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The `robustTest` package: two-sample tests revisited

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Abstract

The R package `robustTest` offers corrected versions of several common tests in bivariate statistics. We point out the limitations of these tests in their classical versions, some of which are well known such as robustness or calibration problems, and provide simple alternatives that can be easily used instead. The classical tests and their robust alternatives are compared through a small simulation study. The latter emphasizes the superiority of robust versions of the test of interest. Finally, an illustration of correlation's tests on a real data set is also provided.

1 Introduction

In this article, we consider several usual tests in bivariate statistics, which are taught in many scientific courses at various levels.

Our first goal is to illustrate, with the help of mathematical considerations and simulations, that many of these tests are either not very robust (i.e. they do not work outside the very strict framework in which they have been defined), or badly calibrated (i.e. we can find simple examples for which the null hypothesis is true, but the type I error rate is not the one announced).

Of course, we are not the only ones or the first ones to have noticed this, and in some cases valid solutions have been proposed (see for example the famous article by Welch [16] about the test of equality of two expectations in the case where the variances are unequal, the Welch test being (asymptotically) robust to non-normality).

Nevertheless, we felt it was important to return to these issues for at least two reasons:

- First, these tests are not only mathematical objects on which students can discover the basic principles of statistical analysis, they are also very often used in practice (for example in biomedical research articles). It is therefore important to come back to the limits of these tests, which are not always well indicated (especially in the case of so-called "non-parametric" tests).

- Secondly, because it is often very easy to modify these tests in order to make them more robust or (asymptotically) well calibrated. The modification presented is each time based on the calculation of the limiting variance of the test statistic under the null hypothesis. Renormalizing by an estimator of the standard deviation, we obtain a robust or asymptotically well-calibrated version thanks to the central limit theorem. Of course, the justification is asymptotic, but we will illustrate that, on simulated examples, it is always interesting to correct these tests, even for relatively small sample sizes. It seems to us that these modified tests, which are easy to describe and to implement, should be systematically pointed out to the students, and also to researchers from other disciplines unfamiliar with these issues.

Our second objective is to present the `robustTest` package, which implements robust modifications of the usual bivariate statistics tests. As we will see, the functions of the `robustTest`

package, as well as their syntax, are very close to the functions of the *stats* package, so that regular users of the R software can use and compare them easily. This package is already used at Université Paris Cité and Université d'Evry Val d'Essonne to illustrate second or third year postgraduate courses on bivariate statistics.

The article is organised as follows: in Section 2, we present the robust versions of the Pearson (see [8]), Kendall (see [12]) and Spearman (see [10]) correlation tests. For completeness, we also provide in the same section the Kolmogorov-Smirnov type test of independence for continuous variables (see [3]). In Section 3, we describe the test of equality of conditional variances of X knowing Y , when Y is a categorical variable with several levels. This procedure is based on the James-Welch ANOVA which we briefly recall (see [11] and [17]). In Section 4, we present the robust version of the two sample Mann-Whitney test for stochastic dominance (see [14]). In Section 5, we describes several ways to test the stochastic dominance in the case of paired two samples (see [15]). At the end of each section or subsection, we provide the R-functions of the *robustTest* package for each corrected test. Finally, in the last section, we illustrate on a real data set the functions of the *robustTest* package to test the correlation, and we compare the outputs to those of the usual tests.

2 Robust tests for testing correlation and independence

In the following four subsections we respectively focus on the correlation tests of Pearson, Kendall and Spearman, and on the independence test of Kolmogorov-Smirnov type.

2.1 Pearson's correlation test

Let $(X_1, Y_1), \dots, (X_n, Y_n)$ be independent and identically distributed (i.i.d) copies of the random pair (X, Y) , where X, Y are real-valued random variables. We suppose that X_i and Y_i have finite moments of order 2 and we denote $\mathbb{E}(X_i) = \mu_X$ and $\mathbb{E}(Y_i) = \mu_Y$. Our aim is to test

the null hypothesis $H_0 : \rho = 0$ against the alternative hypothesis $H_1 : \rho \neq 0$

where ρ is the correlation coefficient between X_i and Y_i . To this end we often use the Pearson test statistic (see for instance Fisher [8]):

$$T_n = \frac{\hat{\rho}_n}{\sqrt{\frac{1-\hat{\rho}_n^2}{n-2}}} \quad \text{where} \quad \hat{\rho}_n = \frac{\sum_{k=1}^n (X_k - \bar{X}_n)(Y_k - \bar{Y}_n)}{\sqrt{\sum_{k=1}^n (Y_k - \bar{Y}_n)^2} \sqrt{\sum_{k=1}^n (X_k - \bar{X}_n)^2}}$$

is the empirical correlation coefficient and where \bar{X}_n, \bar{Y}_n are the empirical means.

The interest of this statistic being that, in the case where the couple (X_i, Y_i) is Gaussian, its exact distribution under H_0 is known: it is the Student $St(n-2)$ -distribution. If the pair (X_i, Y_i) is not Gaussian, we can also easily show that, if X_i is independent of Y_i (and if X_i and Y_i are not constants), then T_n converges in distribution to the standard normal distribution $\mathcal{N}(0, 1)$. Therefore, if X_i and Y_i are independent, the rejection region $R_{n,\alpha} = \{|T_n| > c_\alpha\}$ where c_α is the quantile of order $1 - (\alpha/2)$ of the $\mathcal{N}(0, 1)$ distribution, is such that its probability tends to α as $n \rightarrow \infty$. But this is not the case in general under H_0 , which of course does not imply independence of the variables.

We can therefore make the following conclusion: in a general context, the Pearson test is not well calibrated to test the null hypothesis $\rho = 0$. See also [7] for a similar observation. Outside the strict framework of the linear model (Gaussian or not), it is therefore preferable to use the intuitive statistic

$$T'_n = \frac{\sum_{k=1}^n (X_k - \bar{X}_n)(Y_k - \bar{Y}_n)}{\sqrt{\sum_{k=1}^n (Z_k - \bar{Z}_n)^2}} \quad \text{where} \quad Z_i = (X_i - \bar{X}_n)(Y_i - \bar{Y}_n).$$

Under the assumption $0 < \mathbb{E}((X_i - \mu_X)^2(Y_i - \mu_Y)^2) < \infty$, a direct application of the central limit theorem and Slutsky's lemma provides the convergence in distribution of T'_n under H_0 to the standard normal distribution. The rejection region of H_0 is $R'_{n,\alpha} = \{|T'_n| > c_\alpha\}$ where c_α is the quantile of order $1 - (\alpha/2)$ of the $\mathcal{N}(0, 1)$ distribution, which provides a test asymptotically well calibrated.

Remark. It is easy to see that, in the Gaussian case and under H_0 , the distribution of the statistic T'_n does not depend on the expectation and variance parameters, and can therefore be tabulated. In practice, in the *robustTest* package, we use this table of quantiles for $n < 130$ and the quantiles of the Student $St(n - 2)$ distribution for $n \geq 130$ (simulations show that the quantiles of the Student $St(n - 2)$ distribution are close to those of the distribution of the statistic T'_n under H_0 in the Gaussian case).

