REVIEW OF RANKED SET SAMPLING: MODIFICATIONS AND APPLICATIONS

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ABSTRACT
The problem of estimating the population mean is considered by McIntyre (1952). A new sampling method is suggested, namely; ranked set sampling (RSS) as efficient method compared to the well known simple random sampling (SRS) method. In the last years many authors suggested different modifications of the RSS and used it in wide applications. In this paper, a literature review of the RSS method is presented as well as some its modifications and applications are provided.

KEYWORDS: Ranked set sampling; Population mean; Efficiency.

MSC: 62G05, 94A20, 62P12

RESUMEN
El problema de estimar la media poblacional fue considerado por McIntyre (1952). Este sugirió un nuevo método de muestreo, llamado muestreo por conjuntos ordenados (Ranked Set Sampling, RSS) que consideraba como un método eficaz en comparación con el método del usual muestreo aleatorio simple. En los últimos años muchos autores sugirieron diversas modificaciones del RSS y les han usado en un amplio espectro de aplicaciones. En este trabajo se presenta una revisión de la literatura del método RSS así como algunas de sus modificaciones y aplicaciones.

1. INTRODUCTION

In this paper, we will present some studies where the well known ranked set sampling (RSS) method as well as some of its modifications are applied. The RSS was first suggested by McIntyre (1952) for estimating the population mean of pasture and forage yields. He claimed without proof that RSS was more accurate than simple random sample, its efficiency for estimating the higher population moments is better than that of simple random sampling (SRS) unless if the underlying distribution is rectangular in shape. The usual sampling designs are characterized as follows.

Definition: A randomly selected sample from a larger sample or population, giving all the individuals in the sample an equal chance to be chosen. (Cochran 1977).

RSS may be considered as a “controlled random sampling” design. It can be described as follows:

Step 1: Select randomly m^2 units from the population of interest.
Step 2: Allocate the m^2 units randomly into m sets, each set of size m.
Step 3: Rank the units in each set based on a variable of interest visually or by using any cost free method.
Step 4: The sample is chosen for actual measurement by selecting from the first set the lowest ranked unit, from the second set the second smallest ranked unit, and so forth until from the last set the maximum ranked unit is selected.
Step 5: The above steps can be repeated n cycles to get a sample of size mn.

For fixing some ideas, consider a random sample from a distribution $F(x)$, which admits a density function $f(x)$, with a mean μ and a variance σ^2. With compared to SRS, RSS uses one unit, namely, $X_{1(1:m)}$, the smallest unit from this set, then $X_{2(2:m)}$, the second smallest ranked unit from another independent set of m units, and finally $X_{m(m:m)}$, the largest ranked unit from a last set of m units. This

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process can be explained in Figure 1. It is important to emphasize here, that although RSS require identification of as many as m^2 units, but only m of them are quantified.

The final m units $X_{1(1:m)}, X_{2(2:m)}, \ldots, X_{m(m:m)}$ are used for investigation. These units are independent but not identically distributed, and $X_{i(i:m)}$, is the ith order statistic in a random sample of size m from $F(x)$. Thus making a comparison of a RSS of size m with a SRS of the same size m is meaningful. Obviously, RSS would be a good rival to SRS in case where the collecting of the sampling units is easy and their relative rankings based on the characteristic under study can be done with trivial cost.

$\left(X_{1(1:m)}\right)$	$X_{1(2:m)}$	$\cdots$	$X_{1(m:m)}$
$X_{2(1:m)}$	$\left(X_{2(2:m)}\right)$	$\cdots$	$X_{2(m:m)}$
$\vdots$	$\vdots$	$\vdots$	$\vdots$
$X_{m(1:m)}$	$X_{m(2:m)}$	$\cdots$	$\left(X_{m(m:m)}\right)$

Figure 1: Elucidation of m^2 units in m sets of m each

The efficiency of RSS relies on the sampling allocation, either balanced or unbalanced. In balanced RSS, the rank order statistics has an equal allocation. They proved theoretically and pretended empirically that the balanced RSS estimator has a variance which is no greater than the SRS estimator variance whether errors in ranking or the nature of the parent distribution of the variable of interest.

In SRS the sampler must increase the sample size to increase the chance of coverage the whole range of possible observations values and there is no other chance. While in RSS, one can increase the representativeness based on a specific number of sample observations. Hence, there is a saving considerably on the measurement costs. Thus, based on the measured ranked set sample, we can obtain unbiased estimators of population parameters, as the mean and, and for more than one cycle, the population variance. The relative precision (RP) of RSS relative to SRS based on the same number of measured units is defined as $RP = \frac{Var\left(\bar{X}_{SRS}\right)}{Var\left(\bar{X}_{RSS}\right)} \in \left[1, \frac{m+1}{2}\right]$. Note that the RSS method cannot be worse than the SRS method (Patil 2002; Takahasi and Wakimoto 1968).

