As was shown recently, the measurement errors in regressors affect only the power of the rank test, but not its critical region. Noting that, we study the effect of measurement errors on R-estimators in linear model. It is demonstrated that while an R-estimator admits a local asymptotic bias, its bias surprisingly depends only on the precision of measurements and does neither depend on the chosen rank test score-generating function nor on the regression model error distribution. The R-estimators are numerically illustrated and compared with the LSE and $L_1$ estimators in this situation.

Keywords: contiguity, local asymptotic bias, linear rank statistic, linear regression model, measurement error, R-estimate.

1. Introduction

Measurement technologies are often affected by random errors; if the goal of the experiment is to estimate a parameter, then the estimate is biased, and thus inconsistent. This problem appears in the analytic chemistry, in environmental monitoring, in modeling astronomical data, in biometrics, and practically in all parts of the reality. Moreover, some observations can be undetected, e.g. when the measured flux (light, magnetic) in the experiment falls below some flux limit. In econometrics, the errors can be a result of misreporting by subjects, miscoding by the collectors of the data, or by incorrect transformation from initial reports. An essential part of measuring techniques, used for example in the analytic chemistry, is the construction of a calibration curve - the result for an unknown sample is then determined by interpolation. However, even the calibration can be affected by measurement errors. The mismeasurements make the statistical inference biased, and they distort the trends in the data.

A variety of functional models have been proposed for handling measurement errors in regression models. Either the regressor or the response or both can be affected by random errors. Technicians, geologists and other specialists are aware of this problem, and try to reduce the bias with various ad hoc procedures. The bias cannot be completely eliminated or substantially reduced unless we
have some additional knowledge on the behavior of measurement errors. The papers dealing with practical aspects of measurement error models include [2], [15], [21], [23], [30], among others.

Adcock [1] was probably the first to realize the importance of the situation. There exists a rich literature on the statistical inference in the error-in-variables (EV) models as is evidenced by the monographs of Fuller [9], Carroll et al. [6], and Cheng and van Ness [7], and the references therein. The monographs [9] and [7] deal mostly with classical Gaussian set up while [6] discusses numerous inference procedure under semi-parametric set up. Nonparametric methods in EV models are considered in [4], [5] and in references therein, and in [8], among others. The regression quantile theory in the area of EV models was started by He and Liang [13]. Arias [3] used an instrumental variable estimator for quantile regression, considering biases arising from unmeasured ability and measurement errors. The problem of mismeasurement is also of interest in the econometric literature: [11] and [16] described the recent developments in treating the effect of mismeasurement on econometric models.

The advantage of rank and signed rank procedures in the measurement errors models was discovered recently in [20], [25], [26], [31] and in [32]: the latter made a detailed analysis of rank procedures in the linear model with a nonlinear nuisance regressor and under various kinds of measurement errors. Namely the rank tests can be recommended in this situation: it is shown in [20] that the critical region of the rank test for regression is insensitive to measurement errors in regressors under very general conditions; the errors affect only the power of the test. However, against expectations following from the invariance of the ranks, due to which an estimate of a nuisance parameter in [20] was consistent for every fixed value of the same, we show that the R-estimator of slope parameter \( \beta \) in linear model is biased. More precisely, we show that, unless \( \beta = 0 \), the R-estimator is biased even in a local neighborhood of \( 0 \). Hence, we cannot have an unbiased estimator of any kind in this situation, unless we have some additional information on the measurement errors.

As we further show in the present paper, surprisingly the local asymptotic bias of R-estimator neither depends on the chosen rank test score-generating function nor on the unknown distribution of the model errors. It depends only on value of slope parameter vector and on the covariance matrix of the measurement error distribution of regressors.

2. Model and preliminary considerations

Consider the linear regression model

\[
Y_{ni} = \beta_0 + x_{ni}^T \beta + e_{ni}, \quad i = 1, \cdots, n
\]  

(2.1)

with unknown parameters \( \beta_0 \in \mathbb{R}^1, \beta \in \mathbb{R}^p \). The regressors \( x_{ni} \) are either deterministic or random and affected by additive random measurement errors, so that instead of \( x_{ni} \) we observe \( w_{ni} = x_{ni} + v_{ni}, \ i = 1, \cdots, n \), where \( v_{n1}, \cdots, v_{nn} \) are \( p \)-dimensional random errors, identically distributed with an unknown distribution, and independent of the errors \( e_{ni}, 1 \leq i \leq n \). Moreover, there are additive measurement errors in the responses, thus instead of \( Y_{ni} \) we observe \( Y_{ni}^* = Y_{ni} + u_{ni} \), where \( u_{n1}, \cdots, u_{nn} \) are i.i.d. random variables. Thus in terms of the observable responses and predicting variables, our regression model becomes

\[
Y_{ni}^* = \beta_0 + w_{ni}^T \beta + e_{ni}^*, \quad i = 1, \cdots, n,
\]  

(2.2)
where \( e_{ni}^* = e_{ni}^*(\beta) = e_{ni} + u_{ni} - v_{ni}^\top \beta, \ i = 1, \cdots, n \) are i.i.d random variables.

We are interested in R-estimator of the slope vector \( \beta \), considering \( \beta_0 \) as nuisance parameter. To define these estimators, let \( R_{ni}(b) \) be the rank of the residual

\[
Y_{ni}^* - w_{ni}^\top b = e_{ni} + u_{ni} + x_{ni}^\top \beta - w_{ni}^\top b = e_{ni} + u_{ni} - w_{ni}^\top b + v_{ni}^\top \beta, \quad i = 1, \cdots, n
\]

where \( b^* = b - \beta \). We shall work with the vector of linear rank statistics

\[
S_n(b) = (S_{nj}(b); \ j = 1, \cdots, p)^\top = n^{-1/2} \sum_{i=1}^n (w_{ni} - \bar{w}_n)a_n(R_{ni}(b)), \quad (2.3)
\]

where the scores \( a_n(i), 1 \leq i \leq n \) are nondecreasing in \( i \) and \( \sum_{i=1}^n a_n(i) = 0 \).

Hodges and Lehmann [14] introduced a class of estimators of the location parameter \( \theta \) in one- and two- sample location models, by inverting a class of rank tests for \( \beta \). This methodology was extended to linear regression models without measurement error by Jurečková [19], where an estimator of \( \beta \) is defined as

\[
\hat{\beta}_n = \arg \min_{b \in \mathbb{R}^p} \sum_{j=1}^p |S_{nj}(b)|.
\]

This estimator can be seen to be asymptotically equivalent to an estimator obtained by inverting the equations \( S_{nj}(b) = 0, j = 1, \cdots, p \). Note that this latter estimator is precisely an extension of the Hodges-Lehmann estimator from one- and two- sample location models to linear regression models without measurement error.

