Maximum likelihood estimation of a log-concave density and its distribution function: basic properties and uniform consistency[*]

Lutz Dümbgen and Kaspar Rufibach

University of Bern and University of Zürich

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Abstract

We study nonparametric maximum likelihood estimation of a log-concave probability density and its distribution and hazard function. Some general properties of these estimators are derived from two characterizations. It is shown that the rate of convergence with respect to supremum norm on a compact interval for the density and hazard rate estimator is at least $(\log(n)/n)^{1/3}$ and typically $(\log(n)/n)^{2/5}$ whereas the difference between the empirical and estimated distribution function vanishes with rate $o_{\mathrm{p}}(n^{-1/2})$ under certain regularity assumptions.

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1 Introduction

Two common approaches to nonparametric density estimation are smoothing methods and qualitative constraints. The former approach includes, among others, kernel density estimators, estimators based on discrete wavelets or other series expansions, and estimators based on roughness penalization. Good starting points for the vast literature in this field are Silverman (1982, 1986) and Donoho et al. (1996). A common feature of all these methods is that they involve certain tuning parameters, e.g. the order of a kernel and the bandwidth. A proper choice of these parameters is far from trivial, since optimal values depend on unknown properties of the underlying density f. The second approach avoids such problems by imposing qualitative properties on f, e.g. monotonicity or convexity on certain intervals in the univariate case. Such assumptions are often plausible or even justified rigorously in specific applications.

Density estimation under shape constraints was first considered by Grenander (1956), who found that the nonparametric maximum likelihood estimator (NPMLE) $\hat{f}_n^{\mathrm{mon}}$ of a non-increasing density function f on $[0, \infty)$ is given by the left derivative of the least concave majorant of the empirical cumulative distribution function on $[0, \infty)$. This work was continued by Prakasa Rao (1969) and Groeneboom (1985, 1988), who established asymptotic distribution theory for $n^{1/3}(f - \hat{f}_n^{\mathrm{mon}})(t)$ at a fixed point $t > 0$ under certain regularity conditions and analyzed the non-gaussian limit distribution. For various estimation problems involving monotone functions, the typical rate of convergence is $O_{\mathrm{p}}(n^{-1/3})$ pointwise. The rate of convergence with respect to supremum norm is further decelerated by a factor of $\log(n)^{1/3}$ (Jonker and van der Vaart 2001). For applications of monotone density estimation consult e.g. Barlow et al. (1972) or Robertson et al. (1988).

Monotone estimation can be extended to cover unimodal densities. Remember that a density f on the real line is unimodal if there exists a number $M = M(f)$ such that f is non-decreasing on $(-\infty, M]$ and non-increasing on $[M, \infty)$. If the true mode is known a priori, unimodal density estimation boils down to monotone estimation in a straightforward manner, but the situation is different if M is unknown. In that case, the likelihood is unbounded, problems being caused by observations too close to a hypothetical mode. Even if the mode was known, the density estimator is inconsistent at the mode, a phenomenon called "spiking". Several methods were proposed to remedy this problem, see Wegman (1970), Woodroofe and Sun (1993), Meyer and Woodroofe (2004) or Kulikov and Lopuhaä (2006), but all of them require additional constraints on f.

The combination of shape constraints and smoothing was assessed by Eggermont and La-Riccia (2000). To improve the slow rate of convergence of $n^{-1/3}$ in the space $\mathrm{L}_1(\mathbb{R})$ for arbitrary unimodal densities, they derived a Grenander type estimator by taking the derivative of the least concave majorant of an integrated kernel density estimator rather than the empirical distribution function directly, yielding a rate of convergence of $O_{\mathrm{p}}(n^{-2/5})$.

Estimation of a convex decreasing density on $[0, \infty)$ was pioneered by Anevski (1994, 2003). The problem arose in a study of migrating birds discussed by Hampel (1987). Groeneboom et al.

(2001) provide a characterization of the estimator as well as consistency and limiting behavior at a fixed point of positive curvature of the function to be estimated. They found that the estimator has to be piecewise linear with knots between the observation points. Under the additional assumption that the true density f is twice continuously differentiable on $[0, \infty)$, they show that the MLE converges at rate $O_{\mathrm{p}}(n^{-2/5})$ pointwise, fairly better than in the monotone case. Monotonicity and convexity constraints on densities on $[0, \infty)$ have been embedded into the general framework of k–monotone densities by Balabdaoui and Wellner (2008). See Section 5 for a more thorough discussion of the similarities and differences between k–monotone density estimation and the present work.

In the present paper we impose an alternative and quite natural shape constraint on the density f, namely, log-concavity. That means,

$$f(x) \;=\; \exp\varphi(x)$$

for some concave function $\varphi : \mathbb{R} \to [-\infty, \infty)$. This class is rather flexible in that it generalizes many common parametric densities. These include all nondegenerate normal densities, all Gamma densities with shape parameter ≥ 1, all Weibull densities with exponent ≥ 1, and all beta densities with parameters ≥ 1. Further examples are the logistic and Gumbel densities. Log-concave densities are of interest in econometrics, see Bagnoli and Bergstrom (2005) for a summary and further examples. Barlow and Proschan (1975) describe advantageous properties of log-concave densities in reliability theory, while Chang and Walther (2007) use log-concave densities as ingredient of nonparametric mixture models. In nonparametric Bayesian analysis, log-concavity is of certain relevance, too (Brooks 1998).

Note that log-concavity of a density implies that it is also unimodal. It will turn out that by imposing log-concavity one circumvents the spiking problem mentioned before, which yields a new approach to estimate a unimodal, possibly skewed density. Moreover, the log-concave density estimator is fully automatic in the sense that there is no need to select any bandwidth, kernel function or other tuning parameters. Finally, simulating data from the estimated density is rather easy. All these properties make the new estimator appealing for its use in statistical applications.

Little large sample theory is available for log-concave estimators so far. Sengupta and Paul (2005) considered testing for log-concavity of distribution functions on a compact interval. Walther (2002) introduced an extension of log-concavity in the context of certain mixture models, but his theory doesn't cover asymptotic properties of the density estimators themselves. Pal et al. (2006) proved the log-concave NPMLE to be consistent, but without rates of convergence.

Concerning the computation of the log-concave NPMLE, Walther (2002) and Pal et al. (2006) used a crude version of the iterative convex minorant (ICM) algorithm. A detailed description and comparison of several algorithms can be found in Rufibach (2007), while Dümbgen et al. (2007a) describe an active set algorithm, which is similar to the vertex reduction algorithms presented by Groeneboom et al. (2008) and seems to be the most efficient one by now. The ICM and active

set algorithms are implemented within the R package `"logcondens"`, accessible via `"CRAN"`. Corresponding Matlab code is available from the first author's homepage.

In Section 2 we introduce the log-concave maximum likelihood density estimator, discuss its basic properties and derive two characterizations. In Section 3 we illustrate this estimator with a real data example and explain briefly how to simulate data from the estimated density. Consistency of this density estimator and the corresponding estimator of the distribution function are treated in Section 4. It is shown that the supremum norm between estimated density, $\hat{f}_n$, and true density on compact subsets of the interior of $\{f > 0\}$ converges to zero at rate $O_{\mathrm{p}}\big((\log(n)/n)^\gamma\big)$ with $\gamma \in [1/3, 2/5]$ depending on f's smoothness. In particular, our estimator adapts to the unkown smoothness of f. Consistency of the density estimator entails consistency of the distribution function estimator. In fact, under additional regularity conditions on f, the difference between the empirical c.d.f. and the estimated c.d.f. is of order $o_{\mathrm{p}}(n^{-1/2})$ on compact subsets of the interior of $\{f > 0\}$.

As a by-product of our estimator note the following. Log-concavity of the density function f also implies that the corresponding hazard function $h = f/(1 - F)$ is non-decreasing (cf. Barlow and Proschan 1975). Hence our estimators of f and its c.d.f. F entail a consistent and non-decreasing estimator of h, as pointed out at the end of Section 4.

Some auxiliary results, proofs and technical arguments are deferred to Section A.

2 The estimators and their basic properties

Let X be a random variable with distribution function F and Lebesgue density

$$f(x) \;=\; \exp \varphi(x)$$

for some concave function $\varphi : \mathbb{R} \to [-\infty, \infty)$. Our goal is to estimate f based on a random sample of size $n > 1$ from F. Let $X_1 < X_2 < \cdots < X_n$ be the corresponding order statistics. For any log-concave probability density f on $\mathbb{R}$, the normalized log-likelihood function at f is given by

$$\int \log f \, \mathrm{d}\mathbb{F}_n \;=\; \int \varphi \, \mathrm{d}\mathbb{F}_n \tag{1}$$

where $\mathbb{F}_n$ stands for the empirical distribution function of the sample. In order to relax the constraint of f being a probability density and to get a criterion function to maximize over the convex set of *all* concave functions φ, we employ the standard trick of adding a Lagrange-term to (1), leading to the functional

$$\Psi_n(\varphi) \;:=\; \int \varphi \, \mathrm{d}\mathbb{F}_n - \int \exp \varphi(x) \, \mathrm{d}x$$

(see Silverman, 1982, Theorem 3.1). The nonparametric maximum likelihood estimator of $\varphi = \log f$ is the maximizer of this functional over all concave functions,

$$\hat{\varphi}_n \;:=\; \underset{\varphi \text{ concave}}{\arg\max} \, \Psi_n(\varphi),$$

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and $\hat{f}_n := \exp \hat{\varphi}_n$.

Existence, uniqueness and shape of $\hat{\varphi}_n$. One can easily show that $\Psi_n(\varphi) > -\infty$ if, and only if, φ is real-valued on $[X_1, X_n]$. The following theorem was proved independently by Pal et al. (2006) and Rufibach (2006). It follows also from more general considerations in Dümbgen et al. (2007a, Section 2).

Theorem 2.1. *The NPMLE $\hat{\varphi}_n$ exists and is unique. It is linear on all intervals $[X_j, X_{j+1}]$, $1 \leq j < n$. Moreover, $\hat{\varphi}_n = -\infty$ on $\mathbb{R} \setminus [X_1, X_n]$.*

Characterizations and further properties. We provide two characterizations of the estimators $\hat{\varphi}_n$, $\hat{f}_n$ and the corresponding distribution function $\hat{F}_n$, i.e. $\hat{F}_n(x) = \int_{-\infty}^{x} \hat{f}_n(r) \, dr$. The first characterization is in terms of $\hat{\varphi}_n$ and perturbation functions:

Theorem 2.2. *Let $\widetilde{\varphi}$ be a concave function such that $\{x : \widetilde{\varphi}(x) > -\infty\} = [X_1, X_n]$. Then $\widetilde{\varphi} = \hat{\varphi}_n$ if, and only if,*

$$\int \Delta(x) \, d\mathbb{F}_n(x) \leq \int \Delta(x) \exp \widetilde{\varphi}(x) \, dx \tag{2}$$

for any $\Delta : \mathbb{R} \to \mathbb{R}$ such that $\widetilde{\varphi} + \lambda\Delta$ is concave for some $\lambda > 0$.

Plugging in suitable perturbation functions Δ in Theorem 2.2 yields valuable information about $\hat{\varphi}_n$ and $\hat{F}_n$. For a first illustration, let $\mu(G)$ and $\mathrm{Var}(G)$ be the mean and variance, respectively, of a distribution (function) G on the real line with finite second moment. Setting $\Delta(x) := \pm x$ or $\Delta(x) := -x^2$ in Theorem 2.4 yields:

Corollary 2.3.
$$\mu(\hat{F}_n) = \mu(\mathbb{F}_n) \quad \text{and} \quad \mathrm{Var}(\hat{F}_n) \leq \mathrm{Var}(\mathbb{F}_n).$$

Our second characterization is in terms of the empirical distribution function $\mathbb{F}_n$ and the estimated distribution function $\hat{F}_n$. For a continuous and piecewise linear function $h : [X_1, X_n] \to \mathbb{R}$ we define the set of its "knots" to be

$$\mathcal{S}_n(h) := \{t \in (X_1, X_n) : h'(t-) \neq h'(t+)\} \cup \{X_1, X_n\}.$$

Recall that $\hat{\varphi}_n$ is an example for such a function h with $\mathcal{S}_n(\hat{\varphi}_n) \subset \{X_1, X_2, \ldots, X_n\}$.

