A BAYESIAN APPROACH FOR THE JOINT ESTIMATION OF THE MULTIFRACTALITY PARAMETER AND INTEGRAL SCALE BASED ON THE WHITTLE APPROXIMATION

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ABSTRACT

Multifractal analysis is a powerful standard signal processing tool. Multifractal models are essentially characterized by two parameters, the so-called multifractality parameter \( \alpha_2 \) and the integral scale \( A \) (the time scale beyond which multifractal properties vanish). Yet, most applications concentrate on estimating \( \alpha_2 \) while estimating \( A \) is mostly overlooked, despite of \( A \) potentially conveying important information. Joint estimation of \( \alpha_2 \) and \( A \) is challenging due to the statistical nature of multifractal processes (strong dependence, non-Gaussian), and has barely been considered. The present contribution addresses these limitations and proposes a Bayesian procedure for the joint estimation of \( (\alpha_2, A) \). Its originality resides, first, in the construction of a generic multivariate model for the statistics of wavelet leaders for multifractal multiplicative cascade processes, and second, in the use of a suitable Whittle approximation for the likelihood associated with the model. The resulting model enables Bayesian estimators for \( (\alpha_2, A) \) to be computed also for large sample size. Performance is assessed numerically for synthetic multifractal processes and illustrated for wind-tunnel turbulence data. The proposed procedure significantly improves estimation of \( \alpha_2 \) and yields, for the first time, reliable estimates for \( A \).

Index Terms— Multifractal Analysis, Integral Scale, Wavelet Leaders, Bayesian Estimation, Whittle Likelihood

1. INTRODUCTION

Context. Scale invariance provides practitioners with a powerful concept for real-world data analysis. It has been used in a large variety of applications of very different natures, e.g., biomedical applications (body rhythms [1], infra slow brain activity [2]), hydrodynamic turbulence [3], geophysics [4], finance [5], Internet traffic [6], to name but a few. Scale invariance implies that the temporal dynamics of data are not driven by any particular scale that could play a privileged role in analysis. Instead, a large continuum of time scales equally contributes to the temporal dynamics. From a practical perspective, this translates into power law behaviors of the sample moments of well chosen multi-scale quantities \( T_X(a,t) \) (quantities that depend jointly on time \( t \) and scale \( a \), such as wavelet coefficients), i.e.

\[
S(q,j) \equiv \frac{1}{n_j} \sum_{k=1}^{n_j} |T_X(a,k)|^q \sim a^{\zeta(q)}, \quad a_m \leq a \leq a_M. \tag{1}
\]

The goal of scale invariance is hence to estimate the scaling exponents \( \zeta(q) \) that characterize the mechanisms relating scales.

Multifractal analysis consists of a specific instance of scale invariance analysis (cf. e.g., [7]). It notably enables discrimination between two classes of processes commonly used to model scale invariance: self-similar processes, characterized by \( \zeta(q) \equiv qH \) and an underlying additive structure [8], the fractional Brownian motion (fBM) being the celebrated representative member [9]; multifractal multiplicative cascades (hereafter denoted MMC), characterized by a strictly concave \( \zeta(q) \) and an underlying multiplicative structure [3].

Deciding which model is preferred by data is of utmost importance in applications as it may significantly modify the understanding and interpretations of the underlying mechanisms producing the data. The scaling exponents \( \zeta(q) \) of MMC can be expanded as a function of \( q, \zeta(q) = c_1q + c_2q^2/2 + \ldots \), with strictly negative \( c_2 < 0 \), while \( \zeta(q) = qH \) and \( c_2 \equiv 0 \) for self-similar processes. The discrimination between self-similar processes and MMC can thus be recast into testing \( c_2 \equiv 0 \) versus \( c_2 < 0 \) (cf. e.g., [10, 11]) and \( c_2 \) is therefore often referred to as the multifractality or intermittency parameter. The second fundamental difference between self-similar processes and MMC is that the power law relation in (1) theoretically holds for all scales \( a > 0 \) for self-similar processes, while holds only within a range of scales that is necessary bounded from above, \( 0 < a \leq A \) for MMC. This upper bound is commonly referred to as the integral scale [3, 12].