On the other hand, Jaeckel [17] called an analog of the function

\[
D_n(b) = \sum_{i=1}^n (Y_{ni}^* - w_{ni}^\top b) (a_n(R_{ni}(b)) - \bar{a}_n), \quad (2.4)
\]

as a measure of rank dispersion of residuals, in the case of no measurement error where \( w_{ni} \)'s are replaced by \( x_{ni} \)'s. He showed that \( D_n(b) \) is convex and piecewise linear in \( b \in \mathbb{R}^p \). He also showed that \( -n^{1/2}S_n(b) \) is the subgradient of \( D_n(b) \); hence the estimator defined as a minimizer of \( D_n \) exists and is equivalent to the above estimators based on \( S_n \). Both of these estimators are asymptotically equivalent, and Jaeckel’s definition of R-estimator is now generally used in the literature. We are using this definition of R-estimator throughout this paper.

In the absence of measurement errors, i.e. if \( w_{ni} = x_{ni}, u_{ni} = 0, i = 1, \cdots, n \), the estimator \( \hat{\beta}_n \) is consistent and asymptotically normal. However, \( \hat{\beta}_n \) is biased in the presence of measurement errors, even asymptotically, unless the true \( \beta = 0 \). Furthermore, we show that it is even asymptotically locally biased in the sense that the asymptotic distribution of \( n^{1/2} (\hat{\beta}_n - n^{-1/2} \beta^0) \), with a fixed \( \beta^0 \in \mathbb{R}^p \), converges to a normal distribution with non-zero mean vector and some positive definite covariance matrix.

In the sequel, all limits are taken as \( n \to \infty \), unless mentioned otherwise. \( \overset{P}{\to} \) denotes the convergence in probability. We shall now describe the needed assumptions on the underlying entities.

**A.1** The score generating function \( \varphi : (0, 1) \to \mathbb{R} \) is nondecreasing, square-integrable and skew-symmetric on \( (0, 1) \), i.e. satisfies \( \varphi(1 - t) = -\varphi(t), \ 0 < t < 1 \). The scores \( a_n(i), i = 1, \cdots, n \) are generated by \( \varphi \) in either of the following two ways:

\[
a_n(i) = \varphi \left( \frac{i}{n + 1} \right) \quad \text{or} \quad a_n(i) = \mathbb{E} \varphi (U_{ni}), \ i = 1, \cdots, n
\]
where $U_{n,1} \leq \cdots \leq U_{n,n}$ are order statistics pertaining to the sample of size $n$ from the uniform $(0,1)$ distribution.

\textbf{F.1} Distribution function $F$ of the model errors $e_{ni}$ has an absolutely continuous density $f$ with a.e. derivative $f'$.

\textbf{F.2} For every $u \in \mathbb{R}$, $\int (|f'(x - tu)|^j/f^{j-1}(x))dx \to \int (|f'(x)|^j/f^{j-1}(x))dx < \infty$, as $t \to 0$, $j = 2, 3$.

\textbf{V.1} The measurement errors $\{u_{ni}, 1 \leq i \leq n\}$ are independent of $\{e_{ni}, v_{ni}, 1 \leq i \leq n\}$ and i.i.d. with generally an unknown absolutely continuous density $h$, having finite Fisher’s information for location.

\textbf{V.2} The measurement error $v_{ni}$ is independent of $e_{ni}$ and its $p$-dimensional distribution function $G$ has a continuous density $g$, generally unknown, $i = 1, \cdots, n$.

\textbf{V.3} $E V_n \to V$ where $V_n = n^{-1}\sum_{i=1}^{n}(v_{ni} - \bar{v}_n)(v_{ni} - \bar{v}_n)^\top$ and $V$ is a positive definite $p \times p$ matrix. Moreover, $\sup_{n \geq 1} E(\|v_{ni}\|^3 + \|x_{ni}\|^3) < \infty$.

\textbf{V.4} $E \left[ n^{-1} \sum_{i=1}^{n}(v_{ni} - \bar{v}_n)(x_{ni} - \bar{x}_n)^\top \right] \to 0$.

\textbf{X.1} If the regressors $x_{ni}$ are nonrandom, then assume that $Q_n \to Q$, where

$$Q_n = n^{-1}\sum_{i=1}^{n}(x_{ni} - \bar{x}_n)(x_{ni} - \bar{x}_n)^\top,$$

and $Q$ is positive definite $p \times p$ matrix. Moreover,

$$\frac{1}{n}\max_{1 \leq i \leq n}(x_{ni} - \bar{x}_n)^\top(Q_n)\left(Q_n\right)^{-1}(x_{ni} - \bar{x}_n) \to 0.$$

\textbf{X.2} If the regressors $x_{ni}$ are random, then assume that they are independent of $e_{ni}$, $u_{ni}$, $v_{ni}$, $i = 1, \cdots, n$, and

$$E \left[ n^{-1} \sum_{i=1}^{n}(x_{ni} - \bar{x}_n)(x_{ni} - \bar{x}_n)^\top \right] \to Q,$$

where $Q$ is positive definite $p \times p$ matrix.

Let $m(\cdot)$, $M(\cdot)$ denote the density and distribution function of $e_{ni} + u_{ni}$, $i = 1, \cdots, n$, i.e. $m(z) = \int f(z - t)h(t)dt$. The density is absolutely continuous and has finite Fisher’s information $\mathcal{I}(m)$. We need to define

$$\gamma_m = -\int_{\mathbb{R}} \varphi(M(z))dm(z), \quad A_m^2(\varphi) = \gamma_m^{-2} \int_0^1 \varphi^2(u)du,$$

$$B = -(Q + V)^{-1}V\beta^0.$$

The following theorem gives the asymptotic distribution of the estimator $\hat{\beta}_n$ when the true parameter value is

$$\beta_n = n^{-1/2}\beta^0, \beta^0 \in \mathbb{R}^p \text{ fixed.} \quad (2.6)$$

\textbf{Theorem 2.1}. Assume the conditions \textbf{A.1}, \textbf{F.1} – \textbf{F.2}, \textbf{V.1} – \textbf{V.4}, \textbf{X.1} – \textbf{X.2} hold. When the true parameter value is $\beta_n$, the R-estimator $\hat{\beta}_n$ is asymptotically normally distributed with the bias $B = -(Q + V)^{-1}V\beta^0$, i.e.

$$n^{1/2}(\hat{\beta}_n - \beta_n) \overset{D}{\to} N_p(B, (Q + V)^{-1}A_m^2(\varphi)) \cdot \quad (2.7)$$
Theorem 2.1 will be proved in several steps; the proof is given in Section 3. The numerical illustrations of the results are given in subsequent Section 4.

Corollary 2.1. Under the conditions of Theorem 2.1 and under \( \beta = \beta_n = n^{-1/2}\beta^0 \), the R-estimator \( \hat{\beta}_n \) has asymptotic normal distribution

\[
n^{1/2}(\hat{\beta}_n - (Q + V)^{-1}Q\beta_n) \xrightarrow{D} \mathcal{N}_p(0, (Q + V)^{-1}A^2_n(\varphi)). \tag{2.8}
\]

Notice that the local asymptotic bias cannot be controlled by the choice of the score-generating function \( \varphi \); this choice can only influence the asymptotic variance factor of the estimator. The magnitude of the bias fully depends on the precision of the measurements, namely on the matrix \( V \).

