

Bayesian estimation of the parameters for two-parameter bathtub-shaped lifetime distribution based on ranked set sampling

Mehdi Basikhasteh, Fazlollah Lak* and Mahmoud Afshari

Received November 2020

Revised December 2020

Abstract

Bayesian inference of the parameters for bathtub-shaped lifetime distribution based on simple random sampling (SRS) and ranked set sampling (RSS) are obtained. The Monte Carlo Markov chain is used, especially Metropolis-Hastings and Gibbs sampling. To compare different estimates, the Monte Carlo simulations are used. The results of simulation show that the estimators based on RSS are more efficient than based on SRS. Also, the length of highest posterior density (HPD) credible interval based on RSS is shorter than its SRS counterparts. Finally, a real data set has been analyzed for illustrative purposes.

1 Introduction

Several probability distributions are considered and studied for analyzing the bathtub-shaped failure rate data in literature. For example power exponential (Smith and Bain, 1975; Leemis, 1986), a reliability distribution with increasing, decreasing, and bathtub-shaped failure rate (Hjorth, 1980), Weibull exponentiated distribution (Mudholkar and Srivastava, 1993) and the two-parameter bathtub-shaped lifetime (TPBL) distribution (Chen, 2000; Wu et al., 2004, 2008; Ammar et al., 2012). The cumulative distribution function (cdf), the probability density function (pdf) and the failure rate function of the TPBL distribution with shaped and scale parameters (β, λ) are:

$$(1.1) \quad F(x) = 1 - e^{\lambda(1-e^{x^\beta})}, \quad x > 0, \quad \lambda > 0, \quad \beta > 0,$$

$$(1.2) \quad f(x) = \lambda\beta x^{\beta-1} e^{[x^\beta + \lambda(1-e^{x^\beta})]}, \quad x > 0, \quad \lambda > 0, \quad \beta > 0,$$

and

$$(1.3) \quad h(x) = \lambda\beta x^{\beta-1} e^{x^\beta}, \quad x > 0, \quad \lambda > 0, \quad \beta > 0.$$

The failure rate function $h(x)$ has a bathtub-shaped when $\beta < 1$ and increasing-shaped when $\beta \geq 1$, (chen (2000)).

Ranked set sampling (RSS) proposed by McIntyre (1952) and discussed in detail by Takahasi et al. (1968). This method of sampling have studied in different areas that actual measurements of the variable of study are time consuming or expensive, for example, in environmental (Patil et al., 1993), in reliability (Kvam et al., 1994) and in quality control (Muttalak et al., 2003). Comparing RSS with SRS in Bayesian context considered by Lavine (1999), Sadek and ALharbi(2014) and Sadek et al. (2015).

In order to obtain a sample of size n by RSS method, we first take n simple random samples of size n from the population, and each sample is called a set. The items of each set are ranked by means of an auxiliary variable or visual judgment (without actual measurement). Then the r^{th} observation of r^{th} set is measured for $r = 1, \dots, n$.

*Corresponding author: Department of Statistics, Faculty of Sciences, Persian Gulf University, Bushehr 75169, Iran. lak@pgu.ac.ir

Keywords: Bayesian estimation, Monte Carlo Markov chain, ranked set sampling, two-parameter bathtub-shaped lifetime distribution

MSC(2010): Primary: 62D05; Secondary: 62F07, 62F15



This entire process denoted by $X_{(1)1}, \dots, X_{(n)n}$ and is called a *cycle*. It may be repeated k times to drive a RSS of size $m = nk$ from the underlying population where n is the set size and k is the cycle size. Note that $X_{(r)r}$'s are independent and non-identically distributed (INID). The pdf and cdf of $X_{(r)r}$, denoted by $f_{r:n}(x)$ and $F_{r:n}(x)$, and are given by

$$f_{r:n}(x) = \frac{n!}{(r-1)!(n-r)!} [F(x)]^{r-1} [1-F(x)]^{n-r} f(x), \quad x > 0,$$

and

$$F_{r:n}(x) = \sum_{i=r}^n \binom{n}{i} [F(x)]^i [1-F(x)]^{n-i}, \quad x > 0,$$

respectively.

The organization of this paper is as follows: In Section 2, Bayesian estimation and HPD credible interval for the scale parameter λ of TPBL distribution are considered based on SRS and RSS when β is known using gamma and non-informative prior distributions. Also, Bayesian inference of the parameters (β, λ) are provided in Section 3. Section 4 is devoted to simulation results. In Section 5 a real data set is analyzed for illustrative purposes. Finally, in Section 6 we conclude the paper by numerical examples.

2 Bayesian estimation of λ when β is known

Bayesian estimation of λ is considered when β is known based on SRS and RSS using gamma and non-informative prior distributions.

