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Nonparametric estimation of the conditional tail copula

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Abstract

The tail copula is widely used to describe the amount of extremal dependence of a multivariate distribution. In some situations such as risk management, the dependence structure can be linked with some covariate. The tail copula thus depends on this covariate and is referred to as the conditional tail copula. The aim of this paper is to propose a nonparametric estimator of the conditional tail copula and to establish its asymptotic normality. Some illustrations are presented both on simulated and real datasets.

AMS Subject Classifications: 62G05, 62G20, 62G30, 62G32.

Keywords: Extremal dependence, Tail copula, Covariate.

1 Introduction

In various domains such as insurance, finance and hydrology, there is an increasing interest in understanding the dependence structure between large values of two or more variables. For example, in finance, one can be interested in the link between large values of the market index and large returns of individual stocks, see [2, 28]. The dependence structure of a random vector $Y = (Y_1, \dots, Y_d)^T$ with $d \geq 2$ is commonly described by a copula which is a multivariate distribution function with uniformly distributed margins on the interval $[0, 1]$. Let us denote by \bar{F}_i , $i = 1, \dots, d$ the marginal survival functions of Y . From Sklar theorem [34], if the margins of Y are continuous, there exist a unique copula S defined on $[0, 1]^d$ such that for all $y = (y_1, \dots, y_d)^T \in \mathbb{R}^d$, the multivariate survival function writes $\bar{F}(y) := \mathbb{P}(Y \geq y) = S(\bar{F}_1(y_1), \dots, \bar{F}_d(y_d))$. This function S is then given for all $u = (u_1, \dots, u_d)^T \in [0, 1]^d$ by $S(u) = \bar{F}(q(u))$, where $q(u) := (\bar{F}_1^{\leftarrow}(u_1), \dots, \bar{F}_d^{\leftarrow}(u_d))^T$ is a multivariate quantile. Note that $(a_1, \dots, a_d) \geq (b_1, \dots, b_d)$ corresponds to the componentwise inequality $a_i \geq b_i$ for all $i = 1, \dots, d$. Besides, f^{\leftarrow} denotes the generalized inverse of a decreasing function f *i.e.*, $f^{\leftarrow}(x) = \inf\{t, f(t) \leq x\}$. As mentioned in [2], the tail copula is a useful tool to

describe the upper tail dependence structure. It is defined as follows: if there exists a function $\Lambda : [0, \infty)^d \rightarrow [0, \infty)$ such that

$$\lim_{\alpha \rightarrow 0^+} \frac{1}{\alpha} S(\alpha y) = \Lambda(y), \quad y \in (0, \infty)^d, \quad (1)$$

then Λ is called the tail copula [26]. Let us note that (1) entails first that Λ is an homogeneous positive function and second that S is a multivariate regularly varying function at 0 since

$$\lim_{\alpha \rightarrow 0^+} \frac{S(\alpha y)}{S(\alpha \mathbf{1})} = W(y),$$

where $W(\cdot) = \Lambda(\cdot)/\Lambda(\mathbf{1})$ and $\mathbf{1} = (1, \dots, 1)^T \in \mathbb{R}^d$. We refer to [31, Chapter 5] for a definition of multivariate regularly varying functions and to [4] for a general account on regular variation. In practice, the tail copula is often reduced to the tail-dependence coefficient defined by

$$\lambda := \Lambda(\mathbf{1}) = \lim_{\alpha \rightarrow 0^+} \mathbb{P} \{ Y_i \geq \bar{F}_i^{\leftarrow}(\alpha), \quad i \neq j \mid Y_j \geq \bar{F}_j^{\leftarrow}(\alpha) \}.$$

Clearly, this coefficient does not depend on the index $j \in \{1, \dots, d\}$. It was first introduced by [33] in the bivariate case. The multivariate random vector Y is said to be tail-dependent if $\lambda \in (0, 1]$ and tail-independent if $\lambda = 0$. The case $\lambda = 1$ corresponds to complete tail dependence. Intuitively, when $\lambda = 0$, large values of the margins Y_i are unlikely to happen simultaneously. As an example, it is well known that a bivariate normal distribution with correlation coefficient different from ± 1 is tail-independent.

The bivariate case ($d = 2$) has been widely studied in the literature especially when the distribution function of Y belongs to the maximum domain of attraction of a bivariate extreme-value distribution *i.e.* when there exist positive sequences $(a_{1,n})$, $(a_{2,n})$ and sequences $(b_{1,n})$ and $(b_{2,n})$ such that

$$\lim_{n \rightarrow \infty} \mathbb{P} \left\{ \bigvee_{j=1}^n \frac{Y_{1,j} - b_{1,n}}{a_{1,n}} \leq y_1; \bigvee_{j=1}^n \frac{Y_{2,j} - b_{2,n}}{a_{2,n}} \leq y_2 \right\} = G(y_1, y_2),$$

where $(Y_{1,j}, Y_{2,j})$, $j = 1, \dots, n$ are independent copies of the random vector $Y = (Y_1, Y_2)$. The bivariate extreme-value distribution G can be expressed as $G(y_1, y_2) = D(G_1(y_1), G_2(y_2))$, where G_1 and G_2 are the margins of G and D is the bivariate extreme-value copula defined by:

$$D(u, v) = \exp \left\{ \log(uv) A \left(\frac{\log u}{\log(uv)} \right) \right\}. \quad (2)$$

The marginal distributions G_1 and G_2 are univariate extreme value distributions *i.e.* for $i \in \{1, 2\}$,

$$G_i(x) = \exp\{-(1 + \gamma_i x)_+^{-1/\gamma_i}\},$$

where $y_+ = 0 \vee y$ and $\gamma_i \in \mathbb{R}$ is called the extreme value index. The function $A(t)$ is referred to as the Pickands dependence function. It is a convex function lying in a triangular area *i.e.* such

that $t \vee (1 - t) \leq A(t) \leq 1$ for all $t \in [0, 1]$. It can be shown [31, Chapter 5] that for a bivariate distribution function in the maximum domain of attraction of G , the tail copula exists and is given by

$$\Lambda(y_1, y_2) = (y_1 + y_2) \left\{ 1 - A \left(\frac{y_1}{y_1 + y_2} \right) \right\}. \quad (3)$$

Collecting (2) and (3), the bivariate extreme-value copula can be rewritten in terms of the tail copula as

$$D(u, v) = uv \exp \Lambda(-\log u, -\log v).$$

As a consequence, in the bivariate case, the tail copula and the Pickands dependence function both fully characterise the extreme dependence structure of Y . Finally, one may equivalently consider the so-called stable tail-dependence function given by

$$\ell(y_1, y_2) = y_1 + y_2 - \Lambda(y_1, y_2) \quad (4)$$

to describe the extreme dependence structure.

In the bivariate case, numerous nonparametric estimators of the Pickands dependence function have been introduced [5, 14, 30, 37] thanks to an interpretation of A in terms of exponential distributions. In [24], it is remarked that such estimators may not provide proper dependence functions (convex and lying in the above triangular area) and two new estimators satisfying these constraints are introduced. Similarly, a projection technique is proposed in [15] to force any estimator to be a proper dependence function. The above mentioned estimators are constructed assuming knowledge of the marginal distributions. Rank-based versions of these estimators are proposed for the unknown margins case in [18]. The estimation of A within the class of Archimax copulas is studied in [6]. See [3, Chapter 9] for an overview of parametric and nonparametric inference methods for bivariate Pickands dependence functions. We also refer to [22, 23] for extensions to the general multivariate setting. In view of (4), it is clear that estimating the stable tail-dependence function or the tail copula are equivalent problems. The first estimators of ℓ are introduced in [25] in the bivariate case. They are then adapted to the estimation of the tail copula [32] and compared to nonparametric estimators based on the empirical copula. Parametric estimators of Λ have also been proposed in the bivariate case [12] and for multivariate elliptical copulas [27].

The case where the random vector Y is recorded simultaneously with a predictive variable $X \in \mathbb{R}^p$ has been less considered. In [36], the authors propose to estimate the tail-dependence coefficient in the regression case using a maximum likelihood approach. In this regression framework, S is a conditional copula. Its estimation is addressed in [21, 35] with nonparametric techniques and in [1] using semiparametric approaches. The case of functional covariates is studied in [20]. Up to our knowledge, the estimation of the tail copula in the regression case has not been considered yet despite potential applications: study of the Chinese stock market [36] or modelling of financial returns [16].

From now on, we thus consider the random vector $(X, Y) \in \mathbb{R}^p \times \mathbb{R}^d$. For $i = 1, \dots, d$, the conditional survival function of Y_i given $X = x$ is denoted by $\bar{F}_i(\cdot|x) := \mathbb{P}(Y_i \geq \cdot | X = x)$. The dependence structure of Y conditionally on $X = x$ is then described by the conditional copula:

$$S(u|x) = \bar{F}(q(u|x)|x), \quad u = (u_1, \dots, u_d) \in \mathbb{R}^d,$$

where $\bar{F}(\cdot|x) : \mathbb{R}^d \rightarrow [0, 1]$ is the conditional survival function of Y given $X = x$ and $q(\cdot|x)$ is the multivariate conditional quantile defined as $q(u|x) := (\bar{F}_1^{\leftarrow}(u_1|x), \dots, \bar{F}_d^{\leftarrow}(u_d|x))^T$. Similarly to (1), we thus define the conditional tail copula by

$$\Lambda(y|x) := \lim_{\alpha \rightarrow 0^+} \frac{1}{\alpha} S(\alpha y|x), \quad y \in (0, \infty)^d, \quad (5)$$

if this limit exists. Note that in what follows, we only consider the case $\Lambda(y|x) > 0$ *i.e.* tail-dependent distributions.

