

Semiparametric robust mean estimations based on the orderliness of quantile averages

Tuban Lee

This manuscript was compiled on May 22, 2023

As one of the most fundamental problems in statistics, robust location estimation has many prominent solutions, such as the symmetric trimmed mean, symmetric Winsorized mean, Hodges–Lehmann estimator, Huber M -estimator, and median of means. Recent studies suggest that their maximum biases concerning the mean can be quite different in asymmetric distributions, but the underlying mechanisms and average performance remain largely unclear. In this article, similar to the mean-median-mode inequality, it is proven that within the context of nearly all common unimodal distributions, there is an orderliness of symmetric quantile averages with varying breakdown points. Further deductions explain why the Winsorized mean and median of means typically have smaller biases compared to the trimmed mean. Building on the U -orderliness, the superiority of the median Hodges–Lehmann mean is discussed.

semiparametric | mean-median-mode inequality | asymptotic | unimodal
| Hodges–Lehmann estimator

In 1823, Gauss (1) proved that for any unimodal distribution with a finite second moment, $|m - \mu| \leq \sqrt{\frac{3}{4}}\omega$, where μ is the population mean, m is the population median, and ω is the root mean square deviation from the mode, M . This pioneering work revealed that despite potential bias with respect to the mean in robust estimates, the deviation remains bounded in unit of a scale parameter under certain assumptions. Bernard, Kazzi, and Vanduffel (2020) (2) further derived asymptotic bias bounds of any quantile for unimodal distributions with finite second moments by reducing this optimization problem to a parametric one, which can be solved analytically. They showed that the population median, m , has the smallest maximum distance to the population mean, μ , among all symmetric quantile averages (SQA $_{\epsilon}$). Daniell, in 1920, (3) analyzed a class of estimators, linear combinations of order statistics, and identified that ϵ -symmetric trimmed mean (STM $_{\epsilon}$) belongs to this class. Another popular choice, the ϵ -symmetric Winsorized mean (SWM $_{\epsilon}$), named after Winsor and introduced by Tukey (4) and Dixon (5) in 1960, is also an L -estimator. Bieniek (2016) derived exact bias upper bounds of the Winsorized mean based on Danielak and Rychlik's work (2003) on the trimmed mean for any distribution with a finite second moment and confirmed that the former is smaller than the latter (6, 7). In 1963, Hodges and Lehmann (8) proposed a class of nonparametric location estimators based on rank tests and, from the Wilcoxon signed-rank statistic (9), deduced the median of pairwise means as a robust location estimator for a symmetric population. Both L -statistics and R -statistics achieve robustness essentially by removing a certain proportion of extreme values. In 1964, Huber (10) generalized maximum likelihood estimation to the minimization of the sum of a specific loss function, which measures the residuals between the data points and the model's parameters. Some L -estimators are also M -estimators, e.g., the sample mean is an M -estimator with a squared error loss function, the sample median is an

M -estimator with an absolute error loss function (10). The Huber M -estimator is obtained by applying the Huber loss function that combines elements of both squared error and absolute error to achieve robustness against gross errors and high efficiency for contaminated Gaussian distributions (10). Sun, Zhou, and Fan (2020) examined the concentration bounds of Huber M -estimator (11). Mathieu (2022) (12) further derived the concentration bounds of M -estimators and demonstrated that, by selecting the tuning parameter which depends on the variance, Huber M -estimator can also be a sub-Gaussian estimator. The concept of median of means (MoM $_{k,b} = \frac{n}{k}$, k is the number of size in each block, b is the number of blocks) was implicitly introduced several times in Nemirovsky and Yudin (1983) (13), Jerrum, Valiant, and Vazirani (1986), (14) and Alon, Matias and Szegedy (1996) (15)'s works. Given its good performance even for distributions with infinite second moments, MoM has received increasing attention over the past decade (16–18). Devroye, Lerasle, Lugosi, and Oliveira (2016) showed that MoM nears the optimum of sub-Gaussian mean estimation with regards to concentration bounds when the distribution has a heavy tail (17). For a comparison of concentration bounds of trimmed mean, Huber M -estimator, median of means and other relevant estimators, readers are directed to Gobet, Lerasle, and Métivier's paper (2022) (19). Laforgue, Clemencon, and Bertail (2019) proposed the median of randomized means (MoRM $_{k,b}$) (18), wherein, rather than partitioning, an arbitrary number, b , of blocks are built independently from the sample, and showed that MoRM has a better non-asymptotic sub-Gaussian property compared to MoM. In fact, asymptotically, the Hodges–Lehmann (H-L) estimator is equivalent to MoM $_{k=2,b=\frac{n}{k}}$ and MoRM $_{k=2,b}$, and they can be seen as the pairwise mean distribution is approximated by the sampling without replacement and bootstrap, respectively. For the asymptotic validity, readers are referred to the foundational works of Efron (1979) (20), Bickel and

Significance Statement

In 1964, van Zwet introduced the convex transformation order for comparing the skewness of two distributions. This paradigm shift played a fundamental role in defining robust measures of distributions, from spread to kurtosis. Here, instead of examining the stochastic ordering between two distributions, the orderliness of quantile averages within a distribution is investigated. By classifying distributions through the signs of derivatives, a series of sophisticated robust mean estimators are deduced. Nearly all common nonparametric robust location estimators are found to be special cases thereof.