2.1 Bayes estimate of λ based on SRS

Let $\mathbf{x} = x_1, x_2, \dots, x_n$ be a SRS from TPBL(β, λ) distribution, when β is known. It is assumed that the parameter λ follows a conjugate prior, with density:

$$(2.1) \quad \pi(\lambda) \propto \lambda^{a-1} e^{-b\lambda}, \quad a \geq 0, \quad b \geq 0,$$

where a and b are known. In this case, the posterior density $\pi(\lambda|\mathbf{x})$ is given by

$$(2.2) \quad \pi(\lambda|\mathbf{x}) \propto \lambda^{n+a-1} e^{-\lambda \left\{ b - \sum_{i=1}^n (1 - e^{x_i^\beta}) \right\}}.$$

Therefore, $\lambda|\mathbf{x} \sim \text{Gamma}(n+a, b - \sum_{i=1}^n (1 - e^{x_i^\beta}))$. The posterior mean of λ based on SRS is

$$(2.3) \quad \hat{\lambda}_{SRS} = E(\lambda|\mathbf{x}) = \frac{n+a}{b - \sum_{i=1}^n (1 - e^{x_i^\beta})},$$

If $a = b = 0$, the prior (2.1) becomes the non-informative prior, which is given by $\pi(\lambda) \propto 1/\lambda, \lambda > 0$. Then, the Bayesian estimation of λ under non-informative prior is

$$(2.4) \quad \hat{\lambda}_{SRS}^J = -\frac{n}{\sum_{i=1}^n (1 - e^{x_i^\beta})},$$

Furthermore, a $100(1 - \alpha)\%$ highest posterior density (HPD) Bayesian credible interval of λ can be constructed by using a numerical method as a subset $C \in \Lambda$ defined by $C = \{\lambda : \pi(\lambda|\mathbf{x}) \geq k\}$ where k is the largest number such that

$$(2.5) \quad \int_{\lambda: \pi(\lambda|\mathbf{x}) \geq k} \pi(\lambda|\mathbf{x}) d\lambda = 1 - \alpha.$$

2.2 Bayes estimate of λ based on RSS

Let Y_1, Y_2, \dots, Y_n be a RSS from TPBL(β, λ) distribution, the density of Y_i is (see Arnold et al. (1992))

$$g(y_i|\lambda, \beta) = \frac{n!}{(i-1)!(n-i)!} [F(y_i|\lambda, \beta)]^{i-1} [1 - F(y_i|\lambda, \beta)]^{n-i} f(y_i|\lambda, \beta)$$

$$= \sum_{k=0}^{i-1} i \binom{n}{i} \binom{i-1}{k} (-1)^k \lambda \beta y_j^{\beta-1} e^{y_j^\beta + \lambda(1-e^{y_j^\beta})(n+k-j+1)},$$

where f and F are pdf and cdf of the TPBL distribution. The joint density of $\mathbf{y} = (y_1, \dots, y_n)$ is obtained as

$$g(\mathbf{y}|\lambda, \beta) = \prod_{j=1}^n g(y_j|\lambda, \beta) = \sum_{i_1=0}^0 \sum_{i_2=0}^1 \dots \sum_{i_n=0}^{n-1} \left[\prod_{j=1}^n j \binom{n}{j} \binom{j-1}{i_j} (-1)^{i_j} \right]$$

$$\times (\lambda\beta)^n \left(\prod_{j=1}^n y_j \right)^{\beta-1} e^{\sum_{j=1}^n y_j^\beta + \lambda \sum_{j=1}^n (n+i_j-j+1)(1-e^{y_j^\beta})}.$$

If λ follows prior distribution (2.1), then the posterior density of λ is given by

$$\pi(\lambda|\mathbf{y}, \beta) = \frac{\pi(\lambda)g(\mathbf{y}|\lambda, \beta)}{\int_0^\infty \pi(\lambda)g(\mathbf{y}|\lambda, \beta)d\lambda}$$

$$= \frac{\sum_{i_1=0}^0 \sum_{i_2=0}^1 \dots \sum_{i_n=0}^{n-1} \left[\prod_{j=0}^n j \binom{n}{j} \binom{j-1}{i_j} (-1)^{i_j} \right] \lambda^{n+a-1} e^{-\lambda[b - \sum_{j=1}^n (n+i_j-j+1)(1-e^{y_j^\beta})]}}{\sum_{i_1=0}^0 \sum_{i_2=0}^1 \dots \sum_{i_n=0}^{n-1} \left[\prod_{j=0}^n j \binom{n}{j} \binom{j-1}{i_j} (-1)^{i_j} \right] \Gamma(n+a) [b - \sum_{j=1}^n (n+i_j-j+1)(1-e^{y_j^\beta})]^{-(n+a)}}.$$