The measurement errors in the responses \( Y_{ni} \) affect only the asymptotic variance, not the bias. The result is entirely non-parametric, valid for classes of distributions of model and measurement errors, demanding only finite first moment and finite (and positive) Fisher’s information for location of the model error distributions, and finite third moment for measurement error distributions.

Consider the two measurement methods with the same regressors (random or non-random), with the respective limiting covariance matrices \( V_1, V_2 \). Comparing the biases in (2.7) for \( V_1 \) and \( V_2 \), the first method is considered being more precise than the second one if the matrix \( (V_2 + Q)^{-1} \preceq (V_1 + Q)^{-1} \); otherwise speaking, if \( Q^{-1}V_1 \preceq Q^{-1}V_2 \), where the ordering \( A \prec B \) means that \( B - A \) is a positive definite matrix.

3. Proof of Theorem 2.1

We shall prove Theorem 2.1 in several steps. Notice that if we observe \( Y_{ni}^* = Y_{ni} + u_{ni} \) instead of \( Y_{ni} \), then \( e_n^* = e_{ni} + u_{ni}, \ i = 1, \cdots, n \) are still i.i.d. random variables with density \( m(z) = \int f(z - t)h(t)dt \). The steps of the proof are parallel for both densities \( f \) and \( m \) of model errors; measurement errors in the \( Y_{ni} \) affect only the asymptotic variance of the estimate, not the bias. Noting this, we shall prove the theorem assuming \( u_{ni} = 0, \ i = 1, \cdots, n \). In the sequel, we shall suppress the subscript \( n \) whenever it does not cause a confusion.

The steps of the proof are as follows:

1. Asymptotic representation of the linear rank statistic

\[
S_n(0, 0) = n^{-1/2} \sum_{i=1}^n (w_{ni} - \bar{w}_n)a_n(R_{ni}(0)) \tag{3.1}
\]

with the sum of independent summands. Here \( w_{ni} = x_{ni} + v_{ni}, \ i = 1, \cdots, n \), while \( x_{n1}, \cdots, x_{nn} \) are either i.i.d. random vectors or nonrandom vectors, and \( v_{n1}, \cdots, v_{nn} \) are i.i.d. random vectors.

2. Contiguity of the sequence \( \{Q_n\} \) of distributions of \( (e_{ni} - (w_{ni} - \bar{w}_n)^\top b_n^* - (v_{ni} - \bar{v}_n)^\top \beta_n) \), with \( b_n^* = n^{-1/2}b^0, \ \beta_n = n^{-1/2}\beta^0 \) for \( b^0, \ \beta^0 \in \mathbb{R}^p \) fixed, with respect to the sequence \( \{P_n\} \) of distributions of \( e_{ni}, \ i = 1, \cdots, n \).

3. Asymptotic representation of the linear rank statistic (2.3) under contiguous sequence of distribution \( \{Q_n\} \), and the resulting asymptotic linearity of (2.3) in parameters \( b^0, \beta^0 \).
(4) Uniform asymptotic quadraticity of $D_n$ in parameters $b^0, \beta^0$ under \{$(Q_n)$\}, as a result of (3) and of the convexity of $D_n$.

(5) Resulting asymptotic distribution and bias of $\hat{\beta}_n$ in the case $u_{ni} \equiv 0$, $i = 1, \cdots, n$.

(6) Asymptotic distribution and bias of $\hat{\beta}_n$ in the case of non-zero $u_{ni}$, $i = 1, \cdots, n$.

3.1. Asymptotic representation of $S_n(0, 0)$

Assume that $u_{ni} = 0$, $i = 1, \cdots, n$. That is, for now we assume that there is no measurement error in the response variables $Y_{ni}$. Let

$$Z_n = n^{-1/2} \sum_{i=1}^{n} (w_{ni} - \bar{w}_n) \varphi(F(e_{ni})).$$

We are ready to state and prove

**Lemma 3.1.** Under the conditions of Theorem 2.1, the statistic $S_n(0, 0)$ admits the asymptotic representation

$$S_n(0, 0) = Z_n + o_p(1). \quad (3.2)$$

**Proof.** The proof is adapted from [28]. If $b = \beta = 0$, then $(Y_{n1}, \cdots, Y_{nn}) = (e_{n1}, \cdots, e_{nn})$. Let $R_{n1}, \cdots, R_{nn}$ denote their ranks. Further denote $r_{ni} = a_n(R_{ni}) - \varphi(F(e_{ni}))/2$, $i = 1, \cdots, n$.

Let $\sigma^2_j$ be the variance of $w_{ij}$, $i = 1, \cdots, n$, for $j = 1, \cdots, p$, and let $s^2 = \sum_{j=1}^{p} \sigma^2_j$.

Notice that $(r_{n1}, \cdots, r_{nn})$ and $(w_1, \cdots, w_n)$ are independent. Consider the conditional squared distance

$$E_G \left\{ (S_n - Z_n)^\top (S_n - Z_n) \bigg| e_1, \cdots, e_n \right\}$$

$$= n^{-1} E_G \left\{ \sum_{i=1}^{n} \sum_{k=1}^{n} (w_i - \bar{w}_n)^\top (w_k - \bar{w}_n) r_i r_k \bigg| e_1, \cdots, e_n \right\}$$

$$= n^{-1} \sum_{i=1}^{n} \sum_{k=1}^{n} r_i r_k E_G \left\{ \sum_{j=1}^{p} (w_{ij} - \bar{w}_j)(w_{kj} - \bar{w}_j) \bigg| e_1, \cdots, e_n \right\}$$

$$= n^{-1} \sum_{i=1}^{n} \sum_{k=1}^{n} r_i r_k \left\{ n^{-1/2} \sum_{j=1}^{n} (x_{ij} - \bar{x}_j)(x_{kj} - \bar{x}_j) + s^2 \sum_{i=1}^{n} (r_i - \bar{r})^2 \right\}$$

$$= n^{-1} \sum_{j=1}^{p} \left[ n^{-1/2} \sum_{i=1}^{n} (x_{ij} - \bar{x}_j) r_i \right]^2 + s^2 \sum_{i=1}^{n} (r_i - \bar{r})^2.$$

Then (3.2) follows from [10] [Theorems V.1.4.a,b, V.1.6.a].

3.2. Contiguity

For any two probability measures $P$ and $Q$, absolutely continuous with respect to a $\sigma$-finite measure $\nu$ with $p = dP/\nu$, $q = dQ/\nu$, let

$$H(P, Q) = \left( \int (\sqrt{p} - \sqrt{q})^2 \, d\mu \right)^{1/2} = \left( 2 \int (1 - \sqrt{pq}) \, d\mu \right)^{1/2}.$$
denote the the Hellinger distance between $P$ and $Q$.