T.L. designed research, performed research, analyzed data, and wrote the paper.

The author declares no competing interest.

¹To whom correspondence should be addressed. E-mail: tl@biomathematics.org

71 Freedman (1981, 1984) (21, 22), and Helmers, Janssen, and
72 Veraverbeke (1990) (23).

Here, the ϵ, b -stratified mean is defined as

$$SM_{\epsilon, b, n} := \frac{b}{n} \left(\sum_{j=1}^{\frac{b-1}{2b\epsilon}} \sum_{i_j=\frac{(2bj-b-1)n\epsilon}{b-1}+1}^{\frac{(2bj-b+1)n\epsilon}{b-1}} X_{i_j} \right),$$

73 where $X_1 \leq \dots \leq X_n$ denote the order statistics of a sample
74 of n independent and identically distributed random variables
75 X_1, \dots, X_n . $b \in \mathbb{N}$, $b \geq 3$. The definition was further refined to
76 guarantee the continuity of the breakdown point by incorporat-
77 ing an additional block in the center when $\lfloor \frac{b-1}{2b\epsilon} \rfloor \bmod 2 = 0$,
78 or by adjusting the central block when $\lfloor \frac{b-1}{2b\epsilon} \rfloor \bmod 2 = 1$ (SI
79 Text). If the subscript n is omitted, only the asymptotic
80 behavior is considered. If b is omitted, $b = 3$ is assumed.
81 $SM_{\epsilon, b=3}$ is equivalent to STM_{ϵ} , when $\epsilon > \frac{1}{6}$. The basic idea
82 of the stratified mean, when $\frac{b-1}{2\epsilon} \in \mathbb{N}$, $b \bmod 2 = 1$, is to dis-
83 tribute the data into $\frac{b-1}{2\epsilon}$ equal-sized non-overlapping blocks
84 according to their order, then further sequentially group these
85 blocks into b equal-sized strata and compute the mean of the
86 middle stratum, which is the median of means of each stratum.
87 In situations where $i \bmod 1 \neq 0$, a potential solution is to
88 generate multiple smaller samples that satisfy the equality
89 by sampling without replacement, and subsequently calculate
90 the mean of all estimations. The details of determining the
91 sample size and sampling times are provided in the SI Text.
92 Although the principle resembles that of the median of means,
93 without the random shift, $SM_{\epsilon, b, n}$ is different from $MoM_{k=\frac{n}{b}, b}$.
94 Additionally, the stratified mean differs from the mean of the
95 sample obtained through stratified sampling methods, intro-
96 duced by Neyman (1934) (24) or ranked set sampling (25),
97 introduced by McIntyre in 1952, as these sampling methods
98 aim to obtain more representative samples or improve the
99 efficiency of sample estimates, but the sample means based
100 on them are not robust. When $b \bmod 2 = 1$, the stratified
101 mean can be regarded as replacing the other equal-sized strata
102 with the middle stratum, which, in principle, is analogous to
103 the Winsorized mean that replaces extreme values with less
104 extreme percentiles. Furthermore, while the bounds confirm
105 that the Winsorized mean and median of means outperform
106 the trimmed mean (6, 7, 17, 19) in worst-case performance,
107 the complexity of bound analysis makes it difficult to achieve a
108 complete and intuitive understanding of these results. Also, a
109 clear explanation for the average performance of them remains
110 elusive. The aim of this paper is to define a series of semi-
111 parametric models using the signs of derivatives, reveal their
112 elegant interrelations and connections to parametric models,
113 and show that by exploiting these models, a set of sophisti-
114 cated mean estimators can be deduced, which exhibit strong
115 robustness to departures from assumptions.

116 Quantile average and weighted average

The symmetric trimmed mean, symmetric Winsorized mean, and stratified mean are all L -estimators. More specifically, they are symmetric weighted averages, which are defined as

$$SWA_{\epsilon, n} := \frac{\sum_{i=1}^{\lceil \frac{n}{2} \rceil} \frac{X_i + X_{n-i+1}}{2} w_i}{\sum_{i=1}^{\lceil \frac{n}{2} \rceil} w_i},$$

where w_i s are the weights applied to the symmetric quantile
averages according to the definition of the corresponding L -
estimators. For example, for the ϵ -symmetric trimmed mean,
 $w_i = \begin{cases} 0, & i < n\epsilon \\ 1, & i \geq n\epsilon \end{cases}$, provided that $n\epsilon \in \mathbb{N}$. The mean and
median are indeed two special cases of the symmetric trimmed
mean.