While most research concentrate on the estimation of the sole parameter \( c_2 \), the integral scale \( A \) has mostly been overlooked. Yet, it conveys fundamental information since it subtly reintroduces a notion of typical (decorrelation) time scale within the scale invariance framework (cf. [13] for a review). This may provide insights into the mechanisms underlying data production. The estimation of the integral scale \( A \) constitutes the core topic of the present contribution.

Related work: Joint estimation of \( \alpha_2 \) and \( A \). It is now well understood that the estimation of \( \alpha_2 \) (and the detection of deviations of \( \zeta(q) \) from a linear behavior in \( q \)) should be based on recently proposed refined multi-scale quantities termed wavelet leaders, cf., e.g., [7, 10]. The estimation of \( \alpha_2 \) essentially relies on linear regressions across scales, motivated by (1) and variations (cf., (2) in Section 2). To improve estimation (notably for small sample size), a generalized moment approach has been proposed, relying strongly on fully parametric models [14] and hence is of limited applicability to real-world data. Alternatively, the use of Bayesian models was proposed but remained mostly restricted to the estimation of the self-similarity parameter for Gaussian processes [15–17]. Only recently, new Bayesian models were proposed for the estimation of \( c_2 \), either by considering specific properties of certain processes [18] or by exploiting generic properties of wavelet leaders [19] that are valid for large classes of MMC.

In contrast, the estimation of the integral scale \( A \) has received
limited attention. In certain applications, the order of magnitude of the integral scale can be approximately inferred from a priori available physical parameters (typically flow size, average flow speed, . . . e.g., for climatology and rainfall analysis [20, 21] and for hydrodynamic turbulence [12, 22–24]), while this is not possible in most other applications. At the methodological level, an extension of the generalized moment approach to the estimation of the integral scale was proposed in [25] (see also [24]), yet with limited use in applications due to the requirement of fully parametric models.

Contributions. The present work aims at developing a Bayesian model for the joint estimation of the multifractality parameter $c_2$ and the integral scale $A$. The procedure generalizes [19], which proposed the first wavelet-leader based Bayesian estimator for the sole parameter $c_2$, yet assumes $A \approx n$ and is intractable for sample size larger than $n \sim 10^5$. The main contributions of the present work lie, first, in the generalization of the statistical model proposed in [19] to $A \leq n$, enabling the formulation of a joint estimator for $(c_2, A)$, and second, the use of a suitable Whittle likelihood in the Bayesian procedure [26–29], enabling their use for large sample sizes.

We first propose a semi-parametric model for the statistics of the wavelet leaders of MMC, motivated by the asymptotic covariance of the logarithm of multiscale quantities associated with these processes (c.f. [3]). The model is generically valid for this class of processes, for all values of $A$. It imposes minimal model assumptions on data (essentially, (2) below) and involves few parameters (effectively, $c_2$ and $A$, cf. Section 3.1). From this model, a Bayesian estimation procedure for $(c_2, A)$ is developed by assigning an appropriate prior distribution to the parameters, reflecting the constraints inherent to the multifractal model. To explore the resulting posterior distribution and generate samples used to approximate the Bayesian estimators, a suitable MCMC random-walk Metropolis-Hastings sampling scheme is devised (cf. Section 3).