Let $\{P_n, \cdots, P_{nn}\}$ and $\{Q_n, \cdots, Q_{nn}\}$ be two triangular arrays of probability measures defined on measurable space $(\mathcal{X}, \mathcal{A})$ with densities $p_n, q_n$ with respect to $\sigma$-finite measures $\mu_i$ [which can be also $\mu_i = P_{ni} + Q_{ni}, i = 1, \cdots, n$]. Denote $P_n^{(n)} = \prod_{i=1}^n P_{ni}$ and $Q_n^{(n)} = \prod_{i=1}^n Q_{ni}$ the product measures, $n = 1, 2, \cdots$

Oosterhoff and van Zwet [27] proved that $\{Q_n^{(n)}\}$ is contiguous with respect to $\{P_n^{(n)}\}$ if and only if

\[
\limsup_{n \to \infty} \frac{1}{n} \sum_{i=1}^n H^2(P_{ni}, Q_{ni}) < \infty, \tag{3.3}
\]

\[
\lim_{n \to \infty} \frac{1}{n} \sum_{i=1}^n Q_{ni} \left\{ \frac{q_n(X_{ni})}{p_n(X_{ni})} \geq c_n \right\} = 0, \; \forall \; c_n \to \infty. \tag{3.4}
\]

Note that in the case $P_{ni} = P_n, P_{ni} = p_n$, and $Q_{ni} = Q_n, q_{ni} = q_n$, not depending on $i$, $\sum_{i=1}^n H^2(P_{ni}, Q_{ni}) = \int \left[ \sqrt{q_n(z)} - \sqrt{p_n(z)} \right]^2 dz = n \int \frac{(q_n(z) - p_n(z))^2}{\sqrt{q_n(z)} + \sqrt{p_n(z)}} dz \leq n \int \frac{(q_n(z) - p_n(z))^2}{p_n(z)} dz. \tag{3.5}

Moreover, for $c_n > 1$ and with $d_n = c_n - 1$, $\sum_{i=1}^n Q_{ni} \left\{ \frac{q_n(X_{ni})}{p_n(X_{ni})} \geq c_n \right\} = nQ_n \left\{ \frac{q_n(X_{ni}) - p_n(X_{ni})}{p_n(X_{ni})} \geq d_n \right\} \leq d_n^{-2} n \int \frac{(q_n(x) - p_n(x))^2}{p_n^2(x)} q_n(x) dx \leq d_n^{-2} n \int \frac{(q_n(x) - p_n(x))^3}{p_n^2(x)} dx + d_n^{-2} n \int \frac{(q_n(x) - p_n(x))^2}{p_n(x)} dx.$

Now, let $Y_{ni} = x_{ni}^\top \beta + e_{ni}, i = 1, \cdots, n$, where the $e_{ni}$ are i.i.d. with distribution function $F$ and density $f$, satisfying F.1 and F.2. Consider the residuals

\[
Y_{ni} - (w_{ni} - \bar{w}_n)^\top b_n = e_{ni} + (x_{ni} - \bar{x}_n)^\top \beta_n - (w_{ni} - \bar{w}_n)^\top b_n = e_{ni} - (w_{ni} - \bar{w}_n)^\top b_n^* - (v_{ni} - \bar{v}_n)^\top \beta_n,
\]

$i = 1, \cdots, n$, where $b_n = n^{-1/2}b^0, \beta_n = n^{-1/2}\beta^0, b_n^* = n^{-1/2}b_0^*, b_0^* = b^0 - \beta^0$, with fixed $b^0, \beta^0 \in \mathbb{R}^p$. Using (3.3) and (3.4), we shall prove the following lemma:

**Lemma 3.2.** Under the conditions of Theorem 2.1, the sequence $\{Q_n^{(n)}\}$ is contiguous with respect to $\{P_n^{(n)}\}$, where $Q_n^{(n)} = \prod_{i=1}^n Q_{ni}, P_n^{(n)} = \prod_{i=1}^n P_{ni}$, where $P_{ni}$ is the distribution of $e_{ni}$ and $Q_{ni}$ is the distribution of $(e_{ni} - (w_{ni} - \bar{w}_n)^\top b_n^* - (v_{ni} - \bar{v}_n)^\top \beta_n), i = 1, \cdots, n$.

**Proof.** We shall distinguish the two cases: the $x_{ni}$ are either i.i.d. random vectors or nonrandom vector components.

We start with the first case, where $w_{ni}, \cdots, w_{nn}$ are i.i.d. random vectors. Note that $U_i := (w_{ni} - \bar{w}_n)^\top b_0^* + (v_{ni} - \bar{v}_n)^\top \beta^0, i = 1, \cdots, n$, are i.i.d. r.v.’s. Let $k_1$ denote the common density
function of $U_i$. Then, $Q_{ni}, P_{ni}$ do not depend on $i$ and $q_n(x) \equiv \int f(x - n^{-1/2}u)k_1(u)du$, $p_n(x) \equiv f(x)$. Hence, by the Cauchy-Schwarz inequality, and the Fubini Theorem,

$$
n \int \frac{(q_n(x) - p_n(x))^2}{p_n(x)}dx = n \int \left\{ \int [f(x - n^{-1/2}u) - f(x)] k_1(u)du \right\}^2 \frac{dx}{f(x)}
$$

$$
\leq n \int \left[ \int [f(x - n^{-1/2}u) - f(x)]^2 \frac{k_1(u)}{f(x)}dudx \right]
$$

$$
\leq n \int \left[ \int \left( \int_{-n^{-1/2}}^{n^{-1/2}} |u f'(x - tu)|dt \right)^2 \frac{k_1(u)}{f(x)}dudx \right]
$$

$$
\leq 2n^{1/2} \int \int \left( \int_{-n^{-1/2}}^{n^{-1/2}} |f'(x - tu)|^2 dt u^2 k_1(u) \right)dudx
$$

$$
\leq 2n^{1/2} \int \int \left( \int_{-n^{-1/2}}^{n^{-1/2}} \frac{|f'(x - tu)|^2}{f(x)} dx u^2 k_1(u) \right)du, \quad \forall n \geq 1.
$$

Hence, by (3.5), (F.2) applied with $j = 2$, and by (V.3), which guaranteed $\int u^2 k_1(u)du < \infty$,

$$
\limsup_n \sum_{i=1}^n H^2(P_{ni}, Q_{ni}) \leq 2I(f) \int u^2 k_1(u)du < \infty. \quad (3.7)
$$

Similarly, the bound

$$
n \int \frac{(q_n(x) - p_n(x))^3}{p_n^2(x)}dx \leq 2n^{1/2} \int \int \int_{-n^{-1/2}}^{n^{-1/2}} \frac{|f'(x - tu)|^3}{f^2(x)} dx |u|^3 k_1(u) du dt, \quad \forall n \geq 1
$$

together with (3.6), (F.2) applied with $j = 3$, and (V.3), which guaranteed $\int |u|^3 k_1(u)du < \infty$, yield

$$
\lim_n \sum_{i=1}^n Q_{ni} \left\{ \frac{q_{ni}(Y_{ni})}{p_{ni}(Y_{ni})} \geq c_n \right\} \leq 2 \lim_n d_n^{-2} \left\{ \int \left( \frac{|f'(x)|}{f(x)} \right)^3 f(x)dx \right\} \int |u|^3 k_1(u)du
$$

$$
+ I(f) \int u^2 k_1(u)du \right) = 0.
$$

This ensures the validity of (3.4), and completes the proof of the contiguity in present case.

