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ON ADAPTIVE WAVELET ESTIMATION OF A CLASS OF WEIGHTED DENSITIES

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Abstract

We investigate the estimation of a weighted density taking the form $g = w(F)f$, where f denotes an unknown density, F the associated distribution function and w is a known (non-negative) weight. Such a class encompasses many examples, including those arising in order statistics or when g is related to the maximum or the minimum of N (random or fixed) independent and identically distributed (*i.i.d.*) random variables. We here construct a new adaptive non-parametric estimator for g based on a plug-in approach and the wavelets methodology. For a wide class of models, we prove that it attains fast rates of convergence under the \mathbb{L}_p risk with $p \geq 1$ (not only for $p = 2$ corresponding to the mean integrated squared error) over Besov balls. The theoretical findings are illustrated through several simulations.

Key words and phrases: Weighted density, density estimation, plug-in approach, wavelets, block thresholding, reliability, series system, parallel system.

1 Introduction

1.1 Problem statement

Let $(\Omega, \mathcal{A}, \mathbb{P})$ be a probability space, X be a real random variable with unknown density f and Y be a random variable having the unknown weighted density

$$g(x) = w(F(x))f(x), \quad x \in \mathbb{R}, \quad (1.1)$$

where w denotes a known (of course non-negative) weight and F denotes the distribution function of f . The goal we pursue here is to estimate g from a random number N of *i.i.d.* sample X_1, \dots, X_N of X .

Such an estimation problem arises in many situations, typically when g is related to the maximum¹ of N *i.i.d.* random variables, where N is a discrete random number in \mathbb{N}^* which is independent of the X_i 's. Application fields cover hydrology, meteorology, reliability, investment, management science, insurance business, etc.. For example, when the X_i are non-negative, the random variable $Y = \max(X_1, \dots, X_N)$ (or $Y = \min(X_1, \dots, X_N)$) arises naturally in reliability theory as the lifetime of a parallel (series) system with a random number N of identical components with lifetimes X_1, \dots, X_N .

To make things clearer to the reader, we next give some illustrative examples.

1.2 Motivating examples

Example 1.1 (Order statistics). Let X_1, \dots, X_m be *i.i.d.* random variables with absolutely continuous distribution function F and probability density function (pdf) f . Let $X_{(1)} \leq \dots \leq X_{(m)}$ denote the corresponding order statistics. Then, the pdf $g_{X_{(j)}}$ of the j -th order statistic is

$$g_{X_{(j)}}(x) = \frac{m!}{(j-1)!(m-j)!} (F(x))^{j-1} (1-F(x))^{m-j} f(x), \quad x \in \mathbb{R}.$$

Thus, $X_{(m)}$, for example, is the random variable representing the largest observation of a sample of n and corresponds to the sample maximum and the density $g_{X_{(m)}}$ of $X_{(m)} = \max(X_1, \dots, X_m)$ is given by

$$g_{X_{(m)}}(x) = m (F(x))^{m-1} f(x), \quad x \in \mathbb{R}.$$

¹Since $\min(X_1, \dots, X_N) = -\max(-X_1, \dots, -X_N)$ the results can be easily reformulated for the sample minimum.

The aim is to estimate $g_{X(j)}$ from a m *i.i.d.* sample X_1, \dots, X_m of X .

Example 1.2 (Maximum of a random number N of *i.i.d.* random variables). Let X be a random variable with density f , $\{X_n\}_{n \in \mathbb{N}^*}$ be a sequence of *i.i.d.* random variables with density f and N be a discrete random variable taking values in \mathbb{N}^* with a known probability mass function. Then the density of $Y = \max(X_1, \dots, X_N)$ is

$$g(x) = w(F(x))f(x), \quad x \in \mathbb{R}, \quad (1.2)$$

where

$$w(u) = \sum_{k=1}^{\infty} k u^{k-1} \mathbb{P}(N = k), \quad u \in [0, 1].$$

The goal is again to estimate g from an n *i.i.d.* sample X_1, \dots, X_n of X .

Example 1.3 (Pile-up model). Let us now present the “pile-up model”. Let $\{Y_n\}_{n \in \mathbb{N}^*}$ be a sequence of *i.i.d.* random variables with density g , N be a discrete random variable in \mathbb{N}^* as in the previous example, and let $X = \min(Y_1, \dots, Y_N)$ with density f . Then the density of Y_1 is

$$g(x) = w(F(x))f(x), \quad x \in \mathbb{R},$$

where

$$w(u) = \frac{1}{M'(M^{-1}(1-u))}, \quad u \in [0, 1],$$

$M(u) = \mathbb{E}(u^N)$ and $M'(u) = \mathbb{E}(Nu^{N-1})$. We are seeking an estimate of g from a n *i.i.d.* sample X_1, \dots, X_n of X .

1.3 Previous work

Some distributional properties of the maximum and minimum of random variables have been extensively studied in the literature (see, e.g., [Raghunandanan and Patil \(1972\)](#), [Shaked \(1975\)](#) and [Shaked and Wong \(1997\)](#)). In addition, the literature on order statistics contains a huge work about the maximum. In the context of extreme value theory, various statistical properties and (real data) applications can be found in [Adamidis et al. \(2005\)](#), [Louzada et al. \(2012\)](#) and the references therein.

The estimation of the density function of the maximum of two independent random variables has been considered by [Chen and Hsu \(2004\)](#) via kernel methods. The Pile-up model has been considered by [Comte and Rebafka \(2010\)](#) via model selection methods.

