Bootstrap for log Wavelet Leaders Cumulant based Multifractal Analysis.
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ABSTRACT
Multifractal analysis, which mostly consists of estimating scaling exponents related to the power law behaviors of the moments of wavelet coefficients, is becoming a popular tool for empirical data analysis. However, little is known about the statistical performance of such procedures. Notably, despite their being of major practical importance, no confidence intervals are available. Here, we choose to replace wavelet coefficients with wavelet Leaders and to use a log-cumulant based multifractal analysis. We investigate the potential use of bootstrap to derive confidence intervals for wavelet Leaders log-cumulant multifractal estimation procedures. From numerical simulations involving well-known and well-controlled synthetic multifractal processes, we obtain two results of major importance for practical multifractal analysis: we demonstrate that the use of Leaders instead of wavelet coefficients brings significant improvements in log-cumulant based multifractal estimation, we show that accurate bootstrap designed confidence intervals can be obtained for a single finite length time series.

1. MOTIVATION
Scaling or Multifractal analysis [1, 2, 3] is becoming a standard analysis procedure commonly available in empirical data analysis toolboxes. Scaling, or scale invariance, is indeed a property that has been extensively observed in empirical data produced from numerous applications of very different nature. Multifractal analysis mostly consists of measuring scaling exponents, whose values are then commonly involved in various detection, identification or classification tasks. Despite becoming increasingly popular in data analysis, multifractal estimation procedures remain poorly studied. Questions, that may appear natural or simple, such as should one prefer increments or wavelet coefficients?, should one perform weighted or non-weighted regressions? or what are the typical sizes of the confidence intervals? still remain insufficiently addressed. However, for practical uses and purposes, elements of answers to such issues are crucial. Indeed, in many real life applications, the sizes of confidence intervals are as important as the values of the scaling exponents themselves, as no classification, discrimination or hypothesis testing are possible without them.

In the present contribution, we elaborate on multifractal analysis in combining together three key improvements: wavelet Leaders, log-cumulants and bootstrap.
First, it is now considered as classical and powerful to chose wavelet coefficients as the key multiresolution quantities multifractal analysis should be based on [2, 3]. Very recent findings reported in [4, 5] indicate that an accurate multifractal analysis should be based on wavelet Leaders rather than on wavelet coefficients. Indeed, the former enable to estimate exactly the entire multifractal spectrum and to analyze accurately processes containing oscillating singularities when the later do not. For further details, the reader is referred to [4, 5]. Wavelet Leaders are defined in Section 4.

Second, multifractal estimation procedures are commonly based on structure functions (i.e., power law behaviors of the moments of multiresolution quantities) as in Eqs. (1) or (2) below. However, it has been proposed to use instead the cumulants of the logarithm of the multiresolution quantities. This was originally introduced in the early nineties in [6] and largely developed in [7]. We follow here these promising developments. Cumulant based estimation procedures are described in Section 2.

Third, we investigate on potential benefits resulting from the use of non-parametric bootstrap for multifractal estimation. In bootstrapping, the distribution of an estimator is approximated through repeated resampling with replacement from the available data. The technique was introduced in the eighties [8] and has recently regained interest due to continuously growing computer abilities [9, 10, 11]. Bootstrapping has been used in the wavelet domain after the pioneering work reported in [12]. It has also been considered for the estimation of the Hurst parameter of self-similar processes [13, 14].

In the present work, we intend to explore the use of bootstrap in two respects: estimation procedure enhancement and confidence intervals derivation. Bootstrap procedures are detailed in Section 3.

Therefore, the aims of the present article are to contribute to the answers of the two following questions: Does the use of (log-cumulant) wavelet Leaders improve multifractal estimation procedures? Can bootstrap provide us with reliable confidence intervals? To address these questions, we first compare the statistical performance of estimation procedures based on wavelet coefficients and on wavelet Leaders. Second, we use a coverage procedure to compare confidence intervals obtained from a simple bootstrap approach. Results are derived by applying our procedures to a large number of realizations of synthetic scaling processes with a priori known and controlled multifractal properties (cf. Section 5). In Section 6, we show that the use of wavelet Leaders instead of wavelet coefficients brings substantial improvements in multifractal estimation performance. We also clearly demonstrate that the use of bootstrap procedures enables us to obtain highly reliable confidence intervals. We end up with a practical procedure that provides us with both accurate multifractal estimates and confidence intervals and that can actually be used for analyzing a single run of empirical data with finite observation duration.