Theorem 2.4. *Let $\widetilde{\varphi}$ be a concave function which is linear on all intervals $[X_j, X_{j+1}]$, $1 \leq j < n$, while $\widetilde{\varphi} = -\infty$ on $\mathbb{R} \setminus [X_1, X_n]$. Defining $\widetilde{F}(x) := \int_{-\infty}^{x} \exp \widetilde{\varphi}(r) \, dr$, we assume further that $\widetilde{F}(X_n) = 1$. Then $\widetilde{\varphi} = \hat{\varphi}_n$ and $\widetilde{F} = \hat{F}_n$ if, and only if, for arbitrary $t \in [X_1, X_n]$,*

$$\int_{X_1}^{t} \widetilde{F}(r) \, dr \leq \int_{X_1}^{t} \mathbb{F}_n(r) \, dr \tag{3}$$

with equality in case of $t \in \mathcal{S}_n(\widetilde{\varphi})$.

A particular consequence of Theorem 2.4 is that the distribution function estimator $\hat{F}_n$ is very close to the empirical distribution function $\mathbb{F}_n$ on $\mathcal{S}_n(\hat{\varphi}_n)$:

Corollary 2.5.

$$\mathbb{F}_n - n^{-1} \ \leq \ \hat{F}_n \ \leq \ \mathbb{F}_n \quad on \quad \mathcal{S}_n(\hat{\varphi}_n).$$

Figure 1 illustrates Theorem 2.4 and Corollary 2.5. The upper plot displays $\mathbb{F}_n$ and $\hat{F}_n$ for a sample of $n = 25$ random numbers generated from a Gumbel distribution with density $f(x) = e^{-x} \exp(-e^{-x})$ on $\mathbb{R}$. The dotted vertical lines indicate the "kinks" of $\hat{\varphi}_n$, i.e. all $t \in \mathcal{S}_n(\hat{\varphi}_n)$. Note that $\hat{F}_n$ and $\mathbb{F}_n$ are indeed very close on the latter set with equality at the right endpoint X_n. The lower plot shows the process

$$D(t) \ := \ \int_{X_1}^t (\hat{F}_n - \mathbb{F}_n)(r) \, \mathrm{d}r$$

for $t \in [X_1, X_n]$. As predicted by Theorem 2.4, this process is nonpositive and equals zero on $\mathcal{S}_n(\hat{\varphi}_n)$.

3 A data example

In a recent consulting case, a company asked for Monte Carlo experiments to predict the reliability of a certain device they produce. The reliability depends in a certain deterministic way on five different and independent random input parameters. For each input parameter a sample was available, and the goal was to fit a suitable distribution to simulate from. Here we just focus on one of these input parameters.

At first we considered two standard approaches to estimate the unknown density f, namely, (i) fitting a gaussian density $\hat{f}_{\mathrm{par}}$ with mean $\mu(\mathbb{F}_n)$ and variance $\hat{\sigma}^2 := n(n-1)^{-1}\mathrm{Var}(\mathbb{F}_n)$, and (ii) the kernel density estimator

$$\hat{f}_{\mathrm{ker}}(x) \ := \ \int \phi_{\hat{\sigma}/\sqrt{n}}(x - y) \, \mathrm{d}\mathbb{F}_n(y),$$

where ϕ_σ denotes the density of $\mathcal{N}(0, \sigma^2)$. This very small bandwidth $\hat{\sigma}/\sqrt{n}$ was chosen to obtain a density with variance $\hat{\sigma}^2$ and to avoid putting too much weight into the tails.

Looking at the data, approach (i) is clearly inappropriate because our sample of size $n = 787$ revealed a skewed and significantly non-gaussian distribution. This can be seen in Figure 2, where the multimodal curve corresponds to $\hat{f}_{\mathrm{ker}}$, while the dashed line depicts $\hat{f}_{\mathrm{par}}$. Approach (ii) yielded Monte Carlo results agreeing well with measured reliabilities, but the engineers questioned the multimodality of $\hat{f}_{\mathrm{ker}}$. Choosing a kernel estimator with larger bandwidth would overestimate the variance and put too much weight into the tails. Thus we agreed on a third approach and estimated f by a slightly smoothed version of $\hat{f}_n$,

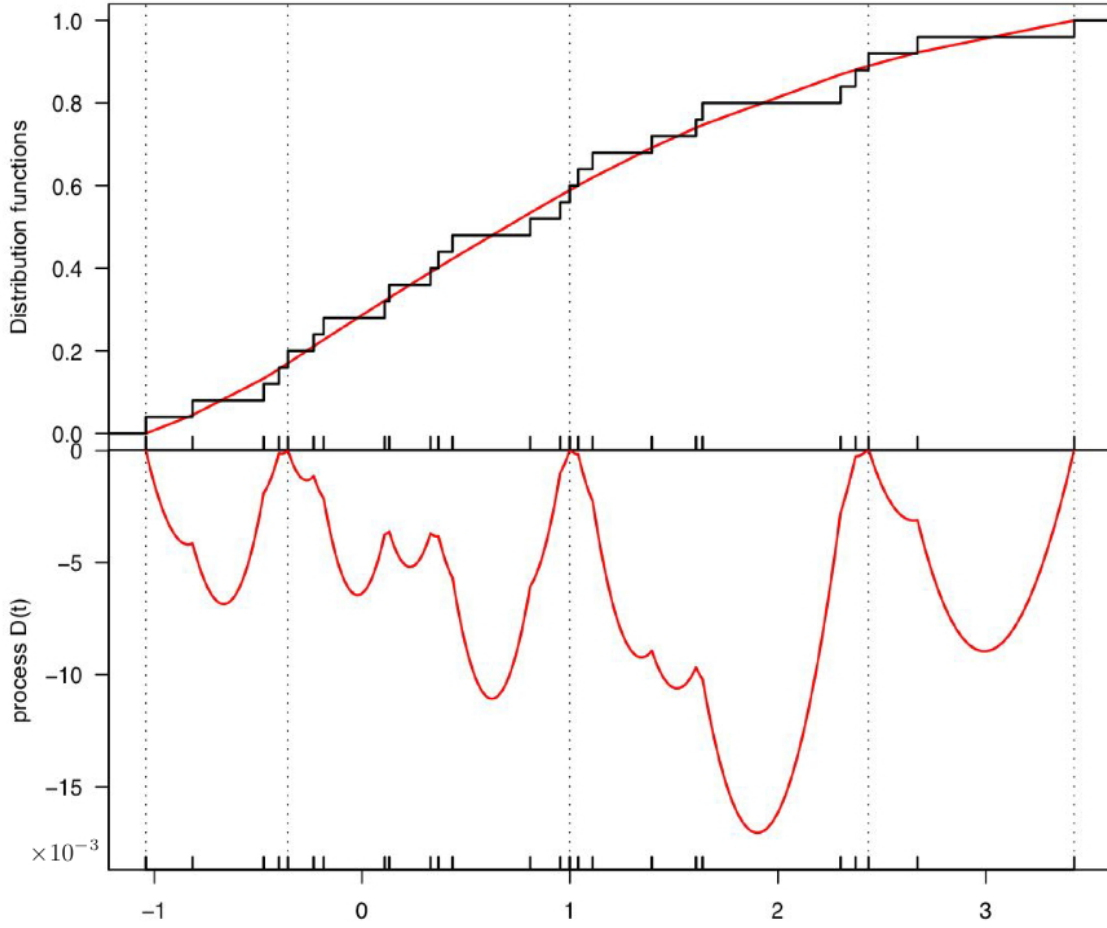$$\hat{f}_n^* \ := \ \int \phi_{\hat{\gamma}}(x - y) \, \mathrm{d}\hat{F}_n(y)$$

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Figure 1: Distribution functions and the process $D(t)$ for a Gumbel sample.

with $\hat{\gamma}^2 := \hat{\sigma}^2 - \text{Var}(\hat{F}_n)$, so that the variance of $\hat{f}_n^*$ coincides with $\hat{\sigma}^2$. Since log-concavity is preserved under convolution (cf. Prékopa 1971), $\hat{f}_n^*$ is log-concave, too. For the explicit computation of $\text{Var}(\hat{F}_n)$, see Dümbgen et al. (2007a). By smoothing we also avoid the small discontinuities of $\hat{f}_n$ at X_1 and X_n. This density estimator is the skewed unimodal curve in Figure 2. It yielded convincing results in the Monte Carlo simulations, too.

Note that both estimators $\hat{f}_n$ and $\hat{f}_n^*$ are fully automatic. Moreover it is very easy to sample from these densities: Let $\mathcal{S}_n(\hat{\varphi}_n)$ consist of $x_0 < x_1 < \cdots < x_m$, and consider the data X_i temporarily as fixed. Now

(a) generate a random index $J \in \{1, 2, \ldots, m\}$ with $\mathbb{P}(J = j) = \hat{F}_n(x_j) - \hat{F}_n(x_{j-1})$,

(b) generate

$$ X := x_{J-1} + (x_J - x_{J-1}) \cdot \begin{cases} \log\big(1 + (e^\Theta - 1)U\big)/\Theta & \text{if } \Theta \neq 0, \\ U & \text{if } \Theta = 0, \end{cases} $$

where $\Theta := \hat{\varphi}_n(x_J) - \hat{\varphi}_n(x_{J-1})$ and $U \sim \text{Unif}[0, 1]$, and

(c) generate

$$X^* := X + \hat{\gamma}Z \quad \text{with } Z \sim \mathcal{N}(0,1),$$

where J, U and Z are independent. Then $X \sim \hat{f}_n$ and $X^* \sim \hat{f}_n^*$.
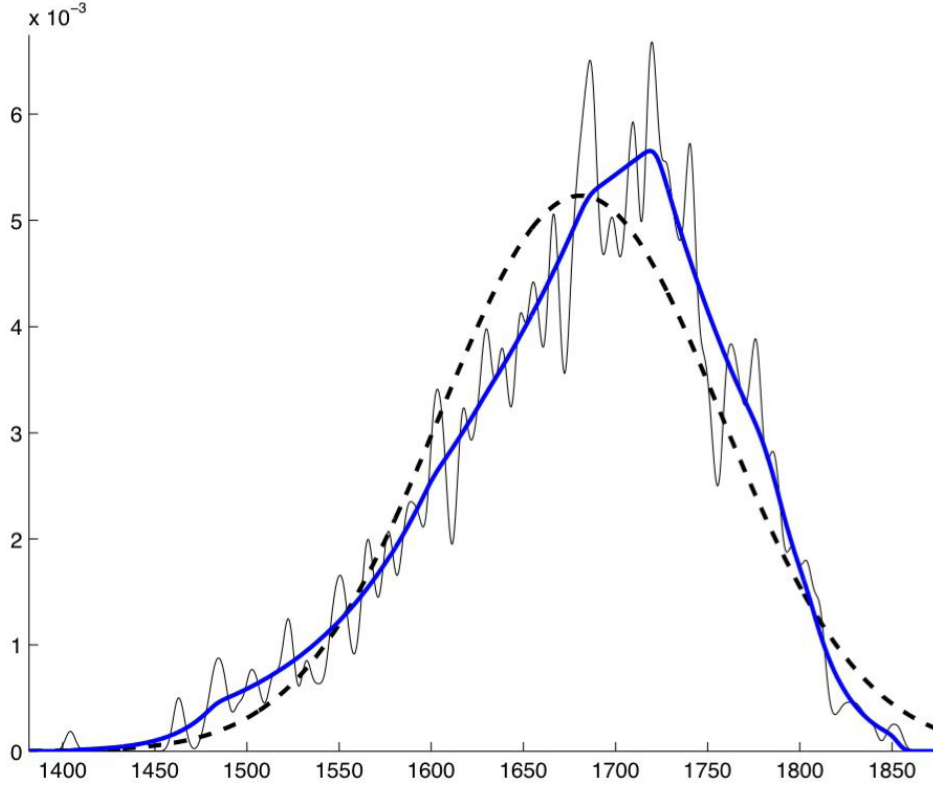

Figure 2: Three competing density estimators.

