



**HAL**  
open science

## Bregman superquantiles. Estimation methods and applications.

Tatiana Labopin-Richard, Fabrice Gamboa, Aurélien Garivier, Bertrand Iooss

► **To cite this version:**

Tatiana Labopin-Richard, Fabrice Gamboa, Aurélien Garivier, Bertrand Iooss. Bregman superquantiles. Estimation methods and applications.. 2014. hal-00996440v5

**HAL Id: hal-00996440**

**<https://hal.science/hal-00996440v5>**

Preprint submitted on 19 Nov 2014 (v5), last revised 6 Jan 2016 (v8)

**HAL** is a multi-disciplinary open access archive for the deposit and dissemination of scientific research documents, whether they are published or not. The documents may come from teaching and research institutions in France or abroad, or from public or private research centers.

L'archive ouverte pluridisciplinaire **HAL**, est destinée au dépôt et à la diffusion de documents scientifiques de niveau recherche, publiés ou non, émanant des établissements d'enseignement et de recherche français ou étrangers, des laboratoires publics ou privés.

# BREGMAN SUPERQUANTILES. ESTIMATION METHODS AND APPLICATIONS

T. LABOPIN-RICHARD, F. GAMBOA, A. GARIVIER, AND B. IOOSS

ABSTRACT. In this work, we extend some quantities introduced by Rockafellar to the case where the proximity between real numbers is measured by using a Bregman divergence. This leads to the definition of the Bregman superquantile. Axioms of a coherent measure of risk discussed previously by Rockafellar are studied in the case of Bregman superquantile. Furthermore, we deal with asymptotic properties of a Monte Carlo estimator of the Bregman superquantile. Several numerical tests confirm the theoretical results and an application illustrates the potential interests of the Bregman superquantile.

## 1. INTRODUCTION

**1.1. Aim and scope.** The aim of this article is to define and to study properties and estimation procedures for Bregman extension of the superquantile defined in [25] (see also [18], [21] and references therein). In the introduction we first recall the necessary conditions for a measure of risk to be coherent. Further in Section 2 we present the superquantile as a partial response to this problem. We also introduce the Bregman superquantile and study axioms of a coherent measure of risk for this quantity. In Section 3 we are interested in estimating this Bregman superquantile, we introduce a plug-in estimator and study its convergence and asymptotic normality. All the proofs are deferred to Section 4. Some numerical simulations are shown in Section 5. An application on real data of radiological exposure is given in Section 6.

**1.2. Coherent measures of risk.** Let  $X$  be a real-valued random variable and let  $F_X$  be its cumulative distribution function. We define for  $u \in ]0, 1[$ , the quantile function

$$F_X^{-1}(u) := \inf\{x : F_X(x) \geq u\}.$$

A usual way to quantify the risk associated with  $X$  is to consider, for a given number  $\alpha \in ]0, 1[$  close to 1, its lower quantile  $q_\alpha^X := F_X^{-1}(\alpha)$ .

Nevertheless, the quantile is not a sub-additive function of  $X$ , a property considered important in some applications (e.g finance, see [2]). Thus Rockafellar introduces in [20] a new quantity called therein superquantile that satisfies this property. The superquantile is defined by

$$Q_\alpha = \mathbb{E}(X|X \geq q_\alpha) = \mathbb{E}(X|X \geq F_X^{-1}(\alpha)) = \mathbb{E}\left(\frac{X\mathbf{1}_{X \geq F_X^{-1}(\alpha)}}{1 - \alpha}\right).$$

Notice that this quantity is also called conditional value at risk in older references ([25], [23], [21], ...).

Sub-additivity is not the sole interesting property for a measure of risk (for example for financial application). Following Rockaffellar in [20] we define:

**Definition 1.1.** *Let  $\mathcal{R}$  be a measure of risk that is a numerical function defined on random variables and  $X$  and  $X'$  be two real-valued random variables. We say that  $\mathcal{R}$  is coherent if, and only if, it satisfies the five following properties :*

- i) *Constant invariance : let  $C \in \mathbb{R}$ , If  $X = C$  (a.s.) then  $\mathcal{R}(C) = C$ .*
- ii) *Homogeneity :  $\forall \lambda > 0$ ,  $\mathcal{R}(\lambda X) = \lambda \mathcal{R}(X)$ .*
- iii) *Subadditivity :  $\mathcal{R}(X + X') \leq \mathcal{R}(X) + \mathcal{R}(X')$ .*
- iv) *Non decreasing : If  $X \leq X'$  (a.s.) then  $\mathcal{R}(X) \leq \mathcal{R}(X')$ .*
- v) *Closeness : Let  $(X_h)_{h \in \mathbb{R}}$  be a collection of random variables. If  $\mathcal{R}(X_h) \leq 0$  and  $\lim_{h \rightarrow 0} \|X_h - X\|_2 = 0$  then  $\mathcal{R}(X) \leq 0$ .*

The superquantile is a coherent measure of risk (see [17], [?], [1]).

## 2. BREGMAN SUPERQUANTILES

In this section the aim is to build a general measure of risk that satisfies some of the regularity axioms stated in Definition 1.1. These quantities will be built by using a dissimilarity measure between real numbers, the Bregman divergence (see [5]).

**2.1. Bregman divergence, mean and superquantile.** In this section we first recall the definition of the Bregman mean of a probability measure  $\mu$  (see [4]) and define the measure of risk that we will study. To begin with, we recall the definition of the Bregman divergence that will be used to build the Bregman mean. Let  $\gamma$  be a strictly convex function,  $\overline{\mathbb{R}}$ -valued on  $\mathbb{R}$ . As usual we set

$$\text{dom}\gamma := \{x \in \mathbb{R} : \gamma(x) < +\infty\}.$$

For sake of simplicity we assume that  $\text{dom}\gamma$  is a non empty open set and that  $\gamma$  is a closed proper differentiable function on the interior of  $\text{dom}\gamma$  (see [19]). From now we always consider function  $\gamma$  satisfying this assumption. The Bregman divergence  $d_\gamma$  associated to  $\gamma$  (see [5]) is a function defined on  $\text{dom}\gamma \times \text{dom}\gamma$  by

$$d_\gamma(x, x') := \gamma(x) - \gamma(x') - \gamma'(x')(x - x') \quad ; (x, x' \in \text{dom}\gamma).$$

The Bregman divergence is not a distance as it is not symmetric. Nevertheless, as it is non negative and vanishes, if and only if, the two arguments are equal, it quantifies the proximity of points in  $\text{dom}\gamma$ . Let us recall some classical examples of such a divergence.

- Euclidean.  $\gamma(x) = x^2$  on  $\mathbb{R}$ , we obviously obtain, for  $x, x' \in \mathbb{R}$ ,

$$d_\gamma(x, x') = (x - x')^2.$$

- Geometric.  $\gamma(x) = x \ln(x) - x + 1$  on  $\mathbb{R}_+^*$  we obtain, for  $x, x' \in \mathbb{R}_+^*$ ,

$$d_\gamma(x, x') = x \ln \frac{x}{x'} + x' - x.$$

- Harmonic.  $\gamma(x) = -\ln(x) + x - 1$  on  $\mathbb{R}_+^*$  we obtain, for  $x, x' \in \mathbb{R}_+^*$ ,

$$d_\gamma(x, x') = -\ln \frac{x}{x'} + \frac{x}{x'} - 1.$$

Let  $\mu$  be a probability measure whose support is included in  $\text{dom}\gamma$  and that does not weight the boundary of  $\text{dom}\gamma$ . Assume further that  $\gamma'$  is integrable with respect to  $\mu$ . Following [4], we first define the Bregman mean as the unique point  $b$  in the support of  $\mu$  satisfying

$$(1) \quad \int d_\gamma(b, x)\mu(dx) = \min_{m \in \text{dom}\gamma} \int d_\gamma(m, x)\mu(dx).$$

In fact, we replace the  $L^2$  minimization in the definition of the mathematical expectation by the minimization of the Bregman divergence. Existence and uniqueness come from the convexity properties of  $d_\gamma$  with respect to its first argument. By differentiating it is easy to see that

$$b = \gamma'^{-1} \left[ \int \gamma'(x)\mu(dx) \right].$$

Hence, coming back to our three previous examples, we obtain the classical mean in the first example (Euclidean case), the geometric mean ( $\exp \int \ln(x)\mu(dx)$ ), in the second one and the harmonic mean ( $[\int x^{-1}\mu(dx)]^{-1}$ ), in the third one. Notice that, as the Bregman divergence is not symmetric, we have to pay attention to the definition of the Bregman mean. Indeed, we have

$$\int d_\gamma(x, \mathbb{E}(X))\mu(dx) = \min_{m \in \text{dom}\gamma} \int d_\gamma(x, m)\mu(dx).$$

We turn now to the definition of our new measure of risk.

**Definition 2.1.** Let  $\alpha \in ]0, 1[$ , the Bregman superquantile  $Q_\alpha^{d_\gamma}$  is defined by

$$Q_\alpha^{d_\gamma} := \gamma'^{-1} \left( \mathbb{E}(\gamma'(X) | X \geq F_X^{-1}(\alpha)) \right) = \gamma'^{-1} \left[ \mathbb{E} \left( \frac{\gamma'(X) \mathbf{1}_{X \geq F_X^{-1}(\alpha)}}{1 - \alpha} \right) \right].$$

In words  $Q_\alpha^{d_\gamma}$  satisfies (1) taking for  $\mu$  the distribution of  $X$  conditionally to

$X \geq F_X^{-1}(1 - \alpha)$ . We now denote  $Q_\alpha^{d_\gamma}$  the Bregman superquantile of the law  $X$  when there is no ambiguity and  $Q_\alpha^{d_\gamma}(X)$  if we need to distinguish Bregman superquantile of different laws.

**2.2. Coherence of Bregman superquantile.** The following proposition gives some conditions under which the Bregman superquantile is a coherent measure of risk.

**Proposition 2.1.** Fix  $\alpha$  in  $]0, 1[$ .

- i) Any Bregman superquantile always satisfies the properties of constant invariance and non decreasing.
- ii) The Bregman superquantile associated to  $\gamma$  is homogeneous if and only if  $\gamma''(x) = \beta x^\delta$  for some real numbers  $\beta$  and  $\delta$ .
- iii) If  $\gamma'$  is concave and sub-additive, then subadditivity and closeness axioms both hold.

The proof of this proposition, like all the others, is deferred to Section 5.

To conclude, under some regularity assumptions on  $\gamma$ , the Bregman superquantile is a coherent measure of risk. Let us take some examples.

### 2.2.1. Examples and counter-examples.

- Example 1 :  $x \mapsto x^2$  satisfies all the hypothesis but it is already known that the classical superquantile is sub-additive.
- Example 2 : The Bregman geometric and harmonic functions satisfies the hypothesis i) and ii). Moreover, their derivatives are respectively  $x \mapsto \gamma'(x) = \ln(x)$  and  $x \mapsto \gamma'(x) = \frac{x-1}{x}$  which are concave but sub-additive only on  $[1, +\infty[$ . Then the harmonic and geometric functions satisfy iii) not for all pairs of random variables but only for pairs  $(X, X')$  such that, denoting  $Z := X + X'$  we have

$$\min \left( q_\alpha^X(\alpha), q_\alpha^{X'}(\alpha), q_\alpha^Z(\alpha) \right) > 1$$

- Counter-example 3 : The sub-additivity is not true in the general case. Indeed, let  $\gamma(x) = \exp(x)$  be our convex function and let us consider the uniform law on  $[0, 1]$ .

$$\mathbb{E}(\gamma'(X)\mathbf{1}_{X \geq F_X^{-1}(\alpha)}) = \int_\alpha^1 \exp(x) dx = e - \exp(\alpha).$$

Then

$$\mathcal{R}(X) = \ln \left( \frac{e - \exp(\alpha)}{1 - \alpha} \right).$$

Moreover,

$$\mathbb{E} \left( \gamma'(\lambda X) \mathbf{1}_{\lambda X \geq F_{\lambda X}^{-1}(\alpha)} \right) = \int_\alpha^1 \exp(\lambda x) dx = \frac{\exp(\lambda) - \exp((\alpha)\lambda)}{\lambda}.$$

So

$$\mathcal{R}(\lambda X) = \ln \left( \frac{\exp(\lambda) - \exp((\alpha)\lambda)}{\lambda(1 - \alpha)} \right).$$

For  $\alpha = 0.95$  and  $\lambda = 2$ , we obtain

$$\mathcal{R}(2X) - 2\mathcal{R}(X) = \mathcal{R}(X + X) - (\mathcal{R}(X) + \mathcal{R}(X)) = 0.000107 > 0,$$

and subadditivity fails.

We can also notice that for  $\lambda = 4$

$$\frac{\mathcal{R}(4X)}{4\mathcal{R}(X)} = 1,000321$$

and the homogeneity is not true. It is coherent with the Proposition 2.1 since the derivative of  $\gamma$  does not fulfilled the assumption.

### 3. ESTIMATION AND ASYMPTOTICS OF THE BREGMAN SUPERQUANTILE

In this section the aim is to make estimation of the Bregman superquantile. We introduce a Monte Carlo estimator and study its asymptotics properties. Under regularity assumptions on the functions  $\gamma$  and  $F_X^{-1}$ , the Bregman superquantile is consistent and asymptotically Gaussian. All along this section, we consider a function  $\gamma$  satisfying our usual properties and a real-valued random variable  $X$  such that  $\gamma'(X)\mathbf{1}_{X \geq 0}$  is integrable.