To extend the symmetric quantile average to the asymmet-
ric case, there are two possible definitions for the ϵ, γ -quantile
average (QA(ϵ, γ, n)), i.e.,

$$\frac{1}{2}(\hat{Q}_n(\gamma\epsilon) + \hat{Q}_n(1 - \epsilon)), \quad [1]$$

and

$$\frac{1}{2}(\hat{Q}_n(\epsilon) + \hat{Q}_n(1 - \gamma\epsilon)), \quad [2]$$

where $\gamma \geq 0$ and $0 \leq \epsilon \leq \frac{1}{1+\gamma}$, $\hat{Q}_n(p)$ is the empirical quantile
function. For trimming from both sides, [1] and [2] are essen-
tially equivalent. [1] is assumed in this article unless otherwise
specified, since many common asymmetric distributions are
right-skewed, and [1] allows trimming only from the right side
by setting $\gamma = 0$.

Analogously, the weighted average can be defined as

$$WA_{\epsilon, \gamma, n} := \frac{\int_{\epsilon_0=0}^{\frac{1}{1+\gamma}} QA(\epsilon_0, \gamma, n) w_{\epsilon_0}}{\int_{\epsilon_0=0}^{\frac{1}{1+\gamma}} w_{\epsilon_0}}.$$

For instance, the ϵ, γ -trimmed mean ($TM_{\epsilon, \gamma, n}$) is a weighted
average with a left trim size of $\gamma\epsilon n$ and a right trim size of ϵn ,
where $w_{\epsilon_0} = \begin{cases} 0, & \epsilon_0 < \epsilon \\ 1, & \epsilon_0 \geq \epsilon \end{cases}$. Using this definition, even $\gamma\epsilon n \notin \mathbb{N}$
or $\epsilon n \notin \mathbb{N}$, the TM computation remains unaltered since this
definition is based on the empirical quantile function. However,
considering the computational cost in practice, here, the non-
asymptotic definitions of various types of weighted averages,
in most cases, are essentially based on order statistics. The
solution to the decimal issue of them is the same as that in
SM, unless stated otherwise.

145 Classifying distributions by the signs of derivatives

Let \mathcal{P}_k denote the set of all distributions over \mathbb{R} whose mo-
ments, from the first to the k th, are all finite. Without loss
of generality, the discussion of all the classes outlined below
is restricted to the intersection with the nonparametric class
of distributions $\mathcal{P}_7^k := \{\text{All continuous distribution } P \in \mathcal{P}_k\}$.
Besides fully and smoothly parameterizing by a Euclidean pa-
rameter or just assuming regularity conditions, there are many
ways to classify distributions. In 1956, Stein initiated the
problem of estimating parameters in the presence of an infinite
dimensional nuisance shape parameter (26). A notable exam-
ple discussed in his groundbreaking work was the estimation
of the center of symmetry for an unknown symmetric distribu-
tion. In 1993, Bickel, Klaassen, Ritov, and Wellner published
an influential semiparametrics textbook (27) which system-
atically categorized many common models into three classes:
parametric, nonparametric, and semiparametric. Yet, there
is another old and commonly encountered class of distribu-
tions that receives little attention in semiparametric literature:
the unimodal distribution. It is a very unique semiparametric

model because its definition is based on the signs of derivatives, i.e., for a continuous distribution, $(f'(x) > 0 \text{ for } x \leq M) \wedge (f'(x) < 0 \text{ for } x \geq M)$. Let \mathcal{P}_U denote the set of all unimodal distributions. There was a widespread misbelief that the median of an arbitrary unimodal distribution always lies between its mean and mode until Runnenburg (1978) and van Zwet (1979) (28, 29) endeavored to determine sufficient conditions for the inequality to hold, thereby implying the possibility of its violation. The class of distributions that satisfy the mean-median-mode inequality constitutes a subclass of \mathcal{P}_U . By analogy, a right-skewed distribution is called γ -ordered, if and only if

$$\forall 0 \leq \epsilon_1 \leq \epsilon_2 \leq \frac{1}{1+\gamma}, \text{QA}_{\epsilon_1, \gamma} \geq \text{QA}_{\epsilon_2, \gamma}.$$

The necessary and sufficient condition below hints at the relation between the mean-median-mode inequality and the γ -orderliness.