Therefore, the Bayesian estimation of λ based on RSS is

$$\hat{\lambda}_{RSS} = E(\lambda|\mathbf{y}, \beta)$$

$$= \frac{\sum_{i_1=0}^0 \sum_{i_2=0}^1 \dots \sum_{i_n=0}^{n-1} \left[\prod_{j=0}^n j \binom{n}{j} \binom{j-1}{i_j} (-1)^{i_j} \right] (n+a) [b - \sum_{j=1}^n (n+i_j-j+1)(1-e^{y_j^\beta})]^{-(n+a+1)}}{\sum_{i_1=0}^0 \sum_{i_2=0}^1 \dots \sum_{i_n=0}^{n-1} \left[\prod_{j=0}^n j \binom{n}{j} \binom{j-1}{i_j} (-1)^{i_j} \right] [b - \sum_{j=1}^n (n+i_j-j+1)(1-e^{y_j^\beta})]^{-(n+a)}},$$

The Bayes estimates of λ under non-informative prior are obtained as follows:

$$\hat{\lambda}_{RSS}^J = \frac{\sum_{i_1=0}^0 \sum_{i_2=0}^1 \dots \sum_{i_n=0}^{n-1} \left[\prod_{j=0}^n j \binom{n}{j} \binom{j-1}{i_j} (-1)^{i_j} \right] [-\sum_{j=1}^n (n+i_j-j+1)(1-e^{y_j^\beta})]^{-(n+1)}}{\sum_{i_1=0}^0 \sum_{i_2=0}^1 \dots \sum_{i_n=0}^{n-1} \left[\prod_{j=0}^n j \binom{n}{j} \binom{j-1}{i_j} (-1)^{i_j} \right] [-\sum_{j=1}^n (n+i_j-j+1)(1-e^{y_j^\beta})]^{-n}}.$$

The form of the posterior pdf of λ is so complex to solve the equation (2.5). We use MCMC technique to generate samples from it. Then, a $100(1-\alpha)\%$ HPD Bayesian credible interval of λ can be constructed by applying the method proposed by Chen and Shao (1999) as

$$(\lambda_{[(\alpha/2)N]}, \lambda_{[(1-\alpha/2)N]}),$$

where $\lambda_{[(\alpha/2)N]}$ and $\lambda_{[(1-\alpha/2)N]}$ are the $[(\alpha/2)N]$ th smallest integer and the $[(1-\alpha/2)N]$ th smallest integer of $\{\lambda_i, i = 1, \dots, N\}$.

3 When β and λ are unknown

Bayesian estimation of both parameters of TPBL(β, λ) distribution are discussed based on SRS and RSS using gamma and non-informative priors for each parameter.

3.1 Bayes estimates of β, λ based on SRS

Assuming that β and λ have the joint prior density

$$(3.1) \quad \pi(\beta, \lambda) = \pi(\beta).\pi(\lambda) \propto \beta^{a_1-1}e^{-b_1\beta}\lambda^{a_2-1}e^{-b_2\lambda},$$

using equations (1.2) and (3.1), the joint posterior density $\pi(\beta, \lambda|\mathbf{x})$ is given by

$$(3.2) \quad \pi(\beta, \lambda|\mathbf{x}) \propto \beta^{n+a_1-1}\lambda^{n+a_2-1}e^{-\beta(b_1-\sum_{i=1}^n \log x_i)+\sum_{i=1}^n x_i^\beta-\lambda(b_2-n+\sum_{i=1}^n e^{x_i^\beta})}.$$

Clearly, (3.2) may not reduced to a well known distribution. Therefore, we use Gibbs sampling method to generate values of β and λ from it. The full conditional density of λ is obtained as

$$(3.3) \quad (\lambda|\mathbf{x}, \beta) \sim \text{Gamma} \left(n + a_2, b_2 - n + \sum_{i=1}^n e^{x_i^\beta} \right),$$

and the conditional posterior density of β , satisfies

$$(3.4) \quad \pi(\beta|\mathbf{x}, \lambda) \propto \beta^{n+a_1-1}e^{-\beta(b_1-\sum_{i=1}^n \log x_i)+\sum_{i=1}^n x_i^\beta-\lambda\sum_{i=1}^n e^{x_i^\beta}}.$$

One can easily generates samples of λ from (3.3). But since (3.4) is not a well known pdf, we use the Metropolis-Hastings (M-H) algorithm (Metropolis et al. (1953) and Hastings (2003)) to generate samples of β with the half normal proposal density $q(\beta) \propto N(\beta^{(i-1)}, \tau^2)I(\beta > 0)$, where the accept probability is

$$\rho = \min \left\{ \frac{\pi(\beta^{(i)}, \lambda^{(i)}|\mathbf{x})}{\pi(\beta^{(i-1)}, \lambda^{(i)}|\mathbf{x})} \frac{q(\beta^{(i-1)}|\beta^{(i)}, \mathbf{x})}{q(\beta^{(i)}|\beta^{(i-1)}, \mathbf{x})}, 1 \right\}.$$