1.4 Contributions and relation to prior work

In this paper, we develop a new *non-linear* adaptive estimator for g in model (1.1) based on a plug-in method, wavelets and the block thresholding rule introduced by [Cai \(1999\)](#). Wavelet-based thresholding estimators are attractive for non-parametric function estimation because of their virtues from the viewpoints of spatial adaptivity, computational efficiency and asymptotic optimality properties. In the case of simple density estimation, wavelet thresholding is probably one of the most attractive nonlinear methods. We refer to e.g., [Antoniadis \(1997\)](#), [Härdle et al. \(1998\)](#) and [Vidakovic \(1999\)](#) for a detailed discussion of the performances of wavelet estimators and some of their advantages over traditional methods such as kernel-based or projection estimators.

We here explore the theoretical performance of our estimator under the \mathbb{L}_p risk with $p \geq 1$ over a very rich class of function spaces, namely Besov spaces. Sharp rates of convergence are obtained. Application of our theory on Example 1.2 above is described in detail. Finally, numerical experiments are carried out to illustrate the practical performance of our estimator. In particular, the numerical tests indicate that the block thresholding estimators compare very favorably to standard kernel methods.

1.5 Paper organization

The paper is structured as follows. Our wavelet estimator is described in Section 2. Section 3 presents our theoretical results. Simulations are detailed in Section 4. The proofs are postponed to Section 5.

2 Wavelet estimators

First of all, we briefly recall some key facts on wavelets and Besov spaces that will be essential to us in the sequel. Then we develop our nonlinear adaptive wavelet block thresholding estimator.

2.1 Wavelets and Besov balls

Let $b > 0$, $p > 0$ and $\mathbb{L}_p([-b, b]) = \left\{ h : [-b, b] \rightarrow \mathbb{R}; \|h\|_p^p = \int_{-b}^b |h(x)|^p dx < +\infty \right\}$.

For the purposes of this paper, we use compactly supported wavelet bases on $[-b, b]$. More precisely, we consider the Daubechies family db_{2R} with the scaling and wavelet functions ϕ and ψ , where $R \geq 2$ is a fixed integer, see e.g., [Mallat \(2009\)](#). Define the scaled and translated version of the ϕ and ψ

$$\phi_{j,k}(x) = 2^{j/2} \phi(2^j x - k), \quad \psi_{j,k}(x) = 2^{j/2} \psi(2^j x - k).$$

Then there exists an integer τ and a set of consecutive integers Λ_j such that $\text{Card}(\Lambda_j) = C2^j$ for a $C > 0$ and, for any integer $\ell \geq \tau$, the collection

$$\mathcal{B} = \{\phi_{\ell,k}, k \in \Lambda_\ell; \psi_{j,k}; j \in \mathbb{N} - \{0, \dots, \ell - 1\}, k \in \Lambda_j\},$$

is an orthonormal basis of $\mathbb{L}_2([-b, b])$.

Consequently, for any integer $\ell \geq \tau$, any $h \in \mathbb{L}_2([-b, b])$ can be expanded on \mathcal{B} as

$$h(x) = \sum_{k \in \Lambda_\ell} \alpha_{\ell,k} \phi_{\ell,k}(x) + \sum_{j=\ell}^{\infty} \sum_{k \in \Lambda_j} \beta_{j,k} \psi_{j,k}(x),$$

where

$$\alpha_{\ell,k} = \int_{-b}^b h(x) \phi_{\ell,k}(x) dx, \quad \beta_{j,k} = \int_{-b}^b h(x) \psi_{j,k}(x) dx. \quad (2.1)$$

As is traditional in the wavelet estimation literature, we will investigate the performance of our estimator by assuming that the unknown density f belongs to a Besov ball. The Besov norm for a function can be related to a sequence space norm on its wavelet coefficients. More precisely, let $M > 0$, $s \in (0, R)$, $q \geq 1$ and $r \geq 1$. A function $h \in \mathbb{L}_p([-b, b])$ belongs to $B_{q,r}^s(M)$ if and only if there exists a constant $M^* > 0$ (depending on M) such that the associated wavelet coefficients (2.1) satisfy

$$\left(\sum_{k \in \Lambda_\tau} |\alpha_{\tau,k}|^q \right)^{1/q} + \left(\sum_{j=\tau}^{\infty} \left(2^{j(s+1/2-1/q)} \left(\sum_{k \in \Lambda_j} |\beta_{j,k}|^q \right)^{1/q} \right)^r \right)^{1/r} \leq M^*,$$

with the usual modifications if $q = \infty$ or $r = \infty$.

In this expression, s is a smoothness parameter and q and r are norm parameters. They include many traditional smoothness spaces such as Hölder and Sobolev spaces. A comprehensive account on Besov spaces can be found in e.g., [Devore and Popov \(1988\)](#); [Meyer \(1992\)](#); [Härdle et al. \(1998\)](#).

2.2 Plug-in block wavelet estimator

Let us consider the general statistical framework described in Section 1 with *i.i.d.* sample X_1, \dots, X_n . First of all, we investigate the estimation of f via the so-called wavelet block hard thresholding estimator. We suppose that $\text{supp}(f) \subseteq [-b, b]$ with $b > 0$.

