

For estimation of the integrated quarticity $\int_t^{t+T} \sigma_s^4 \mathrm{d}s$, the literature focused on the following quantities:

$$\begin{aligned} \mathsf{QPV}_\delta(X)_t &= \mu_1^{-4} \cdot \mathsf{MPV}_\delta(X)_t^{[1,1,1,1]} \\ &= \mu_1^{-4} \frac{1}{\delta} \sum_{j=4}^{[T/\delta]} |\Delta_{j-3}X| \cdot |\Delta_{j-2}X| \cdot |\Delta_{j-1}X| \cdot |\Delta_j X| \end{aligned} \tag{A.1}$$

$$\begin{aligned} \mathsf{TriPV}_\delta(X)_t &= \mu_{\frac{4}{3}}^{-3} \cdot \mathsf{MPV}_\delta(X)_t^{\left[\frac{4}{3},\frac{4}{3},\frac{4}{3}\right]} \\ &= \mu_{\frac{4}{3}}^{-3} \frac{1}{\delta} \sum_{j=3}^{[T/\delta]} |\Delta_{j-2}X|^{\frac{4}{3}} \cdot |\Delta_{j-1}X|^{\frac{4}{3}} \cdot |\Delta_j X|^{\frac{4}{3}} \end{aligned} \tag{A.2}$$

where $\mu_{\frac{4}{3}} \simeq 0.8309$, and

$$\mathsf{TQV}_\delta(X)_t = \frac{1}{3\delta} \sum_{j=1}^{[T/\delta]} |\Delta_j X|^4 I_{\left\{|\Delta_j X|^2 \le \Theta(\delta)\right\}} \tag{A.3}$$

where $\Theta(\delta)$ satisfies (2.9). We propose the counterparts $\mathsf{TQPV} = \mu_1^{-4} \cdot \mathsf{TMPV}_\delta(X)_t^{[1,1,1,1]}$, $\mathsf{TTriPV} = \mu_{\frac{4}{3}}^{-3} \cdot \mathsf{TMPV}_\delta(X)_t^{\left[\frac{4}{3},\frac{4}{3},\frac{4}{3}\right]}$ (see Eq. (2.10)), and their corrected version

$$\mathsf{C\text{-}TQV}_\delta(X)_t = \mu_1^{-4} \cdot \mathsf{C\text{-}TMPV}_\delta(X)_t^{[1,1,1,1]} \tag{A.4}$$

$$\mathsf{C\text{-}TTriPV}_\delta(X)_t = \mu_{\frac{4}{3}}^{-3} \cdot \mathsf{C\text{-}TMPV}_\delta(X)_t^{\left[\frac{4}{3},\frac{4}{3},\frac{4}{3}\right]} \tag{A.5}$$

(see definition (3.2)). We also define a corrected version of (A.3):

$$\mathsf{C\text{-}TQV}_\delta(X)_t = \frac{1}{3\delta} \sum_{j=1}^{[T/\delta]} Z_4(\Delta_j X, \vartheta_j). \tag{A.6}$$

A feasible jump test can be constructed using multipower estimators of the integrated quarticity given by Eq. (A.1)–(A.2). In their study, ABD use the test

Table 7
OLS (partial) estimates for daily ($h = 1$), weekly ($h = 5$), monthly ($h = 22$) HAR, HAR-CJ and HAR-TCJ volatility forecast regressions for US Bonds from January 1990 to December 2004 (3736 observations). The significant daily jumps are computed using a critical value of $\alpha = 99.9\%$ and the C-Tz statistics. Reported in parentheses are the t-statistics based on Newey–West correction. The performance measures are the Mincer–Zarnowitz R^2, the *HRMSE* as in Eq. (5.10) and the *QLIKE* as in (5.11), computed unconditionally, conditionally on having had a jump at time $t-1$ (J-R^2, J-*HRMSE*, J-*QLIKE*) and conditionally on no jump at time $t-1$(C-R^2, C-*HRMSE*, C-*QLIKE*), using the Diebold–Mariano test at the 5% confidence level.