The value of τ should be fixed so that the optimized value of ρ is about 0.4. Therefore, the approximates of Bayesian estimation of β and λ based on SEL and LINEX are given by

$$(3.5) \quad \hat{\beta}_{SEL} = \hat{E}(\beta|\mathbf{x}) = \frac{1}{N-M} \sum_{i=M+1}^N \beta^{(i)},$$

$$(3.6) \quad \hat{\beta}_{LINEX} = -\frac{1}{c} \ln[\hat{E}(e^{-c\beta}|\mathbf{x})] = -\frac{1}{c} \ln \left[\frac{1}{N-M} \sum_{i=M+1}^N e^{-c\beta^{(i)}} \right],$$

$$(3.7) \quad \hat{\lambda}_{SEL} = \hat{E}(\lambda|\mathbf{x}) = \frac{1}{N-M} \sum_{i=M+1}^N \lambda^{(i)},$$

$$(3.8) \quad \hat{\lambda}_{LINEX} = -\frac{1}{c} \ln[\hat{E}(e^{-c\lambda}|\mathbf{x})] = -\frac{1}{c} \ln \left[\frac{1}{N-M} \sum_{i=M+1}^N e^{-c\lambda^{(i)}} \right],$$

where M first iterations are taken as burn-in period. Note that if $a_1 = a_2 = b_1 = b_2 = 0$, the prior (3.1) becomes non-informative.

3.2 Bayes estimates of β , λ based on RSS

Using equations (2.6) and (3.1) the joint posterior distribution of β and λ based on RSS is obtained as

$$\begin{aligned} \pi(\beta, \lambda | \mathbf{y}) &\propto \sum_{i_1=0}^0 \sum_{i_2=0}^1 \cdots \sum_{i_n=0}^{n-1} \prod_{j=0}^n j \binom{n}{j} \binom{j-1}{i_j} (-1)^{i_j} \beta^{n+a_1-1} \\ &\times \lambda^{n+a_2-1} \left(\prod_{j=1}^n y_j \right)^{\beta-1} e^{-b_1\beta + \sum_{j=1}^n y_j^\beta - \lambda [b_2 - \sum_{j=1}^n (n+i_j-j+1)(1-e^{y_j^\beta})]} \end{aligned} \tag{3.9}$$

Similarly (3.2), since (3.9) does not have a well known distribution form, we use MCMC technique to generate samples from it. Note that

$$\begin{aligned} \pi(\beta | \mathbf{y}, \lambda) &\propto \sum_{i_1=0}^0 \sum_{i_2=0}^1 \cdots \sum_{i_n=0}^{n-1} \left[\prod_{j=0}^n c_{i_j}(j) \right] \beta^{n+a_1-1} \left(\prod_{j=1}^n y_j \right)^{\beta-1} \\ &\times \exp \left[-b_1\beta + \sum_{j=1}^n y_j^\beta + \lambda \sum_{j=1}^n (n+i_j-j+1)(1-e^{y_j^\beta}) \right] \end{aligned} \tag{3.10}$$

and

$$\begin{aligned} \pi(\lambda | \mathbf{y}, \beta) &\propto \sum_{i_1=0}^0 \sum_{i_2=0}^1 \cdots \sum_{i_n=0}^{n-1} \left[\prod_{j=0}^n c_{i_j}(j) \right] \lambda^{n+a_2-1} \\ &\times \exp \left[-\lambda \left(b_2 - \sum_{j=1}^n (n+i_j-j+1)(1-e^{y_j^\beta}) \right) \right]. \end{aligned} \tag{3.11}$$

where $c_{i_j}(j) = j \binom{n}{j} \binom{j-1}{i_j} (-1)^{i_j}$.

Since, no full conditional densities are known analytically, we rely again on M-H algorithm with two half normal proposals $q(\beta) \propto N(\beta^{(i-1)}, \tau^2)I(\beta > 0)$ and $q(\lambda) \propto N(\lambda^{(i-1)}, \tau^2)I(\lambda > 0)$.

By using equations (3.5) to (3.8) the approximations of Bayesian estimation of β and λ are derived.

4 Simulation Study

Monte Carlo simulations are used to approximate the Bayes estimators based on RSS and SRS for TPBL distribution.

4.1 Simulation for λ when β is known

Table 1 presents the mean squared error(MSE) and the bias of the estimators, for $\lambda = 0.5$ at different sizes of the sample [$n = 4$, $n = 5$ and $n = 6$], when $\beta = 0.5$. We set $a = 2$ and $b = 3$ for Gamma prior. Note that the Bayesian estimates based on RSS have mean squared errors smaller than based on SRS. Also, in terms of bias it is observed that the bias via RSS is smaller than its SRS counterparts.

Furthermore, the mean squared errors of all estimates decrease when n increases. The 95% HPD credible intervals of the Bayesian estimates for λ are presented in Table 2. It is clear that, the HPD length based on RSS is shorter than its SRS counterparts.

