

Oscillations of Empirical Distribution Functions under Dependence*

Wei Biao Wu

University of Chicago

Abstract: We obtain an almost sure bound for oscillation rates of empirical distribution functions for stationary causal processes. For short-range dependent processes, the oscillation rate is shown to be optimal in the sense that it is as sharp as the one obtained under independence. The dependence conditions are expressed in terms of physical dependence measures which are directly related to the data-generating mechanism of the underlying processes and thus are easy to work with.

1. Introduction

Let $\varepsilon'_0, \varepsilon_i, i \in \mathbb{Z}$, be independent and identically distributed (iid) random variables on the same probability space $(\Omega, \mathcal{A}, \mathbb{P})$. For $k \in \mathbb{Z}$ let

$$X_k = g(\dots, \varepsilon_{k-1}, \varepsilon_k), \quad (1)$$

where g is a measurable function such that X_k is a well-defined random variable. Then $\{X_k\}_{k \in \mathbb{Z}}$ forms a stationary sequence. The framework (1) is very general. See [12, 15, 19, 23] among others. The process (X_k) is causal or non-anticipative in the sense that X_k does not depend on future innovations $\varepsilon_{k+1}, \varepsilon_{k+2}, \dots$. Causality is a reasonable assumption in practice. The Wiener-Rosenblatt conjecture says that, for every stationary and ergodic process X_k , there exists a measurable function g and iid innovations ε_i such that the distributional equality $(X_k)_{k \in \mathbb{Z}} =_{\mathcal{D}} (g(\dots, \varepsilon_{k-1}, \varepsilon_k))_{k \in \mathbb{Z}}$ holds; see [13, 20]. For an overview of the Wiener-Rosenblatt conjecture see [9].

Let F be the cumulative distribution function of X_k . Assume throughout the paper that F has a square integrable density f with square integrable derivative f' . In this paper we are interested in the oscillatory behavior of empirical distribution function

$$F_n(x) = \frac{1}{n} \sum_{i=1}^n \mathbf{1}_{X_i \leq x}, \quad x \in \mathbb{R}. \quad (2)$$

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In particular, we shall obtain an almost sure bound for the modulus of continuity for the function $G_n(x) = \sqrt{n}[F_n(x) - F(x)]$:

$$\Delta_n(b) = \sup_{|x-y|\leq b} |G_n(x) - G_n(y)|, \tag{3}$$

where $b = b_n$ is a sequence of positive numbers satisfying

$$b_n \rightarrow 0 \text{ and } nb_n \rightarrow \infty. \tag{4}$$

Under the assumption that X_i are iid, there exists a huge literature on the asymptotic behavior of $\Delta_n(\cdot)$; see Chapter 14 in [14] and the references cited therein. A powerful tool to deal with the empirical distribution F_n is strong approximation [1]. In comparison, the behavior of $\Delta_n(\cdot)$ has been much less studied under dependence. In this paper we shall implement the new dependence measures proposed in [23] and obtain an almost sure bound for Δ_n . For a recent account of empirical processes for dependent random variables see the monograph edited by Dehling et al [4].

The rest of the paper is structured as follows. Main results on $\Delta_n(b)$ are presented in Section 2 and proved in Section 3. Section 4 contains comparisons with results obtained under independence. Some open problems are also posed in Section 4.

2. Main results

We first introduce some notation. For a random variable Z write $Z \in \mathcal{L}^p$ ($p > 0$) if $\|Z\|_p := (\mathbb{E}|Z|^p)^{1/p} < \infty$. Write $\|\cdot\| = \|\cdot\|_2$. Let for $k \in \mathbb{Z}$,

$$\xi_k = (\dots, \varepsilon_{k-1}, \varepsilon_k)$$

and for $k \geq 0$ let ξ_k^* be a coupled process of ξ_k with ε_0 replaced by ε'_0 , i.e.,

$$\xi_k^* = (\xi_{-1}, \varepsilon'_0, \varepsilon_1, \dots, \varepsilon_k).$$

We shall write $X_k^* = g(\xi_k^*)$. Let $k \geq 1$ and define the conditional cumulative distribution function $F_k(\cdot|\xi_0)$ by