2. MULTIFRACTAL, CUMULANTS AND WAVELETS
2.1. Definitions
Wavelet Coefficients. Let \( X(t), t \in [0, n) \) denote the process under analysis and \( n \) its observation duration. \( \psi_0(t) \) is a reference pattern with fast exponential decay, called the mother-wavelet. It is characterized by its number of vanishing moments, a strictly positive integer \( N_\psi \geq 1 \) defined as: \( \forall k = 0, 1, \ldots, N_\psi - 1, \int_R \psi(t)^k \psi_0(t)dt \equiv 0 \) and \( \int_R \psi(t)^{N_\psi} \psi_0(t)dt \neq 0 \). Let us further denote by \( \{ \psi_{j,k}(t) = 2^{-j/2} \psi_0(2^{-j}t - k), j \in \mathbb{Z}, k \in \mathbb{Z} \} \) templates of \( \psi_0 \) dilated to scales \( 2^j \), and translated to time positions \( 2^j k \). The
wavelet coefficients of $X$ are defined as $d_X(j,k) = \langle \psi_{j,k} | X \rangle$.

**Scaling and Multifractal.** A process $X$ is said to possess **scale invariance** or scaling properties if, for some statistical orders $q \in [q_-^*, q_+^*]$ (cf. [15]), the time averages of $|d_X(j,k)|^q$ taken at fixed scales display power law behaviors with respect to scales $a = 2^j$:

$$\langle |d_X(j,k)|^q \rangle = F_q |a|^{\zeta(q)},$$

(1)

over a wide range of scales $a \in [a_m, a_M]$, $a_M/a_m > 1$. The $\zeta(q)$ are referred to as the scaling exponents of $X$ and are closely related to its multifractal spectrum [5].

When $\zeta(q)$ is linear in $q$, i.e., $\zeta(q) = qH$, the process $X$ is said to be multifractal. This is, for instance, the case for finite variance self similar processes such as fractional Brownian motion. When $\zeta(q) \neq qH$, $X$ is said to be multifractal. This is clearly only a poor and operational definition of multifractality. However, for the purposes of this article it is sufficient. We refer the reader to, e.g., [1], for a thorough introduction to multifractal analysis.

**Cumulants.** [6, 7] For some processes, Eq. (1) is equivalent to

$$\mathbb{E}[|d_X(j,k)|^q] = F_q |a|^{2j \zeta(q)},$$

(2)

Using a second characteristic function type expansion, one can rewrite Eq. (2) as:

$$\ln \mathbb{E}[|d_X(j,k)|^q] = \sum_{p=1}^{\infty} C_p^q \frac{q^p}{p!} = \ln F_q + \zeta(q) \ln 2^j,$$

(3)

where $\zeta(q)$ stand for the cumulant of order $p$ $\geq 1$ of the random variable $\ln |d_X(j,k)|$. Combining Eqs (2) and (3) yields that the cumulants of $\ln |d_X(j,k)|$ have to be of the form:

$$\forall p \geq 1 : \quad C_p^q = c_p^0 + c_p \ln 2^j,$$

(4)

and therefore that

$$\ln \mathbb{E}[|d_X(j,k)|^q] = \sum_{p=1}^{\infty} C_p^q \frac{q^p}{p!} + \sum_{p=1}^{\infty} C_p^0 \frac{q^p}{p!} \ln 2^j,$$

(5)

where $c_p^0$ and $c_p$ do not depend on the scale $2^j$.

Thus, the measurements of the scaling exponents $\zeta(q)$ can be interestingly replaced by those of the log-cumulants $c_p$. This is mainly motivated by the fact that it emphasizes the difference between monofractal ($\forall p \geq 2 : c_p \equiv 0$) and multifractal processes [6, 7]. The next section describes estimation procedures for the $c_p$'s.

### 2.2. Estimation Procedures

Commonly, the scaling exponents $\zeta(q)$ are estimated by linear fits performed in $\ln \{ |d_X(j,k)|^q \}$ vs. $\ln |2^j|$ plots (see e.g., [2]). In the present work, we explore the alternative estimation of the equivalent quantities $c_p$.