**3.1. Monte Carlo estimator.** Assume that we have at hand  $(X_1, \dots, X_n)$  an i.i.d sample with same distribution as  $X$ . If we wish to estimate  $Q_\alpha^{d_\gamma}$ , we may use the following empirical estimator :

$$\hat{Q}_\alpha^{d_\gamma} = \gamma'^{-1} \left[ \frac{1}{1-\alpha} \left( \frac{1}{n} \sum_{i=[n\alpha]+1}^n \gamma'(X_{(i)}) \right) \right]$$

where  $X_{(1)} \leq X_{(2)} \leq \dots \leq X_{(n)}$  is the re-ordered sample built with  $(X_1, \dots, X_n)$ .

**3.2. Asymptotics.** We give the following theorem which study the asymptotic behaviour of the Bregman superquantile.

**Theorem 3.1.** *Let  $\frac{1}{2} < \alpha < 1$  and  $X$  be a real-valued random variable such that the cumulative function  $F_X$  is absolutely continuous and the density function  $f_X$  is strictly non-negative on its support. Let  $(X_1, \dots, X_n)$  be an independant sample with the same distribution as  $X$ .*

- i) *We assume that  $\gamma$  is twice differentiable that the derivative of  $(\gamma' \circ F_X^{-1})$  that we denote by  $l_\gamma$  satisfies  $l_\gamma = o((1-t)^{-2})$  and is non-decreasing in when  $t$  goes to  $1^-$ .*

*Then, the estimator  $\hat{Q}_\alpha^{d_\gamma}$  is consistent in probability.*

- ii) *Assume that  $\gamma$  is three times differentiable, and the second derivative of  $(\gamma' \circ F_X^{-1})$  that we denote  $L_\gamma$  satisfies  $L_\gamma = O((1-t)^{-m_L})$  and is non decreasing and when  $t$  goes to  $1^-$ .*

*Then,  $\hat{Q}_\alpha^{d_\gamma}$  the estimator is asymptotically normal*

$$\sqrt{n} \left( \hat{Q}_\alpha^{d_\gamma} - Q_\alpha^{d_\gamma}(X) \right) \implies \mathcal{N} \left( 0, \frac{\sigma_\gamma^2}{\left( \gamma'' \left( Q_\alpha^{d_\gamma}(X) \right) \right)^2 (1-\alpha)^2} \right)$$

where

$$\sigma_\gamma^2 := \int_\alpha^1 \int_\alpha^1 \frac{(\min(x, y) - xy)}{f_Z(F_Z^{-1}(x))f_Z(F_Z^{-1}(y))}.$$

and  $Z := \gamma'(X)$ .

**Remark 3.1.** *Easy calculations show that we have the following equalities*

$$l_\gamma := \frac{\gamma'' \circ F_X^{-1}}{f_X \circ F_X^{-1}},$$

$$L_\gamma := \frac{(\gamma''' \circ F_X^{-1}) \times (f_X \circ F_X^{-1}) - (f - X' \circ F_X^{-1}) \times (\gamma'' \circ F_X^{-1})}{(f_X \circ F_X^{-1})^3}$$

and

$$f_Z = \frac{f_X \circ \gamma'^{-1}}{\gamma'' \circ \gamma'^{-1}}.$$

**Remark 3.2.** *The second part of the theorem shows the asymptotic normality of the Bregman superquantile. We can then use the Slutsky's lemma to find confident interval.*

Indeed, since our estimator  $\hat{Q}_\alpha^{d\gamma} := \gamma'^{-1} \left( (n(1-\alpha))^{-1} \sum_{i=\lfloor n\alpha \rfloor + 1}^n Z_{(i)} \right)$  is consistent

$$\frac{\sqrt{n}}{(\gamma'' \circ \hat{Q}_\alpha^{d\gamma})} \left( \hat{Q}_\alpha^{d\gamma} - Q_\alpha^{d\gamma}(X) \right) \implies \mathcal{N} \left( 0, \frac{\sigma_\gamma^2}{(1-\alpha)} \right)$$

To prove this Theorem 3.1 we use results on the asymptotic behavior of the superquantile. We have the fundamental link between this two quantities

$$(2) \quad Q_\alpha^{d\gamma}(X) = \gamma'^{-1} (Q_\alpha(\gamma'(X))).$$

Indeed, as  $\gamma' \left( F_{(\gamma')^{-1}(Z)}^{-1}(\alpha) \right) = F_Z^{-1}(\alpha)$ , we have

$$\mathbb{E} \left( \gamma'(X) \mathbf{1}_{X > F_X^{-1}(\alpha)} \right) = \mathbb{E} \left( Z \mathbf{1}_{Z > F_Z^{-1}(\alpha)} \right).$$

So that, we will first prove the following proposition on the asymptotic behavior of the superquantile (which is equivalent to deal with the Bregman superquantile with the function  $\gamma$  equals to identity). Then, we apply this proposition to the sample  $(Z_1, \dots, Z_n)$  where  $Z_i := \gamma'(X_i)$ . We conclude by applying the continuous map theorem for the consistency and the delta-method for the asymptotic normality (see for example [27]), both with the regular function  $\gamma'^{-1}$ .

**Proposition 3.1.** *Let  $\frac{1}{2} < \alpha < 1$ , and  $X$  be a real-valued random variable such that the cumulative function  $F_X$  is absolutely continuous and the density function  $f_X$  is strictly non-negative on its support. Let  $(X_1, \dots, X_n)$  be an independant sample with the same distribution as  $X$ .*

- i) *We suppose that the derivative of the quantile function  $F_X^{-1}$  denoting  $l$  satisfies  $l = o((1-t)^{-2})$  and is non-decreasing when  $t$  goes to  $1^-$ .*

*Then, the estimator  $\hat{Q}_\alpha := ((1-\alpha)n)^{-1} \sum_{i=\lfloor n\alpha \rfloor}^n X_{(i)}$  of  $Q_\alpha^X$  is consistent in*

*probability.*

- ii) *Assume that the second derivative of the quantile function that we denote  $L$  satisfy  $L = O((1-t)^{-m_L})$  for an  $1 < m_L < \frac{5}{2}$  and is non decreasing when  $t$  goes to  $1^-$ . Then the estimator  $\hat{Q}_\alpha$  asymptotically normal*

$$\sqrt{n} \left( \frac{1}{n(1-\alpha)} \sum_{i=[n\alpha]+1}^n X_{(i)} - Q_\alpha \right) \Rightarrow \mathcal{N} \left( 0, \frac{\sigma^2}{(1-\alpha)^2} \right)$$

where

$$\sigma^2 := \int_\alpha^1 \int_\alpha^1 \frac{(\min(x, y) - xy)}{f(F^{-1}(x))f(F^{-1}(y))}.$$

**3.3. Examples of asymptotic behavior for the classical superquantile.** Our hypothesis are easy to check in practice. Let us show examples of the asymptotic behaviour of the superquantile : the exponential distribution of parameter 1 and the Pareto distribution.

**3.3.1. Exponential distribution.** In this case, we have on  $\mathbb{R}_*^+$   $f(t) = \exp(-t)$ ,  $F(t) = 1 - \exp(-x)$ . Then  $F^{-1}(t) = -\ln(1-t)$ .

- Consistency :

In the neighborhood of 1,  $l(t) = (1-t)^{-1} = o((1-t)^{-2})$  and  $l$  is non decreasing then the superquantile is consistent.

- Asymptotic normality :

In the neighborhood of 1,  $L(t) = (1-t)^{-2} = O((1-t)^{-m_L})$  for  $2 < m_L < \frac{5}{2}$  and  $L$  is non-decreasing. So that asymptotic normality holds.

**3.3.2. Pareto law.** Here, we consider the Pareto law of parameter  $a > 0$  : on  $\mathbb{R}_*^+$ ,

$$F(t) = 1 - x^{-a}, f(t) = ax^{-a-1}, \text{ and } F^{-1}(t) = (1-t)^{-\frac{1}{a}}.$$

- Consistency :

$l(t) = (a(1-t)^{-1-\frac{1}{a}})$  thus, in the neighborhood of 1  $l(t) = o((1-t)^{-2})$  as soon as  $a > 1$ . Then,  $l$  is non-decreasing near 1. The consistency is true.

- Asymptotic normality :

$L(t) = C(a)(1-x)^{-\frac{1}{a}-2}$  thus, as soon as  $a > 2$ , there exists  $\frac{3}{2} < m_L < \frac{5}{2}$  such that  $L(t) = O\left(\frac{1}{(1-t)^\delta}\right)$  in the neighborhood of 1 and  $L$  is non decreasing around 1. The asymptotic normality is true if and only if  $a > 2$ .

**3.4. Examples of asymptotic behaviour of the Bregman superquantile.** Let us now study the same examples for the Bregman superquantile. For the exponential distribution, the conclusion is the same. Nevertheless, we can find a function  $\gamma$  such that the estimator of the Bregman superquantile is asymptotically normal without conditions on  $a$ . So, the Bregman superquantile is a more interesting measure of risk than the superquantile.

**3.4.1. Exponential law.** Let us show the example of the exponential distribution and the geometric Bregman function. We have  $\gamma'(x) = (x-1)x^{-1}$  and  $F^{-1}(x) = -\ln(1-t)$ . Then

$$F_Z^{-1}(t) = 1 + \frac{1}{\ln(1-t)},$$

- Consistency. In this case, we have,

$$l_\gamma(t) = \frac{1}{(1-t)(\ln(1-t))^2}.$$

So, in the neighborhood of 1,  $l_\gamma$  is non decreasing and  $o((1-t)^{-2})$ .

The estimator  $\hat{Q}_\alpha^{d_\gamma}$  is consistent.

- Asymptotical normality.

$$L_\gamma(t) = \frac{(\ln(1-t))^2 + 2\ln(1-t)}{(1-t)^2(\ln(1-t))^4},$$

Then in the neighborhood of 1  $L_\gamma$  is non decreasing and  $O((1-t)^{-m_L})$ , for  $2 < m_L < \frac{5}{2}$ .

Our estimator is asymptotically Gaussian.

3.4.2. *Pareto law.* Let us now study the case of the Pareto law with the geometric Bregman function. We have  $F^{-1}(t) = (1-t)^{\frac{-1}{a}}$  and  $\gamma'(t) = \ln(t)$ . Then

$$F_Z^{-1}(t) = -\frac{1}{a} \ln(1-t).$$

- Consistency.

$$l_\gamma(t) = \frac{1}{a} \frac{1}{1-t} = o\left(\frac{1}{(1-t)^2}\right)$$

and the non-decreasing is true. The estimator is consistent.

- Asymptotic normality.

$$L_\gamma(t) = \frac{1}{a} \frac{1}{(1-t)^2} = O((1-t)^{-m_L}),$$

for  $2 < m_L < \frac{5}{2}$ .

The estimator is consistent and normally asymptotic for every  $a > 0$ .

## 4. PROOFS

### 4.1. Proof of the Proposition 2.1 : coherence of the Bregman superquantile.

*Proof.* For the first part of the point i), we obviously have  $\mathcal{R}(C) = \gamma'^{-1}(\gamma'(C)) = C$ .

Let us show the non decreasing property. We first show that the superquantile is non-decreasing. Then, (2) and the monotony of  $\gamma'^{-1}$  and  $\gamma$  will allow us to conclude. In [16] it is shown that the Superquantile of the law of  $X$  satisfy :

$$Q_\alpha = \min_{a \in \mathbb{R}} \left\{ a + \frac{1}{1-\alpha} \mathbb{E}([X - a]^+) \right\} = q_\alpha + \frac{1}{1-\alpha} \mathbb{E}([X - q_\alpha]^+)$$

Let  $X$  and  $X'$  be two random variables such that  $X \leq X'$  (a.s.). Using the previous results we have

$$Q_\alpha(X) \leq q_\alpha(X') + \frac{1}{1-\alpha} \mathbb{E}([X - q_\alpha(X')]^+).$$

Then the monotony of the function  $x \mapsto (x - c)^+$  when  $c$  is fixed gives

$$Q_\alpha(X) \leq q_\alpha(X') + \frac{1}{1-\alpha} \mathbb{E} \left( [X' - q_\alpha(X')]^+ \right) = Q_\alpha(X').$$

Let us now deal with ii). For every (measurable) function  $f$ , we denote

$$\mathbb{E}_\alpha[f(X)] = \mathbb{E}[f(X)|X \geq F_X^{-1}(\alpha)].$$

Let  $X$  and  $X'$  be two real-valued random variable. The Bregman superquantile associated to  $\gamma$  is:

$$\gamma'^{-1} (\mathbb{E}_\alpha[\gamma'(X)]) .$$

According to Definition 1.1, it is *homogeneous* if, for every  $\lambda > 0$ ,

$$\gamma'^{-1} (\mathbb{E}_\alpha[\gamma'(\lambda X)]) = \lambda \gamma'^{-1} (\mathbb{E}_\alpha[\gamma'(X)]) .$$

As  $\gamma'$  and  $x \mapsto (\gamma'(x) - \gamma'(1))/\gamma''(1)$  yield the same superquantiles, one may assume without loss of generality that  $\gamma'(1) = 0$  and that  $\gamma''(1) = 1$