$$F_k(x|\xi_0) = \mathbb{P}(X_k \leq x|\xi_0). \tag{5}$$

Note that (X_i, ξ_i) is stationary. Then for almost every $\xi \in \dots \times \mathbb{R} \times \mathbb{R}$ with respect to \mathbb{P} we have

$$F_k(x|\xi_0 = \xi) = \mathbb{P}(X_k \leq x|\xi_0 = \xi) = \mathbb{P}(X_{k+i} \leq x|\xi_i = \xi) \quad (6)$$

for all $i \in \mathbb{Z}$. In other words, $F_k(\cdot|\xi)$ is the cumulative distribution function of the random variable $g(\xi, \varepsilon_{i+1}, \dots, \varepsilon_{i+k})$. Assume that for all $k \geq 1$ and almost every $\xi \in \dots \times \mathbb{R} \times \mathbb{R}$ with respect to \mathbb{P} that $F_k(x|\xi_0 = \xi)$ has a derivative $f_k(x|\xi_0 = \xi)$, which is the conditional density of X_k at x given $\xi_0 = \xi$. By (6), for any $i \in \mathbb{Z}$, $f_k(x|\xi_i)$ is the conditional density of X_{k+i} at x given ξ_i . Let the conditional characteristic function

$$\varphi_k(\theta|\xi_0 = \xi) = \mathbb{E}(e^{\sqrt{-1}\theta X_k}|\xi_0 = \xi) = \int_{\mathbb{R}} e^{\sqrt{-1}\theta t} f_k(t|\xi_0 = \xi) dt, \quad (7)$$

where $\sqrt{-1}$ is the imaginary unit. Our dependence condition is expressed in terms of the \mathcal{L}^2 norm $\|\varphi_k(\theta|\xi_0) - \varphi_k(\theta|\xi_0^*)\|$.

THEOREM 1. *Assume that $b_n \rightarrow 0$, $\log n = O(nb_n)$ and that there exists a positive constant c_0 for which*

$$\sup_x f_1(x|\xi_0) \leq c_0 \quad (8)$$

holds almost surely. Further assume that

$$\sum_{k=1}^{\infty} \left[\int_{\mathbb{R}} (1 + \theta^2) \|\varphi_k(\theta|\xi_0) - \varphi_k(\theta|\xi_0^*)\|^2 d\theta \right]^{1/2} < \infty. \quad (9)$$

Let

$$\iota(n) = (\log n)^{1/2} \log \log n. \quad (10)$$

Then

$$\Delta_n(b_n) = O_{\text{a.s.}}(\sqrt{b_n \log n}) + o_{\text{a.s.}}[b_n \iota(n)]. \quad (11)$$

Roughly speaking, (9) is a short-range dependence condition. Recall that $f_k(x|\xi_0)$ is the conditional (predictive) density of X_k at x given ξ_0 and $\varphi_k(\theta|\xi_0)$ is the conditional characteristic function. So $\varphi_k(\theta|\xi_0) - \varphi_k(\theta|\xi_0^*)$ measures the degree of dependence of $\varphi_k(\cdot|\xi_0)$

on ε_0 . Hence the summand in (9) quantifies a distance between the conditional distributions $[X_k|\xi_0]$ and $[X_k^*|\xi_0^*]$ and (9) means that the cumulative contribution of ε_0 in predicting future values is finite.

For the two terms in the bound (11), the first one $O_{\text{a.s.}}(\sqrt{b_n \log n})$ has the same order of magnitude as the one that one can obtain under independence. See Chapter 14 in [14] and Section 4. The second term $o_{\text{a.s.}}[b_n \nu(n)]$ is due to the dependence of the process (X_k) . Clearly, if $b_n^{1/2}(\log \log n) = o(1)$, then the first term dominates the bound in (11). The latter condition holds under mild conditions on b_n , for example, if $b_n = O(n^{-\eta})$ for some $\eta > 0$.