Theorem .1. Let P_γ^k represent an arbitrary distribution in the set \mathcal{P}_γ^k . P_γ^k is γ -ordered if and only if the probability density function (pdf) satisfies the inequality $f(Q(\gamma\epsilon)) \geq f(Q(1-\epsilon))$ for all $0 \leq \epsilon \leq \frac{1}{1+\gamma}$ or $f(Q(\gamma\epsilon)) \leq f(Q(1-\epsilon))$ for all $0 \leq \epsilon \leq \frac{1}{1+\gamma}$, where $\gamma \geq 0$.

Proof. Without loss of generality, consider the case of right-skewed continuous distribution. From the definition of γ -orderliness, it is deduced that $\frac{Q(\gamma\epsilon-\delta)+Q(1-\epsilon+\delta)}{2} \geq \frac{Q(\gamma\epsilon)+Q(1-\epsilon)}{2} \Leftrightarrow Q(\gamma\epsilon-\delta)-Q(\gamma\epsilon) \geq Q(1-\epsilon)-Q(1-\epsilon+\delta) \Leftrightarrow Q'(1-\epsilon) \geq Q'(\gamma\epsilon)$, where δ is an infinitesimal positive quantity. Observing that the quantile function is the inverse function of the cumulative distribution function (cdf), $Q'(1-\epsilon) \geq Q'(\gamma\epsilon) \Leftrightarrow F'(Q(\gamma\epsilon)) \geq F'(Q(1-\epsilon))$, thereby completing the proof, given that the derivative of cdf is pdf. \square

According to Theorem .1, if a probability distribution is right-skewed and monotonic, it will always be γ -ordered, provided $\gamma \geq 0$. For a right-skewed continuous unimodal distribution, if $Q(\gamma\epsilon) > M$, the inequality $f(Q(\gamma\epsilon)) \geq f(Q(1-\epsilon))$ holds. The principle is extendable to unimodal-like distributions. Suppose there is a right-skewed continuous multimodal distribution following the mean- γ -median-first mode inequality with several smaller modes on the right side, with the first mode, M , having the greatest probability density, and the γ -median, $Q(\frac{\gamma}{1+\gamma})$, falling within the first dominant mode (i.e., if $x > Q(\frac{\gamma}{1+\gamma})$, $f(Q(\frac{\gamma}{1+\gamma})) \geq f(x)$), then if $Q(\gamma\epsilon) > M$, the inequality $f(Q(\gamma\epsilon)) \geq f(Q(1-\epsilon))$ also holds. In other words, while a distribution following the mean- γ -median-mode inequality may not be strictly γ -ordered, the inequality that defines γ -orderliness remains valid for most quantile averages. The mean- γ -median-mode inequality can also indicate possible bounds for γ in practice, e.g., for any distributions, when $\gamma \rightarrow \infty$, the γ -median will be greater than the mean and the mode, when $\gamma \rightarrow 0$, the γ -median will be smaller than the mean and the mode.

Consider the sign of the derivative of the quantile average with respect to the breakdown point, the above definition of γ -orderliness can also be expressed as

$$\forall 0 \leq \epsilon \leq \frac{1}{1+\gamma}, \frac{\partial \text{QA}_{\epsilon, \gamma}}{\partial \epsilon} \leq 0.$$

The left-skewed case can be obtained by reversing the inequality $\frac{\partial \text{QA}_{\epsilon, \gamma}}{\partial \epsilon} \leq 0$ to $\frac{\partial \text{QA}_{\epsilon, \gamma}}{\partial \epsilon} \geq 0$ and employing the second definition of QA, as given in [2]; for simplicity, it will be omitted in the following discussion. If $\gamma = 1$, the γ -ordered distribution is referred to as ordered.

Furthermore, many common right-skewed distributions are partial bounded, indicating a convex behavior of the QA function when $\epsilon \rightarrow 0$. By further assuming convexity, the second γ -orderliness can be defined as follows for a right-skewed distribution,

$$\forall 0 \leq \epsilon \leq \frac{1}{1+\gamma}, \frac{\partial^2 \text{QA}_{\epsilon, \gamma}}{\partial \epsilon^2} \geq 0 \wedge \frac{\partial \text{QA}_{\epsilon, \gamma}}{\partial \epsilon} \leq 0.$$

