Geometric aspects of robust testing for normality and sphericity

This item was submitted to Loughborough University’s Institutional Repository by the/an author.


Additional Information:

• This is an Accepted Manuscript of an article published by Taylor & Francis in Stochastic Analysis and Applications on 06 Feb 2017, available online: http://dx.doi.org/10.1080/07362994.2016.1273785

Metadata Record: https://dspace.lboro.ac.uk/2134/23634

Version: Accepted for publication

Publisher: © Taylor & Francis

Rights: This work is made available according to the conditions of the Creative Commons Attribution-NonCommercial-NoDerivatives 4.0 International (CC BY-NC-ND 4.0) licence. Full details of this licence are available at: https://creativecommons.org/licenses/by-nc-nd/4.0/

Please cite the published version.
Geometric aspects of robust testing for normality and sphericity

Wolf-Dieter Richtera, Luboš Střelecb, Hamid Ahmadinezhadc and Milan Stehlíkd,e

aInstitute of Mathematics, University of Rostock, Rostock, Germany;
bDepartment of Statistics and Operation Analysis, Mendel University in Brno, Brno, Czech Republic;
cDepartment of Mathematical Sciences, Loughborough University, Loughborough, UK;
dDepartment of Applied Statistics, Johannes-Kepler-University Linz, Austria;
eInstitute of Statistics, University of Valparaíso, Valparaíso, Chile

ARTICLE HISTORY
Compiled October 30, 2016

CONTACT: Milan Stehlík. Email: Milan.Stehlik@jku.at
Geometric aspects of robust testing for normality and sphericity

ABSTRACT
Stochastic Robustness of Control Systems under random excitation motivates challenging developments in geometric approach to robustness. The assumption of normality is rarely met when analyzing real data and thus the use of classic parametric methods with violated assumptions can result in the inaccurate computation of p-values, effect sizes, and confidence intervals. Therefore, quite naturally, research on robust testing for normality has become a new trend. Robust testing for normality can have counter-intuitive behavior, some of the problems have been introduced in [46]. Here we concentrate on explanation of small-sample effects of normality testing and its robust properties, and embedding these questions into the more general question of testing for sphericity. We give geometric explanations for the critical tests. It turns out that the tests are robust against changes of the density generating function within the class of all continuous spherical sample distributions.

KEYWORDS
Huberization, trimming, Lehman-Bickel functional, Monte Carlo simulations, power comparison, robust tests for normality, normality, sphericity

Classification codes: 62H15, 62H10

1. Introduction

Classic parametric statistical significance tests, such as analysis of variance and least squares regression, are widely used by researchers in many disciplines of chemistry, economics or social sciences. For classic parametric tests to produce accurate results, the assumptions underlying them (e.g. normality and homoscedasticity) must be satisfied. These assumptions are rarely met when analyzing real data. The use of classic parametric methods with violated assumptions can result in the inaccurate computation of p-values, effect sizes, and confidence intervals. This may lead to substantive errors in the interpretation of data.

Recently, several articles on robust testing for normality have been written (see for example [6], [9], [14], [15], [45] and [46]). However, deeper understanding of geometry of normality tests, normality and deviations from normality is still ongoing problem. To address these topics is the main objective of this paper. In particular we compare the recently introduced robust tests on the base of RT class, thus justifying RT class to be a good base for robust tests based on robustification of first four moments. In particular we show that selected RT class of tests accommodate the alternatives which are problematic for JB test: bimodal, Weibull and uniform alternatives – see [49]. It is worth a note that since JB test has no finite exponential moments, there were no asymptotical efficiency assessment to the best knowledge of the authors. From systematic study of RT class the superiority of Geary and Uthoff tests has been found for bimodal alternatives.

We emphasize that in this paper we use the assumption that possible contamination is only due to outliers. Thus we use three techniques for outliers filtering: trimming, huberization and functional approach introduced by [3]. For normality testing when the whole distribution may be contaminated see e.g. [1].

Testing for normality is a very important issue, for example in chemometrics. It is also notable that several books in the literature, e.g. in analytical chemistry, devote a section to normality testing; see for example [27]. However, robust testing for normality is still not completely communicated to the community of chemometricians. In the next section we illustrate necessity of robust testing for normality in chemometrics.
In section 3 the M-estimates are discussed. In section 4 the general RT class of tests is presented. We also list already known tests belonging to RT class, together with the most significant new tests. We also show on convenient examples that tests from RT class improve substantially the weaknesses of classical JB test. Sections 5, 6 and 7 contain theoretical explanation on geometrical behavior of tests and algebraic manipulations on main quadratic. In section 8 the comparative study of powers of robust tests is conducted. Discussion concludes the paper. Therein also practical guidelines for robust testing of normality are given. To maintain the continuity of explanation proofs and technicalities are put into Appendix.

2. Illustrative examples – necessity of robust testing for normality in chemometrics

In this section we illustrate the necessity of robust testing for normality in chemometrics. We present a few illustrative examples where testing for normality plays an important role.

2.1. Catalytic Isomerization: influence of truncation

In several chemical application truncation of normal distribution can have a severe impact on statistical decision. For illustration, let us consider kinetics of catalytic isomerization of n-Pentane (see [7]). Therein we have rates $r_1 = \frac{\gamma K_2(x_2-x_3)K}{1+K_1 x_1+K_2 x_2+K_3 x_3}$ and $r_2 = +K_2 x_2+K_3 x_3)$, where $K_i$ are adsorption equilibrium constants, $x_i$ are partial pressures, $K$ is equilibrium constant, $\gamma$ is a constant dependent on catalyst and temperature $T$ and $i = 1, 2, 3$ are indices for hydrogen, n-pentane and isopentane, respectively.

[7] introduced variables $y_1 = \left(\frac{x_2-x_3}{r_1}\right)$, $y_2 = (y_1)^{1/2}$ and tests for importance of higher factors, later is this discussed in [22]. In [44] we pointed out that we should be aware of the fact, that the values $y_i$, which are realizations of $y = \gamma K_2/(1 + \sum K_i x_i)$, are not only positive but strictly over some positive constant (see their measured values in table III of [7] or realize their chemical meaning, i.e. $k_i$ are equilibrium adsorption constants, $x_i$ particular pressures of hydrogen, n-pentane, isopentane, respectively). Thus the proper test should be based on truncated normals. Therefore alongside the ”outlier importance” discussed in [29], there is a further issue concerning the distributional deviations from $\mathcal{F}$-distribution. For more discussion from point of view of chemometrics see [46].

2.2. Example: Control Charts for Chemical processes (positive kurtosis)

Here, we suppose some alternative distributions with zero skewness and positive excess (kurtosis). For the purpose of tests comparison we suppose the following alternatives commonly used in chemometrics and related fields:

- Laplace ($t, s$) distribution, where $t$ and $s$ are the location and scale parameters, respectively, defined as Laplace(0;1), with $SK = 0$ and $K_{exc} = 3$ where $SK$ is skewness and $K_{exc}$ is excess (kurtosis).
- $t$-Student($\nu$) distribution, where $\nu$ is the number of degrees of freedom, with $SK = 0$ for $\nu > 3$ and $K_{exc} = \frac{6}{\nu-4}$ for $\nu > 4$ where $SK$ is skewness and $K_{exc}$ is...
kurtosis.
• Logistic \((t, s)\) distribution, where \(t\) and \(s\) are the location and scale parameters, respectively, defined as \(\text{Logistic}(0;1)\), with \(SK = 0\) and \(K_{\text{exc}} = 6/5\) where \(SK\) is skewness and \(K_{\text{exc}}\) is kurtosis.

Table 1 presents the power of selected classical and robust tests for normality against the mentioned alternatives (note that these results are based on results presented in [46]). Based on these results we can conclude that the best tests are \(SJ_{\text{dir}}\), \(RJB\) and from \(RT\) class \(TTRT2\) tests, which powerfull outperform the other tests. For example, power of the \(SJ_{\text{dir}}\) test against Laplace distribution with \(SK = 0\) and \(K_{\text{exc}} = 3\) for \(n = 20\) is 0.391, \(RJB\) test has power 0.357 and finally the power of \(TTRT2\) test is 0.345. On the other hand, commonly used the \(SW\) test has power only 0.260 and the classical \(JB\) test 0.307. If we suppose large sample size \(n = 100\) the power of mentioned tests is following: \(SJ_{\text{dir}} = 0.942\), \(RJB = 0.889\), \(TTRT2 = 0.862\), \(SW = 0.796\) and finally \(JB = 0.802\).

Similarly, if we suppose \(t\)-Student(\(\nu\)) distribution \(t(7)\) with \(SK = 0\) and \(K_{\text{exc}} = 2\) and \(n = 100\), the results are following: \(SW\) test has power 0.363, \(JB = 0.454\), \(SJ_{\text{dir}} = 0.445\), \(TTRT2 = 0.468\) and finally \(RJB\) test has power 0.477. Similar results are also achieved for \(t\)-Student(\(\nu\)) distribution \(t(5)\) and logistic distribution.

From the mentioned results we can see that power between the most powerful \(RJB\), \(SJ_{\text{dir}} = 0.445\) and \(TTRT1\) tests vanished for distributions with excess kurtosis close to kurtosis of Gaussian normal distribution.