The corresponding R function of the *robustTest* package is: `cortest(,method="pearson")`.

Remark. Starting from the quantity

$$\frac{\sum_{k=1}^n (X_k - \bar{X}_n)(Y_k - \bar{Y}_n) - n\text{Cov}(X, Y)}{\sqrt{\sum_{k=1}^n (Z_k - \bar{Z}_n)^2}},$$

which converges to the $\mathcal{N}(0, 1)$ distribution as $n \rightarrow \infty$, we can easily obtain a confidence interval for $\text{Cov}(X, Y)$, the covariance between X and Y . As an output of the command `cortest(,method="pearson")`, we propose instead a confidence interval for Pearson's ρ coefficient. This confidence interval is based on the central limit theorem for the empirical estimator of $(\text{Cov}(X, Y), \text{Var}(X), \text{Var}(Y))$ and the delta method applied to the function $h(x, y, z) = x/\sqrt{yz}$ from \mathbb{R}^3 to \mathbb{R} . There may be a difference between the result based on the robust Pearson test and whether or not 0 is in the confidence interval of ρ . In this case, the test should be preferred since the confidence interval based on the central limit theorem + the delta method is a priori less precise.

2.2 Kendall's correlation test

The context is the same as in the previous paragraph $(X_1, Y_1), \dots, (X_n, Y_n)$ are i.i.d. copies of the pair (X, Y) , where X, Y are real-valued random variables. We assume moreover that the variables are continuous. We want to know if X_i and Y_i tend to vary in the same direction or in the opposite direction. Let then

$$\tau = 2(\mathbb{P}((X_2 - X_1)(Y_2 - Y_1) > 0) - 0.5)$$

be Kendall's correlation coefficient [12]. As for Pearson's correlation coefficient, τ is between -1 and 1 . X_i and Y_i are positively correlated in the sense of Kendall when $\tau > 0$ and negatively if $\tau < 0$. To test

$$H_0 : \tau = 0 \quad \text{against} \quad H_1 : \tau \neq 0$$

Kendall [12] therefore proposed to count the number of concordant pairs (i.e. for which the product $(X_i - X_j)(Y_i - Y_j)$ is strictly positive), which leads to the statistic

$$T_n = \frac{1}{n(n-1)} \sum_{i=1}^n \sum_{j=1, j \neq i}^n (\mathbf{1}_{(X_i - X_j)(Y_i - Y_j) > 0} - 0.5).$$

It is quite easy to see that, if X_i and Y_i are independent, then the distribution of T_n is distribution-free (i.e. does not depend on the distribution of (X_i, Y_i)), by reducing to two independent sequences U_1, \dots, U_n and V_1, \dots, V_n of i.i.d. random variables with standard uniform distribution $\mathcal{U}([0, 1])$. Consequently, if X_i and Y_i are independent, T_n is distributed according

to a known and tabulated distribution. Kendall's test is constructed from the quantiles of this distribution. However if X_i and Y_i are not independent, the distribution of the statistic T_n has no reason to be distribution free under H_0 (it depends on the joint distribution of (X_i, Y_i)). We can therefore make the following conclusion: in a general context, Kendall's test is not well calibrated to test $\tau = 0$.

We can nevertheless solve (asymptotically) this problem, by considering the limiting distribution of $\sqrt{n}T_n$ under H_0 . Starting from the Hoeffding decomposition of the U -statistic T_n (see [9] or [15], example 12.5), we see that, under H_0 , $\sqrt{n}T_n$ converges in distribution to the $\mathcal{N}(0, V)$ distribution, with

$$V = 4\text{Var}(F(X_i, Y_i) + H(X_i, Y_i)),$$

where $F(x, y) = \mathbb{P}(X_i < x, Y_i < y)$ and $H(x, y) = \mathbb{P}(X_i > x, Y_i > y)$. The empirical estimator of V is then

$$V_n = \frac{4}{n-1} \sum_{k=1}^n (F_n(X_k, Y_k) + H_n(X_k, Y_k) - \bar{F}_n - \bar{H}_n)^2$$

where

$$F_n(x, y) = \frac{1}{n} \sum_{k=1}^n \mathbf{1}_{X_k < x, Y_k < y}, \quad H_n(x, y) = \frac{1}{n} \sum_{k=1}^n \mathbf{1}_{X_k > x, Y_k > y}$$

$$\bar{F}_n = \frac{1}{n} \sum_{k=1}^n F_n(X_k, Y_k), \quad \bar{H}_n = \frac{1}{n} \sum_{k=1}^n H_n(X_k, Y_k).$$

Finally, under H_0 ,

$$K_n = \frac{\sqrt{n}T_n}{\sqrt{V_n}} \text{ converges in distribution to the } \mathcal{N}(0, 1) \text{ distribution.}$$

The rejection region of the corrected Kendall test is then $R_{n,\alpha} = \{|K_n| > c_\alpha\}$ where c_α is the quantile $1 - (\alpha/2)$ of the $\mathcal{N}(0, 1)$ distribution, which provides a test asymptotically well calibrated (see [1] for more details).

The corresponding R function of the *robustest* package is: `cortest(,method="kendall")`.

2.3 Spearman's correlation test

In the same context as for the Kendal test statistic of the previous paragraph, we can also correct Spearman's correlation test (as described by Hotelling and Pabst [10]), which tests

$$H_0 : \rho_S = 0 \quad \text{against} \quad H_1 : \rho_S \neq 0$$

where ρ_S is the correlation coefficient between the variables $F_X(X_i)$ and $F_Y(Y_i)$ uniformly distributed over $[0, 1]$ (here $F_X(x) = \mathbb{P}(X \leq x)$ and $F_Y(y) = \mathbb{P}(Y \leq y)$). Like the Kendall and Pearson tests, this test is not well calibrated if the variables X_i and Y_i are not independent.

As with Kendall's test, Hoeffding [9] showed that Spearman's test statistic can be expressed using a U -statistic, from which he deduced a central limit theorem for the normalized statistic, with an exact expression of the limiting variance. As in the previous paragraph we can estimate this limit variance (by taking the empirical estimator), and then obtain a corrected test which is asymptotically well calibrated. The expression of the limit variance being more complicated than that of Kendall's statistic, we do not give all the details here.

The corresponding R function of the *robustest* package is: `cortest(,method="spearman")`.