**Cumulant estimations.** Given $n_j$ coefficients $d(j,k)$ and thus samples $Y_j(k) = \ln |d_X(j,k)|$, the asymptotically unbiased and consistent standard estimators (see e.g., [16]) are employed to obtain estimates $\hat{C}_p^q$ for the cumulants of $\ln |d_X(j,k)|$.

**Linear regressions.** From these $\hat{C}_p^q$’s, the $c_p$ can then be estimated by linear regression (cf. equation (4)),

$$\hat{c}_p = \log_2 e \sum_{j=1}^{j_1} w_j \hat{C}_p^q.$$

(6)

Theoretical performance. Since the $\hat{C}_p^q$'s are asymptotically unbiased and consistent, the $\hat{c}_p$ are asymptotically unbiased. As detailed in, e.g., [2], the $d_X(j,k)$ of scaling processes are weakly correlated. Hence, one can approximate the variance of $\hat{c}_p$ as:

$$\text{Var} \hat{c}_p \approx (\log_2 e)^2 \sum_{j=1}^{j_2} w_j^2 \text{Var} \hat{C}_p^q.$$

Thus, the $\hat{c}_p$'s are as well consistent.

**Weights.** The weights $w_j$ have to satisfy the constraints $\sum_{j=1}^{j_2} j w_j = 1$ and $\sum_{j=1}^{j_2} w_j = 0$ and can be expressed as $w_j = \frac{1}{b_j S_0 S_2 - S_j^2}$, with $S_i = \sum_{j=1}^{j_2} j^i / b_j, i = 0, 1, 2$. The positive numbers $b_j$ are freely selectable and reflect the confidence granted to each $\hat{C}_p^q$. We have chosen to come three cases, corresponding respectively to i) non-weighted regression, ii) the $d_X(j,k)$ can be idealized to independent random variables (cf. [2]), and iii) the confidence level is set proportional to the inverse of the estimated variances $\hat{C}_p^q$.

(8)

**3. BOOTSTRAP**

We use bootstrap generated nonparametric empirical distributions (see e.g. [8, 9, 10, 11]) for first estimating the variance $\hat{\sigma}_p^2(j)$ of $\hat{C}_p^q$, and second for constructing confidence intervals for $\hat{c}_p$. As the wavelet coefficients of scaling processes at a given scale are weakly correlated, we adopt a moving blocks bootstrap with overlapping blocks of length $L$. At each scale $a = 2^j$, the bootstrap resamples $D_j^{(1)}, \ldots, D_j^{(R)}$ are generated from the original sample $D_j = \{d_X(j,1), \ldots, d_X(j,n_j)\}$. Each resample $D_j^{(r)}$ represents an unsorted collection of $n_j$ sample points, drawn blockwise and with replacement from the original sample. These collections $D_j^{(r)}$ are used to compute $R$ bootstrap cumulant estimates $r = 1, \ldots, R, \hat{C}_p^{(r)}$. In turns, these $\hat{C}_p^{(r)}$’s are used for obtaining i) variance estimates for $\hat{C}_p^{(r)}$, and ii) $R$ bootstrap $\hat{c}_p$:

$$\hat{\sigma}_p^2(j) = \text{var} \hat{C}_p^{(r)};$$

(7)

$$\hat{c}_p^{(r)} = \log_2 e \sum_{j=1}^{j_2} w_j \hat{C}_p^{(r)}.$$
wavelet coefficients. Let us introduce the indexing \( \lambda_{j,k} = [2^j, (k+1)2^j) \) and the union \( 3\lambda_{j,k} = \lambda_{j-1} \cup \lambda_{j,k} \cup \lambda_{j,k+1} \). The wavelet Leaders \( L_X(j,k) \) are defined as

\[
L_X(j,k) = \sup_{X \in 3\lambda_{j,k}} |d_X|, \tag{10}
\]

where the supremum is taken on the discrete wavelet coefficients \( d_X(\cdot, \cdot) \) in the time neighborhood \( 3\lambda_{j,k} \) over all finer scales \( 2^j \prec 2^k \). All relations in Subsection 2.1, in particular Eqs. (1-5) can be rewritten replacing the \( d_X(j,k) \) by \( L_X(j,k) \). Thus, the estimation procedures detailed in Subsection 2.2 and in Section 3 can be rewritten, mutatis mutandis. For convenience, we introduce the superscript \( d \) and \( L \) to distinguish between estimates involving wavelet coefficients and Leaders, i.e., \( \hat{C}_p^{d} \), \( \hat{c}_p^{d} \), \( \hat{C}_p^{L} \) involve the \( d_X(j,k) \), while \( \hat{C}_p^{d} \), \( \hat{c}_p^{L} \), \( \text{CI}_p \) involve the \( L_X(j,k) \).