<table>
<thead>
<tr>
<th></th>
<th>(n = 20)</th>
<th></th>
<th></th>
<th>(n = 100)</th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Laplace</td>
<td>(t_5)</td>
<td>(t_7)</td>
<td>logistic</td>
<td>Laplace</td>
<td>(t_5)</td>
</tr>
<tr>
<td>(AD)</td>
<td>0.276</td>
<td>0.172</td>
<td>0.116</td>
<td>0.107</td>
<td>0.824</td>
<td>0.483</td>
</tr>
<tr>
<td>(DT)</td>
<td>0.290</td>
<td>0.223</td>
<td>0.160</td>
<td>0.142</td>
<td>0.725</td>
<td>0.587</td>
</tr>
<tr>
<td>(JB)</td>
<td>0.307</td>
<td>0.233</td>
<td>0.168</td>
<td>0.148</td>
<td>0.802</td>
<td>0.651</td>
</tr>
<tr>
<td>(JBU)</td>
<td>0.320</td>
<td>0.238</td>
<td>\textbf{0.170}</td>
<td>0.152</td>
<td>0.814</td>
<td>0.661</td>
</tr>
<tr>
<td>(LT)</td>
<td>0.232</td>
<td>0.139</td>
<td>0.093</td>
<td>0.083</td>
<td>0.699</td>
<td>0.329</td>
</tr>
<tr>
<td>(MC_{\text{LR}})</td>
<td>0.058</td>
<td>0.051</td>
<td>0.052</td>
<td>0.049</td>
<td>0.130</td>
<td>0.058</td>
</tr>
<tr>
<td>(RJB)</td>
<td>0.357</td>
<td>\textbf{0.241}</td>
<td>0.169</td>
<td>\textbf{0.153}</td>
<td>0.889</td>
<td>\textbf{0.678}</td>
</tr>
<tr>
<td>(SJ_{\text{dir}})</td>
<td>\textbf{0.391}</td>
<td>\textbf{0.231}</td>
<td>0.160</td>
<td>\textbf{0.144}</td>
<td>\textbf{0.942}</td>
<td>0.659</td>
</tr>
<tr>
<td>(SW)</td>
<td>0.260</td>
<td>0.186</td>
<td>0.133</td>
<td>0.116</td>
<td>0.796</td>
<td>0.569</td>
</tr>
</tbody>
</table>

\(M_{\text{MRT1}}\) | 0.268 | 0.208 | 0.148 | 0.130 | 0.790 | 0.631 | 0.436 | 0.375  |
| \(M_{\text{MRT2}}\) | 0.263 | 0.204 | 0.142 | 0.126 | 0.787 | 0.626 | 0.429 | 0.368  |
| \(T_{\text{TRT1}}\) | 0.099 | 0.068 | 0.064 | 0.063 | 0.341 | 0.306 | 0.156 | 0.107  |
| \(T_{\text{TRT2}}\) | \textbf{0.345} | \textbf{0.235} | \textbf{0.165} | \textbf{0.151} | \textbf{0.862} | \textbf{0.667} | \textbf{0.468} | \textbf{0.413} |

\(\text{Table 1.}\) Power of the selected tests for normality against various heavy-tailed alternatives commonly used in chemometrics and related fields for \(n = 20\) and \(n = 100\)

3. M-estimator of location

First of all, the M-estimates are a generalisation of the maximum likelihood (ML) estimates (see Maronna et al. 2006). Some popular robust M-estimates are Huber-estimate, Hampel estimate, Andrews’ wave, Tukey’s biweight. A lot of literature is available for robust M-estimates, for example, Huber (1981), Hampel et al. (1986), Rousseeuw and Leroy (1987), Staudte and Sheather (1990), Wilcox (1997) and
Maronna et al. (2006). We will use a function \( \psi(x, C) = x \) for \(|x| < C\) and \( \psi(x, C) = C\text{sign}(x) \) for \(|x| \geq C\).

We recall, that the most efficient location estimator, i.e. MLE, corresponds typically to the highest power at the alternative. Notice that MLE is a special form of M-estimate for a location parameter (since it is obtained for \( \psi := \ln f \) and \( \Psi = f'/f \), where \( \ln f \) is log likelihood and the M-estimator can be expressed as \( \arg\min \sum_{i=1}^{n} \psi(x_i - M) \), where \( M \) is the location parameter).

The robust inference emphasises the concepts of breakdown point and influence function of an estimator. The latter is then used to explore the efficiency and robustness properties of the estimator. Pure location functional approach, which has been introduced by P. E. Bickel and E. L. Lehmann as well as P. J. Rousseeuw and A. M. Verbeek has been introduced by P. E. Bickel and E. L. Lehmann to the sample

\[
k_j(x) = \frac{1}{n-2r} \sum_{m=r+1}^{n-r} \varphi_j(X_{(m)} - T(F_n, s))
\]

where \( M_j(r, T(F_n, s)) = \frac{1}{n-2r} \sum_{m=r+1}^{n-r} \varphi_j(X_{(m)} - T(F_n, s)) \) for \( j \in \{0, 1, 2, 3, 4\} \), where \( \varphi_j \) is a tractable and continuous function, where \( \varphi_0(x) = \sqrt{\pi/2} |x| \) and \( \varphi_j(x) = x^j \) for \( j \in \{1, 2, 3, 4\} \), \( X_{(m)} \) is the order statistic, \( T(F_n, s) \) is a location functional applied to the sample \( X_1, X_2, \ldots, X_n \), \( r \) and \( s \) are the trimming constants for moments and location, respectively, \( K_1 \) and \( K_2 \) are small-sample variants of mean corrections, \( C_1 \) and \( C_2 \) are asymptotic constants, \( \alpha_1, \alpha_2, \alpha_3 \) and \( \alpha_4 \) are exponents, and finally, \( k_1(n) \) and \( k_2(n) \) are functions of sample size \( n \).

The reason so many classical procedures are nonrobust to outliers is that the parameters of the model are expressed in terms of moments, and their classical estimators are very sensitive to outliers. Another approach to robustness is to concentrate on the parameters of interest suggested by the problem under study. It may well turn out that these parameters can be expressed as functions of the underlying distribution independently of a particular parametric model; that is as descriptive measures. If these descriptive measures are judiciously chosen, their naturally induced estimators are robust to aberration in the data (see e.g. [3]).

4. General RT class

The general RT class is based on robustification of the classical Jarque-Bera test introduced by [4]. The general RT class test statistic is defined by [45] for purpose of robust testing for normality against Pareto tails and has the following general form

\[
RT = \frac{k_1(n)}{C_1} \left( \frac{M_{j_1}^{\alpha_1}(r_1, T_{(i_1)}(s_1))}{M_{j_1}^{\alpha_2}(r_2, T_{(i_2)}(s_2))} - K_1 \right)^2 + \frac{k_2(n)}{C_2} \left( \frac{M_{j_4}^{\alpha_3}(r_3, T_{(i_3)}(s_3))}{M_{j_4}^{\alpha_4}(r_4, T_{(i_4)}(s_4))} - K_2 \right)^2 \tag{1}
\]

where \( M_j(r, T(F_n, s)) = \frac{1}{n-2r} \sum_{m=r+1}^{n-r} \varphi_j(X_{(m)} - T(F_n, s)) \) for \( j \in \{0, 1, 2, 3, 4\} \), where \( \varphi_j \) is a tractable and continuous function, where \( \varphi_0(x) = \sqrt{\pi/2} |x| \) and \( \varphi_j(x) = x^j \) for \( j \in \{1, 2, 3, 4\} \), \( X_{(m)} \) is the order statistic, \( T(F_n, s) \) is a location functional applied to the sample \( X_1, X_2, \ldots, X_n \), \( r \) and \( s \) are the trimming constants for moments and location, respectively, \( K_1 \) and \( K_2 \) are small-sample variants of mean corrections, \( C_1 \) and \( C_2 \) are asymptotic constants, \( \alpha_1, \alpha_2, \alpha_3 \) and \( \alpha_4 \) are exponents, and finally, \( k_1(n) \) and \( k_2(n) \) are functions of sample size \( n \).

Note, that for the construction of the general RT class test statistic it was used location functional approach, which has been introduced by P. E. Bickel and E. L. Lehmann in a series of papers (e.g. [3]) and as was shown in [45] and [46] it looks to be playing a crucial role also by robust testing for normality. As it will be seen later, the power of RT class test mimics the effectiveness of location estimator in typical cases. Thus trade off between power and robustness is a typical issue here.

**Definition 1.** (Location Functional, see [3]) Let \( T(F) \) be a function defined on the
set of distribution functions. We say that $T(F)$ is a location functional if the following conditions hold:

1. if $G$ is stochastically larger than $F$ then $T(G) \geq T(F)$,
2. $T(F_{aX+b}) = aT(F_X) + b$,
3. $T(F_{-X}) = -T(F_X)$.

Then we will call $\theta = T(F)$ a location parameter of $F$. Having a sample, we work with its empirical variant $\theta_n = T(F_n)$, $F_n$ being an empirical cdf of a sample.