Analogously, the ν th γ -orderliness of a right-skewed distribution can be defined as $(-1)^\nu \frac{\partial^\nu \text{QA}_{\epsilon, \gamma}}{\partial \epsilon^\nu} \geq 0 \wedge \dots \wedge -\frac{\partial \text{QA}_{\epsilon, \gamma}}{\partial \epsilon} \geq 0$. If $\gamma = 1$, the ν th γ -orderliness is referred as ν th orderliness. Let \mathcal{P}_O denote the set of all distributions that are ordered and \mathcal{P}_{O_ν} and $\mathcal{P}_{\gamma O_\nu}$ represent the sets of all distributions that are ν th ordered and ν th γ -ordered, respectively. When the shape parameter of the Weibull distribution, α , is smaller than 3.258, it can be shown that the Weibull distribution belong to $\mathcal{P}_U \cap \mathcal{P}_O \cap \mathcal{P}_{O_2} \cap \mathcal{P}_{O_3}$ (SI Text). At $\alpha \approx 3.602$, the Weibull distribution is symmetric, and as $\alpha \rightarrow \infty$, the skewness of the Weibull distribution reaches 1. Therefore, the parameters that let it not be included in the set correspond to cases when it is near-symmetric, as shown in the SI Text. Nevertheless, computing the derivatives of the QA function is often intricate and, at times, challenging. The following theorems establish the relationship between \mathcal{P}_O , \mathcal{P}_{O_ν} , and $\mathcal{P}_{\gamma O_\nu}$, and a wide range of other semi-parametric distributions. They can be used to quickly identify some parametric distributions in \mathcal{P}_O , \mathcal{P}_{O_ν} , and $\mathcal{P}_{\gamma O_\nu}$.

Theorem .2. For any random variable X whose probability distribution function belongs to a location-scale family, the distribution is ν th γ -ordered if and only if the family of probability distributions is ν th γ -ordered.

Proof. Let Q_0 denote the quantile function of the standard distribution without any shifts or scaling. After a location-scale transformation, the quantile function is $Q(p) = \lambda Q_0(p) + \mu$, where λ is the scale parameter and μ is the location parameter. According to the definition of the ν th γ -orderliness, the signs of derivatives of the QA function are invariant after this transformation. As the location-scale transformation is reversible, the proof is complete. \square

Theorem .2 demonstrates that in the analytical proof of the ν th γ -orderliness of a parametric distribution, both the location and scale parameters can be regarded as constants. It is also instrumental in proving other theorems, as illustrated below.

Theorem .3. Any symmetric distribution with a finite second moment is ν th ordered.

Proof. Without loss of generality, assuming continuity and $m = 0$. A symmetric distribution is a probability distribution such that for all x , $f(x) = f(-x)$. Its cdf satisfies $F(x) = 1 - F(-x)$. Let $x = Q(p)$, then, $F(Q(p)) = p = 1 - F(-Q(p))$ and $F(Q(1-p)) = 1 - p \Leftrightarrow p = 1 - F(Q(1-p))$. Therefore, $F(-Q(p)) = F(Q(1-p))$. Since the cdf is monotonic, $-Q(p) = Q(1-p) \Leftrightarrow Q(p) + Q(1-p) = 0$. As a result, all symmetric

quantile averages coincide; the ν th order derivative is zero. The case of $m \neq 0$ follows directly from Theorem 2. \square

As a consequence of Theorem 3 and the fact that generalized Gaussian distribution is symmetric around the median, it is ν th ordered.

Theorem 4. Any continuous right-skewed distribution whose quantile function Q satisfies $Q^{(\nu)}(p) \geq 0 \wedge \dots \wedge Q^{(i)}(p) \geq 0 \wedge \dots \wedge Q^{(2)}(p) \geq 0$, $i \bmod 2 = 0$, is ν th γ -ordered, provided that $0 \leq \gamma \leq 1$.

Proof. Since $(-1)^i \frac{\partial^i Q_{A_{\epsilon}, \gamma}}{\partial \epsilon^i} = \frac{1}{2}((- \gamma)^i Q^i(\gamma \epsilon) + Q^i(1 - \epsilon))$, for $0 \leq \epsilon \leq \frac{1}{1+\gamma}$ and $1 \leq i \leq \nu$, when $i \bmod 2 = 0$, $(-1)^i \frac{\partial^i Q_{A_{\epsilon}, \gamma}}{\partial \epsilon^i} \geq 0$ for all $\gamma \geq 0$. When $i \bmod 2 = 1$, if further assuming $0 \leq \gamma \leq 1$, $(-1)^i \frac{\partial^i Q_{A_{\epsilon}, \gamma}}{\partial \epsilon^i} \geq 0$, since $Q^{(i+1)}(p) \geq 0$. \square

It is now straightforward to prove that the Pareto distribution follows the ν th γ -orderliness, provided that $0 \leq \gamma \leq 1$, since the quantile function of the Pareto distribution is $Q(p) = x_m(1-p)^{-\frac{1}{\alpha}}$, where $x_m > 0$, $\alpha > 0$, and so $Q^{(\nu)}(p) \geq 0$ for all $\nu \in \mathbb{N}$ according to the chain rule.

Theorem 5. A right-skewed continuous distribution with a monotonic decreasing pdf is second γ -ordered.

