

# Robust inference under the beta regression model with application to health care studies

Abhik Ghosh

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## Abstract

Data on rates, percentages, or proportions arise frequently in many different applied disciplines like medical biology, health care, psychology, and several others. In this paper, we develop a robust inference procedure for the beta regression model, which is used to describe such response variables taking values in  $(0, 1)$  through some related explanatory variables. In relation to the beta regression model, the issue of robustness has been largely ignored in the literature so far. The existing maximum likelihood-based inference has serious lack of robustness against outliers in data and generate drastically different (erroneous) inference in the presence of data contamination. Here, we develop the robust minimum density power divergence estimator and a class of robust Wald-type tests for the beta regression model along with several applications. We derive their asymptotic properties and describe their robustness theoretically through the influence function analyses. Finite sample performances of the proposed estimators and tests are examined through suitable simulation studies and real data applications in the context of health care and psychology. Although we primarily focus on the beta regression models with a fixed dispersion parameter, some indications are also provided for extension to the variable dispersion beta regression models with an application.

## Keywords

Robustness, beta regression model, rates and proportions data, minimum density power divergence estimator, Wald-type tests

## 1 Introduction

In many biological experiments, medical research including health care studies and psychology, survey research in sociology and marketing, and several other applied sciences, we often come across data on rates, ratios, percentages or proportions, taking values in the unit interval  $(0, 1)$ . Examples include the “body fat percentage” or any similar health condition measured in percentage, health assessment questionnaire (HAQ) data or similar rating data, accuracy percentage of any treatment in clinical trials, experimental scores measuring stress, depression, etc. in psychology, proportion of a certain group of patients (for some particular disease) in a region and many more. Such data can be modeled individually by a beta distribution having support  $(0, 1)$ . However, in order to better understand the underlying data-generating mechanism for more detailed inference, it is often required to relate their values with some other associated explanatory variables through a suitable regression structure; this also enables us to do prediction. The beta regression model (BRM) is designed to help in this situation, which models a response variable  $y$  taking values in  $(0, 1)$  through a set of explanatory variables  $x$ .

There are several recent specifications of the BRM, e.g. see Paolino,<sup>1</sup> Kieschnick and McCullough,<sup>2</sup> Ferrari and Cribari-Neto,<sup>3</sup> and Vasconcellos and Cribari-Neto,<sup>4</sup> among others. In this paper, we follow the most popular specification provided by Ferrari and Cribari-Neto.<sup>3</sup> This is because this specification (i) models the “mean” of the response variable on  $(0, 1)$  to depend on a linear combination of available covariates through a suitable link function, (ii) is closely related to the popular class of generalized linear models,<sup>5</sup> (iii) allows many different possible

Indian Statistical Institute, Kolkata, India

### Corresponding author:

Abhik Ghosh, Indian Statistical Institute, 203, B.T. Road, Kolkata 700108, West Bengal, India.

Email: abhianik@gmail.com

link functions to model various structures within the data, and (iv) the inference methodologies are well developed for this specification and are available in the standard statistical software R (package *betareg*) for practitioners.

Mathematically, suppose  $y_1, \dots, y_n$  are  $n$  independent responses each taking value in  $(0, 1)$  and are associated with  $p$ -dimensional covariate values  $\mathbf{x}_1, \dots, \mathbf{x}_n$ , respectively. Then, in the beta regression model (BRM) of Ferrari and Cribari-Neto,<sup>3</sup> each  $y_i$  follows a beta distribution having density  $f(\cdot; \mu_i, \phi)$ , where

$$f(y; \mu, \phi) = \frac{1}{B(\mu\phi, (1-\mu)\phi)} y^{\mu\phi-1} (1-y)^{(1-\mu)\phi-1}, \quad 0 < y < 1 \quad (1)$$

with  $B(\cdot, \cdot)$  being the (complete) beta function, and  $E(y_i) = \mu_i \in (0, 1)$  is related to the (given)  $i$ th value  $\mathbf{x}_i \in \mathbb{R}^p$  of the explanatory variables through a suitable link function  $g$  (defined on  $(0, 1)$ ). Note that,  $\text{Var}(y_i) = \frac{\mu_i(1-\mu_i)}{1+\phi}$ . Given  $g$ , the BRM of Ferrari and Cribari-Neto<sup>3</sup>—with fixed precision parameter  $\phi$  or dispersion parameter  $\sigma^2 = (1 + \phi)^{-1}$ —assumes the regression structure

$$y_i \sim f(y_i; \mu_i, \phi) \text{ independently, with } g(\mu_i) = \mathbf{x}_i^T \boldsymbol{\beta}, \quad i = 1, \dots, n \quad (2)$$

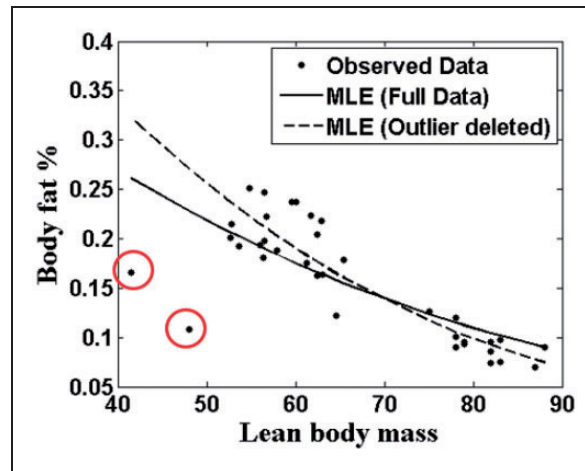
where  $\boldsymbol{\beta} \in \mathbb{R}^p$  is the vector of unknown regression coefficients. Our objective then is to make inference about the parameter of interest  $\boldsymbol{\theta} = (\boldsymbol{\beta}^T, \phi)^T \in \mathbb{R}^{p+1}$  based on the available data  $\{(y_i, \mathbf{x}_i) : i = 1, \dots, n\}$ . This BRM has later been extended to cover the cases of heterogeneous precision parameter  $\phi_i$  (or, dispersion parameter  $\sigma_i^2$ ) for  $y_i$  by Smithson and Verkuilen,<sup>6</sup> Simas et al.,<sup>7</sup> Rocha and Simas,<sup>8</sup> Cribari-Neto and Souza,<sup>9</sup> and Melo et al.,<sup>10</sup> where  $\phi_i$  depends on another set of covariates through a (possibly different) regression structure. To keep a clear focus in our presentations, we restrict our attention primarily to the fixed dispersion BRM (2) in the present paper. Our methodology, however, is not critically dependent on the fixed dispersion assumption, and we also briefly indicate the possible extension to a general class of non-linear, variable dispersion BRMs towards the end of the paper. Indeed, our illustrations will show that the extension of the proposed methodology to such a general class of BRMs has exactly the same structure and robustness implications in relation to the fixed dispersion results presented in this paper.

The BRM (2) has become very useful in several recent applications, since it can also be applied to data within any finite interval. If  $y$  takes values in any other open interval, say  $(a, b)$ , we can apply the BRM (2) with the transformed response  $\frac{y-a}{b-a} \in (0, 1)$ . Further, if the response  $y$  also takes values in the end-points 0 and 1, rather than using sophisticated and complicated modifications, we can simply apply the BRM (2) with the widely used ad hoc transformation  $\frac{1}{n}[y(n-1) + 0.5]$ ,  $n$  being the sample size.<sup>6</sup>

The existing inference procedures under the BRM (2) are primarily based on the classical maximum likelihood approach. The point estimator of  $\boldsymbol{\theta} = (\boldsymbol{\beta}^T, \phi)^T$  is obtained by maximizing the likelihood function with respect to  $\boldsymbol{\theta}$ , generating the maximum likelihood estimator (MLE), and any hypothesis testing problem can be solved by the likelihood ratio test or the Wald test based on the MLE; see Ferrari and Cribari-Neto<sup>3</sup> for more details. The R package “*betareg*” provides the inferential solution for the BRM (2) based on this standard maximum likelihood approach, which possesses many asymptotic optimality properties. However, a serious problem with maximum likelihood-based inference is the high degree of sensitivity to potential outliers in the data. This lack of robustness often leads to drastically different (erroneous) inference in the presence of even a small amount of data contamination. Since such outliers are not uncommon in practical datasets, we need to be very cautious before using maximum likelihood-based inference (and also using the R package “*betareg*”). To illustrate this issue, let us present a motivating example from an Australian health care study.

## A motivating example (AIS Data)

Consider the data on health measurements of several athletes collected at the Australian Institute of Sport (AIS), which is publicly available in the R package “*sn*”. Bayes et al.<sup>11</sup> have recently studied a subset of these data corresponding to the 37 rowing athletes to predict their body fat percentages (BFP) from their lean body masses (LBM) using Bayesian inference. Since the response variable BFP takes values within  $(0, 1)$ , we here fit the BRM (2) with response  $y = \text{BFP}$ , covariate  $\mathbf{x}_i = (1, \text{LBM})^T$  and a logit link function,  $\text{logit}(E[\text{BFP}]) = \beta_1 + \beta_2 \text{LBM}$ . Applying the existing maximum likelihood procedure using “*betareg*”, the MLE of the parameter of interest  $\boldsymbol{\theta} = (\beta_1, \beta_2, \phi)^T$  turns out to be  $(0.097, -0.027, 95.472)^T$ . Further, applying the existing Wald test based on this MLE, the  $p$ -values of the significance of two regression coefficients become 0.699 and 0 respectively, which indicates that the intercept component ( $\beta_1$ ) is not significant in the model.



**Figure 1.** The AIS Data along with the fitted lines based on the MLE for the full data and the outlier deleted data. (The two outlying observations are marked with red circles.)

