

Near-consistent robust estimations of moments for unimodal distributions

Tuban Lee^{a,1}

^aInstitute of Biomathematics, Macau SAR 999078, China

This manuscript was compiled on March 20, 2023

Descriptive statistics for parametric models currently heavily rely on the accuracy of distributional assumptions. Here, based on the invariant structures of unimodal distributions, a series of sophisticated yet efficient estimators, robust to both gross errors and departures from parametric assumptions, are proposed for estimating mean and central moments with insignificant asymptotic biases for common unimodal distributions. This article also illuminates the understanding of the common nature of probability distributions and the measures of them.

orderliness | invariant | unimodal | adaptive estimation | U -statistics

The asymptotic inconsistencies between sample mean (\bar{x}) and nonparametric robust location estimators in asymmetric distributions on the real line have been noticed for more than two centuries (1), yet remain unsolved. Strictly speaking, it is unsolvable as by trimming, some information about the original distribution is removed, making it impossible to estimate the values of the removed parts without distributional assumptions. Newcomb (1886, 1912) provided the first modern approach to robust parametric estimation by developing a class of estimators that gives "less weight to the more discordant observations" (2, 3). In 1964, Huber (4) used the minimax procedure to obtain M-estimator for the contaminated normal distribution, which has played a pre-eminent role in the later development of robust statistics. However, as previously demonstrated, under growing asymmetric departures from normality, the bias of the Huber M-estimator increases rapidly. This is a common issue in parameter estimations. For example, He and Fung (1999) constructed (5) a robust M-estimator for the two-parameter Weibull distribution, from which all moments can be calculated. Nonetheless, it is inadequate for the gamma, Perato, lognormal, and the generalized Gaussian distributions (SI Dataset S1). Another old and interesting approach is arithmetically computing the parameters using one or more L -statistics as inputs, such as percentile estimators. Examples of percentile estimators for the Weibull distribution, the reader is referred to Menon (1963) (6), Dubey (1967) (7), Hassanein (1971) (8), Marks (2005) (9), and Boudt, Caliskan, and Croux (2011) (10)'s works. At the outset of the study of percentile estimators, it was known that they arithmetically utilizes the invariant structures of probability distributions (6, 11, 12). Maybe such estimators can be named as I -statistics. Formally, an estimator is classified as an I -statistic if it asymptotically satisfies $I(LE_1, \dots, LE_l) = (\theta_1, \dots, \theta_q)$ for the distribution it is consistent, where LEs are calculated with the use of L -statistics, I is defined using arithmetic operations and constants, but it may also incorporate other functions, and θ s are the population parameters it estimates. A subclass of I -statistics, arithmetic I -statistics, is defined as LEs are L -statistics, I is solely defined using arithmetic operations and constants.

Since some percentile estimators use the logarithmic function to transform all random variables before computing the L -statistics, a percentile estimator might not always be an arithmetic I -statistic (7). In this article, two subclasses of I -statistics are introduced, arithmetic I -statistics and quantile I -statistics. Examples of quantile I -statistics will be discussed later. Based on L -statistics, I -statistics are naturally robust. Compared to probability density functions (pdfs) and cumulative distribution functions (cdfs), the quantile functions of many parametric distributions are more elegant. Since the expectation of an L -statistic can be expressed as an integral of the quantile function, I -statistics are often analytically obtainable. However, the performance of the aforementioned examples is often worse than that of the robust M -statistics when the distributional assumption is violated (SI Dataset S1). Even when distributions such as the Weibull and gamma belong to the same larger family, the generalized gamma distribution, a misassumption can still result in substantial biases, rendering the approach ill-suited.

In previous research on semiparametric robust mean estimation, the binomial mean (BM_ϵ) is still inconsistent for any skewed distribution, despite having much smaller asymptotic biases than other weighted averages. All robust location estimators commonly used are symmetric due to the universality of the symmetric distributions. One can construct an asymmetric weighted average that is consistent for a semiparametric class of skewed distributions. This approach has been investigated previously, but its lack of symmetry makes it suitable only for certain applications (13). Shifting from semiparametrics to parametrics, an ideal robust location estimator would have a non-sample-dependent breakdown point (defined in Subsection ??) and be consistent for any symmetric distribution and a skewed distribution with finite second moments. This is called an invariant mean. Based on the mean-symmetric weighted

Significance Statement

Bias, variance, and contamination are the three main errors in statistics. Consistent robust estimation is unattainable without parametric assumptions. Here, based on a paradigm shift inspired by mean-median-mode inequality, Bickel-Lehmann spread, and adaptive estimation, invariant moments are proposed as a means of achieving near-consistent and robust estimations of moments, even in scenarios where moderate violations of distributional assumptions occur, while the variances are sometimes smaller than those of the sample moments.

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

average-median inequality, the recombined mean is defined as

$$rm_{d,\epsilon,n} := \lim_{c \rightarrow \infty} \left(\frac{(\text{SWA}_{\epsilon,n} + c)^{d+1}}{(\text{median} + c)^d} - c \right),$$

where d is the key factor for bias correction, $\text{SWA}_{\epsilon,n}$ is $\text{BM}_{\epsilon,n}$ in the first three Subsections, but other symmetric weighted averages can also be used in practice as long as the inequalities hold. The following theorem shows the significance of this arithmetic I -statistic.

Theorem .1. *If the second moments are finite, $rm_{d \approx 0.375, \epsilon = \frac{1}{8}}$ is a consistent mean estimator for the exponential and any symmetric distributions and the Pareto distribution with quantile function $Q(p) = x_m(1-p)^{-\frac{1}{\alpha}}$, $x_m > 0$, when $\alpha \rightarrow \infty$.*

Proof. Finding d and ϵ that make $rm_{d,\epsilon}$ a consistent mean estimator is equivalent to finding the solution of $E[rm_{d,\epsilon}] = E[X]$. Rearranging the definition, $rm_{d,\epsilon} = \lim_{c \rightarrow \infty} \left(\frac{(\text{BM}_{\epsilon} + c)^{d+1}}{(\text{median} + c)^d} - c \right) = (d+1)\text{BM}_{\epsilon} - d\text{median} = \mu$. So, $d = \frac{\mu - \text{BM}_{\epsilon}}{\text{BM}_{\epsilon} - \text{median}}$. The quantile function of the exponential distribution is $Q(p) = \ln\left(\frac{1}{1-p}\right)\lambda$. $E[x] = \lambda$. $E[\text{median}] = Q\left(\frac{1}{2}\right) = \ln 2\lambda$. For the exponential distribution, the expectation of $\text{BM}_{\frac{1}{8}}$ is $E\left[\text{BM}_{\frac{1}{8}}\right] = \lambda \left(1 + \ln\left(\frac{46656}{8575\sqrt{35}}\right)\right)$. Obviously, the scale parameter λ can be canceled out, $d \approx 0.375$. The proof of the second assertion follows directly from the coincidence property. For any symmetric distribution with a finite second moment, $E[\text{BM}_{\epsilon}] = E[\text{median}] = E[X]$. Then $E[rm_{d,\epsilon}] = \lim_{c \rightarrow \infty} \left(\frac{(E[X] + c)^{d+1}}{(E[X] + c)^d} - c \right) = E[X]$. The proof for the Pareto distribution is more general. The mean of the Pareto distribution is given by $\frac{\alpha x_m}{\alpha - 1}$. The d value with two unknown percentiles p_1 and p_2 for the Pareto distribution is $d_{\text{Pareto}} = \frac{\frac{\alpha x_m}{\alpha - 1} - x_m(1-p_1)^{-\frac{1}{\alpha}}}{x_m(1-p_1)^{-\frac{1}{\alpha}} - x_m(1-p_2)^{-\frac{1}{\alpha}}}$. Since any weighted average can be expressed as an integral of the quantile function, $\lim_{\alpha \rightarrow \infty} \frac{\frac{\alpha}{\alpha-1} - (1-p_1)^{-1/\alpha}}{(1-p_1)^{-1/\alpha} - (1-p_2)^{-1/\alpha}} = -\frac{\ln(1-p_1)+1}{\ln(1-p_1)-\ln(1-p_2)}$, the d value for the Pareto distribution approaches that of the exponential distribution as $\alpha \rightarrow \infty$, regardless of the type of weighted average used. This completes the demonstration. \square