Let $k \geq 1$. Observe that $\varphi_k(\theta|\xi_0) = \mathbb{E}[\varphi_1(\theta|\xi_{k-1})|\xi_0]$ and

$$\mathbb{E}[\varphi_1(\theta|\xi_{k-1})|\xi_{-1}] = \mathbb{E}[\varphi_1(\theta|\xi_{k-1}^*)|\xi_{-1}] = \mathbb{E}[\varphi_1(\theta|\xi_{k-1}^*)|\xi_0]. \quad (12)$$

To see (12), write $h(\xi_{k-1}) = \varphi_1(\theta|\xi_{k-1})$. Note that $\varepsilon_i, \varepsilon'_0, i \in \mathbb{Z}$, are iid and $k-1 \geq 0$. Then we have $\mathbb{E}[h(\xi_{k-1})|\xi_{-1}] = \mathbb{E}[h(\xi_{k-1}^*)|\xi_{-1}]$ since ξ_{k-1}^* is a coupled version of ξ_{k-1} with ε_0 replaced by ε'_0 . On the other hand, we have $\mathbb{E}[h(\xi_{k-1}^*)|\xi_{-1}] = \mathbb{E}[h(\xi_{k-1}^*)|\xi_0]$ since ε_0 is independent of ξ_{k-1}^* . So (12) follows. Define the projection operator \mathcal{P}_k by

$$\mathcal{P}_k Z = \mathbb{E}(Z|\xi_k) - \mathbb{E}(Z|\xi_{k-1}), \quad Z \in \mathcal{L}^1.$$

By the Jensen and the triangle inequalities,

$$\begin{aligned} \|\varphi_k(\theta|\xi_0) - \varphi_k(\theta|\xi_0^*)\| &\leq \|\varphi_k(\theta|\xi_0) - \mathbb{E}[\varphi_k(\theta|\xi_0)|\xi_{-1}]\| + \|\mathbb{E}[\varphi_k(\theta|\xi_0)|\xi_{-1}] - \varphi_k(\theta|\xi_0^*)\| \\ &= 2\|\mathcal{P}_0 \varphi_1(\theta|\xi_{k-1})\| \leq 2\|\varphi_1(\theta|\xi_{k-1}) - \varphi_1(\theta|\xi_{k-1}^*)\|. \end{aligned}$$

Then a sufficient condition for (9) is

$$\sum_{k=0}^{\infty} \left[\int_{\mathbb{R}} (1 + \theta^2) \|\varphi_1(\theta|\xi_k) - \varphi_1(\theta|\xi_k^*)\|^2 d\theta \right]^{1/2} < \infty. \quad (13)$$

In certain applications it is easier to work with (13). In Theorem 2 below we show that (13) holds for processes (X_k) with the structure

$$X_k = \varepsilon_k + Y_{k-1}, \quad (14)$$

where Y_{k-1} is $\xi_{k-1} = (\dots, \varepsilon_{k-2}, \varepsilon_{k-1})$ measurable. It is also a large class. The widely used linear process $X_k = \sum_{i=0}^{\infty} a_i \varepsilon_{k-i}$ is of the form (14). Nonlinear processes of the form

$$X_k = m(X_{k-1}) + \varepsilon_k \tag{15}$$

also fall within the framework of (14) if (15) has a stationary solution. A prominent example of (15) is the threshold autoregressive model [19]

$$X_k = a \max(X_{k-1}, 0) + b \min(X_{k-1}, 0) + \varepsilon_k,$$

where a and b are real parameters. For processes of the form (14), condition (9) can be simplified. Let φ be the characteristic function of ε_1 .

THEOREM 2. *Let $0 < \alpha \leq 2$. Assume (14),*

$$\int_{\mathbb{R}} |\varphi(t)|^2 (1 + t^2) |t|^\alpha dt < \infty \tag{16}$$

and

$$\sum_{k=0}^{\infty} \|X_k - X_k^*\|_\alpha^{\alpha/2} < \infty \tag{17}$$

Then (9) is satisfied.

Condition (16) does not appear to be overly restrictive. It is satisfied if $|\varphi(t)| = O(|t|^{-\eta})$ as $|t| \rightarrow \infty$, where $\eta > (3 + \alpha)/2$. It is also satisfied for symmetric- α -stable distributions, an important class of distributions with heavy tails. Let ε_k have standard symmetric α stable distribution with index $0 < \iota \leq 2$. Then its characteristic function $\varphi(t) = \exp(-|t|^\iota)$ and (16) trivially holds.