Next, consider the case where $x_{n1}, \cdots, x_{nn}$ are nonrandom, and we observe $w_{ni} = x_{ni} + v_{ni}$, $i = 1, \cdots, n$. Let $k_2$ denote the density of $(v_{ni} - \bar{v}_n)^\top b^0$, $i = 1, \cdots, n$. Again, by (3.5),

$$
\sum_{i=1}^n H^2(P_{ni}, Q_{ni})
$$

$$
\leq \sum_{i=1}^n \int \left\{ \int [f(e - n^{-1/2}u) - f(e)] k_2(u + (x_{ni} - \bar{x}_n)^\top b^0)du \right\}^2 \frac{de}{f(e)}
$$

$$
\leq \sum_{i=1}^n \int \left\{ \int [f(e - n^{-1/2}u) - f(e)]^2 k_2(u - (x_{ni} - \bar{x}_n)^\top b^0)du \right\} \frac{de}{f(e)}
$$

$$
\leq 2n^{1/2} \int \int \frac{|f'(e - tu)|^2}{f(e)} de dt \int n^{-1} \sum_{i=1}^n u^2 k_2(u - (x_{ni} - \bar{x}_n)^\top b^0)du.
$$

Hence, by (F.2) and by the change of variable formula,

$$
\limsup_n \sum_{i=1}^n H^2(P_{ni}, Q_{ni}) \leq C \left[ \int u^2 k_2(u)du + b^{0*\top} Q_n b^{0*} \right] < \infty. \quad (3.8)
$$
Similarly one verifies (3.4) here.

Lemmas 3.1 and 3.2 enable us to extend the approximation of the rank statistic $S_n(b_n^*, \beta_n)$ by a sum of independent r.v.‘s under the contiguous sequence of distributions. Let

$$ T_n(b_n^*, \beta_n) = n^{-1/2} \sum_{i=1}^{n} (w_{ni} - \bar{w}_n) \varphi \left( F(e_{ni} - (w_{ni} - \bar{w}_n)^\top b_n^* - (v_{ni} - \bar{v}_n)^\top \beta_n) \right). $$

We have the following

**Corollary 3.1.** Under the conditions of Theorem 2.1, and under $\{Q_n^{(n)}\}$,

$$ S_n(b_n^*, \beta_n) = n^{-1/2} \sum_{i=1}^{n} (w_{ni} - \bar{w}_n) a_n \left( R(e_{ni} - (w_{ni} - \bar{w}_n)^\top b_n^* - (v_{ni} - \bar{v}_n)^\top \beta_n) \right) = T_n(b_n^*, \beta_n) + o_p(1). $$ \hfill (3.9)

Hence,

$$ S_n(b_n^*, \beta_n) - S_n(0, 0) = T_n(b_n^*, \beta_n) - T_n(0, 0) + o_p(1). $$

### 3.3. Asymptotic linearity of $S_n(b_n^*, \beta_n)$

**Lemma 3.3.** Under the conditions of Theorem 2.1,

$$ \|S_n(b_n^*, \beta_n) - S_n(0, 0) + \gamma [(Q + V)b_0^* + V\beta_0] \| \xrightarrow{P} 0, $$ \hfill (3.10)

where

$$ \gamma = \int_0^1 - \frac{f'(F^{-1}(u))}{f(F^{-1}(u))} \varphi(u) du = - \int_{R^1} \varphi(F(z)) df(z). $$ \hfill (3.11)

**Proof.** Consider the sequence of functions $\{\varphi^{(k)}(\cdot)\}_{k=1}^{\infty}$

$$ \varphi^{(k)}(u) = \varphi \left( \frac{1}{k + 1} \right) \mathbb{I} \left[ u < \frac{1}{k} \right] + \varphi(u) \mathbb{I} \left[ \frac{i - 1}{k + 1} < u \leq \frac{i}{k + 1} \right], \quad i = 2, \ldots, k. $$ \hfill (3.12)

Then, by Lemma V.1.6.a [10], $\varphi^{(k)}$ is nondecreasing and bounded on $(0, 1)$ and

$$ \lim_{n \to \infty} \int_0^1 [\varphi^{(k)}(u) - \varphi(u)]^2 du = 0. $$ \hfill (3.13)

The function $\varphi^{(k)}$ has at most countable set $B_k$ of discontinuity points. Observe that assumption V.3 implies that $n^{-1/2} \max_{1 \leq i \leq n} \{\|w_{ni} - \bar{w}_n\| + \|v_{ni} - \bar{v}_n\|\} \xrightarrow{P} 0$. This fact together with the uniform continuity of $F$ implies that

$$ \sup_{e \in R, 1 \leq i \leq n} |F(e - n^{-1/2}(w_{ni} - \bar{w}_n)^\top b_0^* - n^{-1/2}(v_{ni} - \bar{v}_n)^\top \beta_0) - F(e)| \xrightarrow{P} 0. $$

Hence

$$ \varphi^{(k)} \left( F(e - n^{-1/2}(w_{ni} - \bar{w}_n)^\top b_0^* - n^{-1/2}(v_{ni} - \bar{v}_n)^\top \beta_0) \right) $$

converges to $\varphi^{(k)}(F(e))$, in probability, uniformly in $i = 1, \ldots, n$. It, in turn, implies that the conditional expectation

$$ \mathbb{E} \left[ \left( \varphi^{(k)} \left( F(e_{ni} - n^{-1/2}(w_{ni} - \bar{w}_n)^\top b_0^* - n^{-1/2}(v_{ni} - \bar{v}_n)^\top \beta_0) \right) - \varphi^{(k)}(F(e_{ni})) \right)^2 \right] v_{ni}, x_{ni} $$
converges to 0, in probability, uniformly in $i = 1, \ldots, n$ and $k$.

Let $S_n^{(k)}(b^*, \beta)$ and $T_n^{(k)}(b^*, \beta)$ be analogous to $S_n(b^*, \beta)$, $T_n(b^*, \beta)$ respectively, with $\varphi$ replaced with $\varphi^{(k)}$. Then we can bound the norm of the covariance matrix of $T_n^{(k)}(b^*, \beta) - T_n^{(k)}(0, 0)$ for any fixed $k$ in the following way: Denote

$$A_n^{(k)} = \mathbb{E} \left\{ \left[ T_n^{(k)}(b_n^*, \beta_n) - T_n^{(k)}(0, 0) \right]\left[ T_n^{(k)}(b_n^*, \beta_n) - T_n^{(k)}(0, 0) \right]^T \right\}.$$ 

Then

$$A_n^{(k)} = \mathbb{E} \left\{ n^{-1} \sum_{i=1}^{n} (w_{ni} - \bar{w}_n)(w_{ni} - \bar{w}_n)^T \right\} \cdot \left[ \varphi^{(k)}(F(e_{ni}) - (w_{ni} - \bar{w}_n)^T b_n^* - (v_{ni} - \bar{v}_n)^T \beta_n) - \varphi^{(k)}(F(e_{ni})) \right]^2 \right\},
$$