2.4 Kolmogorov–Smirnov statistic for testing independence

The context is the same as in the previous subsection: $(X_1, Y_1), \dots, (X_n, Y_n)$ are independent copies of a couple (X, Y) of real-valued random variables, and we assume that the variables X and Y are continuous. We want to test

$$H_0 : X \text{ and } Y \text{ are independent} \quad \text{against} \quad H_1 : X \text{ and } Y \text{ are not independent}$$

To answer this question, one can use the Kolmogorov-Smirnov statistic (as described for instance in Blum, Kiefer and Rosenblatt [3])

$$KS_n = \sup_{s, t \in \mathbb{R}} \sqrt{n} \left| \frac{1}{n} \sum_{i=1}^n \mathbf{1}_{X_i \leq t, Y_i \leq s} - F_{n,X}(t) F_{n,Y}(s) \right|,$$

where

$$F_{n,X}(t) = \frac{1}{n} \sum_{i=1}^n \mathbf{1}_{X_i \leq t}, \quad F_{n,Y}(s) = \frac{1}{n} \sum_{i=1}^n \mathbf{1}_{Y_i \leq s}.$$

It is clear that, under the hypothesis H_0 : X_i is independent of Y_i , the statistic KS_n is distribution-free (since the variables are continuous, one can go back to the case where X_i and Y_i are uniformly distributed over $[0, 1]$). On another hand, one can easily check that KS_n is asymptotically equivalent to

$$KS'_n = \sup_{s, t \in \mathbb{R}} \sqrt{n} \left| \frac{1}{n} \sum_{i=1}^n (\mathbf{1}_{X_i \leq t} - F_X(t)) (\mathbf{1}_{Y_i \leq s} - F_Y(s)) \right|.$$

Now, as proved in [3], KS'_n (and hence KS_n) converges in distribution under H_0 to the supremum of a Gaussian process. The rejection region of the Kolmogorov-Smirnov test of independence is then $R_{n,\alpha} = \{KS_n > c_\alpha\}$ where c_α is the exact quantile of order $1 - \alpha$ of the distribution of KS_n under H_0 (note that, for small n , the quantity $\mathbb{P}_{H_0}(R_{n,\alpha})$ is in general not exactly equal to α , because the distribution of KS_n under H_0 is a discrete distribution). These quantiles (or the p -value of the test) can be easily estimated via a basic Monte-Carlo procedure.

The corresponding R function of the *robustest* package is: `indeptest()`.

Remark: When the continuous variables X_i and Y_i are observed with too rough a rounding, the statistics of Kendall, Spearman or Kolmogorov-Smirnov can behave badly (because they involve quantities of the type $\mathbf{1}_{X_i > Y_j}$). In the *robustest* package, we add the possibility to correct this problem by a simple randomization procedure (if there is a tie, we toss heads to see if the indicator is 0 or 1); to do so, it suffices to use the argument `ties.break="random"`. This remark is also valid for the Mann-Whitney test which will be presented in one of the following sections.

In order to highlight the differences between the classical tests and the corrected tests, we will consider in the next section two simulation scenarios.

2.5 Simulation study

First scenario: we simulate, for different values of n , i.i.d. pairs $(X_i, Y_i)_{1 \leq i \leq n}$ according to the model

$$Y_i = X_i^2 + 0.3\varepsilon_i \tag{1}$$

where the $(X_i)_{1 \leq i \leq n}$ and the $(\varepsilon_i)_{1 \leq i \leq n}$ are two independent sequences of i.i.d variables with $\mathcal{N}(0, 1)$ distribution (see Figure 1). One can easily see that, for this model, $\rho = \tau = \rho_S = 0$.

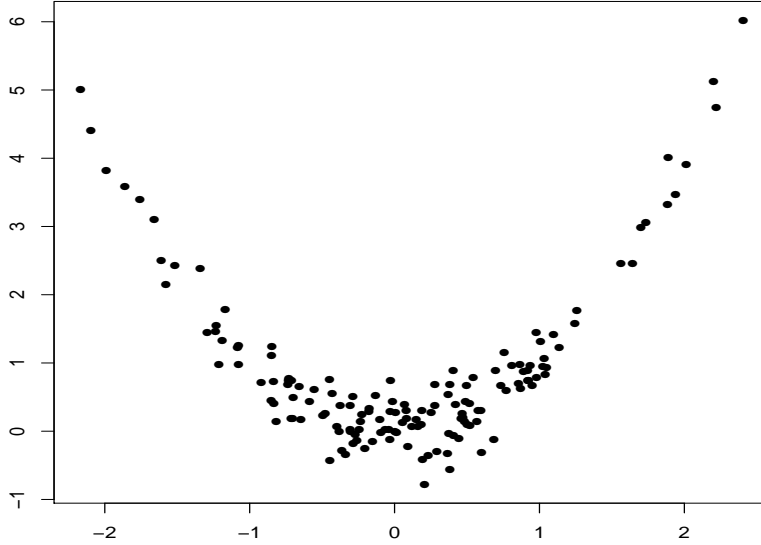


Figure 1: Scatter plot of 150 couples (x_i, y_i) drawn according to model (1)

Hence X_i and Y_i are not correlated in the sense of Pearson, Kendall or Spearman; but of course, they are not independent.

We simulate $N = 3000$ samples of size n according to model (1), and for each test we indicate the frequency of rejection of H_0 at level 5%. We consider seven tests : the usual Pearson (usual P), Kendall (usual K) and Spearman (usual S), and the Robust Pearson (robust P), Kendall (robust K) and Spearman (robust S), to test $H_0 : \rho = 0$, $H_0 : \tau = 0$ and $H_0 : \rho_S = 0$ respectively; the Kolmogorov-Smirnov independence test (KS indep) to test $H_0 : X$ and Y are independent. The results are given in Table 1 (level $\alpha = 5\%$).

From Table 1, we can notice that the three usual correlation tests are poorly calibrated, with rejection frequencies of H_0 around 37% for usual P, 19% for usual K, and 13% for usual S, instead of the expected 5%. We can also remark on this set of simulations, that it is always preferable to use robust tests, whose rejection frequencies are always lower than those of the

n	30	40	50	60	70	80	90	100	200	300
usual P	0.361	0.376	0.365	0.369	0.38	0.353	0.366	0.362	0.374	0.384
robust P	0.055	0.047	0.048	0.052	0.055	0.05	0.049	0.049	0.051	0.054
usual K	0.174	0.184	0.183	0.201	0.201	0.202	0.187	0.194	0.203	0.194
robust K	0.077	0.067	0.069	0.066	0.064	0.057	0.057	0.052	0.052	0.051
usual S	0.123	0.124	0.132	0.135	0.145	0.131	0.131	0.126	0.139	0.137
robust S	0.113	0.096	0.089	0.082	0.078	0.071	0.068	0.061	0.057	0.056
KS indep	0.732	0.934	0.989	0.999	1	1	1	1	1	1

Table 1: Frequencies of rejection of the seven tests at level 5% for model (1).

usual tests, and close to 5% for $n \geq 100$. In fact, for robust P, the frequencies of rejection are always between 4.7% and 5.5% for $n \geq 30$; for robust K, the frequencies of rejection are always between 5% and 7% for $n \geq 40$; for robust S, the frequencies of rejection are always between 5% and 7% for $n \geq 90$. The fact that the rejection frequencies of robust S converge more slowly towards 5% can perhaps be explained by the fact that the asymptotic variance term of the Spearman statistic is quite complicated, and is therefore more difficult to estimate. To conclude, unsurprisingly, the independence test detects very well the non-independence of X and Y , systematically as soon as $n \geq 70$.