2. REVIEW ON SOME PREVIOUS APPLICATIONS OF RSS

Fortunately, in many fields, such as in medicine, environment, biology and agriculture, the study variable is not easily measured, but its ranking can be done easily with cheap or free cost. The RSS can be implemented to yield more efficient estimators of the population parameters as compared to SRS using the same number of quantified observations. Here, some examples on reported applications of RSS in real situations will be given.

Evans (1967) applied the RSS to regeneration surveys in areas direct-seeded to longleaf pine. He noted that the means based on both of RSS and SRS methods were not significantly different, but the computed variances of the means were very different. Martin et al. (1980) applied the RSS procedure for estimating shrub phytomass in Appalachian Oak forests. Cobby et al. (1985) conducted four experiments at Hurley (UK) during 1983 to investigate the performance of RSS relative to SRS for estimation of herbage mass in pure grass swards, and of herbage mass and clover content in mixed grass-clover swards. Johnson et al. (1993) applied RSS method to estimate the mean of forest, grassland and other vegetation resources. Mode et al. (1999) investigated under which conditions the RSS becomes a cost-effective sampling method for ecological and environmental field studies where the rough but cheap measurement has a cost. They have introduced formula for the total cost for both RSS and SRS, and present cost ratios for a real data set consisting of judgment estimated and physically measured stream. Al-Saleh and Al-Shrafat (2001) studied the performance of RSS in estimation milk yield based on 402 sheep. Al-Saleh and Al-Omari (2002) used the multistage RSS to estimate the average of Olives yields in a field in West of Jordan. Husby et al. (2005) investigated on the use of the RSS in estimating of the mean and median of a population using the crop production dataset from the United State Department of Agriculture. They found that the gain in efficiency for mean estimation using RSS is better for symmetric distribution than asymmetric distribution, and vice versa in the case of median estimation. Kowalczyk (2005) applied the RSS procedure in market and consumer surveys.

Ganeslingam and Ganesh (2006) applied the RSS method to estimate the population mean and the ratio using a real data set on body measurement. The authors used the data of the weight and height of 507 individuals. Halls and Dell (1966) coined McIntyre's method as RSS and applied it for estimating the weights of browse and herbage in a pine-hardwood forest of east Texas, USA. Wang et al. (2009) used the RSS in fisheries research. Tiwari and Pandey. (2013) considered an application of ranked set sampling design in environmental investigations for real data set. For more about applications of RSS see Dong et al. (2012).

3. ESTIMATION USING RSS

3.1. Estimation of the population mean

Assume, that the population under consideration has a density function $f(x)$ with mean μ and variance σ^2.

Theorem 1: An unbiased estimator of a population mean μ using SRS is given by $\bar{X}_{SRS} = \frac{1}{m}\sum_{i=1}^{m} X_i$, with variance $\operatorname{Var}\left(\bar{X}_{SRS}\right) = \frac{\sigma^2}{m}$.

McIntyre (1952) claimed without proof that:

1) Regardless of any errors in ranking, the RSS estimator of the population mean $\bar{X}_{RSS}$ is unbiased.
2) Under perfect ranking the efficiency of RSS with respect to SRS is nearly $\frac{m+1}{2}$ for estimating the mean of typical unimodal distributions based on the same number of quantified units.
3) In estimating the higher population moments, the efficiency of RSS is better than that of SRS unless if the underlying distribution has a rectangular shape.
4) The ranking errors reduce the efficiency of RSS and these errors increase as sample size increases.
5) If there are enough information about the underlying distribution the unequal allocations may improve the performance of RSS.

Takahasi and Wakimoto (1968) have given the mathematical properties of RSS and obtain the same method independently of McIntyre (1952).

Theorem 2: Let $X_{(i:m)}$ be *i*th order statistics of a sample of size *m* from a probability density function $f_{(i:m)}(x)$, with mean $\mu_{(i:m)}$ and variance $\sigma^2_{(i:m)}$. An unbiased estimator of the population mean using RSS is given by $\bar{X}_{RSS} = \frac{1}{m}\sum_{i=1}^{m} X_{i(i:m)}$, with variance $\operatorname{Var}\left(\bar{X}_{RSS}\right) = \frac{\sigma^2}{m} - \frac{1}{m^2}\sum_{i=1}^{m}\left(\mu_{i(i:m)} - \mu\right)^2$. Also,

$$f(x) = \frac{1}{m}\sum_{i=1}^{m} f_{(i:m)}(x), \quad \mu = \frac{1}{m}\sum_{i=1}^{m}\mu_{(i:m)}, \quad \sigma^2 = \frac{1}{m}\sum_{i=1}^{m}\sigma^2_{(i:m)} + \frac{1}{m}\sum_{i=1}^{m}(\mu_{(i:m)} - \mu)^2.$$

Let nm units be randomly chosen from the population of interest and randomly allocated into n sets, each of size m units. From each set of size *m* one unit will be drawn to get n measured units. The units to be measured were chosen as in the following steps. First, let $l_1, l_2, \ldots, l_m$ be positive integers such that $l_1 + l_2 + \ldots + l_m = n.$ After ranking the units within each set with respect to the variable of study, the smallest ranked unit is measured from the first l_1 sets; the second smallest ranked unit is measured from the next l_2 sets, and so forth until the largest ranked unit is measured from the last l_m sets. Let T_i be the sum of measurements of the *i*th ranked units for $i = 1, 2, \ldots, m$. So that, T_i are independent. Therefore, the unbiased RSS estimator of μ is $\bar{X}_{RSS} = \frac{1}{m}\sum_{i=1}^{m}\frac{T_i}{l_i}$.