Theorem .1 implies that for the Weibull, gamma, Pareto, lognormal and generalized Gaussian distribution, $rm_{d \approx 0.375, \epsilon = \frac{1}{8}}$ is consistent for at least one particular case. The biases of $rm_{d \approx 0.375, \epsilon = \frac{1}{8}}$ for distributions with skewness between those of the exponential and symmetric distributions are tiny (SI Dataset S1). $rm_{d \approx 0.375, \epsilon = \frac{1}{8}}$ exhibits excellent performance for all these common unimodal distributions (SI Dataset S1).

Besides introducing the concept of invariant mean, the purpose of this paper is to demonstrate that, in light of previous works, the estimation of central moments can be transformed into a location estimation problem by using U -statistics, the central moment kernel distributions possess desirable properties, and a series of sophisticated yet efficient robust estimators can be constructed whose biases are typically smaller than the variances (as seen in Table ?? for $n = 5400$) for unimodal distributions.

Background and Main Results

A. Invariant mean. It has long been known that a theoretical model can be adjusted to fit the first two moments of the observed data. A continuous distribution belonging to a location-scale family takes the form $F(x) = F_0\left(\frac{x-\mu}{\lambda}\right)$, where F_0 is a "standard" distribution. Therefore, $F(x) = Q^{-1}(x) \rightarrow x = Q(p) = \lambda Q_0(p) + \mu$. Thus, any weighted average can be expressed as $\lambda \text{WA}_0(\epsilon) + \mu$, where $\text{WA}_0(\epsilon)$ is an integral of $Q_0(p)$ according to the definition of the weighted average. The simultaneous cancellation of μ and λ in $\frac{(\lambda\mu_0 + \mu) - (\lambda\text{BM}_0(\epsilon) + \mu)}{(\lambda\text{BM}_0(\epsilon) + \mu) - (\lambda\text{median}_0 + \mu)}$ assures that d is a constant. Consequently, the roles of BM_{ϵ} and median in $rm_{d,\epsilon}$ can be replaced by any weighted averages, although only symmetric weighted averages are considered in defining the invariant mean.

The performance in heavy-tailed distributions can be further improved by constructing the quantile mean as

$$qm_{d,\epsilon,n} := \hat{Q}_n \left(\left(\hat{F}_n(\text{SWA}_{\epsilon,n}) - \frac{1}{2} \right) d + \hat{F}_n(\text{SWA}_{\epsilon,n}) \right),$$

provided that $\hat{F}_n(\text{SWA}_{\epsilon,n}) \geq \frac{1}{2}$, where $\hat{F}_n(x)$ is the empirical cumulative distribution function of the sample, \hat{Q}_n is the sample quantile function. The most popular method for computing the sample quantile function was proposed by Hyndman and Fan in 1996 (14). To minimize the finite sample bias, here, $\hat{F}_n(x) := \frac{1}{n} \left(\frac{x - X_{sp}}{X_{sp+1} - X_{sp}} + sp \right)$, where $sp = \sum_{i=1}^n \mathbf{1}_{X_i \leq x}$, $\mathbf{1}_A$ is the indicator of event A . The solution of $\hat{F}_n(\text{SWA}_{\epsilon,n}) < \frac{1}{2}$ is reversing the percentile by $1 - \hat{F}_n(\text{SWA}_{\epsilon,n})$, the obtained percentile is also reversed. Without loss of generality, in the following discussion, only the case where $\hat{F}_n(\text{SWA}_{\epsilon,n}) \geq \frac{1}{2}$ is considered. Moreover, in extreme heavy-tailed distributions, the calculated percentile can exceed the breakdown point of SWA_{ϵ} , so the percentile will be modified to $1 - \epsilon$ if this occurs. The quantile mean uses the location-scale invariant in a different way as shown in the following proof.

Theorem A.1. *$qm_{d \approx 0.321, \epsilon = \frac{1}{8}}$ is a consistent mean estimator for the exponential, Pareto ($\alpha \rightarrow \infty$) and any symmetric distributions provided that the second moments are finite.*

Proof. Similarly, rearranging the definition, $d = \frac{F(\mu) - F(\text{BM}_{\epsilon})}{F(\text{BM}_{\epsilon}) - \frac{1}{2}}$. The cdf of the exponential distribution is $F(x) = 1 - e^{-\lambda^{-1}x}$, $\lambda \geq 0$, $x \geq 0$, the expectation of BM_{ϵ} can be expressed as $\lambda \text{BM}_0(\epsilon)$, so $F(\text{BM}_{\epsilon})$ is free of λ . When $\epsilon = \frac{1}{8}$, $d = \frac{-e^{-1} + e^{-\left(1 + \ln\left(\frac{46656}{8575\sqrt{35}}\right)\right)}}{\frac{1}{2} - e^{-\left(1 + \ln\left(\frac{46656}{8575\sqrt{35}}\right)\right)}} \approx 0.321$. The proof of the symmetric case is similar. Since for any symmetric distribution with a finite second moment, $F(E[\text{BM}_{\epsilon}]) = F(\mu) = \frac{1}{2}$. Then, the expectation of the quantile mean is $qm_{d,\epsilon} = F^{-1}\left(\left(F(\mu) - \frac{1}{2}\right)d + F(\mu)\right) = F^{-1}\left(0 + F(\mu)\right) = \mu$.

For the assertion related to the Pareto distribution, the cdf of it is $1 - \left(\frac{x_m}{x}\right)^{\alpha}$. So, the d value with two unknown percentile p_1 and p_2 is

$$d_{\text{Pareto}} = \frac{1 - \left(\frac{x_m}{\frac{\alpha x_m}{\alpha-1}}\right)^{\alpha} - \left(1 - \left(\frac{x_m}{x_m(1-p_1)^{-\frac{1}{\alpha}}}\right)^{\alpha}\right)}{\left(1 - \left(\frac{x_m}{x_m(1-p_1)^{-\frac{1}{\alpha}}}\right)^{\alpha}\right) - \left(1 - \left(\frac{x_m}{x_m(1-p_2)^{-\frac{1}{\alpha}}}\right)^{\alpha}\right)} = \frac{1 - \left(\frac{\alpha-1}{\alpha}\right)^{\alpha-p_1}}{p_1 - p_2}. \text{ When } \alpha \rightarrow \infty, \left(\frac{\alpha-1}{\alpha}\right)^{\alpha} = \frac{1}{e}. \text{ The } d \text{ value for the exponential distribution is identical, since } d_{\text{exp}} =$$

$$\frac{(1-e^{-1}) - \left(1-e^{-\ln\left(\frac{1}{1-p_1}\right)}\right)}{\left(1-e^{-\ln\left(\frac{1}{1-p_1}\right)}\right) - \left(1-e^{-\ln\left(\frac{1}{1-p_2}\right)}\right)} = \frac{1-\frac{1}{e}-p_1}{p_1-p_2}. \quad \text{All results are now proven.} \quad \square$$