(3.14)

$$= n^{-1} \sum_{i=1}^{n} \mathbb{E} \left\{ (w_{ni} - \bar{w}_n)(w_{ni} - \bar{w}_n)^T \right\} \cdot \mathbb{E} \left\{ \left[ \varphi^{(k)}(F(e_{ni}) - (w_{ni} - \bar{w}_n)^T b_n^* - (v_{ni} - \bar{v}_n)^T \beta_n) - \varphi^{(k)}(F(e_{ni})) \right]^2 \right\}.$$ 

Hence,

$$\|A_n^{(k)}\| \leq \left\{ n^{-1} \sum_{i=1}^{n} (w_{ni} - \bar{w}_n)(w_{ni} - \bar{w}_n)^T - (Q + V) \right\} \cdot o(1),
$$

$$= \{\|Q + V\| + o(1) \} \cdot o(1).$$

This, together with the fact $\mathbb{E} T_n^{(k)}(0, 0) = 0$, implies

$$\|T_n^{(k)}(b_n^*, \beta_n) - T_n^{(k)}(0, 0) - \mathbb{E} T_n^{(k)}(b_n^*, \beta_n)\| \Rightarrow 0.
$$

(3.15)

Furthermore, for any fixed $k$ and for fixed $b^{0*}, \beta^0$,

$$T_n^{(k)}(b_n^*, \beta_n) - T_n^{(k)}(0, 0) + \gamma_k[(Q + V)b^{0*} + V\beta^0] \Rightarrow 0,$n

(3.16)

where

$$\gamma_k = -\int_{\mathbb{R}} \varphi^{(k)}(F(e))f'(e)de = -\int_{0}^{1} \varphi^{(k)}(u)f'(F^{-1}(u))du.$$ 

Indeed, (we put $\bar{x}_n = \bar{v}_n = 0$, for the sake of brevity)

$$n^{-1/2} \sum_{i=1}^{n} \mathbb{E} \left\{ w_{ni} \left[ \mathbb{E} \left\{ \varphi^{(k)}(F(e_{ni}) - n^{-1/2}(w_{ni}^T b_n^{0*} - n^{-1/2}v_{ni}^T \beta^0) - \varphi^{(k)}(F(e_{ni})) \right\} - \gamma_k(n^{-1/2}[w_{ni}^T b_n^{0*} + v_{ni}^T \beta^0]) \right\} \right\} \mid v_{ni}, x_{ni} \}
$$

$$= n^{-1/2} \sum_{i=1}^{n} \mathbb{E} \left\{ w_{ni} \left[ \int_{\mathbb{R}} \varphi^{(k)}(F(z))dF(z)F(z + n^{-1/2}w_{ni}^T b_n^{0*} + n^{-1/2}v_{ni}^T \beta^0) - F(z) \right] - n^{-1/2}[w_{ni}^T b_n^{0*} + v_{ni}^T \beta^0] \int_{\mathbb{R}} \varphi^{(k)}(F(z))f'(z)dz \right\}
$$

$$= n^{-1/2} \sum_{i=1}^{n} \mathbb{E} \left\{ w_{ni} \left[ \int_{\mathbb{R}} \varphi^{(k)}(F(z))dF(z) \right] \cdot d\left[ F(z + n^{-1/2}w_{ni}^T b_n^{0*} + n^{-1/2}v_{ni}^T \beta^0) - F(z) - n^{-1/2}(w_{ni}^T b_n^{0*} + v_{ni}^T \beta^0)f(z) \right] \right\} \Rightarrow 0.
Moreover, we have
\[
(\gamma_k - \gamma)^2 = \left( (\varphi_k - \varphi) - \frac{f'(F^{-1}(\cdot))}{f(F^{-1}(\cdot))} \right)^2
\]
(3.17)
\[
\leq \|\varphi_k - \varphi\|^2 - \frac{f'(F^{-1}(\cdot))}{f(F^{-1}(\cdot))} \|^2 = I(f) \|\varphi_k - \varphi\|^2 \to 0 \text{ as } k \to \infty.
\]

Using (3.16), (3.17), Lemma 3.1, Lemma 3.2, Corollary 3.1 and Lemma 3.5 in [18], we obtain that
\[
P \left( \|S_n(b_n^*, \beta_n) - S_n^{(k)}(b_n^*, \beta_n)\| > \varepsilon \right) < \varepsilon
\]
for \( \forall k > k(\varepsilon), \forall n > n(k) \), and finally we arrive at (3.10).

3.4. Uniform asymptotic quadarticity of the Jaeckel dispersion

Recall that \( \bar{a}_n = 0 \) under (A.1). Rewrite the Jaeckel dispersion in the presence of measurement errors in the form
\[
D_n(b) = \sum_{i=1}^{n} \left[ Y_{ni} - w_{ni}^T b \right] a_n(R_{ni}(b))
\]
or eventually in the form
\[
D_n(b^*, \beta) = \sum_{i=1}^{n} \left[ e_{ni} - (w_{ni} - \bar{w})^T b^* - (v_{ni} - \bar{v})^T \beta \right] a_n(R(e_{ni} - w_{ni}^T b^* - v_{ni}^T \beta)
\]
where \( b^* = b - \beta \). It is a piecewise linear, convex function of \( b \) and \( b^* \), respectively. Hence, the minimum \( \hat{\beta}_n = \arg \min_{\beta \in \mathbb{R}^p} D_n(b) \) exists, and is considered as an estimate of \( \beta \) in model (2.1).

By [17], the partial derivatives of \( D_n(b) \) exist for almost all \( b \), and where they exist, are equal to
\[
\frac{\partial}{\partial b_{nj}} D_n(b) = -n^{1/2} S_{nj}(b) = -\sum_{i=1}^{n} (w_{nj} - \bar{w}_j) a_n(Y_i - w_{ni} b), \ j = 1, \ldots, p.
\]

Otherwise speaking,
\[
\nabla D_n(b^*, \beta) = -n^{1/2} S_n(b^*, \beta) = -\sum_{i=1}^{n} (w_{ni} - \bar{w}_i) a_n(R(e_{ni} - w_{ni}^T b^* - v_{ni}^T \beta))
\]
where \( \nabla \) denotes the subgradient.