In the general $RT$ class we used the following four different location estimators:

- mean: $T_{(0)} = \frac{1}{n} \sum_{i=1}^{n} X_i$,
- median: $T_{(1)} = F^{-1}(1/2)$,
- trimmed mean: $T_{(2)}(s) = \frac{1}{n-2s} \sum_{i=s+1}^{n-s} X_i$, where $X_{(i)}$ is the $i$-th order statistic of the sample and $s$ is the trimming constant for location,
- pseudo-median: $T_{(3)} = \text{median}_{i \leq j}(X_i + X_j)/2$, i.e. the median of the set $\{(X_1 + X_1)/2, (X_1 + X_2)/2, (X_1 + X_3)/2, \ldots, (X_1 + X_n)/2, (X_2 + X_2)/2, (X_2 + X_3)/2, \ldots, (X_2 + X_n)/2, \ldots, (X_{n-1} + X_n)/2, (X_n + X_n)/2\}$.

Note, that some theoretical results on consistency and asymptotic $\chi^2$-distribution of the general $RT$ class test statistic can be found in [45] and [46].

In the $RT$ class we have also a direct relation to M-estimation. From the classical M-estimation ([19], page 6) is proposed the odd function $\psi(x, c) = x$, for $|x| < c$ and $c \text{ signum}(x)$ otherwise. Notice, that a special case of $RT_{JB}$ class fits to M-procedure based on

$$a_n \left[ \sum_{i=1}^{n} \psi^3 \left( \frac{X_i - M_n}{\hat{\sigma}}, c \right) \right]^2 + b_n \left[ \sum_{i=1}^{n} \psi^4 \left( \frac{X_i - M_n}{\hat{\sigma}}, c \right) - k \right]^2$$

(2)

for a well choice of censoring, function $\psi$ and $\hat{\sigma}$ being a robust estimator of variance.

The form of (2) has been already studied by [16]. Furthermore, we can use also variance functional construction given by Bickel-Lehman, similarly to the mean functional.

Remark 1. Constants $C_1$ and $C_2$ used in general $RT$ class test statistic

As was noted in [46] choosing of appropriate constants $C_1$ and $C_2$ is the hardest aspect of the variants of $RT$ class tests, because to obtain the constants $C_1$ and $C_2$ we need to find the expressions for $E(M_{h_n u_n}^k)$ for a finite sample size. Such calculations are very tedious and therefore we obtained these constants from Monte Carlo simulations (see [45]). As was also mentioned in [46] the critical constant (for small and mid samples) under the trimming of moments ($r > 0$) are different from critical constants without trimming ($r = 0$), since only asymptotical distribution is normal (see [48]) in this case.

Remark 2. Trimming

Notice that two levels of trimming enter $RT$ class test statistic: first trimming (with trimming constant $s$) enters trimming in the location estimator $T(F_n)$, the latter on trimming (with trimming constant $r$) enters $M_j(r, T(s))$. Amazing property of $RT$ class and robust tests in general is, that power of $RT$ class mimics the effectiveness of location estimator. Thus practitioner can tune how much of robustness is needed,
of course at price of the power. One should be really careful here: mechanical down-weighting of peculiar observations may divert attention from important clues to new discovery. Based on our simulation study we can suggest the suitable choosing of these constants: \( s = r = 0.05n \).

**Remark 3. Robustness of RT class**

A statistical procedure is called robust, if its performance is insensitive to small deviations of the actual situation from the idealized theoretical model (see [20]). In particular, a robust procedure should be insensitive to the presence of a few "bad" observations (but at the same time the discordant minority of the observations might be prime source of information for improving the theoretical model). Here we assume that possible contamination is due to presence of outliers. Based on our research, we would like to point out the following robust aspects of RT class of tests:

- The classical JB test is not robust: it has a 0-breakdown point (see [6]).
- Note, that our tests are constructed on the robust Bickel-Lehman construction of location ([3]).
- The 2nd "version" of robust estimation is trimming of location with trimming constant \( s \) (so called "trimmed-mean tests") and moments with trimming constant \( r \) (so called "trim-trim tests"). Note that selected trim-trim tests have the same power level as Medcouple tests introduced by [6], so they are very robust.
- Normality testing procedure is typically a pre-test preceding some further testing or inference. Therefore one interesting issue is robustness with respect to other procedures, assuming that RT tests did not rejected normality. The influence of different shapes of distribution with the same first 4 moments on robustness has been discussed for sequential \( t \)-test by [30].

**Remark 4. Trade off between power and robustness**

Two typical extremal behaviors occur in robust testing: the tests which are more robust have smaller power (since they are not affected by single outliers) and tests with higher power are typically less robust (because they are affected by single outliers). An example of the first extreme case are the Medcouple tests and selected RT class tests based on trimming and an example of the second extreme case are the RT tests based on mean-median combination. To be more precise, for instance the robust test of normality could be also obtained by removing the outliers from the data, using an outlier detection rule such as provided by the boxplot or a rule based on robust estimators of location and scale. When the majority of the data are instead normally distributed, this is a valuable alternative to the robust tests based on medcouple as both the boxplot and the most popular robust estimators of location and scale (such as M-estimators) are based on this normal assumptions and thus will indicate the correct set of outliers. However it becomes more complicated when even majority of the data points do not come from a normal distribution (see [6]). However, probably it will be in many practical situations more worth to conduct test with clear outliers adaptively, i.e. recognizing and deleting in the first step and then using the zero breakdown robust version of JB for normality testing. To illustrate this general framework, we use the right-location-standard normal distribution based on mixture of two normal distributions with various parameters. For this purpose we assume the right-contamination with distribution function: \( F = (1 - p)N(\mu_1, \sigma_1^2) + pN(\mu_2, \sigma_2^2) \) for \( p = 0.05, \mu_1 = 0, \mu_2 = 3, \sigma_1^2 = \sigma_2^2 = 1 \), i.e. \( F = 0.95N(0, 1) + 0.05N(3, 1) \), for large sample size \( n = 100 \). For mentioned contamination and sample size the power of the classical Jarque-Bera test is 0.744, power of SW test is 0.685 and MMRRT1 has power 0.686. On the other
hand the $TTRT1$ test has power only 0.190, $MC_{LR}$ test only 0.076, the Jarque-Bera and Shapiro-Wilk tests with deleting outliers based on boxplot in the first step have power 0.110 and 0.055, respectively. For complete comparison for right and central contamination see Table 1 in [46].

4.1. A special known tests of RT class

As can be seen from (1) there exist a vast amount of RT class tests, which we can obtain for a different settings of $r_i, T_{(i)}(s_i)$ etc. Next we refer to the RT class tests which has been studied already in the literature. These are i) classical JB test, ii) test of Urzua, iii) robust JB ($RJB$) test, iv) skewness test $b_1$, v) the kurtosis test $\sqrt{b_2}$, vi) the Geary test $a$, vii) the Utho test only 0.076, the Jarque-Bera test behaves well in comparison with some other tests for normality if the alternatives belong to the Pearson family. However, the JB test behaves very badly for distributions with short tails and bimodal shape, sometimes it is even biased (see [49]).

i) The classical Jarque-Bera test is a special case of RT test without trimming for $K_1 = 0, K_2 = 3, k_1(n) = n, k_2(n) = n$, $C_1 = 6, C_2 = 24, \alpha_1 = 1, \alpha_2 = 3/2, \alpha_3 = 1, \alpha_4 = 2, T_{(i)} = T_{(i_2)} = T_{(i_3)} = T_{(i_4)} = T_{(0)}, j_1 = 3, j_2 = 2, j_3 = 4, j_4 = 2$. The Jarque-Bera test statistic $JB$ is defined as

$$JB = \frac{n}{6} \left( \frac{\hat{\mu}_3}{\sqrt{\hat{\mu}_2^3/2}} \right)^2 + \frac{n}{24} \left( \frac{\hat{\mu}_4}{\hat{\mu}_2^2} - 3 \right)^2.$$

As pointed out by several authors (see for example [51]) the classical JB test behaves well in comparison with some other tests for normality if the alternatives belong to the Pearson family. However, the JB test behaves very badly for distributions with short tails and bimodal shape, sometimes it is even biased (see [49]).