Based on both balanced allocations with $l_1 = l_2 = \ldots = l_m$, and Neyman allocations when each l_i proportional to $\sigma_{(i:m)}$ they compared the performance of $\bar{X}_{RSS}$ with respect to $\bar{X}_{SRS}$ using the relative precision (RP), $RP=\frac{\operatorname{Var}\left(\bar{X}_{SRS}\right)}{\operatorname{Var}\left(\bar{X}_{RSS}\right)}$, or the equivalent relative savings (RS), $RS=1-\frac{1}{RP}$. For balanced and Neyman allocations we have

$$\operatorname{Var}\left(\bar{X}_{RSS}\right)=\frac{1}{m^2}\sum_{i=1}^{m}\frac{\sigma_{(i:m)}^2}{l_i},\ \text{and}\ \operatorname{Var}\left(\bar{X}_{RSS}\right)=\frac{1}{n}\left(\frac{1}{m}\sum_{i=1}^{m}\sigma_{(i:m)}^2\right)^2,$$

respectively. Also, for balanced allocations, the relative saving is $RS=\frac{1}{m\sigma^2}\sum_{i=1}^{m}\left(\mu_{(i:m)}-\mu\right)^2$, where $0\le RS\le\frac{m-1}{m+1}$, and $1\le RP\le\frac{m+1}{2}$. The lower bound is holds if and only if the underlying distribution is degenerate, and the upper bound is holds if and only if the parent distribution is rectangular. But for Neyman allocations they showed that $0\le RS\le\frac{m-1}{m}$, and $1\le RP\le m.$ See Takahasi and Wakimoto (1968) for more details.

Al-Saleh and Al-Omari (2002) introduced a multistage ranked set sampling (MSRSS) as a generalization of the double RSS. The MSRSS procedure can be described as:

Step 1: Randomly select m^{r+1} units from the population of interest, where r is the number of stages and m is the sample size.

Step 2: Allocate the m^{r+1} selected units as randomly as possible into m^{r-1} sets, each of size m^2.

Step 3: For every set in Step (2), use the procedure of balanced ranked set sampling as described in Section 1 to have a ranked set sample of size m. This step yields m^{r-1} ranked set samples each of size m.

Step 4: Repeat Step (3) on the m^{r-1} ranked set samples to obtain m^{r-2} second stage RSS samples each of size m. The process continues until we end up with one rth stage RSS of size m.

Suppose that the variable of interest X has mean $\mu,$ and variance σ^2 with a pdf $f(x)$ and cdf $F(x)$. Let $X_1^{(r)},X_2^{(r)},...,X_m^{(r)}$ be a MSRSS of size m at stage r, with mean $\mu_i^{(r)}$, variance $\sigma_i^{2(r)}$, pdf $f_i^{(r)}(x)$ and cdf $F_i^{(r)}(x)$, $i=1,2,...,m$. The MSRSS estimator of the mean is $\bar{X}_{MSRSS}^{(r)}=\frac{1}{m}\sum_{i=1}^{m}X_i^{(r)}$, with variance

$$\operatorname{Var}\left(\bar{X}_{MSRSS}^{(r)}\right)=\frac{1}{m^2}\sum_{i=1}^{m}\sigma_i^{2(r)}=\frac{\sigma^2}{m}-\frac{1}{m^2}\sum_{i=1}^{m}\left(\mu_i^{(r)}-\mu\right)^2.$$

The authors proved the following identities

$$f(x)=\frac{1}{m}\sum_{i=1}^{m}f_i^{(r)}(x),\ \ \mu=\frac{1}{m}\sum_{i=1}^{m}\mu_i^{(r)},\ \ \sigma^2=\frac{1}{m}\sum_{i=1}^{m}\sigma_i^{2(r)}+\frac{1}{m}\sum_{i=1}^{m}\left(\mu_i^{(r)}-\mu\right)^2,$$

and defined the RS and RP at stage r, respectively, as

$$RS^{(r)}=\frac{1}{m\sigma^2}\sum_{i=1}^{m}\left(\mu_i^{(r)}-\mu\right)^2,\ \text{and}\ RP^{(r)}=1+\left(\sum_{i=1}^{m}\left(\mu_i^{(r)}-\mu\right)^2\Big/\sum_{i=1}^{m}\sigma_i^{2(r)}\right).$$

The authors defined a steady state efficiency of RSS at stage r to be as $eff^{(\infty)}=\lim_{r\to\infty}eff^{(r)}$, and they proved the following theorem.