However, by plotting the data (see Figure 1), one can clearly see that there are two outlying observations as also noticed by Bayes et al.;<sup>11</sup> the fitted line based on the above MLE does not yield a good fit to the bulk of the data in the presence of these two outliers. In fact, if we again compute the MLE of the parameter  $\theta$  after removing these two outliers, the resulting estimate becomes  $(0.838, -0.038, 246.305)^T$ , which drastically differs from the previous MLE based on the full data. The change in the fitted line is clearly visible in Figure 1 and the estimate of  $\phi$  changes substantially! Further, after deleting these two outliers, both the  $p$ -values of the MLE-based Wald test for testing the significance of the regression coefficients become 0. Thus, only these two outliers completely hide the significance of  $\beta_1$  reversing the inferential conclusions.  $\square$

As we have seen in the above example, few outliers in a dataset can lead to completely wrong inference through the existing likelihood procedures under the BRM. Several other authors have also recently noticed this nonrobust behavior of the MLE and the instability of the related inference against the outlying observations in the BRMs;<sup>3,12,13</sup> they have developed some diagnostic tools to identify such influential observations or outliers in a BRM and suggested their deletion before doing maximum likelihood-based inference. Although this solution with prior outlier detection is feasible (although rarely advisable) for simple and small datasets, it is quite difficult and needs several additional analyses for most complicated datasets including the big or high-dimensional datasets of recent era. A robust inference procedure that can automatically take care of these outliers to successfully yield stable results is much more logical, efficient and useful in all such complicated cases. However, unlike other inferential set-ups, there exists no such robust inference procedure for the recently developed beta regression model. The only related work is the one by Bayes et al.<sup>11</sup> who have proposed to solve this issue for the BRM under Bayesian paradigm through the use of a modified distribution in place of the simple beta distribution; but it does not really address the nonrobustness problem of the MLE-based inference with respect to the simpler specified BRM.

In this paper, we develop a robust inference procedure for the BRM (2) without changing its original distributional form. Among several approaches of robust inference, we follow the minimum divergence approach where we quantify the discrepancy between the data and the parametric model through a statistical divergence measure and minimize it to estimate the unknown parameters. In particular, we consider the density power divergence (DPD) of Basu et al.,<sup>14</sup> because the resulting estimator has become very popular in recent times due to its high asymptotic efficiency along with strong robustness properties. It has also been applied to many real-life inference problems; see Section 2.1 and Basu et al.<sup>15</sup> for additional details. We develop the robust minimum density power divergence estimator (MDPDE) for the BRM (2) along with its asymptotic properties in Section 2. Based on the proposed MDPDE, we develop a robust Wald-type hypothesis testing procedure in Section 3 and derive its asymptotic properties. We also theoretically illustrate the robustness of both the proposed estimator and the testing procedure through suitable influence function analyses. Finite sample performances of the proposed inference procedures are examined through suitable simulation studies in Section 4. In Section 5, our proposals are applied to reanalyze the motivating example along with two additional real data examples from health care studies (including psychology). Extension of the proposed methodology for performing robust inference under a general

class of (nonlinear) variable dispersion BRM is briefly discussed in Section 6 and illustrated through a real data application. Finally, the paper ends with some concluding remarks in Section 7.

## 2 Robust minimum density power divergence estimators

### 2.1 Background

The density power divergence (DPD) measure between two densities  $f_1$  and  $f_2$  (with respect to some common dominating measure) is defined in terms of a tuning parameter  $\alpha \geq 0$  as<sup>14</sup>

$$\begin{aligned} d_\alpha(f_1, f_2) &= \int f_2^{1+\alpha} - \frac{1+\alpha}{\alpha} \int f_1 f_2^\alpha + \frac{1}{\alpha} \int f_1^{1+\alpha}, & \text{if } \alpha > 0; \\ d_0(f_1, f_2) &= \lim_{\alpha \rightarrow 0} d_\alpha(f_1, f_2) = \int f_1 \log\left(\frac{f_1}{f_2}\right) \end{aligned} \quad (3)$$

Note that the DPD measure at  $\alpha=0$  coincides with the famous likelihood disparity, minimization of which is known to generate the MLE. The DPD family connects the likelihood disparity (at  $\alpha=0$ ) to the  $L_2$ -divergence (at  $\alpha=1$ ) smoothly through the tuning parameter  $\alpha$ . For the sake of completeness and a better understanding, let us start by recalling the minimum DPD estimation under the independent and identically distributed (iid) set-up.

For  $n$  iid observations  $Y_1, \dots, Y_n$  modeled by a parametric family of densities  $\mathcal{F}_\theta = \{f_\theta : \theta \in \Theta \subseteq \mathbb{R}^k\}$ , the MDPDE is obtained by minimizing the estimated DPD measure (3) between the observed data (at  $f_1$ ) and the model density  $f_\theta$  (at  $f_2$ ), or, equivalently by minimizing the quantity

$$\int f_\theta^{1+\alpha} - \frac{1+\alpha}{\alpha} \int f_\theta^\alpha dG_n = \int f_\theta^{1+\alpha} - \frac{1+\alpha}{\alpha} \frac{1}{n} \sum_{i=1}^n f_\theta^\alpha(Y_i)$$

with  $G_n$  being the empirical distribution function based on the observed data.<sup>14,15</sup> Under suitable differentiability assumptions, the estimating equation of  $\theta$  is given by

$$\frac{1}{n} \sum_{i=1}^n \mathbf{u}_\theta(Y_i) f_\theta^\alpha(Y_i) - \int \mathbf{u}_\theta f_\theta^{1+\alpha} = 0$$

where  $\mathbf{u}_\theta = \nabla \ln f_\theta$  is the score function with  $\nabla$  representing gradient with respect to  $\theta$ . Note that, at  $\alpha=0$ , this MDPDE estimating equation coincides with the estimating (score) equation of the MLE. The MDPDE at  $\alpha>0$  yields a generalization of the MLE which down-weights the effect of the outlying observations in the estimating equation by  $\alpha$ -th power of the model density and hence is expected to be more robust. This MDPDE has become very popular in recent days, because (i) it does not need nonparametric kernel estimation unlike many other divergences, (ii) it is a robust estimator having high asymptotic efficiency at properly chosen  $\alpha$ , and (iii) it can be obtained from a simple unbiased estimating equation along with an underlying objective function which helps to avoid the problem of multiple roots.

However, in general, our data for the BRM (2) are NOT iid, and hence the above approach cannot be applied directly. This is because we generally do not make any distributional assumptions on the covariates  $\mathbf{x}_i$ s and treat them as fixed (given) so that, for each  $i$ ,  $y_i \sim \text{Beta}(\mu_i \phi, (1 - \mu_i) \phi)$  with density given by (2). Thus, each  $y_i$  is independent but not identically distributed. Recently, Ghosh and Basu<sup>16</sup> have proposed an extension of the MDPDE for the general independent but nonhomogeneous set-up by considering the average DPD measure over different distributions. Ghosh and Basu<sup>17</sup> have applied this extended approach to develop robust inference for a simple class of canonical generalized linear models (GLMs) with fixed designs including normal, Poisson and logistic regressions; Ghosh<sup>18</sup> has also applied it to an exponential regression model to propose a robust estimator of the tail index. However, unfortunately, the class of GLMs considered in Ghosh and Basu<sup>17</sup> does not directly cover our BRM (2). So, in this paper, we further extend this approach to develop a robust estimator for the BRM (2) with fixed covariates (design). For the sake of completeness, necessary background concepts, assumptions and results from Ghosh and Basu<sup>16</sup> are presented in the online supplement.

## 2.2 The MDPDE for the beta regression model

Consider the BRM set-up as described in Section 1. Let us assume that the responses  $y_1, \dots, y_n$  are independent but  $y_i \sim g_i$  for each  $i = 1, \dots, n$ , where  $g_i$ s are potentially different true densities of  $y_i$ s, respectively, depending on  $x_i$ s. We model  $g_i$  by the BRM (2), i.e. by the model density  $f_i(\cdot, \theta) \equiv \text{Beta}(\mu_i\phi, (1 - \mu_i)\phi)$  density. The unknown parameter of interest is  $\theta = (\beta^T, \phi)^T$ , which is common across the densities. Following Ghosh and Basu,<sup>16</sup> we define the MDPDE of  $\theta$  under the BRM (2) as the minimizer of the average DPD measure with tuning parameter  $\alpha \geq 0$  given by

$$n^{-1} \sum_{i=1}^n d_\alpha(\hat{g}_i(\cdot), f_i(\cdot, \theta)) \quad (4)$$

where  $\hat{g}_i$  is an estimate of  $g_i$  based on the given data. Since the DPD measure is a proper statistical divergence, the resulting minimizer is clearly Fisher consistent for  $\theta$ . For the present case of BRM (2), since we have only one observation from each  $g_i$ , a simple estimate of it is given by the degenerate distribution at  $y_i$  for any  $i = 1, \dots, n$ . Hence, after some simplification, the minimizer of (4) is seen to be the minimizer of the simpler objective function (equation (1) of the online supplement)

$$H_{n,\alpha}(\theta) = n^{-1} \sum_{i=1}^n \left[ K_{i,\alpha}(\theta) - \frac{1+\alpha}{\alpha} f_i(y_i, \theta)^\alpha \right] \quad (5)$$

with  $K_{i,\alpha}(\theta) = \frac{B((1+\alpha)\mu_i\phi - \alpha, (1+\alpha)(1-\mu_i)\phi - \alpha)}{B(\mu_i\phi, (1-\mu_i)\phi)^\alpha}$ . We need to minimize this objective function  $H_{n,\alpha}(\theta)$  with respect to  $\theta = (\beta^T, \phi)^T$  to obtain its MDPDE with tuning parameter  $\alpha$ , say  $\hat{\theta}_{n,\alpha} = (\hat{\beta}_{n,\alpha}^T, \hat{\phi}_{n,\alpha})^T$ . Note that, the above objective function  $H_{n,\alpha}(\theta)$  becomes  $[1 - \log\text{-likelihood}]$  as  $\alpha \rightarrow 0$  and hence the proposed MDPDE at  $\alpha=0$  coincides with the usual MLE of Ferrari and Cribari-Neto,<sup>3</sup> which is known to be nonrobust but fully efficient. Further, the MDPDE at  $\alpha=1$  coincides with the minimum  $L_2$ -distance estimator which is known to be highly robust but inefficient under any general model. Hence, the tuning parameter  $\alpha$  in the proposed MDPDE under the BRM is expected to yield a trade-off between robustness and efficiency of the estimator.