We now discuss Condition (17). Recall $X_k^* = g(\xi_k^*)$. Note that X_k^* and X_k are identically distributed and X_k^* is a coupled version of X_k with ε_0 replaced by ε_0' . If we view (1) as a physical system with $\xi_k = (\dots, \varepsilon_{k-1}, \varepsilon_k)$ being the input, g being a filter or transform and X_i being the output, then the quantity $\|X_k - X_k^*\|_\alpha$ measures the degree of dependence of $g(\dots, \varepsilon_{k-1}, \varepsilon_k)$ on ε_0 . In [23] it is called the *physical* or *functional dependence measure*. With this input/output viewpoint, the condition (17) means that the cumulative impact of ε_0 is finite, and hence suggesting short-range dependence. In many applications it is

easily verifiable since it is directly related to the data-generating mechanism and since the calculation of $\|X_k - X_k^*\|_\alpha$ is generally easy [23]. In the special case of linear process $X_k = \sum_{j=0}^{\infty} a_j \varepsilon_{k-j}$ with $\varepsilon_k \in \mathcal{L}^\alpha$ and $\alpha = 2$, then $\|X_k - X_k^*\|_\alpha = |a_k| \|\varepsilon_0 - \varepsilon_0'\|_\alpha$ and (17) is reduced to $\sum_{k=0}^{\infty} |a_k| < \infty$, which is a classical condition for linear processes to be short-range dependent. It is well-known that, if the latter condition is barely violated, then one enters the territory of long-range dependence. Consequently both the normalization and the bound in (11) will be different; see [8, 21].

For the nonlinear time series (15), assume that $\varepsilon_k \in \mathcal{L}^\alpha$ and $\rho = \sup_x |m'(x)| < 1$. Then (15) has a stationary distribution and $\|X_k - X_k^*\|_\alpha = O(\rho^k)$ (see [24]). Hence (17) holds.

3. Proofs

LEMMA 1. *Let H be a differential function on \mathbb{R} . Then for any $\lambda > 0$,*

$$\sup_{x \in \mathbb{R}} H^2(x) \leq \lambda \int_{\mathbb{R}} H^2(x) dx + \lambda^{-1} \int_{\mathbb{R}} [H'(x)]^2 dx. \quad (18)$$

Proof. By the arithmetic mean geometric inequality inequality, for all $x, y \in \mathbb{R}$,

$$H^2(x) \leq H^2(y) + \left| \int_x^y 2H(t)H'(t)dt \right| \leq H^2(y) + \lambda \int_{\mathbb{R}} H^2(x) dx + \lambda^{-1} \int_{\mathbb{R}} [H'(x)]^2 dx. \quad (19)$$

If $\inf_{y \in \mathbb{R}} |H(y)| > 0$, then $\int_{\mathbb{R}} H^2(x) dx = \infty$ and (18) holds. If $\inf_{y \in \mathbb{R}} |H(y)| = 0$, let $(y_n)_{n \in \mathbb{N}}$ be a sequence such that $H(y_n) \rightarrow 0$. So (19) entails (18). \diamond

Lemma 1 is a special case of the Kolmogorov-type inequalities [18]. The result in the latter paper asserts that $\sup_{x \in \mathbb{R}} H^4(x) \leq \int_{\mathbb{R}} H^2(x) dx \int_{\mathbb{R}} H'(x)^2 dx$. For the sake of completeness, we decide to state Lemma 1 with a simple proof here.

Recall that $F_k(\cdot|\xi_0)$ is the conditional distribution function of X_k given ξ_0 (cf (5) and (6)). Introduce the conditional empirical distribution function

$$F_n^*(x) = \frac{1}{n} \sum_{i=1}^n \mathbb{E}(\mathbf{1}_{X_i \leq x} | \xi_{i-1}) = \frac{1}{n} \sum_{i=1}^n F_1(x | \xi_{i-1}).$$

Write

$$G_n(x) = G_n^\circ(x) + G_n^*(x), \quad (20)$$

where

$$G_n^\diamond(x) = \sqrt{n}[F_n(x) - F_n^*(x)] \text{ and } G_n^*(x) = \sqrt{n}[F_n^*(x) - F(x)]. \quad (21)$$

Then

$$\sqrt{n}G_n^\diamond(x) = \sum_{i=1}^n d_i(x) \quad (22)$$

is a martingale with respect to the filtration $\sigma(\xi_n)$ and the increments $d_i(x) = \mathbf{1}_{X_i \leq x} - \mathbb{E}(\mathbf{1}_{X_i \leq x} | \xi_{i-1})$ are stationary, ergodic and bounded. On the other hand, if the conditional density $f_1(\cdot | \xi_i)$ exists, then G_n^* is differentiable. The latter differentiability property is quite useful.