First, it is easy to check that the condition given is sufficient. For simplicity, we write  $\phi = \gamma'$ . If  $\phi(x) = (x^\beta - 1)/\beta$ , then  $\phi^{-1}(y) = (1 + \beta y)^{1/\beta}$  and

$$\begin{aligned} \phi^{-1} (\mathbb{E}_\alpha[\phi(\lambda X)]) &= \left( 1 + \beta \mathbb{E}_\alpha \left[ \frac{(\lambda X)^\beta - 1}{\beta} \right] \right)^{\frac{1}{\beta}} \\ &= \lambda \left( \mathbb{E}_\alpha [X^\beta] \right)^{1/\beta} \\ &= \lambda \left( 1 + \beta \mathbb{E}_\alpha \left[ \frac{X^\beta - 1}{\beta} \right] \right)^{\frac{1}{\beta}} \\ &= \lambda \phi^{-1} (\mathbb{E}_\alpha[\phi(X)]) . \end{aligned}$$

If  $\phi(x) = \ln(x)$ , then  $\phi^{-1}(y) = \exp(y)$  and

$$\begin{aligned} \phi^{-1} (\mathbb{E}_\alpha[\phi(\lambda X)]) &= \exp (\mathbb{E}_\alpha (\ln(\lambda X))) \\ &= \exp (\mathbb{E}_\alpha (\ln(\lambda)) + \mathbb{E}_\alpha (\ln(X))) \\ &= \lambda \exp (\mathbb{E}_\alpha (\ln(X))) \\ &= \lambda \phi^{-1} (\mathbb{E}_\alpha[\phi(X)]) . \end{aligned}$$

For the other implication, let  $y > 0$ . Let  $Y$  be a random variable with distribution  $\mathbb{P}$  such that, denoting  $a = y \wedge 1$ ,  $\mathbb{P}(du) = \alpha a^{-1} 1_{[0,a]}(u) du + (1-\alpha)p\delta_y + (1-\alpha)(1-p)\delta_1$ . Its quantile of order  $\alpha$  is  $F_Y^{-1}(\alpha) = a$ . The conditional distribution of  $Y$  given  $Y \geq F_Y^{-1}(\alpha)$  is  $(1-p)\delta_1 + p\delta_y$ , and  $\mathbb{E}_\alpha[\phi(Y)] = (1-p)\phi(1) + p\phi(y)$ . The homogeneity property and the assumption  $\phi(1) = 0$  imply that

$$\phi^{-1} ((1-p)\phi(\lambda) + p\phi(\lambda y)) = \lambda \phi^{-1} (p\phi(y)) .$$

By assumption, the expressions on both sides are smooth in  $p$  and  $y$ . Taking the derivative in  $p$  at  $p = 0$  yields

$$\frac{\phi(\lambda y) - \phi(\lambda)}{\phi'(\lambda)} = \lambda \frac{\phi(y)}{\phi'(1)} ,$$

and hence, as  $\phi'(1) = 1$ ,

$$\phi(\lambda y) - \phi(\lambda) = \lambda \phi'(\lambda) \phi(y).$$

By differentiating with respect to  $y$ , one gets:

$$(3) \quad \phi'(\lambda y) = \phi'(\lambda) \phi'(y).$$

Let  $\psi$  be defined on  $\mathbb{R}$  by  $\psi(z) = \ln(\phi'(\exp(z)))$ . One readily checks that Equation (3) yields:

$$\psi(\ln(y) + \ln(\lambda)) = \psi(\ln(y)) + \psi(\ln(\lambda)).$$

This equation holds for every  $y, \lambda > 0$ . This is well known to imply the linearity of  $\psi$ : there exists a real number  $\beta$  such that for all  $z \in \mathbb{R}$ ,

$$\ln(\phi'(\exp(z))) = \psi(z) = \beta z.$$

Thus,  $\phi'(\exp(z)) = \exp(\beta \exp(z))$ , that is  $\phi'(y) = y^\beta$  for all  $y > 0$ . For  $\alpha = -1$ , one obtains  $\phi(y) = \ln(y)$ . Otherwise, taking into account the constraint  $\phi(1) = 0$ , this yields

$$\phi(y) = \frac{y^{1+\beta} - 1}{1 + \beta}.$$

Let us show iii). Let  $X$  and  $X'$  be two real-valued random variable. Since  $\gamma$  is convex,  $\gamma'$  is non-decreasing. So, to deal with expectation, it is the same thing to show the subadditivity or that

$$\gamma' \left( \frac{\mathcal{R}(X + X')}{2} \right) \leq \gamma' \left( \frac{\mathcal{R}(X) + \mathcal{R}(X')}{2} \right).$$

We set  $S := X + X'$ . Using the concavity of  $\gamma$ , we have

$$\begin{aligned} \gamma' \left( \frac{\mathcal{R}(X) + \mathcal{R}(X')}{2} \right) &\leq \frac{1}{2} \left[ \gamma'(\mathcal{R}(X)) + \gamma'(\mathcal{R}(X')) \right] \\ &= \frac{1}{2(1-\alpha)} \mathbb{E} \left( \gamma'(X) \mathbf{1}_{X \geq q_\alpha^X} + \gamma'(X') \mathbf{1}_{X' \geq q_\alpha^{X'}} \right) \end{aligned}$$

But

$$\gamma' \left( \frac{\mathcal{R}(S)}{2} \right) = \frac{1}{2(1-\alpha)} \mathbb{E} \left( \gamma'(S) \mathbf{1}_{S \geq q_\alpha^S} \right).$$

So we want to show that

$$\mathbb{E} \left( \gamma'(X) \mathbf{1}_{X \geq q_\alpha^X} + \gamma'(X') \mathbf{1}_{X' \geq q_\alpha^{X'}} - \gamma'(S) \mathbf{1}_{S \geq q_\alpha^S} \right) \geq 0.$$

The hypothesis of sub-additivity is now useful because it allows us to use the same argument that in the proof of the sub-additivity of the classical superquantile (see [2]).

$$\begin{aligned}
& \mathbb{E}\left(\gamma'(X)\mathbf{1}_{X \geq q_\alpha^X} + \gamma'(X')\mathbf{1}_{X' \geq q_\alpha^{X'}} - \gamma'(S)\mathbf{1}_{S \geq q_\alpha^S}\right) \\
& \geq \mathbb{E}\left(\gamma'(X)\mathbf{1}_{X \geq q_\alpha^X} + \gamma'(X')\mathbf{1}_{X' \geq q_\alpha^{X'}} - \gamma'(X)\mathbf{1}_{S \geq q_\alpha^S} - \gamma'(X')\mathbf{1}_{S \geq q_\alpha^S}\right) \\
& \geq \gamma'(q_\alpha^X)\mathbb{E}\left(\mathbf{1}_{X \geq q_\alpha^X} - \mathbf{1}_{S \geq q_\alpha^S}\right) + \gamma'(q_\alpha^{X'})\mathbb{E}\left(\mathbf{1}_{X' \geq q_\alpha^{X'}} - \mathbf{1}_{S \geq q_\alpha^S}\right) \\
& = 0.
\end{aligned}$$

Finally, we show the closeness under the same assumption as just before. Let be  $(X_h)_h > 0$  satisfying the hypothesis. By subadditivity we have

$$\mathcal{R}(X) \leq \mathcal{R}(X_h) + \mathcal{R}(X_h - X) \leq 0 + \mathcal{R}(X_h - X).$$

Then denoting  $Y_h = X_h - X$ , it is enough to show that

$$Y_h \xrightarrow{L^2, n \rightarrow +\infty} 0 \implies \mathcal{R}(Y_h) \xrightarrow{n \rightarrow +\infty} 0$$

to conclude. Thanks to the concavity of  $\gamma'$  we can use Jensen inequality

$$\gamma'^{-1} \left[ \frac{\mathbb{E}\left(\gamma'(Y_h)\mathbf{1}_{Y_h \geq F_{Y_h}^{-1}(\alpha)}\right)}{1 - \alpha} \right] \leq \frac{\mathbb{E}\left(Y_h\mathbf{1}_{Y_h \geq F_{Y_h}^{-1}(\alpha)}\right)}{1 - \alpha}.$$

We conclude with Cauchy-Schwartz inequality

$$\frac{\mathbb{E}\left(Y_h\mathbf{1}_{Y_h \geq F_{Y_h}^{-1}(\alpha)}\right)}{1 - \alpha} \leq \mathbb{E}((Y_h)^2)^{\frac{1}{2}} \frac{\mathbb{E}\left(\mathbf{1}_{Y_h \geq F_{Y_h}^{-1}(\alpha)}^2\right)^{\frac{1}{2}}}{1 - \alpha} = \|Y_h\|^2 \sqrt{1 - \alpha} \xrightarrow{h \rightarrow 0} 0.$$

□

## 4.2. Proof of Proposition 3.1 : asymptotic behavior of the plug-in estimator of the superquantile.

**4.3. Mathematical tools.** We first give some technical or classical results we will use in the forthcoming proofs.

**4.3.1. Ordered statistics and Beta function.** Let us recall some results of ordered statistics (see [11]). Let  $(Y_i)_{i=1 \dots n+1}$  be an independent sample having the standard exponential distribution. It's well known that

$$(4) \quad U_{(i)} := \sum_{j=1}^i Y_j \left( \sum_{j=1}^{n+1} Y_j \right)^{-1}$$

as the same distribution as the  $i^{\text{th}}$  ordered statistics of an i.i.d sample of size  $n$  uniformly distributed on  $[0, 1]$ . That is Beta distribution of parameter  $i, n - i + 1$  denoted  $\mathcal{B}(i, n - i + 1)$ . It is also known that this equality in law holds

$$X_{(i)} \stackrel{\mathcal{L}}{=} F^{-1}(U_{(i)}).$$

Recall that the  $\mathcal{B}(a, b)$  distribution has the following density

$$f_{\mathcal{B}(a,b)}(x) = \frac{x^{a-1}(x-1)^{b-1}}{B(a,b)}, \quad a, b > 0$$

where

$$(5) \quad B(a, b) = \int_0^1 t^{a-1}(1-t)^{b-1} dt = \frac{\Gamma(a)\Gamma(b)}{\Gamma(a+b)}.$$

A classical property of the Beta function is

$$(6) \quad \forall (x, y) \in \mathbb{R}^+, \quad B(x+1, y) = \frac{x}{x+y} B(x, y)$$

Generalizing the definition of the factorials, we set for  $n \in \mathbb{N}^*$

$$\left(n - \frac{1}{2}\right)! := \left(n - \frac{1}{2}\right) \left(n - 1 - \frac{1}{2}\right) \dots \left(\frac{1}{2}\right),$$

we have for  $i \in \mathbb{N}^*, n \geq i + 2$  :

$$(7) \quad B(i, n - i - \frac{5}{2} + 1) = \frac{(i-1)! (n - i - 2 - \frac{1}{2})!}{(n - 2 - \frac{1}{2})!}$$

$$(8) \quad \left(n - \frac{1}{2}\right)! = \frac{(2n)!}{(2^n)2n!}$$

Indeed, the first equation is the definition (5). To proof the second equation, let us fix  $k = \frac{1}{2}$ .

$$2^n(n-k)! = (2n-1)(2n-3)\dots 3 \times 1 = \frac{(2n)!}{2n(2n-2)\dots 6 \times 4 \times 2} = \frac{(2n)!}{2^n n!}$$

4.3.2. *Technical lemma.*

**Lemma 4.1.** *Let  $\delta > 1$ . Then  $n^{-1} \sum_{i=\lfloor n\alpha \rfloor}^{n-1} \left(1 - \frac{i}{n+1}\right)^{-\delta} = \mathcal{O}(\sqrt{n})$  if and only if  $\delta \leq \frac{3}{2}$ .*

*Proof.* Let  $\delta > 1$ . We have to characterize the  $\delta$  for which

$$n^{-\frac{3}{2}} \sum_{i=\lfloor n\alpha \rfloor}^{n-1} \left(1 - \frac{i}{n+1}\right)^{-\delta}$$

is bounded when  $n$  goes to infinity. We set  $j := n + 1 - i$ .

The sum becomes

$$n^{-\frac{3}{2}} \sum_{i=2}^{n+1-\lfloor n\alpha \rfloor} \left( \frac{j}{n+1} \right)^\delta \underset{n \rightarrow +\infty}{\sim} n^{\delta-\frac{3}{2}} \sum_{i=2}^{n+1-\lfloor n\alpha \rfloor} \frac{1}{j^\delta} \underset{n \rightarrow +\infty}{\sim} n^{\delta-\frac{3}{2}} \zeta(\delta)$$

where  $\zeta$  denote the Zeta function. The conclusion of the lemma holds.

Finally, our problem is equivalent to characterize the set of  $\delta$  which are superior than 1 and such that  $n^{\delta-\frac{1}{2}}$  is bounded. This set is clearly  $]1, \frac{3}{2}]$ .  $\square$

**4.3.3. A corollary of Lindenberg-Feller theorem.** To prove the asymptotic normality, we use a central limit theorem which is a corollary of the Lindeberg-Feller theorem (see lemma 1 in [8]).