The definitions of location and scale parameters are such that they must satisfy $F(x; \lambda, \mu) = F\left(\frac{x-\mu}{\lambda}; 1, 0\right)$. By recalling $x = \lambda Q_0(p) + \mu$, it follows that the percentile of any weighted average is free of λ and μ , which guarantees the validity of the quantile mean. The quantile mean is a quantile I -statistic. Specifically, an estimator is classified as a quantile I -statistic if LEs are percentiles of a distribution obtained by plugging L -statistics into a cumulative distribution function and I is defined with arithmetic operations, constants and quantile functions. $qm_{d \approx 0.321, \epsilon = \frac{1}{8}}$ works better in the fat-tail scenarios (SI Dataset S1). Theorem .1 and A.1 show that $rm_{d \approx 0.375, \epsilon = \frac{1}{8}}$ and $qm_{d \approx 0.321, \epsilon = \frac{1}{8}}$ are both consistent mean estimators for any symmetric distribution and a skewed distribution with finite second moments. It's obvious that the breakdown points of $rm_{d \approx 0.375, \epsilon = \frac{1}{8}}$ and $qm_{d \approx 0.321, \epsilon = \frac{1}{8}}$ are both $\frac{1}{8}$. Therefore they are all invariant means.

To study the impact of the choice of SWAs in rm and qm , it is constructive to recall that a symmetric weighted average is a linear combination of symmetric quantile averages. While using a less-biased symmetric weighted average can generally enhance performance (SI Dataset S1), there is a greater risk of violation in the semiparametric framework. However, the mean-SWA-median inequality is robust to slight fluctuations of the SQA function of the underlying distribution. Suppose the SQA function is generally decreasing in $[0, u]$, but increasing in $[u, \frac{1}{2}]$, since $1-2\epsilon$ of the symmetric quantile averages will be included in the computation of SWA_ϵ , as long as $\frac{1}{2} - u \ll 1-2\epsilon$, and other portions of the SQA function satisfy the inequality constraints that define the ν th orderliness on which the SWA_ϵ is based, the mean-SWA-median inequality will still hold. This is due to the violation being bounded (15) and therefore cannot be extreme for unimodal distributions. For instance, the SQA function is non-monotonic when the shape parameter of the Weibull distribution $\alpha > \frac{1}{1-\ln(2)} \approx 3.259$ as shown in the previous article, the violation of the third orderliness starts near this parameter as well, yet the mean-BM $_{\frac{1}{8}}$ -median inequality is still valid when $\alpha \leq 3.322$. Another key factor in determining the risk of violation is the skewness of the distribution. Previously, it was demonstrated that in a family of distributions differing by a skewness-increasing transformation in van Zwet's sense, the violation of orderliness, if it happens, often only occurs when the distribution is nearly symmetrical (16). The over-corrections in rm and qm are dependent on the SWA_ϵ -median difference, which can be a reasonable measure of skewness (17, 18), implying that the over-correction is often tiny with a moderate d . This qualitative analysis provides another perspective, in addition to the bias bounds (15), that rm and qm based on the mean-SWA-median inequality are generally safe.

B. Robust estimations of the central moments. In 1979, Bickel and Lehmann, in their final paper of the landmark series *Descriptive Statistics for Nonparametric Models* (19), generalized a class of estimators called "measures of spread," which "does not require the assumption of symmetry." From that, a popular

efficient scale estimator, the Rousseeuw-Croux scale estimator (20), was derived in 1993, but the importance of tackling the symmetry assumption has been greatly underestimated. While they had already considered one version of the trimmed standard deviation in the third paper of that series (21), in the final section of that paper (19), they explored another two possible versions, which were modified here for comparison,

$$\left[n\left(\frac{1}{2} - \epsilon\right)\right]^{-\frac{1}{2}} \left[\sum_{i=\frac{n}{2}}^{n(1-\epsilon)} [X_i - X_{n-i+1}]^2\right]^{\frac{1}{2}}, \quad [1]$$

and

$$\left[\binom{n}{2} (1 - \epsilon - \gamma\epsilon)\right]^{-\frac{1}{2}} \left[\sum_{i=\binom{n}{2}\gamma\epsilon}^{\binom{n}{2}(1-\epsilon)} (X - X')_i^2\right]^{\frac{1}{2}}, \quad [2]$$

where $(X - X')_1 \leq \dots \leq (X - X')_{\binom{n}{2}}$ are the order statistics of the "pseudo-sample", $X_i - X_j$, $i < j$. The paper ended with, "We do not know a fortiori which of the measures [1] or [2] is preferable and leave these interesting questions open."

Observe that the kernel of the unbiased estimation of the second central moment by using U -statistic is $\psi_2(x_1, x_2) = \frac{1}{2}(x_1 - x_2)^2$. If adding the $\frac{1}{2}$ term in [2], as $\epsilon \rightarrow 0$, the result is equivalent to the standard deviation estimated by using U -statistic (also noted by Janssen, Serfling, and Veraverbeke in 1987) (22). In fact, they also showed that, when ϵ is 0, [2] is $\sqrt{2}$ times the standard deviation.

To address their open question, the nomenclature used in this paper is introduced as follows:

Nomenclature. Given a robust estimator $\hat{\theta}$ with an adjustable breakdown point which can be infinitesimal. The name of $\hat{\theta}$ is composed of two parts: the first part denotes the type of estimator, and the second part is the name of the population parameter θ that the estimator is consistent with as $\epsilon \rightarrow 0$. The abbreviation of the estimator is formed by combining the initial letter(s) of the first part with the common abbreviation of the consistent estimator that measures the population parameter. If the estimator is symmetric, ϵ is indicated in the subscript of the abbreviation of the estimator, except the median. For asymmetric estimators based on quantile average, the corresponding γ is also indicated after ϵ . Note that ϵ is the right breakdown point (defined in Subsection ??), while the left breakdown point should be further calculated.

In the previous article on semiparametric robust mean estimation, it was shown that the bias of a robust estimator with an adjustable breakdown point is often monotonic with respect to the breakdown point in a semiparametric distribution. Naturally, the estimator's name should correspond to the population parameter with which it is consistent as $\epsilon \rightarrow 0$. The trimmed standard deviation following this nomenclature is $Tsd_{\epsilon=1-\sqrt{1-\epsilon_0}, \gamma, n} := \left[TM_{\epsilon_0, \gamma} \left((\psi_2(X_{N_1}, X_{N_2}))_{N=1}^{\binom{n}{2}}\right)\right]^{-\frac{1}{2}}$, where $TM_{\epsilon_0, \gamma}(Y)$ denotes the ϵ_0, γ -trimmed mean with the sequence $(\psi_2(X_{N_1}, X_{N_2}))_{N=1}^{\binom{n}{2}}$ as an input, the proof of the breakdown point is given in Subsection ??. Removing the square root yields the trimmed variance ($Tvar_{\epsilon, \gamma, n}$). It is now very clear that this definition, essentially the same as [2], should be preferable. Not only because it is essentially a

trimmed U -statistic for the standard deviation but also because the γ -orderliness of the second central moment kernel distribution is ensured by the next exciting theorem.

Theorem B.1. *The second central moment kernel distribution generated from any unimodal distribution is γ -ordered.*

Proof. The monotonic increasing of the pairwise difference distribution was first implied in its unimodality proof done by Hodges and Lehmann in 1954 (23). Whereas they used induction to get the result, Dharmadhikari and Jogdeo in 1982 (24) provided a modern proof of the unimodality using Khintchine's representation (25). Assuming absolute continuity, Purkayastha (26) introduced a much simpler proof in 1998. Transforming the pairwise difference distribution by squaring and multiplying by $\frac{1}{2}$ does not change the monotonicity, making the pdf become monotonically decreasing with mode at zero. In the previous article, it was proven that a right skewed distribution with a monotonic decreasing pdf is always γ -ordered, which gives the desired result. \square

Previously, it was shown that any symmetric distribution with a finite second moment is ν th ordered, indicating that orderliness does not require unimodality, e.g., a symmetric bimodal distribution is also ordered. An analysis of the Weibull distribution showed that unimodality does not guarantee orderliness. Theorem B.1 reveals another profound relationship between unimodality and orderliness, which is sufficient for trimming inequality.