In Section 2, we propose nonparametric estimators of $\Lambda(y|x)$ for a fixed vector $(x, y) \in \mathbb{R}^p \times [0, \infty)^d$ when the margins are first known and then unknown. Their asymptotic properties are established in Section 3. A simulation study is presented in Section 4 and an illustration on real data is proposed in Section 5. Finally, proofs of the main results are given in Section 6.

2 Estimation of the conditional tail copula

Let (X_j, Y_j) , $j = 1, \dots, n$ be independent copies of the random vector (X, Y) . The estimator of the conditional tail copula is based on a nonparametric estimator of the conditional copula function of Y given $X = x$. In the following, we propose to use a classical kernel estimator. The first step consists in estimating the probability density function g of X by:

$$\hat{g}_h(x) = \frac{1}{nh^p} \sum_{j=1}^n K\left(\frac{x - X_j}{h}\right), \quad x \in \mathbb{R}^p,$$

where $h = h_n$ is a non random sequence tending to zero as n goes to infinity and K is a bounded density on \mathbb{R}^p with support \mathcal{S} included in the unit ball of \mathbb{R}^p . As mentioned in the introduction, in many related papers the margins are often supposed to be known. In this situation, the conditional copula $S(\cdot|x)$ can be estimated by

$$\tilde{S}_h(u|x) := \frac{1}{nh^p \hat{g}_h(x)} \sum_{j=1}^n \mathbb{I}\{Y_j \geq q(u|x)\} K\left(\frac{x - X_j}{h}\right), \quad u \in [0, 1]^d, \quad x \in \mathbb{R}^p,$$

where $\mathbb{I}\{\cdot\}$ is the indicator function. An estimator of the tail copula is then given by:

$$\tilde{\Lambda}_h(y|x) = \frac{1}{\alpha} \tilde{S}_h(\alpha y|x), \quad (6)$$

where $y \in (0, \infty)^d$ and $\alpha = \alpha_n$ is a sequence tending to zero as n goes to infinity. In this paper, we consider the more realistic case where nothing is known on the margins. In this situation, the

estimation of the marginal survival functions of Y given $X = x$ is obviously required. For all $u \in [0, 1]^d$, let us denote by

$$\hat{q}(u|x) = (\hat{q}_1(u_1|x), \dots, \hat{q}_d(u_d|x))^T$$

an estimator of the vector $q(u|x) \in \mathbb{R}^d$. We propose the following plug-in estimator of the conditional copula:

$$\hat{S}_h(u|x) := \frac{1}{nh^p \hat{g}_h(x)} \sum_{j=1}^n \mathbb{I}\{Y_j \geq \hat{q}(u|x)\} K\left(\frac{x - X_j}{h}\right), \quad u \in [0, 1]^d, \quad x \in \mathbb{R}^p.$$

The conditional tail copula is then estimated by:

$$\hat{\Lambda}_h(y|x) = \frac{1}{\alpha} \hat{S}_h(\alpha y|x). \quad (7)$$

As an example of an estimator of $q(u|x)$, one can propose the statistic

$$\check{q}_k(u|x) := (\hat{F}_{k,1}^{\leftarrow}(u_1|x), \dots, \hat{F}_{k,d}^{\leftarrow}(u_d|x))^T, \quad (8)$$

where $k = k_n$ is a non random sequence tending to zero as n goes to infinity and, for $i = 1, \dots, d$, $\hat{F}_{k,i}^{\leftarrow}(\cdot|x)$ is the usual kernel estimator with bandwidth k for a conditional survival function defined as

$$\hat{F}_{k,i}^{\leftarrow}(z|x) = \frac{1}{nk^p \hat{g}_k(x)} \sum_{j=1}^n \mathbb{I}\{Y_{i,j} \geq z\} K\left(\frac{x - X_j}{k}\right). \quad (9)$$

The obtained estimators of the conditional copula and of the tail copula are respectively denoted by $\check{S}_{h,k}(\cdot|x)$ and $\check{\Lambda}_{h,k}(\cdot|x)$. Let us highlight that these estimators depend on two bandwidth h and k . The first one tune the smoothness of the estimator of the conditional copula, while the second one controls the smoothness of the estimator of the marginal distributions. In the following, the asymptotic behavior of $\check{\Lambda}_{h,k}(\cdot|x)$ is analysed in the two situations $h = k$ and $h \neq k$.

3 Main results

In this section, the asymptotic properties of the estimators $\tilde{\Lambda}_h(y|x)$ and $\hat{\Lambda}_h(y|x)$ are established for a fixed vector $(x, y) \in \mathbb{R}^p \times [0, \infty)^d$ such that $\Lambda(y|x) > 0$. Some regularity and extreme-value assumptions are required. The first one is a classical second-order condition on the conditional copula function $S(\cdot|x)$, see for instance [32].

H.1 There exist a function $b(\cdot|x)$ with $b(t|x) \rightarrow 0$ as $t \rightarrow \infty$ and a function $c_\Lambda(\cdot|x)$ such that for all $y \in [0, \infty)^d$,

$$\lim_{t \rightarrow \infty} \frac{\Lambda(y|x) - tS(t^{-1}y|x)}{b(t|x)} = c_\Lambda(y|x) < \infty.$$

Condition **H.1** establishes that rate of convergence in (5) is driven by a given function $b(\cdot|x)$. Let us recall that necessarily $|b(\cdot|x)|$ is a regularly varying function at infinity with index $\rho(x) < 0$, see [13].

In the extreme-value literature, $\rho(x)$ is referred to as the conditional second order parameter. As an example, let us consider the multivariate extreme-value copula defined for all $u = (u_1, \dots, u_d)^T$ by

$$\mathbb{P}(Y \leq q(1-u|x)|X=x) = \exp \left\{ \kappa(u) A \left(\frac{\log u_i}{\kappa(u)}, i = 1, \dots, d-1 \middle| x \right) \right\}, \quad (10)$$

where

$$\kappa(u) = \sum_{l=1}^d \log u_l,$$

and $A(\cdot|x) : \mathbb{R}^{d-1} \rightarrow [1/2, 1]$ is the conditional multivariate Pickands dependence function [30]. Clearly, (10) is the extension of the bivariate extreme-value copula (2) to a dimension $d > 2$. It can be shown that, in the dependent case (*i.e.* $A(\cdot|x) \neq 1$) and if $A(\cdot|x)$ is continuously differentiable (denoted by $A(\cdot|x) \in \mathcal{C}^1$), then condition **H.1** is satisfied with $\rho(x) = -1$.

A Lipschitz condition on the probability density function g of X is also required. Denoting by $\tilde{d}(x, x')$ the Euclidean distance between $x \in \mathbb{R}^p$ and $x' \in \mathbb{R}^p$, it is assumed that

H.2 There exists a positive constant c_g such that $|g(x) - g(x')| \leq c_g \tilde{d}(x, x')$.

For $x \in \mathbb{R}^p$ and $u \in [0, 1]^d$, the oscillation of the conditional survival function $\bar{F}(\cdot|x)$ is controlled in a neighbourhood of $q(u|x)$ by the following quantity:

$$\omega_{h,\zeta}(u|x) = \sup \left\{ \left| \frac{\bar{F}(q(v|x)|t)}{S(v|x)} - 1 \right| \text{ for } t \in B(x, h), (1-\zeta)u \leq v \leq (1+\zeta)u \right\}, \quad (11)$$

where $\zeta \in [0, 1)$ and $B(x, h)$ is the ball of \mathbb{R}^p centred at x with radius h .

In a first time, conditions **H.1** and **H.2** are sufficient to establish the asymptotic normality of the estimator $\tilde{\Lambda}_h(y|x)$ in the (unrealistic) situation where the margins are known.

Theorem 1. *Suppose **H.1** and **H.2** hold. Let $x \in \mathbb{R}^p$ and $y \in [0, \infty)^d$ such that $g(x) \neq 0$ and $\Lambda(y|x) \neq 0$. If α in (6) is such that $\alpha \rightarrow 0$, $nh^p\alpha \rightarrow \infty$ and*

$$nh^p\alpha(h \vee \omega_{h,0}(\alpha y|x) \vee b(\alpha^{-1}|x))^2 \rightarrow 0 \quad (12)$$

as $n \rightarrow \infty$, then,

$$(nh^p\alpha)^{1/2} \left(\tilde{\Lambda}_h(y|x) - \Lambda(y|x) \right) \xrightarrow{d} \mathcal{N} \left(0, \frac{\|K\|_2^2 \Lambda(y|x)}{g(x)} \right).$$

As pointed out in [9], $nh^p\alpha \rightarrow \infty$ is a necessary and sufficient condition for the almost sure presence of at least one sample point in the region $B(x, h) \times (q(\alpha|x), \infty)$ of \mathbb{R}^{p+d} . Condition (12) forces the bias to be neglectible compared to the standard-deviation of the estimator. Two main sources of bias appear. The first contribution to the bias is due to the kernel smoothing and to the oscillations of order $\mathcal{O}(h)$ and $\mathcal{O}(\omega_{h,0}(\alpha y|x))$ of g and $\bar{F}(\cdot|x)$ respectively. The second source of bias is due

to the approximation of the conditional tail copula $\Lambda(y|x)$ by $\alpha S(\alpha^{-1}y|x)$. This entails a bias of order $\mathcal{O}(b(\alpha^{-1}|x))$, see assumption **H.1**.