ii) Test of Urzua (see [51]) is a special case of RT test without trimming for $K_1 = 0, K_2 = 3, k_1(n) = ((n + 1)(n + 3))/(n - 2), k_2(n) = ((n + 1)^2(n + 3)(n + 5))/(n(n - 2)(n - 3))$, $C_1 = 6, C_2 = 24, \alpha_1 = 1, \alpha_2 = 3/2, \alpha_3 = 1, \alpha_4 = 2, T_{(i)} = T_{(i_2)} = T_{(i_3)} = T_{(i_4)} = T_{(0)}, j_1 = 3, j_2 = 2, j_3 = 4, j_4 = 2$. The Urza’s Jarque-Bera test statistic $JBU$ is defined as

$$JBU = \frac{(n+1)(n+3)}{6(n-2)} \left( \frac{\hat{\mu}_3}{\sqrt{\hat{\mu}_2^3/2}} \right)^2 + \frac{(n+1)^2(n+3)(n+5)}{n(n-2)(n-3)} \left( \frac{\hat{\mu}_4}{\hat{\mu}_2^2} - 3 \right)^2.$$

iii) The robust Jarque-Bera test (see [14]) is a special case of RT test without trimming for $K_1 = 0, K_2 = 3, k_1(n) = n, k_2(n) = n$, $C_1 = 6, C_2 = 64, \alpha_1 = 1, \alpha_2 = 3, \alpha_3 = 1, \alpha_4 = 4, T_{(i)} = T_{(i_2)} = T_{(0)}, T_{(i_3)} = T_{(i_4)} = T_{(i_4)} = T_{(1)}, j_1 = 3, j_2 = 0, j_3 = 4, j_4 = 0$. The robust Jarque-Bera test statistic $RJB$ is defined as

$$RJB = n \left( \frac{\hat{\mu}_3}{\hat{\mu}_2} \right)^2 + n \left( \frac{\hat{\mu}_4}{\hat{\mu}_2^2} - 3 \right)^2.$$
\( K_2 = 0, k_1(n) = n, k_2(n) = 0, C_1 = 6, \alpha_1 = 1, \alpha_2 = 3/2, T_{(i_1)} = T_{(i_2)} = T_{(0)}, j_1 = 3 \) and \( j_2 = 2 \). The skewness test statistic \( b_1 \) is defined as

\[
    b_1 = \frac{n}{6} \left( \frac{\hat{\mu}_3}{\hat{\mu}_2^{3/2}} \right)^2.
\]

v) The kurtosis test \( \sqrt{b_2} \) is a special case of RT test without trimming for \( K_1 = 3, K_2 = 0, k_1(n) = n, k_2(n) = 0, C_1 = 24, \alpha_1 = 1, \alpha_2 = 2, T_{(i_1)} = T_{(i_2)} = T_{(0)}, j_1 = 4 \) and \( j_2 = 2 \). The kurtosis test statistic \( \sqrt{b_2} \) is defined as

\[
    \sqrt{b_2} = \frac{n}{24} \left( \frac{\hat{\mu}_4}{\hat{\mu}_2^2} - 3 \right)^2.
\]

vi) The Geary’s test \( a \) (see [13]) is a special case of RT test without trimming for \( K_1 = 0, K_2 = 0, k_1(n) = 1/n, k_2(n) = 0, C_1 = 1, \alpha_1 = 1/2, \alpha_2 = 1/4, T_{(i_1)} = T_{(i_2)} = T_{(0)}, \varphi_0(x) = |x|, j_1 = 0, j_2 = 2 \) as an alternative to \( b_2 \), because the small sample properties where more tractable for \( a \). The Geary’s test statistic \( a \) (originally denoted \( w_n^0 \)) is defined as

\[
    a = \frac{1}{n} \sum_{i=1}^{n} \frac{|X_i - \bar{X}|}{\sqrt{m_2}}.
\]

vii) The Uthoff’s test \( U \) (see [52]) is a special case of RT test without trimming for \( K_1 = 0, K_2 = 0, k_1(n) = 1/n, k_2(n) = 0, C_1 = 1, \alpha_1 = 1/2, \alpha_2 = 1/4, T_{(i_1)} = T_{(i_2)} = T_{(1)}, \varphi_0(x) = |x|, j_1 = 0 \) and \( j_2 = 2 \). The Uthoff’s test statistic \( U \) is defined as

\[
    U = \frac{1}{n} \sum_{i=1}^{n} \frac{|X_i - M_n|}{\sqrt{m_2}}.
\]

viii) \( SJ \) test (see [15]) is a special case of RT test without trimming for \( K_1 = 0, K_2 = 0, k_1(n) = \sqrt{\pi/2}/n, k_2(n) = 0, C_1 = 1, \alpha_1 = 1/4, \alpha_2 = 1/2, T_{(i_1)} = T_{(i_2)} = T_{(1)}, \varphi_0(x) = |x|, j_1 = 2 \) and \( j_2 = 0 \). The \( SJ \) test statistic \( SJ \) is defined as

\[
    SJ = \frac{\sqrt{\pi/2}}{n} \sum_{i=1}^{n} \sqrt{m_2} \frac{|X_i - M_n|}{|X_i - M_n|}.
\]

4.2. Tractable approach to RT class

In [45] we introduced general RT class test statistic (see (1)) as well as \( RT_{JB} \) and \( RT_{RJB} \) subclasses. The \( RT_{JB} \) subclass test statistic is defined as follows
where is a special case of RT test statistics for \( k_1(n) = n, k_2(n) = n, \alpha_1 = 1, \alpha_2 = 3/2, \alpha_3 = 1, \alpha_4 = 2, j_1 = 3, j_2 = 2, j_3 = 4, j_4 = 2. \)

Similarly, \( RT_{RJB} \) subclass test statistic is based on robustification of the robust Jarque-Bera test introduced by [15] and is defined as follows

\[
RT_{RJB} = \frac{n}{C_1} \left( \frac{M_3(r_1, T_{(1)}(s_1))}{M_0^3(r_2, T_{(2)}(s_2))} - K_1 \right)^2 + \frac{n}{C_2} \left( \frac{M_4(r_3, T_{(1)}(s_3))}{M_0^4(r_4, T_{(4)}(s_4))} - K_2 \right)^2, \tag{4}
\]

which is a special case of RT test statistics for \( k_1(n) = n, k_2(n) = n, \alpha_1 = 1, \alpha_2 = 3, \alpha_3 = 1, \alpha_4 = 4, j_1 = 3, j_2 = 0, j_3 = 4, j_4 = 0. \)

By clustering based on power values from all analyzed tests of RT class the following representatives with good properties against most common alternatives has been obtained (for more details of these test statistics see section 3.2 in [46]):

- The mean-median \( MMRT_1 \) test which is suitable for testing of normality against heavy- and light-tailed asymmetric alternatives.
- The mean-median \( MMRT_2 \) test which is suitable for testing of normality against heavy- and light-tailed asymmetric alternatives as well as bimodal and short-tailed symmetric alternatives.
- The trim-trim \( TTTRT_1 \) test with trimming \( s = r = 0.05n \) which is suitable for testing of normality against short-tailed symmetric alternatives and which is also more robust than most other tests of normality.
- The trim-trim \( TTTRT_2 \) test with trimming \( s = r = 0.05n \) which is suitable for testing of normality against heavy-tailed symmetric alternatives and has good robust properties.

Our pilot simulation study showed that mean-median \( MMRT_1 \) and \( MMRT_2 \) tests have comparable power with the most powerful tests (\( SW \) and \( AD \)) against heavy- and light-tailed asymmetric alternatives like one side Cauchy, Weibull and exponential. Similarly, \( TTTRT_2 \) test has comparable power with the most powerful tests (\( SJ_{dir} \) and \( RJB \)) against heavy-tailed symmetric alternatives like Cauchy and \( t_5 \). Consequently, \( MMRT_2 \) and \( TTTRT_1 \) tests have comparable power with the most powerful tests (\( SW, AD \) and \( DT \)) against short-tailed symmetric alternatives like beta and uniform. Finally, \( MMRT_2 \) test is also suitable for bimodal alternatives. Note that especially \( TTTRT_1 \) test has good robustness properties – this test is more robust than most other tests of normality.

Power of analyzed normality tests against symmetric and asymmetric heavy-, light- and short-tailed alternatives was presented in [46]. We found that \( SJ_{dir}, RJB \) and \( TTTRT_2 \) tests are the best tests against symmetric heavy-tailed alternatives like Cauchy, Laplace and \( t_5 \). For moderately heavy-tailed symmetric alternatives like \( t_5 \), \( t_7 \) and logistic the \( RJB \) and \( TTTRT_2 \) test are most powerful. If we suppose the light-tailed asymmetric alternatives like lognormal, exponential, Burr and Weibull, the \( SW, AD, MMRT_1 \) and \( MMRT_2 \) tests perform well. For the short-tailed symmetric alternatives
like beta and uniform, the \( SW, AD, MMRT2 \) and \( TTRT1 \) tests have reasonably high power. On the other hand, the \( JB, JBU, RJB \) and \( SJ_{dir} \) tests are mostly biased for small sample size \( n = 20 \) \( (SJ_{dir} \) is biased even for large sample size \( n = 100 \)).

5. Geometric interpretation of skewness and kurtosis based test statistics

This section is aimed to prepare the reader for the derivation of the exact cumulative distribution functions (cdfs) of skewness and kurtosis based statistics which will be given in the next section.