Let $p \geq 1$, and j_1 and j_2 be the integers corresponding to the finest and coarsest scales defined as

$$j_1 = \lfloor (\max(p, 2)/2) \log_2(\log n) \rfloor, \quad j_2 = \lfloor \log_2(n/\log n) \rfloor,$$

where $\lfloor a \rfloor$ denotes the whole number part of $a \in \mathbb{R}^+$. For any $j \in \{j_1, \dots, j_2\}$, let A_j and $U_{j,K}$ be given such that $\cup_{K \in A_j} U_{j,K} = \Lambda_j$, $|U_{j,K}| = L = \lfloor (\log n)^{\max(p, 2)/2} \rfloor$ and $U_{j,K} \cap U_{j,K'} = \emptyset$ for any $K \neq K'$ with $K, K' \in A_j$. In a nutshell, at each scale j , each $U_{j,K}$ is the set containing position indices of the wavelet coefficients inside block $K \in A_j$.

We define the wavelet block hard thresholding estimator of f by

$$\hat{f}(x) = \sum_{k \in \Lambda_{j_1}} \hat{\alpha}_{j_1,k} \phi_{j_1,k}(x) + \sum_{j=j_1}^{j_2} \sum_{K \in A_j} \sum_{k \in U_{j,K}} \hat{\beta}_{j,k} \mathbb{1}_{\left\{ \left(\sum_{k \in U_{j,K}} |\hat{\beta}_{j,k}|^p / L \right)^{1/p} \geq \kappa n^{-1/2} \right\}} \psi_{j,k}(x), \quad x \in [-b, b], \quad (2.2)$$

where $\mathbb{1}_{\{\cdot\}}$ denotes the indicator function, and

$$\hat{\alpha}_{j_1,k} = \frac{1}{n} \sum_{i=1}^n \phi_{j_1,k}(X_i), \quad \hat{\beta}_{j,k} = \frac{1}{n} \sum_{i=1}^n \psi_{j,k}(X_i)$$

and κ denotes a large enough constant.

The estimator \hat{f} was initially developed by Cai (1999) for the regression model under the \mathbb{L}_2 risk with equispaced deterministic samples. The \mathbb{L}_p risk version was studied in Picard and Tribouley (2000) for the standard density estimation problem and by Chesneau (2010) for the biased density estimation problem.

The idea underlying \hat{f} (2.2) is to operate a group/block selection: it keeps intact the large groups of unknown wavelet coefficients of f (2.1) and removes the others. Wavelet block thresholding is one of the most attractive non-linear thresholding methods, since it is both numerically straightforward to implement and asymptotically optimal for a large variety of Sobolev or Besov classes. Detailed references on the subject for various models include, but are not limited to, Cai (1999, 2002), Li and Xiao (2008); Li (2008), Picard and Tribouley (2000) and Chesneau (2008, 2010).

Finally, plugging (2.2) into (1.1) leads to the following estimator of g :

$$\hat{g}(x) = w(\hat{F}(x))\hat{f}(x), \quad x \in [-b, b], \quad (2.3)$$

where

$$\hat{F}(x) = \frac{1}{n} \sum_{i=1}^n \mathbb{1}_{\{X_i \leq x\}}. \quad (2.4)$$

The rest of the paper explores the theoretical and practical performances of \hat{g} .

3 Main results

In this section, we discuss the asymptotic properties of the proposed estimator. Rates of \mathbb{L}_p convergence are investigated under the following assumptions:

(A.1) Compact support: $\text{supp}(f) \subseteq [-b, b]$ with $b > 0$.

(A.2) Uniform boundedness: there exists a constant $C_1 > 0$ such that

$$\sup_{x \in [-b, b]} f(x) \leq C_1. \quad (3.1)$$

(A.3) Uniform Lipschitz continuity of w (with Lipschitz constant C_2):

$$|w(u) - w(v)| \leq C_2|u - v|, \quad \text{for all } (u, v) \in [0, 1]^2, \quad (3.2)$$

3.1 Estimator convergence rates

Theorem 3.1 studies the \mathbb{L}_p risk of \hat{g} (2.3) over Besov balls and Assumptions A.1-A.3 on f and w .

Theorem 3.1. *Consider the general statistical framework described in Section 1 (the estimation of g (1.1) is of interest). Suppose that Assumptions A.1-A.3 hold. Let $p \geq 1$ and \hat{g} be given by (2.3). Then, for any $f \in B_{q,r}^s(M)$, $q \geq 1$, $r \geq 1$ and $s \in (1/p, R)$, there exists a constant $C > 0$ such that*

$$\mathbb{E}(\|\hat{g} - g\|_p^p) \leq C\varphi_{n,p},$$

where

$$\varphi_{n,p} = \begin{cases} n^{-sp/(2s+1)}, & \text{if } q \geq p, \\ \left(\frac{\log n}{n}\right)^{sp/(2s+1)}, & \text{if } \{p > q, qs > (p-q)/2\}, \\ \left(\frac{\log n}{n}\right)^{(s-1/q+1/p)p/(2(s-1/q)+1)}, & \text{if } qs < (p-q)/2 \text{ or } \{qs = (p-q)/2, p \leq q/r\}, \\ \left(\frac{\log n}{n}\right)^{(s-1/q+1/p)p/(2(s-1/q)+1)} (\log n)^{p-q/r}, & \text{if } \{qs = (p-q)/2, p > q/r\}. \end{cases} \quad (3.3)$$

Note that the rate $\varphi_{n,p}$ in (3.3) is the near optimal (or optimal in some regimes) one in the minimax sense for f . See, e.g., Donoho et al. (1996) and Härdle et al. (1998). In plain words, the near optimality of the estimator of f is transferred to that of g through the plug-in principle.