Proof. A monotonic decreasing pdf implies $f'(x) = F^{(2)}(x) \leq 0$. Since $Q'(p) \geq 0$, let $x = Q(F(x))$, then by differentiating both sides of the equation twice, one can obtain $0 = Q^{(2)}(F(x))(F'(x))^2 + Q'(F(x))F^{(2)}(x) \Leftrightarrow Q^{(2)}(F(x)) = -\frac{Q'(F(x))F^{(2)}(x)}{(F'(x))^2} \geq 0$. The desired result is derived from Theorem 1 and 4. \square

Theorem 5 provides valuable insights into the relation between modality and orderliness. The conventional definition states that a distribution with a monotonic pdf is still considered unimodal. However, within its supported interval, the mode number is zero. The number of modes and their magnitudes within a distribution are closely related to the possibility of orderliness being valid, although counterexamples can always be constructed for non-monotonic distributions. A proof of the second γ -orderliness, if $\gamma > 0$, can be easily established for the gamma distributions when $\alpha \leq 1$ as the pdf of the gamma distribution is $f(x) = \frac{\lambda^{-\alpha} x^{\alpha-1} e^{-\frac{x}{\lambda}}}{\Gamma(\alpha)}$, where $x \geq 0$, $\lambda > 0$, $\alpha > 0$, Γ is the gamma function, it is a product of two monotonic decreasing functions under constraints. For $\alpha > 1$, the proof becomes challenging. Numerical results show that the orderliness is valid until $\alpha > 140$, the second orderliness is valid until $\alpha > 78$, and the third orderliness is valid until $\alpha > 55$ (SI Text). It is instructive to consider that when $\alpha \rightarrow \infty$ the gamma distribution converges to a Gaussian distribution with mean $\mu = \alpha\lambda$ and variance $\sigma = \alpha\lambda^2$. The skewness of the gamma distribution, $\frac{\alpha+2}{\sqrt{\alpha(\alpha+1)}}$, is monotonic with respect to α , since $\frac{\partial \mu_3(\alpha)}{\partial \alpha} = \frac{-3\alpha-2}{2(\alpha(\alpha+1))^{3/2}} < 0$. When $\alpha = 55$, $\mu_3(\alpha) = 1.027$. Therefore, similar to the Weibull distribution, the parameters that let the distribution not be included in $\mathcal{P}_U \cap \mathcal{P}_O \cap \mathcal{P}_{O_2} \cap \mathcal{P}_{O_3}$ also correspond to cases when it is near-symmetric.

Theorem 6. Consider a symmetric random variable X . Let it be transformed using a function $\phi(x)$ such that $\phi^{(2)}(x) \geq 0 \wedge \phi'(x) \geq 0$ over the interval supported, the resulting convex transformed distribution is ordered. Moreover, if the quantile function of X satisfies $Q^{(2)}(\epsilon) \leq 0$, the convex transformed distribution is second ordered.

Proof. Let $\phi\text{SQA}(\epsilon) = \frac{1}{2}(\phi(Q(\epsilon)) + \phi(Q(1-\epsilon)))$, $0 \leq \epsilon \leq \frac{1}{2}$. Then, $\frac{d\phi\text{SQA}}{d\epsilon} = \frac{1}{2}(\phi'(Q(\epsilon))Q'(\epsilon) - \phi'(Q(1-\epsilon))Q'(1-\epsilon)) = \frac{1}{2}Q'(\epsilon)(\phi'(Q(\epsilon)) - \phi'(Q(1-\epsilon))) \leq 0$, since for a symmetric distribution, $m - Q(\epsilon) = Q(1-\epsilon) - m$, differentiating both sides, $-Q'(\epsilon) = -Q'(1-\epsilon)$, where $Q'(\epsilon) \geq 0$, $\phi^{(2)}(x) \geq 0$. If further differentiating the equality, $Q^{(2)}(\epsilon) = -Q^{(2)}(1-\epsilon)$. Since $\frac{d^{(2)}\phi\text{SQA}}{d\epsilon^{(2)}} = \frac{1}{2}(\phi^2(Q(\epsilon))(Q'(\epsilon))^2 + \phi^2(Q(1-\epsilon))(Q'(1-\epsilon))^2) + \frac{1}{2}(\phi'(Q(\epsilon))(Q^2(\epsilon)) + \phi'(Q(1-\epsilon))(Q^2(1-\epsilon))) = \frac{1}{2}((\phi^{(2)}(Q(\epsilon)) + \phi^{(2)}(Q(1-\epsilon)))(Q'(\epsilon))^2) + \frac{1}{2}((\phi'(Q(\epsilon)) - \phi'(Q(1-\epsilon)))Q^{(2)}(\epsilon))$. If $Q^{(2)}(\epsilon) \leq 0$, $\frac{d^{(2)}\phi\text{SQA}}{d\epsilon^{(2)}} \geq 0$. \square