Let us recall that if \( x_1, \ldots, x_n \) is a concrete sample of size \( n \) then the skewness test statistic based upon this sample allows the representation

\[
b_1 = \frac{n}{6} \left( \frac{\hat{\mu}_3}{\hat{\mu}_2^2} \right)^2 = \frac{n}{6} \left[ \frac{\frac{1}{n} \sum_{i=1}^{n} (x_i - \bar{x})^2}{\left( \frac{1}{n} \sum_{i=1}^{n} (x_i - \bar{x})^2 \right)^{3/2}} \right]^2. \tag{5}
\]

Replacing the sample mean in nominator and denominator of this ratio with different location estimations of Lehman-Bickel type, we consider in the following statistic

\[
\tilde{b}_1 = \frac{n}{6} \left[ \frac{\frac{1}{n} \sum_{i=1}^{n} (x_i - \tilde{x})^3}{\left( \frac{1}{n} \sum_{i=1}^{n} (x_i - \tilde{x})^2 \right)^{3/2}} \right]^2, \tilde{x} = \sum_{i=1}^{n} w_i x_i, w_i \geq 0, \sum_{i=1}^{n} w_i = 1, \tilde{x} = \sum_{i=1}^{n} a_i x_i. \tag{6}
\]

Let \( \tilde{B}_1 \) the population based version of the sample based quantity \( \tilde{b}_1 \). Because we are interested in the probability \( P(\tilde{B}_1 < t) \), we are interested in a geometric understanding of the subset \( \{ x \in \mathbb{R}^n : \tilde{b}_1 < t \}, t \in \mathbb{R} \) of the \( n \)-dimensional sample space \( \mathbb{R}^n \). It turns out that the set \( \{ x \in \mathbb{R}^n : \tilde{b}_1 < t \}, t \in \mathbb{R} \) is a cone having its vertex in the origin. However, for making things as much visible as possible, our discussion of the exact distribution of \( \tilde{B}_1 \) is restricted here to the case

\[
n = 2, w_1 = \Theta, w_2 = (1 - \Theta), 0 < \Theta < 1, a_1 = 1, a_2 = a > 0 \text{ and } x_1 = x, x_2 = y. \tag{7}
\]

The cone \( \{ x \in \mathbb{R}^2 : \tilde{b}_1 < t \}, t \in \mathbb{R} \) can be closer described then in an easy way because the statistic \( \tilde{b}_1 \) can be reformulated then as follows

\[
\tilde{B}_1 = \frac{1}{3} \left[ \frac{\frac{1}{2} \left( (x - \Theta x - (1 - \Theta)y)^3 + (y - \Theta x - (1 - \Theta)y)^3 \right)}{\left( \frac{1}{2} ((x - x - ay)^2 + (y - x - ay)^2) \right)^{3/2}} \right]^2 = \frac{2}{3} \left[ (1 - \Theta)^3 (x - y)^3 + \Theta^3 (x - y)^3 \right]^2 \left[ a^2 y^2 + ((1 - a)y - x)^2 \right]^3 = \frac{2}{3} (x - y)^6 \left[ (1 - \Theta)^3 - \Theta^3 \right]^2 \left[ a^2 y^2 + ((1 - a)y - x)^2 \right]^3 = \frac{2}{3} (1 - \Theta)^6 \left( (1 - \Theta)^3 - \Theta^3 \right)^2 \left[ a^2 z^2 + ((1 - a)z - 1)^2 \right]^3.
\]
where \( z = \frac{y}{x} \). Analogously, the excess (or kurtosis) test statistic based upon the sample \( x_1, ..., x_n \) allows the representations

\[
\sqrt{b_2} = \frac{n}{24} \left( \frac{\hat{\mu}_4}{\hat{\mu}_2^2} - 3 \right)^2 = \frac{n}{24} \left[ \frac{1}{n} \sum_{i=1}^{n} (x_i - \bar{x}_n)^4}{\frac{1}{n} \sum_{i=1}^{n} (x_i - \bar{x}_n)^2} - 3 \right]^2. \tag{8}
\]

Replacing the sample mean with \( \tilde{x}_z \) and \( \tilde{x}_n \) as before, we consider now the statistic

\[
\sqrt{\tilde{b}_2} = \frac{n}{24} \left[ \frac{1}{n} \sum_{i=1}^{n} (x_i - \tilde{x}_z)^4}{\frac{1}{n} \sum_{i=1}^{n} (x_i - \tilde{x}_n)^2} - 3 \right]^2. \tag{9}
\]

Under the restrictions in (7), this statistic reads as

\[
\sqrt{\tilde{b}_2} = \frac{1}{12} \left[ \frac{1}{\gamma} \left( (x - (1 - \Theta)x - (1 - \Theta)y)^4 + (y - \Theta x - (1 - \Theta)y)^4 \right)}{(\gamma (x - x - ay)^2 + (y - x - ay)^2))^2} - 3 \right]^2 = \frac{1}{12} \left[ \frac{(1 - z)^4}{(a^2 y^2 + ((1 - a)y - x))^2} - 3 \right]^2
\]

where

\[
\gamma = 2 \left[ (1 - \Theta)^4 + \Theta^4 \right], \quad z = \frac{y}{x}.
\]

The sets

\[
\tilde{C}_1(t) = \{(x, y) \in R^2 : \sqrt{b}_1 < t \} \quad \text{and} \quad \tilde{C}_2(t) = \{(x, y) \in R^2 : \sqrt{\tilde{b}_2} < t \}, \quad t > 0 \quad (10)
\]

can be proved to be cones in \( R^2 \) having their vertex in the origin.

Let us remark the interesting fact that it is known that a value of Student’s distribution can be interpreted as being the value of the standard Gaussian measure taken at a suitably defined cone having vertex in the origin, see in [34]. Moreover, such a value does not depend on the df of a spherical measure, and is therefore the same for all spherical measures. We will follow this idea and represent the exact distribution of \( \tilde{B}_1 \) in Section 7. For a \( p \)-generalization of this observation, see [35].

We determine now the boundary lines of the cone \( \tilde{C}_1(t) \). The points of these lines satisfy the equation

\[
(1 - z)^2 = \sqrt{\frac{t}{\beta}} \left[ (a^2 z^2 + ((1 - a)z - 1)^2 \right] \tag{11}
\]

where \( \beta = \frac{3}{2}(1 - \Theta)^3 - \Theta^3 \). The points which belong the the inner of this cone satisfy the corresponding inequality

\[
(1 - z)^2 < \sqrt{\frac{t}{\beta}} \left[ (a^2 z^2 + ((1 - a)z - 1)^2 \right. \tag{12}
\]

Solving equation (11) leads to

\[ 1 - 2z + z^2 = \sqrt{\frac{t}{\beta}} \left[ (a^2z^2 + ((1 - a)^2z^2 - 2(1 - a)z + 1) \right] \]

and

\[ z^2 \left[ 1 - (a^2 + (1 - a)^2) \sqrt{\frac{t}{\beta}} \right] + z \left[ -2 + 2(1 - a) \sqrt{\frac{t}{\beta}} \right] + 1 - \sqrt{\frac{t}{\beta}} = 0. \]

Before division by \( \frac{1}{h} \left( a^2 + (1 - a)^2 \right) \sqrt{\frac{t}{\beta}} \) we need to take care that \( a^6 = q/4 \) and \( a^7 = q/2 \). The solutions of the equation \( p_1(z) = 0 \) where

\[ p_1(z) = \frac{z^2 + z \left( -2 + 2(1 - a) \sqrt{\frac{t}{\beta}} \right) + 1 - \sqrt{\frac{t}{\beta}}}{1 - (a^2 + (1 - a)^2) \sqrt{\frac{t}{\beta}}} \]

are

\[ z_{1/2} = -p/2 \pm \sqrt{p^2/4 - q} \]

where

\[ p = \frac{-2 + 2(1 - a) \sqrt{\frac{t}{\beta}}}{1 - (a^2 + (1 - a)^2) \sqrt{\frac{t}{\beta}}} \quad \text{and} \quad q = \frac{1 - \sqrt{\frac{t}{\beta}}}{1 - (a^2 + (1 - a)^2) \sqrt{\frac{t}{\beta}}}. \] (13)

Thus the boundary lines of the cone \( \tilde{C}_1(t) \) are

\[ y_1 = (-p/2 + \sqrt{p^2/4 - q})x \quad \text{and} \quad y_2 = (-p/2 - \sqrt{p^2/4 - q})x, \] (14)

and the points \((x, y)^T\) from the inner part of this cone satisfy the inequality

\[ p_1(y/x) < 0. \]

If we denote the angle between the boundary lines of the cone \( \tilde{C}_1(t) \) by \( \alpha_1(t) \) then

\[ \alpha_1(t) = |\arctan(-p/2 + \sqrt{p^2/4 - q}) - \arctan(-p/2 - \sqrt{p^2/4 - q})|. \] (15)

The following Figure 1 displays the boundary lines of cone \( \tilde{C}_1(t) \) for \( \theta = 1/3, \ a = 1, \ t = \frac{12250}{59599} \).

It can be shown in a similar manner that the boundary lines of the cone \( \tilde{C}_2(t) \) are
Figure 1. the boundary lines of cone $\tilde{C}_1(t)$ for $\theta = 1/3$, $a = 1$, $t = \frac{12250}{140049}$.

Thus, the angle between the boundary lines of $\tilde{C}_2(t)$ allows the representation (15), but with $p$ and $q$ from (16). Notice, that for same values for parameters $\theta, a$ and $t$ there is no real cone $\tilde{C}_2(t)$.

6. Exact distributions of the skewness and kurtosis based test statistics

Let us assume throughout this section that the sample vector follows a multivariate Gaussian distribution. According to this assumption, the derivation of the exact distributions of the skewness and kurtosis based test statistics presented here is based upon an application of the following geometric measure formula in [33]

$$\Phi(A) = \frac{2^{1-n/2}}{\Gamma(n/2)} \int_0^\infty \tilde{\Phi}(A,r)r e^{-r^2/2}dr \quad (17)$$

where the function

$$\tilde{\Phi}(A,r) = \frac{AL([\frac{1}{2}A] \cap C)}{AL(C)}, r > 0$$
is called the Euclidean intersection proportion function of the set $A$. Here, $C$ denotes the Euclidean unit circle and $AL$ means Euclidean arc-length. Formula (17) has been proved for the special case of Gaussian distribution in [32] and was used later on in a series of subsequent papers.