The proof of Theorem 3.1 uses a suitable decomposition of the \mathbb{L}_p risk and capitalizes on results on the performances of \hat{f} (2.2) and \hat{F} (2.4) established in Chesneau (2010). See Section 5 for details.

3.2 An illustrative application

Let's recall Example 1.2, where $\{X_n\}_{n \in \mathbb{N}^*}$ is a sequence of *i.i.d.* random variables with pdf f and N be a discrete random variable of values in \mathbb{N}^* independent of this sequence. The density of $Y = \max(X_1, \dots, X_N)$ is given by (1.2).

Suppose that Assumptions A.1-A.2 hold. Thus, several examples for the distribution of N can be considered.

(a) Degenerate distribution. $\mathbb{P}(N = m) = 1$. Then

$$w(u) = mu^{m-1}, \quad u \in [0, 1], \quad (3.4)$$

(b) Geometric distribution. $N \sim G(\eta)$ ($\mathbb{P}(N = k) = \eta(1 - \eta)^{k-1}$, $k \in \mathbb{N}^*$). Then

$$w(u) = \frac{\eta}{(1 - u(1 - \eta))^2}, \quad u \in [0, 1], \quad (3.5)$$

(c) Poisson plus 1 distribution. $N = P + 1$ with $P \sim \mathcal{P}(\lambda)$ ($\mathbb{P}(N = k) = e^{-\lambda} \frac{\lambda^{k-1}}{(k-1)!}$, $k \in \mathbb{N}^*$). Then

$$w(u) = e^{-\lambda} e^{\lambda u} (1 + \lambda u), \quad u \in [0, 1].$$

Remark 3.2. In examples (a)–(c) above it is clear that Assumption 3 is satisfied; more precisely, in example (a), we have $C_2 = m(m-1)$; in example (b), we have $C_2 = 2(1 - \eta)/\eta^2$; in example (c), we have $C_2 = \lambda(2 + \lambda)$.

In this context, Theorem 3.1 can be applied. Let $p \geq 1$ and \hat{g} be the estimator given in (2.3). Then, for any $f \in B_{q,r}^s(M)$, $q \geq 1$, $r \geq 1$ and $s \in (1/p, R)$ there exists a constant $C > 0$ such that

$$\mathbb{E}(\|\hat{g} - g\|_p^p) \leq C\varphi_{n,p},$$

where $\varphi_{n,p}$ is given by (3.3).

Remark 3.3. Taking $m = 2$, the obtained rate is similar to the one attained by the kernel estimators developed by Chen and Hsu (2004); the only difference is the extra-logarithmic term $(\log n)^{2s/(2s+1)}$. However, unlike kernel estimators Chen and Hsu (2004), our procedure \hat{g} is adaptive and our rate of convergence holds for a wider class of functions f including Hölder class, Sobolev class, etc..

4 Simulation results

We now illustrate these theoretical results by a simulation study within the context described in Section 3.2. That is, we consider the problem of estimating the density g of the maximum of a random number N of *i.i.d.* random variables. From a reliability study standpoint, this problem corresponds to a parallel system with N identical components. Thereby, we have considered two numerical examples. They complement the asymptotic results of 3.1.

Computational aspects. In the sequel, we will refer to our adaptive wavelet estimator (2.3) simply as Block. We have compared its performance to alternatives from the literature on several densities. We have considered the uniform distribution, as well as a family of normal mixture densities (“SeparatedBimodal”, “Kurtotic” and “StronglySkewed”, initially introduced in Marron and Wand (1992)) representing different degrees of smoothness (see Figure 1). We assumed that the density function f of the X_i 's has a compact support included in $[-b, b]$. We have used the formulae given by Marron and Wand (1992) to simulate such densities so that

$$\min_l(\mu_l - 3\sigma_l) = -3, \quad \max_l(\mu_l + 3\sigma_l) = 3,$$

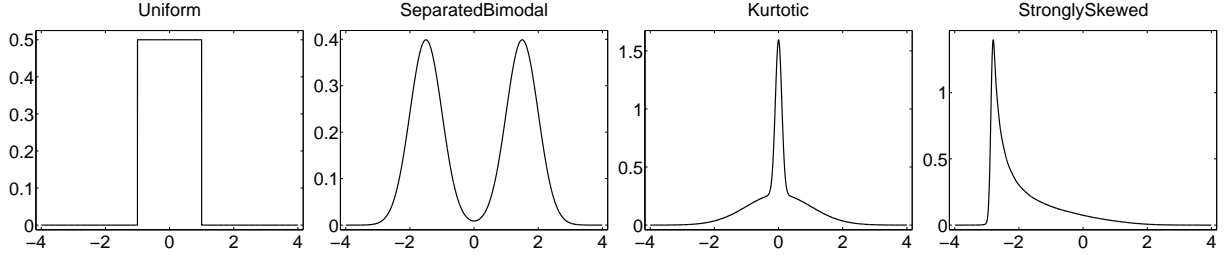


Figure 1: Test densities.