In order to deal with the case where the margins are estimated, an extreme-value assumption on the conditional marginal distributions of Y given $X = x$ is required. The following von-Mises condition (see [11], equation (1.1.30)) is considered:

H.3 For $i = 1, \dots, d$, the marginal survival function $\bar{F}_i(\cdot|x)$ is twice differentiable and such that

$$\lim_{y \uparrow y_{F_i}(x)} \frac{\bar{F}_i(y|x) \bar{F}_i''(y|x)}{(\bar{F}_i'(y|x))^2} = \gamma_i(x) + 1,$$

where $y_{F_i}(x) := \bar{F}_i^{\leftarrow}(0|x) \in (-\infty, \infty]$ is the right endpoint of the i -th margin, $\bar{F}_i'(\cdot|x)$ and $\bar{F}_i''(\cdot|x)$ are respectively the first and the second derivative of $\bar{F}_i(\cdot|x)$.

From [11], Theorem 1.1.8, condition **H.3** entails that for all $i = 1, \dots, d$, the conditional distribution of Y_i given $X = x$ is in the domain of attraction of the univariate extreme-value distribution with conditional extreme-value index $\gamma_i(x) \in \mathbb{R}$. Note also that from [10], Lemma 2, condition **H.3** implies that for all $i = 1, \dots, d$ and for all sequences $\alpha \rightarrow 0$ and $t_{n,i}(x) \rightarrow 0$ as $n \rightarrow \infty$, there exists a positive auxiliary function $a_i(\cdot|x)$ such that

$$\lim_{n \rightarrow \infty} \alpha^{-1} \bar{F}_i(\bar{F}_i^{\leftarrow}(\alpha|x) + t_{n,i}(x) a_i(\bar{F}_i^{\leftarrow}(\alpha|x)|x)|x) = 1. \quad (13)$$

The next assumption permits to control the rate of convergence of the estimator $\hat{q}(\cdot|x)$ towards the true quantile function $q(\cdot|x) = \bar{F}_i^{\leftarrow}(\cdot|x)$.

H.4 There exists a sequence (η_n) satisfying $\eta_n \rightarrow \infty$ such that for all $y = (y_1, \dots, y_d)^T \in (0, \infty)^d$,

$$\eta_n \bigvee_{i=1}^d \frac{\bar{F}_i^{\leftarrow}(y_i \alpha | x)}{a_i(\bar{F}_i^{\leftarrow}(y_i \alpha | x) | x)} \left| \frac{\hat{q}_i(y_i \alpha | x)}{\bar{F}_i^{\leftarrow}(y_i \alpha | x)} - 1 \right| = \mathcal{O}_{\mathbb{P}}(1).$$

A condition on the gradient $\nabla S(\cdot|x)$ of the conditional copula function is also needed. It states that convergence (5) still holds when differentiating.

H.5 The conditional copula $S(\cdot|x)$ is of class \mathcal{C}^1 . Furthermore, for $y \in (0, \infty)^d$ and for all vectors $\vartheta(\alpha, y)$ such that $\vartheta(\alpha, y) = \alpha y(1 + o(1))$ as $\alpha \rightarrow 0^+$,

$$\lim_{\alpha \rightarrow 0^+} y^T \nabla S(\vartheta(\alpha, y)|x) = \Lambda(y|x).$$

This assumption is for instance satisfied by the multivariate extreme-value copula (10) with $A(\cdot|x) \neq 1$ and $A(\cdot|x) \in \mathcal{C}^1$.

We are now in position to state our main result on the asymptotic behavior of $\hat{\Lambda}_h(y|x)$.

Theorem 2. Suppose **H.1** - **H.5** hold. Let $x \in \mathbb{R}^p$ and $y \in [0, \infty)^d$ such that $g(x) \neq 0$ and $\Lambda(y|x) \neq 0$.

(i) If α in (7) is such that $\alpha \rightarrow 0$, $nh^p\alpha\eta_n^{-1} \rightarrow \infty$ and, for some $\zeta > 0$,

$$nh^p\alpha \left(h \vee \eta_n^{1/2} \omega_{h,\zeta}(\alpha y|x) \vee b(\alpha^{-1}|x) \right)^2 \rightarrow 0, \quad (14)$$

as $n \rightarrow \infty$, then

$$\hat{\Lambda}_h(y|x) - \Lambda(y|x) = (nh^p\alpha)^{-1/2} \xi_n + \mathcal{O}_{\mathbb{P}}(\eta_n^{-1}),$$

where ξ_n converges in distribution to a $\mathcal{N}(0, \|K\|_2^2 \Lambda(y|x)/g(x))$ random variable.

(ii) If, moreover, $(nh^p\alpha)^{1/2} \eta_n^{-1} \rightarrow 0$, then

$$(nh^p\alpha)^{1/2} (\hat{\Lambda}_h(y|x) - \Lambda(y|x)) \xrightarrow{d} \mathcal{N}\left(0, \frac{\|K\|_2^2 \Lambda(y|x)}{g(x)}\right).$$

Note that Theorem 2(i) yields $\hat{\Lambda}_h(y|x) \xrightarrow{\mathbb{P}} \Lambda(y|x)$. In situation (ii), condition $(nh^p\alpha)^{1/2} \eta_n^{-1} \rightarrow 0$ entails that the estimator of $q(\cdot|x)$ converges faster than the one of the conditional copula $S(\cdot|x)$. Thus, the estimator $\hat{\Lambda}_h(y|x)$ has the same asymptotic behavior as the estimator $\tilde{\Lambda}_h(y|x)$. In this case, the estimation of the margins does not change the asymptotic distribution.

As an illustration, let us consider the situation where the statistic (8) is used to compute the estimator of the conditional tail copula. To this end, the regularity of $\bar{F}_i(\cdot|x)$ is controlled by

$$\bar{\omega}_{h,\zeta}^{(i)}(u|x) = \sup \left\{ \left| \frac{\bar{F}_i(\bar{F}_i^{\zeta-}(v_i|x)|t)}{v_i} - 1 \right| \text{ for } t \in B(x, h) \text{ and } (1 - \zeta)u_i \leq v_i \leq (1 + \zeta)u_i \right\}$$

where $u \in (0, 1)^d$ and $\zeta \in [0, 1)$. Introducing (e_1, \dots, e_d) the canonical basis of \mathbb{R}^d , the above quantity can be rewritten using (11) as $\bar{\omega}_{h,\zeta}^{(i)}(u|x) = \omega_{h,\zeta}^{(i)}((u-1)e_i + \mathbf{1}|x)$. Let us first focus on the situation where the same bandwidth h is used to estimate the conditional copula and the margins. The asymptotic behavior of the estimator $\tilde{\Lambda}_{h,h}(y|x)$ is a consequence of Theorem 2(i).

Corollary 1. Suppose **H.1**, **H.2**, **H.3** and **H.5** hold. Let $x \in \mathbb{R}^p$ and $y \in [0, \infty)^d$ such that $g(x) \neq 0$ and $\Lambda(y|x) \neq 0$. If α in (7) is such that $\alpha \rightarrow 0$, $nh^p\alpha \rightarrow \infty$ and, for some $\zeta > 0$,

$$nh^p\alpha \left(h \vee b(\alpha^{-1}|x) \vee \omega_{h,\zeta}^{2/3}(\alpha y|x) \vee \bigvee_{i=1}^d \bar{\omega}_{h,\zeta}^{(i)}(y_i\alpha|x) \right)^2 \rightarrow 0,$$

as $n \rightarrow \infty$, then

$$\tilde{\Lambda}_{h,h}(y|x) - \Lambda(y|x) = \mathcal{O}_{\mathbb{P}}\left((nh^p\alpha)^{-1/2}\right).$$

Here, the estimators of $q(\cdot|x)$ and $S(\cdot|x)$ share the same bandwidth and thus the same rate of convergence $(nh^p\alpha)^{-1/2}$. Hence, plugging-in (8) in $\tilde{\Lambda}_h(y|x)$ does not change this rate of convergence. Finally, we consider the situation where two different bandwidth sequences k and h are used to estimate the marginal distributions and the conditional copula. The asymptotic behavior of the estimator $\tilde{\Lambda}_{h,k}(y|x)$ is a consequence of Theorem 2(ii).