Table 1: MSEs and biases of the Bayesian estimates based on SRS and RSS for λ (when $\beta = 0.5$).

n	MSE				bias			
	non-informative		Gamma prior		non-informative		Gamma prior	
	SRS	RSS	SRS	RSS	SRS	RSS	SRS	RSS
4	0.1784	0.0413	0.0596	0.0325	0.1577	0.0518	0.1233	0.0809
5	0.1502	0.0314	0.0565	0.0247	0.1566	0.0524	0.0651	0.0496
6	0.0894	0.0206	0.0473	0.0204	0.0756	0.0247	0.0925	0.0489

Table 2: The 95% HPD credible intervals of the Bayesian estimates based on SRS and RSS for λ (when $\beta = 0.5$).

n	HPD			
	non-informative prior		Gamma prior	
	SRS	RSS	SRS	RSS
4	(0.1233, 1.2810)	(0.2521, 0.9624)	(0.1835, 1.1260)	(0.2904, 0.9771)
5	(0.1614, 1.2261)	(0.2970, 0.8832)	(0.1904, 0.9892)	(0.3113, 0.8574)
6	(0.1697, 1.0388)	(0.3147, 0.7917)	(0.2316, 1.0601)	(0.3397, 0.8091)

4.2 Simulation for β and λ

The simulation study are based on 1000 Monte Carlo runs, and 5000 samples generated by using the M-H algorithm and discard the first $M = 1000$ values as burn-in period and $\tau = 0.5$.

The Bias and mean squared error of estimators, under SEL and LINEX loss function, for values $\beta = 0.5$ and $\lambda = 0.5$ at the sample sizes of $n = 5, 6$ and 7 are presented in Tables 3, when $a_1 = a_2 = 3$ and, $b_1 = b_2 = 2$ for Gamma prior. In Table 4, the bias and MSE of estimators are shown for non-informative prior.

From Tables 3 and 4, it is observed that the Bayesian estimates for β and λ based on RSS are in most cases less biased based on SRS. Also, the MSEs of estimators via RSS are obviously smaller than its SRS counterparts.

In context of loss function, from Tables 3 and 4, it can seen that MSE of the estimators under $LINEX_{\{C=1\}}$ are slightly smaller than the others. Furthermore, the MSE of all estimates decreases when n increases.

The 95% HPD credible intervals of the Bayesian estimates based on SRS and RSS for β and λ under Gamma and non-informative priors are presented in Tables 5 and 6 respectively. It is observed that, the confidence length of HPDs based on RSS are shorter than its SRS counterparts.

We observe no much variation between estimators based on Gamma prior and non-informative prior especially based on RSS from the Tables 3-6.

The relative efficiency of the estimates based on RSS with respect to estimates based on SRS are calculated as follows:

$$efficiency = \frac{MSE(\hat{\theta}_{SRS})}{MSE(\hat{\theta}_{RSS})}$$

and their results are shown in Table 7. Clearly, estimators of β and λ based on RSS are more efficient than based on SRS, especially under $LINEX_{\{C=-1\}}$ loss function that the estimators via SRS have a relatively larger MSEs.

5 Real data analysis

In order to compare the performance of Bayes estimates of the parameters based on SRS and RSS, taking two samples of size 5 from the following data set that are display the number of 1000s of cycles to failure for electrical appliances in a life test originally reported by Lawless (2003)

0.014, 0.034, 0.059, 0.061, 0.069, 0.08, 0.123, 0.142, 0.165, 0.21, 0.381, 0.464, 0.479, 0.556, 0.574, 0.839, 0.917, 0.969, 0.991, 1.064, 1.088, 1.091, 1.174, 1.27, 1.275, 1.35, 1.397, 1.477, 1.578, 1.649, 1.702, 1.893, 1.932, 2.001, 2.161, 2.292, 2.326, 2.337, 2.628, 2.785, 2.811, 2.886, 2.993, 3.122, 3.248, 3.715, 3.79, 3.857, 3.912, 4.1, 4.106, 4.116, 4.315, 4.51, 4.58, 5.267, 5.299, 5.583, 6.065, 9.701.

Table 3: MSEs and biases (in parenthesis) of the Bayesian estimates based on SRS and RSS for β and λ under Gamma prior.