4 Uniform consistency

Let us introduce some notation. For any integer $n > 1$ we define

$$\rho_n := \log(n)/n,$$

and the uniform norm of a function $g : I \to \mathbb{R}$ on an interval $I \subset \mathbb{R}$ is denoted by

$$\|g\|_\infty^I := \sup_{x \in I} |g(x)|.$$

We say that g belongs to the Hölder class $\mathcal{H}^{\beta,L}(I)$ with exponent $\beta \in [1,2]$ and constant $L > 0$ if for all $x, y \in I$ we have

$$
\begin{aligned}
|g(x) - g(y)| &\leq L|x - y| &&\text{if } \beta = 1, \\
|g'(x) - g'(y)| &\leq L|x - y|^{\beta-1} &&\text{if } \beta > 1.
\end{aligned}
$$

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Uniform consistency of $\hat{\varphi}_n$. Our main result is the following theorem:

Theorem 4.1. *Assume for the log-density $\varphi = \log f$ that $\varphi \in \mathcal{H}^{\beta,L}(T)$ for some exponent $\beta \in [1, 2]$, some constant $L > 0$ and a subinterval $T = [A, B]$ of the interior of $\{f > 0\}$. Then,*

$$\max_{t \in T} (\hat{\varphi}_n - \varphi)(t) = O_{\mathrm{p}}\left(\rho_n^{\beta/(2\beta+1)}\right),$$

$$\max_{t \in T(n,\beta)} (\varphi - \hat{\varphi}_n)(t) = O_{\mathrm{p}}\left(\rho_n^{\beta/(2\beta+1)}\right),$$

where $T(n, \beta) := \left[A + \rho_n^{1/(2\beta+1)}, B - \rho_n^{1/(2\beta+1)}\right]$.

Note that the previous result remains true when we replace $\hat{\varphi}_n - \varphi$ with $\hat{f}_n - f$. It is well-known that the rates of convergence in Theorem 4.1 are optimal, even if β was known (cf. Khas'minskii 1978). Thus our estimators adapt to the unknown smoothness of f in the range $\beta \in [1, 2]$.

Note also that concavity of φ implies that it is Lipschitz-continuous, i.e. belongs to $\mathcal{H}^{1,L}(T)$ for some $L > 0$, on any interval $T = [A, B]$ with $A > \inf\{f > 0\}$ and $B < \sup\{f > 0\}$. Hence one can easily deduce from Theorem 4.1 that $\hat{f}_n$ is consistent in $L_1(\mathbb{R})$ and that $\hat{F}_n$ is uniformly consistent:

Corollary 4.2.

$$\int \left|\hat{f}_n(x) - f(x)\right| \mathrm{d}x \to_{\mathrm{p}} 0 \quad \text{and} \quad \|\hat{F}_n - F\|_\infty^{\mathbb{R}} \to_{\mathrm{p}} 0.$$

Distance of two consecutive knots and uniform consistency of $\hat{F}_n$. By means of Theorem 4.1 we can solve a "gap problem" for log-concave density estimation. The phrase "gap problem" was first used by Balabdaoui and Wellner (2008) to describe the problem of computing the distance between two consecutive knots of certain estimators.

Theorem 4.3. *Suppose that the assumptions of Theorem 4.1 hold. Assume further that $\varphi'(x) - \varphi'(y) \geq C(y - x)$ for some constant $C > 0$ and arbitrary $A \leq x < y \leq B$, where φ' stands for $\varphi'(\cdot -)$ or $\varphi'(\cdot +)$. Then*

$$\sup_{x \in T} \min_{y \in \mathcal{S}_n(\hat{\varphi}_n)} |x - y| = O_{\mathrm{p}}\left(\rho_n^{\beta/(4\beta+2)}\right).$$

Theorems 4.1 and 4.3, combined with a result of Stute (1982) about the modulus of continuity of empirical processes, yield a rate of convergence for the maximal difference between $\hat{F}_n$ and $\mathbb{F}_n$ on compact intervals:

Theorem 4.4. *Under the assumptions of Theorem 4.3,*

$$\max_{t \in T(n,\beta)} \left|\hat{F}_n(t) - \mathbb{F}_n(t)\right| = O_{\mathrm{p}}\left(\rho_n^{3\beta/(4\beta+2)}\right).$$

In particular, if $\beta > 1$, then

$$\max_{t \in T(n,\beta)} \left|\hat{F}_n(t) - \mathbb{F}_n(t)\right| = o_{\mathrm{p}}(n^{-1/2}).$$

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Thus, under certain regularity conditions, the estimators $\hat{F}_n$ and $\mathbb{F}_n$ are asymptotically equivalent on compact sets. Conclusions of this type are known for the Grenander estimator (cf. Kiefer and Wolfowitz 1976) and the least squares estimator of a convex density on $[0, \infty)$ (cf. Balabdaoui and Wellner 2007).

The result of Theorem 4.4 is also related to recent results of Giné and Nickl (2007, 2008). In the latter paper they devise kernel density estimators with data–driven choice of bandwidth which are also adaptive with respect to β in a certain range while the integrated density estimator is asymptotically equivalent to $\mathbb{F}_n$ on the whole real line. However, if $\beta \geq 3/2$, they have to use kernel functions of higher order, i.e. no longer being non-negative, and simulating data from the resulting estimated density is not straightforward.

Example. Let us illustrate Theorems 4.1 and 4.4 with simulated data, again from the Gumbel distribution with $\varphi(x) = -x - e^{-x}$. Here $\varphi''(x) = -e^{-x}$, so the assumptions of our theorems are satisfied with $\beta = 2$ for any compact interval T. The upper panels of Figure 3 show the true log-density φ (dashed line) and the estimator $\hat{\varphi}_n$ (line) for samples of sizes $n = 200$ (left) and $n = 2000$ (right). The lower panels show the corresponding empirical processes $n^{1/2}(\mathbb{F}_n - F)$ (jagged curves) and $n^{1/2}(\hat{F}_n - F)$ (smooth curves). First of all, the quality of the estimator $\hat{\varphi}_n$ is quite good, even in the tails, and the quality increases with sample size, as expected. Looking at the empirical processes, the similarity between $n^{1/2}(\mathbb{F}_n - F)$ and $n^{1/2}(\hat{F}_n - F)$ increases with sample size, too, but rather slowly. Note also that the estimator $\hat{F}_n$ outperforms $\mathbb{F}_n$ in terms of supremum distance from F, which leads us to the next paragraph.

Marshall's Lemma. In all simulations we looked at, the estimator $\hat{F}_n$ satisfied the inequality

$$\|\hat{F}_n - F\|_\infty^{\mathbb{R}} \ \leq \ \|\mathbb{F}_n - F\|_\infty^{\mathbb{R}}, \tag{4}$$

provided that f is indeed log-concave. Figure 3 shows two numerical examples for this phenomenon. In view of such examples and Marshall's (1970) lemma about the Grenander estimator $\hat{F}_n^{\mathrm{mon}}$, we first tried to verify that (4) is correct almost surely and for any $n > 1$. However, one can construct counterexamples showing that (4) may be violated, even if the right hand side is multiplied with any fixed constant $C > 1$. Nevertheless our first attempts resulted in a version of Marshall's lemma for *convex* density estimation; see Dümbgen et al. (2007b). For the present setting, we conjecture that (4) is true with asymptotic probability one as $n \to \infty$, i.e.

$$\mathbb{P}\Big(\|\hat{F}_n - F\|_\infty^{\mathbb{R}} \leq \|\mathbb{F}_n - F\|_\infty^{\mathbb{R}}\Big) \ \to \ 1.$$

A monotone hazard rate estimator. Estimation of a monotone hazard rate is described, for instance, in the book by Robertson et al. (1988). They solve directly an isotonic estimation problem similar to that for the Grenander density estimator. For this setting, Hall et al. (2001) and Hall and van Keilegom (2005) consider methods based upon suitable modifications of kernel estimators.
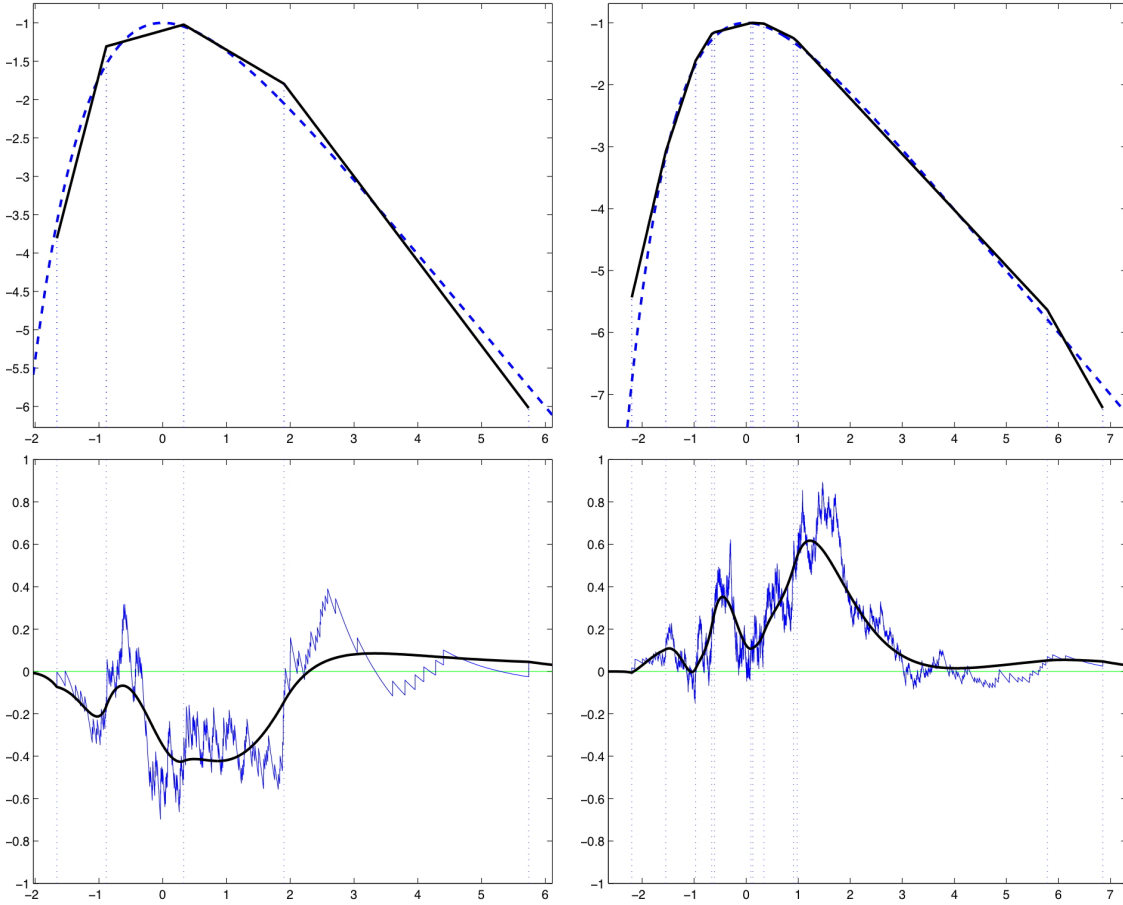
Figure 3: Density functions and empirical processes for Gumbel samples of size $n = 200$ and $n = 2'000$.