The direct evaluation of the likelihood in the MCMC scheme would require, at each iteration, the inversion of a dense matrix of essentially the order of the sample size $n$, which is prohibitive both numerically and computationally for large $n$. To overcome this difficulty, we propose to approximate the exact likelihood induced by the multifractal model by an appropriate Whittle likelihood (cf. Section 4). The resulting algorithm for the joint estimation of $c_2$ and $A$ is effective for both small and large sample sizes. Its performance is assessed by means of Monte Carlo simulations, showing the clear benefits of the Bayesian estimator over linear regressions for the estimation of $c_2$, and its effectiveness for the reliable estimation of the integral scale $A$ (cf. Section 5). Finally, we illustrate the potential of the proposed Bayesian procedure for the analysis of a high-quality real-world data set of wind-tunnel turbulence.

## 2. MULTIFRACTAL ANALYSIS

### Discrete wavelet transform.

A mother wavelet $\psi_0(t)$ is a reference pattern with narrow supports in both time and frequency domains. It is characterized by its number of vanishing moments $N_0 \geq 1$ ($\forall k = 0, 1, \ldots, N_0 - 1, \int_{R} t^{k} \psi_0(t) dt = 0$ and $\int_{R} t^{N_0} \psi_0(t) dt \neq 0$). Also, it is chosen such that the collection $\{\psi_{j,k}(t) \equiv 2^{-j/2} \psi_0(2^{-j} t - k), j \in \mathbb{N}, k \in \mathbb{N}\}$ forms a basis of $L^2(R)$. The discrete wavelet transform (DWT) coefficients of $X$ are defined as $d_X(j, k) = \langle X, \psi_{j,k} \rangle$, cf., e.g., [30] for further details.

**Wavelet leaders.** Let denote $\lambda_j(k) = [k2^j, (k + 1)2^j)$ the dyadic interval of size $2^j$ and $3\lambda_j(k)$ the union of $\lambda_j(k)$ with its 2 neighbors. The wavelet leaders are defined as the largest wavelet coefficient in the neighborhood $3\lambda_j(k)$ over all finer scales, $\ell(j, k) := \sup_{\lambda_j \subset 3\lambda_j(k)} |d_X(\lambda_j)|$ [7, 31].

### Multifractal formalism.

The wavelet leader scaling function is defined as $\zeta(q) = \lim_{j \to \infty} \inf_{k \to \infty} \left[ \ln|S(j, k)| / \ln 2 \right]$ where $S(j, k) = 2^j \sum_{l} |\ell(l, k)|$ are the empirical moments of order $q$ of the wavelet leaders of $X$ at scale $j$. The function $\zeta(q)$ is intimately tied to the multifractal spectrum $D(h)$ via a Legendre transform, $D(h) \leq \zeta(h) = \inf_{q \in \mathbb{R}} [1 + qh - \zeta(q)]$. It can be shown that this inequality is strict for large classes of multifractal processes. The Legendre spectrum $\zeta(h)$ is thus practically often confounded with the theoretical spectrum $D(h)$, see [10, 31].

### Log-cumulant expansion.

It is often advantageous in applications to work with the leading order coefficients of the polynomial expansion $\zeta(q) = \sum_{m \geq 1} \epsilon_m q^m / m!$ of the scaling function. This expansion directly translates to $D(h)$, see [31]. In particular, the first log-cumulant $c_1$ is identical to the position of the mode of $D(h)$ (i.e., the average smoothness), and the second log-cumulant $c_2$ is directly related to its width (i.e., the degree of regularity fluctuations). The seminal work [32] shows that the $c_{2m}$ are directly related to the cumulants of order $m$ of the wavelet leaders, $\mu_{2m}[\ln \ell(j, k)] = c_{2m} + c_m \ln 2^j$ and specifically,

$$C_2(j) = \mu_{2}[\ln \ell(j, k)] = c_0^2 + c_2 \ln 2^j.$$ (2)

The parameter $c_2$ can thus be estimated by linear regression of the sample variance (denoted $\mu_2$) of $\ln \ell(j, k)$ against scale $j \in [j_1, j_2]$

$$\hat{c}_2 = \frac{1}{\ln 2} \sum_{j=j_1}^{j_2} w_j \mu_2[\ln \ell(j, k)]$$ (3)

where $w_j$ are suitable regression weights, see [6, 10] for details.