LEMMA 2. Recall (10) for $\iota(n)$. Assume (9). Let $g_n^*(x) = dG_n^*(x)/dx$. Then

$$\sup_x |g_n^*(x)| = o_{\text{a.s.}}[\iota(n)]. \quad (23)$$

Proof. Let $k \geq 1$. Recall that $f_1(x | \xi_{k-1})$ is the one-step-ahead conditional density of X_k at x given ξ_{k-1} . By (7), we have

$$\mathcal{P}_0 \varphi_1(t | \xi_{k-1}) = \int_{\mathbb{R}} e^{\sqrt{-1}xt} \mathcal{P}_0 f_1(x | \xi_{k-1}) dt.$$

By Parseval's identity, we have

$$\int_{\mathbb{R}} |\mathcal{P}_0 \varphi_1(t | \xi_{k-1})|^2 dt = \frac{1}{2\pi} \int_{\mathbb{R}} |\mathcal{P}_0 f_1(x | \xi_{k-1})|^2 dx$$

and

$$\int_{\mathbb{R}} |\mathcal{P}_0 \varphi_1(t | \xi_{k-1})|^2 t^2 dt = \frac{1}{2\pi} \int_{\mathbb{R}} |\mathcal{P}_0 f_1'(x | \xi_{k-1})|^2 dx.$$

Let

$$\alpha_k = \int_{\mathbb{R}} \|\mathcal{P}_0 f_1(x | \xi_{k-1})\|^2 dx$$

and

$$\beta_k = \int_{\mathbb{R}} \|\mathcal{P}_0 f_1'(x | \xi_{k-1})\|^2 dx.$$

By (12) and Jensen's inequality, $\|\mathcal{P}_0\varphi_1(\theta|\xi_{k-1})\| \leq \|\varphi_k(\theta|\xi_0) - \varphi_k(\theta|\xi_0^*)\|$. So (9) implies that

$$\sum_{k=1}^{\infty} \sqrt{\alpha_k + \beta_k} < \infty. \quad (24)$$

Let $\Lambda = \sum_{k=1}^{\infty} \sqrt{\alpha_k}$ and for $k \geq 0$

$$H_k(x) = \sum_{i=1}^k [f_1(x|\xi_{i-1}) - f(x)].$$

Then $H_k(x) = \sqrt{k}g_k(x)$. Note that for fixed $l \in \mathbb{N}$, $\mathcal{P}_{i-l}f_1(x|\xi_{i-1})$, $i = 1, 2, \dots$, are stationary martingale differences with respect to the filtration $\sigma(\xi_{i-l})$. By the Cauchy-Schwarz inequality and Doob's maximal inequality, since $H_k(x) = \sum_{l=1}^{\infty} \sum_{i=1}^k \mathcal{P}_{i-l}f_1(x|\xi_{i-1})$,

$$\begin{aligned} \mathbb{E} \int_{\mathbb{R}} \max_{k \leq n} H_k^2(x) dx &\leq \mathbb{E} \int_{\mathbb{R}} \sum_{l=1}^{\infty} \frac{\max_{k \leq n} |\sum_{i=1}^n \mathcal{P}_{i-l}f_1(x|\xi_{i-1})|^2}{\sqrt{\alpha_l}} \Lambda dx \\ &\leq \int_{\mathbb{R}} \sum_{l=1}^{\infty} \frac{4n \|\mathcal{P}_0 f_1(x|\xi_{l-1})\|^2}{\sqrt{\alpha_l}} \Lambda dx = 4n\Lambda^2 = O(n). \end{aligned}$$