5. NUMERICAL SIMULATIONS

Monte Carlo Simulation. We evaluate the performance of the proposed estimation procedures by applying them to a large number \( N_{MC} \) of realizations of synthetic stochastic multifractal processes with known and controlled multifractal properties. From averages \( \langle \cdot \rangle \) over Monte Carlo realizations, we compute the standard deviations \( \hat{\sigma}_p = \sqrt{\langle c_p^2(i) \rangle - \langle c_p(i) \rangle^2} \), the biases \( \hat{\beta}_p = \langle c_p(i) \rangle - \hat{c}_p(i) \) and mean-square errors \( \text{MSE}_p = \hat{\sigma}_p^2 + \hat{\beta}_p^2 \) of the proposed estimators.

To evaluate the reliability of the confidence intervals obtained from bootstrap, we investigate the coverages produced by re-centered confidence intervals,

\[
\text{CI}_{p,R}(i) = \text{CI}_{p}(i) - \hat{\beta}_p, \quad i = 1, \ldots, N_{MC}
\]

i.e., confidence intervals that are corrected by the Monte Carlo estimates of the bias of the estimators. This allows us to determine the quality of the confidence intervals independently of the influence of a possible bias of the estimators. The empirical coverages of re-centered confidence intervals are then calculated as:

\[
\text{Cov}_{\text{emp}} = \langle \epsilon(c_p, \text{CI}_{p,R}(i)) \rangle.
\]

Here, \( \epsilon(c_p, \text{CI}_{p,R}(i)) = 1 \) if \( c_p \in \text{CI}_{p,R}(i) \) and 0 otherwise: i.e., the empirical coverages \( \text{Cov}_{\text{emp}} \) equals the percentage of MC realizations for which the true \( c_p \) fall within the corresponding re-centered confidence intervals.

Scaling Processes. We use two well known scaling processes, Fractional Brownian motion (FBM) and Multifractal random walk (MRW), chosen because they provide us with simple yet representative examples of Gaussian monotofractal processes and non Gaussian multifractal processes respectively. FBM is defined as the only Gaussian exactly self-similar process with stationary increments. Its full definition as well as that of self-similarity can be found in e.g., [17]. The statistical properties of FBM are entirely determined by the parameter \( H \). FBM possesses scaling properties as in Eq. (2), with \( \zeta(q) = qH \), for \( q \in (-1, \infty) \). Thus, \( c_1 = H \) and \( c_p \equiv 0 \) for all \( p \geq 2 \). MRW has been introduced in [18] as a simple multifractal (hence non Gaussian) process with stationary increments: \( X(k) = \sum_{k=1}^{n} G_H(k)^{\omega(k)} \), where \( G_H(k) \) consists of the increments of FBM with parameter \( H \). The process \( \omega \) is independent of \( G_H \), Gaussian, with non trivial covariance: \( \text{cov}(\omega(k_1), \omega(k_2)) = \lambda \ln \left( \frac{k_1}{k_2} \right)^{1/2} \) when \( k_1, k_2 \) are comparable, \( |k_1 - k_2| < L \) and 0 otherwise. MRW has interesting scaling properties as in Eqs. (1) or (2) for \( q \in [-\sqrt{2/\lambda}, \sqrt{2/\lambda}] \) (cf. [15]), with \( \zeta(q) = (H + \lambda)q - \lambda^2 q^2/2 \), hence \( c_1 = H + \lambda \), \( c_2 = -\lambda^2 \) and \( c_p \equiv 0 \) for all \( p \geq 3 \). One sees that the departure from a linear behavior in \( q \) is fully controlled by \( \lambda \) (or \( c_2 \)).