In 1928, Fisher constructed k -statistics as unbiased estimators of cumulants (27). Halmos (1946) proved that the functional θ admits an unbiased estimator if and only if it is a regular statistical functional of degree k and showed a relation of symmetry, unbiasedness and minimum variance (28). In 1948, Hoeffding generalized U -statistics (29) which enable the derivation of a minimum-variance unbiased estimator from each unbiased estimator of an estimable parameter. Heffernan (1997) (30) obtained an unbiased estimator of the k th central moment by using U -statistics and demonstrated that it is the minimum variance unbiased estimator for distributions with finite moments (31, 32). In 1976, Saleh generalized the Hodges-Lehmann (H-L) estimator (33) to the trimmed H-L mean (which he named "Wilcoxon one-sample statistic") (34). In 1984, Serfling pointed out the speciality of Hodges-Lehmann estimator, which is neither a simple L -statistic nor a U -statistic, and considered the generalized L -statistics and U -statistic structure (35). Also in 1984, Janssen and Serfling and Veraverbeke (36) showed that the Bickel-Lehmann spread also belongs to the same class. It gradually became clear that the Hodges-Lehmann estimator, trimmed H-L mean and trimmed standard deviation are all trimmed U -statistics (37-39).

Extending the trimmed U -statistic to weighted U -statistic, i.e., replacing the trimmed mean with weighted average. The weighted k th central moment ($k \leq n$) is defined as,

$$Wkm_{\epsilon=1-(1-\epsilon_0)\frac{1}{k}, \gamma, n} := WA_{\epsilon_0, \gamma, n} \left((\psi_k(X_{N_1}, \dots, X_{N_k}))_{N=1}^{(n)} \right),$$

where X_{N_1}, \dots, X_{N_k} are the n choose k elements from X , $\psi_k(x_1, \dots, x_k) = \sum_{j=0}^{k-2} (-1)^j \binom{1}{k-j} \sum (x_{i_1}^{k-j} \dots x_{i_{j+1}}) + (-1)^{k-1} (k-1) x_1 \dots x_k$, the second summation is over $i_1, \dots, i_{j+1} = 1$ to k with $i_1 < \dots < i_{j+1}$ (30). Despite the complexity, the structure of the k th central moment kernel distributions can be elucidated by decomposing.

Theorem B.2. *For each pair $(Q(p_i), Q(p_j))$ of the original distribution such that $Q(p_i) < Q(p_j)$, let $x_1 = Q(p_i)$ and $x_k = Q(p_j)$, $\Delta = Q(p_i) - Q(p_j)$, the k th central moment kernel distribution, $k > 2$, can be seen as a mixture distribution and each of the components has the support $(-\frac{k}{3+(-1)^k})^{-1}(-\Delta)^k, \frac{1}{k}(-\Delta)^k$.*

Proof. Without loss of generality, generating the distribution of the k -tuple $(Q(p_{i_1}), \dots, Q(p_{i_k}))$ under continuity, $k > 2$, $i_1 < \dots < i_k$, $p_{i_1} < \dots < p_{i_k}$, the corresponding probability density is $f_{X, \dots, X}(Q(p_{i_1}), \dots, Q(p_{i_k})) = k! f(Q(p_{i_1})) \dots f(Q(p_{i_k}))$. Transforming the distribution of the k -tuple by the function $\psi_k(x_1, \dots, x_k)$, denoting $\bar{\Delta} = \psi_k(Q(p_{i_1}), \dots, Q(p_{i_k}))$. The probability $f_{\bar{\Delta}}(\bar{\Delta}) = \sum_{\bar{\Delta} = \psi_k(Q(p_{i_1}), \dots, Q(p_{i_k}))} f_{X, \dots, X}(Q(p_{i_1}), \dots, Q(p_{i_k}))$ is the summation of the probabilities of all k -tuples such that $\bar{\Delta}$ is equal to $\psi_k(Q(p_{i_1}), \dots, Q(p_{i_k}))$. The following Ξ_k is equivalent.

Ξ_k : Every pair with a difference equal to $\Delta = Q(p_{i_1}) - Q(p_{i_k})$ can generate a pseudodistribution (but the integral is not equal to 1, so "pseudo") such that x_2, \dots, x_{k-1} exhaust all combinations under the inequality constraints, i.e., $Q(p_{i_1}) = x_1 < x_2 < \dots < x_{k-1} < x_k = Q(p_{i_k})$. The combination of all the pseudodistributions with the same Δ is ξ_Δ . The combination of ξ_Δ , i.e., from $\Delta = 0$ to $Q(0) - Q(1)$, is Ξ_k .

The support of ξ_Δ is the extrema of ψ_k subject to the inequality constraints. Using the Lagrange multiplier, one can easily determine the only critical point at $x_1 = \dots = x_k = 0$, where $\psi_k = 0$. Other candidates are within the boundaries, i.e., $\psi_k(x_1 = x_1, x_2 = x_k, \dots, x_k = x_k)$, \dots , $\psi_k(x_1 = x_1, \dots, x_i = x_1, x_{i+1} = x_k, \dots, x_k = x_k)$, \dots , $\psi_k(x_1 = x_1, \dots, x_{k-1} = x_1, x_k = x_k)$. $\psi_k(x_1 = x_1, \dots, x_i = x_1, x_{i+1} = x_k, \dots, x_k = x_k)$ can be divided into k groups. If $\frac{k+1-i}{2} \leq j \leq \frac{k-1}{2}$, from $j+1$ st to $k-j$ th group, the g th group has $i \binom{i-1}{g-j-1} \binom{k-i}{j}$ terms having the form $(-1)^{g+1} \frac{1}{k-g+1} x_1^{k-j} x_k^j$, from $k-j+1$ th to $i+j$ th group, the g th group has $i \binom{i-1}{g-j-1} \binom{k-i}{j} + (k-i) \binom{k-i-1}{j-k+g-1} \binom{i}{k-j}$ terms having the form $(-1)^{g+1} \frac{1}{k-g+1} x_1^{k-j} x_k^j$. If $j < \frac{k+1-i}{2}$, from $j+1$ st to $i+j$ th group, the g th group has $i \binom{i-1}{g-j-1} \binom{k-i}{j}$ terms having the form $(-1)^{g+1} \frac{1}{k-g+1} x_1^{k-j} x_k^j$. If $j \geq \frac{k}{2}$, from $k-j+1$ st to j th group, the g th group has $(k-i) \binom{k-i-1}{j-k+g-1} \binom{i}{k-j}$ terms having the form $(-1)^{g+1} \frac{1}{k-g+1} x_1^{k-j} x_k^j$, from $j+1$ th to $j+i$ th group, $i+j < k$, the g th group has $i \binom{i-1}{g-j-1} \binom{k-i}{j} + (k-i) \binom{k-i-1}{j-k+g-1} \binom{i}{k-j}$ terms having the form $(-1)^{g+1} \frac{1}{k-g+1} x_1^{k-j} x_k^j$. The final k th group is the term $(-1)^{k-1} (k-1) x_1^i x_k^{k-i}$. So, if $i+j = k$, $j \geq \frac{k}{2}$, $i \leq \frac{k}{2}$, the summed coefficient of $x_1^i x_k^{k-i}$ is $(-1)^{k-1} (k-1) + \sum_{g=i+1}^{k-1} (-1)^{g+1} \frac{1}{k-g+1} (k-i) \binom{k-i-1}{g-i-1} + \sum_{g=k-i+1}^{k-1} (-1)^{g+1} \frac{1}{k-g+1} i \binom{i-1}{g-k+i-1} = (-1)^{k-1} (k-1) + (-1)^{k+1} + (k-i)(-1)^k + (-1)^k (i-1) = (-1)^{k+1}$. The summation identities are $\sum_{g=i+1}^{k-1} (-1)^{g+1} \frac{1}{k-g+1} (k-i) \binom{k-i-1}{g-i-1} = (k-i) \int_0^1 \sum_{g=i+1}^{k-1} (-1)^{g+1} \binom{k-i-1}{g-i-1} t^{k-g} dt = (k-i) \int_0^1 ((-1)^i (t-1)^{k-i-1} - (-1)^{k+1}) dt = (k-i) \left(\frac{(-1)^k}{i-k} + (-1)^k \right) = (-1)^{k+1} + (k-i)(-1)^k$ and $\sum_{g=k-i+1}^{k-1} (-1)^{g+1} \frac{1}{k-g+1} i \binom{i-1}{g-k+i-1} =$