n	SEL				C	LINEX			
	β		λ			β		λ	
	SRS	RSS	SRS	RSS		SRS	RSS	SRS	RSS
5	0.1892 (0.2897)	0.0896 (0.1245)	0.3023 (0.3094)	0.0345 (0.0941)	1	0.1577 (0.2617)	0.0795 (0.1133)	0.2214 (0.2542)	0.0303 (0.0811)
					-1	0.2282 (0.3203)	0.1011 (0.1361)	0.3919 (-0.3105)	0.0395 (0.1079)
6	0.0751 (0.1345)	0.0276 (0.0815)	0.2540 (0.2906)	0.0319 (0.0561)	1	0.0640 (0.1188)	0.0252 (0.0742)	0.2115 (0.2381)	0.0293 (0.0473)
					-1	0.0885 (0.1512)	0.0302 (0.0888)	0.3404 (-0.3267)	0.0349 (0.0654)
7	0.0689 (0.1613)	0.0261 (0.0598)	0.1257 (0.1443)	0.0186 (0.0451)	1	0.0607 (0.1479)	0.0239 (0.0540)	0.1029 (0.1162)	0.0174 (0.0386)
					-1	0.0784 (0.1752)	0.0295 (0.0656)	0.1575 (0.1769)	0.0199 (0.0519)

Table 4: MSEs and biases (in parenthesis) of the Bayesian estimates based on SRS and RSS for β and λ under non-informative prior.

n	SEL				C	LINEX			
	β		λ			β		λ	
	SRS	RSS	SRS	RSS		SRS	RSS	SRS	RSS
5	0.1248 (0.1323)	0.0435 (0.1005)	0.1604 (0.1117)	0.0265 (-0.0167)	1	0.0946 (0.1052)	0.0378 (0.0868)	0.1026 (0.0595)	0.0249 (-0.0281)
					-1	0.1709 (0.1631)	0.0496 (0.1145)	0.3718 (0.1967)	0.0287 (-0.0043)
6	0.0463 (0.0878)	0.0301 (0.0214)	0.0870 (0.0167)	0.0162 (-0.0193)	1	0.0400 (0.0696)	0.0279 (0.0140)	0.0701 (-0.0143)	0.0156 (-0.0270)
					-1	0.0541 (0.1072)	0.0324 (0.0289)	0.3565 (-0.3370)	0.0168 (-0.0112)
7	0.0430 (0.0650)	0.0203 (0.0796)	0.0771 (0.0485)	0.0142 (-0.0014)	1	0.0357 (0.0511)	0.0190 (0.0727)	0.0620 (0.0191)	0.0139 (-0.0074)
					-1	0.0522 (0.0797)	0.0218 (0.0866)	0.1135 (0.0837)	0.0147 (0.0048)

Table 5: The 95% HPD credible intervals of the Bayesian estimates based on SRS and RSS for β and λ under Gamma prior.

n	HPD			
	β		λ	
	SRS	RSS	SRS	RSS
5	(0.3964, 1.2513)	(0.3691, 0.9048)	(0.3204, 1.5472)	(0.3231, 0.9430)
6	(0.3345, 0.9792)	(0.3657, 0.8084)	(0.3294, 1.4674)	(0.3309, 0.8494)
7	(0.3758, 0.9793)	(0.3648, 0.7666)	(0.2841, 1.1684)	(0.3469, 0.7914)

Table 6: The 95% HPD credible intervals of the Bayesian estimates based on SRS and RSS for β and λ under non-informative prior.

n	HPD			
	β		λ	
	SRS	RSS	SRS	RSS
5	(0.2613, 1.0939)	(0.3090, 0.9136)	(0.1735, 1.3267)	(0.2432, 0.8208)
6	(0.2657, 0.9776)	(0.3018, 0.7612)	(0.1644, 1.0932)	(0.2718, 0.7490)
7	(0.2791, 0.8940)	(0.3637, 0.8096)	(0.1961, 1.0927)	(0.3093, 0.7377)

Table 7: Relative efficiency when β and λ are unknown.

n	non-informative					Gamma prior				
	SEL		LINEX			SEL		LINEX		
	β	λ	C	β	λ	β	λ	C	β	λ
5	2.8690	6.0528	1	2.5026	4.1205	2.1116	8.7623	1	1.9836	7.3069
			-1	3.4456	12.9547			-1	2.2572	9.9215
6	1.5382	5.3704	1	1.4337	4.4936	2.7210	7.9624	1	2.5397	7.2184
			-1	1.6697	21.2202			-1	2.9305	9.7536
7	2.1182	5.4296	1	1.8789	4.4604	2.6398	6.7581	1	2.5397	5.9138
			-1	2.3945	7.7211			-1	2.6576	7.9146

In the context of parameter estimation, Ammar et al.(2012) analyzed this data set based on the TPBL distribution. They showed that the TPBL model fits reasonably well to the data set. Therefore, we do not reproduce the results here for the sake of brevity. The drawn samples based on SRS and RSS are given in Table 8.

Now we compute the Bayesian estimations of β and λ based on SRS and RSS under SEL. Also, 95% HPD credible intervals of the Bayes estimates and the length of HPDs are constructed. All the results are based on the 5000 replications as reported in Table 9. It is obvious from this table that the length of HPD credible interval of β based on RSS is smaller than the length of HPD credible interval of β based on SRS. Also, the length of HPD of λ via RSS is smaller than its SRS counterparts.