If $n = 2, A = \tilde{C}_1(t)$ then

$$P(\tilde{B}_1 < t) = \int_0^\infty \mathfrak{P}(\tilde{C}_1(t), r) r e^{-r^2/2} dr. \quad (18)$$

We recall that, for arbitrary $r > 0$, $\tilde{C}_1(t)$ is a cone. Thus, the ipf of $\tilde{C}_1(t)$ does not depend on the radius $r$. In consequence,

$$P(\tilde{B}_1 < t) = \mathfrak{P}(\tilde{C}_1(t), 1) = \alpha_1(t)/(2\pi) \quad \text{and} \quad P(\sqrt{B}_2 < t) = \mathfrak{P}(\tilde{C}_2(t), 1) = \alpha_2(t)/(2\pi) \quad (19)$$

where $\alpha_1$ and $\alpha_2$ are chosen according to (15) with (13) and (16), respectively.

### 7. Representing the distribution of $\tilde{B}_1$ in terms of a Student distribution

As we have remarked in Section 5, the distribution of $\tilde{B}_1$ may be represented in terms of Student’s distribution. For seeing that, we transform the representation of the cone into a representation which was used in [34]. Because the transformations $\mathfrak{T}_1, \mathfrak{T}_2$ where

$$\mathfrak{T}_1^{-1} = \frac{1}{\sqrt{2}} \begin{pmatrix} 1 & 1 \\ -1 & 1 \end{pmatrix} \quad \text{and} \quad \mathfrak{T}_2 = \begin{pmatrix} 1 & 0 \\ 0 & -1 \end{pmatrix}$$

do not change the spherical measure of a set, it holds true that $\Phi(A) = \Phi(\mathfrak{T}_1 \mathfrak{T}_2 A)$.

We note that

$$C_1^0(t) = \mathfrak{T}_1 \mathfrak{T}_2 \tilde{C}_1(t) = \{ (\psi, \nu)^T \in \mathbb{R}^2 : \frac{|\psi|}{|\psi + \nu|} < \delta(t) \} \quad \text{where} \quad \delta(t) = \sqrt{a(a - 1)(t/\beta)^{1/6}}.$$

Thus

$$\Phi(C_1^0(t)) = 2\Phi(\{ (\psi, \nu)^T \in \mathbb{R}^2 : \frac{\psi}{\nu} < \frac{\delta(t)}{1 - \delta(t)} \}).$$

Note that $\Phi(\{ (\psi, \nu)^T \in \mathbb{R}^2 : \frac{\psi}{\nu} < \frac{\delta(t)}{1 - \delta(t)} \}) = F_1(\frac{\delta(t)}{1 - \delta(t)}) \quad \text{where} \quad F_1$ denotes the cdf of the Student cdf with one dgf.

### 8. Truncated data

In Section 6, we assumed that the sample vector $X$ follows the multivariate standard Gaussian law. Here, we assume that the distribution law of $X$ is a multivariate truncated Gaussian law having the pdf

$$f(x) = \frac{1}{\beta^n} (\prod_{i=1}^{n} \varphi(x_i)) I_B(x)$$
where \( B \) is a truncation set, \( \beta^* = \Phi(B) \), and \( I_B \) denotes the indicator function of the set \( B \).

In the special case that \( B = [a, b)^n \) for some \( a < b \) this means that \( X \) has i.i.d. marginal variables following the univariate truncated normal density \( h(x) = \frac{\varphi(x)}{Z} \) where \( Z = \Phi(b) - \Phi(a) \).

In the general truncation case it follows for the cdf of the test statistic \( \tilde{B}_1(t) \) that

\[
P(\tilde{B}_1 < t) = \frac{1}{\beta^*} \int_0^\infty \tilde{g}(\tilde{C}_1(t) \cap B, r)re^{-r^2/2}dr, \tag{20}
\]

Note that

\[
\tilde{g}(\tilde{C}_1(t) \cap B, r) = \begin{cases} 
F(\tilde{C}_1(t), r) & \text{if } 0 < r \leq b \\
F(\tilde{C}_1(t), r) & \text{if } b < r \leq \sqrt{b^2 + a^2} \\
0 & \text{if } \sqrt{b^2 + a^2} < r
\end{cases}
\]

Thus the distribution of \( \tilde{B}_1 \) is as in Section 6.

In the special case that \( B \) is an Euclidean ball of radius \( \rho \) having its center at the origin it follows that \( \tilde{g}(\tilde{C}_1(t) \cap B, r) = F(\tilde{C}_1(t), r) \), \( r > 0 \). Thus the distribution of \( \tilde{B}_1 \) differs from that in Section 6.

### 9. Exact distributions under non-standard model assumptions

A well known more general assumption w.r.t. the distribution the sample vector follows is that to assume it follows a spherical distribution. Recognize that a yet more general assumption is to let the sample vector follow an \( l_{n,p} \)-symmetric distribution.

A general method of deriving exact distributions of statistics under non-standard model assumptions is outlined in [39]. Here, we follow the approach described there. Let the random vector \((X_1, X_2)^T\) follow the \( l_{2,p} \)-symmetric density

\[
\varphi_{g,p}(x) = C_{g,p}g(|x|^p), x \in \mathbb{R}^2, p > 0
\]

where \(|x|^p = (|x_1|^p + |x_2|^p)^{1/p}\) is a norm if \( p \geq 1 \) and an antinorm (see [28]) if \( 0 < p \leq 1 \), \( g : [0, \infty) \rightarrow [0, \infty) \) is a density generating function (dgf) which satisfies \( 0 < I_{g,p} < \infty \) for \( I_{g,p} = \int_0^\infty rg(r^p)dr \), and \( C_{g,p} \) is a suitable normalizing constant making \( \varphi_{g,p} \) a density. If \( \Phi_{g,p} \) denotes the probability law corresponding to the density \( \varphi_{g,p} \),

\[
\Phi_{g,p}(A) = \int_A \varphi_{g,p}(x)dx, A \text{ is a measurable subset of } \mathbb{R}^2,
\]

then formula(17) is a special case of the following geometric measure representation

\[
\Phi_{g,p}(A) = \frac{1}{I_{g,p}} \int_0^\infty \tilde{g}(A, r)rg(r^p)dr \tag{21}
\]
which was proved for the \( n \)-dimensional case in [38]. Note that

\[ I_{g,p} = \frac{1}{2\pi(p)C_{g,p}} \]

where the constants \( \pi(p) \) are generalizations of the famous circle number \( \pi \) which have been introduced and studied for convex and radially concave \( p \)-generalized circles (or \( l_{2,p} \)-circles) in [36] and [37], respectively.

The function \( x \to |x|_p \) is the Euclidean norm \( |x|_2 = \sqrt{x_1^2 + x_2^2} \) if \( p = 2 \), the taxi cap norm \( |x|_1 = |x_1| + |x_2| \) if \( p = 1 \), and, formally, the maximum norm \( |x| = \max\{|x_1|, |x_2|\} \) if \( p = \infty \).

The \( p \)-circles \( C_p(r) = rC_p, C_p = \{ x \in \mathbb{R}^2 : |x|_p = r \} \) are the level sets of the density of the random vector \((X_1, X_2)^T\). The function \( r \to \Phi_p(A, r) \) is a generalization of the ipf \( r \to \Phi_p(A, r) \) considered in Section 6. It is defined as

\[ \Phi_p(A, r) = \frac{\Omega_p(\{\frac{1}{r}A\} \cap C_p)}{2\pi(p)} \]

where the \( p \)-generalized arc-length measure \( \Omega_p \) has been considered in [36, 37]. In the easiest non-Euclidean case, i.e. if \( p = 1 \), it holds

\[ \Phi_1(A, r) = \frac{AL(\{\frac{1}{r}A\} \cap C_p)}{2\pi}. \]

We restrict our consideration therefore here as in [24] to the two cases \( p = 2 \) and \( p = 1 \).

Special density generating functions have been considered, e.g. in [24]. If \( g(r) = r^M e^{-\beta r^\gamma} \) with positive \( M, \beta, \gamma \) then \( \Phi_{g,1} = \frac{\gamma^\frac{M+1}{\gamma}}{\Gamma(M+1)} \), and if \( g(r^2) = r^{2M-2} e^{-\beta r^{2\gamma}} \) with positive \( M, \beta, \gamma \) then \( \Phi_{g,2} = \frac{\gamma^\frac{M}{\gamma} \Gamma(M/\gamma)}{\Gamma(M+1/\gamma)} \). Note, however, that in the given cases of test statistics \( \tilde{B}_1 \) and \( \sqrt{B}_2 \), the distributions do not depend of the dgf. This is due to the following much more general fact. If the ipf of a set \( A \) does not depend on the radius variable \( r \), \( \Phi_p(A, r) = \Phi_p(0) \), then it follows from (21) that \( \Phi_{g,p}(A) = \Phi_{p,0} \).