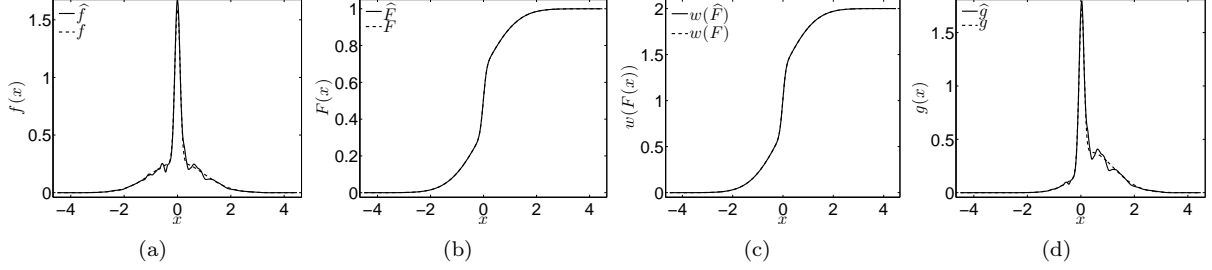


Figure 2: Typical reconstructions from a single simulation with $n = 1000$ for the Kurtotic density. The dashed line depicts the original density and the solid one depicts its wavelet block estimate. (a): $\hat{f}(x)$. (b): $\hat{F}(x)$. (c): $w(\hat{F}(x))$. (d): $\hat{g}(x) = w(\hat{F}(x))\hat{f}(x)$

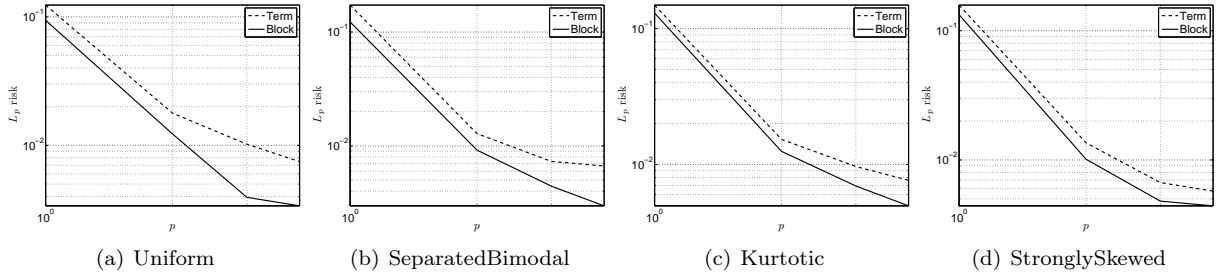


Figure 3: The influence of p in the numerical values of the L_p risk (in log-log scale) of Block (solid) and term-by-term (dashed) thresholding ($L = 1$).

where $l = 1, \dots, d$ with d the number of densities in the mixture (see, (Marron and Wand, 1992, Section 4, Table 1), for the values of the parameters). Thereby, it is very unlikely to have values outside the interval $[-4, 4]$ and we loose little by assuming compact support. The same kind of assumption was made in the context of wavelet density estimation by Vannucci and Vidakovic (1997). In order to simplify the presentation of the results, one can simply rescale the data such that they fall into $[-b, b]$ (which covered the full range of all observed data). Thus, the density was evaluated at $T = 2^J$ equispaced points $t_i = 2ib/T$, $i = -T/2, \dots, T/2 - 1$ between $-b$ and b , where J is the index of the highest resolution level and T is the number of discretization points. The primary level $j_1 = 3$, $T = 512$ and the Symmlet wavelet with 6 vanishing moments were used throughout all experiments. All simulations have been implemented under Matlab.

Results and discussion. In order to illustrate Theorem 3.1, we study the influence of p on the numerical performances of the Block and the term-by-term ($L = 1$) thresholding estimator. Let us first consider a parallel system with $m = 2$ identical independent components. Then, the corresponding weighted function is (3.4) and the goal is to estimate g in (1.2) from X_1, \dots, X_n sample simulated from one of the test densities. A typical example of estimation for the Kurtotic density (for $p = 2$), with $n = 1000$ is given in Figure 2. The mean L_p risk of \hat{g} i.e., $R_p(\hat{g}, g) = (1/T) \sum_{i=-T/2}^{T/2-1} |\hat{g}(t_i) - g(t_i)|^p$, is obtained with 10 samples for $n = 1000$, and it is plotted as a function of p in Figure 3. As predicted by Theorem 3.1,