### 3. BAYESIAN ESTIMATION

#### 3.1. Model for the multivariate statistics of log-wavelet leaders

Let $l(j, k) = \ln \ell(j, k)$ denote the log-wavelet leaders. We propose a model for the multivariate statistics of $l(j, k)$ of MMC that generalizes the model in [19] to MMC with integral scale $A \leq n$.

**Marginal distributions.** It has been shown in [19] that the marginal distribution of $l(j, k)$ of MMC can be approximated by a Gaussian distribution. This is illustrated in Fig. 1 (top row) for different integral scale values for the process described in Section 5.

**Covariance.** The numerical studies reported in [19] suggest that the covariance of the logarithm of wavelet leaders of MMC at fixed scale $j$, $\Sigma_j(\Delta k) := \text{Cov}[l(j, k), l(j, k + \Delta k)]$, is characterized by a logarithmic decay controlled by the parameter $c_2$

$$\Sigma_j(\Delta k) \approx \eta + c_2(\ln \Delta k + \ln 2^j) =: \gamma^{(j)}(\Delta r; c_2, \eta)$$ (4)

for $3 \leq \Delta k \leq n_j \frac{1}{n}$, where $n_j \approx [n/2^j]$ denotes the number of wavelet leaders at scale $j$ (see also [33] for results obtained for log-wavelet coefficients of 1D random wavelet cascades). Second, the theoretical variance of the log-wavelet leaders is given by $C_2(j) = C_2(j; c_2, c_2^0)$ defined in (2). Finally, we propose to model the short-term covariance as a logarithmic decay from $C_2(j; c_2, c_2^0)$ at $\Delta k = 0$ to $\gamma^{(j)}(\Delta r; c_2, \eta)$ at $\Delta k = 3$.

$$\gamma^{(j)}(\Delta k; c_2, c_2^0) = C_2(j; c_2, c_2^0) + \left( \ln(\Delta k + 1) / \ln 4 \right) \left( \gamma^{(j)}(3; c_2, \eta) - C_2(j; c_2^0) \right).$$ (5)

By combining (3), (4) and (5), we obtain the following model for the covariance of log-wavelet leaders, parametrized by $\theta = [c_2, c_2^0, \eta]^T$

$$\gamma_j(\Delta k; \theta) = \begin{cases} C_2(j; c_2, c_2^0) & \Delta k = 0 \\ \gamma^{(j)}(\Delta k; c_2, c_2^0, \eta) & 0 < \Delta k \leq 3 \\ \max[0, \gamma^{(j)}(\Delta k; c_2, \eta)] & 3 < \Delta k \leq n_j. \end{cases}$$ (6)
The posterior distribution and Bayesian estimators.

3.3. Gibbs sampler

We define the admissible set $\{c_{2}, c_{2}^\star\}$ of the integral scale $A$ with entries by the parametric covariance (4), respectively.

Prior distribution. The parameter vector $\theta = \{c_{2}, c_{2}^\star, \eta\}^T$ must be chosen such that the variances of $l(j, k)$ are positive, i.e., $C_{2}(j, j) > 0$. We define the admissible set $T = (T^+ \cup T^-) \cap T^m$, where $T^- = \{\theta \in \mathbb{R}^3 | c_{2} < 0$ and $c_{2} + c_{2} j \neq 2 > 0\}$, $T^+ = \{\theta \in \mathbb{R}^3 | c_{2} > 0$ and $c_{2} + c_{2} j \neq 2 > 0\}$, $T^m = \{\theta \in \mathbb{R}^3 | c_{2}^m < c_{2} < c_{2}^m, |\eta| < \eta^m\}$ and $c_{2}^m, c_{2}^m, \eta^m$ are the largest admissible values for $c_{2}, c_{2}^\star$ and $\eta$. Without additional prior information regarding $\theta$, a uniform prior distribution on the set $T$ is assigned to $\theta$, i.e., $P(\theta) = U(\mathbb{R}) \propto 1_{T}(\theta)$, where $1_{T}$ is the indicator function of $T$.