Consider the quadratic function
\[
C_n(b^*, \beta) = \frac{1}{2} \gamma b^T (Q + V) b^* - b^*^T S_n(0) + \gamma b^* V \beta + D_n(0).
\]

Then \( D_n(b) \) and \( C_n(b) \) are both convex functions and \( D_n(0) = C_n(0) \). Moreover,
\[
\nabla [D_n(b^*, \beta) - C_n(b^*, \beta)] = -n^{1/2} [S_n((b^*, \beta)) - S_n(0, 0)] + \gamma (Q + V) b^* + \gamma V \beta.
\]

Hence it follows from (3.10) that for \( b_n^* = n^{-1/2} b^{0*}, \beta_n = n^{-1/2} \beta^0 \) with \( b^{0*}, \beta^0 \in \mathbb{R}^p \) fixed that
\[
\|\nabla (D_n(n^{-1/2} b^{0*}, n^{-1/2} \beta^0) - C_n(n^{-1/2} b^{0*}, n^{-1/2} \beta^0))\| \leq 0.
\]

Using the convexity arguments in [12] (Appendix) and [29] (Convexity Lemma), we conclude that
\[ \sup \left| D_n(n^{-1/2}b^0*, n^{-1/2}\beta^0) - \frac{1}{2} \gamma b^{0*\top}(Q + V)b^{0*} + b^{0*\top}S_n(0) - \gamma b^{0*V\beta^0} + D_n(0) \right| = o_p(1), \]

where the supremum is taken over the set \( \{ \|b^{0*}\| \leq C, \|\beta^0\| \leq C \} \). Hence, following the arguments in the proof of Theorem 1 in [29], we conclude that, under the local alternative \( \beta_n = n^{-1/2}\beta^0 \),

\[
\arg\min_{b^{0*}} D_n(n^{-1/2}b^{0*}, n^{-1/2}\beta^0)
\]

is asymptotically equivalent to

\[
\arg\min_{b^{0*}} \left[ \frac{1}{2} \gamma b^{0*\top}(Q + V)b^{0*} - b^{0*\top}S_n(0) + \gamma b^{0*V\beta^0} \right].
\] (3.18)

The solution of (3.18) equals to

\[
b^{0*} = b^0 - \beta^0 = n^{1/2}(\hat{\beta}_n - \beta_n) = \gamma^{-1}(Q + V)^{-1}S_n(0, 0) - (Q + V)^{-1}V\beta^0.
\]

Hence, in the linear model with local value of regression parameter \( \beta \), when

\[ Y_{ni} = x_{ni}^\top\beta_n + e_{ni}, \beta_n = n^{-1/2}\beta^0, \]

when we observe only \( w_{ni} = x_{ni} + v_{ni} \) instead of \( x_{ni}, i = 1, \ldots, n \), the R-estimator is asymptotically normally distributed with a bias \( B = -(Q + V)^{-1}V\beta^0 \), i.e.

\[
n^{1/2}(\hat{\beta}_n - n^{-1/2}\beta^0) \xrightarrow{D} N_p(B, (Q + V)^{-1}A^2(\varphi)), \quad B = -(Q + V)^{-1}V\beta^0.
\] (3.19)

Finally, as we have already mentioned, all of the above arguments and motivations are valid when we replace \( e_{ni} \) with \( e_{ni} + u_{ni}, i = 1, \ldots, n \). This completes the proof of Theorem 2.1.

4. Numerical illustration

The following simulation study illustrates the effect of measurement errors in regressors on the finite-sample performance of R-estimates. Empirical bias (and variance) of R-estimates are computed and compared for various measurement error models. For the sake of comparison, the biases and variances are also computed for the least squares estimate (LSE) and the least absolute deviation (L1) estimate, under the same setup. Moreover, we compare the deterministic and random regressors.

All the simulations were performed in the statistical software R using standard tools and libraries. For minimization of (2.4) functions \texttt{optimize} and \texttt{optim} with initial estimate \( 0.5 \) - regression quantile were used. The random numbers generator was setup with the initial value \texttt{set.seed(15)}.

The results illustrate that the bias of R-estimate is surprisingly stable with respect to the sample size; the bias corresponding to small \( n \) is comparable to the asymptotic one derived in Theorem 2.1.

Notice that the bias of R-estimator only slightly differs from the biases of LSE and L1-estimators.
4.1. Regression line

Consider first the model of regression line

\[ Y_i = \beta_0 + x_i \beta_1 + e_i, \quad i = 1, \ldots, n \]

where the \( Y_i \) are measured accurately, while instead of \( x_i \) we observe only \( w_i = x_i + v_i, i = 1, \ldots, n \). The R-estimator of parameter \( \beta_1 \) is based on Wilcoxon scores generated by score function \( \varphi(u) = u - 1/2 \).

All the simulation results are based on 10 000 replications, parameters were chosen as \( \beta_0 = 1, \beta_1 = 2 \), and model errors \( e_i \) follow the logistic distribution. In Tables 1 and 2, the empirical bias of R-estimator based on Wilcoxon scores is compared for various sample sizes \((n = 10, \ldots, 1000)\) and with the theoretical asymptotic result \((n = \infty)\). The regressors \( x_i \) are deterministic in Table 1; they were generated from uniform \( U(-3, 9) \) distribution once for all experiment and then considered as fixed. The regressors in Table 2 are random; each time they were generated also from uniform distribution \( U(-3, 9) \). This enables to see the difference between deterministic and random regressors: The bias differs more from its asymptotic value in case of deterministic regressors than in case of random regressors; it can be caused by the slower rate of convergence. The measurement errors \( v_i \) are either uniformly or normally distributed \((i = 1, \ldots, n)\).

Table 1. Empirical bias of R-estimator for various \( n \) and measurement errors \( v_i \); nonrandom regressors \( x_i \).

<table>
<thead>
<tr>
<th>( n )</th>
<th>10</th>
<th>20</th>
<th>50</th>
<th>100</th>
<th>200</th>
<th>500</th>
<th>1000</th>
<th>( \infty )</th>
</tr>
</thead>
<tbody>
<tr>
<td>0</td>
<td>0.002</td>
<td>0.001</td>
<td>0.000</td>
<td>0.000</td>
<td>0.000</td>
<td>0.000</td>
<td>0.000</td>
<td>0.000</td>
</tr>
<tr>
<td>( U(-5, 0) )</td>
<td>-0.264</td>
<td>-0.295</td>
<td>-0.305</td>
<td>-0.297</td>
<td>-0.302</td>
<td>-0.306</td>
<td>-0.307</td>
<td>-0.296</td>
</tr>
<tr>
<td>( U(0, 9) )</td>
<td>-0.684</td>
<td>-0.727</td>
<td>-0.732</td>
<td>-0.714</td>
<td>-0.719</td>
<td>-0.727</td>
<td>-0.728</td>
<td>-0.729</td>
</tr>
<tr>
<td>( U(-3, 9) )</td>
<td>-0.982</td>
<td>-1.013</td>
<td>-1.006</td>
<td>-0.983</td>
<td>-0.986</td>
<td>-0.995</td>
<td>-0.995</td>
<td>-1.000</td>
</tr>
<tr>
<td>( \mathcal{N}(0, 1) )</td>
<td>-0.128</td>
<td>-0.148</td>
<td>-0.150</td>
<td>-0.146</td>
<td>-0.148</td>
<td>-0.151</td>
<td>-0.152</td>
<td>-0.154</td>
</tr>
<tr>
<td>( \mathcal{N}(0, 2) )</td>
<td>-0.440</td>
<td>-0.485</td>
<td>-0.485</td>
<td>-0.476</td>
<td>-0.480</td>
<td>-0.487</td>
<td>-0.488</td>
<td>-0.500</td>
</tr>
<tr>
<td>( \mathcal{N}(0, 3) )</td>
<td>-0.790</td>
<td>-0.836</td>
<td>-0.837</td>
<td>-0.819</td>
<td>-0.822</td>
<td>-0.832</td>
<td>-0.833</td>
<td>-0.857</td>
</tr>
</tbody>
</table>

Table 2. Empirical bias of R-estimator for various \( n \) and measurement errors \( v_i \); random regressors \( x_i \).