Corollary 2. *Suppose **H.1**, **H.2**, **H.3** and **H.5** hold. Let $x \in \mathbb{R}^p$ and $y \in [0, \infty)^d$ such that $g(x) \neq 0$ and $\Lambda(y|x) \neq 0$. If $h/k \rightarrow 0$ and α in (7) is such that $\alpha \rightarrow 0$, $nh^p\alpha \rightarrow \infty$, $nk^{p+2}\alpha \rightarrow 0$ and, for some $\zeta > 0$,*

$$nh^p\alpha \left(b(\alpha^{-1}|x) \vee \omega_{h,\zeta}^{2/3}(\alpha y|x) \vee \left(\frac{k}{h}\right)^{p/2} \bigvee_{i=1}^d \bar{\omega}_{h,\zeta}^{(i)}(y_i\alpha|x) \right)^2 \rightarrow 0,$$

as $n \rightarrow \infty$, then

$$(nh^p\alpha)^{1/2}(\check{\Lambda}_{h,k}(y|x) - \Lambda(y|x)) \xrightarrow{d} \mathcal{N}\left(0, \frac{\|K\|_2^2 \Lambda(y|x)}{g(x)}\right).$$

Note that condition $h/k \rightarrow 0$ entails that the rate of convergence of the margin estimators $(nk^p\alpha)^{1/2}$ is faster than the one of the conditional copula estimator $(nh^p\alpha)^{1/2}$. The asymptotic normality of the plug-in estimator can thus be established and the obtained rate of convergence is $(nh^p\alpha)^{1/2}$.

4 Illustration on simulated data

The behavior of the conditional tail copula estimator $\check{\Lambda}_{h,k}(\cdot|x)$ is investigated on a two-dimensional simulated dataset when the marginal distributions are supposed to be unknown. The smoothness of the estimators of the conditional copula and of the marginal distributions are tuned by the same bandwidth $h = k$ and by the same bi-quadratic kernel given by:

$$K(x_1, x_2) := \frac{3}{\pi} (1 - (x_1^2 + x_2^2))^2 \mathbb{I}\{x_1^2 + x_2^2 \leq 1\}.$$

The simulated model is described in Paragraph 4.1, implementation details are given in Paragraph 4.2 and results are reported in Paragraph 4.3.

4.1 Model

The covariate $X \in [0, 1]$ follows a standard uniform distribution. The conditional distribution of Y_1 given $X = x$ is an exponential distribution with mean parameter $\mu(x) > 0$. Besides, Y_2 given $X = x$ follows a Pareto distribution with tail-index $\gamma(x) > 0$, i.e. $\mathbb{P}(Y_2 > y|X = x) = y^{-1/\gamma(x)}$ for all $y \geq 1$. The dependence structure between the margins Y_1 and Y_2 is controlled by the Gumbel copula defined for $(u_1, u_2) \in (0, 1)^2$ by:

$$S(u_1, u_2|x) = \exp \left\{ - \left[(-\log(1 - u_1))^{\theta(x)} + (-\log(1 - u_2))^{\theta(x)} \right]^{1/\theta(x)} \right\} + u_1 + u_2 - 1,$$

where for all $x \in [0, 1]$, $\theta(x) > 1$ tunes the dependence level. If $\theta(x) = 1$, then $S(u_1, u_2|x) = u_1 u_2$ is the independent copula. At the opposite, if $\theta(x) \rightarrow +\infty$, then $S(u_1, u_2|x) \rightarrow u_1 \wedge u_2$ which is the Fréchet upper bound copula for modelling comonotone random variables. The conditional tail copula is given by:

$$\Lambda(y_1, y_2|x) = y_1 + y_2 - \left(y_1^{\theta(x)} + y_2^{\theta(x)} \right)^{1/\theta(x)},$$

for all $(y_1, y_2) \in (0, \infty)^2$ and $x \in [0, 1]$. The function θ is taken as $\theta(x) = 12x^2 - 12x + 5$ for $x \in [0, 1]$, its minimum is $\theta(0.5) = 2$ while its maximum is $\theta(0) = \theta(1) = 5$. Two cases are considered for the functions μ and γ driving the margin distributions:

$$\mu(x) = \gamma(x) = 1 \text{ (Case 1)} \quad \text{and} \quad \mu(x) = \gamma(x) + 1/2 = \frac{1}{2} \sin(2\pi x) + 1 \text{ (Case 2)} .$$

Let us note that, in Case 1, the margins do not depend on the covariate. For visualization purposes, only two different values of x are considered: $x = 0.1$ corresponding to a strong dependence and $x = 0.5$ corresponding to a weak dependence. The associated conditional tail copulas are depicted on Figure 1.

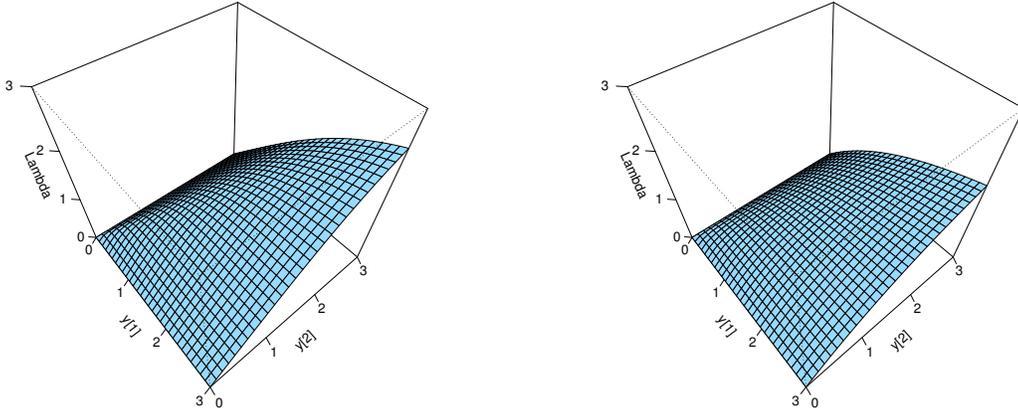


Figure 1: Representations of the conditional tail copula $\Lambda(y_1, y_2|x)$ for $(y_1, y_2) \in [0, 3]^2$ at $x = 0.1$ (left panel) and $x = 0.5$ (right panel).

4.2 Hyper-parameters selection

The considered conditional tail copula estimator $\check{\Lambda}_{h,h}(\cdot|x)$ depends on two hyper-parameters h and α . We propose to select these parameters by using an homogeneity criterion

$$\mathcal{H}(h, \alpha) := \sum_{t \in \mathcal{T}} (\check{\Lambda}_{h,h}(t\alpha|x) - t\check{\Lambda}_{h,h}(\alpha|x))^2, \quad (15)$$

where $\mathcal{T} \subset (0, \infty)$ is a finite set. The selected parameters h_{sel} and α_{sel} are given by

$$(h_{sel}, \alpha_{sel}) := \arg \min_{\alpha \in \mathcal{A}, h \in \mathcal{B}} \mathcal{H}(h, \alpha),$$

where \mathcal{A} and \mathcal{B} are finite sets of possible values for α and h respectively. In what follows, we take $\mathcal{T} = \{1/3, 2/3, 1, 4/3, 5/3\}$, $\mathcal{A} = \{0.05, 0.06, \dots, 0.5\}$ and $\mathcal{B} = \{0.03, 0.04, \dots, 0.33\}$.

4.3 Results

In this experiment, $N = 100$ independent copies of a sample of size $n = 1000$ are simulated from the model presented in Paragraph 4.1. In each situation, the median as well as the mean-squared error of the N estimations of $\Lambda(y_1, y_2|x)$ are represented for $(y_1, y_2) \in [0, 3]^2$ respectively on Figures 2 and 3. In all cases, the estimations of Figure 2 allow to visually recover the shape of the true conditional tail copula, see Figure 1. Focussing on Figure 3, it appears that the mean-squared error gets larger when the conditional tail copula increases. This phenomena can be explained thanks to Theorem 2(ii): The asymptotic variance is proportional to the conditional tail copula. One can also note that the Case 1 where the margins do not depend on the covariate leads to smaller mean-squared errors than Case 2. Finally, the “boundary effect” due to kernel estimation does not appear in this simulated example. In other situations, one may consider the use of mirror-reflection type estimators [19] or kernel local linear estimators [8].

5 Illustration on real data

In this section, we illustrate our approach on the so-called `pima-indians-diabetes` dataset¹ consisting of nine variables measured on $n = 768$ subjects aged from 21 to 81 years. This population was studied in order to understand the high prevalence of diabetes in the Pima Indian population. For visualization sakes, only two variables are considered: the diastolic blood pressure (DBP, denoted by Y_1) and the body mass index (BMI, denoted by Y_2). The age is considered as the univariate covariable X . It is indeed suspected that high values of DBP and BMI may represent significant risk factors for diabetes [7]. The estimation of the conditional tail copula $\Lambda(y_1, y_2|x)$ can show how the dependence structure between high values of DBP and BMI may be affected by the age.

Similarly to Section 4, the conditional tail copula estimator $\check{\Lambda}_{h,h}(\cdot|x)$ is used with the bi-quadratic kernel and an unique bandwidth h . This bandwidth h and the sample fraction α are selected using the procedure described in Paragraph 4.2 with $\mathcal{T} = \{1/3, 2/3, 1, 4/3, 5/3\}$, $\mathcal{A} = \{0.05, 0.06, \dots, 0.5\}$ and $\mathcal{B} = \{1, 2, \dots, 10\}$. The resulting conditional tail copula estimator $\check{\Lambda}_{h,h}(y_1, y_2|x)$ is depicted for $(y_1, y_2) \in [0, 3]^2$ at $x = 30$ years and $x = 60$ years on Figure 4. It appears that the dependence between large values of BMI and DBP is weaker for young subjects than for old ones.