Alternatively, in our setting it follows from Lemma A.2 in Section A that

$$\hat{h}_n(x) \; := \; \frac{\hat{f}_n(x)}{1 - \hat{F}_n(x)}$$

defines a simple plug-in estimator of the hazard rate on $(-\infty, X_n)$ which is non-decreasing as well. By virtue of Theorem 4.1 and Corollary 4.2 it is uniformly consistent on any compact subinterval of the interior of $\{f > 0\}$. Theorems 4.1 and 4.4 entail even a rate of convergence:

Corollary 4.5. *Under the assumptions of Theorem 4.3,*

$$\max_{t \in T(n,\beta)} \left| \hat{h}_n(t) - h(t) \right| \; = \; O_{\mathrm{p}}\left(\rho_n^{\beta/(2\beta+1)} \right).$$

5 Estimation of convex or k–monotone densities on $[0, \infty)$

Under assumptions comparable to ours for $\beta = 2$, Groeneboom et al. (2001) proved uniform consistency of the maximum likelihood estimator $\hat{p}_{n,2}$ of a convex density p_2 on $[0, \infty)$ as well

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as a rate of convergence of $O_{\mathrm{p}}(n^{-2/5})$ at a fixed point $x_o > 0$. Using these results, they further provided the limiting distribution of $\hat{p}_{n,2}$ at a fixed point x_o.

Monotone and convex densities are members of the broader class of k–monotone densities. A density function $p_k : [0, \infty) \to [0, \infty)$ is 1–monotone if it is non–increasing. It is 2–monotone if it is non–increasing and convex, and k–monotone for $k \geq 3$ if, and only if, $(-1)^j p_k^{(j)}$ is non–negative, non–increasing, and convex for $j = 0, ..., k-2$. Balabdaoui and Wellner (2008) generalized the results of Groeneboom et al. (2001) to these k–monotone densities. However, only by assuming that a so far unverified conjecture about the upper bound on the error in a particular Hermite interpolation via odd–degree splines holds true.

Similarly to $\hat{\varphi}_n$, the maximum likelihood estimators $\hat{p}_{n,k}$ of p_k are splines of order $k-1$. However, for any $k > 1$ the knots of $\hat{p}_{n,k}$ fall strictly between observations, with probability equal to one. This property makes it considerably difficult to obtain a result analogous to Theorem 4.3.

Remarkably, the characterization of $\hat{f}_n$ in Theorem 2.4 by means of integrated distribution functions coincides with that of the least squares estimator of a convex density on $[0, \infty)$, see Lemma 2.2 of Groeneboom et al. (2001). This turns out to be crucial in finding the limiting distribution of $n^{\ell/(2\ell+1)}(\hat{f}_n(x_o) - f(x_o))$ for any $x_o \in \mathbb{R}$, see Balabdaoui et al. (2008). Here, ℓ indicates the first non–vanishing higher order derivative of φ at x_o. That means, $\ell = 2$ if $\varphi''(x_o) \neq 0$. Otherwise, $\ell \geq 4$ is the smallest even integer such that $\varphi^{(j)}(x_o) = 0$ for $2 \leq j < \ell$ while $\varphi^{(\ell)}(x_o) \neq 0$.

6   Outlook

Starting from the results presented here, Balabdaoui et al. (2008) derived recently the pointwise limiting distribution of $\hat{f}_n$. They also consider the limiting distribution of $\mathrm{argmax}_{x \in \mathbb{R}} \hat{f}_n(x)$ as an estimator of the mode of f. Empirical findings of Müller and Rufibach (2006) show that the estimator $\hat{f}_n$ is even useful for extreme value statistics. Log-concave densities have also potential as building blocks in more complex models (e.g. regression or classification) or when handling censored data (cf. Dümbgen et al. 2007a).

Unfortunately, our proofs work only for fixed compact intervals, whereas simulations suggest that the estimators perform well on the whole real line. Right now the authors are working on a different approach where $\hat{\varphi}_n$ is represented locally as a parametric maximum likelihood estimator of a log-linear density. Presumably this will deepen our understanding of the log-concave NPMLE's consistency properties, particularly in the tails. For instance, we conjecture that $\mathbb{F}_n$ and $\hat{F}_n$ are asymptotically equivalent on any interval T on which φ' is strictly decreasing.

A   Auxiliary results and proofs

A.1   Two facts about log-concave densities

The following two results about a log-concave density $f = \exp \varphi$ and its distribution function F are of independent interest. The first result entails that the density f has at least subexponential tails:

Lemma A.1. *For arbitrary points $x_1 < x_2$,*

$$\sqrt{f(x_1)f(x_2)} \leq \frac{F(x_2) - F(x_1)}{x_2 - x_1}.$$

Moreover, for $x_o \in \{f > 0\}$ and any real $x \neq x_o$,

$$\frac{f(x)}{f(x_o)} \leq \begin{cases} \left(\dfrac{h(x_o, x)}{f(x_o)|x - x_o|}\right)^2, \\ \exp\left(1 - \dfrac{f(x_o)|x - x_o|}{h(x_o, x)}\right) & \text{if } f(x_o)|x - x_o| \geq h(x_o, x), \end{cases}$$

where

$$h(x_o, x) := F(\max(x_o, x)) - F(\min(x_o, x)) \leq \begin{cases} F(x_o) & \text{if } x < x_o, \\ 1 - F(x_o) & \text{if } x > x_o. \end{cases}$$

A second well-known result (Barlow and Proschan 1975, Lemma 5.8), provides further connections between the density f and the distribution function F. In particular, it entails that $f/(F(1 - F))$ is bounded away from zero on $\{x : 0 < F(x) < 1\}$.

Lemma A.2. *The function f/F is non-increasing on $\{x : 0 < F(x) \leq 1\}$, and the function $f/(1 - F)$ is non-decreasing on $\{x : 0 \leq F(x) < 1\}$.*

Proof of Lemma A.1.   To prove the first inequality, it suffices to consider the nontrivial case of $x_1, x_2 \in \{f > 0\}$. Then concavity of φ entails that

$$\begin{aligned}
F(x_2) - F(x_1) &\geq \int_{x_1}^{x_2} \exp\left(\frac{x_2 - t}{x_2 - x_1}\varphi(x_1) + \frac{t - x_1}{x_2 - x_1}\varphi(x_2)\right) dt \\
&= (x_2 - x_1)\int_0^1 \exp\left((1 - u)\varphi(x_1) + u\varphi(x_2)\right) du \\
&\geq (x_2 - x_1)\exp\left(\int_0^1 \left((1 - u)\varphi(x_1) + u\varphi(x_2)\right) du\right) \\
&= (x_2 - x_1)\exp\left(\varphi(x_1)/2 + \varphi(x_2)/2\right) \\
&= (x_2 - x_1)\sqrt{f(x_1)f(x_2)},
\end{aligned}$$

where the second inequality follows from Jensen's inequality.

We prove the second asserted inequality only for $x > x_o$, i.e. $h(x_o, x) = F(x) - F(x_o)$, the other case being handled analogously. The first part entails that

$$\frac{f(x)}{f(x_o)} \leq \left(\frac{h(x_o, x)}{f(x_o)(x - x_o)}\right)^2,$$

and the right hand side is not greater than one if $f(x_o)(x - x_o) \geq h(x_o, x)$. In the latter case, recall that

$$h(x_o, x) \geq (x - x_o) \int_0^1 \exp\big((1 - u)\varphi(x_o) + u\varphi(x)\big) \, du = f(x_o)(x - x_o) J\big(\varphi(x) - \varphi(x_o)\big)$$

with $\varphi(x) - \varphi(x_o) \leq 0$, where $J(y) := \int_0^1 \exp(uy) \, du$. Elementary calculations show that $J(-r) = (1 - e^{-r})/r \geq 1/(1 + r)$ for arbitrary $r > 0$. Thus

$$h(x_o, x) \geq \frac{f(x_o)(x - x_o)}{1 + \varphi(x_o) - \varphi(x)},$$

which is equivalent to $f(x)/f(x_o) \leq \exp\big(1 - f(x_o)(x - x_o)/h(x_o, x)\big)$. $\qquad\square$

A.2 Proofs of the characterizations

Proof of Theorem 2.2. In view of Theorem 2.1 we may restrict our attention to concave and real-valued functions φ on $[X_1, X_n]$ and set $\varphi := -\infty$ on $\mathbb{R} \setminus [X_1, X_n]$. The set $\mathcal{C}_n$ of all such functions is a convex cone, and for any function $\Delta : \mathbb{R} \to \mathbb{R}$ and $t > 0$, concavity of $\varphi + t\Delta$ on $\mathbb{R}$ is equivalent to its concavity on $[X_1, X_n]$.

One can easily verify that Ψ_n is a concave and real-valued functional on $\mathcal{C}_n$. Hence, as well-known from convex analysis, a function $\widetilde{\varphi} \in \mathcal{C}_n$ maximizes Ψ_n if, and only if,

$$\lim_{t \downarrow 0} \frac{\Psi_n(\widetilde{\varphi} + t(\varphi - \widetilde{\varphi})) - \Psi_n(\widetilde{\varphi})}{t} \leq 0$$

for all $\varphi \in \mathcal{C}_n$. But this is equivalent to the requirement that

$$\lim_{t \downarrow 0} \frac{\Psi_n(\widetilde{\varphi} + t\Delta) - \Psi_n(\widetilde{\varphi})}{t} \leq 0$$

for any function $\Delta : \mathbb{R} \to \mathbb{R}$ such that $\widetilde{\varphi} + \lambda\Delta$ is concave for some $\lambda > 0$. Now the assertion of the theorem follows from

$$\lim_{t \downarrow 0} \frac{\Psi_n(\widetilde{\varphi} + t\Delta) - \Psi_n(\widetilde{\varphi})}{t} = \int \Delta \, d\mathbb{F}_n - \int \Delta(x) \exp \widetilde{\varphi}(x) \, dx. \qquad\square$$

Proof of Theorem 2.4. We start with a general observation. Let G be some distribution (function) with support $[X_1, X_n]$, and let $\Delta : [X_1, X_n] \to \mathbb{R}$ be absolutely continuous with L_1–derivative Δ'. Then it follows from Fubini's theorem that

$$\int \Delta \, dG = \Delta(X_n) - \int_{X_1}^{X_n} \Delta'(r) G(r) \, dr. \tag{5}$$

Now suppose that $\widetilde{\varphi} = \hat{\varphi}_n$, and let $t \in (X_1, X_n]$. Let Δ be absolutely continuous on $[X_1, X_n]$ with L_1–derivative $\Delta'(r) = 1\{r \leq t\}$ and arbitrary value of $\Delta(X_n)$. Clearly, $\widetilde{\varphi} + \Delta$ is concave, whence (2) and (5) entail that

$$\Delta(X_n) - \int_{X_1}^t \mathbb{F}_n(r)\,\mathrm{d}r \;\leq\; \Delta(X_n) - \int_{X_1}^t \widetilde{F}(r)\,\mathrm{d}r,$$

which is equivalent to inequality (3). In case of $t \in \mathcal{S}_n(\widetilde{\varphi}) \setminus \{X_1\}$, let $\Delta'(r) = -1\{r \leq t\}$. Then $\widetilde{\varphi} + \lambda\Delta$ is concave for some $\lambda > 0$, so that

$$\Delta(X_n) + \int_{X_1}^t \mathbb{F}_n(r)\,\mathrm{d}r \;\leq\; \Delta(X_n) + \int_{X_1}^t \widetilde{F}(r)\,\mathrm{d}r,$$

which yields equality in (3).