The mean-median-mode inequality for distributions of the powers and roots of the variates of a given distribution was investigated by Henry Rietz in 1927 (30), but the most straightforward solution is the exponential transformation since the derivatives are invariably positive. An application of Theorem 6 is that the lognormal distribution is ordered as it is exponentially transformed from the Gaussian distribution. The quantile function of the Gaussian distribution meets the condition $Q^{(2)}(\epsilon) = -2\sqrt{2}\pi\sigma e^{2\text{erfc}^{-1}(2\epsilon)^2} \text{erfc}^{-1}(2\epsilon) \leq 0$, where σ is the standard deviation, erfc denotes the complementary error function. Thus, the lognormal distribution is second ordered. Numerical results suggest that it is also third ordered, although an analytical proof is challenging.

Theorem 6 also reveals a relation between convex transformation and orderliness, since ϕ is the non-decreasing convex function in van Zwet's trailblazing work *Convex transformations of random variables* (31). Consider a near-symmetric distribution S , such that SQA_ϵ as a function of ϵ fluctuates from 0 to $\frac{1}{2}$, with $\mu = m$. By definition, S is not ordered. Let s be the pdf of S . Applying the transformation $\phi(x)$ to S decreases $s(Q_S(\epsilon))$, and the decrease rate, due to the order, is much smaller for $s(Q_S(1-\epsilon))$. As a consequence, as the second derivative of $\phi(x)$ increases, eventually, after a point, $s(Q_S(\epsilon))$ becomes greater than $s(Q_S(1-\epsilon))$ even if it was not previously. Thus, the SQA_ϵ function becomes monotonically decreasing, and S becomes ordered. Accordingly, in a family of distributions that differ by a skewness-increasing transformation in van Zwet's sense, violations of orderliness typically occur only when the distribution is near-symmetric.

Pearson proposed using the mean-median difference $\mu - m$ as a measure of skewness after standardization in 1895 (32). Bowley (1926) proposed a measure of skewness based on the SQA -median difference $\text{SQA}_\epsilon - m$ (33). Groeneveld and Meeden (1984) (34) generalized these measures of skewness based on van Zwet's convex transformation (31) while exploring their properties. A distribution is called monotonically right-skewed if and only if $\forall 0 \leq \epsilon_1 \leq \epsilon_2 \leq \frac{1}{2}$, $\text{SQA}_{\epsilon_1} - m \geq \text{SQA}_{\epsilon_2} - m$. Since m is a constant, the monotonic skewness is equivalent to the orderliness. For a nonordered distribution, the signs of

343 $SQA_\epsilon - m$ with different breakdown points might be different,
 344 implying that some skewness measures indicate left-skewed
 345 distribution, while others suggest right-skewed distribution.
 346 Although it seems reasonable that the distribution is likely be
 347 generally near-symmetric, however, counterexamples can be
 348 constructed. For example, consider the Weibull distribution,
 349 when $\alpha > \frac{1}{1-\ln(2)}$, it is near-symmetric and nonordered, the
 350 non-monotonicity of the SQA function arises when ϵ is close to
 351 $\frac{1}{2}$. Replacing the third quartile with one from a right-skewed
 352 heavy-tailed distribution leads to a right-skewed, heavy-tailed,
 353 and nonordered distribution. Therefore, the validity of robust
 354 measures of skewness based on the SQA-median difference is
 355 closely related to the orderliness of the distribution.

356 **Data Availability.** Data for Figure ?? are given in SI Dataset
 357 S1. All codes have been deposited in [GitHub](#).

358 **ACKNOWLEDGMENTS.** I sincerely acknowledge the insightful
 359 comments from the editor which considerably elevated the lucidity
 360 and merit of this paper.