**Proposition 4.1.** *Let  $(Y_1, \dots, Y_n)$  be an independent sample of exponential variables of parameter 1 and  $(\alpha_{j,n})_{j \leq n, n \geq 2}$  be a triangle array of real numbers.*

*If  $Q_n = n^{-1} \sum_{j=1}^n \alpha_{j,n}(Y_j - 1)$  and  $\sigma_n^2 = \frac{1}{n} \sum_{j=1}^n \alpha_{j,n}^2$ , then*

$$\frac{\sqrt{n}Q_n}{\sigma_n} \implies \mathcal{N}(0, 1)$$

*if and only if  $\max_{1 \leq j \leq n} |\alpha_{j,n}| = o(n^{\frac{1}{2}} \sigma_n)$ .*

*If furthermore  $\sigma_n$  converges in probability to  $\sigma$  then by Slutsky's lemma*

$$\sqrt{n}Q_n \implies \mathcal{N}(0, \sigma^2)$$

#### 4.4. Proof of i) of Theorem 3.1 : consistency of the plug-in estimator.

*Proof.* We aim to show consistency of the estimator.

$$\frac{1}{(1-\alpha)n} \sum_{i=\lfloor n\alpha \rfloor}^n X_{(i)}.$$

Let us first notice that

$$Q_\alpha = \frac{\mathbb{E} \left( X \mathbf{1}_{X \geq F_X^{-1}(x)(\alpha)} \right)}{1-\alpha} = \frac{\int_{\mathbb{R}} x \mathbf{1}_{X \geq F_X^{-1}(\alpha)} f_X(x) dx}{1-\alpha} = \frac{\int_\alpha^1 F_X^{-1}(y) dy}{1-\alpha}$$

Thus, we need to show that

$$\frac{1}{n} \sum_{i=\lfloor n\alpha \rfloor}^n X_{(i)} - \int_\alpha^1 F_X^{-1}(y) dy \xrightarrow[n \rightarrow +\infty]{} 0 \text{ a.s.}$$

In the sequel we omit the index  $X$  in  $F_X^{-1}$  because there is no ambiguity. Let us introduce the two following quantities.

$$A_n = \frac{1}{n} \sum_{i=\lfloor n\alpha \rfloor}^n X_{(i)} - \frac{1}{n} \sum_{i=\lfloor n\alpha \rfloor}^n F^{-1} \left( \frac{i}{n+1} \right)$$

and

$$B_n = \frac{1}{n} \sum_{i=\lfloor n\alpha \rfloor}^n F^{-1}\left(\frac{i}{n+1}\right) - \int_{\alpha}^1 F^{-1}(y) dy.$$

$A_n$  converges to 0 in  $L^1$ . Indeed, denoting by  $F_n$  the empirical cumulative distribution function, we have  $X_{(i)} = F_n^{-1}\left(\frac{i}{n}\right)$  and

$$\begin{aligned} \mathbb{E}(|A_n|) &= \mathbb{E}\left(\left|\frac{1}{n} \sum_{i=\lfloor n\alpha \rfloor}^n F_n^{-1}\left(\frac{i}{n+1}\right) - F^{-1}\left(\frac{i}{n+1}\right)\right|\right) \\ &\leq \frac{1}{n} \sum_{i=\lfloor n\alpha \rfloor}^n \mathbb{E}\left(\left|F_n^{-1}\left(\frac{i}{n+1}\right) - F^{-1}\left(\frac{i}{n+1}\right)\right|\right) \end{aligned}$$

We know thanks to recalling 4.3.1 that  $X_{(i)} = F_n^{-1}\left(\frac{i}{n+1}\right)$  is distributed like the  $i^{\text{th}}$  ordered statistic of a uniform sample. Thus, defining  $U_{(i)}$  with law  $\mathcal{B}(i, n+1-i)$ , it holds that,

$$\mathbb{E}(|A_n|) = \mathbb{E}\left(\left|F^{-1}(U_{(i)}) - F^{-1}\left(\frac{i}{n+1}\right)\right|\right).$$

By the mean value theorem, there exists  $w_{(i)}^n \in ]U_{(i)}, \frac{i}{n+1}[$  such that for  $n$  large enough:

$$\begin{aligned} \mathbb{E}\left(\left|F_n^{-1}\left(\frac{i}{n+1}\right) - F^{-1}\left(\frac{i}{n+1}\right)\right|\right) &\leq \mathbb{E}\left(\left|U_{(i)} - \frac{i}{n+1}\right| \frac{1}{f(F^{-1}(w_{(i)}^n))}\right) \\ (9) \quad &\leq \sqrt{\text{Var}(U_{(i)})} \sqrt{E\left[\left(\frac{1}{f(F^{-1}(w_{(i)}^n))}\right)^2\right]} \\ &\leq \sqrt{\frac{i(n-i+1)}{(n+1)^2(n+2)}} \max\left(\sqrt{E\left[\left(\frac{1}{f(F^{-1}(\frac{i}{n+1}))}\right)^2\right]}, \sqrt{E\left[\left(\frac{1}{f(F^{-1}(U_{(i)}))}\right)^2\right]}\right) \end{aligned}$$

Indeed,  $l$  (the derivative of  $F^{-1}$ ) is non decreasing in the neighborhood of 1. Let us now deal with the two terms in the maximum. As  $l = o((1-t)^2)$  in the neighborhood of 1, for  $\epsilon > 0$ , there exists  $N$  such that for  $n \geq N$  and  $i \in [\lfloor n\alpha \rfloor, n]$

$$\begin{aligned} \sqrt{E\left[\left(\frac{1}{f(F^{-1}(U_{(i)}))}\right)^2\right]} &\leq \epsilon \mathbb{E}\left[\frac{1}{(1-U_{(i)})^4}\right] \\ &= \frac{\epsilon}{B(i, n+1-i)} \int_0^1 x^{i-1} (1-x)^{n+1-i-4} dx \end{aligned}$$

But

$$\begin{aligned} \sqrt{\frac{i(n-i+1)}{(n+1)^2(n+2)}} \sqrt{E \left[ \left( \frac{1}{f(F^{-1}(U_{(i)}))} \right)^2 \right]} &= \sqrt{\frac{i(n-i+1)(i-1)!}{B(i, n+1-i)(n-4-(i-2)) \dots (n-4)(n+1)^2(n+2)}} \\ &\underset{n \rightarrow +\infty}{\sim} \frac{\epsilon}{\sqrt{n}} \frac{1}{\sqrt{\frac{i}{n+1}} \left(1 - \frac{i}{n+1}\right)^{\frac{3}{2}}}. \end{aligned}$$

The second term of (9) can be calculated in the same way.

$$\sqrt{E \left[ \left( \frac{1}{f(F^{-1})(\frac{i}{n})} \right)^2 \right]} \leq \epsilon \mathbb{E} \left[ \frac{1}{\left(1 - \frac{i}{n}\right)^4} \right] \underset{n \rightarrow +\infty}{\sim} \frac{\epsilon}{\sqrt{n}} \frac{\sqrt{\frac{i}{n+1}}}{\left(1 - \frac{i}{n+1}\right)^{\frac{3}{2}}}$$

Finally, the two sums :

$$P_{n,1} = \epsilon \frac{1}{n} \sum_{i=[n\alpha]}^n \frac{1}{\sqrt{n} \sqrt{\frac{i}{n+1}} \left(1 - \frac{i}{n+1}\right)^{\frac{3}{2}}}, \quad P_{n,2} = \epsilon \frac{1}{n} \sum_{i=[n\alpha]}^n \frac{\sqrt{\frac{i}{n+1}}}{\sqrt{n} \left(1 - \frac{i}{n+1}\right)^{\frac{3}{2}}}$$

convergence to 0 when  $n$  goes to infinity thanks to Lemma 4.1.

So,  $A_n$  converge in  $L^1$  to 0 when  $n$  goes to infinity, then it converges to 0 in probability.

Let us study the term  $B_n = \frac{1}{n} \sum_{i=[n\alpha]}^n F^{-1}\left(\frac{i}{n+1}\right) - \int_{\alpha}^1 F^{-1}(y) dy$  to show its a.s convergence to 0.

**Remark 4.1.** *To begin with, it is easy to show the convergence when the sum and the integral are truncated at  $1 - \epsilon$ ,  $\forall 0 < \epsilon < 1$  (convergence of a Riemann sum for a continuous functions).*

Let us fix  $\epsilon > 0$ . We split the forthcoming sum in two parts denoting the first part  $S_n^1$  and the second part  $S_n^2$ .

$$\frac{1}{n+1} \sum_{i=[n\alpha]}^{[n(1-\epsilon)]} F^{-1}\left(\frac{i}{n+1}\right) + \frac{1}{n+1} \sum_{[n(1-\epsilon)]+1}^{n-1} F^{-1}\left(\frac{i}{n+1}\right) := S_n^1 + S_n^2$$

Since the quantile function is non-decreasing on  $[\alpha, 1]$ , we have :

$$\begin{aligned}
& \int_{\frac{\lfloor n\alpha \rfloor - 1}{n+1}}^{\frac{\lfloor n(1-\epsilon) \rfloor}{n+1}} F^{-1}(t) dt + \int_{\frac{\lfloor n(1-\epsilon) \rfloor}{n+1}}^{\frac{n-1}{n+1}} F^{-1}(t) dt \\
& \leq \frac{1}{n+1} \sum_{i=\lfloor n\alpha \rfloor}^{\lfloor n(1-\epsilon) \rfloor} F^{-1}\left(\frac{i}{n+1}\right) + \frac{1}{n+1} \sum_{\lfloor n(1-\epsilon) \rfloor + 1}^{n-1} F^{-1}\left(\frac{i}{n+1}\right) \\
& \leq \int_{\frac{\lfloor n\alpha \rfloor}{n+1}}^{\frac{\lfloor n(1-\epsilon) \rfloor}{n+1}} F^{-1}(t) dt + \int_{\frac{\lfloor n(1-\epsilon) \rfloor}{n+1}}^{\frac{n}{n+1}} F^{-1}(t) dt
\end{aligned}$$

Let  $C_n^1$  and  $C_n^2$  be the two terms of the previous line and  $D_n^1$  and  $D_n^2$  those of the last one line.

Then, we have :

$$(C_n^1 - S_n^1) + (C_n^2 - D_n^2) \leq (S_n^2 - D_n^2) \leq (D_n^1 - S_n^1)$$

Let us show that  $C_n^2 - D_n^2$  converge to 0, to conclude (the convergence of  $D_n^1 - S_n^1$  and  $C_n^1 - S_n^1$  to 0 is true thanks to the Remark 4.1).

As in the neighborhood of 1  $l(t) = o((1-t)^{-2})$ , we also have  $F^{-1}(t) = o((1-t)^{-1})$ . Then, for  $\epsilon > 0$ , there exist  $N$  such that for  $n \geq N$  :

$$C_n^2 - D_n^2 = - \int_{\frac{n-1}{n+1}}^{\frac{n}{n+1}} F^{-1}(t) dt \leq \epsilon \int_{\frac{n-1}{n+1}}^{\frac{n}{n+1}} \frac{1}{1-t} dt = \epsilon \ln(2)$$

Finally,  $S_n^2 - D_n^2$  converges to 0 a.s. So that, the same holds for  $B_n$ .

We have shown that  $A_n + B_n$  converge to 0 in probability. So under our hypothesis, the superquantile is consistent in probability.

**Remark 4.2.** *using the same arguments, we can show that under stronger hypothesis on the quantile function  $F^{-1}(t) = o\left(\frac{1}{(1-t)^{\frac{1}{2}}}\right)$  (that is the case in ii) of Proposition 3.1), we have :*

$$- \int_{\frac{n-1}{n+1}}^{\frac{n}{n+1}} F^{-1}(t) dt \leq \epsilon \int_{\frac{n-1}{n+1}}^{\frac{n}{n+1}} \frac{1}{(1-t)^{\frac{1}{2}}} dt = \epsilon - 2(1-\sqrt{2}) \frac{1}{\sqrt{n}}$$

Then

$$\sqrt{n} \left( \frac{1}{n} \sum_{i=\lfloor n\alpha \rfloor}^n F^{-1}\left(\frac{i}{n+1}\right) - \int_{\alpha}^1 F^{-1}(y) dy \right) \xrightarrow{n \rightarrow +\infty} 0$$

We will use this result in the next part.

□

4.4.1. *Proof of ii) of Proposition 3.1 : asymptotic normality of the plug-in estimator.* Let us prove the asymptotic normality of the estimator of the superquantile. To begin with, we make some technical remarks.

**Remark 4.3.** *The assumption on  $L$  implies that there exists  $m_l < \frac{3}{2}$  and  $m_{F^{-1}} < \frac{1}{2}$  such that  $l(t) = O((1-t)^{-m_l})$ , and  $F^{-1}(t) = O((1-t)^{-m_{F^{-1}}})$ . It also implies that in the neighborhood of 1,  $L(t) = o((1-t)^{-\frac{5}{2}})$ .*

*Proof.* The proof stands in three steps. First we reformulate and simplify the problem and apply the Taylor Lagrange formula. Then, we show that the second order term converges to 0 in probability. In the third step, we identify the limit of the first order term.