Equivalently, we can also obtain the MDPDE  $\hat{\theta}_{n,\alpha} = (\hat{\beta}_{n,\alpha}^T, \hat{\phi}_{n,\alpha})^T$  by solving the estimating equations obtained by differentiating the objective function  $H_{n,\alpha}(\theta)$  with respect to  $\theta = (\beta^T, \phi)^T$ . For the BRM, these estimating equations simplify to (from equation (2) of the online supplement)

$$\sum_{i=1}^n \left[ \gamma_{1,i}^{(\alpha)}(\theta) - (y_{1,i}^* - \mu_{1,i}^*) \frac{\phi}{g'(\mu_i)} f_i(y_i, \theta)^\alpha \right] x_i = \mathbf{0}_p, \quad (6)$$

$$\sum_{i=1}^n \left[ \gamma_{2,i}^{(\alpha)}(\theta) - \left\{ \mu_i (y_{1,i}^* - \mu_{1,i}^*) + (y_{2,i}^* - \mu_{2,i}^*) \right\} f_i(y_i, \theta)^\alpha \right] = 0, \quad (7)$$

where  $\mathbf{0}_p$  is the zero vector of length  $p$ ,  $g'$  denotes the derivative of  $g$  and

$$\begin{aligned} y_{1,i}^* &= \log \frac{y_i}{1-y_i}, & \mu_{1,i}^* &= E(y_{1,i}^*) = \psi(\mu_i\phi) - \psi((1-\mu_i)\phi), \\ y_{2,i}^* &= \log(1-y_i), & \mu_{2,i}^* &= E(y_{2,i}^*) = \psi((1-\mu_i)\phi) - \psi(\phi), \\ \gamma_{1,i}^{(\alpha)}(\theta) &= \left( \psi(a_{i,\alpha}) - \psi(b_{i,\alpha}) - \mu_{1,i}^* \right) \frac{\phi K_{i,\alpha}(\theta)}{g'(\mu_i)}, \\ \gamma_{2,i}^{(\alpha)}(\theta) &= \left[ \mu_i \left( \psi(a_{i,\alpha}) - \psi(b_{i,\alpha}) - \mu_{1,i}^* \right) + \left( \psi(b_{i,\alpha}) - \psi(a_{i,\alpha} + b_{i,\alpha}) - \mu_{2,i}^* \right) \right] K_{i,\alpha}(\theta) \end{aligned}$$

with  $a_{i,\alpha} = (1+\alpha)\mu_i\phi - \alpha$ ,  $b_{i,\alpha} = (1+\alpha)(1-\mu_i)\phi - \alpha$  and  $\psi(\cdot)$  being the digamma function. Clearly, the estimating equations are unbiased at the model for any  $\alpha \geq 0$ . Also, at  $\alpha=0$ , we have  $\gamma_{1,i}^{(\alpha)}(\theta) = 0 = \gamma_{2,i}^{(\alpha)}(\theta)$  for all  $i = 1, \dots, n$  and these MDPDE estimating equations then coincide with the MLE estimating (score) equations as expected.

The asymptotic distribution of this proposed MDPDE can be derived from the general results of Ghosh and Basu<sup>16</sup> under Assumptions (A1)–(A7) of their paper, also presented in the online supplement. In particular,



whenever the model assumption (2) holds with true parameter value  $\theta_0$ , i.e.  $g_i(\cdot) = f_i(\cdot, \theta_0)$  for all  $i$ , we have the following from Result R1 of the online supplement.

- (1) There exists a consistent sequence  $\hat{\theta}_{n,\alpha}$  of roots to the estimating equations (6) and (7).
- (2) Asymptotically  $\Omega_n(\theta_0)^{-1/2} \Psi_n(\theta_0) [\sqrt{n}(\hat{\theta}_{n,\alpha} - \theta_0)] \sim N_{p+1}(\mathbf{0}_{p+1}, \mathbf{I}_{p+1})$ , where  $\mathbf{I}_{p+1}$  is identity matrix of order  $(p+1)$  and

$$\Psi_n(\theta) = \frac{1}{n} \sum_{i=1}^n \begin{bmatrix} \gamma_{11,i}^{(\alpha)}(\theta) \mathbf{x}_i \mathbf{x}_i^T & \gamma_{12,i}^{(\alpha)}(\theta) \mathbf{x}_i \\ \gamma_{12,i}^{(\alpha)}(\theta) \mathbf{x}_i^T & \gamma_{22,i}^{(\alpha)}(\theta) \end{bmatrix},$$

$$\Omega_n(\theta) = \frac{1}{n} \sum_{i=1}^n \begin{bmatrix} \{\gamma_{11,i}^{(2\alpha)}(\theta) - \gamma_{1,i}^{(\alpha)}(\theta)^2\} \mathbf{x}_i \mathbf{x}_i^T & \{\gamma_{12,i}^{(2\alpha)}(\theta) - \gamma_{1,i}^{(\alpha)}(\theta) \gamma_{2,i}^{(\alpha)}(\theta)\} \mathbf{x}_i \\ \{\gamma_{12,i}^{(2\alpha)}(\theta) - \gamma_{1,i}^{(\alpha)}(\theta) \gamma_{2,i}^{(\alpha)}(\theta)\} \mathbf{x}_i^T & \{\gamma_{22,i}^{(2\alpha)}(\theta) - \gamma_{2,i}^{(\alpha)}(\theta)^2\} \end{bmatrix}$$

with explicit forms of  $\gamma_{jk,i}^{(\alpha)}(\theta)$  being given by

$$\gamma_{11,i}^{(\alpha)}(\theta) = \frac{\phi^2 K_{i,\alpha}(\theta)}{g'(\mu_i)^2} \left[ \psi_1(a_{i,\alpha}) + \psi_1(b_{i,\alpha}) + (\psi(a_{i,\alpha}) - \psi(b_{i,\alpha}) - \mu_{1,i}^*)^2 \right]$$

$$\gamma_{12,i}^{(\alpha)}(\theta) = \frac{\phi K_{i,\alpha}(\theta)}{g'(\mu_i)} \left[ \mu_i \left\{ \psi_1(a_{i,\alpha}) + \psi_1(b_{i,\alpha}) + (\psi(a_{i,\alpha}) - \psi(b_{i,\alpha}) - \mu_{1,i}^*)^2 \right\} \right. \\ \left. + \left\{ -\psi_1(b_{i,\alpha}) + (\psi(a_{i,\alpha}) - \psi(b_{i,\alpha}) - \mu_{1,i}^*)(\psi(b_{i,\alpha}) - \psi(a_{i,\alpha} + b_{i,\alpha}) - \mu_{2,i}^*) \right\} \right]$$

$$\gamma_{22,i}^{(\alpha)}(\theta) = K_{i,\alpha}(\theta) \left[ \mu_i^2 \left\{ \psi_1(a_{i,\alpha}) + \psi_1(b_{i,\alpha}) + (\psi(a_{i,\alpha}) - \psi(b_{i,\alpha}) - \mu_{1,i}^*)^2 \right\} \right. \\ \left. + 2\mu_i \left\{ -\psi_1(b_{i,\alpha}) + (\psi(a_{i,\alpha}) - \psi(b_{i,\alpha}) - \mu_{1,i}^*)(\psi(b_{i,\alpha}) - \psi(a_{i,\alpha} + b_{i,\alpha}) - \mu_{2,i}^*) \right\} \right. \\ \left. + \left\{ \psi_1(b_{i,\alpha}) - \psi_1(a_{i,\alpha} + b_{i,\alpha}) + (\psi(b_{i,\alpha}) - \psi(a_{i,\alpha} + b_{i,\alpha}) - \mu_{2,i}^*)^2 \right\} \right]$$

and  $\psi_1$  being the trigamma function. The required conditions (A1)–(A7) of Ghosh and Basu<sup>16</sup> can be verified to hold under mild boundedness conditions on the given covariate values (fixed design). However, the form of the above asymptotic variance matrix indicates that, given any fixed design, the asymptotic relative efficiency of the proposed MDPDE decreases as  $\alpha$  increases but this loss in efficiency is not significant at small positive values of  $\alpha$ . We will verify this property empirically again in Section 4.1; but this small loss in efficiency leads to increased robustness of the proposed estimator over the nonrobust MLE which we justify through the influence function analysis in the next subsection.

## 2.3 Influence function of the MDPDE under the BRM

The influence function (IF) is a classical tool to measure the theoretical robustness property of any estimator under the iid set-up.<sup>19</sup> It measures the asymptotic bias due to infinitesimal contamination in the data. The concept has been suitably extended and applied to the case of nonhomogeneous observations by Ghosh and Basu,<sup>16,17</sup> Huber,<sup>20</sup> Aerts and Haesbroeck,<sup>21</sup> where the corresponding statistical functional and the IF both depend on the sample size  $n$  (unlike the iid case). Note that, for such nonhomogeneous cases the contamination can be in any of the distributions indexed by  $i$  or in all of them. We use this concept to illustrate the robustness of our proposed MDPDE under the BRM.

Assuming  $G_i$  to be the true distribution function of  $y_i$  corresponding to the density  $g_i$  for each  $i$ , the statistical functional corresponding to the MDPDE of  $\theta$  under the BRM (2) is defined as

$$T_\alpha(G_1, \dots, G_n) = \arg \min_{\theta} n^{-1} \sum_{i=1}^n d_\alpha(g_i(\cdot), f_i(\cdot, \theta)) \quad (8)$$

whenever the minimum exists. This is a Fisher consistent functional at the assumed BRM by the definition of the DPD measure. Suppose first, for simplicity, the contamination is in only the  $i_0$ th distribution through

$G_{i_0,\epsilon} = (1 - \epsilon)G_{i_0} + \epsilon \wedge_{t_{i_0}}$ , where  $\epsilon$  is the contamination proportion and  $\wedge_{t_{i_0}}$  is the degenerate distribution at the contamination point  $t_{i_0}$ . The corresponding (first order) influence function (IF) of the proposed MDPDE functional  $T_\alpha$  is defined as

$$\begin{aligned} \mathcal{IF}(t_{i_0}, T_\alpha; G_1, \dots, G_n) &= \left| \frac{\partial T_\alpha(G_1, \dots, G_{i_0,\epsilon}, \dots, G_n)}{\partial \epsilon} \right|_{\epsilon=0} \\ &= \lim_{\epsilon \downarrow 0} \frac{T_\alpha(G_1, \dots, G_{i_0,\epsilon}, \dots, G_n) - T_\alpha(G_1, \dots, G_n)}{\epsilon} \end{aligned}$$

Note that, whenever this IF is bounded in  $t_{i_0}$ , the asymptotic bias due to infinitesimal contamination at  $G_{i_0}$  remains bounded, implying the robustness of the corresponding estimator. On the other hand, if this IF is unbounded in  $t_{i_0}$ , then the same bias may tend to infinity for distant contaminations implying the nonrobust nature of the estimator.