Theorem 3: Suppose that the variable of interest has an absolutely continuous distribution function and let $X_1^{(r)},X_2^{(r)},...,X_m^{(r)}$ be a MSRSS of size m, then

$$F_i^{(\infty)}(x)=\lim_{r\to\infty}F_i^{(r)}(x)=\begin{cases}0, & x<Q_{(i-1)/m},\\ mF(x)-(i-1), & Q_{(i-1)/m}\le x<Q_{i/m},\\ 1, & x\ge Q_{i/m},\end{cases}$$

where Q_β is the quantity which satisfies $\int_{-\infty}^{Q_\beta}f(x)dx=\beta$, $\beta\in(0,1)$.

So that $f_i^{(r)}(x) \to f_i^{(\infty)}(x) = mf(x)$ for $Q_{(i-1)/m} \le x \le Q_{i/m}$, and zero otherwise. If the variable of interest X has a uniform distribution, $U(0,\theta)$, then $f_i^{(\infty)}(x) = \frac{m}{\theta}$ for $\left(i - \frac{1}{m}\right)\theta < x < \left(\frac{i}{m}\right)\theta,$ and zero otherwise, and $eff^{(\infty)} = m^2$. For the special case, if we sample from the standard uniform distribution, it can be shown that an approximate integrals for the above function can be found by Monte Carlo methods. This idea was considered by Al-Saleh and Samawi (2000). It can be noted that the efficiency is increasing in terms of the sample size and the number of stages. In Table 3.1 some values of the efficiency of the MSRSS with respect to the SRS estimators are summarized. See Al-Saleh and Al-Omari (2002).

Table 3.1: The efficiency of MSRSS relative to SRS for *m*=2,3

Distribution	$m=2$				$m=3$
	$r=1$	$r=2$	$r=3$	$r\to\infty$	$r\to\infty$
$U(0,\theta)$	1.500	1.923	2.269	4.000	9.000
$N(\theta,1)$	1.467	1.797	2.006	2.752	4.839
Logistic $(\theta,1)$	1.436	1.705	1.877	2.400	3.831
Exp(θ)	1.333	1.516	1.625	1.923	2.843
LogN$(\theta,1)$	1.186	1.257	1.293	1.371	2.708
Pareto $(\theta,3)$	1.135	1.182	1.205	1.249	1.437

For mean estimation based on RSS and its modifications: Samawi et al. (1996) suggested a variety of extreme RSS. Muttlak (1997) introduced a median ranked set sampling. Samawi (2002) suggested double extreme ranked set sampling. Yu and Tam (2002) proposed the RSS in the presence of concord data. Al-Saleh and Al-Hadhrami (2003) investigated moving extremes RSS parametrically for estimating the location parameter of symmetric distribution. Muttlak (2003a,b) suggested percentile and quartile RSS methods. Rahimov and Muttlak (2003) extend the random selection in RSS suggested by Li et al. (1999) for estimating the population mean. Barabisi and Pisani (2002) investigated steady state RSS for replicated sampling protocols in order to estimate the objective parameter using Horvitz-Thompson estimator. For asymmetric distribution Muttlak and Abu-Dayyeh (2004) suggested weighted modified RSS to overcome the bias of several estimators based on modified RSS methods such as ERSS, MRSS and PRSS methods. Bouza (2008) considered the mean estimation when some observations are missing using product type estimators. Bouza (2009) investigated the mean estimation of a sensitive quantitative character based on RSS and randomized response procedures. Bouza (2010) considered the mean estimation using RSS in non-responses case. Jeelani et al. (2014) considered a role of rank set sampling in improving the estimates of population mean under stratification. Jozani et al. (2012) introduced unbiased ratio estimators of the population mean using ranked set sampling. Bani-Mustafa et al (2011) suggested folded RSS for estimating the population mean of asymmetric distributions. Singh et al. (2014) introduced a general procedure for the mean estimation using RSS.
Patil et al. (1997) investigated the effect of the sample size upon the performance of the balanced RSS for estimating the population mean. Hossain (2001) suggested a nonparametric approach for the modified RSS method for the population mean estimation, namely, nonparametric selected ranked set sampling. Unlike the usual RSS where we chose only one unit from each ranked set of size *m*, Wang et al. (2004) proposed an estimator of the population mean using the general RSS. In which more than one unit can be chosen from each ranked set.

Al-Saleh and Al-Kadiri (2000) proposed the double RSS procedure (DRSS) for the mean estimation. They showed that at the second stage the ranking is easier than ranking at the first stage, and also the DRSS estimator is more efficient than that using RSS with respect to SRS based on the same sample size.

Al-Saleh et al. (2000) considered Bayesian estimation of the parameter of the underlying distribution using RSS. In terms of the Bayes risk, the Bayes risk of the Bayes estimator using RSS method is less than the Bayes risk of the Bayes estimator using SRS. The procedure was used for estimating the average milk yield of 402 sheep.