Second scenario: we simulate, for different values of n , i.i.d. pairs $(X_i, Y_i)_{1 \leq i \leq n}$ according to the model

$$Y_i = (X_i \cdot 2(\varepsilon_i - 0.5))^3 \quad (2)$$

where $(X_i)_{1 \leq i \leq n}$ and $(\varepsilon_i)_{1 \leq i \leq n}$ are two independent sequences, the X_i 's being i.i.d. random variables with $\mathcal{U}([0, 1])$ distribution, the ε_i 's being i.i.d. random variables with a $\mathcal{B}(0.5)$ Bernoulli distribution (see Figure 2). Here, again, one can easily see that $\rho = \tau = \rho_S = 0$. Hence X_i and Y_i are not correlated in the sense of Pearson, Kendall or Spearman; but of course, they are not independent.

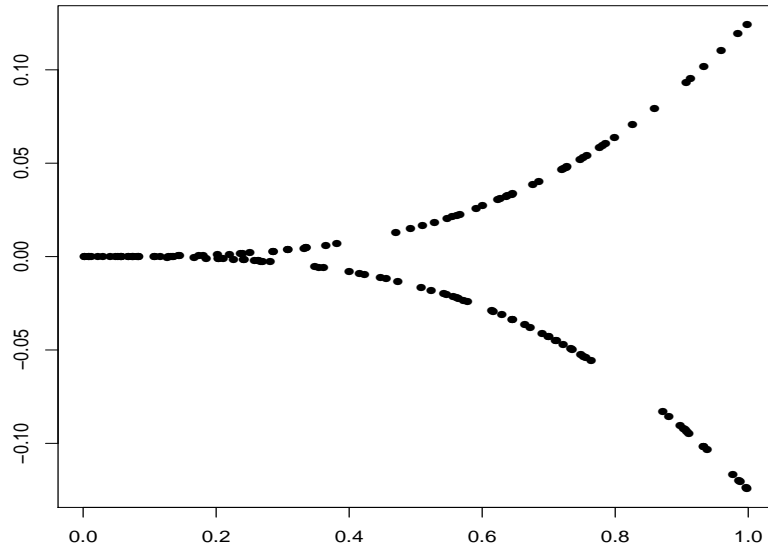


Figure 2: Scatter plot of 150 couples (x_i, y_i) drawn according to model (2)

We simulate $N = 3000$ samples of size n according to model (2), and for each test we indicate the frequency of rejection of H_0 at level 5%. We consider the same seven tests as in the first scenario. The results are given in Table 2 (level $\alpha = 5\%$).

From Table 2 we can make much the same comments and observations as for scenario 1 (although models (1) and (2) are quite different). The correlation tests are all poorly calibrated. Robust tests always have a lower rejection frequency than usual tests. In fact, for robust P, the frequencies of rejection are always between 5% and 6% for $n \geq 50$; for robust K, the frequencies of rejection are always between 5% and 7% for $n \geq 50$; for robust S, the frequencies of rejection

n	30	40	50	60	70	80	90	100	200	300
usual P	0.162	0.159	0.158	0.154	0.148	0.146	0.154	0.153	0.141	0.142
robust P	0.072	0.064	0.06	0.057	0.056	0.058	0.058	0.05	0.049	0.052
usual K	0.231	0.236	0.236	0.242	0.249	0.248	0.255	0.247	0.245	0.246
robust K	0.082	0.074	0.063	0.06	0.06	0.059	0.06	0.059	0.054	0.05
usual S	0.138	0.143	0.133	0.132	0.142	0.139	0.148	0.14	0.142	0.135
robust S	0.113	0.099	0.085	0.079	0.071	0.071	0.07	0.068	0.059	0.056
KS indep	0.956	1	1	1	1	1	1	1	1	1

Table 2: Frequencies of rejections of the seven tests at level 5% for model (2).

are always between 5% and 7% for $n \geq 90$. Our last observation is that, unsurprisingly, the independence test detects very well the non-independence of X and Y , systematically as soon as $n \geq 40$.

Conclusion: We have seen, through two really different scenarios of simulations, that if we want to test non-correlation, it is preferable to use robust tests. Indeed the usual tests can lead to a false positive rate that is much too high. One could object that the usual correlation tests can be used to test the null hypothesis H_0 : X and Y are independent. But our simulation study reveal that the usual correlation tests are not consistent to test this hypothesis (their rejection frequency does not tend to 1 when n tends to infinity), whereas the Kolmogorov-Smirnov type test is consistent.

Remark: Concerning the particular case of Pearson's test of correlation: if we know that the couple (X, Y) is Gaussian, then the usual test will be well calibrated, and more powerful than the robust version. Note however that, to our knowledge, there is no practical way to test whether the pair (X, Y) is Gaussian (which is a much stronger assumption than assuming that X and Y are Gaussian).

3 Robust test for the equality of variances

Let $(X_1, Y_1), \dots, (X_n, Y_n)$ be independent and identically distributed (i.i.d) copies of the random pair (X, Y) , where X is real-valued random variable and Y is a categorical random variable with p different levels L_1, \dots, L_p . Assume that $\mathbb{E}(X^2) < \infty$, and let $\mu_i = \mathbb{E}(X|Y = L_i)$ and $V_i = \text{Var}(X|Y = L_i)$. The problem of testing the hypothesis $H_0 : \mu_1 = \dots = \mu_p$ is called ANOVA (Analysis of Variance) in the literature. If we do not assume that the variances V_i are all equal, this problem has been solved satisfactorily by James [11] and Welch [17]. The procedure proposed by James and Welch works for any sample size n when the conditional distribution of X knowing Y is Gaussian, but also without any additional assumptions when n is large enough (the test is asymptotically well calibrated). More precisely, denoting by JW_n the James-Welch test statistic, we know that : under H_0 , $(p-1)JW_n$ converges in distribution as $n \rightarrow \infty$ to a $\chi^2(p-1)$ distribution. For the ANOVA test of James-Welch (not assuming equal variances), the corresponding R function of the *stats* package is: `oneway.test()`.