Now suppose that $\widetilde{\varphi}$ satisfies inequality (3) for all t with equality if $t \in \mathcal{S}_n(\widetilde{\varphi})$. In view of Theorem 2.1 and the proof of Theorem 2.2, it suffices to show that (2) holds for any function Δ defined on $[X_1, X_n]$ which is linear on each interval $[X_j, X_{j+1}]$, $1 \leq j < n$, while $\widetilde{\varphi} + \lambda\Delta$ is concave for some $\lambda > 0$. The latter requirement is equivalent to Δ being concave between two consecutive knots of $\widetilde{\varphi}$. Elementary considerations show that the L_1–derivative of such a function Δ may be written as

$$\Delta'(r) \;=\; \sum_{j=2}^n \beta_j 1\{r \leq X_j\}$$

with real numbers $\beta_2, \ldots, \beta_n$ such that

$$\beta_j \;\geq\; 0 \quad \text{if } X_j \notin \mathcal{S}_n(\widetilde{\varphi}).$$

Consequently, it follows from (5) and our assumptions on $\widetilde{\varphi}$ that

$$\begin{aligned}
\int \Delta\,\mathrm{d}\mathbb{F}_n &\;=\; \Delta(X_n) - \sum_{j=2}^n \beta_j \int_{X_1}^{X_j} \mathbb{F}_n(r)\,\mathrm{d}r \\
&\;\leq\; \Delta(X_n) - \sum_{j=2}^n \beta_j \int_{X_1}^{X_j} \widetilde{F}(r)\,\mathrm{d}r \\
&\;=\; \int \Delta\,\mathrm{d}\widetilde{F}. \qquad \square
\end{aligned}$$

Proof of Corollary 2.5. For $t \in \mathcal{S}_n(\hat{\varphi}_n)$ and $s < t < u$, it follows from Theorem 2.4 that

$$\frac{1}{u-t}\int_t^u \hat{F}_n(r)\,\mathrm{d}r \;\leq\; \frac{1}{u-t}\int_t^u \mathbb{F}_n(r)\,\mathrm{d}r \quad \text{and}$$

$$\frac{1}{t-s}\int_s^t \hat{F}_n(r)\,\mathrm{d}r \;\geq\; \frac{1}{t-s}\int_s^t \mathbb{F}_n(r)\,\mathrm{d}r.$$

Letting $u \downarrow t$ and $s \uparrow t$ yields

$$\hat{F}_n(t) \;\leq\; \mathbb{F}_n(t) \quad \text{and} \quad \hat{F}_n(t) \;\geq\; \mathbb{F}_n(t-) = \mathbb{F}_n(t) - n^{-1}. \qquad \square$$

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A.3 Proof of $\hat{\varphi}_n$'s consistency

Our proof of Theorem 4.1 is a refinement and modification of methods introduced by Dümbgen et al. (2004). A first key ingredient is an inequality for concave functions due to Dümbgen (1998) (see also Dümbgen et al. 2004 or Rufibach 2006):

Lemma A.3. *For any $\beta \in [1,2]$ and $L > 0$ there exists a constant $K = K(\beta, L) \in (0,1]$ with the following property: Suppose that g and $\hat{g}$ are concave and real-valued functions on a compact interval $T = [A, B]$, where $g \in \mathcal{H}^{\beta,L}(T)$. Let $\epsilon > 0$ and $0 < \delta \leq K \min\{B - A, \epsilon^{1/\beta}\}$. Then*

$$\sup_{t \in T} (\hat{g} - g) \geq \epsilon \quad \text{or} \quad \sup_{t \in [A+\delta, B-\delta]} (g - \hat{g}) \geq \epsilon$$

implies that

$$\inf_{t \in [c, c+\delta]} (\hat{g} - g)(t) \geq \epsilon/4 \quad \text{or} \quad \inf_{t \in [c, c+\delta]} (g - \hat{g})(t) \geq \epsilon/4$$

for some $c \in [A, B - \delta]$.

Starting from this lemma, let us first sketch the idea of our proof of Theorem 4.1: Suppose we had a family $\mathcal{D}$ of measurable functions Δ with finite seminorm

$$\sigma(\Delta) := \left(\int \Delta^2 \, dF \right)^{1/2},$$

such that

$$\sup_{\Delta \in \mathcal{D}} \frac{\left| \int \Delta \, d(\mathbb{F}_n - F) \right|}{\sigma(\Delta) \rho_n^{1/2}} \leq C \tag{6}$$

with asymptotic probability one, where $C > 0$ is some constant. If, in addition, $\varphi - \hat{\varphi}_n \in \mathcal{D}$ and $\varphi - \hat{\varphi}_n \leq C$ with asymptotic probability one, then we could conclude that

$$\left| \int (\varphi - \hat{\varphi}_n) \, d(\mathbb{F}_n - F) \right| \leq C\sigma(\varphi - \hat{\varphi}_n) \rho_n^{1/2},$$

while Theorem 2.2, applied to $\Delta := \varphi - \hat{\varphi}_n$, entails that

$$\begin{aligned}
\int (\varphi - \hat{\varphi}_n) \, d(\mathbb{F}_n - F) &\leq \int (\varphi - \hat{\varphi}_n) \, d(\hat{F} - F) \\
&= -\int \Delta \left(1 - \exp(-\Delta) \right) dF \\
&\leq -(1 + C)^{-1} \int \Delta^2 \, dF \\
&= -(1 + C)^{-1} \sigma(\varphi - \hat{\varphi}_n)^2,
\end{aligned}$$

because $y(1 - \exp(-y)) \geq (1 + y_+)^{-1} y^2$ for all real y, where $y_+ := \max(y, 0)$. Hence with asymptotic probability one,

$$\sigma(\varphi - \hat{\varphi}_n)^2 \leq C^2 (1 + C)^2 \rho_n.$$

Now suppose that $|\varphi - \hat{\varphi}_n| \geq \epsilon_n$ on a subinterval of $T = [A, B]$ of length $\epsilon_n^{1/\beta}$, where $(\epsilon_n)_n$ is a fixed sequence of numbers $\epsilon_n > 0$ tending to zero. Then $\sigma(\varphi - \hat{\varphi}_n)^2 \geq \epsilon_n^{(2\beta+1)/\beta} \min_T(f)$, so that

$$\epsilon_n \leq \widetilde{C} \rho_n^{2\beta/(2\beta+1)}$$

with $\widetilde{C} = \left(C^2(1 + C)^2 / \min_T(f)\right)^{\beta/(2\beta+1)}$.

The previous considerations will be modified in two aspects to get a rigorous proof of Theorem 4.1: For technical reasons we have to replace the denominator $\sigma(\Delta)\rho_n^{1/2}$ of inequality (6) with $\sigma(\Delta)\rho_n^{1/2} + W(\Delta)\rho_n^{2/3}$, where

$$W(\Delta) := \sup_{x \in \mathbb{R}} \frac{|\Delta(x)|}{\max\left(1, |\varphi(x)|\right)}.$$

This is necessary to deal with functions Δ with small values of $F(\{\Delta \neq 0\})$. Moreover, we shall work with simple "caricatures" of $\varphi - \hat{\varphi}_n$, namely, functions which are piecewise linear with at most three knots. Throughout this section, piecewise linearity does *not* necessarily imply continuity. A function being piecewise linear with at most m knots means that the real line may be partitioned into $m + 1$ nondegenerate intervals on each of which the function is linear. Then the m real boundary points of these intervals are the knots.

The next lemma extends inequality (2) to certain piecewise linear functions:

Lemma A.4. *Let $\Delta : \mathbb{R} \to \mathbb{R}$ be piecewise linear such that each knot q of Δ satisfies one of the following two properties:*

$$q \in \mathcal{S}_n(\hat{\varphi}_n) \quad and \quad \Delta(q) = \liminf_{x \to q} \Delta(x), \tag{7}$$

$$\Delta(q) = \lim_{r \to q} \Delta(r) \quad and \quad \Delta'(q-) \geq \Delta'(q+). \tag{8}$$

Then

$$\int \Delta \, d\mathbb{F}_n \leq \int \Delta \, d\hat{F}_n. \tag{9}$$

Now we can specify the "caricatures" mentioned before:

Lemma A.5. *Let $T = [A, B]$ be a fixed subinterval of the interior of $\{f > 0\}$. Let $\varphi - \hat{\varphi}_n \geq \epsilon$ or $\hat{\varphi}_n - \varphi \geq \epsilon$ on some interval $[c, c + \delta] \subset T$ with length $\delta > 0$, and suppose that $X_1 < c$ and $X_n > c + \delta$. Then there exists a piecewise linear function Δ with at most three knots each of which satisfies condition (7) or (8) and a positive constant $K' = K'(f, T)$ such that*

$$|\varphi - \hat{\varphi}_n| \geq \epsilon|\Delta|, \tag{10}$$

$$\Delta(\varphi - \hat{\varphi}_n) \geq 0, \tag{11}$$

$$\Delta \leq 1, \tag{12}$$

$$\int_c^{c+\delta} \Delta^2(x) \, dx \geq \delta/3, \tag{13}$$

$$W(\Delta) \leq K'\delta^{-1/2}\sigma(\Delta). \tag{14}$$

Our last ingredient is a surrogate for (6):

Lemma A.6. *Let $\mathcal{D}_m$ be the family of all piecewise linear functions on $\mathbb{R}$ with at most m knots. There exists a constant $K'' = K''(f)$ such that*

$$\sup_{m \geq 1, \Delta \in \mathcal{D}_m} \frac{\left| \int \Delta \, \mathrm{d}(\mathbb{F}_n - F) \right|}{\sigma(\Delta) m^{1/2} \rho_n^{1/2} + W(\Delta) m \rho_n^{2/3}} \leq K''$$

with probability tending to one as $n \to \infty$.

Before we verify all these auxiliary results, let us proceed with the main proof.

Proof of Theorem 4.1. Suppose that

$$\sup_{t \in T} (\hat{\varphi}_n - \varphi)(t) \geq C \epsilon_n$$

or

$$\sup_{t \in [A + \delta_n, B - \delta_n]} (\varphi - \hat{\varphi}_n)(t) \geq C \epsilon_n$$

for some constant $C > 0$, where $\epsilon_n := \rho_n^{\beta/(2\beta+1)}$ and $\delta_n := \rho_n^{1/(2\beta+1)} = \epsilon_n^{1/\beta}$. It follows from Lemma A.3 with $\epsilon := C\epsilon_n$ that in case of $C \geq K^{-\beta}$ and for sufficiently large n, there is a (random) interval $[c_n, c_n + \delta_n] \subset T$ on which either $\hat{\varphi}_n - \varphi \geq (C/4)\epsilon_n$ or $\varphi - \hat{\varphi}_n \geq (C/4)\epsilon_n$. But then there is a (random) function $\Delta_n \in \mathcal{D}_3$ fulfilling the conditions stated in Lemma A.5. For this Δ_n it follows from (9) that

$$\int_{\mathbb{R}} \Delta_n \, \mathrm{d}(F - \mathbb{F}_n) \geq \int_{\mathbb{R}} \Delta_n \, \mathrm{d}(F - \hat{F}_n) = \int_{\mathbb{R}} \Delta_n \big(1 - \exp\big[-(\varphi - \hat{\varphi}_n)\big]\big) \, \mathrm{d}F. \quad (15)$$