Simulation Setup. The results presented here are obtained using Daubechies wavelets with \( N_\psi = 3 \). Parameters were set to \( N_{MC} = 1000, n = 2^{15}, R = 200, L = 6, H = c_1 = 0.8 \) for FBM and \( (H, \lambda) = (0.72, \sqrt{0.08}) \), i.e. \( c_1 = 0.8 \) and \( c_2 = -0.08 \), for MRW.

6. RESULTS

6.1. Statistical Performance. Tables 1 and 2 compare the biases and MSEs (respectively) of \( \hat{c}_p^d \) and \( \hat{c}_p^L \) for \( p = 1, \ldots, 5 \), obtained for 1000 realizations of FBM and MRW.

Bias. Table 1 shows that, while \( \hat{c}_p^d \) and \( \hat{c}_p^L \) have comparable biases, for \( p \geq 2 \), \( \hat{c}_p^L \) systematically exhibits smaller biases. Note that this discrepancy increases with \( p \) and that for \( p = 4, 5 \), the difference counts at least 3 orders of magnitude! Clearly, \( \hat{c}_p^d \) become useless in practise for \( p = 4, 5 \), whereas \( \hat{c}_p^L \) continue to give estimates of high accuracy, similar to those produced for \( p = 1 \). Also, \( \hat{c}_p^d \) has a bias smaller than that of \( \hat{c}_p^L \) for (monofractal) FBM, but a larger bias for (multifractal) MRW. Together, these arguments clearly indicate that when the deviations from linearity of \( \zeta(q) \) are of interest, Leaders must be preferred to wavelet coefficients. From the weight choice point of view, biases are equivalent.

Mean Square Error. Standard deviations are an order of magnitude larger than biases, so that they mostly contribute to MSEs. Hence MSEs only are reported. Table 2 shows that the MSEs of \( \hat{c}_p^L \) are systematically much smaller than those of the \( \hat{c}_p^d \) for both processes, all weights and all orders. Again, this difference grows with \( p \) (roughly as \( 10^{p-1} \)), showing that estimates for the cumulant of order 3 can no longer be obtained from wavelet coefficients and require the use of Leaders. For the weight choice issue, we note that the MSEs of \( \hat{c}_p^d \) based on \( w_{1, j} \) and \( w_{2, j} \) are of comparable order of magnitude in all cases and much smaller than those obtained with \( w_{0, j} \). Four conclusions can be drawn: i) linear regressions for the cumulants must be weighted, ii) the independence assumption of the \( d_X(j,k) \) underlying the choice \( w_{1, j} \) remains valid for Leaders, iii) bootstrap does a good job in estimating the variances of the \( \hat{c}_p^L(s) \), iv) as the choice \( w_{2, j} \) involves a much higher computational cost for results equivalent to those obtained with the choice \( w_{1, j} \) this latter is preferred. Fig. 1 displays histograms of \( \hat{c}_p^d \) and \( \hat{c}_p^L \) for \( p = 1, 2, 3 \) (weights \( w_{1, j} \)) obtained from 1000 realizations of MRW. Whereas the empirical distributions of \( \hat{c}_p^L \) and \( \hat{c}_p^d \) have similar shape and center, those of \( \hat{c}_p^L \) have smaller spread, in particular for \( p = 2 \). Again, this suggests the use of \( \hat{c}_p^L \) rather than \( \hat{c}_p^d \).

Conclusion. These results lead us to conclude that weighted wavelet Leader based estimators produce the most accurate estimates for log-cumulants. In particular, highly relevant estimates of log-cumulants of higher order are obtained. Hence, we recommend the use of \( \hat{c}_p^L \) with weights \( w_{1, j} \) for practical log-cumulant based multifractal analysis.

Block Bootstrap. From the definition of the Leaders \( L_X(j,k) \) in Eq. (10), it is clear that they may display complicated intra- and inter-scale (i.e., joint time and scale) correlation structures. Therefore, one may expect poor results for leader-based simple bootstrap estimates. Our results show that this is not the case. Moreover, in Fig. 2, histograms for \( \hat{c}_p^L \), based on 200 realizations of MRW (cf. below), and for 200 corresponding bootstrap estimates \( \hat{c}_p^L \), based on a single realization, are shown. The closeness of Monte Carlo and bootstrap empirical distributions for \( \hat{c}_p^L \) provides us with a clear
From these analyzes, we have designed a multifractal analysis. Second, the choice of wavelet Leaders instead of wavelet coefficients brings substantial improvements resulting from the use of pivotal statistics (see e.g., Table 2). MSE estimates for $c_p$ and $d_p$ for $p = 1 - 5$. Best results are marked in bold.