371 $\int_0^1 \sum_{g=k-i+1}^{k-1} (-1)^{g+1} i \binom{i-1}{g-k+i-1} t^{k-g} dt =$
372 $\int_0^1 (i(-1)^{k-i} (t-1)^{i-1} - i(-1)^{k+1}) dt = (-1)^k (i-1).$
373 If $j < \frac{k+1-i}{2}$, $i > k-1$, if $i = k$, $\psi_k = 0$, if $\frac{k+1-i}{2} \leq j \leq \frac{k-1}{2}$,
374 $\frac{k+1}{2} \leq i \leq k-1$, the summed coefficient of $x_1^i x_k^{k-i}$ is
375 $(-1)^{k-1} (k-1) + \sum_{g=k-i+1}^{k-1} (-1)^{g+1} \frac{1}{k-g+1} i \binom{i-1}{g-k+i-1} +$
376 $\sum_{g=i+1}^{k-1} (-1)^{g+1} \frac{1}{k-g+1} (k-i) \binom{k-i-1}{g-i-1}$, the same as above. If
377 $i+j < k$, since $\binom{i}{k-j} = 0$, the related terms can be ignored, so,
378 using the binomial theorem and beta function, the summed co-
379 efficient of $x_1^{k-j} x_k^j$ is $\sum_{g=j+1}^{i+j} (-1)^{g+1} \frac{1}{k-g+1} i \binom{i-1}{g-j-1} \binom{k-i}{j} =$
380 $i \binom{k-i}{j} \int_0^1 \sum_{g=j+1}^{i+j} (-1)^{g+1} \binom{i-1}{g-j-1} t^{k-g} dt =$
381 $\binom{k-i}{j} i \int_0^1 \left((-1)^j t^{k-j-1} \left(\frac{t}{1-t} \right)^{i-1} \right) dt =$
382 $\binom{k-i}{j} i \frac{(-1)^{j+i+1} \Gamma(i) \Gamma(k-j-i+1)}{\Gamma(k-j+1)} = \frac{(-1)^{j+i+1} i! (k-j-i)! (k-i)!}{(k-j)! j! (k-j-i)!} =$
383 $(-1)^{j+i+1} \frac{i! (k-i)!}{k!} \frac{k!}{(k-j)! j!} = \binom{k}{i}^{-1} (-1)^{1+i} \binom{k}{j} (-1)^j.$
384 The coefficient of $x_1^i x_k^{k-i}$ in $\binom{k}{i}^{-1} (-1)^{1+i} (x_1 - x_k)^k$
385 is $\binom{k}{i}^{-1} (-1)^{1+i} \binom{k}{i} (-1)^{k-i} = (-1)^{k+1}$, same as the
386 summed coefficient if $i+j = k$. If $i+j <$
387 k , the coefficient of $x_1^{k-j} x_k^j$ is $\binom{k}{i}^{-1} (-1)^{1+i} \binom{k}{j} (-1)^j$,
388 same as the corresponding summed coefficient. There-
389 fore, $\psi_k(x_1 = x_1, \dots, x_i = x_1, x_{i+1} = x_k, \dots, x_k = x_k) =$
390 $\binom{k}{i}^{-1} (-1)^{1+i} (x_1 - x_k)^k$, the maximum and minimum of ψ_k
391 follow directly from the properties of the binomial coeffi-
392 cient. \square

393 ξ_Δ is closely related to $f_\Xi(\Delta)$, which is the pairwise differ-
394 ence distribution, since the probability density of ξ_Δ can be ex-
395 pressed as $f_{\Xi_k}(\bar{\Delta}|\Delta)$ and $\sum_{\bar{\Delta}=-\left(\frac{k}{3+(-1)^k}\right)^{-1}(-\Delta)^k}^{\frac{1}{k}(-\Delta)^k} f_{\Xi_k}(\bar{\Delta}|\Delta) =$
396 $f_\Xi(\Delta)$. Recall that $f_\Xi(\Delta)$ is monotonic increasing with a mode
397 at the origin if the original distribution is unimodal. Thus, in
398 general, ignoring the shape of ξ_Δ , Ξ_k is monotonic left and
399 right around zero. In fact, the median of Ξ_k also exhibits
400 a strong tendency to be close to zero, as it can be cast as
401 a weighted mean of the medians of ξ_Δ . When Δ is small,
402 all values of ξ_Δ are close to zero, resulting in the median of
403 ξ_Δ being close to zero as well. When Δ is large, the me-
404 dian of ξ_Δ depends on its skewness, but the corresponding
405 weight is much smaller, so even if ξ_Δ is highly skewed, the
406 median of Ξ_k will only be slightly shifted from zero. Denote
407 the median of Ξ_k as m_{Ξ_k} , for the five parametric distribu-
408 tions here, $|m_{\Xi_k}|$ s are all $\leq 0.1\sigma$ for Ξ_3 and Ξ_4 (SI Dataset
409 S1). Assuming $m_{\Xi_k} = 0$, for the even ordinal central moment
410 kernel distribution, the average probability density on the
411 left side of zero is greater than that on the right side, since
412 $\frac{\frac{1}{2}}{\binom{k}{2}^{-1} (Q(0)-Q(1))^k} > \frac{\frac{1}{2}}{\frac{1}{k} (Q(0)-Q(1))^k}$. This means that, on aver-
413 age, the inequality $f(Q(\epsilon)) \geq f(Q(1-\epsilon))$ holds. For the odd
414 ordinal distribution, the discussion is more challenging since
415 it is generally symmetric. Just consider Ξ_3 , let $x_1 = Q(p_i)$
416 and $x_3 = Q(p_j)$, changing the value of x_2 from $Q(p_i)$
417 to $Q(p_j)$ will monotonically change the value of $\psi_3(x_1, x_2, x_3)$,
418 since $\frac{\partial \psi_3(x_1, x_2, x_3)}{\partial x_2} = -\frac{x_1^2}{2} - x_1 x_2 + 2x_1 x_3 + x_2^2 - x_2 x_3 - \frac{x_3^2}{2},$
419 $-\frac{3}{4}(x_1 - x_3)^2 \leq \frac{\partial \psi_3(x_1, x_2, x_3)}{\partial x_2} \leq -\frac{1}{2}(x_1 - x_3)^2 \leq 0$. If the
420 original distribution is right-skewed, ξ_Δ will be left-skewed,
421 so, for Ξ_3 , the average probability density of the right side of
422 zero will be greater than that of the left side, which means,
423 on average, the inequality $f(Q(\epsilon)) \leq f(Q(1-\epsilon))$ holds (the
424 same result can be inferred from the definition of central mo-

ments, where the positivity of the odd order central moment
is directly related to the left-skewness of the corresponding
kernel distribution). In all, the monotonicity of the pairwise
difference distribution guides the general shape of the k th
central moment kernel distribution, $k > 2$, forcing it to be
unimodal-like with the mode and median close to zero, then,
the inequality $f(Q(\epsilon)) \leq f(Q(1-\epsilon))$ or $f(Q(\epsilon)) \geq f(Q(1-\epsilon))$
holds in general. If a distribution is ordered and all of its cen-
tral moment kernel distributions are also ordered, it is called
completely ordered. Although strict complete orderliness is
difficult to prove, even if the inequality may be violated in
a small range, as discussed in Subsection A, the mean-SWA-
median inequality remains valid, in most cases, for the central
moment kernel distribution.