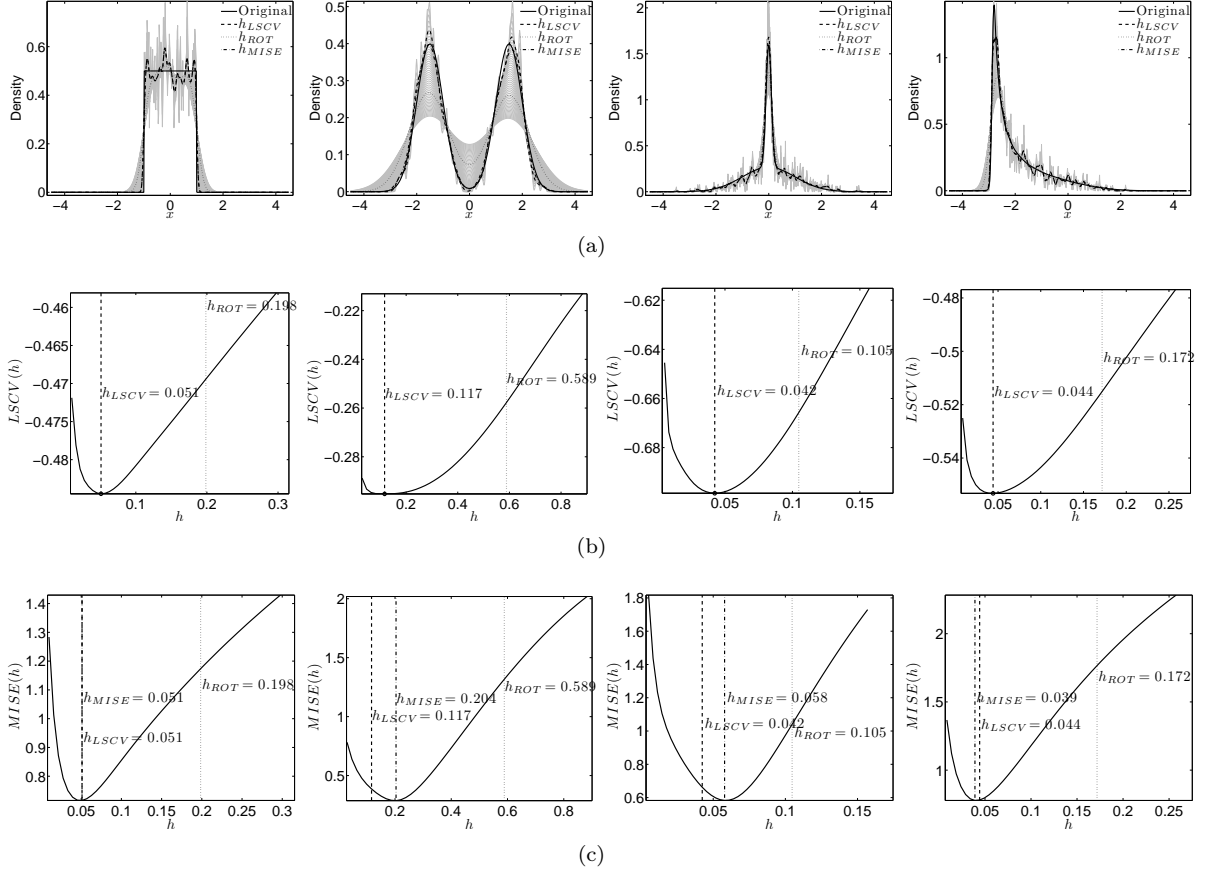


Figure 4: (a): Density estimates. (b): Graph of the the LSCV function versus the kernel bandwidth h for each of the tested densities, the vertical dashed lines represent the value of h that minimizes $LSCV(h)$. (c): The solid line depicts the estimated MISE as a function of h , the vertical dashed-dotted lines represent the true MISE-minimizing bandwidth h_{MISE} and the vertical dotted lines represent the pilot bandwidth h_{ROT} .

the larger p , the smaller L_p risk of \hat{g} . We can see that our estimation procedure provides better results than the term-by-term thresholding ($L = 1$) in all cases. In particular, the risk improvement achieved by the block estimators upon the term-by-term estimator is significant for the non-smooth Uniform density. This is in agreement with the predictions of our theoretical findings.

In our second example, the adaptive estimator described in Section 3.2 is tested when N follows a Geometric distribution, so that the weight function is that given by (3.5). This example is devoted to a simulation study comparing the performance of the block hard thresholding estimator with that of the traditional kernel defined as follows

$$\hat{f}_h(x) = \frac{1}{nh} \sum_{i=1}^n K\left(\frac{x - X_i}{h}\right), \quad (4.1)$$

where the positive kernel K satisfies $\int K(x)dx = 1$ and the smoothing parameter h is known as the bandwidth.

Many procedures of bandwidth selection for kernel density estimation have been developed in the literature (see, e.g., Silverman (1986)). We use least-squares cross-validation (LSCV) (Rudemo (1982), Bowman (1984)) where the bandwidth is defined as

$$h_{LSCV} = \underset{h}{\operatorname{argmin}} \int_{-b}^b \hat{f}_h(x)^2 dx - 2n^{-1} \sum_{i=1}^n \hat{f}_{-i}(X_i),$$

and \hat{f}_{-i} is the *leave-one out* kernel estimator constructed from the data without the observation X_i . It

is motivated by the fact that for independent data

$$\text{LSCV}(h) = \int_{-b}^b \hat{f}_h(x)^2 dx - 2n^{-1} \sum_{i=1}^n \hat{f}_{-i}(X_i)$$

is an unbiased estimator of $\text{MISE}(h) = \int_{-b}^b f^2(x) dx$. One frequently used cross-validation (CV) procedure is the K-fold CV (as described e.g. in (Hastie et al., 2009, Section 7.10)) in which the data set X_1, \dots, X_n is randomly partitioned into K approximately equal-sized and non-overlapping subsets S_1, \dots, S_K . To obtain the bandwidths h_{LSCV} , we have performed a 10-fold CV, using a Gaussian kernel, with a simple “rule-of-thumb” pilot bandwidth h_{ROT} . Figure 4(b) contains a plot of the LSCV function versus the kernel bandwidth h and Figure 4(c) the estimated MISE as a function of h . For each density, it is clear from this figure (Figure 4(b)), that the value of h_{LSCV} is the unambiguous minimizer of $\text{LSCV}(h)$. We see that h_{LSCV} provides a decent approximation, close to h_{MISE} for all test densities. For the StronglySkewed density, the bandwidth which minimizes $\text{MISE}(h)$ in this case is $h_{\text{MISE}} = 0.039$ and $h_{\text{LSCV}} = 0.044$. In this case, for the Uniform density, $h_{\text{MISE}} = h_{\text{LSCV}} = 0.051$.