Similarly, since $\sum_{k=1}^{\infty} \sqrt{\beta_k} < \infty$,

$$\mathbb{E} \int_{\mathbb{R}} \max_{k \leq n} |H'_k(x)|^2 dx = O(n).$$

By Lemma 1 with $\lambda = 1$, we have

$$\begin{aligned} \sum_{d=1}^{\infty} \frac{\mathbb{E}[\max_{k \leq 2^d} \sup_{x \in \mathbb{R}} |H_k(x)|^2]}{2^d \iota^2(2^d)} &\leq \sum_{d=1}^{\infty} \frac{\mathbb{E}[\max_{k \leq 2^d} \int_{\mathbb{R}} |H_k(x)|^2 + |H'_k(x)|^2 dx]}{2^d \iota^2(2^d)} \\ &\leq \sum_{d=1}^{\infty} \frac{\mathbb{E} \int_{\mathbb{R}} \max_{k \leq 2^d} |H_k(x)|^2 + \max_{k \leq 2^d} |H'_k(x)|^2 dx}{2^d \iota^2(2^d)} \\ &= \sum_{d=1}^{\infty} \frac{O(2^d)}{2^d \iota^2(2^d)} < \infty. \end{aligned}$$

By the Borel-Cantelli lemma, as $d \rightarrow \infty$, $\sup_x \max_{k \leq 2^d} |H_k(x)| = o_{\text{a.s.}}[2^{d/2} \iota(2^d)]$. For any $n \geq 2$ there is a $d \in \mathbb{N}$ such that $2^{d-1} < n \leq 2^d$. Note that $\max_{k \leq n} |H_k(x)| \leq \max_{k \leq 2^d} |H_k(x)|$ and $\iota(n)$ is slowly varying. So (23) follows. \diamond

LEMMA 3. Assume $\log n = O(nb_n)$ and $X_k \in \mathcal{L}^\alpha$ for some $\alpha > 0$. Then for any $\tau > 2$, there exists $C = C_\tau > 0$ such that

$$\mathbb{P} \left[\sup_{|x-y| \leq b_n} |G_n^\diamond(x) - G_n^\diamond(y)| > C \sqrt{b_n \log n} \right] = O(n^{-\tau}). \quad (25)$$

Proof. We can adopt the argument of Lemma 5 in [22]. Let $x_0 = n^{(3+\tau)/\alpha}$. Then for any $C > 0$, by Markov's inequality,

$$\begin{aligned} \mathbb{P}[n\{F_n(-x_0) + 1 - F_n(x_0)\} > c\sqrt{b_n \log n}] &\leq \frac{n\mathbb{E}\{F_n(-x_0) + 1 - F_n(x_0)\}}{C\sqrt{b_n \log n}} \\ &\leq \frac{nx_0^{-\alpha}\mathbb{E}(|X_0|^\alpha)}{C\sqrt{b_n \log n}} = O(n^{-\tau}). \end{aligned} \quad (26)$$

It is easily seen that the preceding inequality also holds if F_n is replaced by F_n^* .

Recall (22) for $G_n^\diamond(x)$ and $d_i(x) = \mathbf{1}_{X_i \leq x} - \mathbb{E}(\mathbf{1}_{X_i \leq x} | \xi_{i-1})$. Let $x \leq y \leq x + b_n$. By (8),

$$\sum_{i=1}^n \mathbb{E}[(d_i(y) - d_i(x))^2 | \xi_{i-1}] \leq \sum_{i=1}^n \mathbb{E}(\mathbf{1}_{x \leq X_i \leq y} | \xi_{i-1}) \leq nb_n c_0.$$

By Freedman's inequality in [6], if $|x - y| \leq b_n$, we have

$$\mathbb{P} \left[\sqrt{n} |G_n^\diamond(x) - G_n^\diamond(y)| > C \sqrt{nb_n \log n} \right] \leq 2 \exp \left[\frac{-C^2 nb_n \log n}{C \sqrt{nb_n \log n} + nb_n c_0} \right].$$