US Bond regressions (C-Tz statistics)

HAR: $\mathrm{RV}_{t:t+h-1} = \beta_0 + \beta_d \mathrm{RV}_{t-1} + \beta_w \mathrm{RV}_{t-5:t-1} + \beta_m \mathrm{RV}_{t-22:t-1} + \varepsilon_t$
HAR-CJ: $\mathrm{RV}_{t:t+h-1} = \beta_0 + \beta_d \widehat{C}_{t-1} + \beta_w \widehat{C}_{t-5:t-1} + \beta_m \widehat{C}_{t-22:t-1} + \beta_j \widehat{J}_{t-1} + \varepsilon_t$
HAR-TCJ: $\mathrm{RV}_{t:t+h-1} = \beta_0 + \beta_d \widehat{TC}_{t-1} + \beta_w \widehat{TC}_{t-5:t-1} + \beta_m \widehat{TC}_{t-22:t-1} + \beta_j \widehat{TJ}_{t-1} + \varepsilon_t$

	$\mathrm{RV}_{t:t+h-1}$			$\mathrm{RV}^{1/2}_{t:t+h-1}$			$\log \mathrm{RV}_{t:t+h-1}$		
	HAR	HAR-CJ	HAR-TCJ	HAR	HAR-CJ	HAR-TCJ	HAR	HAR-CJ	HAR-TCJ
Daily forecasts ($h = 1$)									
β_j		−0.108	0.030		−0.041	0.029		−0.012	0.016
		(−2.633)	(1.445)		(−1.932)	(1.798)		(−1.393)	(2.015)
R^2	0.124	0.144	0.143	0.204	0.216	0.214	0.250	0.256	0.252
HRMSE	0.907	0.872[a]	0.866[a]	0.754	0.737[a]	0.733[a]	0.657	0.649[a]	0.645[a]
QLIKE	5.508	5.491[a]	5.492[a]	5.362	5.356[a]	5.356[a]	5.256	5.254	5.254
J-R^2	0.096	0.119	0.108	0.160	0.169	0.161	0.200	0.201	0.192
J-*HRMSE*	1.326	1.053	0.890[a,b]	1.113	0.906	0.763[b]	0.934	0.796	0.677[a,b]
J-*QLIKE*	5.537	5.455[a]	5.398[a,b]	5.355	5.294[a]	5.236[a,b]	5.222	5.175[a]	5.125[a,b]
C-R^2	0.129	0.144	0.142	0.209	0.218	0.216	0.255	0.261	0.256
C-*HRMSE*	0.888	0.865[a]	0.866[a]	0.738	0.731[a]	0.732	0.645	0.642	0.643
C-*QLIKE*	5.502	5.499	5.512	5.364	5.370	5.382	5.264	5.271	5.282
Weekly forecasts ($h = 5$)									
β_j		−0.063	0.028		−0.030	0.029		−0.010	0.018
		(−3.119)	(2.794)		(−2.625)	(3.582)		(−2.010)	(4.036)
R^2	0.295	0.333	0.332	0.415	0.438	0.436	0.472	0.486	0.480
HRMSE	0.445	0.422[a]	0.419[a]	0.363	0.353[a]	0.351[a]	0.331	0.327[a]	0.325[a]
QLIKE	5.370	5.361[a]	5.363	5.250	5.248	5.249	5.162	5.162	5.163
J-R^2	0.197	0.233	0.232	0.330	0.351	0.344	0.400	0.411	0.400
J-*HRMSE*	0.629	0.498[a]	0.417[a,b]	0.469	0.386[a]	0.338[a,b]	0.367	0.326[a]	0.306[a]
J-*QLIKE*	5.364	5.310[a]	5.270[a,b]	5.221	5.181[a]	5.144[a,b]	5.115	5.086[a]	5.056[a,b]
C-R^2	0.315	0.347	0.347	0.428	0.448	0.446	0.482	0.493	0.488
C-*HRMSE*	0.437	0.419[a]	0.420[a]	0.358	0.352[a]	0.352[a]	0.329	0.326[a]	0.325[a]
C-*QLIKE*	5.371	5.372	5.383	5.257	5.262	5.271	5.173	5.178	5.186
Monthly forecasts ($h = 22$)									
β_j		−0.058	0.010		−0.034	0.010		−0.013	0.006
		(−3.986)	(1.586)		(−3.871)	(1.743)		(−3.259)	(1.875)
R^2	0.333	0.387	0.395	0.433	0.473	0.481	0.488	0.514	0.519
HRMSE	0.321	0.305[a]	0.303[a]	0.264	0.258[a]	0.257[a]	0.262	0.260	0.260
QLIKE	5.351	5.344	5.345	5.233	5.231	5.231	5.144	5.143	5.144
J-R^2	0.296	0.353	0.365	0.403	0.442	0.452	0.461	0.484	0.484
J-*HRMSE*	0.398	0.313[a]	0.271[a,b]	0.281	0.240[a]	0.235[a]	0.234	0.229	0.250
J-*QLIKE*	5.342	5.300[a]	5.267[a,b]	5.204	5.173[a]	5.142[a,b]	5.100	5.077[a]	5.052[a,b]
C-R^2	0.339	0.387	0.393	0.436	0.471	0.478	0.489	0.513	0.517
C-*HRMSE*	0.318	0.305[a]	0.304[a]	0.263	0.259[a]	0.258[a]	0.262	0.260	0.260
C-*QLIKE*	5.353	5.354	5.361	5.240	5.243	5.250	5.153	5.158	5.164