Also, see Jemain and Al-Omari (2006a, 2006b, 2006c, 2007a, 2007b). Al-Omari, (1999, 2011), Al-Omari and Al-Saleh (2012), Jemain et al. (2007a, 2007b, 2008a, 2008b), Chen et al. (2004), Al-Omari and Raqab (2013), Kominiak and Mahdizadeh (2014), Takahasi (1969), Bouza (2002, 2013), Al-Nasser and Al-Omari (2014), Haq et al. (2014b), Al-Nasser (2007), Syam et al. (2012), Ohyama et al. (2008), Haq et al. (2013), Mehta and Mandowara (2013), Alodat and Al-Saleh (2001), Wang et al. (2008), Al-Nasser and Bani-Mustafa (2009), Al-Hadhrami and Al-Omari (2014), Sinha (2005), Wolfe (2012), and Syam et al. (2013a, 2013b).

3.2. Estimation of the variance

The SRS estimator of the population variance σ^2 is given by

$$\hat{\sigma}^2_{SRS} = \frac{1}{m-1}\sum_{i=1}^{m}\left(X_{(i:m)} - \bar{X}_{SRS}\right)^2.$$

An earlier work for estimating the population variance is considered by Stokes (1980a). Based on judgment ordered using balanced RSS she defined

$$\hat{\sigma}^2_{RSS} = \frac{1}{nm-1}\sum_{h=1}^{n}\sum_{i=1}^{m}\left(X_{[i:m]h} - \bar{X}_{RSS}\right)^2,$$

where $X_{[i:m]h}$ is the quantification of the *i*th ranked unit in a set of size *m* in the *h*th replicate. She showed that

$$E\left(\hat{\sigma}^2_{RSS}\right) = \sigma^2 + \frac{1}{nm-1}\sum_{i=1}^{m}\left(\mu_{[i:m]} - \mu\right)^2,$$

i.e., $\hat{\sigma}^2_{RSS}$ is a biased estimator of the population variance. However, the bias approach to zero as the number of measurements nm becomes large. Based on the ratio $RP = \frac{\mathrm{Var}\left(\hat{\sigma}^2_{SRS}\right)}{\mathrm{MES}\left(\hat{\sigma}^2_{RSS}\right)}$ the performance of RSS is investigated and proved that $\lim_{n\to\infty} RP \geq 1$. The author concluded that the gain in efficiency of RSS over SRS is little when estimating higher moments.
A nonparametric study is considered by Perron et al. (2004) for the estimation of the population variance σ^2 under ranked set sample. Biswis et al. (2013) considered variance estimation using Jackknife technique in ranked set sampling based on finite population framework. Also, see MacEachern et al. (2002), Al-Hadhrami and Al-Omari (2006), Al-Hadhrami (2010a), Chen and Lim (2011), Abu-Dayyeh and Al-Subh (2013).

3.3. Estimation of the population ratio

The population ratio of two variables *X* and *Y* is defined as $R = \frac{\mu_Y}{\mu_X}$. The SRS estimator of the population ratio is

$$R_{SRS} = \frac{\sum_{i=1}^{m} y_i}{\sum_{i=1}^{m} x_i} = \frac{\bar{y}}{\bar{x}}.$$

This estimator is biased since the denominator $\bar{x}$ as well as the numerator $\bar{y}$ are random variables. Samawi and Muttlak (1996) suggested an estimator of the population ratio based on ranked set sampling as

$$\hat{R}_{RSS} = \frac{\frac{1}{m}\sum_{i=1}^{m} X_{i(i)}}{\frac{1}{m}\sum_{i=1}^{m} Y_{i[i]}} = \frac{\bar{Y}_{RSS}}{\bar{X}_{RSS}},$$

where $(\bullet)$ and $[\bullet]$ denote that the ranking of *X* is perfect and the ranking of *Y* has errors. The variance of $\hat{R}_{RSS}$ is given by

$$\operatorname{Var}\left(\hat{R}_{RSS}\right) \cong \frac{R^2}{m}\left[\frac{\sigma_X^2}{\mu_X^2}+\frac{\sigma_Y^2}{\mu_Y^2}-2\rho\frac{\sigma_X\sigma_Y}{\mu_X\mu_Y}-\left(\frac{1}{m\mu_X^2}\sum_{i=1}^{m}\tau_{X(i)}^2+\frac{1}{m\mu_Y^2}\sum_{i=1}^{m}\tau_{Y[i]}^2-\frac{2}{m\mu_X\mu_Y}\sum_{i=1}^{m}\tau_{XY(i)}^2\right)\right],$$

where

$$\tau_{X(i)}=\mu_{X(i)}-\mu_X,\ \tau_{Y[i]}=\mu_{Y[i]}-\mu_Y \text{ and } \tau_{XY(i)}=\left(\mu_{X(i)}-\mu_X\right)\left(\mu_{Y[i]}-\mu_Y\right).$$

Based on Table 3.2 it is clear that the RSS is more efficient than SRS for estimating the population ratio.