Posterior distribution and Bayesian estimators. The posterior distribution of $\theta$ is obtained from the Bayes rule

$$f(\theta | \mathcal{L}) \propto f(\mathcal{L} | \theta) P(\theta)$$

and can be used to define the maximum a posteriori (MAP) and minimum mean squared error (MMSE) estimators in (9) below.

3.3. Gibbs sampler

The following Gibbs sampler enables the generation of samples $\{\theta^{(i)}\}_{i=1}^{N_{MC}}$ that are asymptotically distributed according to the posterior distribution (8). The Gibbs sampling strategy consists of successively sampling according to the conditional distributions associated with $f(\theta | \mathcal{L})$. To sample according to the conditional distributions, a Metropolis-within-Gibbs procedure is used, defined by random walks with Gaussian instrumental distributions. More precisely, at iteration $#t$, the three following moves are considered.

Sampling according to $f(c_{2}^2 | c_{2}^2, \eta^{(t)}, \mathcal{L})$. A candidate $c_{2}^2$ is generated according to the proposal distribution $q_{c_{2}^2}(c_{2}^2 | \mathcal{L}) = \mathcal{N}(c_{2}^2, \sigma_{c_{2}^2}^2)$. It is accepted (c_{2}^2 = c_{2}^2) or rejected (c_{2}^2 = c_{2}^2(t)) according to the Metropolis-Hastings ratio $r_{c_{2}^2}$.

Sampling according to $f(\eta^{(t)} | c_{2}^2, \mathcal{L})$. A candidate $\eta^{(t)}$ is generated according to the proposal distribution $q_{\eta}(\eta^{(t)} | \mathcal{L}) = \mathcal{N}(\eta^{(t)}, \sigma_{\eta}^2)$ and accepted (c_{2}^2 = c_{2}^2) or rejected (c_{2}^2 = c_{2}^2(t)) according to the Metropolis-Hastings ratio $r_{\eta}$.

The Metropolis-Hastings acceptance ratios are defined by $r_{c_{2}^2} = \frac{f(\theta^{(t+1)} | \mathcal{L}) q_{c_{2}^2}(c_{2}^2 | \mathcal{L})}{f(\theta^{(t)} | \mathcal{L}) q_{c_{2}^2}(c_{2}^2 | \mathcal{L})}$, and similar for $r_{\eta}$.

4. WHITTE APPROXIMATION

The Gibbs sampler requires inversion of the dense $n_{j} \times n_{j}$ matrices $\Gamma_{i}(\theta)$ in each sampling run. For large sample size, this is practically intractable both for practical (computation time) and numerical (growing condition number of $\Gamma_{i}(\theta)$) reasons. To handle large sample sizes, we replace the exact likelihood (7) with the approximate Whittle likelihood [26, 27]. Up to an additive constant, the Whittle approximation for the negative log-likelihood is given by

$$-\ln L(\theta, \mathcal{L}) \approx \mathcal{W}(\theta, \mathcal{L}) := \frac{1}{2} \sum_{\omega} \ln |\vartheta_{\omega}|(\omega) + \frac{\Pi_{\omega}(\omega)}{n_{j} \vartheta_{\omega}(\omega)}$$

where $\Pi_{\omega}(\omega) = |\omega|^{2} \sum_{k=1}^{n_{j}} (l(j, k) \exp(\omega k))^{2}$ is the periodogram of $\{l(j, k)\}_{k \in F_{j}}$ and $|\vartheta_{\omega}(\omega)| = \sum_{k=1}^{n_{j}} \gamma_{j}(\Delta k; \theta) \exp(\omega k)$ is the Fourier transform of the covariance function (6).