With $\widetilde{\Delta}_n := (C/4)\epsilon_n \Delta_n$, it follows from (10–11) that the right hand side of (15) is not smaller than

$$(4/C)\epsilon_n^{-1} \int \widetilde{\Delta}_n \big(1 - \exp(-\widetilde{\Delta}_n)\big) \, \mathrm{d}F \geq \frac{(4/C)\epsilon_n^{-1}}{1 + (C/4)\epsilon_n} \sigma(\widetilde{\Delta}_n)^2 = \frac{(C/4)\epsilon_n}{1 + o(1)} \sigma(\Delta_n)^2,$$

because $\widetilde{\Delta}_n \leq (C/4)\epsilon_n$ by (12). On the other hand, according to Lemma A.6 we may assume that

$$\begin{aligned}
\int_{\mathbb{R}} \Delta_n \, \mathrm{d}(F - \mathbb{F}_n) &\leq K''\big(3^{1/2}\sigma(\Delta_n)\rho_n^{1/2} + 3W(\Delta_n)\rho_n^{2/3}\big) \\
&\leq K''(3^{1/2}\rho_n^{1/2} + 3K'\delta_n^{-1/2}\rho_n^{2/3})\sigma(\Delta_n) \quad \text{(by (14))} \\
&\leq K''(3^{1/2}\rho_n^{1/2} + 3K'\rho_n^{2/3 - 1/(4\beta+2)})\sigma(\Delta_n) \\
&\leq G\rho_n^{1/2}\sigma(\Delta_n)
\end{aligned}$$

for some constant $G = G(\beta, L, f, T)$, because $2/3 - 1/(4\beta + 2) \geq 2/3 - 1/6 = 1/2$. Consequently,

$$C^2 \leq \frac{16G^2(1 + o(1))\epsilon_n^{-2}\rho_n}{\sigma(\Delta_n)^2} = \frac{16G^2(1 + o(1))}{\delta_n^{-1}\sigma(\Delta_n)^2} \leq \frac{48G^2(1 + o(1))}{\min_T(f)},$$

where the last inequality follows from (13). $\qquad \square$

Proof of Lemma A.4. There is a sequence of continuous, piecewise linear functions Δ_k converging pointwise isotonically to Δ as $k \to \infty$ such that any knot q of Δ_k either belongs to $\mathcal{S}_n(\hat\varphi_n)$, or $\Delta'_k(q-) > \Delta'_k(q+)$. Thus $\hat\varphi_n + \lambda\Delta_k$ is concave for sufficiently small $\lambda > 0$. Consequently, since $\Delta_1 \leq \Delta_k \leq \Delta$ for all k, it follows from dominated convergence and (2) that

$$\int \Delta \, d\mathbb{F}_n \;=\; \lim_{k\to\infty} \int \Delta_k \, d\mathbb{F}_n \;\leq\; \lim_{k\to\infty} \int \Delta_k \, d\hat{F}_n \;=\; \int \Delta \, d\hat{F}_n. \qquad \square$$

Proof of Lemma A.5. The crucial point in all the cases we have to distinguish is to construct a $\Delta \in \mathcal{D}_3$ satisfying the assumptions of Lemma A.4 and (10–13). Recall that $\hat\varphi_n$ is piecewise linear.

Case 1a: $\hat\varphi_n - \varphi \geq \epsilon$ **on** $[c, c+\delta]$ **and** $\mathcal{S}_n(\hat\varphi_n) \cap (c, c+\delta) \neq \emptyset$. Here we choose a continuous function $\Delta \in \mathcal{D}_3$ with knots c, $c+\delta$ and $x_o \in \mathcal{S}_n(\hat\varphi_n) \cap (c, c+\delta)$, where $\Delta := 0$ on $(-\infty, c] \cup [c+\delta, \infty)$ and $\Delta(x_o) := -1$. Here the assumptions of Lemma A.4 and requirements (10–13) are easily verified.

Case 1b: $\hat\varphi_n - \varphi \geq \epsilon$ **on** $[c, c+\delta]$ **and** $\mathcal{S}_n(\hat\varphi_n) \cap (c, c+\delta) = \emptyset$. Let $[c_o, d_o] \supset [c, c+\delta]$ be the maximal interval on which $\varphi - \hat\varphi_n$ is concave. Then there exists a linear function $\widetilde{\Delta}$ such that $\widetilde{\Delta} \geq \varphi - \hat\varphi_n$ on $[c_o, d_o]$ and $\widetilde{\Delta} \leq -\epsilon$ on $[c, c+\delta]$. Next let $(c_1, d_1) := \{\widetilde{\Delta} < 0\} \cap (c_o, d_o)$. Now we define $\Delta \in \mathcal{D}_2$ via

$$\Delta(x) \;:=\; \begin{cases} 0 & \text{if } x \in (-\infty, c_1) \cup (d_1, \infty), \\ \widetilde{\Delta}/\epsilon & \text{if } x \in [c_1, d_1]. \end{cases}$$

Again the assumptions of Lemma A.4 and requirements (10–13) are easily verified; this time we even know that $\Delta \leq -1$ on $[c, c+\delta]$, whence $\int_c^{c+\delta} \Delta(x)^2 \, dx \geq \delta$. Figure 4 illustrates this construction.

Case 2: $\varphi - \hat\varphi_n \geq \epsilon$ **on** $[c, c+\delta]$. Let $[c_o, c]$ and $[c+\delta, d_o]$ be maximal intervals on which $\hat\varphi_n$ is linear. Then we define

$$\Delta(x) \;:=\; \begin{cases} 0 & \text{if } x \in (-\infty, c_o) \cup (d_o, \infty), \\ 1 + \beta_1(x - x_o) & \text{if } x \in [c_o, x_o], \\ 1 + \beta_2(x - x_o) & \text{if } x \in [x_o, d_o], \end{cases}$$

where $x_o := c + \delta/2$ and $\beta_1 \geq 0$ is chosen such that either

$$\Delta(c_o) = 0 \quad \text{and} \quad (\varphi - \hat\varphi_n)(c_o) \geq 0 \quad \text{or}$$
$$(\varphi - \hat\varphi_n)(c_o) < 0 \quad \text{and} \quad \text{sign}(\Delta) = \text{sign}(\varphi - \hat\varphi_n) \text{ on } [c_o, x_o].$$

Analogously, $\beta_2 \leq 0$ is chosen such that

$$\Delta(d_o) = 0 \quad \text{and} \quad (\varphi - \hat\varphi_n)(d_o) \geq 0 \quad \text{or}$$
$$(\varphi - \hat\varphi_n)(d_o) < 0 \quad \text{and} \quad \text{sign}(\Delta) = \text{sign}(\varphi - \hat\varphi_n) \text{ on } [x_o, d_o].$$

Again the assumptions of Lemma A.4 and requirements (10–13) are verified easily. Figure 5 depicts an example.

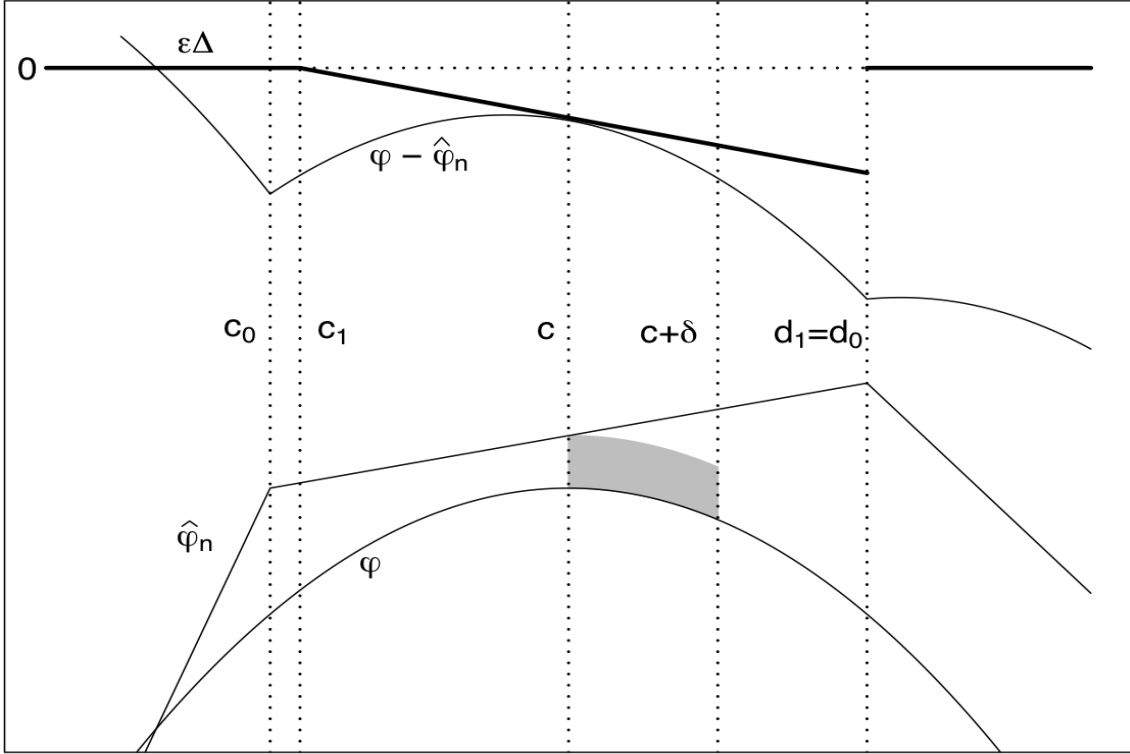Figure 4: The perturbation function Δ in Case 1b.

It remains to verify requirement (14) for our particular functions Δ. Note that by our assumption on $T = [A, B]$, there exist numbers $\tau, C_o > 0$ such that $f \geq C_o$ on $T_o := [A - \tau, B + \tau]$.

In Case 1a, $W(\Delta) \leq \|\Delta\|_\infty^{\mathbb{R}} = 1$, whereas $\sigma(\Delta)^2 \geq C_o \int_c^{c+\delta} \Delta(x)^2 \, \mathrm{d}x = C_o \delta^2/3$. Hence (14) is met if $K' \geq (3/C_o)^{1/2}$.

For Cases 1b and 2 we start with a more general consideration: Let $h(x) := 1\{x \in Q\}(\alpha + \gamma x)$ for real numbers α, γ and a nondegenerate interval Q containing some point in $(c, c+\delta)$. Let $Q \cap T_o$ have endpoints $x_o < y_o$. Then elementary considerations reveal that

$$\sigma(h)^2 \geq C_o \int_{x_o}^{y_o} (\alpha + \gamma x)^2 \, \mathrm{d}x \geq \frac{C_o}{4}(y_o - x_o)\left(\|h\|_\infty^{T_o}\right)^2.$$