361 1. CF Gauss, *Theoria combinationis observationum erroribus minimis obnoxiae*. (Henricus
 362 Dieterich), (1823).
 363 2. C Bernard, R Kazzi, S Vanduffel, Range value-at-risk bounds for unimodal distributions under
 364 partial information. *Insur. Math. Econ.* **94**, 9–24 (2020).
 365 3. P Daniell, Observations weighted according to order. *Am. J. Math.* **42**, 222–236 (1920).
 366 4. JW Tukey, A survey of sampling from contaminated distributions in *Contributions to probability*
 367 *and statistics*. (Stanford University Press), pp. 448–485 (1960).
 368 5. WJ Dixon, Simplified Estimation from Censored Normal Samples. *The Annals Math. Stat.* **31**,
 369 385–391 (1960).
 370 6. K Danielak, T Rychlik, Theory & methods: Exact bounds for the bias of trimmed means. *Aust.*
 371 *& New Zealand J. Stat.* **45**, 83–96 (2003).
 372 7. M Bieniek, Comparison of the bias of trimmed and winsorized means. *Commun. Stat. Methods*
 373 **45**, 6641–6650 (2016).
 374 8. J Hodges Jr, E Lehmann, Estimates of location based on rank tests. *The Annals Math. Stat.*
 375 **34**, 598–611 (1963).
 376 9. F Wilcoxon, Individual comparisons by ranking methods. *Biom. Bull.* **1**, 80–83 (1945).
 377 10. PJ Huber, Robust estimation of a location parameter. *Ann. Math. Stat.* **35**, 73–101 (1964).
 378 11. Q Sun, WX Zhou, J Fan, Adaptive huber regression. *J. Am. Stat. Assoc.* **115**, 254–265 (2020).
 379 12. T Mathieu, Concentration study of m-estimators using the influence function. *Electron. J. Stat.*
 380 **16**, 3695–3750 (2022).
 381 13. AS Nemirovskij, DB Yudin, *Problem complexity and method efficiency in optimization*. (Wiley-
 382 Interscience), (1983).
 383 14. MR Jerrum, LG Valiant, VV Vazirani, Random generation of combinatorial structures from a
 384 uniform distribution. *Theor. computer science* **43**, 169–188 (1986).
 385 15. N Alon, Y Matias, M Szegedy, The space complexity of approximating the frequency moments
 386 in *Proceedings of the twenty-eighth annual ACM symposium on Theory of computing*. pp.
 387 20–29 (1996).
 388 16. PL Bühlmann, Bagging, subagging and bragging for improving some prediction algorithms
 389 in *Research report/Seminar für Statistik, Eidgenössische Technische Hochschule (ETH)*.
 390 (Seminar für Statistik, Eidgenössische Technische Hochschule (ETH), Zürich), Vol. 113,
 391 (2003).
 392 17. L Devroye, M Lerasle, G Lugosi, RI Oliveira, Sub-gaussian mean estimators. *The Annals Stat.*
 393 **44**, 2695–2725 (2016).
 394 18. P Laforgue, S Cléménçon, P Bertail, On medians of (randomized) pairwise means in *Inter-
 395 national Conference on Machine Learning*. (PMLR), pp. 1272–1281 (2019).
 396 19. E Gobet, M Lerasle, D Métivier, Mean estimation for Randomized Quasi Monte Carlo method.
 397 working paper or preprint (2022).
 398 20. B Efron, Bootstrap methods: Another look at the jackknife. *The Annals Stat.* **7**, 1–26 (1979).
 399 21. PJ Bickel, DA Freedman, Some asymptotic theory for the bootstrap. *The annals statistics* **9**,
 400 1196–1217 (1981).
 401 22. PJ Bickel, DA Freedman, Asymptotic normality and the bootstrap in stratified sampling. *The*
 402 *annals statistics* **12**, 470–482 (1984).
 403 23. R Helmers, P Janssen, N Veraverbeke, *Bootstrapping U-quantiles*. (CWI. Department of
 404 Operations Research, Statistics, and System Theory [BS]), (1990).
 405 24. J Neyman, On the two different aspects of the representative method: The method of stratified
 406 sampling and the method of purposive selection. *J. Royal Stat. Soc.* **97**, 558–606 (1934).
 407 25. G McIntyre, A method for unbiased selective sampling, using ranked sets. *Aust. journal*
 408 *agricultural research* **3**, 385–390 (1952).
 409 26. C Stein, , et al., Efficient nonparametric testing and estimation in *Proceedings of the third*
 410 *Berkeley symposium on mathematical statistics and probability*. Vol. 1, pp. 187–195 (1956).
 411 27. P Bickel, CA Klaassen, Y Ritov, JA Wellner, *Efficient and adaptive estimation for semiparamet-
 412 ric models*. (Springer) Vol. 4, (1993).
 413 28. JT Runnenburg, Mean, median, mode. *Stat. Neerlandica* **32**, 73–79 (1978).
 414 29. Wv Zwet, Mean, median, mode ii. *Stat. Neerlandica* **33**, 1–5 (1979).
 415 30. H Rietz, On certain properties of frequency distributions of the powers and roots of the variates
 416 of a given distribution. *Proc. Natl. Acad. Sci.* **13**, 817–820 (1927).
 417 31. WR van Zwet, *Convex Transformations of Random Variables: Nebst Stellingen*. (1964).

32. K Pearson, X. contributions to the mathematical theory of evolution.—ii. skew variation in 418
 homogeneous material. *Philos. Transactions Royal Soc. London.(A.)* **186**, 343–414 (1895). 419
 33. AL Bowley, *Elements of statistics*. (King) No. 8, (1926). 420
 34. RA Groeneveld, G Meeden, Measuring skewness and kurtosis. *J. Royal Stat. Soc. Ser. D* 421
(The Stat. **33**, 391–399 (1984). 422