### Step 1 : Taylor-Lagrange formula

Let us first omit  $(1-\alpha)^{-1}$ . We have to study the convergence in distribution of

$$\sqrt{n} \left( \frac{1}{n} \sum_{i=\lfloor n\alpha \rfloor}^n X_{(i)} - \int_{\alpha}^1 F^{-1}(y) dy \right).$$

We have already noticed (Remarks 4.2 and 4.3) that

$$\sqrt{n} \left[ \frac{1}{n} \sum_{i=\lfloor n\alpha \rfloor}^n F^{-1}\left(\frac{i}{n+1}\right) - \int_{\alpha}^1 F^{-1}(y) dy \right] \xrightarrow[n \rightarrow +\infty]{} 0.$$

Thus, Slutsky's lemma, allows us to study only the convergence in law of

$$\sqrt{n} \left[ \frac{1}{n} \sum_{i=\lfloor n\alpha \rfloor}^n X_{(i)} - \frac{1}{n} \sum_{i=\lfloor n\alpha \rfloor}^n F^{-1}\left(\frac{i}{n+1}\right) \right].$$

The quantile function  $F^{-1}$  is  $\mathcal{C}^2$  so that we may apply the order 1 Taylor-Lagrange formula. Using the same argument in the proof of i), we introduce  $U_{(i)}$  a random variable distributed as a  $\mathcal{B}(i, n+1-i)$ . Considering an equality in law we then have

$$\begin{aligned} \sqrt{n} \left( \frac{1}{n} \sum_{i=\lfloor n\alpha \rfloor+1}^n \left[ X_{(i)} - F^{-1}\left(\frac{i}{n+1}\right) \right] \right) &\stackrel{\mathcal{L}}{=} \sqrt{n} \left[ \frac{1}{n} \sum_{i=\lfloor n\alpha \rfloor+1}^n \left( U_{(i)} - \frac{i}{n+1} \right) \frac{1}{f\left(F^{-1}\left(\frac{i}{n+1}\right)\right)} \right] \\ &\quad + \frac{1}{\sqrt{n}} \sum_{i=\lfloor n\alpha \rfloor+1}^n \left[ \int_{\frac{i}{n+1}}^{U_{(i)}} \frac{f'(F^{-1}(t))}{(f(F^{-1}(t)))^3} (U_{(i)} - t) dt \right]. \end{aligned}$$

Let us call  $\sqrt{n}Q_n$  the first-order term and  $R_n$  the second-order one.

### Step 2 : The second-order term converges to 0 in probability

Let us show that  $R_n$  converge to 0 in probability. Thanks to Markov's inequality, we have

$$\mathbb{P}(R_n > \epsilon) \leq \frac{1}{\epsilon\sqrt{n}} \sum_{i=\lfloor n\alpha \rfloor + 1}^n \mathbb{E} \left( \left| \int_{\frac{i}{n+1}}^{U(i)} L(t)(U(i) - t) dt \right| \right)$$

Since the function  $L$  is non decreasing in the neighborhood of 1, for  $n$  large enough, we have,

$$\begin{aligned} \mathbb{P}(R_n > \epsilon) &\leq \frac{1}{\epsilon\sqrt{n}} \sum_{i=\lfloor n\alpha \rfloor + 1}^n \mathbb{E} \left( \max \left( \left| L(U(i)) \right|, \left| L\left(\frac{i}{n+1}\right) \right| \right) \int_{\frac{i}{n+1}}^{U(i)} (U(i) - t) dt \right) \\ &= \frac{1}{\epsilon\sqrt{n}} \sum_{i=\lfloor n\alpha \rfloor + 1}^n \mathbb{E} \left( \max \left( \left| L(U(i)) \right|, \left| L\left(\frac{i}{n+1}\right) \right| \right) \frac{\left( U(i) - \frac{i}{n+1} \right)^2}{2} \right) \end{aligned}$$

As before, we study the two terms in the maximum separately. First, using the variance of the Beta distribution,

$$\frac{1}{\epsilon\sqrt{n}} \sum_{i=\lfloor n\alpha \rfloor + 1}^n \left| L\left(\frac{i}{n+1}\right) \right| \mathbb{E} \left( \frac{\left( U(i) - \frac{i}{n+1} \right)^2}{2} \right) = \frac{1}{\epsilon\sqrt{n}} \sum_{i=\lfloor n\alpha \rfloor + 1}^n \left| L\left(\frac{i}{n+1}\right) \right| \frac{i(n+1-i)}{2(n+1)^2(n+2)}$$

But according to Remark 4.3, our assumption on  $L$ , gives that for  $\epsilon'$ , there exists an integer  $N$  such that for  $n \geq N$ , for  $i \in [\lfloor n\alpha \rfloor, n]$ ,

$$\begin{aligned} \mathbb{P}(R_n > \epsilon) &\leq \frac{1}{\epsilon\sqrt{n}} \sum_{i=\lfloor n\alpha \rfloor + 1}^n \left| L\left(\frac{i}{n+1}\right) \right| \frac{i(n+1-i)}{2(n+1)^2(n+2)} \\ &\leq \frac{\epsilon'}{2\epsilon\sqrt{n}(n+2)} \sum_{i=\lfloor n\alpha \rfloor + 1}^n \frac{\frac{i}{n+1} \left( 1 - \frac{i}{n+1} \right)}{\left( 1 - \frac{i}{n+1} \right)^{\frac{5}{2}}} \\ &= \frac{\epsilon'}{2\epsilon\sqrt{n}(n+2)} \sum_{i=\lfloor n\alpha \rfloor + 1}^n \frac{\frac{i}{n+1}}{\left( 1 - \frac{i}{n+1} \right)^{\frac{3}{2}}} \end{aligned}$$

Further the convergence to 0 when  $n$  goes to infinity of the first term in the maximum holds thanks to the Lemma 4.1.

We have now to deal with the second term in the maximum. Using our assumptions, we have, for  $\epsilon'$  and  $n$  large enough,

$$\frac{1}{\epsilon\sqrt{n}} \sum_{i=\lfloor n\alpha \rfloor + 1}^n \mathbb{E} \left( \left| L(U_{(i)}) \right| \frac{\left( U_{(i)} - \frac{i}{n+1} \right)^2}{2} \right) \leq \frac{\epsilon'}{2\epsilon\sqrt{n}} \sum_{i=\lfloor n\alpha \rfloor + 1}^n \mathbb{E} \left( \frac{\left( U_{(i)} - \frac{i}{n+1} \right)^2}{\left( U_{(i)} - 1 \right)^{\frac{5}{2}}} \right).$$

As  $U_{(i)}$  has a Beta distribution, we can write that,

$$\begin{aligned} \mathbb{E} \left( \frac{\left( U_{(i)} - \frac{i}{n+1} \right)^2}{\left( U_{(i)} - 1 \right)^{\frac{5}{2}}} \right) &= \frac{1}{B(i, n+1-i)} \int_0^1 x^{i-1} (1-x)^{n-i-\frac{5}{2}} \left( x - \frac{i}{n+1} \right)^2 dx \\ &= \frac{1}{B(i, n+1-i)} \left( \int_0^1 x^{i+1} (1-x)^{n-i-\frac{5}{2}} dx - 2 \frac{i}{n+1} \int_0^1 x^i (1-x)^{n-i-\frac{5}{2}} dx \right. \\ &\quad \left. + \left( \frac{i}{n+1} \right)^2 \int_0^1 x^{i-1} (1-x)^{n-i-\frac{5}{2}} dx \right) \end{aligned}$$

Let us call this last quantity  $I_n^i$ . To have  $n - i - \frac{5}{2} + 1$  non negative, we cut the sum and deal only with the terms for  $i$  from  $\lfloor n\alpha \rfloor$  to  $n - 2$ . Then, we obtain :

$$\begin{aligned} I_n^i &= \frac{1}{B(i, n+1-i)} \left( \int_0^1 x^{i+1} (1-x)^{n-i-\frac{5}{2}} dx - 2 \frac{i}{n+1} \int_0^1 x^i (1-x)^{n-i-\frac{5}{2}} dx \right. \\ &\quad \left. + \left( \frac{i}{n+1} \right)^2 \int_0^1 x^{i-1} (1-x)^{n-i-\frac{5}{2}} dx \right) \\ &= \frac{1}{B(i, n+1-i)} \left( B(i+2, n-i-\frac{5}{2}+1) - 2 \frac{i}{n+1} B(i+1, n-i-\frac{5}{2}+1) \right. \\ &\quad \left. + \left( \frac{i}{n+1} \right)^2 B(i, n-i-\frac{5}{2}+1) \right) \end{aligned}$$

So that using (6) we obtain

$$\begin{aligned} I_n^i &= \frac{1}{B(i, n+1-i)} \left( B(i+2, n-i-\frac{5}{2}+1) - 2 \frac{i}{n+1} B(i+1, n-i-\frac{5}{2}+1) \right. \\ &\quad \left. + \left( \frac{i}{n+1} \right)^2 B(i, n-i-\frac{5}{2}+1) \right) \\ &= \frac{B(i, n-i-\frac{5}{2}+1)}{B(i, n+1-i)} \left( \frac{i(i+1)}{(n-\frac{5}{2}+2)(n-\frac{5}{2}+1)} \right. \\ &\quad \left. - 2 \frac{i^2}{(n-\frac{5}{2}+1)(n+1)} \frac{i(i+1)}{(n-\frac{5}{2}+2)(n-\frac{5}{2}+1)} + \left( \frac{i}{n+1} \right)^2 \frac{i(i+1)}{(n-\frac{5}{2}+2)(n-\frac{5}{2}+1)} \right) \end{aligned}$$

Let  $E_n^i$  be such that  $I_n^i = \frac{B(i, n-i-\frac{5}{2}+1)}{B(i, n+1-i)} E_n^i$ . expanding  $E_n^i$  gives when  $n$  goes to infinity :

$$E_n^i \sim \frac{1}{n} \frac{i}{n+1} \left(1 - \frac{i}{n+1}\right).$$

Let us study the term  $\frac{B(i, n-i-\frac{5}{2}+1)}{B(i, n+1-i)}$ . Using (5) and (8), we obtain

$$\begin{aligned} \frac{B(i, n-i-\frac{5}{2}+1)}{B(i, n+1-i)} &= \frac{n!}{(n-2-\frac{1}{2})!} \frac{(n-i-2-\frac{1}{2})!}{(n-i)!} \\ &= \frac{n(n-1)}{(n-i-1)(n-i)} \frac{(2(n-i-2))!((n-2)!)^2 2^{2i}}{((n-i-2)!)^2 (2(n-2))!} \end{aligned}$$

Since each  $i$  can be written as  $i = \lfloor n\beta \rfloor$  with  $\beta < 1$ ,  $n-i$  goes to infinity when  $n$  goes to infinity and we can apply the Stirling formula:

$$\begin{aligned} \frac{(2(n-i-2))!}{((n-i-2)!)^2} &\underset{n \rightarrow +\infty}{\sim} \frac{\sqrt{2(n-i-2)2\pi} \left(\frac{2(n-i-2)}{e}\right)^{2(n-i-2)}}{2\pi(n-i-2) \left(\frac{n-i-2}{e}\right)^{n-i-2}} \\ &\underset{n \rightarrow +\infty}{\sim} \frac{2^{2(n-i-2)}}{\sqrt{\pi(n-i-2)}} \end{aligned}$$

Likewise,

$$\frac{(2(n-2))!}{((n-2)!)^2} \underset{n \rightarrow +\infty}{\sim} \frac{2^{2(n-2)}}{\sqrt{\pi(n-2)}}.$$

Then, when  $n$  goes to infinity

$$\frac{B(i, n-i-\frac{5}{2}+1)}{B(i, n+1-i)} \underset{n \rightarrow +\infty}{\sim} \frac{1}{\left(1 - \frac{i}{n+1}\right)^{\frac{5}{2}}}.$$

Hence we obtain

$$I_n^i \underset{n \rightarrow +\infty}{\sim} \frac{1}{n} \frac{\frac{i}{n+1}}{\left(1 - \frac{i}{n+1}\right)^{\frac{3}{2}}}.$$

Finally, for  $n$  large enough, we have

$$I_n^i \leq 2 \frac{1}{n} \frac{\frac{i}{n+1}}{\left(1 - \frac{i}{n+1}\right)^{\frac{3}{2}}}$$

and thanks to our assumptions on  $L$ , for  $\epsilon'$  and  $n$  large enough, we have

$$\begin{aligned} \frac{1}{\epsilon\sqrt{n}} \sum_{i=[n\alpha]+1}^{n-2} \mathbb{E} \left( \left| L(U_{(i)}) \right| \frac{\left( U_{(i)} - \frac{i}{n+1} \right)^2}{2} \right) &\leq \frac{\epsilon'}{2\epsilon\sqrt{n}} \sum_{i=[n\alpha]+1}^{n-2} \mathbb{E} \left( \frac{\left( U_{(i)} - \frac{i}{n+1} \right)^2}{\left( U_{(i)} - 1 \right)^{\frac{5}{2}}} \right) \\ &\leq \frac{\epsilon'}{\epsilon\sqrt{n}} \frac{1}{n} \sum_{i=[n\alpha]+1}^{n-2} \frac{\frac{i}{n+1}}{\left( 1 - \frac{i}{n+1} \right)^{\frac{3}{2}}}. \end{aligned}$$

To conclude the second term in the maximum converges to 0 thanks the Lemma 4.1.

**Remark 4.4.** *The terms for  $i = 1$  and  $i = n - 1$  are of leaser order of the sum, so they converge to 0 too.*

Finally, the remaining term  $R_n$  converges to 0 in probability so we can now focus on the first order term.

### Step 3 : Identification of the limit

Our goal is to find the limit of  $\sqrt{n}Q_n$ . Let us reorganize the expression of  $Q_n$  to have a more classical form (sum of independent random variables) and to allows the use of the Proposition 4.1.