Consider now the problem of testing the equality of conditional variances, that is

$$H_0 : V_1 = \dots = V_p \quad \text{against} \quad H_1 : \exists i, j \text{ such that } V_i \neq V_j$$

We will need some notations in order to proceed. Denote by $L(Y_i)$ the level of the random

variable Y_i , by $n(L_k)$ the number of variables Y_i for which $L(Y_i) = L_k$, and by

$$\bar{X}_{L_k} = \frac{1}{n(L_k)} \sum_{i=1}^n X_i \mathbf{1}_{Y_i=L_k}.$$

If $p = 2$, to test the equality of conditional variances, one often uses the Fisher statistic

$$F_{n(L_1), n(L_2)} = \frac{S_{n(L_1)}^2}{S_{n(L_2)}^2}, \quad \text{where} \quad S_{n(L_k)}^2 = \frac{1}{n(L_k) - 1} \sum_{i=1}^n (X_i - \bar{X}_{L_k})^2 \mathbf{1}_{Y_i=L_k}.$$

The interest of this statistic being that, in the case where the conditional distribution of X knowing Y is Gaussian, its exact conditional distribution given $(n(L_1), n(L_2))$ under H_0 is known: it is the Fisher distribution $F(n(L_1) - 1, n(L_2) - 1)$. Fisher's test for equality of variances is constructed from the quantiles of this distribution. But if the conditional distribution of X knowing Y is not Gaussian, the distribution of $F_{n(L_1), n(L_2)}$ under H_0 is not known. For $p \geq 2$, one can use Bartlett's test for equality of variances [2], but again this test (based on the Gaussian likelihood) is not robust to non normality. We can therefore make the following conclusion: in a general context, Fisher's and Bartlett's tests are not well calibrated to test the equality of variances. See also [4] for a similar observation.

An alternative to Bartlett's test often mentioned in the literature is Levene's test (modified by Brown and Forsythe [5]), valid for any $p \geq 2$. This test consists of performing a classic ANOVA on the variables $|X_i - \text{Med}_n(X|L(Y_i))|$, where $\text{Med}_n(X|L_k)$ is the empirical estimator of the conditional median of X given $Y = L_k$. It is quite easy to see that this procedure does not test the equality of conditional variances, but the equality of absolute deviations from the conditional medians (as it is clearly indicated in the R documentation of the package *lawstat*).

Assume now that the random variable X has a finite moment of order 4. In order to test

$$H_0 : V_1 = \dots = V_p \quad \text{against} \quad H_1 : \exists i, j \text{ such that } V_i \neq V_j$$

we propose to perform a James-Welch ANOVA on the variables $Z_i = (X_i - \bar{X}_{L(Y_i)})^2$. Let T_n be the test statistic that we obtain by proceeding in this way. It is then easy to deduce from the James-Welch test statistic that, under H_0 , $(p - 1)T_n$ converges in distribution as $n \rightarrow \infty$ to a $\chi^2(p - 1)$ distribution.

The corresponding R function of the *robustTest* package is: `vartest()`.

Remark: The procedure described above still works within the framework of the "fixed design" ANOVA, i.e. when we observe p independent sequences $(X_{i,1})_{1 \leq i \leq n_1}, \dots, (X_{i,p})_{1 \leq i \leq n_p}$ of i.i.d. variables (in this case $V_j = \text{Var}(X_{1,j})$), provided that all n_i are large enough.

In the next subsection, we will illustrate the differences between the classical test for the equality of variance and the corrected test by a small simulation study.

3.1 Simulation study

We simulate, for different values of n , i.i.d. pairs $(X_i, Y_i)_{1 \leq i \leq n}$ according to the following model:

$$Y_i \sim \mathcal{B}(2/3), \tag{3}$$

where the conditional distribution of X_i given $Y_i = 0$ is the $\mathcal{N}(0, 1)$ distribution and the conditional distribution of X_i given $Y_i = 1$ is the $\chi^2(2)/2$ distribution. The boxplots of the conditional distributions of (x_i) given (y_i) based on 180 observations drawn according to model (3) are shown in Figure 3.

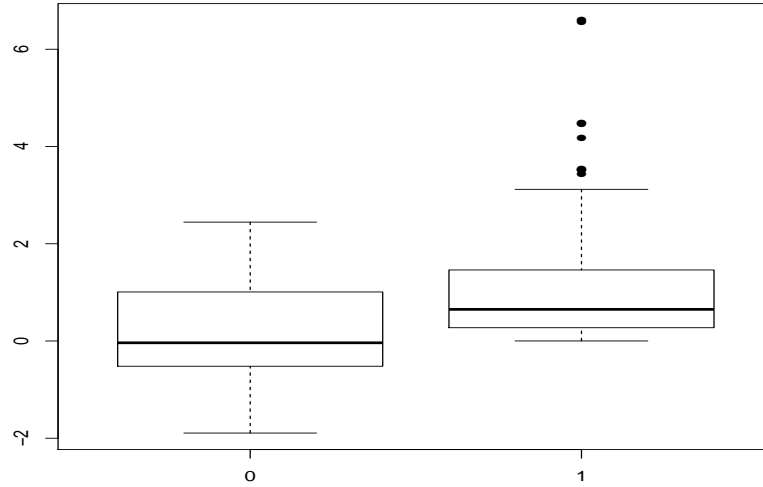


Figure 3: Boxplots of the conditional distributions of (x_i) given (y_i) based on 180 observations drawn according to model (3).

Note that the conditional variances of X given $Y = 0$ or $Y = 1$ are equal; that is, with the notations above, $V_1 = V_2 = 1$ and the hypothesis $H_0 : V_1 = V_2$ is satisfied.

We simulate $N = 3000$ samples of size n according to model (3), and for each test we indicate the frequency of rejection of $H_0 : V_1 = V_2$ at level 5%. We consider four tests : Fisher's test, Bartlett's test, Levene's test, and our test based on the James-Welch procedure (VWelch). The results are given in Table 3 (level $\alpha = 5\%$).

n	60	70	80	90	100	150	200	250	300
Fisher	0.145	0.151	0.144	0.154	0.152	0.153	0.161	0.16	0.161
Bartlett	0.142	0.147	0.143	0.153	0.152	0.153	0.161	0.16	0.161
Levene	0.08	0.083	0.091	0.097	0.106	0.119	0.155	0.172	0.196
VWelch	0.057	0.054	0.056	0.056	0.055	0.053	0.056	0.055	0.053

Table 3: Frequency of rejection of the 4 tests at level 5% for model (3)

From Table 3, our first observation is that, for this two-sample scenario, Fisher's test and Bartlett's test behave similarly, with much too high rejection frequencies hovering around 15%. The frequencies of rejection of Levene's test increases with n , from 8% when $n = 60$ to 19.6% when $n = 300$. This is not surprising, because for model (3), the absolute deviations from the conditional medians are slightly different, depending on whether $Y = 0$ or $Y = 1$. The test VWelch, based on the James-Welch procedure is well calibrated, with frequencies of rejection between 5% and 6% as soon as $n \geq 60$.

Conclusion: We have seen that Fisher's and Bartlett's tests of equality of variances are not robust to non normality. Levene's test is well suited to test equality of absolute deviations from the median, but not to test the equality of variances. For testing equality of variances, it is then

preferable to use the function `vartest()` of the *robustTest* package.