Let $\Theta_n = \{-x_0 + k/n^3 : k = 0, \dots, \lfloor 2x_0 n^3 \rfloor\}$. Since $\log n = O(nb_n)$, it is easily seen that there exists a $C = C_\tau$ such that

$$\begin{aligned} &\mathbb{P} \left[\sup_{x, y \in \Theta_n, |x-y| \leq b_n} \sqrt{n} |G_n^\diamond(x) - G_n^\diamond(y)| > C \sqrt{nb_n \log n} \right] \\ &= O(x_0^2 n^2) \exp \left[\frac{-C^2 nb_n \log n}{C \sqrt{nb_n \log n} + nb_n c_0} \right] = O(n^{-\tau}). \end{aligned} \quad (27)$$

For every $x \in [-x_0, x_0]$, there exists a $\theta \in \Theta_n$ such that $\theta < x \leq \theta + 1/n^3$. So (27) implies that

$$\mathbb{P} \left[\sup_{|x| \leq x_0, |y| \leq x_0, |x-y| \leq b_n} \sqrt{n} |G_n^\diamond(x) - G_n^\diamond(y)| > (C+1) \sqrt{nb_n \log n} \right] = O(n^{-\tau}) \quad (28)$$

in view of the monotonicity of $F_n(\cdot)$ and the fact that, if $\theta \leq \phi \leq \theta + 1/n^3$,

$$\sum_{i=1}^n \mathbb{E}(\mathbf{1}_{\theta \leq X_i \leq \phi} | \xi_{i-1}) \leq |\phi - \theta| n c_0 = c_0/n^2 = o(\sqrt{nb_n \log n}).$$

Combining (26) and (28), we have (25). \diamond

Proof of Theorem 1. By (20), it easily follows from Lemmas 2 and 3. \diamond

Proof of Theorem 2. By Lemma 1 and Parseval's identity, (16) implies

$$\sup_x f^2(x) \leq \int_{\mathbb{R}} \{f^2(x) + [f'(x)]^2\} dx = \frac{1}{2\pi} \int_{\mathbb{R}} |\varphi(t)|^2 (1 + t^2) dt < \infty.$$

Since ε_k and Y_{k-1} are independent,

$$\mathbb{E}[e^{\sqrt{-1}tX_k} | \xi_{k-1}] = e^{\sqrt{-1}tY_{k-1}} \varphi(t).$$

Note that

$$\begin{aligned} \|e^{\sqrt{-1}tY_{k-1}} \varphi(t) - e^{\sqrt{-1}tY_{k-1}^*} \varphi(t)\|^2 &= |\varphi(t)|^2 \|e^{\sqrt{-1}tY_{k-1}} - e^{\sqrt{-1}tY_{k-1}^*}\|^2 \\ &\leq 4|\varphi(t)|^2 \|\min(1, |tY_{k-1} - tY_{k-1}^*|)\|^2 \\ &\leq 4|\varphi(t)|^2 \mathbb{E}(|tY_{k-1} - tY_{k-1}^*|^\alpha). \end{aligned}$$

Hence (9) follows from (17). \diamond

4. Conclusion and open problems

Let X_i be iid standard uniform random variables. Stute [16] obtained the following interesting result. Assume that $b_n \rightarrow 0$ is a sequence of positive numbers such that

$$\log n = o(nb_n) \text{ and } \log \log n = o(\log b_n^{-1}). \quad (29)$$

Then the convergence result holds:

$$\lim_{n \rightarrow \infty} \frac{\Delta_n(b_n)}{\sqrt{b_n \log b_n^{-1}}} = \sqrt{2} \text{ almost surely.} \quad (30)$$

If there exists $\eta > 0$ such that $b_n + (nb_n)^{-1} = O(n^{-\eta})$, then the bound in (11) becomes $\sqrt{b_n \log n}$, which has the same order of magnitude as $\sqrt{b_n \log b_n^{-1}}$, the bound asserted by (30). It is unclear whether there exists an almost sure limit for $\Delta_n(b_n)/\sqrt{b_n \log b_n^{-1}}$ if the dependence among observations is allowed. Mason et al [11] considered almost sure limit for $\Delta_n(b_n)/\sqrt{b_n \log b_n^{-1}}$ when (29) is violated. Deheuvels and Mason [3] (see also [2])

proved functional laws of the iterated logarithm for the increments of empirical processes. Local empirical processes in high dimensions have been studied in [5, 7, 17]. It is an open problem whether similar results hold for stationary causal processes. We expect that our decomposition (21) will be useful in establishing comparable results.