Denoting by

$$\bar{Y} = \frac{\sum_{j=1}^{n+1} Y_j}{n+1}$$

we have thanks to (4)

$$\begin{aligned} Q_n &= \frac{1}{n} \sum_{i=[n\alpha]+1}^n \left( \frac{\sum_{j=1}^i Y_j}{\frac{n+1}{\sum_{j=1}^i Y_j} - \frac{i}{n+1}} \right) l\left(\frac{i}{n+1}\right) \\ &= \frac{1}{n} \sum_{j=1}^n \left[ \left( \frac{Y_j}{\frac{n+1}{\sum_{j=1}^i Y_j} - \frac{1}{n+1}} \right) \sum_{i=\sup([n\alpha]+1, j)} l\left(\frac{i}{n+1}\right) \right] \\ &= \frac{n+1}{n+1} \frac{1}{n(n+1)} \left[ \sum_{j=1}^{[n\alpha]+1} \left( (Y_j - \bar{Y}) \sum_{i=[n\alpha]+1}^n l\left(\frac{i}{n+1}\right) \right) + \sum_{j=[n\alpha]+2}^n \left( (Y_j - \bar{Y}) \sum_{i=j}^n l\left(\frac{i}{n+1}\right) \right) \right]. \end{aligned}$$

where we have permuted the two sums. The law of large numbers gives that

$n+1 \left( \sum_{j=1}^{n+1} Y_j \right)^{-1}$  converges a.s to 1 when  $n$  goes to infinity. Then, thanks to Slutsky's lemma, we only need to study

$$\frac{1}{n(n+1)} \left[ \sum_{j=1}^{\lfloor n\alpha \rfloor + 1} \left( (Y_j - \bar{Y}) \sum_{i=\lfloor n\alpha \rfloor + 1}^n l\left(\frac{i}{n+1}\right) \right) + \sum_{j=\lfloor n\alpha \rfloor + 2}^n \left( (Y_j - \bar{Y}) \sum_{i=j}^n l\left(\frac{i}{n+1}\right) \right) \right].$$

We set  $\forall j \leq n$ ,  $G_j^n := \sum_{i=j}^n l\left(\frac{i}{n+1}\right)$ ,  $G_{n+1}^n := 0$ ,  $H^n := \sum_{j=\lfloor n\alpha \rfloor + 2}^n G_j^n$ . Then

$$\begin{aligned} Q_n &= \frac{1}{n(n+1)} \left[ \sum_{j=1}^{\lfloor n\alpha \rfloor + 1} \left( \frac{(n - \lfloor n\alpha \rfloor) G_{\lfloor n\alpha \rfloor + 1}^n - H^n}{n+1} \right) Y_j + \sum_{\lfloor n\alpha \rfloor + 2}^{n+1} \left( G_j^n - \frac{H^n}{n+1} + G_{\lfloor n\alpha \rfloor + 1}^n \frac{-1 - \lfloor n\alpha \rfloor}{n+1} \right) Y_j \right] \\ &= \frac{1}{n(n+1)} \left[ \sum_{j=1}^{\lfloor n\alpha \rfloor + 1} \left( \frac{(n - \lfloor n\alpha \rfloor) G_{\lfloor n\alpha \rfloor + 1}^n - H^n}{n+1} \right) (Y_j - 1) \right. \\ &\quad \left. + \sum_{\lfloor n\alpha \rfloor + 2}^{n+1} \left( G_j^n - \frac{H^n}{n+1} + G_{\lfloor n\alpha \rfloor + 1}^n \frac{-1 - \lfloor n\alpha \rfloor}{n+1} \right) (Y_j - 1) \right] \end{aligned}$$

because

$$\begin{aligned} &\frac{1}{n(n+1)} \left[ \sum_{j=1}^{\lfloor n\alpha \rfloor + 1} \left( \frac{(n - \lfloor n\alpha \rfloor) G_{\lfloor n\alpha \rfloor + 1}^n - H^n}{n+1} \right) + \sum_{\lfloor n\alpha \rfloor + 2}^{n+1} \left( G_j^n - \frac{H^n}{n+1} + G_{\lfloor n\alpha \rfloor + 1}^n \frac{-1 - \lfloor n\alpha \rfloor}{n+1} \right) \right] \\ &= \frac{1}{n(n+1)} \left[ \left( \frac{G_{\lfloor n\alpha \rfloor}^n (n - \lfloor n\alpha \rfloor) - H^n}{n+1} \right) (\lfloor n\alpha \rfloor + 1) \right. \\ &\quad \left. + \left( \frac{G_{\lfloor n\alpha \rfloor}^n (-1 - \lfloor n\alpha \rfloor)}{n+1} (n - \lfloor n\alpha \rfloor) - \frac{H^n}{n+1} (n - \lfloor n\alpha \rfloor) + H^n \right) \right] \\ &= 0. \end{aligned}$$

Finally, we obtain

$$\begin{aligned}
Q_n &= \frac{1}{n(n+1)} \left[ \sum_{j=1}^{\lfloor n\alpha \rfloor + 1} \left( \frac{(n - \lfloor n\alpha \rfloor) G_{\lfloor n\alpha \rfloor + 1}^n - H^n}{n+1} \right) (Y_j - 1) \right. \\
&\quad \left. + \sum_{j=\lfloor n\alpha \rfloor + 2}^{n+1} \left( G_j^n - \frac{H^n}{n+1} + G_{\lfloor n\alpha \rfloor + 1}^n \frac{-1 - \lfloor n\alpha \rfloor}{n+1} \right) (Y_j - 1) \right] \\
&= \frac{1}{n+1} \sum_{j=1}^{n+1} \alpha_{j,n} (Y_j - 1)
\end{aligned}$$

where :

$$\alpha_{j,n} = \left( \frac{(n - \lfloor n\alpha \rfloor) G_{\lfloor n\alpha \rfloor + 1}^n - H^n}{n(n+1)} \right), \quad \forall j \leq \lfloor n\alpha \rfloor + 1$$

and

$$\alpha_{j,n} = \left( \frac{G_j^n (n+1) - H^n - G_{\lfloor n\alpha \rfloor + 1}^n (1 + \lfloor n\alpha \rfloor)}{n(n+1)} \right), \quad \forall j \geq \lfloor n\alpha \rfloor + 2$$

Let us check the assumptions of the Proposition 4.1. To begin with, let us show that  $\sigma_n^2$  converges. We have

$$\begin{aligned}
\sigma_n^2 &= \frac{1}{n+1} \sum_{j=1}^{n+1} \alpha_{j,n}^2 \\
&= \frac{\lfloor n\alpha \rfloor + 1}{n+1} \left[ \frac{(n - \lfloor n\alpha \rfloor) G_{\lfloor n\alpha \rfloor + 1}^n - H^n}{n(n+1)} \right]^2 \\
&\quad + \frac{1}{n+1} \sum_{j=\lfloor n\alpha \rfloor + 2}^{n+1} \left( \frac{G_j^n (n+1)}{n(n+1)} \right)^2 + 2 \frac{1}{n+1} \sum_{j=\lfloor n\alpha \rfloor + 2}^{n+1} \frac{G_j^n (n+1) (-H^n - G_{\lfloor n\alpha \rfloor + 1}^n (1 + \lfloor n\alpha \rfloor))}{n^2 (n+1)^2} \\
&\quad + \frac{1}{n+1} \sum_{j=\lfloor n\alpha \rfloor + 2}^{n+1} \left( \frac{-H^n - G_{\lfloor n\alpha \rfloor + 1}^n (1 + \lfloor n\alpha \rfloor)}{n(n+1)} \right)^2
\end{aligned}$$

Let us work with the two terms which depend on  $G_j^n$ . The first term can be expanded as

$$\begin{aligned}
\frac{1}{n+1} \sum_{j=\lfloor n\alpha \rfloor + 2}^{n+1} \left( \frac{G_j^n(n+1)}{n(n+1)} \right)^2 &= \frac{1}{n^2(n+1)} \sum_{j=\lfloor n\alpha \rfloor + 2}^{n+1} \left( G_j^n \right)^2 \\
&= \frac{1}{n^2(n+1)} \sum_{j=\lfloor n\alpha \rfloor + 2}^{n+1} \left( \sum_{i=j}^n l \left( \frac{i}{n+1} \right) \right)^2 \\
&= \frac{1}{n(n+1)} \sum_{i_1=\lfloor n\alpha \rfloor + 2}^{n+1} \sum_{i_2=\lfloor n\alpha \rfloor + 2}^{n+1} \left( \frac{-1 + \min(i_1, i_2) - \lfloor n\alpha \rfloor}{n+1} \right) l \left( \frac{i_1}{n+1} \right) l \left( \frac{i_2}{n+1} \right).
\end{aligned}$$

The second term may be rewritten as

$$2 \frac{1}{n+1} \sum_{j=\lfloor n\alpha \rfloor + 2}^{n+1} \frac{G_j^n(n+1)(-H^n - G_{\lfloor n\alpha \rfloor + 1}^n(1 + \lfloor n\alpha \rfloor))}{n^2(n+1)^2} = -2 \frac{(H^n)^2}{n^2(n+1)^2} - 2 \frac{(1 + \lfloor n\alpha \rfloor)}{(n+1)^2 n^2} H^n G_{\lfloor n\alpha \rfloor + 1}^n.$$

Finally,

$$\begin{aligned}
\sigma_n^2 &= \frac{\lfloor n\alpha \rfloor + 1}{n+1} \left[ \frac{(n - \lfloor n\alpha \rfloor) G_{\lfloor n\alpha \rfloor + 1}^n - H^n}{n(n+1)} \right]^2 \\
&+ \frac{1}{n(n+1)} \sum_{i_1=\lfloor n\alpha \rfloor + 2}^{n+1} \sum_{i_2=\lfloor n\alpha \rfloor + 2}^{n+1} \left( \frac{-1}{n+1} + \frac{\min(i_1, i_2) - \lfloor n\alpha \rfloor}{n+1} \right) l \left( \frac{i_1}{n+1} \right) l \left( \frac{i_2}{n+1} \right) \\
&- 2 \frac{(H^n)^2}{n^2(n+1)^2} - 2 \frac{(1 + \lfloor n\alpha \rfloor)}{(n+1)^3 n^2} H^n G_{\lfloor n\alpha \rfloor + 1}^n + \frac{n - \lfloor n\alpha \rfloor - 1}{n+1} \left( \frac{H^n + G_{\lfloor n\alpha \rfloor + 1}^n(1 + \lfloor n\alpha \rfloor)}{n(n+1)} \right)^2
\end{aligned}$$

Let us first notice that, if we denote

$$K^n = \sum_{i_1=\lfloor n\alpha \rfloor + 2}^{n+1} \sum_{i_2=\lfloor n\alpha \rfloor + 2}^{n+1} \frac{\min(i_1, i_2)}{n+1} l \left( \frac{i_1}{n+1} \right) l \left( \frac{i_2}{n+1} \right)$$

and

$$T^n = \sum_{i=\lfloor n\alpha \rfloor}^n \frac{i}{n} l \left( \frac{i}{n+1} \right)$$

then

$$H^n = nT^n - (\lfloor n\alpha \rfloor + 1)G_{\lfloor n\alpha \rfloor + 1}^n$$

So that

$$\begin{aligned} \sigma_n^2 &\underset{n \rightarrow +\infty}{\sim} \alpha \frac{(G^n - T^n)^2}{n^2} + \frac{K^n - \alpha(G^n)^2}{n^2} - \frac{-2(T^n - \alpha G^n)^2}{n^2} - 2 \frac{\alpha(G^n T^n - \alpha(G^n)^2)}{n^2} + \frac{(1 - \alpha)(T^n)^2}{n^2} \\ &\underset{n \rightarrow +\infty}{\sim} \frac{K^n - (T^n)^2}{n^2}. \end{aligned}$$

Let us show that this last quantity converges to  $\sigma^2 = \int_{\alpha}^1 \int_{\alpha}^1 \frac{\min(x,y) - xy}{f(F^{-1}(x))f(F^{-1}(y))} < \infty$ . Indeed it is a generalized Riemann sum. First, we show that the function

$$g : (x, y) \mapsto \frac{\min(x, y) - xy}{f(F^{-1}(x))f(F^{-1}(y))}$$

is integrable on  $] \alpha, 1[ \times ] \alpha, 1[$ . Indeed, around 1,

$$g(x, y) = O\left(\frac{\min(x, y) - xy}{(1-x)^{m_l}(1-y)^{m_l}}\right)$$

which is integrable on this domain because for  $\beta$  close to 1

$$\int_{\alpha}^{\beta} \int_{\alpha}^{\beta} \frac{\min(x, y) - xy}{(1-x)^{m_l}(1-y)^{m_l}} dx dy \sim C(\alpha)\beta(1-\beta)^{\frac{-2}{m_l}+3}.$$

and  $m_l < \frac{3}{2}$ .