For our beta regression model with  $g_i(\cdot) = f_i(\cdot, \theta)$  for all  $i$ , some calculations, based on the general Result R2(i) of the online supplement, yield the simplified form of the above IF as given by

$$\mathcal{IF}(t_{i_0}, T_\alpha; F_1, \dots, F_n) = \Psi_n(\theta)^{-1} \left[ \begin{aligned} & \left( t_{1,i_0}^* - \mu_{1,i_0}^* \right) \frac{\phi}{g'(\mu_{i_0})} f_{i_0}(t_{i_0}, \theta)^\alpha - \gamma_{1,i_0}^{(\alpha)}(\theta) \\ & \left\{ \mu_i \left( t_{1,i_0}^* - \mu_{1,i_0}^* \right) + \left( t_{2,i_0}^* - \mu_{2,i_0}^* \right) \right\} f_{i_0}(t_{i_0}, \theta)^\alpha - \gamma_{2,i_0}^{(\alpha)}(\theta) \end{aligned} \right]$$

where  $t_{1,i_0}^* = \log \frac{t_{i_0}}{1-t_{i_0}}$ ,  $t_{2,i_0}^* = \log(1-t_{i_0})$  and  $F_i$  is the distribution function of  $f_i(\cdot, \theta)$  for each  $i = 1, \dots, n$ , respectively. Clearly this IF of the proposed MDPDE is bounded for all  $\alpha > 0$  but unbounded at  $\alpha = 0$ . This implies that the proposed MDPDE with  $\alpha > 0$  is robust against contamination in data, whereas that at  $\alpha = 0$  (existing MLE) is clearly nonrobust. Further, it can also be verified that, given any fixed design, the supremum of this IF decreases as  $\alpha$  increases, which in turn implies the increase in their robustness. This fact will be further seconded through empirical illustrations in Section 4.1.

Similar results can also be obtained if there are contaminations in all the  $G_S$  (see Result R2(ii) of the online supplement). The resulting influence function is then the sum of the previous IFs for individual componentwise contaminations and hence the implication is again the same indicating robustness at  $\alpha > 0$  and nonrobustness at  $\alpha = 0$ .

### 3 Robust hypothesis testing: A Wald-type test statistic

Let us now consider the second important aspect of statistical inference, namely the testing of statistical hypothesis. As noted previously, the existing MLE-based likelihood ratio tests or Wald tests are highly nonrobust against data contamination in any set-up including the BRM. Suitable robust hypothesis testing procedures under the general nonhomogeneous set-up have been developed in Ghosh and Basu<sup>22</sup> and Basu et al.<sup>23</sup> by extending the likelihood ratio and the Wald-type tests, respectively. In this section, we develop a robust hypothesis testing procedure based on the proposed MDPDE for the BRM; here we restrict ourselves only to the Wald-type tests which are easy to implement in practice. Related background results from Basu et al.<sup>23</sup> are again provided in the online supplement for the sake of completeness.

Consider the BRM (2) with the set-up as discussed in the previous sections. Consider the most common class of general linear hypotheses given by

$$H_0 : M\beta = m_0 \quad \text{against} \quad H_1 : M\beta \neq m_0 \quad (9)$$

where  $M$  is a known matrix of order  $r \times p$  and  $m_0$  is a known  $r$ -vector of reals. We make the standard assumption that  $\text{rank}(M) = r$  so that there exists a true null parameter value  $\beta_0 \neq \theta_p$  (say) satisfying  $M\beta_0 = m_0$ . Suppose  $\hat{\theta}_{n,\alpha} = (\hat{\beta}_{n,\alpha}^T, \hat{\phi}_{n,\alpha})^T$  denotes the MDPDE of  $\theta = (\beta^T, \phi)^T$  under the BRM (2). We define the Wald-type test statistic for testing hypothesis (9) as

$$W_{n,\alpha} = n(\hat{M}\hat{\beta}_{n,\alpha} - m_0)^T \left[ M\Psi_n^{11}(\hat{\theta}_{n,\alpha})^{-1} \Omega_n^{11}(\hat{\theta}_{n,\alpha}) \Psi_n^{11}(\hat{\theta}_{n,\alpha})^{-1} M^T \right]^{-1} (\hat{M}\hat{\beta}_{n,\alpha} - m_0) \quad (10)$$

where  $\Psi_n^{11}$  and  $\Omega_n^{11}$  are the  $p \times p$  principal sub-matrices of  $\Psi_n$  and  $\Omega_n$  respectively and are given by  $\Psi_n^{11}(\theta) = \frac{1}{n} \sum_{i=1}^n \gamma_{11,i}^{(\alpha)}(\theta) \mathbf{x}_i \mathbf{x}_i^T$  and  $\Omega_n^{11}(\theta) = \frac{1}{n} \sum_{i=1}^n \left\{ \gamma_{11,i}^{(2\alpha)}(\theta) - \gamma_{11,i}^{(\alpha)}(\theta)^2 \right\} \mathbf{x}_i \mathbf{x}_i^T$ . Note that, since the MDPDE at  $\alpha = 0$  coincides with the MLE, the test statistic  $W_{n,0}$  is nothing but the nonrobust MLE-based classical Wald test. So, the proposed test statistics  $W_{n,\alpha}$  are robust generalization of the Wald tests and hence referred to as the Wald-type tests.

In particular, for testing the significance of individual regression coefficient  $\beta_j$ , i.e. testing

$$H_0 : \beta_j = 0 \quad \text{against} \quad H_1 : \beta_j \neq 0 \quad (11)$$

for any  $j = 1, \dots, p$ , the proposed test statistic (10) simplifies to  $W_{n,\alpha} = \frac{n\hat{\beta}_{n,\alpha,j}^2}{\sigma_j(\hat{\beta}_{n,\alpha})}$ , where  $\hat{\beta}_{n,\alpha,j}$  is the MDPDE of  $\beta_j$  and  $\sigma_j(\theta)$  is the asymptotic variance of  $\sqrt{n}\hat{\beta}_{n,\alpha,j}$ .

### 3.1 Asymptotic properties

The first property that we need for any proposed test statistic is its null distribution to find out the critical region of the test. Although the exact null distribution is not easy to obtain in general, the asymptotic distribution of our proposed test statistic  $W_{n,\alpha}$  can be derived directly from that of the MDPDE. We assume that the matrices involved in the asymptotic variance of the MDPDE of  $\beta$ , namely  $\Psi_n^{11}$  and  $\Omega_n^{11}$ , are continuous in  $\theta$ . Then, it is straightforward from the results of Section 2.2 that the asymptotic null distribution of  $W_{n,\alpha}$  for hypothesis (9) is  $\chi_r^2$ , the chi-square distribution with  $r$  degrees of freedom (see Result R3(i) in the online supplement). So, the critical region of the proposed testing procedure at  $\alpha_0$ -level of significance is given by  $\left\{ W_{n,\alpha} > \chi_{r,\alpha_0}^2 \right\}$ , where  $\chi_{r,\alpha_0}^2$  is the  $(1 - \alpha_0)$ -th quantile of the  $\chi_r^2$  distribution. For the particular case of the hypothesis (11), the corresponding null asymptotic distribution of  $W_{n,\alpha}$  is  $\chi_1^2$ . So, we can also perform the one-sided testing for the significance of  $\beta_j$  by considering the test statistic  $W_{n,\alpha}^+ = \frac{\sqrt{n}\hat{\beta}_{n,\alpha,j}}{\sqrt{\sigma_j(\hat{\beta}_{n,\alpha})}}$ , which has an asymptotic standard normal distribution at the null hypothesis in (11).

Further we can apply suitable results from Basu et al.<sup>23</sup> on the Wald-type tests for the general nonhomogeneous set-up to obtain useful power approximations for our proposal in the BRM. In particular, by Result R3(ii) of the online supplement, the tests based on  $W_{n,\alpha}$  are consistent at any fixed alternative for every  $\alpha \geq 0$ ; this fact also follows from the Fisher consistency of the MDPDEs used in the construction of test statistics and we leave the details for the reader.

So, for the purpose of comparison, we need to compute the asymptotic power under the contiguous sequence of alternatives  $H_{1,n} : \beta_n = \beta_0 + \frac{d}{\sqrt{n}}$  for  $d \in \mathbb{R}^p - \{\mathbf{0}_p\}$ , where  $\theta_0 = (\beta_0^T, \phi_0^T)^T$  is the null parameter value satisfying  $M\beta_0 = \mathbf{m}_0$ . However, using the asymptotic distribution of the MDPDE from Section 2.2, one can obtain the asymptotic distribution of our test statistics  $W_{n,\alpha}$  under the hypothesis  $H_{1,n}$  to be  $\chi_r^2(\delta)$ , the noncentral  $\chi^2$  with degrees of freedom  $r$  and noncentrality parameter  $\delta = d^T M [M\Psi_n^{11}(\theta_0)^{-1} \Omega_n^{11}(\theta_0) \Psi_n^{11}(\theta_0)^{-1} M^T]^{-1} M^T d$  (see Result R3(iii) of the online supplement). The asymptotic contiguous power of the proposed testing procedure can then be computed as  $\left[ 1 - F_{\chi_r^2(\delta)}(\chi_{r,\alpha_0}^2) \right]$ , where  $F_{\chi_r^2(\delta)}$  denotes the distribution function of  $\chi_r^2(\delta)$ . In particular, the pitman's asymptotic relative efficiencies of  $W_{n,\alpha}$ -based Wald-type tests at  $\alpha > 0$  with respect to the most powerful (but nonrobust) classical Wald test (at  $\alpha = 0$ ) depend on the noncentrality parameter  $\delta$  and are directly proportional to the ratio of the inverse variance matrix of the MDPDE and the MLE. Hence, they are indeed directly proportional to the asymptotic efficiency of the MDPDE itself. So, for any given fixed design, the asymptotic power under contiguous alternative decreases slightly with increasing  $\alpha$  but the loss is not quite significant at small positive  $\alpha$  as in the case of efficiency of the MDPDEs; see Section 4.2 for corresponding empirical illustrations.

### 3.2 Robustness analysis

We theoretically study the robustness of the proposed Wald-type tests through the corresponding influence function analysis.<sup>19</sup> Considering the set-up of Section 2.3, we define the statistical functional corresponding to the proposed test statistics  $W_{n,\alpha}$  (ignoring the multiplier  $n$ ) as

$$W_\alpha(G_1, \dots, G_n) = (MT_\alpha(G_1, \dots, G_n) - \mathbf{m}_0)^T \Sigma(G_1, \dots, G_n)^{-1} (MT_\alpha(G_1, \dots, G_n) - \mathbf{m}_0) \quad (12)$$



where  $\Sigma(G_1, \dots, G_n) = [\mathbf{M}\Psi_n^{11}(\mathbf{T}_\alpha(G_1, \dots, G_n))^{-1}\mathbf{\Omega}_n^{11}(\mathbf{T}_\alpha(G_1, \dots, G_n))\Psi_n^{11}(\mathbf{T}_\alpha(G_1, \dots, G_n))^{-1}\mathbf{M}^T]$  and  $\mathbf{T}_\alpha(G_1, \dots, G_n)$  is the functional for the MDPDE as defined in (8). We can define its influence function as in the case of estimation by assuming contamination in any fixed distribution or in all distributions.