### Table 2. MSE ($\times 10^3$) of estimators $\hat{c}_p$ for FBM (left) and MRW (right) and $p = 1 - 5$. Best results are marked in bold.

<table>
<thead>
<tr>
<th>Estimator</th>
<th>FBM</th>
<th>MRW</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\hat{c}_{1}^p$</td>
<td>32.8</td>
<td>19.8</td>
</tr>
<tr>
<td>$\hat{c}_{2}^p$</td>
<td>77.6</td>
<td>6.5</td>
</tr>
<tr>
<td>$\hat{c}_{3}^p$</td>
<td>340.0</td>
<td>3.1</td>
</tr>
<tr>
<td>$\hat{c}_{4}^p$</td>
<td>1822.4</td>
<td>2.0</td>
</tr>
<tr>
<td>$\hat{c}_{5}^p$</td>
<td>9622.5</td>
<td>2.6</td>
</tr>
</tbody>
</table>

This significantly improves already available practical multifractal estimation procedures as for most applications error bars are as important as estimates themselves. We see this as a major result of the present contribution. Fig. 3 illustrates this procedure at work and shows logscale diagrams (for $p = 1, 2, 3$): $\hat{C}_p^1$ (or $\hat{C}_p^2$) as a function of $j$ for a single realization of MRW, together with corresponding regression lines and $\pm 1.96\hat{\sigma}_p(j)$ error bar estimates obtained from bootstrap. Whereas all $\hat{C}_p^j$ display a highly linear behaviour over a large range of scales $j$, even for large $p$, a zone of linearity is more difficult to find for $\hat{C}_2^d$ and $\hat{C}_3^d$, suggesting estimates of poorer quality in these cases. Clearly, the selection of the regression range of octaves is a key practical issue that adds an extra complication to the analyses reported here. Again, one sees that for Leader based multifractal, selecting the regression range should be easier.

### Table 3. Empirical coverage of re-centered 95% confidence interval for log-cumulant estimates from wavelet coefficients (top) and wavelet Leaders (bottom) for FBM and MRW and $p = 1 - 5$. Results closest to target coverage are marked in bold.

<table>
<thead>
<tr>
<th>Estimate</th>
<th>FBM</th>
<th>MRW</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\hat{C}_{1,LR}^1$</td>
<td>85.7</td>
<td>92.1</td>
</tr>
<tr>
<td>$\hat{C}_{2,LR}^1$</td>
<td>79.4</td>
<td>92.3</td>
</tr>
<tr>
<td>$\hat{C}_{3,LR}^1$</td>
<td>72.3</td>
<td>91.2</td>
</tr>
<tr>
<td>$\hat{C}_{4,LR}^1$</td>
<td>64.3</td>
<td>87.8</td>
</tr>
<tr>
<td>$\hat{C}_{5,LR}^1$</td>
<td>85.7</td>
<td>87.8</td>
</tr>
</tbody>
</table>

6.2. Confidence Intervals.

Table 3 summarizes the empirical coverges of the re-centered confidence intervals (9) for the estimators $\hat{c}_p$ and $\hat{d}_p$ of order $p = 1 - 5$. The targeted coverage is 95%.

6.3. Practical procedure and regression range.

From these analyzes, we have designed a MATLAB routine that implement wavelet coefficient and wavelet Leader based log-cumulant multifractal analysis together with bootstrap based confidence intervals. Therefore, it enables us to obtain from a single observed times series with finite length, both estimates for the $c_p$s and error bars.
knowledge, this is the first time that a nonparametric confidence interval estimation procedure, with excellent performance, is obtained for multifractal analysis. Wavelet Leaders and bootstrap confidence intervals together lead to the design of a multifractal analysis procedure (available in MATLAB upon request) of primary interest for the exploration of empirical data with possibly multifractal properties. A key feature of the results obtained in this work lies in the fact that they hold for both Gaussian monofractal and non-Gaussian multifractal processes. This is very promising as it opens the track for the design of hypothesis tests aiming at discriminating between mono- and multi-fractal processes, and between different multi-fractal processes, two major practical issues. These ideas are currently under investigation.

8. REFERENCES