Table 3.2: The efficiency and bias values of $\hat{R}_{RSS}$ with respect to $\hat{R}_{SRS}$ for $m=5,7$.

ρ		$m=5$	$m=7$	ρ	$m=5$	$m=7$
0.99	*Efficiency*	2.9231	3.7002	**-0.99**	3.0626	3.7815
	Bias RSS	0.0033	0.0016		0.0084	0.0043
	Bias SRS	0.0085	0.0058		0.0252	0.0187
0.80	*Efficiency*	2.1716	2.4952	**-0.80**	2.7928	3.4399
	Bias RSS	0.0032	0.0015		0.0071	0.0042
	Bias SRS	0.0095	0.0070		0.0234	0.0170
0.60	*Efficiency*	2.0774	2.3413	**-0.60**	2.6175	3.1533
	Bias RSS	0.0036	0.0019		0.0071	0.0034
	Bias SRS	0.0108	0.0088		0.0212	0.0159
0.50	*Efficiency*	2.0722	2.3104	**-0.50**	2.4993	3.0319
	Bias RSS	0.0045	0.0024		0.0074	0.0044
	Bias SRS	0.0133	0.0090		0.0200	0.0150
0.20	*Efficiency*	2.1142	2.3685	**-0.20**	2.3399	2.6955
	Bias RSS	0.0066	0.0026		0.0063	0.0022
	Bias SRS	0.0147	0.0115		0.0182	0.0131
0.10	*Efficiency*	2.1938	2.4334	**-0.10**	2.2632	2.5744
	Bias RSS	0.0057	0.0033		0.0057	0.0036
	Bias SRS	0.0154	0.0104		0.0166	0.0122

Samawi and Muttlak (2001) used the median RSS to estimate the population ratio. Samawi and Tawalbeh (2002) introduced a double median RSS for estimating the population mean and ratio. Al-Omari and Bouza (2014) considered ratio estimators of the population mean with missing values using ranked set sampling For more about ratio estimation in RSS see Samawi and Saeid (2004), Al-Omari et al. (2009), Mandowara and Mehta (2014), Kadilar et al. (2009), and Al-Omari (2012).

3.4. Estimation of the quantile

Let *X* be a random variable with cumulative distribution function $F(x)$. The *p*th quantile is defined as, $\xi_p=\inf\{x:F(x)\geq p\}$ for $0\leq p\leq 1$. The following authors have done works to estimate the *p*th quantile by different procedures as given below:

Chen (2000) studied quantile estimation based on balanced RSS data and concluded that the RSS scheme can substantially develop the efficiency of quantile estimation. Chen (2001b) further generalized the results in Chen (2000) from balanced to unbalanced scheme. Indeed, the quantile estimator proposed in both Chen (2000, 2001b) is construced in terms of the empirical distribution of the pooled RSS data. Kaur et al. (2002) proposed RSS sign test for population quantiles and identifies the optimal allocation, based on the quantile obtained, but not based on the underlying distribution.
Adatia and Saleh (2004) applied the generalized RSS method in estimating quantiles of the uniform distribution. Zhu and Wang (2004) considered quantile estimation using RSS under perfect ranking. Also, see Samawi (2001).

3.5. Estimation of the population proportion

Lacayo et al. (2002) investigated a population proportion estimation using RSS in the situations where the binary variable of study is selected from a continuous variable. Terpstra (2004) used RSS procedure to estimate a population proportion and considered two estimators, the sample proportion and maximum likelihood estimator of the RSS data. Chen et al. (2005) investigated the use of RSS in estimating the population proportion. Chen et al. (2006b) was the first to suggest the use of the unbalanced ranked set sampling for estimating a population proportion under perfect ranking. Also, see Terpstra and Nelson (2005), and Bouza (2013).

3.6. Estimation of the distribution function

This section is devoted for estimating the distribution function under RSS. Indeed, it is of interest to estimate the distribution function to investigate the parent distribution properties such as the skewness and multimodality. Also, the estimator of the underlying distribution is required for constructing the confidence intervals and hypothesis testing about some parameters. Some results for the distribution function estimation under RSS are given as follows.

Let $F_{SRS}(x)$ denote the empirical distribution function of a SRS X_1, X_2, …, X_m from $F(x)$ defined as

$$F_{SRS}(x) = \frac{1}{m}\sum_{i=1}^{m} I\left(X_i \le x\right),$$

where $I(\cdot)$ is an indicator function. It is clear that $E\left[F_{SRS}(x)\right] = F(x)$, with

$$Var\left(F_{SRS}(x)\right) = \frac{1}{m}F(x)\left[1 - F(x)\right].$$

Also, $F_{SRS}(x)$ is a consistent estimator of $F(x)$, (See Bahadur (1996)).