4 Robust test for testing stochastic dominance

Let (X_1, \dots, X_{n_1}) and (Y_1, \dots, Y_{n_2}) be two independent sequences of i.i.d. real-valued random variables (the X_i 's are independent copies of X , and the Y_i 's are independent copies of Y , the variables X and Y being independent). Assume moreover that the variables X and Y are continuous. We want to know if the variables Y tend to take larger or smaller values than the variables X . Let $\text{Med}(Y - X)$ be the median of $X - Y$. To answer the question, one can for example test the hypothesis

$$H_0 : \text{Med}(Y - X) = 0 \quad \text{against} \quad H_1 : \text{Med}(Y - X) \neq 0.$$

Mann and Whitney [14] proposed the test statistic

$$T_{n_1, n_2} = \frac{1}{n_1 n_2} \sum_{i=1}^{n_1} \sum_{j=1}^{n_2} (\mathbf{1}_{X_i < Y_j} - 0.5).$$

It is quite easy to see, that if X_i and Y_i have the same distribution, then the distribution of T_{n_1, n_2} is distribution-free (i.e. does not depend on the common distribution of X_i and Y_i): it suffices to reduce to two independent sequences U_1, \dots, U_{n_1} and V_1, \dots, V_{n_2} of i.i.d. random variables with $\mathcal{U}([0, 1])$ distributions. Consequently, if X_i and Y_i have the same distribution, T_{n_1, n_2} follows a known and tabulated distribution. The Mann-Whitney test is constructed from the quantiles of this distribution. It is well known that this test is in fact equivalent to Wilcoxon rank-sum test (see [18]). But if X_i and Y_i are not identically distributed, the distribution of statistics T_{n_1, n_2} has no reason to be distribution-free under H_0 . We can therefore make the following conclusion: in a general context, the Mann-Whitney test is not well calibrated to test $\text{Med}(Y - X) = 0$.

We can nevertheless solve (asymptotically) this problem, by considering the limiting distribution of $a_{n_1, n_2} T_{n_1, n_2}$ under H_0 , for a suitable normalisation a_{n_1, n_2} . Starting from Hoeffding's decomposition of the U -statistic T_{n_1, n_2} (see for instance [15], example 12.7), we see that, under H_0 ,

$$\frac{T_{n_1, n_2}}{\sqrt{\frac{V_1}{n_1} + \frac{V_2}{n_2}}} \quad \text{converges in distribution as } n_1, n_2 \rightarrow \infty \text{ to the } \mathcal{N}(0, 1) \text{ distribution,}$$

with

$$V_1 = \text{Var}(H_Y(X_1)) \quad \text{and} \quad V_2 = \text{Var}(F_X(Y_1)),$$

where $F_X(x) = \mathbb{P}(X_1 < x)$ and $H_Y(x) = \mathbb{P}(Y_1 > x)$. The empirical estimators of V_1 and V_2 are then

$$V_{1, n_1, n_2} = \frac{1}{n_1 - 1} \sum_{k=1}^{n_1} (H_{n_2}(X_k) - \bar{H}_{n_2})^2 \quad \text{and} \quad V_{2, n_1, n_2} = \frac{1}{n_2 - 1} \sum_{k=1}^{n_2} (F_{n_1}(Y_k) - \bar{F}_{n_1})^2$$

where

$$\begin{aligned} F_{n_1}(x) &= \frac{1}{n_1} \sum_{k=1}^{n_1} \mathbf{1}_{X_k < x} & H_{n_2}(y) &= \frac{1}{n_2} \sum_{k=1}^{n_2} \mathbf{1}_{Y_k > y} \\ \bar{F}_{n_1} &= \frac{1}{n_2} \sum_{k=1}^{n_2} F_{n_1}(Y_k) & \bar{H}_{n_2} &= \frac{1}{n_1} \sum_{k=1}^{n_1} H_{n_2}(X_k). \end{aligned}$$

Finally, under H_0 ,

$$MW_{n_1, n_2} = \frac{T_{n_1, n_2}}{\sqrt{\frac{V_{1, n_1, n_2}}{n_1} + \frac{V_{2, n_1, n_2}}{n_2}}}$$

converges in distribution as $n_1, n_2 \rightarrow \infty$ to the $\mathcal{N}(0, 1)$ distribution.

The rejection region of H_0 of the corrected Mann-Whitney test is therefore of the form $R_{n_1, n_2, \alpha} = \{|MW_{n_1, n_2}| > c_\alpha\}$ where c_α is the quantile of order $1 - (\alpha/2)$ of the $\mathcal{N}(0, 1)$ distribution, which provides a test asymptotically well calibrated.

The corresponding R function of the *robustTest* package is: `wilcoxtest(,paired=FALSE)`.

Remark: The procedure described above still works when both samples are obtained from n i.i.d random variables $(X_i, Y_i)_{1 \leq i \leq n}$, where the X_i 's are continuous random variables and the Y_i 's are categorical variables with two levels L_1, L_2 , considering the two sub-samples of continuous variables obtained by conditioning with respect to $Y_i = L_1$ or $Y_i = L_2$.

In the next subsection, we will illustrate the differences between the classical test for testing stochastic dominance and the corrected test by a small simulation study.

4.1 Simulations

We simulate $N = 3000$ samples of size n_1 according to the distribution $\mathcal{U}([-0.5, 0.5])$, and 3000 samples of size n_2 according to the distribution $\mathcal{N}(0, (0.04)^2)$ (with standard deviation equal to 0.04), with $n_2 = 3n_1$ (see Figure 4). For each test, we indicate the frequency of rejection of H_0 at level 5%. We consider here 4 tests: the robust Mann-Whitney (robust M-W) test based on the MW_{n_1, n_2} statistic, the uncorrected Mann-Whitney (M-W) test (both for testing $H_0 : \text{Med}(Y - X) = 0$, which is satisfied here), Welch's test (which tests the equality of the expectations without assuming the equality of the variances, see the previous section), and the two-sample Kolmogorov-Smirnov (K-S) test for testing the equality of distributions (see for instance [6], pages 309-314). The results are given in Table 4 (level $\alpha = 5\%$).

$n_1; n_2$	20;60	30;90	40;120	50;150	60;180	70;210	80;240	90;270	100;300
robust M-W	0.061	0.064	0.057	0.058	0.054	0.056	0.055	0.054	0.051
M-W	0.172	0.171	0.18	0.166	0.17	0.186	0.168	0.179	0.161
Welch	0.057	0.048	0.051	0.057	0.49	0.57	0.053	0.051	0.051
K-S	0.998	1	1	1	1	1	1	1	1

Table 4: Frequency of rejection of the four tests at level 5% for n_1 pairs drawn according to the distribution $\mathcal{U}([-0.5, 0.5])$, and n_2 couples according to the distribution $\mathcal{N}(0, (0.04)^2)$.