Now we deduce an upper bound for $W(h)/\|h\|_\infty^{T_o}$. If $Q \subset T_o$ or $\gamma = 0$, then $W(h)/\|h\|_\infty^{T_o} \leq 1$. Now suppose that $\gamma \neq 0$ and $Q \not\subset T_o$. Then $x_o, y_o \in T_o$ satisfy $y_o - x_o \geq \tau$, and without loss of generality let $\gamma = -1$. Now

$$
\begin{aligned}
\|h\|_\infty^{T_o} &= \max\left(|\alpha - x_o|, |\alpha - y_o|\right) \\
&= (y_o - x_o)/2 + |\alpha - (x_o + y_o)/2| \\
&\geq \tau/2 + |\alpha - (x_o + y_o)/2|.
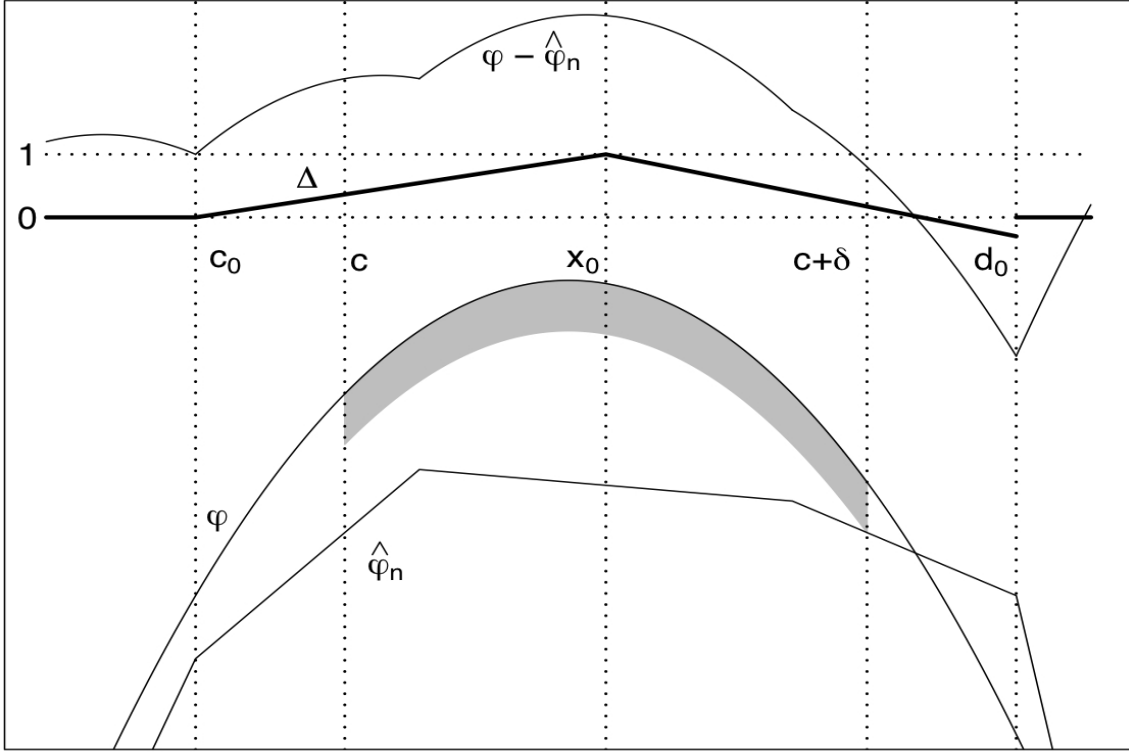\end{aligned}
$$

Figure 5: The perturbation function Δ in Case 2.

On the other hand, since $\varphi(x) \leq a_o - b_o|x|$ for certain constants $a_o, b_o > 0$,

$$
\begin{aligned}
W(h) &\leq \sup_{x \in \mathbb{R}} \frac{|\alpha - x|}{\max(1, b_o|x| - a_o)} \\
&\leq \sup_{x \in \mathbb{R}} \frac{|\alpha| + |x|}{\max(1, b_o|x| - a_o)} \\
&= |\alpha| + (a_o + 1)/b_o \\
&\leq |\alpha - (x_o + y_o)/2| + (|A| + |B| + \tau)/2 + (a_o + 1)/b_o.
\end{aligned}
$$

This entails that

$$
\frac{W(h)}{\|h\|_\infty^{T_o}} \leq C_* := \frac{(|A| + |B| + \tau)/2 + (a_o + 1)/b_o}{\tau/2}.
$$

In Case 1b, our function Δ is of the same type as h above, and $y_o - x_o \geq \delta$. Thus

$$
W(\Delta) \leq C_* \|h\|_\infty^{T_o} \leq 2C_* C_o^{-1/2} \delta^{-1/2} \sigma(\Delta).
$$

In Case 2, Δ may be written as $h_1 + h_2$, with two functions h_1 and h_2 of the same type as h above having disjoint support and both satisfying $y_o - x_o \geq \delta/2$. Thus

$$
\begin{aligned}
W(\Delta) &= \max\big(W(h_1), W(h_2)\big) \\
&\leq 2^{3/2} C_* C_o^{-1/2} \delta^{-1/2} \max\big(\sigma(h_1), \sigma(h_2)\big) \\
&\leq 2^{3/2} C_* C_o^{-1/2} \delta^{-1/2} \sigma(\Delta). \qquad \square
\end{aligned}
$$

To prove Lemma A.6, we need a simple exponential inequality:

Lemma A.7. *Let Y be a random variable such that $\mathbb{E}(Y) = 0$, $\mathbb{E}(Y^2) = \sigma^2$ and $C := \mathbb{E}\exp(|Y|) < \infty$. Then for arbitrary $t \in \mathbb{R}$,*

$$\mathbb{E}\exp(tY) \leq 1 + \frac{\sigma^2 t^2}{2} + \frac{C|t|^3}{(1 - |t|)_+}.$$

Proof of Lemma A.7.

$$\mathbb{E}\exp(tY) = \sum_{k=0}^{\infty} \frac{t^k}{k!} \mathbb{E}(Y^k) \leq 1 + \frac{\sigma^2 t^2}{2} + \sum_{k=3}^{\infty} \frac{|t|^k}{k!} \mathbb{E}(|Y|^k).$$

For any $y \geq 0$ and integers $k \geq 3$, $y^k e^{-y} \leq k^k e^{-k}$. Thus $\mathbb{E}(|Y|^k) \leq \mathbb{E}\exp(|Y|)k^k e^{-k} = Ck^k e^{-k}$. Since $k^k e^{-k} \leq k!$, which can be verified easily via induction on k,

$$\sum_{k=3}^{\infty} \frac{|t|^k}{k!} \mathbb{E}(|Y|^k) \leq C\sum_{k=3}^{\infty} |t|^k = \frac{C|t|^3}{(1 - |t|)_+}. \qquad \square$$

Lemma A.7 entails the following result for finite families of functions:

Lemma A.8. *Let $\mathcal{H}_n$ be a finite family of functions h with $0 < W(h) < \infty$ such that $\#\mathcal{H}_n = O(n^p)$ for some $p > 0$. Then for sufficiently large D,*

$$\lim_{n\to\infty} \mathbb{P}\left(\max_{h\in\mathcal{H}_n} \frac{\left|\int h\,d(\mathbb{F}_n - F)\right|}{\sigma(h)\rho_n^{1/2} + W(h)\rho_n^{2/3}} \geq D\right) = 0.$$

Proof of Lemma A.8. Since $W(ch) = cW(h)$ and $\sigma(ch) = c\sigma(h)$ for any $h \in \mathcal{H}_n$ and arbitrary constants $c > 0$, we may assume without loss of generality that $W(h) = 1$ for all $h \in \mathcal{H}_n$. Let X be a random variable with log–density φ. Since

$$\limsup_{|x|\to\infty} \frac{\varphi(x)}{|x|} < 0$$

by Lemma A.1, the expectation of $\exp(t_o w(X))$ is finite for any fixed $t_o \in (0,1)$, where $w(x) := \max(1, |\varphi(x)|)$. Hence

$$\mathbb{E}\exp\big(t_o|h(X) - \mathbb{E}\,h(X)|\big) \leq C_o := \exp(t_o\,\mathbb{E}\,w(X))\,\mathbb{E}\exp(t_o w(X)) < \infty.$$

Lemma A.7, applied to $Y := t_o(h(X) - \mathbb{E}\,h(X))$, implies that

$$\mathbb{E}\exp\big[t\big(h(X) - \mathbb{E}\,h(X)\big)\big] = \mathbb{E}\big((t/t_o)Y\big) \leq 1 + \frac{\sigma(h)^2 t^2}{2} + \frac{C_1|t|^3}{(1 - C_2|t|)_+}$$

for arbitrary $h \in \mathcal{H}_n$, $t \in \mathbb{R}$ and constants C_1, C_2 depending on t_o and C_o. Consequently,

$$
\begin{aligned}
\mathbb{E}\exp\left(t\int h\,\mathrm{d}(\mathbb{F}_n - F)\right) &= \mathbb{E}\exp\left((t/n)\sum_{i=1}^{n}(h(X_i) - \mathbb{E}\,h(X))\right) \\
&= \left(\mathbb{E}\exp\left((t/n)(h(X) - \mathbb{E}\,h(X))\right)\right)^n \\
&\leq \left(1 + \frac{\sigma(h)^2 t^2}{2n^2} + \frac{C_1|t|^3}{n^3(1 - C_2|t|/n)_+}\right)^n \\
&\leq \exp\left(\frac{\sigma(h)^2 t^2}{2n} + \frac{C_1|t|^3}{n^2(1 - C_2|t|/n)_+}\right).
\end{aligned}
$$

Now it follows from Markov's inequality that

$$
\mathbb{P}\left(\left|\int h\,\mathrm{d}(\mathbb{F}_n - F)\right| \geq \eta\right) \leq 2\exp\left(\frac{\sigma(h)^2 t^2}{2n} + \frac{C_1 t^3}{n^2(1 - C_2 t/n)_+} - t\eta\right) \tag{16}
$$

for arbitrary $t, \eta > 0$. Specifically let $\eta = D(\sigma(h)\rho_n^{1/2} + \rho_n^{2/3})$ and set

$$
t := \frac{n\rho_n^{1/2}}{\sigma(h) + \rho_n^{1/6}} \leq n\rho_n^{1/3} = o(n).
$$

Then the bound (16) is not greater than

$$
\begin{aligned}
&2\exp\left(\frac{\sigma(h)^2 \log n}{2(\sigma(h) + \rho_n^{1/6})^2} + \frac{C_1\rho_n^{1/2}\log n}{(\sigma(h) + \rho_n^{1/6})^3(1 - C_2\rho_n^{1/3})_+} - D\log n\right) \\
&\leq 2\exp\left[\left(\frac{1}{2} + \frac{C_1}{(1 - C_2\rho_n^{1/3})_+} - D\right)\log n\right] = 2\exp\left((O(1) - D)\log n\right).
\end{aligned}
$$

Consequently, for sufficiently large $D > 0$,

$$
\begin{aligned}
&\mathbb{P}\left(\max_{h \in \mathcal{H}_n} \frac{\left|\int h\,\mathrm{d}(\mathbb{F}_n - F)\right|}{\sigma(h)\rho_n^{1/2} + W(h)\rho_n^{2/3}} \geq D\right) \\
&\leq \#\mathcal{H}_n 2\exp\left((O(1) - D)\log n\right) = O(1)\exp\left((O(1) + p - D)\log n\right) \to 0. \qquad \square
\end{aligned}
$$

Proof of Lemma A.6. Let $\mathcal{H}$ be the family of all functions h of the form

$$
h(x) = 1\{x \in Q\}(c + dx)
$$

with any interval $Q \subset \mathbb{R}$ and real constants c, d such that h is nonnegative. Suppose that there exists a constant $C = C(f)$ such that

$$
\mathbb{P}\left(\sup_{h \in \mathcal{H}} \frac{\left|\int h\,\mathrm{d}(\mathbb{F}_n - F)\right|}{\sigma(h)\rho_n^{1/2} + W(h)\rho_n^{2/3}} \leq C\right) \to 1. \tag{17}
$$

For any $m \in \mathbb{N}$, an arbitrary function $\Delta \in \mathcal{D}_m$ may be written as

$$
\Delta = \sum_{i=1}^{M} h_i
$$

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with $M = 2m + 2$ functions $h_i \in \mathcal{H}$ having pairwise disjoint supports. Consequently,

$$\sigma(\Delta) = \Big(\sum_{i=1}^{M} \sigma(h_i)^2\Big)^{1/2} \geq M^{-1/2} \sum_{i=1}^{M} \sigma(h_i)$$

by the Cauchy-Schwarz inequality, while

$$W(\Delta) = \max_{i=1,\dots,M} W(h_i) \geq M^{-1} \sum_{i=1}^{M} W(h_i).$$

Consequently, (17) entails that

$$\Big| \int \Delta \, \mathrm{d}(\mathbb{F}_n - F) \Big| \leq \sum_{i=1}^{M} \Big| \int h_i \, \mathrm{d}(\mathbb{F}_n - F) \Big|$$

$$\leq C\Big(\sum_{i=1}^{M} \sigma(h_i)\rho_n^{1/2} + \sum_{i=1}^{M} W(h_i)\rho_n^{2/3} \Big)$$

$$\leq 4C\big(\sigma(\Delta)m^{1/2}\rho_n^{1/2} + W(\Delta)m\rho_n^{2/3} \big)$$

uniformly in $m \in \mathbb{N}$ and $\Delta \in \mathcal{D}_m$ with probability tending to one as $n \to \infty$.