Stokes and Sager (1998) used RSS to estimate $F(x)$ for fixed *x* and suggested the estimator as

$$F_{RSS}(x) = \frac{1}{m}\sum_{i=1}^{m} I\left(X_{i(i:m)} \le x\right).$$

They proved that $F_{RSS}(x)$ is an unbiased estimator for $F(x)$, with variance

$$Var\left(F_{RSS}(x)\right) = \frac{1}{m^2}\sum_{i=1}^{m} F_{(i:m)}(x)\left[1 - F_{(i:m)}(x)\right],$$

and $\dfrac{F_{RSS}(x) - E\left(F_{RSS}(x)\right)}{\sqrt{Var\left(F_{RSS}(x)\right)}}$ converges in distribution to the standard normal distribution as $m \to \infty$ when *x* and *n* are fixed. The relative efficiency of RSS with respect to SRS for estimating the distribution function is defined as

$$Eff\left[F_{RSS}(x), F_{SRS}(x)\right] = \frac{\mathrm{Var}\left(F_{SRS}(x)\right)}{\mathrm{Var}\left(F_{RSS}(x)\right)}.$$

Table 3.3: The relative precision of $F_{RSS}(x)$ with respect to $F_{SRS}(x)$ with $m = 4,5,6,7,8$

$F(x)$	$m=4$	$m=5$	$m=6$	$m=7$	$m=8$
0.01	0.9886	1.0851	1.0388	1.0644	1.1355
0.30	1.6880	1.8933	2.0402	2.2060	2.3306
0.40	1.7891	1.9969	2.2099	2.2982	2.4696
0.50	1.8487	2.0644	2.2449	2.3500	2.5601
0.60	1.7841	1.9545	2.2061	2.3337	2.4787
0.70	1.6816	1.9025	2.0456	2.1952	2.3528
0.99	1.1206	1.0146	1.0458	1.0781	1.0523

From Table 3.3 we can see that $F_{RSS}(x)$ is more efficient than $F_{SRS}(x)$. See Stokes and Sager (1998) for more details.

Samawi and Al-Sageer (2001) studied the distribution function estimation using extreme and median RSS methods. Barabesi and Fattorini (2002) suggested kernel estimators of probability density functions based on RSS. Kim et al. (2005) mixed extreme RSS and median RSS to produce extreme median RSS for estimating the distribution function. Lam et al. (2002) suggested nonparametric estimators for the distribution function and the mean using the auxiliary information and concomitant variable in RSS process. Ozturk (2002a) suggested a new estimator of the distribution function and the center of symmetry. Gulati (2004) have considered smooth non-parametric estimates of the distribution function based on RSS. Frey. (2014) considered bootstrap confidence bands for the CDF using ranked-set sampling. Lim et al. (2014) investigated a kernel density estimation using RSS. Abu-Dayyeh et al. (2002) considered the distribution function estimation using double ranked set sampling method. Also, see Wolfe (2004), Baraneso and Fattorini (2002), and Huang (1972).

3.7. Regression in ranked set sampling

Philip and Lam (1997) assumed that the ranking is done on the basis of a concomitant variable *X* associated with the dependent variable *Y* which is expensive to measure. In the case of the mean of *X* is known, and both variables are positively correlated, the authors follow the same idea of Stokes (1977) where the ranking is done by means of the concomitant variable which can be measured easily compared to the main variable *Y*. According to Stokes (1997) they assumed that the regression of Y on X is linear as

$$Y = \mu_Y + \rho \frac{\sigma_Y}{\sigma_X}\left(X - \mu_X \right) + \varepsilon,$$

where *X* and ε are independent, ε has zero mean and variance $\sigma_Y^2\left(1-\rho^2\right)$. For known μ_X, they considered the difference estimator

$$\bar{Y}_D = \bar{Y}_{RSS} + B\left(\mu_X - \bar{X}_{RSS}\right),$$

with variance

$$\text{Var}\left(\bar{Y}_D\right) = B^2 \text{Var}\left(\bar{X}_{RSS}\right) - 2B\rho \frac{\sigma_Y}{\sigma_X} \text{Var}\left(\bar{X}_{RSS}\right) + \text{Var}\left(\bar{Y}_{RSS}\right),$$

and

$$\hat{B} = \frac{\sum_{h=1}^{n}\sum_{i=1}^{m}\left(X_{(i:m)h} - \bar{X}_{RSS}\right)\left(Y_{[i:m]h} - \bar{Y}_{RSS}\right)}{\sum_{h=1}^{n}\sum_{i=1}^{m}\left(X_{(i:m)h} - \bar{X}_{RSS}\right)^2},$$

where $X_{(i:m)h}$ is the *i*th smallest and $Y_{[i:m]h}$ is the corresponding value of *Y* get from the *i*th sample in the *h*th cycle. The RSS regression estimator for μ_Y is given by

$$\bar{Y}_{reg} = \bar{Y}_{RSS} + \hat{B}\left(\mu_X - \bar{X}_{RSS}\right),$$

which is an unbiased of μ_Y with variance

$$\text{Var}\left(\bar{Y}_{reg}\right) = \frac{\sigma_Y^2}{mn}\left(1-\rho^2\right)\left[1 + E\left(\frac{\bar{Z}_{RSS}^2}{s_z^2}\right)\right],$$

where, $Z_{(i:m)h} = \dfrac{X_{(i:m)h} - \mu_X}{\sigma_X}$, $\bar{Z}_{RSS} = \dfrac{1}{mn}\sum_{h=1}^{n}\sum_{i=1}^{m} Z_{(i:m)h}$ and $S_Z^2 = \dfrac{1}{mn}\sum_{h=1}^{n}\sum_{i=1}^{m}\left(Z_{(i:m)h} - \bar{Z}_{RSS}\right)^2$.