**Remark 4.5.** *It is here that we need that  $L(t) = O((1-t)^{-m_L})$  with  $m_L < \frac{5}{2}$  instead of  $L(t) = o((1-t)^{-\frac{5}{2}})$ . Indeed, when  $m_L = \frac{5}{2}$  and so  $m_l = \frac{3}{2}$  the integral is not finite.*

As we already have seen, the results on Riemann's sum in dimension 2, give by the continuity of the function  $(x, y) \mapsto \frac{\min(x,y) - xy}{f(F^{-1}(x))f(F^{-1}(y))}$  that for all  $\beta < 1$  :

$$\sigma_{n,\beta}^2 := \frac{1}{n^2} \sum_{i_1=\lfloor n\alpha \rfloor}^{\lfloor n\beta \rfloor} \sum_{i_2=\lfloor n\alpha \rfloor}^{\lfloor n\beta \rfloor} \frac{\frac{\min(i_1, i_2)}{n} - \frac{i_1 i_2}{n^2}}{f\left(F^{-1}\left(\frac{i_1}{n+1}\right)\right)f\left(F^{-1}\left(\frac{i_2}{n+1}\right)\right)} \longrightarrow \int_{\alpha}^{\beta} \int_{\alpha}^{\beta} \frac{\min(x, y) - xy}{f(F^{-1}(x))f(F^{-1}(y))} dx dy$$

We have to study the remaining part of the sum to conclude. Let us fix  $\beta$  close to 1 and deal with

$$r_{n,\beta}^2 := \frac{1}{n^2} \sum_{i_1=\lfloor n\beta \rfloor}^{n+1} \sum_{i_2=\lfloor n\beta \rfloor}^{n+1} \frac{\frac{\min(i_1, i_2)}{n} - \frac{i_1 i_2}{n^2}}{f\left(F^{-1}\left(\frac{i_1}{n+1}\right)\right)f\left(F^{-1}\left(\frac{i_2}{n+1}\right)\right)}.$$

In this case, the monotony will allowed us to conclude with the Lebesgue theorem. First of all, let us notice that

$$r_{n,\beta}^2 = \int_{\beta}^1 \int_{\beta}^1 g\left(\frac{\lfloor nx \rfloor}{n}, \frac{\lfloor ny \rfloor}{n}\right) dx dy$$

Then, we aim to permute integration and limit.

1) Let  $(x, y)$  be fixed in  $[\beta, 1[ \times [\beta, 1[$  and  $n$ . Then

$$g\left(\frac{\lfloor nx \rfloor}{n}, \frac{\lfloor ny \rfloor}{n}\right) \longrightarrow g(x, y)$$

by continuity. And  $g$  is integrable on  $[\beta, 1[ \times [\beta, 1[$  as we saw before.

2) Let  $(x, y)$  be fixed in  $[\beta, 1[ \times [\beta, 1[$  and  $n$ . Let us denote  $x_n = \frac{\lfloor nx \rfloor}{n}$  and  $y_n = \frac{\lfloor ny \rfloor}{n}$ .  
By hypothesis

$$g\left(\frac{\lfloor nx \rfloor}{n}, \frac{\lfloor ny \rfloor}{n}\right) \leq C \frac{\min(x_n, y_n) - x_n y_n}{(1 - x_n)^{m_i} (1 - y_n)^{m_i}}.$$

By separating the two cases and using monotony we obtain that

$$g\left(\frac{\lfloor nx \rfloor}{n}, \frac{\lfloor ny \rfloor}{n}\right) \leq Ch(x, y)$$

where

$$h : (x, y) \mapsto \frac{\min(x, y)}{(1 - \min(x, y))^{m_i} (1 - \max(x, y))^{m_i - 1}}$$

which is integrable on  $[\beta, 1[ \times [\beta, 1[$ .

Then, the Lebesgue theorem allow us to permute integration and limit so that, we have shown that  $\sigma_n^2 \longrightarrow \sigma^2$  and the first assumption of Proposition 4.1 holds.

Let now deal with the second assumption about the maximum of the  $\alpha_{i,n}$ .

For  $j \leq \lfloor n\alpha \rfloor + 1$ , we have

$$\alpha_{j,n} = \frac{(n - \lfloor n\alpha \rfloor + 1)G_{\lfloor n\alpha \rfloor + 1}^n - H^n}{n(n+1)}$$

Using the previous computations, for  $n$  large enough we have

$$\frac{(\alpha_{j,n})^2}{n\sigma_n^2} \underset{n \rightarrow +\infty}{\sim} \frac{(G_{\lfloor n\alpha \rfloor + 1}^n - \frac{T^n}{n})^2}{K_n - \frac{T_n^2}{n^2}} \frac{1}{n}$$

But the convergence

$$\frac{(K^n - (\frac{T^n}{n}))^2}{n^4} \underset{n \rightarrow +\infty}{\longrightarrow} \int_{\alpha}^1 \int_{\alpha}^1 (\min(x, y) - xy)l(x)l(y)dx dy$$

implies the convergence

$$\frac{(G^n - \frac{T^n}{n})^2}{n^4} \underset{n \rightarrow +\infty}{\longrightarrow} \int_{\alpha}^1 (1-x)l(x)dx.$$

Indeed :

$$\int_{\alpha}^1 \int_{\alpha}^1 (\min(x, y) - xy)l(x)l(y)dx dy = \int_{\alpha}^1 \int_{\alpha}^1 (y(1-x))l(x)l(y)dx dy + \int_{\alpha}^1 xl(x) \int_{\alpha}^1 (1-y)l(y)dy dx$$

So that,

$$\frac{(\alpha_{j,n})^2}{n\sigma_n^2} \underset{n \rightarrow +\infty}{\sim} \frac{C}{n} \underset{n \rightarrow +\infty}{\rightarrow} 0$$

when  $n$  goes to infinity. If  $j \geq \lfloor n\alpha \rfloor + 2$  the same property holds as

$$\alpha_{j,n} = \frac{(n+1)G_j^n - H^n - G_{\lfloor n\alpha \rfloor}^n(n\alpha) + 1}{n(n+1)} \underset{n \rightarrow +\infty}{\sim} \frac{(n+1)G_j^n - T^n}{n^2}$$

Hence, we may apply Proposition 4.1 and conclude that

$$\sqrt{n}Q_n \implies \mathcal{N}(O, \sigma^2).$$

where  $\sigma^2 = \int_{\alpha}^1 \int_{\alpha}^1 (\min(x, y) - xy)l(x)l(y)dxdy$ . Finally, just multiply by  $(1 - \alpha)^{-1}$  to get the final result.

#### Step 4 : Conclusion

The Slutsky lemma allows to conclude using the results of steps 1 and 3. □

## 5. NUMERICAL SIMULATIONS

Our numerical tests consist in simulating values from a known theoretical distributions and computing the 0.95-quantiles and superquantiles. For each estimated quantity, the reference value is given via a  $10^6$ -size random sample and a convergence study is performed from a 1000-size sample to a  $10^5$  size sample (with a step of 500). In order to erase the effect of randomness, 50 repetitions of each numerical experiment are made. Then, we compute

- The mean value of the 50 estimations to be compared to the reference value,
- The standard deviation of the 50 estimations. It allows to compute an experimental 95%-confidence interval (CI) to be compared to the theoretical 95%-CI (given by the central limit theorem).

Each is composed with four plots of convergence for the following quantities: quantile (up left), classical superquantile (up right), geometrical superquantile (bottom left) and harmonic superquantile (bottom right). Each convergence plot is composed with the following curves: Reference value (dotted black line), mean estimated values (red circles), theoretical 95%-CI (dashed black line) and experimental 95%-CI (solid blue line).

Figure 1 gives the results for an exponential distribution of parameter  $\lambda = 1$ . As predicted by the theory (see Section 3.3), for the three different superquantiles, the consistency is verified while the experimental CI perfectly fits the theoretical CI (given by the central limit theorem).

We then test the Pareto distribution (see Section 3.3) with three different shape parameters:  $a = 0.5$ ,  $a = 1.5$  and  $a = 2.5$ . Figures 2 ( $a = 0.5$ ), 3 ( $a = 1.5$ ) and 4 ( $a = 2.5$ ) give the convergence results. For the geometrical and harmonic superquantiles, as predicted by the theory (see Section 3.3), the consistency of the Monte Carlo estimation is verified while the experimental CI perfectly fits the theoretical CI (asymptotic normality). For the classical superquantile, we distinguish three different behaviors:

- No consistency for  $a = 0.5$  (Figure 2) (theory predicts consistency only if  $a > 1$ ),

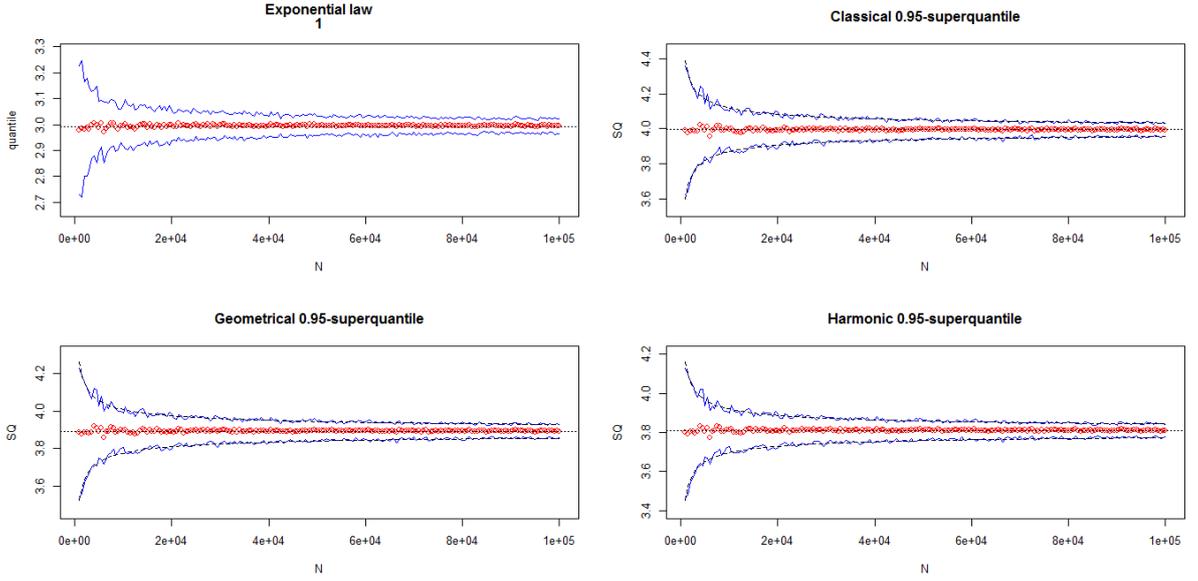


FIGURE 1. Numerical convergence test for the exponential distribution.

- Consistency but no asymptotic normality for  $a = 1.5$  (Figure 3) (theory predicts asymptotic normality only if  $a > 2$ ),
- Consistency and asymptotic normality for  $a = 2.5$  (Figure 4),

## 6. APPLICATIONS TO A NUCLEAR SAFETY EXERCISE

GASCON is a software developed by CEA (French Atomic Energy Commission) to study the potential chronological atmospheric releases and dosimetric impact for nuclear facilities safety assessment [13]. It evaluates, from a fictitious radioactive release, the doses received by a population exposed to the cloud of radionuclides and through the food chains. It takes into account the interactions which exist between humans, plants and animals, the different pathways of transfer (wind, rain, ...), the distance between emission and observation, and the time from emission.

As GASCON is relatively costly in computational time, [13] have built metamodels (of polynomial form) of GASCON outputs in order to perform uncertainty and sensitivity analysis. As in [14], we focus on one output of GASCON, the annual effective dose in  $^{129}\text{I}$  received in one year by an adult who lives in the neighborhood of a particular nuclear facility. Instead of the GASCON software, we will use here the metamodel of this output which depends on 10 input variables, each one modelled by a log-uniform random variable (bounds are defined in [13]). The range of the model output stands on several decades ( $10^{-14}$  to  $10^{-11}$  Sv/year) as shown by Figure 5 which represents the histogram (in logarithmic scale) of  $10^6$  simulated values.

For this kind of numerical simulation exercises, we can be typically interested by safety criteria as 95%-quantile and its associated superquantiles. The idea is to compare these

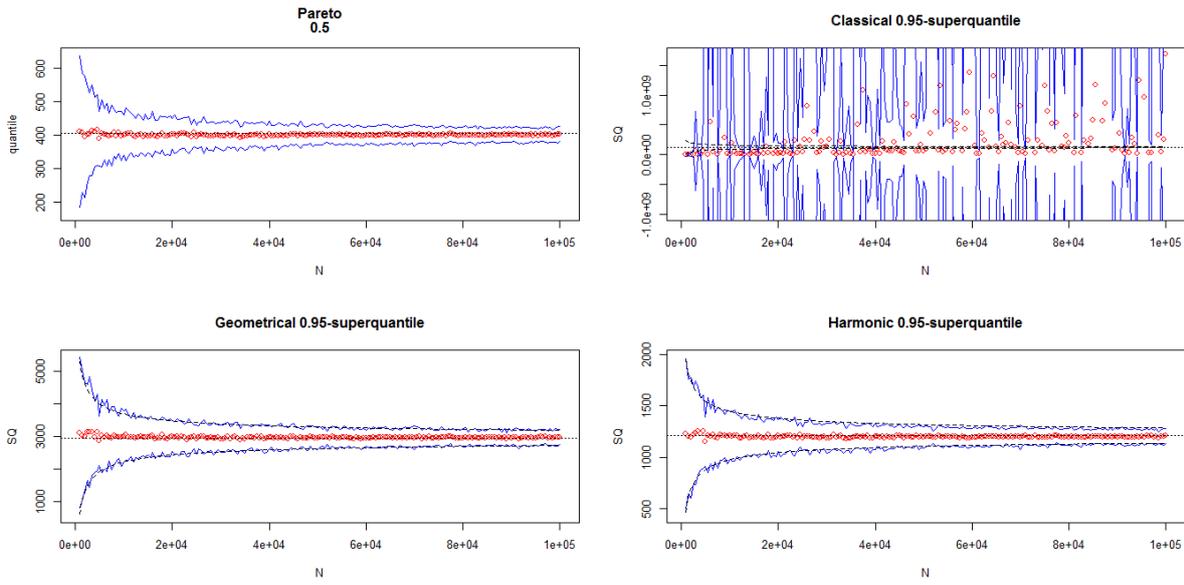


FIGURE 2. Numerical convergence test for the Pareto distribution ( $a = 0.5$ ).