On this set of simulations (Table 4), we see that, for the robust M-W test and all the proposed sample sizes, the rejection frequency of H_0 is below 7%. It is between 5% and 6% as soon as $n_1 \geq 40, n_2 \geq 120$. Welch's test is also well calibrated, with rejection frequencies all between 4.8% and 5.7%. The uncorrected Mann-Whitney test is poorly calibrated, with rejection frequencies around 17%. As expected, the two-sample Kolmogorov-Smirnov test detects very well the difference between the two distributions, systematically as soon as $n_1 \geq 30, n_2 \geq 90$.

Conclusion If we want to test if $\text{Med}(Y - X) = 0$, it is a priori preferable to use the robust Mann-Whitney test, the usual test can lead to a false positive rate that is much too high. One could object that the usual Mann-Whitney test can be used to test the hypothesis $H_0: X$ and Y have the same distribution. But our simulations reveal that the usual Mann-Whitney test is not consistent to test this hypothesis (its rejection frequency does not tend to 1 when n tends

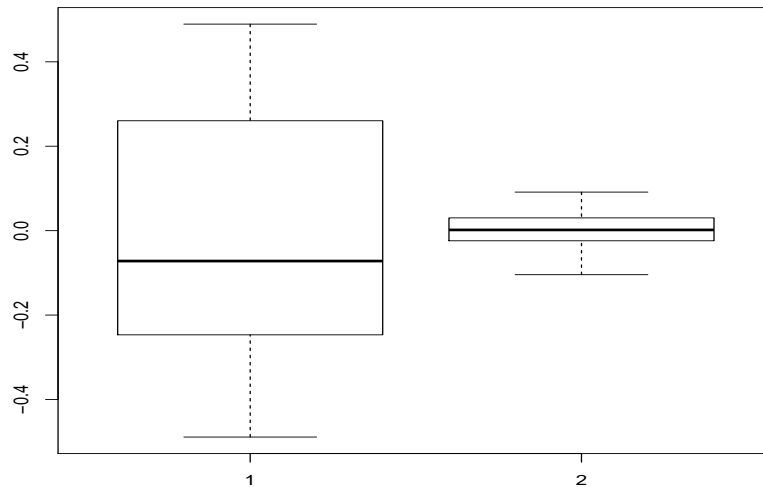


Figure 4: Boxplots of 1: 60 observations drawn according to the distribution $\mathcal{U}([-0.5, 0.5])$, and 2: 180 observations drawn according to the distribution $\mathcal{N}(0, (0.04)^2)$.

to infinity), whereas the two-sample Kolmogorov-Smirnov type test is consistent.

We also investigated how to test stochastic dominance in the case of paired samples, which will be the subject of the next section.

5 Other tests of the *robustTest* package for paired two-sample

In the *robustTest* package, we also provide functions to test stochastic dominance in the case of paired samples. The context is as follows: let $(X_1, Y_1), \dots, (X_n, Y_n)$ be independent copies of the pair (X, Y) , where X, Y are real-valued random variables. We also assume that the variables are continuous. As in the previous section, we want to know if Y tends to take larger or smaller values than X . But here, the variables X and Y are a priori not independent. Classically, we consider the series of differences $D_i = Y_i - X_i$, and we are therefore reduced to a problem of univariate statistics. We can then use the confidence interval for the median $\text{Med}(D)$ described in Example 21.8 of the book by van der Vaart [15], and based on the order statistics $(D_{(i)})_{1 \leq i \leq n}$. Let us describe this confidence interval: for $\alpha \in (0, 1)$, let

$$k_{n,\alpha} = \left\lceil -c_\alpha \frac{\sqrt{n}}{2} + \frac{n}{2} \right\rceil, \quad \ell_{n,\alpha} = \left\lceil c_\alpha \frac{\sqrt{n}}{2} + \frac{n}{2} \right\rceil,$$

the square brackets denoting the integer part, and c_α being the quantile of order $1 - (\alpha/2)$ of the $\mathcal{N}(0, 1)$ distribution. Then, the interval

$$IC_{1-\alpha} = [D_{(k_{n,\alpha})}, D_{(\ell_{n,\alpha})}], \tag{4}$$

is a confidence interval of $\text{Med}(D)$ with asymptotic confidence level $1 - \alpha$. As usual, a test of $H_0 : \text{Med}(D) = 0$ against $H_1 : \text{Med}(D) \neq 0$ can be deduced from (4). This test is an alternative to the sign test.

For these confidence interval and test, the R function of the package *robustTest* is `mediantest`. This function can be used for a single sample, as well as in the case of paired two-sample (in this case, the confidence interval and the test are for the quantity $\text{Med}(Y - X)$).

Note that we could have implemented the non-asymptotic confidence interval for $\text{Med}(D)$ (which is also described in Example 21.8 of [15]). However, some simulations suggest that, even for small samples ($25 \leq n \leq 50$), the asymptotic confidence interval behaves as well as the non asymptotic one.

As another alternative to the sign test, Wilcoxon's sign and rank test is often proposed, whose statistic is written

$$W_n = \sum_{i=1}^n R_i \mathbf{1}_{D_i > 0},$$

where R_i is the rank of $|D_i|$ in the sample $(|D_k|)_{1 \leq k \leq n}$. If the variable D is symmetric (which means that D has the same distribution as $-D$), Wilcoxon [18] showed that $|D|$ and $\mathbf{1}_{D > 0}$ are independent; we can then easily deduce that the distribution of W_n is distribution free (i.e. does not depend on the distribution of D). Therefore, if D is symmetric, the distribution of W_n is known and tabulated. Wilcoxon's sign and rank test is constructed from the quantiles of this distribution. But if the variable D is not symmetric, the distribution of W_n has no reason to be distribution free under $H_0 : \text{Med}(D) = 0$. We can therefore make the following conclusion: in a general context, Wilcoxon's sign and rank test is not correctly calibrated to test $\text{Med}(D) = 0$. In fact, the sign and rank test allows to test

$$H'_0 : \text{Med}(D_1 + D_2) = 0 \quad \text{against} \quad H_1 : \text{Med}(D_1 + D_2) \neq 0.$$

But then again, it is not well calibrated to test this hypothesis. We can nevertheless solve (asymptotically) this problem, by considering the limiting distribution of $\sqrt{n}(2U_n/(n(n-1)) - 0.5)$ under H'_0 , with

$$U_n = \sum_{i=1}^n \sum_{j=1}^{i-1} \mathbf{1}_{D_i + D_j > 0}.$$

Note that

$$W_n = U_n + \sum_{i=1}^n \mathbf{1}_{D_i > 0},$$

the second term on the right in the equality being asymptotically negligible with respect to U_n . Starting from the Hoeffding decomposition of the U -statistic U_n (see for example [15], example 12.4), we see that the variable $\sqrt{n}(2U_n/(n(n-1)) - 0.5)$ converges in distribution under H'_0 to the distribution $\mathcal{N}(0, V)$, where

$$V = 4\text{Var}(F(-D))$$

and F is the distribution function of the variables D_i . The empirical estimator of V is therefore

$$V_n = \frac{1}{n-1} \sum_{i=1}^n (F_n(-D_i) - \bar{F}_n)^2 \quad \text{where} \quad F_n(x) = \frac{1}{n} \sum_{i=1}^n \mathbf{1}_{D_i \leq x} \quad \text{and} \quad \bar{F}_n = \frac{1}{n} \sum_{i=1}^n F_n(-D_i).$$

Finally, under H'_0 ,

$$W'_n = \frac{\sqrt{n}}{\sqrt{V_n}} \left(\frac{2U_n}{n(n-1)} - 0.5 \right)$$

converges in distribution when $n \rightarrow \infty$ to the distribution $\mathcal{N}(0, 1)$.