Let us again consider the contamination only in the distribution  $G_{i_0}$  at the contamination point  $t_{i_0}$ . Then, using the Fisher consistency of  $\mathbf{T}_\alpha$ , a routine differentiation yields the (first order) influence function of the test functional  $W_\alpha$  to be identically zero at the model, i.e.

$$\mathcal{IF}(t_{i_0}, W_\alpha; F_1, \dots, F_n) = 0$$

Therefore, this first-order influence function cannot indicate the robustness of our proposed Wald-type tests, which is expected from the literature of similar quadratic tests.<sup>22–25</sup> So, we need to consider the second-order influence function for  $W_\alpha$  defined analogously with the second-order partial derivative as

$$\mathcal{IF}_2(t_{i_0}, W_\alpha; G_1, \dots, G_n) = \left| \frac{\partial^2 W_\alpha(G_1, \dots, G_{i_0, \epsilon}, \dots, G_n)}{\partial^2 \epsilon} \right|_{\epsilon=0}$$

It indicates a second-order approximation to the asymptotic bias due to infinitesimal contamination in contrast to the first-order approximation provided by the first-order IF. For the present BRM some calculations, based on Result R4(i) of the online supplement, yield the form of this second-order IF for the proposed Wald-type test functional  $W_\alpha$  at the model as given by

$$\mathcal{IF}_2(t_{i_0}, W_\alpha; F_1, \dots, F_n) = \mathcal{IF}(t_{i_0}, \mathbf{T}_\alpha; F_1, \dots, F_n) \mathbf{M}^T \Sigma(F_1, \dots, F_n)^{-1} \mathbf{M} \mathcal{IF}(t_{i_0}, \mathbf{T}_\alpha; F_1, \dots, F_n)$$

Therefore, this second order influence function is bounded if and only if the IF of the MDPDE  $\mathbf{T}_\alpha$ , derived in Section 2.3, is bounded and this holds only for all  $\alpha > 0$ . Hence, the proposed Wald-type test statistics are expected to be robust for  $\alpha > 0$  but nonrobust at  $\alpha = 0$  (which is the classical MLE-based Wald test); further numerical illustrations are given in Section 4.2.

We can also examine the influence of the contamination on the level and power of the proposed Wald-type tests through the level and power influence function analysis.<sup>19,22</sup> For this purpose, we can directly apply the corresponding results for the general nonhomogeneous cases from Basu et al.,<sup>23</sup> described in Result R4(ii) of the online supplement, to conclude that the power influence function is indeed a matrix multiple of the IF of the MDPDE  $\mathbf{T}_\alpha$ . Therefore, the proposed test is robust in asymptotic contiguous power whenever the IF of  $\mathbf{T}_\alpha$  is bounded, i.e. for all  $\alpha > 0$ , but is nonrobust at  $\alpha = 0$ . However, following the same result,<sup>23</sup> the level influence function of this type of tests under nonhomogeneous data is identically zero whenever the IF of  $\mathbf{T}_\alpha$  is bounded, indicating the robustness of asymptotic level for all  $\alpha > 0$  against infinitesimal contiguous contamination at the null hypothesis.

## 4 Simulation studies

### 4.1 Performance of the MDPDE

Let us now study the finite-sample behavior of the proposed estimator, MDPDE, through suitable simulation studies and compare them with theoretical (asymptotic) results. Consider a sample size  $n$  and fix  $n$  covariate values  $x_1, \dots, x_n$  being independent observations from  $U(0, 1)$ . We generate 1000 samples from the BRM (2) with  $p = 2$ , one intercept ( $\beta_1$ ) and one slope ( $\beta_2$ ) corresponding to the covariates  $x_i$ , along with the logit link function. The true value of the parameter  $\boldsymbol{\theta} = (\beta_1, \beta_2, \phi)^T$  is taken as  $(-1, 1, 5)^T$ . For each of the samples, we compute the MDPDEs with different  $\alpha$  and derive their empirical bias and mean square error (MSE) over these 1000 replications (without any outlier); the results are reported in Table 1 for sample sizes  $n = 50, 100$ . Clearly, MLE has the minimum absolute bias and MSE under pure data as expected and the bias and MSE of the proposed MDPDE increase slightly with increasing  $\alpha$ . But this increase in bias or MSE is not quite significant at small positive  $\alpha$  like 0.3, 0.4, which is consistent with the asymptotic efficiency described in Section 2.2.

Next, to study the finite-sample robustness behavior of the proposed MDPDEs, we repeat the previous simulation study, but after contaminating each sample through two different schemes. In the contamination scheme (I), we have randomly changed 10% of the response values  $y$  to  $(1 - y)$  and recalculated the MDPDEs

**Table 1.** Empirical bias and MSE of the MDPDEs with different  $\alpha$  under pure data.

$\alpha$	$n = 50$						$n = 100$					
	Bias			MSE			Bias			MSE		
	$\beta_1$	$\beta_2$	$\phi$	$\beta_1$	$\beta_2$	$\phi$	$\beta_1$	$\beta_2$	$\phi$	$\beta_1$	$\beta_2$	$\phi$
0 (MLE)	-0.010	0.011	0.332	0.058	0.172	1.124	-0.0042	0.004	0.202	0.032	0.097	0.537
0.1	-0.010	0.011	0.325	0.058	0.174	1.123	-0.0033	0.003	0.196	0.032	0.097	0.541
0.2	-0.012	0.013	0.338	0.059	0.178	1.180	-0.003	0.003	0.200	0.033	0.099	0.566
0.3	-0.014	0.015	0.367	0.062	0.185	1.293	-0.0031	0.003	0.211	0.034	0.102	0.607
0.4	-0.017	0.018	0.410	0.064	0.193	1.464	-0.0036	0.004	0.228	0.035	0.106	0.663
0.5	-0.020	0.022	0.464	0.067	0.203	1.696	-0.0042	0.004	0.248	0.037	0.110	0.731
0.6	-0.024	0.026	0.526	0.071	0.214	1.990	-0.005	0.005	0.271	0.038	0.115	0.810
0.7	-0.028	0.031	0.593	0.074	0.225	2.347	-0.0059	0.006	0.296	0.040	0.120	0.895

MLE: maximum likelihood estimator; MSE: mean square error.

**Table 2.** Empirical bias and MSE of MDPDEs with different  $\alpha$  under contamination scheme (I).

$\alpha$	$n = 50$						$n = 100$					
	Bias			MSE			Bias			MSE		
	$\beta_1$	$\beta_2$	$\phi$	$\beta_1$	$\beta_2$	$\phi$	$\beta_1$	$\beta_2$	$\phi$	$\beta_1$	$\beta_2$	$\phi$
0 (MLE)	0.232	-0.240	-0.332	0.141	0.283	1.066	0.2042	-0.192	-0.559	0.082	0.133	0.693
0.1	0.216	-0.218	-0.300	0.132	0.270	1.037	0.1845	-0.166	-0.508	0.074	0.122	0.644
0.2	0.199	-0.197	-0.247	0.125	0.262	1.050	0.1665	-0.144	-0.452	0.067	0.116	0.616
0.3	0.184	-0.177	-0.182	0.121	0.259	1.100	0.1511	-0.125	-0.397	0.063	0.113	0.607
0.4	0.170	-0.160	-0.111	0.119	0.260	1.196	0.1381	-0.109	-0.347	0.060	0.111	0.613
0.5	0.158	-0.145	-0.036	0.118	0.264	1.343	0.1272	-0.096	-0.301	0.059	0.112	0.630
0.6	0.147	-0.133	0.043	0.119	0.269	1.550	0.1181	-0.085	-0.259	0.058	0.113	0.656
0.7	0.138	-0.122	0.120	0.120	0.276	1.790	0.1104	-0.076	-0.220	0.057	0.115	0.689

MLE: maximum likelihood estimator; MSE: mean square error.

based on the contaminated samples. The corresponding bias and MSE are reported in Table 2. Following the suggestion of a referee, in the second contamination scheme (II), we have replaced 5% of the response ( $y$ ) values associated with the minimum  $x$ -values to the extreme point  $y = 0.99$ ; the empirical bias and MSE for this extreme case of contamination are reported in Table 3. It can be observed that the absolute bias and MSE of the MLE are the worst, since it is the most nonrobust one. As  $\alpha$  increases, both the absolute bias and MSE decrease significantly providing more accurate results; these become more prominent in the extreme contamination scheme (II). Thus, the robustness of the proposed MDPDE under contamination significantly improves with increasing values of  $\alpha$ ; this is again consistent with the theoretical influence function analysis discussed in Section 2.3.

Similar results are observed in several simulation studies with different contamination schemes and different sample sizes; so those are not repeated here for brevity.

## 4.2 Performance of the Wald-type tests

In this section, we illustrate the empirical levels and powers of the proposed Wald-type tests based on the MDPDEs through simulation studies. For the sake of consistency, let us consider the same simulation set-up as described in the previous section; with each simulated sample of size  $n = 50$  or  $100$ , both with and without

**Table 3.** Empirical bias and MSE of MDPDEs with different  $\alpha$  under contamination scheme (II).

$\alpha$	$n = 50$						$n = 100$					
	Bias			MSE			Bias			MSE		
	$\beta_1$	$\beta_2$	$\phi$	$\beta_1$	$\beta_2$	$\phi$	$\beta_1$	$\beta_2$	$\phi$	$\beta_1$	$\beta_2$	$\phi$
0 (MLE)	1.120	-1.588	-2.382	1.292	2.635	5.772	0.9779	-1.331	-2.271	0.977	1.829	5.205
0.1	1.067	-1.493	-2.348	1.186	2.363	5.612	0.8868	-1.193	-2.177	0.816	1.498	4.798
0.2	0.869	-1.203	-1.886	0.956	1.873	5.145	0.6326	-0.841	-1.681	0.509	0.925	3.392
0.3	0.450	-0.619	-0.690	0.575	1.136	3.635	0.2536	-0.335	-0.588	0.187	0.352	1.521
0.4	0.222	-0.307	-0.032	0.339	0.693	2.696	0.11	-0.146	-0.087	0.083	0.171	0.897
0.5	0.126	-0.175	0.243	0.216	0.464	2.301	0.0753	-0.101	0.065	0.059	0.131	0.778
0.6	0.092	-0.132	0.361	0.164	0.371	2.216	0.0686	-0.093	0.115	0.053	0.121	0.761
0.7	0.077	-0.114	0.426	0.130	0.309	2.250	0.0716	-0.097	0.129	0.052	0.121	0.782

MLE: maximum likelihood estimator; MSE: mean square error.