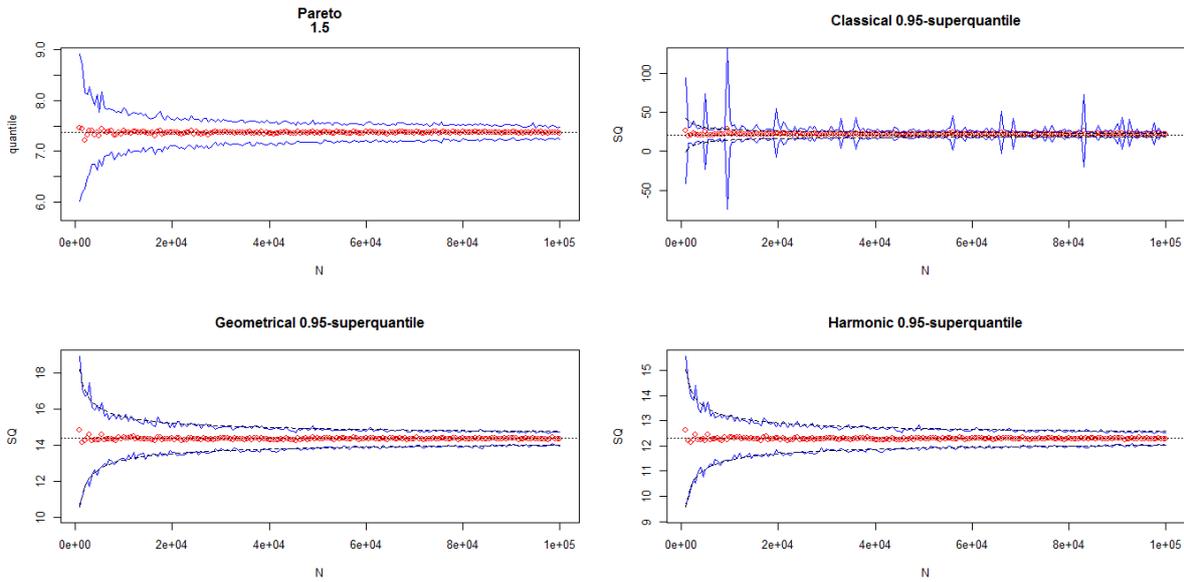


FIGURE 3. Numerical convergence test for the Pareto distribution ( $a = 1.5$ ).

values to regulatory limits or to results coming from other scenarios. In practice, the

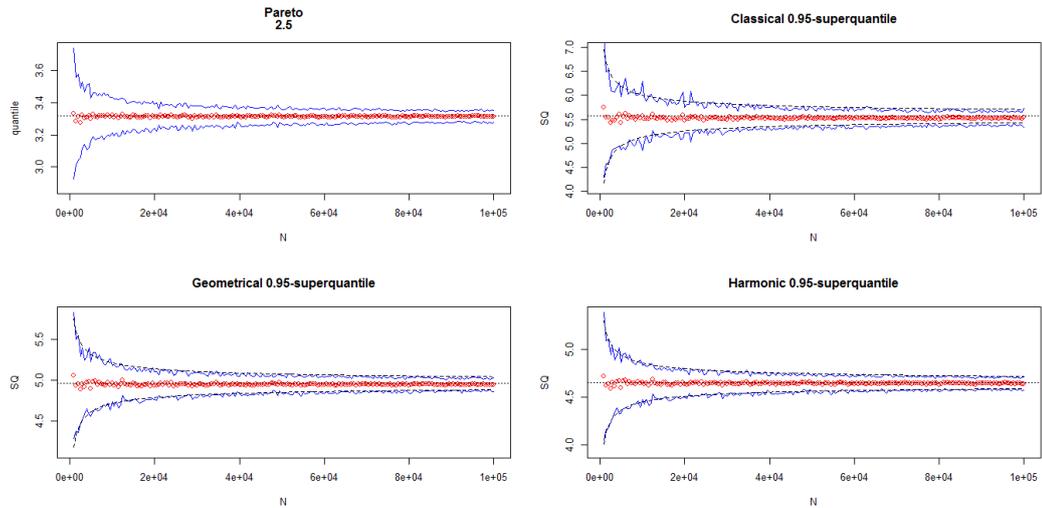


FIGURE 4. Numerical convergence test for the Pareto distribution ( $a = 2.5$ ).

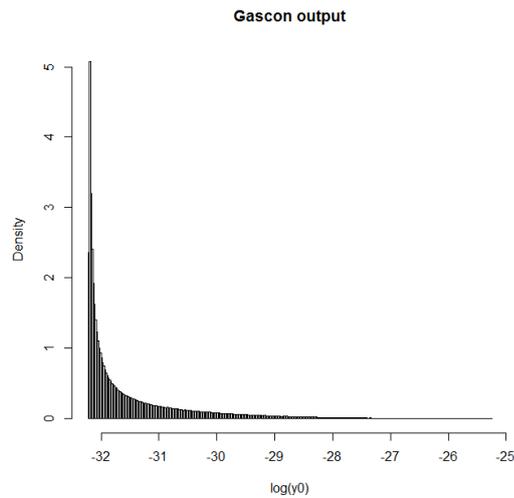
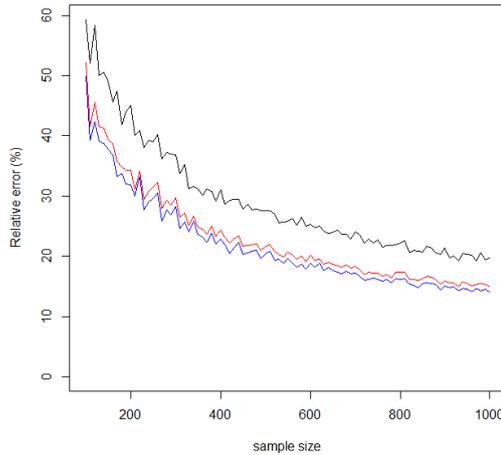


FIGURE 5. Distribution of the GASCON output variable.

number of simulations performed with the GASCON model is several hundreds. Table 1 gives the estimated values of the quantile and superquantiles for 1000 metamodel simulations. Figure 6 shows the relative errors (computed by averaging 1000 different estimations) which are made when estimating the superquantiles using 3 different Bregman divergences and with different sampling sizes. We observe that geometrical and harmonic superquantiles are clearly more precise than the classical one.

TABLE 1. Estimated values of 95%-quantile and 95%-superquantiles for 1000 simulations.

Quantile	Classical superquantile	Geometrical superquantile	Harmonic superquantile
$1.304 \times 10^{-13}$	$4.769 \times 10^{-13}$	$3.316 \times 10^{-13}$	$2.637 \times 10^{-13}$

FIGURE 6. Evolution of the relative errors (mean square error divided by the reference value) on the estimated superquantiles in function of the sample size. In black: classical superquantile; in red: geometrical superquantile; in blue: harmonic superquantile. The reference value has been calculated with  $10^7$  simulations.

## 7. CONCLUSION

The superquantile was introduced because the usual quantile was not subadditive. This quantity is interesting because it satisfies the axioms of a coherent measure of risk. In this paper, we have introduced a new coherent measure of risk with the help of the Bregman divergence associated to a strictly convex function  $\gamma$ . This Bregman superquantile are then rich tools because of the diversity of the functions  $\gamma$  we can choose according to the problem we study. Moreover, we have introduced a Monte Carlo estimator of the Bregman superquantile which is statistically powerful thanks to the strictly convex (and so settling) function  $\gamma$ .

The theoretical properties obtained in this paper are confirmed on several numerical test cases. More precisely, geometrical and harmonic superquantiles are more robust than the classical superquantile. This robustness is particularly important in risk assessment studies. Indeed, when dealing with real data, geometrical and harmonic statistics have been proved to be more relevant than classical statistics. For example [7] prove the

usefulness of the geometrical mean and variance for the analysis of air quality measurements. As an illustration, we have applied the geometrical and harmonic superquantiles on real data coming from a radiological impact code used in the nuclear industry.

Further studies will try to apply these criteria in probabilistic assessment of physical components reliability using numerical simulation codes [12]. However, Monte Carlo estimators are no longer applicable in this context and efficient estimators have to be developed. Ideas involving statistical emulators should be developed (see for example [6] for quantile estimation and [3] for rare event probability estimation).

#### REFERENCES

- [1] Carlo Acerbi and Dirk Tasche. On the coherence of expected shortfall. *Journal of Banking and Finance*, 26(7):1487–1503, 2002.
- [2] P. Artzner, F. Delbaen, J.M Eber, and D. Heath. Coherent measures of risk. 1998.
- [3] J. Bect, D. Ginsbourger, L. Li, V. Picheny, and E. Vazquez. Sequential design of computer experiments for the estimation of a probability of failure. *Statistics and Computing*, 22:773–793, 2012.
- [4] A. Ben-Tal, A. Charnes, and M. Teboulle. Entropic means. *Journal of Mathematical Analysis and Applications*, 139(2):537–551, 1989.
- [5] L. M. Bregman. The relaxation method of finding the common point of convex sets and its application to the solution of problems in convex programming. *USSR computational mathematics and mathematical physics*, 7(3):200–217, 1967.
- [6] C. Cannamela, J. Garnier, and B. Iooss. Controlled stratification for quantile estimation. *Annals of Applied Statistics*, 2:1554–1580, 2008.
- [7] J.C. Chang and S.R. Hanna. Air quality model performance evaluation. *Meteorology and Atmospheric Physics*, 87:167–196, 2004.
- [8] H. Chernoff, J.L. Gastwirth, and M.V. Johns. Asymptotic distribution of linear combinations of functions of order statistics with applications to estimation. *Stanford university, The annals of Mathematical Statistics*, 1965.
- [9] V. Chernozhukov and L. Umantsev. Conditional value at risk, aspects of modeling and estimation. *Empirical Economics*, 2001.
- [10] S.Y. Chun, A. Shapiro, and S. Uryasev. Conditional value atrisk and average value-at-risk : estimation and asymptotics. *Munich Personal RePEc Archive*, 2011.
- [11] David and Nagaraja. *Order Statistics*. Wiley, 2003.
- [12] E. de Rocquigny, N. Devictor, and S. Tarantola, editors. *Uncertainty in industrial practice*. Wiley, 2008.
- [13] B. Iooss, F. Van Dorpe, and N. Devictor. Response surfaces and sensitivity analyses for an environmental model of dose calculations. *Reliability Engineering and System Safety*, 91:1241–1251, 2006.
- [14] J. Jacques, C. Lavergne, and N. Devictor. Sensitivity analysis in presence of model uncertainty and correlated inputs. *Reliability Engineering and System Safety*, 91:1126–1134, 2006.
- [15] L. Kuipers and H. Niederreiter. *Uniform Distribution of Sequences*. Wiley-interscience, 1974.
- [16] G. Ch. Pflug. Some remarks on the value-at-risk and the conditional value-at-risk. 2000.
- [17] Iosif Pinelis. An optimal three-way stable and monotonic spectrum of bounds on quantiles: A spectrum of coherent measures of financial risk and economic inequality. *Risks*, 2(3):349–392, 2014.
- [18] R. Terry Rockafellar, Johannes O. Royset, and S. I. Miranda. Superquantile regression with applications to buffered reliability, uncertainty quantification, and conditional value-at-risk. *European Journal of Operational Research*, 234(1):140–154, 2014.
- [19] R. Tyrrell Rockafellar. *Convex analysis*. Princeton University Press, 1997.
- [20] R.T. Rockafellar. Coherent approaches to risk in optimization under uncertainty. *Tutorials in operation research Informs*, 2007.
- [21] R.T. Rockafellar and J.O. Royset. Random variables, monotone relations, and convex analysis. *Mathematical Programming B, forthcoming*, 2013.

- [22] R.T. Rockafellar and S. Uryasev. Optimization of conditional value at risk. *Portfolio Safeguard*, 1999.
- [23] R.T. Rockafellar and S. Uryasev. The fundamental risk quadrangle in risk management, optimization and statistical estimation. *Surveys in Operations Research and Management Science*, 18(1):33–53, 2013.
- [24] R.T. Rockafellar. The fundamental quadrangle of risk in optimization and estimation. *Risk*, 2010.
- [25] S.Uryasev R.T. Rockafellar. Optimization of conditional value-at-risk. *Journal of risk*, 2:21–42, 2000.
- [26] S. Uryasev. Introduction to the theory of probabilistic functions and percentiles. 2000.
- [27] A. W. Van Der Vaart. *Asymptotic statistics*. Cambridge Series in Statistical and Probabilistic Mathematics. Cambridge University Press, 1998.

FG AG TLR AND BI ARE WITH THE INSTITUT DE MATHÉMATIQUES DE TOULOUSE (CNRS UMR 5219). UNIVERSITÉ PAUL SABATIER, 118 ROUTE DE NARBONNE, 31062 TOULOUSE, FRANCE. BI IS ALSO WITH EDF R&D.