[a] Denotes significant improvement in the forecasting performance with respect to the HAR model.
[b] Denotes significant improvement in the forecasting performance with respect to the HAR-CJ model.

$$z = \delta^{-\frac{1}{2}} \frac{(\mathrm{RV}_\delta(X)_T - \mathrm{BPV}_\delta(X)_T) \cdot \mathrm{RV}_\delta(X)_T^{-1}}{\sqrt{\bar{\vartheta} \max\left\{1, \frac{\mathrm{TriPV}_\delta(X)_T}{(\mathrm{BPV}_\delta(X)_T)^2}\right\}}}, \tag{A.7}$$

with $\bar{\vartheta} = \frac{\pi^2}{4} + \pi - 5$.

Appendix B. Auxiliary estimation of the local variance

We estimate local variance with a non-parametric filter of length $2L+1$ (Fan and Yao, 2003) adapted for the presence of jumps by iterating in Z

$$\widehat{V}_t^Z = \frac{\sum\limits_{i=-L,\ i\neq -1,0,1}^{L} K\left(\frac{i}{L}\right) (\Delta_{t+i}X)^2 \, I_{\left\{(\Delta_{t+i}X)^2 \leq c_V^2 \cdot \widehat{V}_{t+i}^{Z-1}\right\}}}{\sum\limits_{i=-L,\ i\neq -1,0,1}^{L} K\left(\frac{i}{L}\right) I_{\left\{(\Delta_{t+i}X)^2 \leq c_V^2 \cdot \widehat{V}_{t+i}^{Z-1}\right\}}}, \quad Z = 1, 2, \ldots \tag{B.1}$$

with the starting value $\widehat{V}^0 = +\infty$, which corresponds to using all observations in the first step, and $c_V = 3$. At each iteration, large returns are eliminated by the condition $(\Delta_t X)^2 > c_V^2 \cdot \widehat{V}_t^{Z-1}$; each estimate of the variance is multiplied by c_V^2 to get the threshold for the following step. For estimation of the local variance at time t, we do not use the adjacent observations ($i \neq -1, 0, 1$). The iterations stop when no further large returns are removed. With high frequency data, this typically happens with $Z = 2, 3$ iterations. The bandwidth parameter L determines the number of adjacent returns included in the estimation of the local variance around point t. In our application, its choice is not crucial and we set $L = 25$. We use a Gaussian kernel $K(y) = \left(1/\sqrt{2\pi}\right) \exp\left(-y^2/2\right)$.

Alternatively, one may use the approach proposed by Lee and Mykland (2008) to detect and remove individual jumps and subsequently estimate the local variance with the remaining observations. Mancini and Renò (forthcoming) use instead a GARCH(1,1) process.

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