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References

- [1] CSÖRGŐ, M. and RÉVÉSZ, P., (1981). *Strong approximations in probability and statistics*, Academic Press, New York.
- [2] DEHEUVELS P. (1997). Strong laws for local quantile processes. *Ann. Probab.* **25** 2007–2054.
- [3] DEHEUVELS P. and MASON, D. M. (1992). Functional laws of the iterated logarithm for the increments of empirical and quantile processes. *Ann. Probab.* **20** 1248–1287.
- [4] DEHLING, H., T. MIKOSCH and M. SØRENSEN (eds), (2002). *Empirical Process Techniques for Dependent Data*, Boston: Birkhäuser.
- [5] EINMAHL, U. and MASON, D. M. (2000). An empirical process approach to the uniform consistency of kernel-type function estimators. *J. Theoret. Probab.* **13** 1–37.
- [6] FREEDMAN, D. A. (1975). On tail probabilities for martingales. *Ann. Probab.* **3** 100–118.
- [7] GINÉ, E. and GUILLOU, A. (2002). Rates of strong uniform consistency for multivariate kernel density estimators. *Ann. Inst. H. Poincaré Probab. Statist.* **38** 907–921.
- [8] HO, H. C. and HSING, T. (1996). On the asymptotic expansion of the empirical process of long-memory moving averages. *Ann. Statist.* **24** 992–1024.
- [9] KALLIANPUR, G. (1981). Some ramifications of Wiener's ideas on nonlinear prediction. In *Norbert Wiener, Collected Works with Commentaries* (MIT Press, MA), pp. 402–424.

- [10] MASON, D. M. (2004). A uniform functional law of the logarithm for the local empirical process *Ann. Probab.* **32** 1391–1418
- [11] MASON, D. M., SHORACK G. R. AND J. A. WELLNER (1983). Strong limit theorems for oscillation moduli of the uniform empirical process. *Z. Wahrsch. verw. Gebiete* **65** 83–97.
- [12] PRIESTLEY, M. B. (1988). *Nonlinear and Nonstationary Time Series Analysis*. Academic Press, London.
- [13] ROSENBLATT, M. (1959). Stationary processes as shifts of functions of independent random variables. *J. Math. Mech.* **8** 665–681.
- [14] SHORACK, G. R. AND WELLNER, J. A. (1986). *Empirical processes with applications to statistics*. John Wiley & Sons, New York.
- [15] STINE, R. A. (1997) Nonlinear time series. In *Encyclopedia of Statistical Sciences*, Wiley, Updated vol. 1, Edited by S. Kotz, C. B. Read and D. L. Banks., pp. 430–437
- [16] STUTE, W. (1982). The oscillation behavior of empirical processes. *Ann. Probab.* **10** 86–107.
- [17] STUTE, W. (1984). The oscillation behavior of empirical processes: the multivariate case. *Ann. Probab.* **12** 361–379.
- [18] ТАЙКОВ, Л. В. (1968). Inequalities of Kolmogorov type and best formulas for numerical differentiation. (Russian) *Mat. Zametki* **4** 233–238.
- [19] TONG, H. (1990). *Nonlinear time series. A dynamical system approach*. Oxford University Press.
- [20] WIENER, N. (1958). *Nonlinear Problems in Random Theory*. MIT Press, Cambridge, MA
- [21] WU, W. B. (2003). Empirical processes of long-memory sequences. **9** *Bernoulli* 809–831
- [22] WU, W. B. (2005a). On the Bahadur representation of sample quantiles for stationary sequences. *Ann. Statist.* **33** 1934–1963
- [23] WU, W. B. (2005b). Nonlinear system theory: another look at dependence. *Proc. Natl Acad. Sci. USA* **102** 14150–14154.
- [24] WU, W. B. and SHAO, X. (2004). Limit theorems for iterated random functions. *J. Appl. Probab.* **41** 425–436.

*Department of Statistics
The University of Chicago
5734 S. University Avenue
Chicago, IL 60637
E-MAIL: wbwu@galton.uchicago.edu*