We then compared the performance of the Block estimator \hat{g} with that of the plug-in kernel estimator, say \hat{g}_{LSCV} , given by $\hat{g}_{\text{LSCV}} = w(\hat{F}(x))\hat{f}_{\text{LSCV}}(x)$, where \hat{F} is defined by (2.4) and \hat{f}_{LSCV} is given by (4.1) with h_{LSCV} . Figure 5 shows the results of \hat{g} and \hat{g}_{LSCV} for $N \sim G(\eta)$, with $\eta = 0.9$, $\eta = 0.5$ and $\eta = 0.1$ respectively. Table 1 presents the MISE for samples sizes $n = 1000, 2000$ and 5000 . For virtually all cases, the Block estimator consistently showed lower \mathbb{L}_2 risk than \hat{g}_{LSCV} , with the exception of the (very smooth) SeparatedBimodal density for which the kernel estimator performs slightly better. This comes at no surprise given that this density is very smooth. Additionally, small discrepancies in the estimate of f may lead to substantial discrepancies for the estimate of g at the locations overweighted by $w(\hat{F}(\cdot))$. It turns out that this is the case for the Geometric distribution where the weights evolve in $O(1/\eta)$ at high values of x , and thus the discrepancies in \hat{g} increase as η gets smaller. However, the kernel estimator \hat{g}_{LSCV} seems to be more concerned (see, Figure 5(c)), confirming that Block generally provides a better estimate of f . Furthermore, as expected, for both methods, and in all cases, the MISE is decreasing as the sample size increases. Without any prior smoothness knowledge on the unknown density, the Block estimator provides very competitive results in comparison to \hat{g}_{LSCV} .

Table 1: $1000 \times \text{MISE}$ values from 50 replications of sample sizes $n = 1000, 2000$ and 5000 , when N follows a Geometric distribution of parameter η .

| Uniform | | | | | | | | | |
|------------------|--------------|-------|------|--------------|-------|-------|--------------|-------|-------|
| 1.0e-03× | $\eta = 0.9$ | | | $\eta = 0.5$ | | | $\eta = 0.1$ | | |
| n | 1000 | 2000 | 5000 | 1000 | 2000 | 5000 | 1000 | 2000 | 5000 |
| Block | 10.49 | 7.03 | 4.18 | 11.15 | 7.61 | 4.85 | 19.15 | 15.05 | 13.39 |
| Kernel | 13.76 | 11.37 | 9.32 | 17.62 | 14.99 | 12.60 | 70.50 | 63.24 | 55.94 |
| SeparatedBimodal | | | | | | | | | |
| 1.0e-03× | $\eta = 0.9$ | | | $\eta = 0.5$ | | | $\eta = 0.1$ | | |
| n | 1000 | 2000 | 5000 | 1000 | 2000 | 5000 | 1000 | 2000 | 5000 |
| Block | 8.53 | 6.29 | 3.69 | 9.09 | 6.89 | 3.90 | 15.27 | 11.08 | 7.12 |
| Kernel | 6.30 | 4.98 | 3.54 | 7.00 | 5.32 | 3.82 | 13.26 | 10.86 | 7.79 |
| Kurtotic | | | | | | | | | |
| 1.0e-03× | $\eta = 0.9$ | | | $\eta = 0.5$ | | | $\eta = 0.1$ | | |
| n | 1000 | 2000 | 5000 | 1000 | 2000 | 5000 | 1000 | 2000 | 5000 |
| Block | 11.31 | 8.03 | 5.52 | 11.93 | 8.04 | 5.57 | 18.08 | 9.22 | 7.01 |
| Kernel | 12.27 | 8.00 | 5.83 | 13.03 | 8.44 | 6.13 | 21.97 | 13.04 | 9.66 |
| StronglySkewed | | | | | | | | | |
| 1.0e-03× | $\eta = 0.9$ | | | $\eta = 0.5$ | | | $\eta = 0.1$ | | |
| n | 1000 | 2000 | 5000 | 1000 | 2000 | 5000 | 1000 | 2000 | 5000 |
| Block | 9.91 | 7.78 | 5.12 | 8.69 | 7.02 | 4.70 | 10.12 | 9.03 | 5.93 |
| Kernel | 10.57 | 8.13 | 5.97 | 11.16 | 8.68 | 6.24 | 19.62 | 15.54 | 10.86 |