Another crucial property of the central moment kernel dis-
tribution, location invariant, is introduced in the next theorem.
The proof is provided in the SI Text.

Theorem B.3. $\psi_k(x_1 = \lambda x_1 + \mu, \dots, x_k = \lambda x_k + \mu) =$
 $\lambda^k \psi_k(x_1, \dots, x_k).$

Consider two continuous distributions belonging to the
same location-scale family, their corresponding k th central
moment kernel distributions only differ in scaling. So d is
invariant, as shown in Subsection A. The recombined k th
central moment, based on rm , is defined by,

$$rkm_{d, \epsilon=1-(1-\epsilon_0)^{\frac{1}{k}}, n} := (d+1) SWkm_{\epsilon, n} - d mkm_n,$$

where $SWkm_{\epsilon, n}$ is using the binomial k th central moment
($Bkm_{\epsilon_0, n}$) here, mkm_n is the median k th central moment.
Since $SWkm_{\epsilon, n}$ is an L -statistic, the resulting $rkm_{d, \epsilon, n}$ is an
arithmetic I -statistic. Similarly, the quantile will not change
after scaling. The quantile k th central moment is thus defined as

$$qkm_{d, \epsilon, n} := \hat{Q}_n \left(\left(pSWkm_{\epsilon, n} - \frac{1}{2} \right) d + pSWkm_{\epsilon, n} \right),$$

where $pSWkm_{\epsilon, n} = \hat{F}_{\psi, n}(SWkm_{\epsilon, n})$, $\hat{F}_{\psi, n}$ is the empirical
cumulative distribution function of the corresponding central
moment kernel distribution. $qkm_{d, \epsilon, n}$ is a quantile I -statistic.

Finally, for standardized moments, quantile skewness and
quantile kurtosis are defined to be $qskew_{d, \epsilon, n} := \frac{qtm_{d, \epsilon, n}}{qsd_{d, \epsilon, n}^3}$

and $qkurt_{d, \epsilon, n} := \frac{qfm_{d, \epsilon, n}}{qsd_{d, \epsilon, n}^4}$. Quantile standard deviation
($qsd_{d, \epsilon, n}$), recombined standard deviation ($rsd_{d, \epsilon, n}$), quantile
third central moment ($qtm_{d, \epsilon, n}$), quantile fourth central mo-
ment ($qfm_{d, \epsilon, n}$), recombined third central moment ($rtm_{d, \epsilon, n}$),
recombined fourth central moment ($rfm_{d, \epsilon, n}$), recombined
skewness ($rskew_{d, \epsilon, n}$), and recombined kurtosis ($rkurt_{d, \epsilon, n}$)
are all defined similarly as above and not repeated here. The
transformation to a location problem can also empower related
statistical tests. From the better performance of the quantile
mean in heavy-tailed distributions, quantile central moments
are generally better than recombined central moments regard-
ing asymptotic bias.

To avoid confusion, it should be noted that the robust
location estimations of the kernel distributions discussed in
this paper differ from the approach taken by Joly and Lugosi
(2016) (40), which is computing the median of all U -statistics
from different disjoint blocks. Compared to bootstrap median
 U -statistics, this approach can produce two additional kinds
of finite sample bias, one arises from the limited numbers of

blocks, another is due to the size of the U -statistics (consider the mean of all U -statistics from different disjoint blocks, it is definitely not identical to the original U -statistic, except when the kernel is the Hodges-Lehmann kernel). Laforgue, Clemencon, and Bertail (2019)'s median of randomized U -statistics (41) is more sophisticated and can overcome the limitation of the number of blocks, but the second kind of bias remains unsolved.

C. Congruent distribution. In the realm of nonparametric statistics, the precise values of robust estimators are of secondary importance. What is of primary importance is their relative differences or orders. Based on this principle, in the absence of contamination, as the parameters of the distribution vary, all reasonable nonparametric location estimates should asymptotically change in the same direction. Otherwise if the results obtained based on the trimmed mean are completely different from those based on the median, a contradiction arises. However, such contradictions are possible, as in the case of the Weibull distribution, $E[m] = \lambda \sqrt[\gamma]{\ln(2)}$, $E[\mu] = \lambda \Gamma(1 + \frac{1}{\alpha})$, then, when $\alpha = 1$, $E[m] = \lambda \ln(2) \approx 0.693\lambda$, $E[\mu] = \lambda$, but when $\alpha = \frac{1}{2}$, $E[m] = \lambda \ln^2(2) \approx 0.480\lambda$, $E[\mu] = 2\lambda$, the mean increases, but the median decreases. To study the conditions that avoid such scenarios by classifying distributions through the signs of derivatives, let the quantile average function of a parametric distribution be denoted as $QA(\epsilon, \gamma, \alpha_1, \dots, \alpha_i, \dots, \alpha_k)$, where α_i represent the parameters of the distribution, then, a distribution is γ -congruent if and only if the sign of $\frac{\partial QA}{\partial \alpha_i}$ remains the same for all $0 \leq \epsilon \leq \frac{1}{1+\gamma}$. If this partial derivative is equal to zero or undefined, it can be considered both positive and negative, and thus does not impact the analysis. Asymptotically, any weighted average can be expressed as an integral of the quantile average function. Since the sign does not change after integration, the sign of $\frac{\partial QA}{\partial \alpha_i}$ remains the same for all $0 \leq \epsilon \leq \frac{1}{1+\gamma}$ implies that all γ -weighted averages change in the same direction as the parameters change, as long as they are not undefined. A distribution is completely γ -congruent if and only if it is γ -congruent and all its central moment kernel distributions are also γ -congruent. Setting $\gamma = 1$ constitutes the definitions of congruence and complete congruence. Chebyshev's inequality implies that, for any probability distribution with finite moments, even if some weighted averages change in a direction different from that of the sample mean, the deviations are bounded. Furthermore, distributions with infinite moments can be γ -congruent, since the definition is based on the quantile average, not the sample mean.

The following theorems show the conditions that a distribution is congruent or γ -congruent.