The rejection region of H'_0 of the corrected sign and rank test is therefore of the form $R_{n,\alpha} = \{|W'_n| > c_\alpha\}$ where c_α is the quantile of order $1 - (\alpha/2)$ of the distribution $\mathcal{N}(0, 1)$,

which provides a test asymptotically well calibrated.

For the robust Wilcoxon sign and rank test, the corresponding R function of the *robustTest* package is: `wilcoxtest(,paired=TRUE)`.

Remark. As we see, the Wilcoxon sign and rank test is not well suited to test $H_0 : \text{Med}(D) = 0$ (calibration and consistency issues) nor $H_0 : "D \text{ is symmetric}"$ (consistency issues). If one wants to test the symmetry, one can rather use the Kolmogorov-Smirnov type statistic:

$$K_n = \sqrt{n} \sup_{t \in \mathbb{R}} |F_n(t) - F_{n,-}(t)|,$$

where

$$F_n(t) = \frac{1}{n} \sum_{i=1}^n \mathbf{1}_{D_i \leq t} \quad \text{and} \quad F_{n,-}(t) = \frac{1}{n} \sum_{i=1}^n \mathbf{1}_{-D_i \leq t}.$$

One can easily check that the test statistic K_n is distribution-free under $H_0 : "D \text{ is symmetric}"$, and that it converges in distribution under H_0 to the supremum of a Gaussian process.

In the last section, we illustrate on a real data set the functions of the *robustTest* package to test the correlation, and we compare the outputs to those of the usual tests.

6 Testing correlation on a real data set

We will illustrate our results on the Evans data set which comes from the *lbref* package (see also [13]). These are data from a cohort study in which white males in Evans County were followed for 7 years, with coronary heart disease as the outcome of interest. The data set is a data frame with 609 rows and 9 variables. We focus on the three variables :

CDH : outcome variable; 1 = coronary disease
 CHL : cholesterol, mg/dl
 DBP : diastolic blood pressure, mmHg

In this example, we restrict ourselves to the sub-sample made up of the 71 men having coronary heart disease. We propose to test the correlation between the CHL and DBP variables in men affected by this disease. The scatter plot is drawn in Figure 5.

To test Pearson's correlation with the `cor.test` function of the package *stats*, type

```
cor.test(CHL[CDH==1],DBP[CDH==1])
```

We get the output

```
Pearson's product-moment correlation
```

```
data: CHL[CDH == 1] and DBP[CDH == 1]
t = 1.706, df = 69, p-value = 0.09251
alternative hypothesis: true correlation is not equal to 0
95 percent confidence interval:
 -0.03370795  0.41500321
sample estimates:
 cor
0.2011779
```

To test Pearson's correlation with the `cortest` function of the package *robustTest*, type

```
cortest(CHL[CDH==1],DBP[CDH==1])
```

We get the output

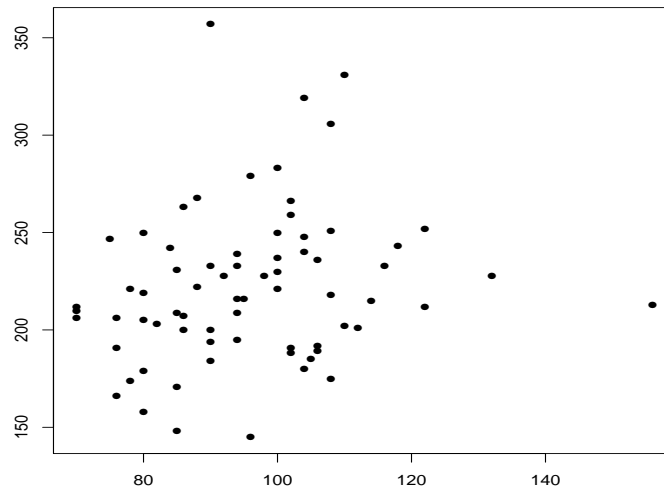


Figure 5: Scatter plot of cholesterol as a function of diastolic blood pressure in patients with coronary artery disease

Corrected Pearson correlation test

```
t = 2.4126, p-value = 0.0174
alternative hypothesis: true correlation is not equal to 0
95 percent asymptotic confidence interval for the correlation coefficient:
 0.0331 0.3692
sample estimates:
 cor
0.2011779
```

We can see that the results of the two tests are very different, with a p-value of 9.25% for the usual Pearson test, and 1.74% for the robust Pearson test. The robust test not requiring any hypothesis on the distribution of the pair of variables, it is a priori the one that must be retained. Note that Shapiro's test on the variables CHL and DBP each time gives a p-value lower than 1%, which implies that the observations do not come a priori from a pair of Gaussian variables. At the 5% risk level, we conclude that there is indeed a significant correlation in the sense of Pearson.

Let us now try to see the differences between the usual versions of the Kendall and Spearman tests and their corrected versions. Using the command `tiebreak` of the package `robustTest`, we see that ties are present in both series of observations. To compare the usual tests and their corrected versions, we get rid of this problem by using

```
X=tiebreak(CHL[CDH==1])
Y=tiebreak(DBP[CDH==1])
```

To test Kendall's correlation with the `cor.test` function of the package `stats`, type

```
cor.test(X,Y, method="kendall")
```

We get the output

Kendall's rank correlation tau

```
data: X and Y
z = 2.0003, p-value = 0.04546
alternative hypothesis: true tau is not equal to 0
sample estimates:
  tau
0.162173
```

To test Kendall's correlation with the `cortest` function of the package `robustTest`, type `cortest(X,Y, method="kendall")`

We get the output

Corrected Kendall correlation test

```
t = 2.4122, p-value = 0.0159
alternative hypothesis: true tau is not equal to 0
95 percent confidence interval:
 0.0304 0.2939
sample estimates:
  tau
0.162173
```

We see a clear difference between the p-values of the two tests. That of the usual test is close to 5%, while that of the robust test is 1.6%. Note that these values remain globally stable when several attempts are made with the tiebreak function. The comments are exactly the same for the usual and robust Spearman tests (using `method="spearman"`): the p-value of the usual test is 4.45% while that of the robust test is 1.72%. Note the the p-values of the three robust tests (Pearson, Kendall, Spearman) are all below 2%.

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