Furthermore, by using the Kolmogorov-Smirnov (K-S) and Anderson-Darling tests of goodness-of-fit we check the validity of the TPBL distribution in Table 10. It indicates that the TPBL distribution that is estimated by RSS, provides p -values larger than via SRS to the real data set. In addition, Figure 1 displays the histogram of the real data set and the fitted density functions, and Figure 2 shows the Q-Q Plot of the different sampling schemes. It is observed from these figures that the TPBL distribution that is estimated by RSS, provide a better fit than by SRS to the data.

6 Conclusions

In this work, we consider Bayesian inference of the parameters for TPBL distribution based on SRS and RSS. We obtain the Bayesian estimation of the scale parameter when the shaped parameter is known. Also, Bayes estimators for the both parameters are derived using MCMC method. Furthermore, we construct the HPD credible intervals for the parameters. By a simulation study difference between the estimation procedures are compared. The Bayes

Table 8: The number of 1000s of cycles to failure data based on SRS and RSS; $n = 5$.

SRS		RSS	
1.893	0.123	0.556	1.355
5.267	2.993		0.210
			0.969
			2.886
			5.267

Table 9: Bayesian estimates, 95% HPD credible intervals and HPD lengths of β and λ .

	$\hat{\beta}$	HPD	HPD length	$\hat{\lambda}$	HPD	HPD length
SRS	0.4840	(0.2139, 0.7730)	0.5592	0.2869	(0.0643, 0.6907)	0.6263
RSS	0.5313	(0.2839, 0.7601)	0.4763	0.2542	(0.1021, 0.4698)	0.3677

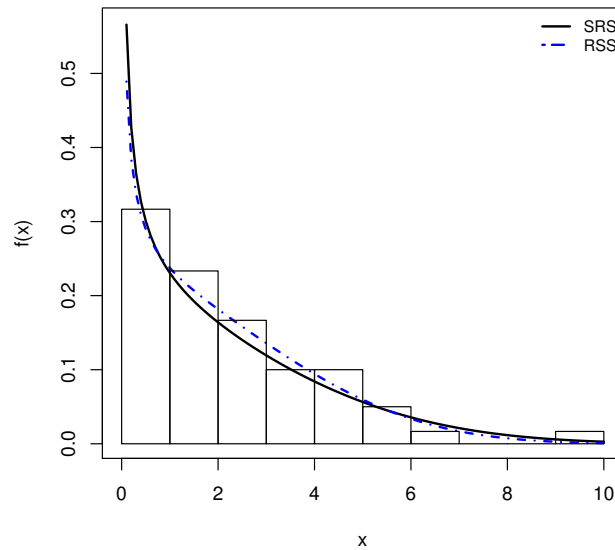


Figure 1: The histogram of real data set with the different fitted density functions.

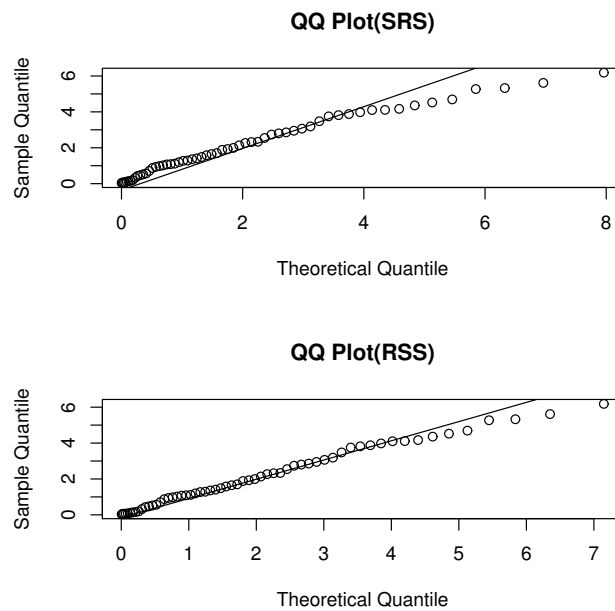


Figure 2: The QQ Plot of different sampling schemes.

Table 10: K-S and A-D tests for real data.

sampling scheme	K-S test		A-D test	
	Statistics D	p -value	Statistics A_n	p -value
SRS	0.1031	0.5135	0.5973	0.6498
RSS	0.0673	0.9316	0.2610	0.9642

estimates based on RSS are perform better than the Bayes estimates based on SRS. In the context of HPD credible interval estimation, the HPDs based on RSS are shorter than the HPDs based on SRS.