**Table 4.** Empirical levels and powers for the MDPDE-based Wald-type tests for different null hypotheses and different  $\alpha$  under pure data.

$\alpha$	$n = 50$						$n = 100$					
	Size			Power			Size			Power		
	$H_0^{L1}$	$H_0^{L2}$	$H_0^{L3}$	$H_0^{P1}$	$H_0^{P2}$	$H_0^{P3}$	$H_0^{L1}$	$H_0^{L2}$	$H_0^{L3}$	$H_0^{P1}$	$H_0^{P2}$	$H_0^{P3}$
0 (Wald)	0.060	0.058	0.055	0.996	0.719	0.995	0.043	0.044	0.061	1.000	0.956	1.000
0.1	0.058	0.057	0.057	0.995	0.712	0.996	0.041	0.044	0.058	1.000	0.957	1.000
0.2	0.058	0.054	0.059	0.994	0.711	0.995	0.043	0.049	0.056	1.000	0.952	1.000
0.3	0.060	0.058	0.066	0.994	0.697	0.995	0.047	0.048	0.057	1.000	0.945	1.000
0.4	0.066	0.063	0.069	0.993	0.691	0.995	0.050	0.052	0.058	1.000	0.934	1.000
0.5	0.071	0.066	0.074	0.992	0.679	0.993	0.053	0.050	0.059	1.000	0.929	1.000
0.6	0.076	0.070	0.082	0.989	0.667	0.993	0.055	0.054	0.057	1.000	0.925	1.000
0.7	0.076	0.069	0.086	0.986	0.657	0.991	0.055	0.054	0.061	1.000	0.958	1.000

contamination as before, we apply the proposed testing procedure for different hypotheses. In particular, we perform the Wald-type tests with different  $\alpha$  for six null hypotheses given by

$$\begin{aligned}
 &H_0^{L1} : \beta_1 = -1; \quad H_0^{L2} : \beta_2 = 1; \quad H_0^{L3} : (\beta_1, \beta_2)^T = (-1, 1)^T; \quad \text{for studying level,} \\
 &\text{and} \quad H_0^{P1} : \beta_1 = 0; \quad H_0^{P2} : \beta_2 = 0; \quad H_0^{P3} : (\beta_1, \beta_2)^T = (0, 0)^T; \quad \text{for studying power}
 \end{aligned}$$

against their respective omnibus alternatives. Note that, all these hypotheses belong to the class of general linear hypotheses (9) considered in Section 3. Based on 1000 replications, we compute the empirical levels and powers at the 5% level of significance for testing these hypotheses under pure data as well as under contaminated data; the results are reported in Tables 4, 5, and 6 respectively.

It can be observed that the levels of the MDPDE-based Wald-type tests increase slightly under pure data for any hypothesis. In fact, most of the empirical levels are slightly inflated due to the use of asymptotic critical values for testing with finite sample sizes. Also, as  $\alpha$  increases, the powers under pure data decrease very little for all the three hypotheses. The changes at small  $\alpha > 0$  under pure data with respect to the classical Wald test at  $\alpha = 0$  are clearly not quite significant. On the other hand, under contaminated data, the levels and powers of the classical Wald-test ( $\alpha = 0$ ) change drastically for all hypotheses. But those for the proposed Wald-type tests at small positive  $\alpha$  remain more stable under both types of contaminations.

**Table 5.** Empirical levels and powers for the MDPDE-based Wald-type tests for different null hypotheses and different  $\alpha$  under (mild) contamination scheme (I).

$\alpha$	$n = 50$						$n = 100$					
	Size			Power			Size			Power		
	$H_0^{L1}$	$H_0^{L2}$	$H_0^{L3}$	$H_0^{P1}$	$H_0^{P2}$	$H_0^{P3}$	$H_0^{L1}$	$H_0^{L2}$	$H_0^{L3}$	$H_0^{P1}$	$H_0^{P2}$	$H_0^{P3}$
0 (Wald)	0.172	0.100	0.163	0.784	0.416	0.831	0.264	0.117	0.237	0.985	0.750	0.993
0.1	0.155	0.098	0.153	0.798	0.428	0.844	0.233	0.103	0.209	0.993	0.784	0.996
0.2	0.144	0.092	0.146	0.806	0.444	0.859	0.207	0.091	0.187	0.993	0.804	0.996
0.3	0.140	0.082	0.137	0.819	0.454	0.863	0.182	0.087	0.175	0.994	0.815	0.996
0.4	0.137	0.078	0.129	0.824	0.462	0.859	0.164	0.078	0.160	0.996	0.816	0.996
0.5	0.133	0.078	0.124	0.828	0.461	0.860	0.156	0.076	0.151	0.996	0.818	0.997
0.6	0.125	0.079	0.120	0.828	0.460	0.856	0.147	0.075	0.148	0.996	0.817	0.997
0.7	0.120	0.077	0.120	0.821	0.453	0.854	0.134	0.073	0.136	0.994	0.809	0.997

**Table 6.** Empirical levels and powers for the MDPDE-based Wald-type tests for different null hypotheses and different  $\alpha$  under (extreme) contamination scheme (II).

$\alpha$	$n = 50$						$n = 100$					
	Size			Power			Size			Power		
	$H_0^{L1}$	$H_0^{L2}$	$H_0^{L3}$	$H_0^{P1}$	$H_0^{P2}$	$H_0^{P3}$	$H_0^{L1}$	$H_0^{L2}$	$H_0^{L3}$	$H_0^{P1}$	$H_0^{P2}$	$H_0^{P3}$
0 (Wald)	0.992	0.945	0.977	0.025	0.199	0.240	0.998	0.996	0.998	0.009	0.133	0.371
0.1	0.983	0.889	0.938	0.022	0.165	0.192	0.996	0.971	0.991	0.023	0.078	0.376
0.2	0.836	0.719	0.767	0.088	0.180	0.235	0.779	0.699	0.739	0.268	0.135	0.570
0.3	0.453	0.401	0.426	0.483	0.371	0.561	0.355	0.279	0.295	0.761	0.552	0.875
0.4	0.238	0.220	0.236	0.754	0.462	0.786	0.157	0.133	0.132	0.931	0.739	0.969
0.5	0.151	0.141	0.150	0.852	0.472	0.901	0.124	0.092	0.091	0.978	0.778	0.992
0.6	0.120	0.103	0.123	0.918	0.627	0.934	0.103	0.074	0.080	0.989	0.785	0.996
0.7	0.100	0.094	0.105	0.916	0.608	0.938	0.095	0.066	0.081	0.993	0.784	0.997

## 5 Applications to real-life data

### 5.1 Application I: AIS Data (the motivating example)

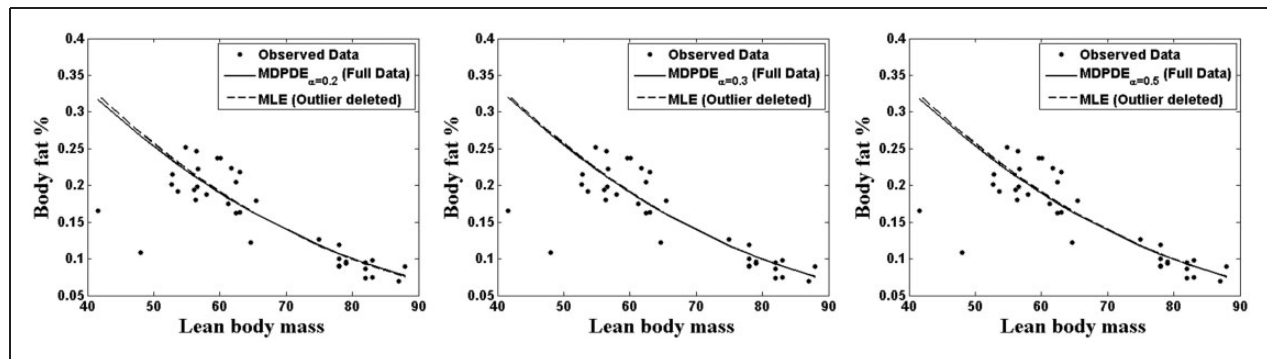
Let us start our illustration with reanalyzing the motivating AIS Dataset described in Section 1. We compute the proposed MDPDEs of the parameter  $\theta = (\beta_1, \beta_2, \phi)^T$  of the fitted BRM for different values of the tuning parameters  $\alpha$  based on the full data and the outlier deleted data. The resulting estimates are reported in Table 7 along with the most commonly used MLE (at  $\alpha = 0$ ). Clearly, unlike the MLE, the proposed MDPDEs with  $\alpha \geq 0.3$  change very little in the presence of two outlying observations. Further, the MDPDEs obtained based on the full data are themselves very close to the outlier deleted MLE (see Figure 2) and so they can be used safely without bothering about the outliers.

Next, let us consider the problem of testing significance of the intercept term, namely  $H_0 : \beta_1 = 0$ . The  $p$ -value of the existing MLE-based Wald test changes drastically due to the presence of two outliers. We apply our proposed Wald-type tests based on the MDPDEs for this testing problem and resulting  $p$ -values are reported in Table 7. Again, the proposed tests with  $\alpha \geq 0.2$  generate stable  $p$ -values (which is zero) even in the full data with outliers. Therefore, the use of the proposed MDPDE and corresponding Wald-type tests with slightly larger  $\alpha > 0$  can successfully tackle the effect of two outliers in the dataset yielding robust estimators and inference even without separately finding and removing these outliers.

**Table 7.** MDPDEs of  $(\beta_1, \beta_2, \phi)^T$  for the AIS data, along with the  $p$ -values for testing  $H_0 : \beta_1 = 0$  using the Wald-type tests.

$\alpha$	Full data				Outlier deleted data			
	$\beta_1$	$\beta_2$	$\phi$	$p$ -value	$\beta_1$	$\beta_2$	$\phi$	$p$ -value
0 (MLE)	0.098	−0.027	96.616	0.699	0.838	−0.038	246.305	0
0.1	0.328	−0.031	116.026	0.158	0.832	−0.038	238.036	0
0.2	0.765	−0.037	206.180	0	0.824	−0.038	231.658	0
0.3	0.807	−0.038	219.286	0	0.815	−0.038	227.072	0
0.4	0.804	−0.038	218.032	0	0.804	−0.038	224.270	0
0.5	0.794	−0.038	216.333	0	0.790	−0.038	223.383	0

MLE: maximum likelihood estimator.