It remains to verify (17). To this end we use a bracketing argument. With the weight function $w(x) = \max\big(1, |\varphi(x)|\big)$ let $-\infty = t_{n,0} < t_{n,1} < \cdots < t_{n,N(n)} = \infty$ such that for $I_{n,j} := (t_{n,j-1}, t_{n,j}]$,

$$(2n)^{-1} \leq \int_{I_{n,j}} w(x)^2 f(x) \, \mathrm{d}x \leq n^{-1} \quad \text{for } 1 \leq j \leq N(n)$$

with equality if $j < N(n)$. Since $1 \leq \int \exp(t_o w(x)) f(x) \, \mathrm{d}x < \infty$, such a partition exists with $N(n) = O(n)$. For any $h \in \mathcal{H}$ we define functions $h_{n,\ell}, h_{n,u}$ as follows: Let $\{j, \dots, k\}$ be the set of all indices $i \in \{1, \dots, N(n)\}$ such that $\{h > 0\} \cap I_{n,i} \neq \emptyset$. Then we define

$$h_{n,\ell}(x) := 1_{\{t_{n,j} < x \leq t_{n,k-1}\}} h(x)$$

and

$$h_{n,u}(x) := h_{n,\ell}(x) + 1\{x \in I_{n,j} \cup I_{n,k}\} W(h)w(x).$$

Note that $0 \leq h_{n,\ell} \leq h \leq h_{n,u} \leq W(h)w$. Consequently, $W(h_{n,\ell}) \leq W(h) = W(h_{n,u})$. Suppose for the moment that the assertion is true for the (still infinite) family $\mathcal{H}_n := \big\{h_{n,\ell}, h_{n,u} :$

$h \in \mathcal{H}\}$ in place of $\mathcal{H}$. Then it follows from $w \geq 1$ that

$$
\begin{aligned}
\int h \, \mathrm{d}(\mathbb{F}_n - F) &\leq \int h_{n,u} \, \mathrm{d}\mathbb{F}_n - \int h_{n,\ell} \, \mathrm{d}F \\
&= \int h_{n,u} \, \mathrm{d}(\mathbb{F}_n - F) + \int (h_{n,u} - h_{n,\ell}) \, \mathrm{d}F \\
&\leq \int h_{n,u} \, \mathrm{d}(\mathbb{F}_n - F) + W(h) \int_{I_{n,j} \cup I_{n,k}} w(x)^2 \, \mathrm{d}F \\
&\leq \int h_{n,u} \, \mathrm{d}(\mathbb{F}_n - F) + 2W(h)n^{-1} \\
&\leq C\big(\sigma(h_{n,u})\rho_n^{1/2} + \rho_n^{2/3}\big) + 2n^{-1} \\
&\leq C\big(\sigma(h)\rho_n^{1/2} + 2^{1/2}W(h)n^{-1/2}\rho_n^{1/2} + \rho_n^{2/3}\big) + 2W(h)n^{-1} \\
&\leq (C + o(1))\big(\sigma(h)\rho_n^{1/2} + W(h)\rho_n^{2/3}\big)
\end{aligned}
$$

uniformly in $h \in \mathcal{H}$ with asymptotic probability one. Analogously,

$$
\begin{aligned}
\int h \, \mathrm{d}(\mathbb{F}_n - F) &\geq \int h_{n,\ell} \, \mathrm{d}(\mathbb{F}_n - F) - 2W(h)n^{-1} \\
&\geq -C\big(\sigma(h_{n,\ell})\rho_n^{1/2} + W(h)\rho_n^{2/3}\big) - 2W(h)n^{-1} \\
&\geq -(C + o(1))\big(\sigma(h)\rho_n^{1/2} + W(h)\rho_n^{2/3}\big)
\end{aligned}
$$

uniformly in $h \in \mathcal{H}$ with asymptotic probability one.

To line up with Lemma A.8, we now have to deal with $\mathcal{H}_n$. For any $h \in \mathcal{H}$ the function $h_{n,\ell}$ may be written as

$$
h(t_{n,j})g_{n,j,k}^{(1)} + h(t_{n,k-1})g_{n,j,k}^{(2)}
$$

with the "triangular functions"

$$
\begin{aligned}
g_{n,j,k}^{(1)}(x) &:= \frac{t_{n,k-1} - x}{t_{n,k-1} - t_{n,j}} \quad \text{and} \\
g_{n,j,k}^{(2)}(x) &:= \frac{x - t_{n,j}}{t_{n,k-1} - t_{n,j}} \quad \text{for } 1 \leq j < k \leq N(n), k - j \geq 2.
\end{aligned}
$$

In case of $k - j \leq 1$ we set $g_{n,j,k}^{(1)} := g_{n,j,k}^{(2)} := 0$. Moreover,

$$
h_{n,u} = h_{n,\ell} + W(h)g_{n,j} + 1\{k > j\}W(h)g_{n,k}
$$

with $g_{n,i}(x) := 1\{x \in I_{n,i}\}w(x)$. Consequently, all functions in $\mathcal{H}_n$ are linear combinations with *nonnegative* coefficients of at most four functions in the finite family

$$
\mathcal{G}_n := \{g_{n,i} : 1 \leq i \leq N(n)\} \cup \{g_{n,j,k}^{(1)}, g_{n,j,k}^{(2)} : 1 \leq j < k \leq N(n)\}.
$$

Since $\mathcal{G}_n$ contains $O(n^2)$ functions, it follows from Lemma A.8 that for some constant $D > 0$,

$$
\left| \int g \, \mathrm{d}(\mathbb{F}_n - F) \right| \leq D\big(\sigma(g)\rho_n^{1/2} + W(g)\rho_n^{2/3}\big)
$$

for all $g \in \mathcal{G}_n$ with asymptotic probability one. Now the assertion about $\mathcal{H}_n$ follows from the basic observation that for $h = \sum_{i=1}^{4} \alpha_i g_i$ with nonnegative functions g_i and coefficients $\alpha_i \geq 0$,

$$\sigma(h) \geq \left(\sum_{i=1}^{4} \alpha_i^2 \sigma(g_i)^2\right)^{1/2} \geq 2^{-1} \sum_{i=1}^{4} \alpha_i \sigma(g_i),$$

$$W(h) \geq \max_{i=1,\dots,4} \alpha_i W(g_i) \geq 4^{-1} \sum_{i=1}^{4} \alpha_i W(g_i). \qquad \square$$

A.4 Proofs for the gap problem and of $\hat{F}_n$'s consistency

Proof of Theorem 4.3. Suppose that $\hat{\varphi}_n$ is linear on an interval $[a,b]$. Then for $x \in [a,b]$ and $\lambda_x := (x-a)/(b-a) \in [0,1]$,

$$\begin{aligned}
\varphi(x) &- (1-\lambda_x)\varphi(a) - \lambda_x\varphi(b) \\
&= (1-\lambda_x)\big(\varphi(x) - \varphi(a)\big) - \lambda_x\big(\varphi(b) - \varphi(x)\big) \\
&= (1-\lambda_x)\int_a^x \varphi'(t)\,\mathrm{d}t - \lambda_x \int_x^b \varphi'(t)\,\mathrm{d}t \\
&= (1-\lambda_x)\int_a^x \big(\varphi'(t) - \varphi'(x)\big)\,\mathrm{d}t + \lambda_x \int_x^b \big(\varphi'(x) - \varphi'(t)\big)\,\mathrm{d}t \\
&\geq C(1-\lambda_x)\int_a^x (x-t)\,\mathrm{d}t + C\lambda_x \int_x^b (t-x)\,\mathrm{d}t \\
&= C(b-a)^2 \lambda_x(1-\lambda_x)/2 \\
&= C(b-a)^2/8 \quad \text{if } x = x_o := (a+b)/2.
\end{aligned}$$

This entails that $\sup_{[a,b]} |\hat{\varphi}_n - \varphi| \geq C(b-a)^2/16$. For if $\hat{\varphi}_n < \varphi + C(b-a)^2/16$ on $\{a,b\}$, then

$$\begin{aligned}
\varphi(x_o) - \hat{\varphi}_n(x_o) &= \varphi(x_o) - (\hat{\varphi}_n(a) + \hat{\varphi}_n(b))/2 \\
&> \varphi(x_o) - (\varphi(a) + \varphi(b))/2 - C(b-a)^2/16 \\
&\geq C(b-a)^2/8 - C(b-a)^2/16 = C(b-a)^2/16.
\end{aligned}$$

Consequently, if $|\hat{\varphi}_n - \varphi| \leq D_n \rho_n^{\beta/(2\beta+1)}$ on $T_n := \left[A + \rho_n^{1/(2\beta+1)}, B - \rho_n^{1/(2\beta+1)}\right]$ with $D_n = O_{\mathrm{p}}(1)$, then the longest subinterval of T_n containing no points from $\mathcal{S}_n$ has length at most $4D_n^{1/2} C^{-1/2} \rho_n^{\beta/(4\beta+2)}$. Since T_n and $T = [A,B]$ differ by two intervals of length $\rho_n^{1/(2\beta+1)} = O\big(\rho_n^{\beta/(4\beta+2)}\big)$, these considerations yield the assertion about $\mathcal{S}_n(\hat{\varphi}_n)$. $\qquad \square$

Proof of Theorem 4.4. Let $\delta_n := \rho_n^{1/(2\beta+1)}$ and $r_n := D\rho_n^{\beta/(4\beta+2)} = D\delta_n^{1/2}$ for some constant $D > 0$. Since $r_n \to 0$ but $nr_n \to \infty$, it follows from boundedness of f and a theorem of Stute (1982) about the modulus of continuity of univariate empirical processes that

$$\begin{aligned}
\omega_n &:= \sup_{x,y \in \mathbb{R}\,:\,|x-y| \leq r_n} \big|(\mathbb{F}_n - F)(x) - (\mathbb{F}_n - F)(y)\big| \\
&= O_{\mathrm{p}}\big(n^{-1/2} r_n^{1/2} \log(1/r_n)^{1/2}\big) \\
&= O_{\mathrm{p}}\big(\rho_n^{(5\beta+2)/(8\beta+4)}\big).
\end{aligned}$$

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If D is sufficiently large, the asymptotic probability that for any point $x \in [A + \delta_n, B - \delta_n]$ there exists a point $y \in \mathcal{S}_n(\hat{\varphi}_n) \cap [A + \delta_n, B - \delta_n]$ with $|x - y| \le r_n$ is equal to one. In that case it follows from Corollary 2.5 and Theorem 4.1 that

$$
\begin{aligned}
\left| (\hat{F}_n - \mathbb{F}_n)(x) \right| &\le \left| (\hat{F}_n - \mathbb{F}_n)(x) - (\hat{F}_n - \mathbb{F}_n)(y) \right| + n^{-1} \\
&\le \left| (\hat{F}_n - F)(x) - (\hat{F}_n - F)(y) \right| + \omega_n + n^{-1} \\
&\le \int_{\min(x,y)}^{\max(x,y)} |\hat{f}_n - f|(x) \, \mathrm{d}x + \omega_n + n^{-1} \\
&\le O_{\mathrm{p}}\big(r_n \rho_n^{\beta/(2\beta+1)} \big) + \omega_n + n^{-1} \\
&= O_{\mathrm{p}}\big(\rho_n^{3\beta/(4\beta+2)} \big). \qquad \square
\end{aligned}
$$

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Institute of Mathematical Statistics
and Actuarial Science
University of Bern
Sidlerstrasse 5
CH-3012 Bern
Switzerland

www.stat.unibe.ch

Abteilung Biostatistik
Institut für Sozial- und Präventivmedizin
Universität Zürich
Hirschengraben 84
CH-8001 Zürich
Switzerland

www.biostat.uzh.ch