It is shown in Section 5, that the ipfs of the sets \( \tilde{C}_i(t), i = 1, 2 \) generated by the statistics \( \tilde{B}_1 \) and \( \sqrt{B}_2 \), respectively, do not depend on the variable \( r \). Thus the distributions of the statistics \( \tilde{B}_1 \) and \( \sqrt{B}_2 \) do not depend on the dgf, and are therefore the same as given in Section 6. Following a definition in [11], such statistics are called \( g \)-robust in [23], and in a more general setting in [39].

10. Exact joint two-dimensional distribution of the skewness and kurtosis based test statistics

Let us assume that the sample vector follows an \( n \)-dimensional spherical distribution \( \Phi_{g,p} \). According to the previous results,

\[ F(t_1, t_2) = P(\tilde{B}_1 < t_1, \sqrt{B}_2 < t_2) = \Phi_{g,p}(\tilde{C}_1(t_1) \cap \tilde{C}_2(t_2)) \]
Note that the intersection of two cones is a cone. The density of \( \left( \frac{\tilde{b}_1}{\sqrt{b_2}} \right) \) is

\[
f(t_1, t_2) = \frac{\partial^2}{\partial t_1 \partial t_2} F(t_1, t_2). \tag{22}
\]

On using this density, we are going now to derive the density of \( \tilde{B}_1 + \sqrt{\tilde{B}_2} \). To this end, let a map

\[
\varphi = \begin{pmatrix} \varphi_1 \\ \varphi_2 \end{pmatrix} : \mathbb{R}^- \times \mathbb{R}^+ \to \mathbb{R}^2
\]

be defined by

\[
\begin{pmatrix} \varphi_1 \left( \frac{\tilde{B}_1}{\sqrt{B_2}} \right) \\ \varphi_2 \left( \frac{\tilde{B}_1}{\sqrt{B_2}} \right) \end{pmatrix} = \begin{pmatrix} Y_1 \\ Y_2 \end{pmatrix} = \begin{pmatrix} \tilde{B}_1 + \sqrt{B_2} \\ \tilde{B}_1 \end{pmatrix}.
\]

Note that \( \varphi \) is \( 1-1 \) with inverse function \( \psi = \varphi^{-1} \) given by

\[
\begin{pmatrix} \tilde{B}_1 \\ \sqrt{B_2} \end{pmatrix} = \begin{pmatrix} Y_2 \\ Y_1 - Y_2 \end{pmatrix} = \begin{pmatrix} \psi_1(Y_1, Y_2) \\ \psi_2(Y_1, Y_2) \end{pmatrix}.
\]

The absolute value of the Jacobian of this transformation is

\[
I = \left| \det \begin{bmatrix} \frac{\partial \psi_1}{\partial y_1} & \frac{\partial \psi_1}{\partial y_2} \\ \frac{\partial \psi_2}{\partial y_1} & \frac{\partial \psi_2}{\partial y_2} \end{bmatrix} \right| = \left| \det \begin{bmatrix} 0 & 1 \\ 1 & -1 \end{bmatrix} \right| = 1.
\]

As a consequence, the density of \( \left( \frac{\tilde{B}_1 + \sqrt{B_2}}{\tilde{B}_1} \right) \) is

\[
f \left( \frac{\tilde{B}_1 + \sqrt{B_2}}{\tilde{B}_1} \right)(y_1, y_2) = f(\psi_1(y_1, y_2), \psi_2(y_1, y_2)) I = f(y_2, y_1 - y_2)
\]

where \( f \) is chosen according to (22). Integrating this joint density w.r.t. the variable \( y_2 \) results in the desired density

\[
f_{\tilde{B}_1 + \sqrt{B_2}}(y_1) = \int_{-\infty}^{\infty} f(y_2, y_1 - y_2) dy_2. \tag{23}
\]

11. Tests based upon higher order moments

In this section we study testing based upon higher order moments. Let
\[ \hat{b}_3 = \frac{1}{n} \sum_{i=1}^{n} (x_i - \tilde{x}_z)^6 \]

be the statistic considered in \([26]\).

If \( n = 2 \), \( \tilde{x}_z = \Theta x + (1 - \Theta)y \), \( \tilde{x}_N = x + ay \) then

\[ \hat{B}_3 = 4 \left[ \frac{(1 - \Theta)^6 - \Theta^6}{a^2 z^2 + ((1 - a)z - 1)^2} \right]^{3/2} z = \frac{y}{x}. \]

For every \( t > 0 \), the set

\[ \tilde{C}_3(t) = \{(x, y) \in \mathbb{R}^2 : \hat{b}_3 < t \} \]

is a cone with vertex in the origin and boundary

\[ \partial \tilde{C}_3(t) = \{(x, y) \in \mathbb{R}^2 : [1 - \Theta)^6 - \Theta^6]^{1/3} (1 - Z)^2 = a^2 z^2 + ((1 - a)z - 1)^2 \}. \]

Because the \( \Phi_{g,p} \)-value of such cones does not depend on the dfg \( g \), the statistic \( \hat{b}_3 \) is \( g \)-robust just like \( \hat{b}_1 \), \( \hat{b}_2 \) and \( \sqrt{\hat{b}_2} \) are \( g \)-robust. Note that the boundary lines of \( \tilde{C}_3(t) \) satisfy the equations \((14)\) but where now \( p \) and \( q \) are

\[ p = \frac{-2C(\theta) + 2(1 - a)t}{C(\theta) - a^2 t - (1 - a)^2 t} \quad \text{and} \quad q = \frac{C(\theta) - t}{C(\theta) - a^2 t - (1 - a)^2 t} \quad \text{with} \quad C(\theta) = ((1 - \theta)^6 - \theta^6)^{1/3}. \]

We conclude this section with the following remark on a non-\( g \)-robust statistic. Let

\[ \hat{b}_k = \frac{1}{2} \left[ \frac{(x - \frac{x+y}{2})^k + (y - \frac{x+y}{2})^k}{\left( \frac{1}{2} \left( (x - \frac{x+y}{2})^2 + (y - \frac{x+y}{2})^2 \right) \right)^{k/2}} \right] = \begin{cases} \frac{|x-y|}{2^{k/2+1}}, & k \text{ is even} \\ 0, & k \text{ is odd} \end{cases} \]

Then, for even \( k \), the critical test region is

\[ H_k(t) = \{(x, y) \in \mathbb{R}^2 : \left| \frac{x-y}{2} \right|^{k/2} > t \} \]

Denoting the acceptance region by \( H_k^c(t) \), it turns out that \( \Phi_{g,p}(H_k^c(t)) \) depends on \( g \) and \( p \), thus \( b_k \) is not \( g \)-robust.

12. Algebraic manipulations on main quadric

Throughout, we denote the set of real numbers by \( \mathbb{R} \). As we can see from simulations, the power depends on choice of location functionals in both numerator and denominator of \( RT \) class tests. We conducted large simulation study for \( n = 2 \) and location functional in nominator given by weighted mean, i.e. \( m(X_1, X_2) = \theta X_1 + (1 - \theta)X_2, \theta \in \)
According to [31] weighted mean is an interesting class of location functionals. However, for \( n = 2 \) in order not to get constant skewness, we chose estimator \( m \ast (X_1, X_2) = X_1 + aX_2, \ a \in (0, 1) \). Notice that \( m \ast \) is biased for \( \mu \neq 0 \), \( X_1, X_2 \sim N(\mu, \sigma^2) \). However it is an unbiased estimator of \( \mu \) when \( X_1, X_2 \sim N(0, \sigma^2) \).

We can relate it here to concept of superefficiency (see [8], page 515 and references therein). This directly relates to Hodges’ estimator (see e.g. [54]). Of course, measure of point \( \mu = 0 \) is zero, interestingly, the rejection cone is also degenerated (in real geometry). Thus, we have one degenerated and one non-degenerated cone, as it is proven in the next lines.

Large simulation study shows that most powerful test was given by \( \theta_1 = 1/3 \) and \( a \in \{1/2, 1\} \). After some algebra for \( \tilde{B}_1 \) we receive the initial equation of the form

\[
28(x^3 - 3x^2y + 3xy^2 - y^3) = 27a(y^2 - 2xy + 2x^2)\sqrt{2y^2 - 4xy + 4x^2} \tag{25}
\]

We are interested in the set of real solutions to this equation. Below, we show that for \( a \) in the interior of the interval \( I = \left[0, \frac{14\sqrt{3}}{27}\right] \) this solution set is two lines passing through the origin and it gives a degenerate answer, i.e. multiple line, on the boundary of \( I \). Moreover, there are no real solutions for \( a \) taken outside \( I \). For simplicity we put \( f = y^2 - 2xy + 2x^2 \). Taking the square of (25) we get

\[
(x - y)^6 = 2\left(\frac{27a}{28}\right)^2 f^3 \tag{26}
\]

which factors through

\[
((x - y)^2 - \alpha f)((x - y)^4 + \alpha f(x - y)^2 + \alpha^2 f^2) = 0 \tag{27}
\]

where

\[
\alpha = \sqrt{\frac{3}{2}\left(\frac{27a}{28}\right)^2}
\]

**Lemma 1.** The equation \((x - y)^4 + \alpha f(x - y)^2 + \alpha^2 f^2 = 0\) has no nontrivial solution in \( \mathbb{R} \) for any nonzero value of \( \alpha \).