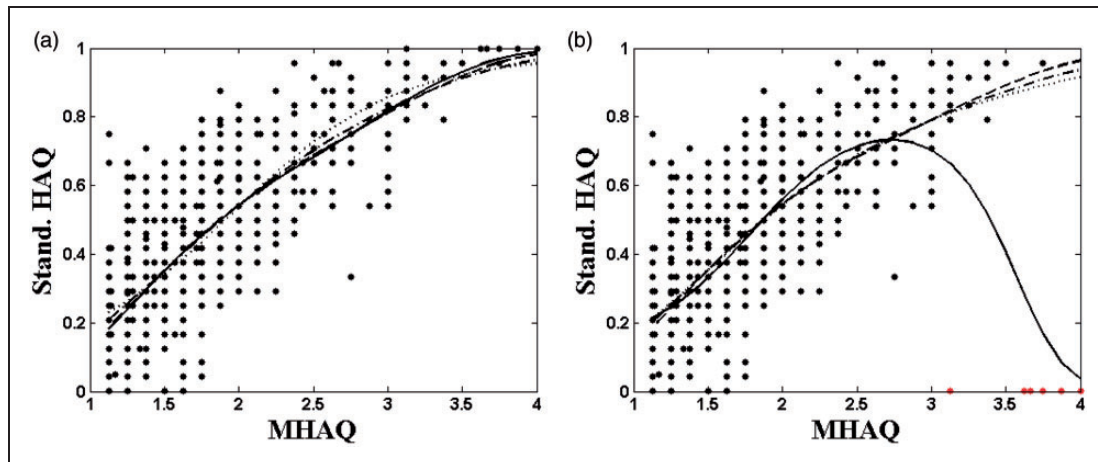
**Figure 2.** The BRM fitted lines for the AIS data based on the MDPDEs at  $\alpha = 0.2, 0.3, 0.5$  for the full data, along with that based on the outlier deleted MLE.

## 5.2 Application 2: HAQ Dataset

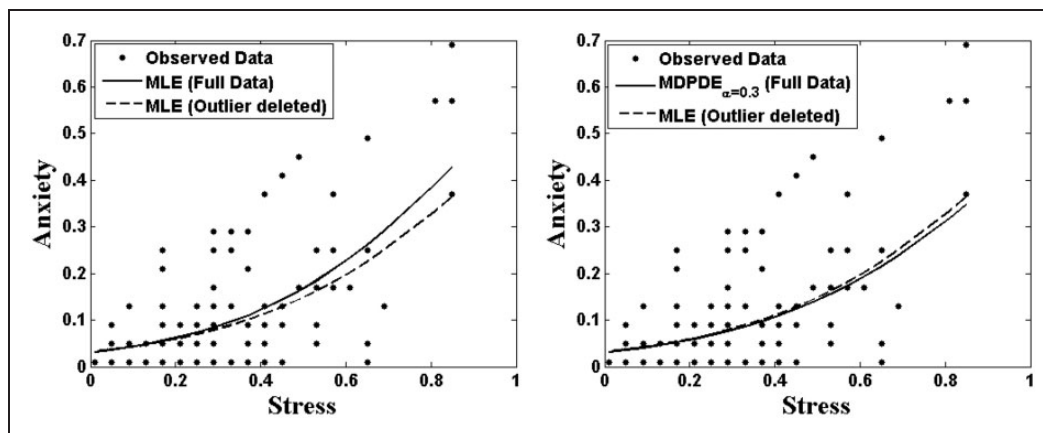
In this example, we consider data on a certain standardized health assessment questionnaire (HAQ) from the Division for Women and Children at the Oslo University Hospital at Ullevål, Oslo, Norway. The data, obtained from Prof. Nils L. Hjort of University of Oslo through personal communication, contain the original (elaborative) HAQ scores along with an easy-to-use modified version (MHAQ) for 1018 patients. These data have been used by Claeskens and Hjort<sup>26</sup> to predict the original HAQ score from the simpler MHAQ scores, after suitable standardization, through a beta regression model. They have argued that the most healthy 219 patients with MHAQ=1 need to be treated separately, but the remaining 799 patients' data can be modelled well by a polynomial BRM with covariates  $x = (1, \text{MHAQ}, \text{MHAQ}^2, \text{MHAQ}^3)^T$  and the logit link function. The corresponding fitted line based on the MLE is plotted in Figure 3(a); clearly there is no outlier in the data. Here, the response variable HAQ takes the values in  $[0, 3]$  inclusive of the end-points and, to get it within the open interval  $(0, 1)$ , we use the popular ad hoc transformation  $y = ((\text{HAQ}/3) \cdot (n - 1) + 0.5)/n$ , where  $n = 799$  is the total sample size.<sup>6,10</sup> Now, let us compute the MDPDEs for this clean dataset to illustrate the behavior of our proposal in pure data. The resulting estimators in fact turn out to be very close to the MLE, which can clearly be seen from the fitted lines in Figure 3(a).

Now, to illustrate the robustness aspect, let us change only six largest HAQ values to  $(1 - \text{HAQ})$  values and again derive the MLE and the MDPDEs; the fitted lines are shown in Figure 3(b) (the artificial outliers are marked as red points). Note that, only due to these 6 outliers, which is about only 0.75% of the total number of observations, the MLE changes to a drastically different fit which clearly gives an erroneous inference. In fact the MLE-based Wald test for testing the significance of the intercept term now gives the  $p$ -value of 0.64 (implying nonsignificance) with these outliers, which was zero (significant) in the original clean data. However, the MDPDE based fits remain very stable for all  $\alpha \geq 0.1$  even in the presence of these outliers as seen from Figure 3(b). Also, the corresponding MDPDE-based Wald-type tests at  $\alpha \geq 0.1$  yield correct  $p$ -value of zero for testing the significance of intercept term both in the clean data and with these outliers.





**Figure 3.** The BRM fitted lines for the HAQ data based on the MDPDEs with different  $\alpha$  (Solid line:  $\alpha = 0$  (MLE); Dashed line:  $\alpha = 0.1$ ; Dash-dotted line:  $\alpha = 0.4$ ; Dotted line:  $\alpha = 0.7$ ): (a) clean data; (b) data with outliers (marked in red).



**Figure 4.** The BRM fitted lines for the Stress-Anxiety data based on the MDPDEs at  $\alpha = 0$  (MLE) and  $\alpha = 0.3$  for the full data, along with that of the outlier deleted MLE.

### 5.3 Application 3: Stress-Anxiety data (psychology)

Our final example is from a psychological trial among 166 nonclinical women in Australia measuring the scores on suitable tests of their anxiety, depression and stress symptoms. The details of the data can be found in Smithson and Verkuilen<sup>6</sup> who have analyzed it with a beta regression model with response as anxiety scores and the covariates being the intercept and the stress scores along with the logit link function. Chien<sup>27</sup> has studied these data to illustrate that there are several groups of highly influential observations affecting the MLE. We consider a set of five such outliers with higher anxiety scores and compute the MLE of the BRM parameters based on the full data and after deleting these outliers. The corresponding fitted lines are shown in Figure 4, which clearly indicate the nonrobust nature of the MLE against the outlying observations.

We have applied our proposed MDPDE for these data and, as before, the MDPDEs with  $\alpha > 0.2$  yield robust estimators. For brevity, we only present the fitted lines corresponding to the MDPDE with  $\alpha = 0.3$  based on the full data in Figure 4; clearly the result is very close to that of the outlier deleted MLE indicating the robustness of our proposal.

### 5.4 On the choice of the tuning parameter $\alpha$

The proposed DPD-based robust estimators and Wald-type tests depend on a tuning parameter  $\alpha$ . We have seen, both theoretically and empirically, that the efficiency of the proposed MDPDE under pure data decreases slightly

as  $\alpha$  increases, but their robustness under contamination increases significantly. Thus, the tuning parameter  $\alpha$  yields a trade-off between efficiency and robustness of the proposed estimator. For hypothesis testing also, the asymptotic contiguous power decreases slightly with increasing  $\alpha$ , but the robustness of its level and power improves significantly under contamination; here  $\alpha$  trades off the contiguous power under pure data with robustness against outliers. Therefore, in either case, this tuning parameter  $\alpha$  needs to be chosen appropriately for a given dataset.

As observed from various simulations and real data analyses, an  $\alpha \approx 0.3, 0.4$  gives sufficiently robust estimator without significant loss in efficiency under pure data and also provides a desired trade-off for the corresponding Wald-type test. So, the empirical suggested value of  $\alpha$  is to be taken around 0.3 to 0.4, which is expected to work well in most of the applications.

However, for a better trade-off based on the amount of contamination in the given dataset, a data-driven choice of this tuning parameter  $\alpha$  could be useful. There are only a few such approaches for the DPD-based inference. We propose to follow the approach presented by Warwick and Jones<sup>28</sup> and Ghosh and Basu<sup>29</sup> for the iid and the nonhomogeneous data respectively. Their approach is mainly based on choosing  $\alpha$  by minimizing an appropriate estimate of the MSE given by

$$E\left[(\hat{\theta}_{n,\alpha} - \theta^*)^T(\hat{\theta}_{n,\alpha} - \theta^*)\right] = (\theta_\alpha - \theta^*)^T(\theta_\alpha - \theta^*) + \frac{1}{n} \text{Trace}[\Psi_n^{-1} \Omega_n \Psi_n^{-1}]$$

where  $\theta^*$  is the target parameter value,  $\theta_\alpha = T_\alpha(G_1, \dots, G_n)$  and  $\hat{\theta}_{n,\alpha}$  is the MDPDE with tuning parameter  $\alpha$ . For the present case of beta regression models, we can estimate this MSE by plugging in the MDPDE  $\hat{\theta}_{n,\alpha}$  for  $\theta_\alpha$  and also in the variance part, but need to use different pilot estimators for  $\theta^*$ . Ghosh and Basu<sup>29</sup> have suggested that the use of the MDPDE with  $\alpha = 0.5$  serves well as the pilot estimator in case of the linear regression model. This suggestion may be followed in the present case of BRM also, but it needs substantial further investigation, which we hope to do in our future research.