¹<http://ftp.ics.uci.edu/pub/machine-learning-databases/pima-indians-diabetes>

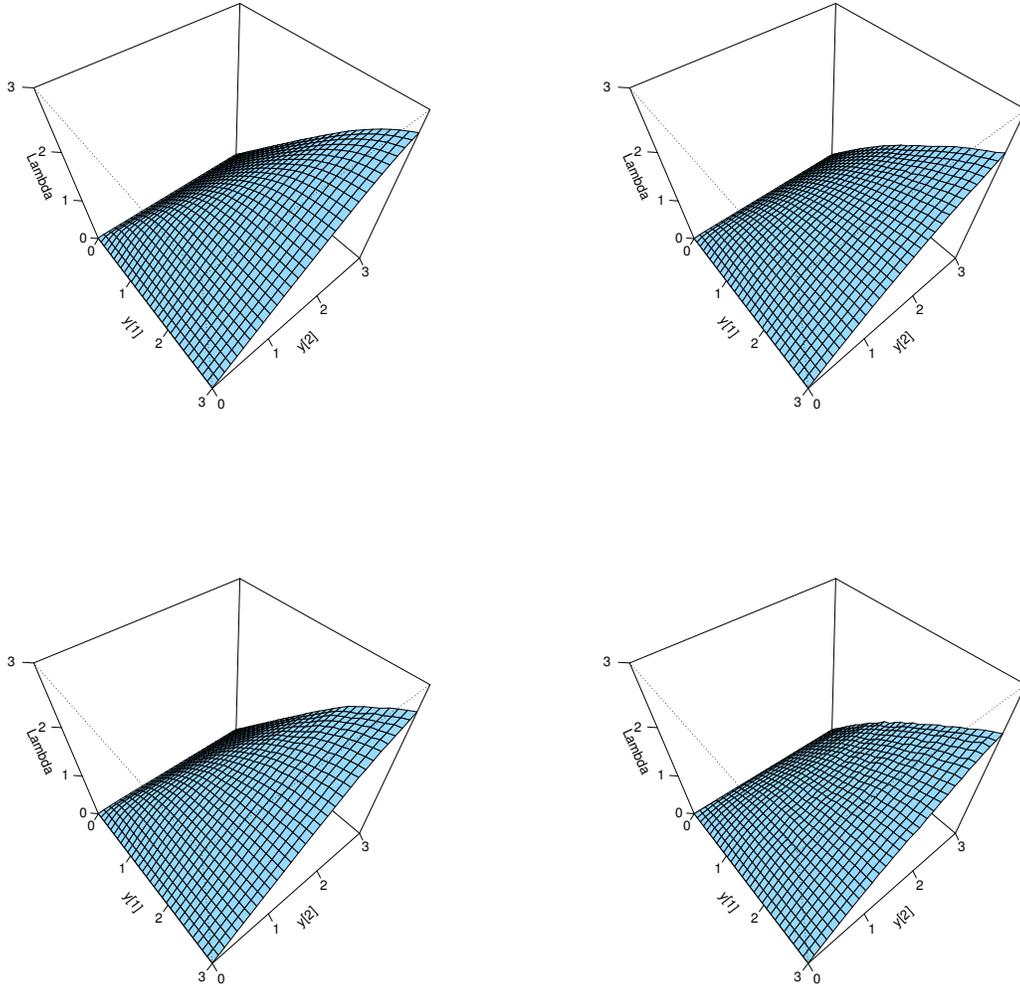


Figure 2: Representation of the median of the conditional tail copula estimates for $(y_1, y_2) \in [0, 3]^2$ in Case 1 (first row) and Case 2 (second row) at $x = 0.1$ (left) and $x = 0.5$ (right).

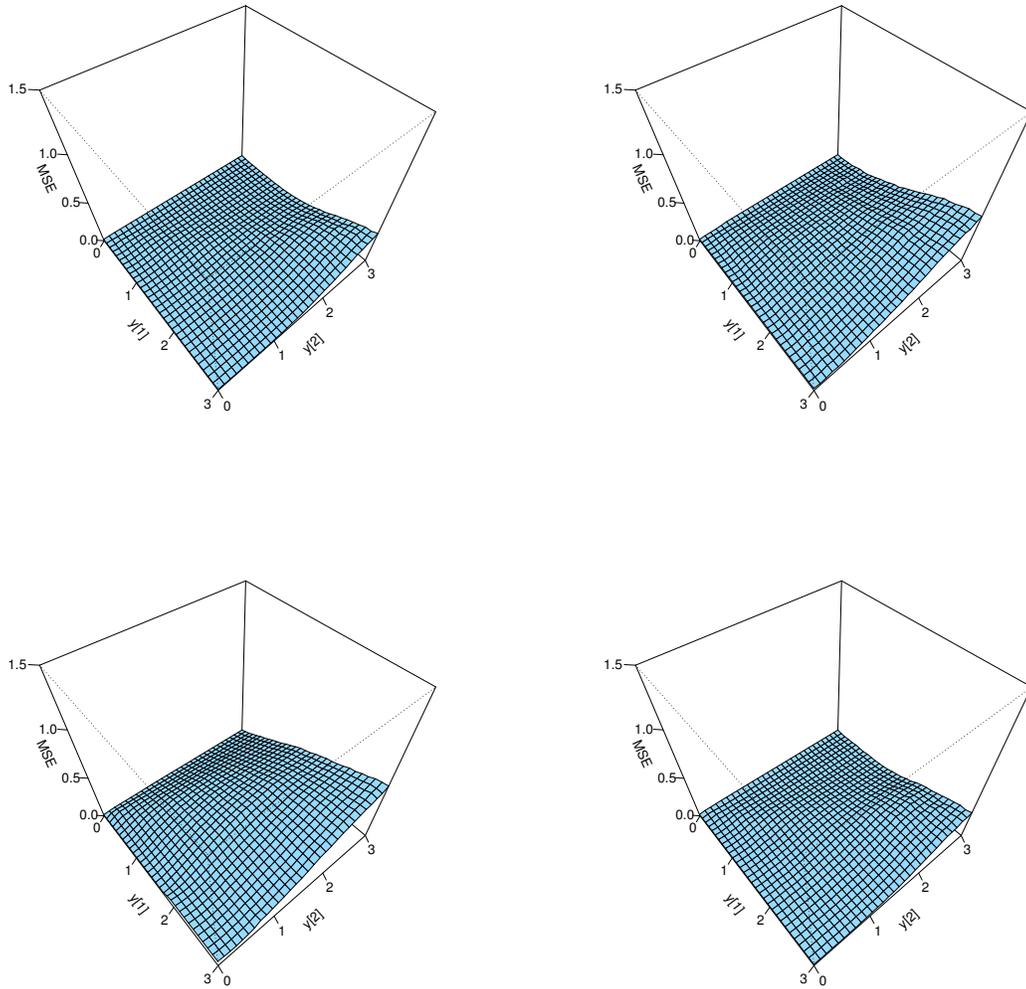


Figure 3: Representation of the mean-squared error of the conditional tail copula estimates for $(y_1, y_2) \in [0, 3]^2$ in Case 1 (first row) and Case 2 (second row) at $x = 0.1$ (left) and $x = 0.5$ (right).

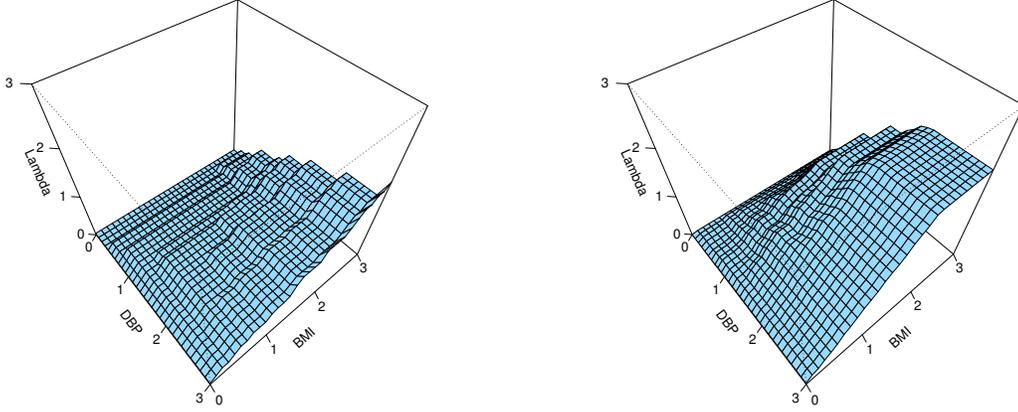


Figure 4: Representation of the conditional tail copula estimator $\check{\Lambda}_{h,h}(y_1, y_2|x)$ for $(y_1, y_2) \in [0, 3]^2$ at $x = 30$ years (left panel) and $x = 60$ years (right panel).

6 Proofs

6.1 Preliminary results

The first lemma is a standard result on the density kernel estimator (see Parzen [29] for a proof).