Theorem C.1. *A symmetric distribution with a finite second moment is always congruent.*

Proof. For any symmetric distribution with a finite second moment, all symmetric quantile averages coincide. The conclusion follows immediately. \square

Theorem C.2. *A positive define location-scale distribution with a finite second moment is always γ -congruent.*

Proof. As shown in discussions in Subsection A, for a location-scale distribution, any weighted average can be expressed as $\lambda WA_0(\epsilon) + \mu$, where $WA_0(\epsilon)$ is an integral of $Q_0(p)$ according

to the definition of the weighted average. Therefore, the derivatives with respect to the parameters λ or μ are always positive. By application of the definition, the desired outcome is obtained. \square

Theorem C.3. *The second central moment kernel distribution derived from a continuous location-scale unimodal distribution with a finite second moment is always γ -congruent.*

Proof. Theorem B.3 shows that the corresponding central moment kernel distribution is also a location-scale family distribution. Theorem B.1 shows that it is positively defined. Implementing Theorem C.2 yields the desired result. \square

For the Pareto distribution, $\frac{\partial Q(p, \alpha)}{\partial \alpha} = \frac{x_m(1-p)^{-1/\alpha} \ln(1-p)}{\alpha^2}$. Since $\ln(1-p) < 0$ for all $0 < p < 1$, $(1-p)^{-1/\alpha} > 0$ for all $0 < p < 1$ and $\alpha > 0$, so $\frac{\partial Q(p, \alpha)}{\partial \alpha} < 0$, and therefore $\frac{\partial QA(\epsilon, \gamma, \alpha)}{\partial \alpha} < 0$, the Pareto distribution is γ -congruent. The derivative for the lognormal distribution is $\frac{\partial SQA(\epsilon, \sigma)}{\partial \sigma} = \frac{-\operatorname{erfc}^{-1}(2\epsilon)e^{\mu - \sqrt{2}\sigma \operatorname{erfc}^{-1}(2\epsilon)} - \operatorname{erfc}^{-1}(2-2\epsilon)e^{\mu - \sqrt{2}\sigma \operatorname{erfc}^{-1}(2-2\epsilon)}}{\sqrt{2}}$. Since the inverse complementary error function is positive when the input is smaller than 1, and negative when the input is larger than 1, $\operatorname{erfc}^{-1}(2\epsilon) = -\operatorname{erfc}^{-1}(2-2\epsilon)$, $e^{\mu - \sqrt{2}\sigma \operatorname{erfc}^{-1}(2-2\epsilon)} > e^{\mu - \sqrt{2}\sigma \operatorname{erfc}^{-1}(2\epsilon)}$, $\frac{\partial SQA(\epsilon, \sigma)}{\partial \sigma} > 0$, the lognormal distribution is congruent. Theorem C.1 implies that the generalized Gaussian distribution is congruent. For the Weibull distribution, when α changes from 1 to $\frac{1}{2}$, the average probability density on the left side of the median increases, since $\frac{1}{\lambda \ln(2)} < \frac{1}{\lambda \ln^2(2)}$, but the mean increases, indicating that the distribution is more heavy-tailed, the probability density of large values will also increase. The main reason for non-congruence of a right-skewed smooth partial bounded probability distribution lies in the simultaneous increase of probability densities on two opposite sides: one approaching the bound and the other approaching infinity. Note that the gamma distribution does not have this issue, it looks to be congruent.

Although some common parametric distributions are not congruent, Theorem C.2 establishes that γ -congruence always holds for a positive define location-scale family distribution and thus for the second central moment kernel distribution generated from a continuous location-scale unimodal distribution as shown in Theorem C.3. Theorem B.2 demonstrates that all their central moment kernel distributions are unimodal-like with mode and median close to zero, as long as they are unimodal distributions. Assuming finite moments and constant $Q(0) - Q(1)$, increasing the mean of the kernel distribution will result in a more heavy-tailed distribution, i.e., the probability density of the values close to $\frac{1}{k}(-\Delta)^k$ increases. While the total probability density on either side of zero remains unchanged as the median is generally close to zero and much less impacted by increasing the mean, the probability density of the values close to zero decreases. This transformation will increase nearly all symmetric weighted averages, in the general sense. Therefore, except for the median, which is assumed to be zero, nearly all symmetric weighted averages for all central moment kernel distributions derived from unimodal distributions should change in the same direction when the parameters change.

D. A shape-scale distribution as the consistent distribution.

Up to this point, in this article, the consistent robust estimation has been limited to a location-scale distribution, with the location parameter often being omitted for simplicity. To construct probability distributions can be made to fit the observed skewness and kurtosis arbitrarily well, in 1894, Pearson (42) introduced a family of continuous probability distributions that are now often characterized by the square of the skewness and the kurtosis. If the skewness and the kurtosis are interrelated by a shape parameter, a distribution specified by a shape parameter (denoted as α) and a scale parameter (denoted as λ) is often referred to as a shape-scale distribution. Weibull, gamma, Pareto, lognormal, and generalized Gaussian distributions (when μ is a constant) are all shape-scale unimodal distributions. Moreover, if α or skewness or kurtosis is a constant, the shape-scale distribution is reduced to a location-scale distribution. The above discussion shows that, due to the invariant property, if a location-scale distribution is chosen as the consistent distribution, the type of invariant moments and their related weighted moments are given, there should exist a unique k -tuple (d_{im}, \dots, d_{ikm}) calibrated by the distribution and the corresponding kernel distributions generated from this distribution. For a right skewed shape-scale distribution, let $D(|skewness|, kurtosis, k, etype, dtype, n) = d_{ikm}$ denote these relations, where the first input is the absolute value of the skewness, the second input is the kurtosis, the third is the order of the central moment (if $k = 1$, the mean), the fourth is the type of estimator, the fifth is the type of consistent distribution, and the sixth input is the sample size. For simplicity, the last three inputs will be omitted in the following discussion. Hold in awareness that since scale is invariant, specifying d values for a shape-scale distribution only requires either skewness or kurtosis, while the other may be also omitted. Since many common shape-scale distributions are always right skewed (if not, only the right skewed or left skewed part is used for calibration, while the other part is omitted), the absolute value of the skewness should be identical to the skewness for them and it can also handle the left skew scenario well.

For recombined moments, the object of using a shape-scale distribution as the consistent distribution is to find solutions for the system of equations

$$\begin{cases} rm(SWA, median, D(|rskew|, rkurt, 1)) = \mu \\ rvar(SWvar, mvar, D(|rskew|, rkurt, 2)) = \mu_2 \\ rtm(SWtm, mtm, D(|rskew|, rkurt, 3)) = \mu_3 \\ rfm(SWfm, mfm, D(|rskew|, rkurt, 4)) = \mu_4 \\ rskew = \frac{\mu_3}{\mu_2} \\ rkurt = \frac{\mu_4}{\mu_2^2} \end{cases},$$

where μ_2 , μ_3 and μ_4 are the population second, third and fourth central moments. $|rskew|$ and $rkurt$ should be the invariant points of the functions $\varsigma(|rskew|) = \frac{rtm(SWtm, mtm, D(|rskew|, 3))}{rvar(SWvar, mvar, D(|rskew|, 2))^{\frac{3}{2}}}$ and $\varkappa(rkurt) = \frac{rfm(SWfm, mfm, D(rkurt, 4))}{rvar(SWvar, mvar, D(rkurt, 2))^2}$. Clearly, this is an overdetermined nonlinear system of equations, given that the skewness and kurtosis are interrelated for a shape-scale distribution. Since an overdetermined system constructed with random coefficients is almost always inconsistent, it is natural to optimize them separately using the fixed-point iteration (see Algorithm 1, only $rkurt$ is provided, others are

the same).

Algorithm 1 $rkurt$ for a shape-scale distribution

Input: D ; $SWvar$; $SWfm$; $mvar$; mfm ; $maxit$; δ

Output: $rkurt_{i-1}$

```

1:  $i = 0$ 
2:  $rkurt_i \leftarrow \varkappa(kurtosis_{max})$   $\triangleright$  Using the maximum kurtosis
   available in  $D$  as an initial guess.
3: repeat
4:    $i = i + 1$ 
    $rkurt_{i-1} \leftarrow rkurt_i$ 
5:    $rkurt_i \leftarrow \varkappa(rkurt_{i-1})$ 
6: until  $i > maxit$  or  $|rkurt_i - rkurt_{i-1}| < \delta$   $\triangleright maxit$  is
   the maximum number of iterations,  $\delta$  is a small positive
   number.
```

The following theorem shows the validity of Algorithm 1.