The Whittle likelihood that replaces (7) in (8) is, up to a multiplicative constant, given by

$$f(l \mid \mathcal{L}) \approx f_{W}(l \mid \theta, \mathcal{L}) := \exp \left(-\sum_{j=1}^{J} W(I_{j}, \theta) \right).$$

5. RESULTS

We quantify the estimation performance of the proposed procedure by applying it to a large number $N_{b}$ of independent realizations of a synthetic multifractal process, the multifractal random walk (MRW), with different prescribed integral scale values. MRW is chosen here as a prominent member of the class of multifractal multiplicative cascade based processes. MRW is a non Gaussian process with stationary increments. Its multifractal properties mimic those of the
considered here. In particular, the bias is found to be significantly
yields consistent estimates of
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correlation coefficient \( \rho(\hat{c}_2, \hat{J}) \) of the estimates of \( (c_2, J) \). For large values
of \( J, c_2 \) and \( \hat{J} \) show relatively strong correlation. For smaller
values of \( J, \rho(\hat{c}_2, \hat{J}) \) decreases since the variances \( C_2(j) \) (con-
trolled by \( c_2 \) only) become more dominant over the covariance term
\( \gamma_j^{(1)}(\Delta k; c_2, \eta) \) (jointly controlled by \( c_2 \) and \( J \)) in (6) (cf. Fig. 1).

**Application to Turbulence Data.** We illustrate the proposed pro-
cedure for a large wind-tunnel turbulence data set consisting of high
sampled simultaneously longitudinal Eulerian velocity signals, measured
with hot-wire anemometry techniques. The dataset, made available to us
by Y. Gagne [23], consists of \( R = 24 \) independent runs of \( n = 2^{25} \)
samples each, with (Taylor scale based) Reynolds number \( R_\lambda \approx
2000 \), integral scale \( A = 2^{13} \), and Taylor scale \( 2^4 \). Estimation
parameters are set as \( [10] \) (i.e., \( |j_1|, j_2| = [6, 10] \) and \( N_\psi = 3 \).

Results are reported in Fig. 4 (c2 (left) and \( J = \log_2 A \) (right))
for each individual run and indicate that the proposed procedure
yields highly consistent estimates for the different runs, both for \( c_2 \)
and \( J \). The averages (standard deviations) of the estimates are found
to be \(-0.016 \) \((0.001) \) for \( c_2 \) and \( 11.78 \) \((0.31) \) for \( J \), respectively.
In view of the above reported results for synthetic data, estimates for \( J \) are
within one standard deviation and hence in agreement with the
value \( J = 13 \) inferred based on Taylor scale in [23].

**6. CONCLUSION**

To the best of our knowledge, this paper studied the first Bayesian
procedure for the joint estimation of the multifractality parameter \( c_2 \)
and (log-) integral scale \( J \) that is operational and can be applied
to real-world data. It relies on a novel generic semiparametric model
for the statistics of wavelet leaders of multifractal cascade based multifractal processes. A Gibbs sampler is designed to
produce samples according to the joint posterior distribution of the
multifractal parameter vector, incorporating the multifractal model
constraints, which are used to approximate the Bayesian estimators.
Computational efficiency of the procedure and applicability to large
sample sizes are made possible by using the approximate Whittle
likelihood in the sampler. The procedure enables reliable estimation
of the integral scale \( J \), previously barely achieved. It signficantly
improves estimation for the multifractality parameter \( c_2 \) over stan-
dard linear regression based estimators, reducing standard deviations
and RMSE values to 25% of those of linear fit based estimation (at
the price of increased computational cost). The procedure is cur-
cently being used in the study of 24 hours long heart rate variability
time series. Future work will include the extension of the proposed
procedure to 2D images and the development of a relevant statistical
framework for the multifractal analysis of multivariate time series.

MATLAB codes implementing the proposed procedure, written
by the authors, are publicly available at http://www.irit.fr/~Herwig.Wendt/.