Lemma 1. *Under H.2, if $nh^p \rightarrow \infty$ then for all $x \in \mathbb{R}^p$ such that $g(x) \neq 0$,*

$$|\hat{g}_h(x) - g(x)| = \mathcal{O}(h) + \mathcal{O}_{\mathbb{P}}\left((nh^p)^{-1/2}\right).$$

The next result is dedicated to the random variable given by $\tilde{S}_h(u|x) - \kappa\tilde{S}_h(v|x)$, where $\kappa \in \{0, 1\}$ and $(u, v) \in (0, 1)^d \times (0, 1)^d$ with $v \leq u$. Clearly, if the margins are known, it can be seen as an estimator of $S(u|x) - \kappa S(v|x)$. The asymptotic normality of this random variable is established in the following result.

Lemma 2. *Suppose H.2 holds and let $x \in \mathbb{R}^p$ such that $g(x) \neq 0$. Consider two sequences $u_n \in (0, 1)^d$ and $v_n \in (0, 1)^d$ converging to zero as $n \rightarrow \infty$ such that $nh^p(S(u_n|x) - \kappa S(v_n|x)) \rightarrow \infty$. For some $\zeta > 0$, if there exists a sequence ν_n such that $(1 - \zeta)\nu_n \leq v_n < u_n < (1 + \zeta)\nu_n$ with*

$$nh^p(S(u_n|x) - \kappa S(v_n|x)) \left(h \vee \frac{S(u_n|x) + \kappa S(v_n|x)}{S(u_n|x) - \kappa S(v_n|x)} w_{h,\zeta}(\nu_n|x) \right)^2 \rightarrow 0,$$

then,

$$(nh^p(S(u_n|x) - \kappa S(v_n|x)))^{1/2} \left(\frac{\tilde{S}_h(u_n|x) - \kappa \tilde{S}_h(v_n|x)}{S(u_n|x) - \kappa S(v_n|x)} - 1 \right) \xrightarrow{d} \mathcal{N} \left(0, \frac{\|K\|_2^2}{g(x)} \right).$$

The proof of this lemma is postponed to the Appendix. Note that if $\kappa = 0$, one can take $u_n = v_n$ and thus Lemma 2 entails that

$$(nh^p S(u_n|x))^{1/2} \left(\frac{\tilde{S}_h(u_n|x)}{S(u_n|x)} - 1 \right) \xrightarrow{d} \mathcal{N} \left(0, \frac{\|K\|_2^2}{g(x)} \right),$$

and the conditions can be simplified as $nh^p S(u_n|x) \rightarrow \infty$ and $nh^p S(u_n|x) (h \vee w_{h,\zeta}(u_n|x))^2 \rightarrow 0$ as $n \rightarrow \infty$. Furthermore, in the particular case where $u_n = \mathbf{1} + (\bar{F}_i(z_n|x) - 1)e_i$, $z_n \in (0, 1)$, the previous convergence result establishes the asymptotic normality of the kernel estimators of the conditional survival functions $\hat{F}_{h,i}(z_n|x)$. Note that this result has also been established in [9, Theorem 1] and [10, Proposition 1]. The case of functional covariates is addressed in [17]. Mimicking the proof of Corollary 1 in [10] thus permits to derive the asymptotic normality of the kernel estimators of the extreme conditional quantiles $\bar{F}_i^{\leftarrow}(\beta_n|x)$.

Lemma 3. *Suppose **H.2**, **H.3** hold and let $x \in \mathbb{R}^p$ such that $g(x) \neq 0$. Consider a sequence $\beta_n \in (0, 1)$ such that, as $n \rightarrow \infty$, $nh^p \beta_n \rightarrow \infty$ and $nh^p \beta_n (h \vee \bar{w}_{h,\zeta}^{(i)}(\beta_n|x))^2 \rightarrow 0$ for some $\zeta > 0$. Then,*

$$(nh^p \beta_n)^{1/2} \frac{\bar{F}_i^{\leftarrow}(\beta_n|x)}{a_i(\bar{F}_i^{\leftarrow}(\beta_n|x)|x)} \left(\frac{\hat{\bar{F}}_{h,i}^{\leftarrow}(\beta_n|x)}{\bar{F}_i^{\leftarrow}(\beta_n|x)} - 1 \right) \xrightarrow{d} \mathcal{N} \left(0, \frac{\|K\|_2^2}{g(x)} \right).$$

The last lemma will be useful to deal with differences between two indicator functions. It is a straightforward consequence of the monotonicity of the indicator function.

Lemma 4. *Let y, u, v, m and M be vectors of \mathbb{R}^d with, for all $i = 1, \dots, d$, $M_i > u_i \vee v_i$ and $m_i < u_i \wedge v_i$. Then,*

$$\begin{aligned} & |\mathbb{I}\{y_i > u_i, i = 1, \dots, d\} - \mathbb{I}\{y_i > v_i, i = 1, \dots, d\}| \\ & \leq \mathbb{I}\{y_i > m_i, i = 1, \dots, d\} - \mathbb{I}\{y_i > M_i, i = 1, \dots, d\}. \end{aligned}$$

6.2 Proof of main results

Proof of Theorem 1 – First, taking $u_n = \alpha y$ and $\kappa = 0$ in Lemma 2, one has under the assumptions of Theorem 1

$$(nh^p \alpha)^{1/2} \left(\frac{\tilde{S}_h(\alpha y|x)}{S(\alpha y|x)} - 1 \right) \xrightarrow{d} \mathcal{N} \left(0, \frac{\|K\|_2^2}{g(x)\Lambda(y|x)} \right), \quad (16)$$

since $S(\alpha y|x) = \alpha \Lambda(y|x)(1+o(1))$. Furthermore, $(nh^p \alpha)^{1/2}(\tilde{\Lambda}_n(y|x) - \Lambda(y|x)) = T_{1,n} + T_{2,n}$, where

$$\begin{aligned} T_{1,n} &= \left(\frac{nh^p}{\alpha}\right)^{1/2} S(\alpha y|x) \left(\frac{\tilde{S}_h(\alpha y|x)}{S(\alpha y|x)} - 1\right) \\ \text{and } T_{2,n} &= \left(\frac{nh^p}{\alpha}\right)^{1/2} S(\alpha y|x) \left(1 - \frac{\alpha \Lambda(y|x)}{S(\alpha y|x)}\right). \end{aligned}$$

A direct use of (16) entails that

$$T_{1,n} \xrightarrow{d} \mathcal{N}\left(0, \frac{\Lambda(y|x) \|K\|_2^2}{g(x)}\right).$$

Finally, from **H.1**,

$$T_{2,n} = \mathcal{O}\left((nh^p \alpha)^{1/2} b(\alpha^{-1}|x)\right) = o(1),$$

which concludes the proof. \blacksquare

Proof of Theorem 2 – We start with the decomposition

$$\hat{\Lambda}_h(y|x) - \Lambda(y|x) = \left(\tilde{\Lambda}_h(y|x) - \Lambda(y|x)\right) + \left(\hat{\Lambda}_h(y|x) - \tilde{\Lambda}_h(y|x)\right).$$

From Theorem 1, the first term of the right-hand side is equal to $(nh^p \alpha)^{-1/2} \xi_n$, where ξ_n converges in distribution to a centred Gaussian variable with variance $\|K\|_2^2 \Lambda(y|x)/g(x)$. Let us focus on the second term. From **H.4**, for all $\varepsilon > 0$, there exists a positive constant c_ε such that $\mathbb{P}(\mathcal{A}_n) < \varepsilon/2$ where

$$\mathcal{A}_n = \left\{ \eta_n \bigvee_{i=1}^d \frac{\bar{F}_i^{\leftarrow}(y_i \alpha|x)}{a_i(\bar{F}_i^{\leftarrow}(y_i \alpha|x)|x)} \left| \frac{\hat{q}_i(y_i \alpha|x)}{\bar{F}_i^{\leftarrow}(y_i \alpha|x)} - 1 \right| \geq c_\varepsilon \right\}.$$

Furthermore, denoting by $\tilde{q}_n(\vartheta) = (\tilde{q}_{n,1}(\vartheta), \dots, \tilde{q}_{n,d}(\vartheta))^T$ with, for $i = 1, \dots, d$,

$$\tilde{q}_{n,i}(\vartheta) := \bar{F}_i^{\leftarrow}(y_i \alpha|x) + \vartheta \eta_n^{-1} a_i(\bar{F}_i^{\leftarrow}(y_i \alpha|x)|x),$$

Lemma 4 entails that under \mathcal{A}_n^C one has for all $j = 1, \dots, n$:

$$|\mathbb{I}\{Y_j \geq \hat{q}_n(\alpha y|x)\} - \mathbb{I}\{Y_j \geq q(\alpha y|x)\}| \leq \mathbb{I}\{Y_j \geq \tilde{q}_n(-c_\varepsilon)\} - \mathbb{I}\{Y_j \geq \tilde{q}_n(c_\varepsilon)\}.$$