Acknowledgment

References

- [1] A.M. Sarhan, D.C. Hamilton, B. Smith, Parameter estimation for a two-parameter bathtub-shaped lifetime distribution, *Appl. Math. Modeling*, **36**, (2012) 5380-5392.
- [2] D. Ardia, *Lecture Notes in Economics and Mathematical Systems*, Springer-Verlag Berlin Heidelberg, 2008.
- [3] B.C. Arnold, N. Balakrishnan, H.N. Nagaraja, *A First Course in Order Statistics*, John Wiley & Sons, New York, 1992.
- [4] M.H. Chen, Q.M. Shao, Monte Carlo estimation of Bayesian credible and HPD intervals, *J. Comp. and Graphical Statistics*, **8**, (1999) 69-92.
- [5] Z. Chen, A new two- parameter lifetime distribution with bathtub shape or increasing failure rate function, *Stat. and Probabil. Letters*, **49**, (2000) 155-161.
- [6] A.E. Gelfand, A.F.M. Smith, Sampling-based approaches to calculating marginal densities, *J. Amer. Stat. Assoc.*, **85**, (1990) 398-409.
- [7] W.K. Hastings, Monte Carlo sampling methods using Markov chains and their applications, *Biometrika*, **57**, (2003) 97-109.
- [8] U. Hjorth, A reliability distribution with increasing, decreasing, and bathtub-shaped failure rate, *Technometrics*, **22** (1980) 99-107.
- [9] J.F. Lawless, *Statistical Models and Methods for Lifetime Data*, Wiley, New York, 2003.
- [10] J.-W. Wu, H.-L. Lu, C.-H. Chen, C.-H. Wu. Statistical inference about the shape parameter of the new two-parameter bathtub-shaped lifetime distribution, *Qual. Reliab. Eng. Int.*, **20** (2004) 607-616.
- [11] P.H. Kvam, F.J. Samaniego, Non-parametric maximum likelihood estimation based on ranked set sampling, *J. Amer. Stat. Assoc.*, **89** (1994) 526-537.
- [12] M. Lavine, The Bayesics of ranked set sampling, *J. Environ. Ecol. Stat.*, **6** (1999) 47-57.
- [13] L. M. Leemis, Lifetime distribution identities, *IEEE Transactions on Reliability*, **35(2)** (1986) 170-174.
- [14] N. Metropolis, A.W. Rosenbluth, M.N. Rosenbluth, A.H. Teller, E. Teller, Equations of state calculations by fast computing machines, *J. Chem. Phys.*, **21** (1953) 1087-1091.
- [15] G. A. McIntyre, A method for unbiased selective sampling using ranked sets, *Aust. J. of Agricultural Research*, **3(4)** (1952) 385-390.
- [16] H.A. Muttlak, W. Al-Sabah, Statistical quality control using ranked set sampling, *J. Appl. Stat.*, **30** (2003) 1055-1078.

- [17] G. Mudholkar, A method for unbiased selective sampling using ranked set, *Aust. J. Agri. Res.*, **3**, (1952) 385-390.
- [18] G. S. Mudholkar, D. K. Srivastava, Exponentiated Weibull family for analysing bathtub failure rate data, *IEEE Transactions on Reliability*, **42(2)** (1993) 299-302.
- [19] G.P. Patil, A.K. Sinha, C. Taillie, Relative precision of ranked set sampling: a comparison with the regression estimator, *Environmetrics*, **4** (1993) 399-412.
- [20] C. Robert, G. Casella, *Monte Carlo Statistical Methods*, Springer Text in Statistics, 2004.
- [21] G. Roberts, A. Smith, Simple Conditions for the Convergence of the Gibbs Sampler and Metropolis-Hastings Algorithm, *Stochastic Processes and their Applications*, **49(2)** (1994) 207-216.
- [22] A. Sadek, F. Alharbi, Weibull-Bayesian analysis based on ranked set sampling, *Int. J. Adv. Stat. Probab.*, **2** (2014) 114-123.
- [23] A. Sadek, K.S. Sultan, N. Balakrishnan, Bayesian estimation based on ranked set sampling using asymmetric loss function, *Bull. Malays. Math. Sci. Soci.*, **38** (2015) 707-718.
- [24] S.-J. Wu. Estimation of the two-parameter bathtub-shaped lifetime distribution with progressive censoring, *J. Appl. Stat.*, **35 (10)** (2008) 1139-1150.
- [25] R. M. Smith, L. J. Bain, An exponential power life-testing distribution, *Communications in Statistics - Theory and Methods*, **4** (1975) 469-481.
- [26] K. Takahasi, K. Wakimoto, On unbiased estimates of the population mean based on the sample stratified by means of ordering, *Ann. Inst. Stat. Math.*, **20** (1968) 1-31.

Mehdi Basikhasteh

Department of Statistics, Faculty of Sciences, Persian Gulf University, Bushehr 75169, Iran.
E-mail: mehdibasikhasteh@gmail.com

Fazlollah Lak

Department of Statistics, Faculty of Sciences, Persian Gulf University, Bushehr 75169, Iran.
E-mail: lak@pgu.ac.ir

Mahmoud Afshari

Department of Statistics, Faculty of Sciences, Persian Gulf University, Bushehr 75169, Iran.
E-mail: afshar.5050@gmail.com