Theorem D.1. $|rskew|$ and $rkurt$, defined as the largest attracting fix points of the functions $\varsigma(|rskew|)$ and $\varkappa(rkurt)$, are consistent estimators of $\tilde{\mu}_3$ and $\tilde{\mu}_4$ for a shape-scale distribution whose central moment kernel distributions are all congruent, as long as they are within the domain of D , where $\tilde{\mu}_3$ and $\tilde{\mu}_4$ are the population skewness and kurtosis.

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

ACKNOWLEDGMENTS. I gratefully acknowledge the constructive comments made by the editor which substantially improved the clarity and quality of this paper.

1. CF Gauss, *Theoria combinationis observationum erroribus minimis obnoxiae*. (Henricus Dieterich), (1823).
2. S Newcomb, A generalized theory of the combination of observations so as to obtain the best result. *Am. Journal Math.* **8**, 343–366 (1886).
3. S Newcomb, Researches on the motion of the moon. part ii, the mean motion of the moon and other astronomical elements derived from observations of eclipses and occultations extending from the period of the babylonians until ad 1908. *United States. Naut. Alm. Off. Astron. paper*; v. **99**, 1 (1912).
4. PJ Huber, Robust estimation of a location parameter. *Ann. Math. Stat.* **35**, 73–101 (1964).
5. X He, WK Fung, Method of medians for lifetime data with weibull models. *Stat. medicine* **18**, 1993–2009 (1999).
6. M Menon, Estimation of the shape and scale parameters of the weibull distribution. *Technometrics* **5**, 175–182 (1963).
7. SD Dubey, Some percentile estimators for weibull parameters. *Technometrics* **9**, 119–129 (1967).
8. KM Hassanein, Percentile estimators for the parameters of the weibull distribution. *Biometrika* **58**, 673–676 (1971).
9. NB Marks, Estimation of weibull parameters from common percentiles. *J. applied Stat.* **32**, 17–24 (2005).
10. K Boudt, D Caliskan, C Croux, Robust explicit estimators of weibull parameters. *Metrika* **73**, 187–209 (2011).
11. SD Dubey, *Contributions to statistical theory of life testing and reliability*. (Michigan State University of Agriculture and Applied Science. Department of statistics), (1960).
12. LJ Bain, CE Antle, Estimation of parameters in the weibull distribution. *Technometrics* **9**, 621–627 (1967).
13. RV Hogg, Adaptive robust procedures: A partial review and some suggestions for future applications and theory. *J. Am. Stat. Assoc.* **69**, 909–923 (1974).
14. RJ Hyndman, Y Fan, Sample quantiles in statistical packages. *The Am. Stat.* **50**, 361–365 (1996).
15. C Bernard, R Kazzi, S Vanduffel, Range value-at-risk bounds for unimodal distributions under partial information. *Insur. Math. Econ.* **94**, 9–24 (2020).
16. WR van Zwet, *Convex Transformations of Random Variables: Nebst Stellingen*. (1964).
17. AL Bowley, *Elements of statistics*. (King) No. 8, (1926).
18. RA Groeneveld, G Meeden, Measuring skewness and kurtosis. *J. Royal Stat. Soc. Ser. D (The Stat.* **33**, 391–399 (1984).
19. PJ Bickel, EL Lehmann, Descriptive statistics for nonparametric models iv. spread in *Selected Works of EL Lehmann*. (Springer), pp. 519–526 (2012).
20. PJ Rousseeuw, C Croux, Alternatives to the median absolute deviation. *J. Am. Stat. association* **88**, 1273–1283 (1993).
21. PJ Bickel, EL Lehmann, Descriptive statistics for nonparametric models. iii. dispersion in *Selected works of EL Lehmann*. (Springer), pp. 499–518 (2012).

- 688 22. P Janssen, R Serfling, N Veraverbeke, Asymptotic normality of u-statistics based on trimmed
689 samples. *J. statistical planning inference* **16**, 63–74 (1987).
- 690 23. J Hodges, E Lehmann, Matching in paired comparisons. *The Annals Math. Stat.* **25**, 787–791
691 (1954).
- 692 24. S Dharmadhikari, K Jogdeo, Unimodal laws and related in *A Festschrift For Erich L. Lehmann*.
693 (CRC Press), p. 131 (1982).
- 694 25. AY Khintchine, On unimodal distributions. *Izv. Nauchno-Issled. Inst. Mat. Mech.* **2**, 1–7 (1938).
- 695 26. S Purkayastha, Simple proofs of two results on convolutions of unimodal distributions. *Stat. &*
696 *probability letters* **39**, 97–100 (1998).
- 697 27. RA Fisher, Moments and product moments of sampling distributions. *Proc. Lond. Math. Soc.*
698 **2**, 199–238 (1930).
- 699 28. PR Halmos, The theory of unbiased estimation. *The Annals Math. Stat.* **17**, 34–43 (1946).
- 700 29. W Hoeffding, A class of statistics with asymptotically normal distribution. *The Annals Math.*
701 *Stat.* **19**, 293–325 (1948).
- 702 30. PM Heffernan, Unbiased estimation of central moments by using u-statistics. *J. Royal Stat.*
703 *Soc. Ser. B (Statistical Methodol.)* **59**, 861–863 (1997).
- 704 31. D Fraser, Completeness of order statistics. *Can. J. Math.* **6**, 42–45 (1954).
- 705 32. AJ Lee, *U-statistics: Theory and Practice*. (Routledge), (2019).
- 706 33. J Hodges Jr, E Lehmann, Estimates of location based on rank tests. *The Annals Math. Stat.*
707 **34**, 598–611 (1963).
- 708 34. A Ehsanes Saleh, Hodges-lehmann estimate of the location parameter in censored samples.
709 *Annals Inst. Stat. Math.* **28**, 235–247 (1976).
- 710 35. RJ Serfling, Generalized l-, m-, and r-statistics. *The Annals Stat.* **12**, 76–86 (1984).
- 711 36. P Janssen, R Serfling, N Veraverbeke, Asymptotic normality for a general class of statistical
712 functions and applications to measures of spread. *The Annals Stat.* **12**, 1369–1379 (1984).
- 713 37. MG Akritas, Empirical processes associated with v-statistics and a class of estimators under
714 random censoring. *The Annals Stat.* **14**, 619–637 (1986).
- 715 38. I Gijbels, P Janssen, N Veraverbeke, Weak and strong representations for trimmed u-statistics.
716 *Probab. theory related fields* **77**, 179–194 (1988).
- 717 39. J Choudhury, R Serfling, Generalized order statistics, bahadur representations, and sequential
718 nonparametric fixed-width confidence intervals. *J. Stat. Plan. Inference* **19**, 269–282 (1988).
- 719 40. E Joly, G Lugosi, Robust estimation of u-statistics. *Stoch. Process. their Appl.* **126**, 3760–3773
720 (2016).
- 721 41. P Laforge, S Cléménçon, P Bertail, On medians of (randomized) pairwise means in *Interna-*
722 *tional Conference on Machine Learning*. (PMLR), pp. 1272–1281 (2019).
- 723 42. K Pearson, Contributions to the mathematical theory of evolution. *Philos. Transactions Royal*
724 *Soc. London. A* **185**, 71–110 (1894).