Thus, under \mathcal{A}_n^C ,

$$\begin{aligned} \left| \hat{\Lambda}_h(y|x) - \tilde{\Lambda}_h(y|x) \right| &\leq \frac{1}{nh^p \alpha \hat{g}_h(x)} \sum_{j=1}^n |\mathbb{I}\{Y_j \geq \tilde{q}_n(-c_\varepsilon)\} - \mathbb{I}\{Y_j \geq \tilde{q}_n(c_\varepsilon)\}| K\left(\frac{x - X_j}{h}\right) \\ &= \alpha^{-1} \left(\tilde{S}_h(\tilde{u}_n(-c_\varepsilon)|x) - \tilde{S}_h(\tilde{u}_n(c_\varepsilon)|x) \right), \end{aligned}$$

with $\tilde{u}_n(\vartheta) = (\bar{F}_1(\tilde{q}_{n,1}(\vartheta)|x), \dots, \bar{F}_d(\tilde{q}_{n,d}(\vartheta)|x))^T$. Let us now treat the difference $\tilde{S}_h(\tilde{u}_n(-c_\varepsilon)|x) - \tilde{S}_h(\tilde{u}_n(c_\varepsilon)|x)$ using Lemma 2 with $\kappa = 1$, $u_n = \tilde{u}_n(-c_\varepsilon)$ and $v_n = \tilde{u}_n(c_\varepsilon)$. We thus need to check the following three conditions: there exists $\zeta > 0$ and a sequence ν_n such that

$$(1 - \zeta)\nu_n \leq \tilde{u}_n(c_\varepsilon) < \tilde{u}_n(-c_\varepsilon) \leq (1 + \zeta)\nu_n, \quad (17)$$

$$nh^p(S(\tilde{u}_n(-c_\varepsilon)|x) - S(\tilde{u}_n(c_\varepsilon)|x)) \rightarrow \infty, \quad (18)$$

and

$$nh^p(S(\tilde{u}_n(-c_\varepsilon)|x) - S(\tilde{u}_n(c_\varepsilon)|x)) \left(h \vee \frac{S(\tilde{u}_n(-c_\varepsilon)|x) + S(\tilde{u}_n(c_\varepsilon)|x)}{S(\tilde{u}_n(-c_\varepsilon)|x) - S(\tilde{u}_n(c_\varepsilon)|x)} w_{h,\zeta}(\nu_n|x) \right)^2 \rightarrow 0. \quad (19)$$

To this end, asymptotic equivalents of the sequences $\tilde{u}_n(\pm c_\varepsilon)$ and $S(\tilde{u}_n(-c_\varepsilon)|x) - S(\tilde{u}_n(c_\varepsilon)|x)$ are required. First, note that a Taylor expansion yields

$$\begin{aligned} S(\tilde{u}_n(\pm c_\varepsilon)|x) &= S(\alpha y|x) \\ &+ (\tilde{u}_n(\pm c_\varepsilon) - \alpha y)^T \nabla S(\alpha y + \theta(\tilde{u}_n(\pm c_\varepsilon) - \alpha y)|x) \end{aligned} \quad (20)$$

where $\theta \in (0, 1)$. Next, for $i = 1, \dots, d$, another first order Taylor expansion entails that, for all $i = 1, \dots, d$,

$$\begin{aligned} \tilde{u}_{n,i}(\pm c_\varepsilon) - \alpha y_i &= \mp F'_i \left(\tilde{q}_{n,i}(\pm \tilde{\theta}_i c_\varepsilon) \middle| x \right) c_\varepsilon \eta_n^{-1} a_i(\bar{F}_i^{\leftarrow}(\alpha y_i|x)|x) \\ &= \mp F'_i \left(\bar{F}_i^{\leftarrow}(\tilde{u}_{n,i}(\pm \tilde{\theta}_i c_\varepsilon)|x) \middle| x \right) c_\varepsilon \eta_n^{-1} a_i(\bar{F}_i^{\leftarrow}(\alpha y_i|x)|x), \end{aligned}$$

for $\tilde{\theta}_i \in (0, 1)$. Since for all $i = 1, \dots, d$, equation (13) entails that $\tilde{u}_{n,i}(\pm \tilde{\theta}_i c_\varepsilon) = y_i \alpha (1 + o(1))$ and since the function $y \mapsto F'_i(\bar{F}_i^{\leftarrow}(y|x)|x)$ is a regularly varying function at zero with index $\gamma_i(x) + 1$ (see [11], Corollary 1.1.10), one has

$$\tilde{u}_{n,i}(\pm c_\varepsilon) - \alpha y_i = \mp F'_i(\bar{F}_i^{\leftarrow}(y_i \alpha|x)|x) c_\varepsilon \eta_n^{-1} a_i(\bar{F}_i^{\leftarrow}(\alpha y_i|x)|x)(1 + o(1)).$$

Now, in view of [11], Theorem 1.2.6 and Remark 1.2.7, for $i = 1, \dots, d$,

$$a_i(y|x) = \frac{\bar{F}_i(y|x)}{F'_i(y|x)}(1 + o(1)) \text{ as } y \rightarrow y_{F_i}(x),$$

and thus, for $i = 1, \dots, d$,

$$\tilde{u}_{n,i}(\pm c_\varepsilon) - \alpha y_i = \mp \alpha c_\varepsilon \eta_n^{-1} y_i (1 + o(1)). \quad (21)$$

Then, condition (17) is satisfied with $\nu_n = \alpha y$. Furthermore, collecting (20) and (21) leads to

$$\begin{aligned} S(\tilde{u}_n(-c_\varepsilon)|x) - S(\tilde{u}_n(c_\varepsilon)|x) &= \alpha c_\varepsilon \eta_n^{-1} y^t \left\{ \begin{aligned} &\nabla S(\alpha y + \theta(\tilde{u}_n(-c_\varepsilon) - \alpha y)|x) \\ &+ \nabla S(\alpha y + \theta(\tilde{u}_n(c_\varepsilon) - \alpha y)|x) \end{aligned} \right\} (1 + o(1)). \end{aligned}$$

Then, under **H.5**,

$$S(\tilde{u}_n(-c_\varepsilon)|x) - S(\tilde{u}_n(c_\varepsilon)|x) = 2c_\varepsilon \Lambda(y|x) \alpha \eta_n^{-1} (1 + o(1)), \quad (22)$$

and, in view of $nh^p \alpha_n \eta_n^{-1} \rightarrow \infty$, condition (18) is clearly satisfied. Finally, since $S(\tilde{u}_n(\pm c_\varepsilon)|x) = \alpha \Lambda(y|x)(1 + o(1))$, one has

$$\frac{S(\tilde{u}_n(-c_\varepsilon)|x) + S(\tilde{u}_n(c_\varepsilon)|x)}{S(\tilde{u}_n(-c_\varepsilon)|x) - S(\tilde{u}_n(c_\varepsilon)|x)} = \mathcal{O}_{\mathbb{P}}(\eta_n),$$

and thus, condition (14) implies (19). Lemma 2 can now be used with (22) to obtain

$$\tilde{S}_h(\tilde{u}_n(-c_\varepsilon)|x) - \tilde{S}_h(\tilde{u}_n(c_\varepsilon)|x) = 2c_\varepsilon\Lambda(y|x)\alpha\eta_n^{-1}(1 + o_{\mathbb{P}}(1)).$$

Hence, for all $\varepsilon > 0$, there exists a positive constant c'_ε such that

$$\mathbb{P}\left(\eta_n^{-1}\left|\hat{\Lambda}_h(y|x) - \tilde{\Lambda}_n(y|x)\right| > c'_\varepsilon \mid \mathcal{A}_n^C\right) \leq \frac{\varepsilon}{2},$$

and thus

$$\mathbb{P}\left(\eta_n^{-1}\left|\hat{\Lambda}_h(y|x) - \tilde{\Lambda}_n(y|x)\right| > c'_\varepsilon\right) \leq \frac{\varepsilon}{2} + \mathbb{P}(\mathcal{A}_n) \leq \varepsilon,$$

which concludes the proof. ■

Proof of Corollary 1 – Under the assumptions of Lemma 3, the estimator $\hat{q}_n(u|x)$ defined in (8) satisfies condition **H.4** with $\eta_n = (nh^p\alpha)^{1/2}$. Applying Theorem 2(i) concludes the proof. ■

Proof of Corollary 2 – Under the assumptions of Lemma 3, the estimator $\hat{q}_n(u|x)$ defined in (8) satisfies condition **H.4** with $\eta_n = (nk^p\alpha)^{1/2}$. Applying Theorem 2(ii) concludes the proof. ■

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Appendix

Proof of Lemma 2 – Let us introduce the random variable

$$\tilde{\Psi}_{h,\kappa}(u_n, v_n|x) := \hat{g}_h(x) \left(\tilde{S}_h(u_n|x) - \kappa \tilde{S}_h(v_n|x) \right),$$

where $\kappa \in \{0, 1\}$. We first focus on its expectation. Since (Y_1, \dots, Y_n) are identically distributed, one has

$$\begin{aligned} \mathbb{E}(\tilde{\Psi}_{h,\kappa}(u_n, v_n|x)) &= h^{-p} \mathbb{E} \left\{ [\mathbb{I}\{Y \geq q(u_n|x)\} - \kappa \mathbb{I}\{Y \geq q(v_n|x)\}] K \left(\frac{x-X}{h} \right) \right\} \\ &= h^{-p} \int_{\mathbb{R}^p} [\bar{F}(q(u_n|x)|t) - \kappa \bar{F}(q(v_n|x)|t)] K \left(\frac{x-t}{h} \right) g(t) dt \\ &= \int_{\mathcal{S}} [\bar{F}(q(u_n|x)|x-hs) - \kappa \bar{F}(q(v_n|x)|x-hs)] K(s) g(x-hs) ds, \end{aligned}$$

and therefore

$$\begin{aligned} \left| \frac{\mathbb{E}(\tilde{\Psi}_{h,\kappa}(u_n, v_n|x))}{S(u_n|x) - \kappa S(v_n|x)} - g(x) \right| &\leq \int_{\mathcal{S}} K(s) |g(x-hs) - g(x)| ds \\ &+ \int_{\mathcal{S}} K(s) \left| \frac{\bar{F}(q(u_n|x)|x-hs) - \kappa \bar{F}(q(v_n|x)|x-hs)}{S(u_n|x) - \kappa S(v_n|x)} - 1 \right| g(x-hs) ds. \end{aligned}$$