**Proof.** Note that if we substitute \( z = (x - y)^2 \) and \( t = x^2 \) in the equation we have

\[
z^2 + \alpha(z + t)z + \alpha^2(z + t)^2 = 0
\]

Obviously \( x = y = 0 \), and consequently \( z = t = 0 \), satisfies the equation. Assume, without loss of generality, that \( z \neq 0 \), and substitute \( r = \frac{z + t}{z} \). The aim is to find real solutions to the quadratic equation

\[
\alpha^2r^2 + \alpha r + 1 = 0
\]

The discriminant of this equation is \( \Delta = -3\alpha^2 \), which is never positive, and is zero if and only if \( \alpha = 0 \). Clearly if \( \alpha = 0 \), the solution set is the line \( x = y \).

**Lemma 2.** The real solutions to the quadratic form \( \{(x - y)^2 - \alpha(y^2 - 2xy + 2x^2) = 0\} \) for different values of \( \alpha \) is
(1) two lines $y = (1 + \sqrt{\frac{\alpha}{1-\alpha}})x$ and $y = (1 - \sqrt{\frac{\alpha}{1-\alpha}})x$, when $\alpha \in (0, 1)$,
(2) the line $x = y$, when $\alpha = 0$,
(3) the line $x = 0$, when $\alpha = 1$,
(4) and it has no nontrivial real solution when $\alpha \notin [0, 1]$.

**Proof.** For simplicity, consider the substitution $z = x - y$. The equation now becomes

$$z^2 - \alpha(z^2 + x^2) = 0$$

which solves to

$$z = \pm \sqrt{\frac{\alpha}{1-\alpha}}x$$

The proof follows from this equality. $\square$

**Corollary 1.** The real solutions to (25) depending on the value of the real parameter $a$ are

(1) two lines $y = (1 + \sqrt{\frac{\sqrt{2}(\frac{27a^2}{28})}{1-\sqrt{2}(\frac{27a^2}{28})})}x$ and $y = (1 - \sqrt{\frac{\sqrt{2}(\frac{27a^2}{28})}{1-\sqrt{2}(\frac{27a^2}{28})}})x$, when $a \in I$,
(2) the line $x = y$, when $a = 0$,
(3) the line $x = 0$, when $a = \frac{14\sqrt{2}}{27}$,
(4) and it has no nontrivial real solution when $\alpha \notin I$.

Note that when we solve $a$ for $\alpha$ for substitution we do not consider the negative values of $a$. This is because we took the square of the equation (25) in the beginning.

### 13. Power comparisons – models for outliers

In this section we compare the power of the classical tests for normality and our omnibus test for normality. We have checked by conducting the thorough simulation study that all considered empirical tests hold the size for $\alpha = 0.05$ and for sample sizes $n = 20$ and $n = 100$. The alternatives we have considered are models for outliers.

Here we consider two outlier models, as used in [2], i.e. location-outlier model and scale-outlier model.

- For the $p$-location-outlier model we consider $X_1, \ldots, X_{n-p}$ to be iid from $N(0,1)$ and $X_{n-p+1}, \ldots, X_n$ to be iid from $N(\lambda, 1)$.
- For the $p$-scale-outlier model we consider $X_1, \ldots, X_{n-p}$ to be iid from $N(0,1)$ and $X_{n-p+1}, \ldots, X_n$ to be iid from $N(0, \tau^2)$.

In this paper we consider only linear location estimators. In particular, median provides best protection against the presence of outlier in term of bias ([2]), but it comes at the cost of a higher MSE than some other estimators. The trimmed mean, linearly weighted mean and modified MLE turn out to be quite robust and efficient in general.

In this paper we consider only median $T_{(1)}$, trimmed mean $T_{(2)}$ and pseudomedian $T_{(3)}$.

The results of our Monte Carlo simulations are summarized in Table 2. As it is shown in mentioned tables, the most commonly used tests of normality, such as the $SW$ and
$JB$ tests, are too strict in rejecting normality in the case of a small number of outliers, even when the sample size is large enough. To illustrate this general framework, notice that one of the most used $SW$ test rejects the hypothesis of normality in 81% cases if we suppose the $p$-location outlier model for $p = 1$, $\lambda = 5$ and sample size $n = 100$. Similarly, the $JB$ and $RJB$ tests are too strict in rejecting normality. If we suppose the same $p$-location outlier model and sample size, the $JB$ and $RJB$ tests reject the hypothesis of normality in 88% and 86% cases, respectively. In contrast, the $TTRT$ and $MC_{LR}$ tests are more robust – $TTRT$ test rejects normality in 29%, and $MC_{LR}$ test even in 5%, which is consistent with chosen significance level.

Now, if we suppose a largest number of outliers in small sample size – e.g. the $p$-location outlier model for $p = 5$, $\lambda = 5$ and sample size $n = 20$ – we need the test with high power, because the hypothesis of normality is not sustainable. Based on our Monte Carlo simulations we can recommend the $AD$, $SW$ and again $TTRT$ tests – these test reject the hypothesis of normality in 92%, 89% and 85% cases, respectively. In contrast, the $JB$, $JBU$ and $RJB$ tests have small power against the mentioned $p$-location outlier model – the classical $JB$ test reject the hypothesis of normality only in approximately 12% cases, the $RJB$ test in 14% cases and finally the $JBU$ reject the hypothesis of normality only in approximately 2% cases.

Based on these simulations we can conclude that especially the $TTRT$ test is more robust against many different types of small numbers of outliers compared to sample size, while the $SW$, $JB$ and $RJB$ tests generally are non-robust. In other words, the $TTRT$ test is more robust in case of small number of outliers than the classical normality tests such as $SW$ and $JB$, and simultaneously have comparable power with the $SW$ test in case of presence of large number of outliers compared to sample size.

### Table 2. Power of the selected tests for normality against outliers models for $n = 20$ and $n = 100$

<table>
<thead>
<tr>
<th>Test</th>
<th>$n = 20$</th>
<th>$n = 100$</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$p = 1$</td>
<td>$p = 5$</td>
</tr>
<tr>
<td></td>
<td>$\lambda = 3$</td>
<td>$\lambda = 5$</td>
</tr>
<tr>
<td>$AD$</td>
<td>0.173</td>
<td>0.630</td>
</tr>
<tr>
<td>$DT$</td>
<td>0.305</td>
<td>0.866</td>
</tr>
<tr>
<td>$JB$</td>
<td>0.314</td>
<td>0.872</td>
</tr>
<tr>
<td>$JBU$</td>
<td>0.319</td>
<td>0.878</td>
</tr>
<tr>
<td>$LT$</td>
<td>0.118</td>
<td>0.407</td>
</tr>
<tr>
<td>$RJB$</td>
<td>0.298</td>
<td>0.850</td>
</tr>
<tr>
<td>$SJ_{dir}$</td>
<td>0.234</td>
<td>0.728</td>
</tr>
<tr>
<td>$SW$</td>
<td>0.226</td>
<td>0.766</td>
</tr>
<tr>
<td>$MC_{LR}$</td>
<td>0.051</td>
<td>0.055</td>
</tr>
<tr>
<td>$MMRT1$</td>
<td>0.273</td>
<td>0.832</td>
</tr>
<tr>
<td>$MMRT2$</td>
<td>0.266</td>
<td>0.829</td>
</tr>
<tr>
<td>$TTRT1$</td>
<td>0.038</td>
<td>0.007</td>
</tr>
<tr>
<td>$TTRT2$</td>
<td>0.275</td>
<td>0.822</td>
</tr>
</tbody>
</table>

14. Discussion and conclusions

This paper introduces the general $RT$ class of robust tests for normality and discuss their properties, especially geometric ones. The further theoretical considerations of class $RT$ will be of interest. In the simulation study we have focused on the power...
study of selected tests from RT class. We have compared these tests with the selected
tests for the normality on the large scale of alternatives. Some of the most important
results (observations) are:

- The RT class includes a large number of tests with different properties, especially
power and robustness.
- Some RT class tests (e.g. the classical JB test, RJB test, SJ test among others)
have already been studied in the literature – these tests are special cases of RT
class tests with/without trimming for the different settings of parameters of the
general RT class test statistic.
- Based on our pilot simulation study we recommend four tests with good prop-
erties for the general use – MMRT1, MMRT2, TTRT1 and TTRT2 tests.
- TTRT1 test is more robust against many different types of small numbers of
outliers.
- Many results of this paper are in the coherence or are extending the results of
previous studies, e.g. [49].

Acknowledgements

Milan Stehlík acknowledges the support of Fondecyt Proyecto Regular No. 1151441.
Luboš Střelec acknowledges the support of the Grant Agency of the Czech Republic
(project No. 16-07089S).

References

seSSing when a sample is mostly normal. Computational Statistics & Data Analysis, Vol.
54, No. 12, pp. 2914–2925.
independence of regression residuals: Monte Carlo evidence. Economics Letters Vol. 7,
No. 4, pp. 313–318
vanced Series, Thomson Learning.
results for the distribution of $\sqrt{b_1}$ and $b_2$. Biometrika, Vol. 60, No. 3, pp. 613–622.