<table>
<thead>
<tr>
<th>( n )</th>
<th>10</th>
<th>20</th>
<th>50</th>
<th>100</th>
<th>200</th>
<th>500</th>
<th>1000</th>
<th>( \infty )</th>
</tr>
</thead>
<tbody>
<tr>
<td>0</td>
<td>0.004</td>
<td>0.001</td>
<td>0.000</td>
<td>0.000</td>
<td>0.000</td>
<td>0.000</td>
<td>0.000</td>
<td>0.000</td>
</tr>
<tr>
<td>( U(-5, 0) )</td>
<td>-0.283</td>
<td>-0.297</td>
<td>-0.305</td>
<td>-0.306</td>
<td>-0.307</td>
<td>-0.309</td>
<td>-0.309</td>
<td>-0.296</td>
</tr>
<tr>
<td>( U(0, 9) )</td>
<td>-0.711</td>
<td>-0.722</td>
<td>-0.728</td>
<td>-0.730</td>
<td>-0.730</td>
<td>-0.732</td>
<td>-0.732</td>
<td>-0.729</td>
</tr>
<tr>
<td>( \mathcal{N}(0, 1) )</td>
<td>-0.998</td>
<td>-1.000</td>
<td>-0.999</td>
<td>-1.000</td>
<td>-0.999</td>
<td>-1.000</td>
<td>-1.000</td>
<td>-1.000</td>
</tr>
<tr>
<td>( \mathcal{N}(0, 2) )</td>
<td>-0.462</td>
<td>-0.481</td>
<td>-0.487</td>
<td>-0.489</td>
<td>-0.491</td>
<td>-0.492</td>
<td>-0.492</td>
<td>-0.500</td>
</tr>
<tr>
<td>( \mathcal{N}(0, 3) )</td>
<td>-0.813</td>
<td>-0.830</td>
<td>-0.833</td>
<td>-0.835</td>
<td>-0.837</td>
<td>-0.837</td>
<td>-0.838</td>
<td>-0.857</td>
</tr>
</tbody>
</table>

Table 3 compares empirical bias and variance (in parenthesis) of R-estimator based on Wilcoxon scores, of LSE and \( L_1 \)-estimate for the sample size \( n = 50 \) and when regressors \( x_i \) are random, generated from uniform \( U(-3, 9) \) distribution; model errors \( e_i \) are generated from normal, logistic, Laplace, Pareto with parameter \( \alpha = 0.9 \) and Cauchy distributions. The measurement errors \( v_i \) follow various distributions, similarly as in Tables 1 and 2.

4.2. Model of two regressors

Consider the model

\[ Y_i = \beta_0 + x_{i1} \beta_1 + x_{i2} \beta_2 + e_i, \quad i = 1, \ldots, n \]
where again the $Y_i$ are measured accurately, but instead of $\mathbf{x}_i$, we observe only $\mathbf{w}_i = \mathbf{x}_i + \mathbf{v}_i$, $i = 1, \ldots, n$. The R-estimator of parameter $\beta = (\beta_1, \beta_2)^T$ is based on Wilcoxon scores generated by score function $\varphi(u) = u - 1/2$.

Here we chose $n = 50$, parameters $\beta_0 = 1, \beta_1 = 2, \beta_2 = 1$, random regressors $\mathbf{x}_i = (x_{i1}, x_{i2})^T$ are generated from 2-dimensional normal distributions $N_2(\mu, \mathbf{S}_\nu)$, $\nu = 1, 2, 3$, where $\mu = (0, 1)^T$ and

$$\mathbf{S}_1 = \begin{pmatrix} 4 & 0.5 \\ 0.5 & 2 \end{pmatrix}, \quad \mathbf{S}_2 = \begin{pmatrix} 2 & 0.2 \\ 0.2 & 2 \end{pmatrix}, \quad \mathbf{S}_3 = \begin{pmatrix} 1 & 0.9 \\ 0.9 & 1 \end{pmatrix}.$$ 

Table 4 compares empirical bias and variance (in parentheses) of R-estimator based on Wilcoxon scores, with those of the LSE and $L_1$-estimator for various distributions of the measurement errors $\mathbf{v}_i$ and model errors $e_i$.

<table>
<thead>
<tr>
<th>$v_i \backslash e_i$</th>
<th>normal</th>
<th>logistic</th>
<th>Laplace</th>
<th>Pareto</th>
<th>Cauchy</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\beta_1$</td>
<td>0.002(0.610)</td>
<td>0.004(0.527)</td>
<td>-0.002(0.254)</td>
<td>0.000(0.416)</td>
<td>0.018(0.672)</td>
</tr>
<tr>
<td>$\beta_2$</td>
<td>0.002(0.274)</td>
<td>0.010(0.708)</td>
<td>-0.002(0.249)</td>
<td>0.000(1.191)</td>
<td>0.018(0.526)</td>
</tr>
</tbody>
</table>

Table 3. Empirical bias (variance) of R-estimator, LSE and $L_1$-estimator for various measurement errors $v_i$ and model errors $e_i$; $n = 50$.

<table>
<thead>
<tr>
<th>$v_i \backslash e_i$</th>
<th>normal</th>
<th>logistic</th>
<th>Laplace</th>
<th>Pareto</th>
<th>Cauchy</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\beta_1$</td>
<td>-0.399(0.155)</td>
<td>-0.398(0.438)</td>
<td>-0.396(0.214)</td>
<td>-0.401(0.422)</td>
<td>-0.404(0.568)</td>
</tr>
<tr>
<td>$\beta_2$</td>
<td>-0.401(0.145)</td>
<td>-0.396(0.466)</td>
<td>-0.394(0.283)</td>
<td>-7.137(6.000)</td>
<td>22.62(400000)</td>
</tr>
</tbody>
</table>

Table 4. Empirical bias (variance) of R-estimator, LSE and $L_1$-estimator for various measurement errors $v_i$ and model errors $e_i$; $n = 50$. 

We have also computed R-estimates generated by other score functions, e.g. van der Waerden, median; also another simulation design was considered – various sample sizes $n$, values of the parameters, distributions of regressors, measurement errors $v_i$ and $u_i$ and model errors. It is of interest that the results for corresponding R-estimates are quite similar to those presented in the previous tables.

The simulation study confirms that R-estimates in measurement error models are biased, as well as other usual estimates. The bias is relatively stable with respect to the sample size and to distribution of model errors. The R-estimates provide meaningful results as long as the $e_i$ have a finite Fisher information; even under the normal errors are their empirical variances only slightly greater than that of LSE. The bias and other properties of R-estimates are comparable with those of the least squares and of $L_1$ estimates unless the distribution of model errors $e_i$ is heavy, where the LSE fails. Generally, the reduction of the bias is rather a matter of measurement precision, of calibration and repeated measurements.

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