Clearly, under **H.2**, the first term of the right hand-side of the previous inequality is a $\mathcal{O}(h)$. Moreover, under **H.2**,

$$\int_{\mathcal{S}} K(s) g(x-hs) ds = g(x)(1 + o(1)),$$

and

$$\left| \frac{\bar{F}(q(u_n|x)|x-hs) - \kappa \bar{F}(q(v_n|x)|x-hs)}{S(u_n|x) - \kappa S(v_n|x)} - 1 \right| \leq \frac{S(u_n|x) + \kappa S(v_n|x)}{S(u_n|x) - \kappa S(v_n|x)} \omega_{h,\zeta}(v_n|x).$$

Hence, the second term is a

$$\mathcal{O} \left(\frac{S(u_n|x) + \kappa S(v_n|x)}{S(u_n|x) - \kappa S(v_n|x)} \omega_{h,\zeta}(v_n|x) \right).$$

As a first conclusion, one has that $\mathbb{E}(\tilde{\Psi}_{h,\kappa}(u_n, v_n|x))$ is equal to:

$$g(x)(S(u_n|x) - \kappa S(v_n|x)) \left(1 + \mathcal{O}(h) + \mathcal{O} \left(\frac{S(u_n|x) + \kappa S(v_n|x)}{S(u_n|x) - \kappa S(v_n|x)} \omega_{h,\zeta}(v_n|x) \right) \right). \quad (23)$$

Let us now focus on the variance of $\tilde{\Psi}_{h,\kappa}(u_n, v_n|x)$. First, remark that

$$\tilde{\Psi}_{h,\kappa}(u_n, v_n|x) - \mathbb{E}(\tilde{\Psi}_{h,\kappa}(u_n, v_n|x)) = \sum_{j=1}^n Z_{j,n},$$

where

$$Z_{j,n} := \frac{1}{nh^p} \left([\mathbb{I}\{Y_j \geq q(u_n|x)\} - \kappa \mathbb{I}\{Y_j \geq q(v_n|x)\}] K \left(\frac{x-X_j}{h} \right) - h^p \mathbb{E}(\tilde{\Psi}_{h,\kappa}(u_n, v_n|x)) \right),$$

are independent, centered and identically distributed random variables. Remarking that for all $\tau > 0$, $[\mathbb{I}\{Y_j \geq q(u_n|x)\} - \kappa\mathbb{I}\{Y_j \geq q(v_n|x)\}]^\tau = \mathbb{I}\{Y_j \geq q(u_n|x)\} - \kappa\mathbb{I}\{Y_j \geq q(v_n|x)\}$, the variance of $\tilde{\Psi}_{h,\kappa}(u_n, v_n|x)$ is equal to:

$$\frac{\|K\|_2^2}{nh^{2p}} \left(\mathbb{E} \left([\mathbb{I}\{Y \geq q(u_n|x)\} - \kappa\mathbb{I}\{Y \geq q(v_n|x)\}] Q \left(\frac{x-X}{h} \right) \right) - \frac{h^{2p}}{\|K\|_2^2} \mathbb{E}^2(\tilde{\Psi}_{h,\kappa}(u_n, v_n|x)) \right),$$

where $Q := K^2/\|K\|_2^2$ has the same properties as K . Thus, (23) entails

$$\begin{aligned} \text{Var}(\tilde{\Psi}_h(u_n|x)) &= \frac{\|K\|_2^2 g(x)}{nh^p} (S(u_n|x) - \kappa S(v_n|x)) \\ &\times \left(1 + o(1) - \frac{g(x)}{\|K\|_2^2} h^p (S(u_n|x) - \kappa S(v_n|x)) (1 + o(1)) \right), \end{aligned}$$

and, in view of $h^p(S(u_n|x) - \kappa S(v_n|x)) \leq h^p \rightarrow 0$, one has

$$\text{Var}(\tilde{\Psi}_h(u_n|x)) = \frac{\|K\|_2^2 g(x)}{nh^p} (S(u_n|x) - \kappa S(v_n|x)) (1 + o(1)). \quad (24)$$

We now use the Lyapounov theorem to establish the asymptotic normality of $\tilde{\Psi}_{h,\kappa}(u_n, v_n|x)$. From (24), the Lyapounov condition reduces to:

$$\lim_{n \rightarrow \infty} \left(\frac{nh^p}{S(u_n|x) - \kappa S(v_n|x)} \right)^{1+\delta/2} n \mathbb{E}|Z_{1,n}|^{2+\delta} = 0, \quad (25)$$

for some $\delta > 0$. Hence, if (25) is satisfied, the following convergence holds

$$\left(\frac{nh^p}{S(u_n|x) - \kappa S(v_n|x)} \right)^{1/2} \left(\tilde{\Psi}_{h,\kappa}(u_n, v_n|x) - \mathbb{E}(\tilde{\Psi}_{h,\kappa}(u_n, v_n|x)) \right) \xrightarrow{d} \mathcal{N}(0, g(x)\|K\|_2^2). \quad (26)$$

To prove (25), remark that for two random variables T_1 and T_2 , $\mathbb{E}|T_1 + T_2|^{2+\delta} \leq 2^{2+\delta}(\mathbb{E}|T_1|^{2+\delta} \vee \mathbb{E}|T_2|^{2+\delta})$. Thus, one has

$$\mathbb{E}|Z_{1,n}|^{2+\delta} \leq \left(\frac{2\|K\|_{2+\delta}}{nh^p} \right)^{2+\delta} \mathbb{E} \left([\mathbb{I}\{Y \geq q(u_n|x)\} - \kappa\mathbb{I}\{Y \geq q(v_n|x)\}] N \left(\frac{x-X}{h} \right) \right),$$

where $N := (K/\|K\|_{2+\delta})^{2+\delta}$ is a bounded density function on \mathbb{R}^p with same support as K . Using (23) yields

$$\mathbb{E}|Z_{1,n}|^{2+\delta} \leq \left(\frac{2\|K\|_{2+\delta}}{nh^p} \right)^{2+\delta} h^p (S(u_n|x) - \kappa S(v_n|x)) (1 + o(1)).$$

Hence, condition (25) is satisfied and convergence (26) holds. Finally, remark that

$$(nh^p(S(u_n|x) - \kappa S(v_n|x)))^{1/2} \left(\frac{\tilde{S}_h(u_n|x) - \kappa\tilde{S}_h(v_n|x)}{S(u_n|x) - \kappa S(v_n|x)} - 1 \right) = T_{1,n} + T_{2,n} + T_{3,n},$$

where

$$\begin{aligned}
T_{1,n} &:= \frac{(nh^p(S(u_n|x) - \kappa S(v_n|x)))^{1/2} \tilde{\Psi}_{h,\kappa}(u_n, v_n|x) - \mathbb{E}(\tilde{\Psi}_{h,\kappa}(u_n, v_n|x))}{\hat{g}_h(x) S(u_n|x) - \kappa S(v_n|x)}, \\
T_{2,n} &:= \frac{(nh^p(S(u_n|x) - \kappa S(v_n|x)))^{1/2} \mathbb{E}(\tilde{\Psi}_{h,\kappa}(u_n, v_n|x)) - g(x)(S(u_n|x) - \kappa S(v_n|x))}{\hat{g}_h(x) S(u_n|x) - \kappa S(v_n|x)}, \\
\text{and } T_{3,n} &:= \frac{(nh^p(S(u_n|x) - \kappa S(v_n|x)))^{1/2}}{\hat{g}_h(x)} (g(x) - \hat{g}_h(x)).
\end{aligned}$$

From (26), $\hat{g}_h(x)T_{1,n} \xrightarrow{d} \mathcal{N}(0, g(x)\|K\|_2^2)$. Since from Lemma 1, $\hat{g}_h(x) \xrightarrow{\mathbb{P}} g(x)$, one has that $T_{1,n} \xrightarrow{d} \mathcal{N}(0, \|K\|_2^2/g(x))$. Furthermore, (23) implies that

$$T_{2,n} = \frac{g(x)}{\hat{g}_h(x)} (nh^p(S(u_n|x) - \kappa S(v_n|x)))^{1/2} \left(\mathcal{O}(h) + \mathcal{O} \left(\frac{S(u_n|x) + \kappa S(v_n|x)}{S(u_n|x) - \kappa S(v_n|x)} w_{h,\zeta}(u_n|x) \right) \right) = o(1),$$

by assumptions. Finally, Lemma 1 entails that $T_{3,n} = o_{\mathbb{P}}(1)$ which completes the proof. \blacksquare