## 6 Extension to nonlinear variable dispersion beta regressions

Although till now we have restricted ourselves to the fixed dispersion (or precision) linear BRM (2) for simplicity, our proposed methodology is in no way limited to such restrictions and can easily be extended to various more complex BRMs. Thus, it is indeed possible to fully exploit the flexibility of the beta regression models through such extensions. To illustrate this claim, in this section, we present the extension of the proposed MDPDE for a general class of nonlinear and variable dispersion BRMs from Simas et al.<sup>7</sup> For this class of BRMs, we allow the precision parameter  $\phi$  (and hence also the dispersion) to be variable for different  $y_i$  so that now we assume  $y_i \sim f(y_i; \mu_i, \phi_i)$  and model  $\phi_i$  by possibly another set of covariates, say  $z_i \in \mathbb{R}^q$ , through suitable link function  $h$  (may be different from the link function  $g$  used in the mean model). We can also avoid the linearity constraint on the predictors to have a larger flexible class of BRMs given by

$$y_i \sim f(y_i; \mu_i, \phi_i) \text{ independently, with } g(\mu_i) = \eta_1(x_i, \beta), h(\phi_i) = \eta_2(z_i, \gamma), \quad i = 1, \dots, n \quad (13)$$

where  $\eta_1$  and  $\eta_2$  are some known functions and  $\beta \in \mathbb{R}^p$  and  $\gamma \in \mathbb{R}^q$  are the vectors of unknown regression coefficients corresponding to the mean and the precision models respectively. Note that, the covariates can have different dimensions compared to the corresponding regression coefficients (although we have kept them the same without any loss of generality), but we need to assume that the derivative matrices of  $\eta_1$  and  $\eta_2$  with respect to  $\beta$  and  $\gamma$ , respectively, have ranks  $p$  and  $q$ . Then, our parameter of interest becomes  $\theta = (\beta^T, \gamma^T)^T \in \mathbb{R}^{p+q}$ .

Note that, under the general class of BRMs (13) also, assuming the covariates to be fixed, the observed responses  $y_i$ s are independent but nonhomogeneous with model density of  $y_i$  being  $f_i(\cdot, \theta) \equiv f(\cdot; \mu_i, \phi_i)$  for  $i = 1, \dots, n$ . So this again belongs to the general set-up of Ghosh and Basu<sup>16</sup> and, as before, we can define the MDPDE of  $\theta$  by minimizing the average DPD measure between the  $i$ th data point and the corresponding model density  $f_i(\cdot, \theta)$  for  $i = 1, \dots, n$ . Following the general theory of Ghosh and Basu,<sup>16</sup> as presented in Section 1 of the online supplement, the MDPDE objective function under the BRMs (13) can again be simplified to have the form

$$H_{n,\alpha}(\theta) = n^{-1} \sum_{i=1}^n \left[ \widetilde{K}_{i,\alpha}(\theta) - \frac{1+\alpha}{\alpha} f_i(y_i, \theta)^\alpha \right] \quad (14)$$

**Table 8.** Relative difference due to outliers, in the MLE and the MDPDE at  $\alpha = 0.3$ , for the Stress-Anxiety data under the variable-dispersion BRM.

	$\beta_1$	$\beta_2$	$\gamma_1$	$\gamma_2$
MLE	1.23%	1.89%	4.47%	13.65%
MDPDE $_{\alpha=0.3}$	0.92%	0.29%	3.41%	6.74%

MLE: maximum likelihood estimator; MDPDE: minimum density power divergence estimator.

where now we have  $f_i(\cdot, \theta) \equiv f(\cdot; \mu_i, \phi_i)$  and  $\widetilde{K}_{i,\alpha}(\theta) = \frac{B((1+\alpha)\mu_i\phi_i - \alpha, (1+\alpha)(1-\mu_i)\phi_i - \alpha)}{B(\mu_i\phi_i, (1-\mu_i)\phi_i)^\alpha}$ . The corresponding estimating equations for the BRMs (13), obtained from the general equation (2) of the online supplement, are again given by

$$\sum_{i=1}^n \left[ \widetilde{\gamma}_{1,i}^{(\alpha)}(\theta) - (y_{1,i}^* - \widetilde{\mu}_{1,i}^*) \frac{\phi_i}{g'(\mu_i)} f_i(y_i, \theta)^\alpha \right] \frac{\partial \eta_1(x_i, \beta)}{\partial \beta} = \mathbf{0}_p \quad (15)$$

$$\sum_{i=1}^n \left[ \widetilde{\gamma}_{2,i}^{(\alpha)}(\theta) - \left\{ \mu_i(y_{1,i}^* - \widetilde{\mu}_{1,i}^*) + (y_{2,i}^* - \widetilde{\mu}_{2,i}^*) \right\} \frac{1}{h'(\phi_i)} f_i(y_i, \theta)^\alpha \right] \frac{\partial \eta_2(z_i, \gamma)}{\partial \gamma} = \mathbf{0}_q \quad (16)$$

where we now have  $\widetilde{\mu}_{1,i}^* = E(y_{1,i}^*) = \psi(\mu_i\phi_i) - \psi((1-\mu_i)\phi_i)$ ,  $\widetilde{\mu}_{2,i}^* = E(y_{2,i}^*) = \psi((1-\mu_i)\phi_i) - \psi(\phi_i)$ ,  $\widetilde{\gamma}_{1,i}^{(\alpha)}(\theta) = \left( \psi(\widetilde{a}_{i,\alpha}) - \psi(\widetilde{b}_{i,\alpha}) - \widetilde{\mu}_{1,i}^* \right) \frac{\phi_i \widetilde{K}_{i,\alpha}(\theta)}{g'(\mu_i)}$  and

$$\widetilde{\gamma}_{2,i}^{(\alpha)}(\theta) = \left[ \mu_i \left( \psi(\widetilde{a}_{i,\alpha}) - \psi(\widetilde{b}_{i,\alpha}) - \widetilde{\mu}_{1,i}^* \right) + \left( \psi(\widetilde{b}_{i,\alpha}) - \psi(\widetilde{a}_{i,\alpha} + \widetilde{b}_{i,\alpha}) - \widetilde{\mu}_{2,i}^* \right) \right] \frac{\widetilde{K}_{i,\alpha}(\theta)}{h'(\phi_i)}$$

with  $\widetilde{a}_{i,\alpha} = (1+\alpha)\mu_i\phi_i - \alpha$ ,  $\widetilde{b}_{i,\alpha} = (1+\alpha)(1-\mu_i)\phi_i - \alpha$ . Proceeding similarly, we can derive all asymptotic and robustness properties of the MDPDEs under the general class of flexible BRMs (13), as before, using the general results from the online supplement. Suitable robust Wald-type tests of any hypothesis under BRMs (13) can also be developed with similar properties based on the general results from Section 2 of the online supplement. Considering the length of the current manuscript, we have decided to keep their details for a future report; but the general interpretations and developments are expected to be exactly similar (as observed in the following example).

### Example: Stress-Anxiety data with variable-dispersion beta regression model

As an illustration of the performance of the MDPDEs under the variable dispersion BMRs, let us reconsider the Stress-Anxiety data studied in Section 5.3. Smithson and Verkuilen<sup>6</sup> have shown that the anxiety scores in this data set can be modeled better with a (linear) variable dispersion beta-regression model than the fixed dispersion BRM (as done in Section 5.3); this is because the variability in anxiety scores clearly depends on the level of stress-scores (see Figure 4). So, we now fit the general model (13) for this dataset with  $y_i = \text{Anxiety Score}_i$ ,  $\mathbf{x}_i = \mathbf{z}_i = (1, \text{Stress Score}_i)^T$  and  $g$  and  $h$  being the “logit” and “log” link functions respectively. Then, we compute the MDPDEs at different  $\alpha > 0$  and the MLEs (at  $\alpha = 0$ ) of the regression parameters  $\theta = (\beta_1, \beta_2, \gamma_1, \gamma_2)^T$  under the full data and the outlier deleted data. Since the changes in the estimators are small, in order to illustrate the extent of robustness, we here study the relative change in the estimators under full data with that under outlier deleted data, which is presented in Table 8 for  $\alpha = 0, 0.3$ . Clearly the change due to outliers is significantly reduced for all parameters, specially for both the slope parameters, while using the newly proposed MDPDEs with  $\alpha \approx 0.3$ . All estimates are statistically significant indicating suitability of the fitted model.

## 7 Concluding remarks

In this paper, we have developed a robust statistical inference procedure under the beta regression model for modeling responses on  $(0, 1)$ . We have proposed the minimum DPD estimator for estimating the parameters in the fixed dispersion BRM and developed a class of Wald-type tests based on them for testing general linear hypotheses in regression coefficients. Besides discussing their asymptotic properties, we have also justified the robustness of the

proposed methodology through appropriate influence function analyses. Suitable numerical illustrations have been provided along with three important real data applications from health care studies. Some indications are also provided, with application, for extending the proposed inference to the variable dispersion beta regression models having nonhomogeneous precisions.

It is worthwhile to note that an important measure of global robustness of an inference procedure is their breakdown point, which is not explored in this paper. Ghosh and Basu<sup>16</sup> have shown that the proposed DPD-based inference with  $\alpha > 0$  has the maximum possible breakdown point of 0.5 under mild boundedness conditions on the covariates in a fixed-design linear regression model. We hope that similar breakdown result can also be derived for the BRM under certain conditions (a mathematical challenge), but we do not have an explicit proof at this moment.

However, our proposed methodology can be directly applied to any complex big dataset to generate robust inference without bothering much about outliers in the data. This is because the proposed estimator has a simple unbiased estimating equation, which can be easily solved efficiently for such big datasets using appropriate numerical techniques and the underline objective function also helps us to avoid any problem in cases with multiple roots to this estimating equation. But, for high-dimensional datasets with more covariates than observations, we need to add a suitable regularization penalty factor (like LASSO or SCAD penalties) in the proposed objective function (5). Such penalized DPD-based approach for robust inference under high-dimensional linear regression model has recently been studied by Zang et al.<sup>30</sup> Similar extension under the present BRM with high-dimensional structure will be an interesting future work.

Besides detailed study of the extension discussed in Section 6, it will also be very useful to further extend it to develop robust inference for the inflated zero or one (or both) BRMs for datasets containing 0 or 1 or both values and the BRMs with repeated measurements; the general theory presented in the online supplement will directly guide in these extensions. Also, the proposed scheme for selection of a data-driven choice of the tuning parameter  $\alpha$  needs more investigation. We plan to pursue some of these extensions in our future works.

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## Supplemental material

Supplemental material is available online for this article.

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