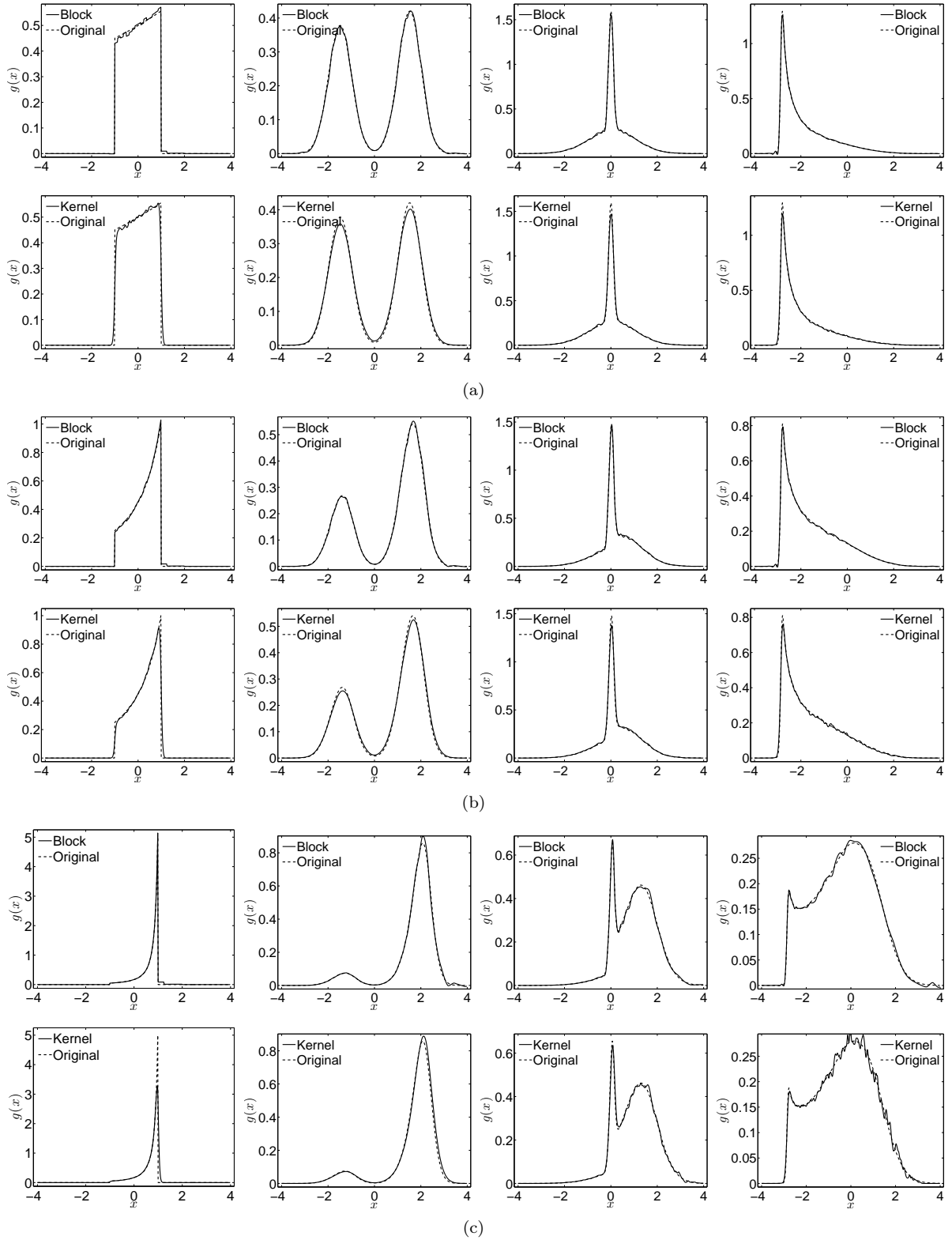


Figure 5: Original densities (dashed), Block thresholding estimator \hat{g} (solid) (1st row), kernel estimator \hat{g}_{LSCV} (solid) (2nd row) from 50 replications of $n = 1000$ samples X_1, \dots, X_n . From left to right Uniform, SeparatedBimodal, Kurtotic and StronglySkewed. $N \sim G(\eta)$, with (a) $\eta = 0.9$, (b) $\eta = 0.5$ and (c) $\eta = 0.1$.

5 Proofs

Proof of Theorem 3.1. Observe that

$$\widehat{g}(x) - g(x) = w(\widehat{F}(x))(\widehat{f}(x) - f(x)) + f(x)(w(\widehat{F}(x)) - w(F(x))).$$

By Assumption A.3 implying $\sup_{x \in [0,1]} w(x) \leq C$, together with Assumptions 2, we have

$$\begin{aligned} |\widehat{g}(x) - g(x)| &\leq C(|\widehat{f}(x) - f(x)| + |w(\widehat{F}(x)) - w(F(x))|) \\ &\leq C(|\widehat{f}(x) - f(x)| + |\widehat{F}(x) - F(x)|). \end{aligned}$$

By the Jensen inequality, we have

$$\mathbb{E}(\|\widehat{g} - g\|_p^p) \leq C(\mathbb{E}(\|\widehat{f} - f\|_p^p) + \mathbb{E}(\|\widehat{F} - F\|_p^p)).$$

It follows from (Chesneau, 2010, Theorem 4.1 with $w(x) = 1 = \mu$) that

$$\mathbb{E}(\|\widehat{f} - f\|_p^p) \leq C\varphi_{n,p},$$

where $\varphi_{n,p}$ is given by (3.3).

Now note that

$$\widehat{F}(x) - F(x) = \frac{1}{n} \sum_{i=1}^n U_i(x),$$

with $U_i(x) = \mathbb{1}_{\{X_i \leq x\}} - F(x)$. Since $U_1(x), \dots, U_n(x)$ are *i.i.d.* with $\mathbb{E}(U_1(x)) = 0$, $|U_1(x)| \leq 2$ and $\mathbb{E}((U_1(x))^2) \leq 4$, the Rosenthal inequality (see Rosenthal (1970)) yields

$$\mathbb{E}(\|\widehat{F} - F\|_p^p) \leq C \sup_{x \in \mathbb{R}} \mathbb{E} \left((\widehat{F}(x) - F(x))^p \right) \leq C \frac{1}{n^{p/2}} \leq C\varphi_{n,p}.$$

Combining the inequalities above, we obtain the desired result i.e.,

$$\mathbb{E}(\|\widehat{g} - g\|_p^p) \leq C\varphi_{n,p}.$$

□

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