If μ_X is unknown, the double sampling scheme can be used to estimate μ_X. If RSS is implemented in the second-phase sampling, the double sampling regression estimator of μ_Y is given by

$$\bar{Y}_{ds} = \bar{Y}_{RSS} + \hat{B}\left(\bar{x}' - \bar{X}_{RSS}\right),$$

with variance

$$\text{Var}\left(\bar{Y}_{ds}\right) = \frac{\sigma_Y^2}{mn} + \frac{\left(\bar{Z}_{RSS} - \bar{Z}\right)^2}{mnS_Z^2} + \frac{1}{n^2 m},$$

where $\bar{Z} = \dfrac{\bar{x}' - \mu_X}{\sigma_X}$. The authors showed that the RSS regression estimator is more efficient than RSS and SRS naive estimators if *X* and *Y* are jointly from a bivariate normal distribution unless $|\rho| < 0.4$.

For more about regression estimation in RSS see Patil et al. (1993), Muttlak (1995), Muttlak (1996), Barreto and Barnett (1999), Barnett and Moore (1997), Chen (2001a), Ozturk (2002b), Badmus et al. (2012), Alodat et al. (2010), Murff and Sager (2006), Al-Odat et al. (2009), Al-Odat and Jetschke (2011), Samawi and Ababneh (2001), Alodat et al. (2009), Samawi and Abu-Dayyeh (2002), Demir and Çingi (2000), and Samawi and Al-Saleh (2002).

3.8. Estimation of the distribution parameters

This section summarizes some of the works where the RSS is used to estimate the unknown parameters of the distribution function.
Stokes (1976) used RSS for estimating the scale and location parameters, variance, interval, and for estimating a correlation coefficient and test of correlation.
Stokes (1980b) investigated the performance of RSS for correlation coefficient estimation in a bivariate normal distribution using the maximum likelihood estimator.
Fei et al. (1994) investigated the performance of RSS in the estimation of the parameters of Weibull distribution.
Kvam and Samaniego (1993,1994) proposed estimation of the population mean and population distribution function under unbalanced ranked-set samples.
Lam et al. (1994) used RSS to estimate the location, scale and quantile of the exponential distribution.
Li et al. (1999) introduced the concept of random selection in RSS with application to estimate the mean and the variance of the normal distribution, and also applied the method when the parent distribution is exponential and logistic.
Shen (1994) investigated the performance of ranked set sampling for estimating the mean of a lognormal distribution with a known coefficient of variation.
Stokes (1995) has considered a best linear unbiased and the maximum likelihood estimation of μ and σ for the location-scale family of random variables with distribution function $F\left((y-\mu)/\sigma\right)$, with known *F*.
Adatia (2000) generalized the ranked set sampling and used it for estimating the location and scale parameters of the half-logistic distribution.
Raqap et al. (2002) extend the works of Sinha et al. (1996) and Stokes (1995) by proposing best linear invariant estimators in RSS (RSS BLIEs) of the scale and location parameters separately, provided thata the nuisance parameter is unknown for symmetric and symmetric distributions.
Al-Saleh (2004a) proposed balanced ranked set sampling as a non-parametric method for estimating the population parameters other than the mean.
Al-Saleh and Samawi (2005) investigated the correlation coefficient estimation using bivariate RSS with an application to the bivariate distribution.
Al-Saleh (2004b) introduced steady state ranked set sampling and investigated its relation to stratified sampling.
Abu-Dayyeh et al. (2004) studied the estimation of the logistic distribution parameters using SRS and RSS with some of its modifications as extreme RSS and median RSS.
Modarres and Zheng (2004) used RSS to investigate maximum likelihood estimation of the dependence parameter of a general bivariate distribution.
Sengupta and Mukhuti (2004) studied the unbiased estimation of an exponential distribution variance using ranked set sampling.
Ozturk (2005) has considered joint estimation of the location and scale parameters of a location-scale family using ranked set sampling method.
Hanandeh and Al-Saleh (2013) considered the inference on Downton's bivariate exponential distribution based on moving extreme ranked set sampling
Hussian. (2014) investigated Bayesian and maximum likelihood estimation for Kumaraswamy distribution based on ranked set sampling.
Chen et al. (2013) used a parametric estimation for the scale parameter for scale distributions using moving extremes RSS.

Tahmasebi and Jafari (2012) studied the scale parameter of Morgenstern type bivariate uniform distribution using ranked set sampling. Omar and Ibrahim (2013) estimated the shape and scale parameters of the Pareto distribution based on extreme RSS.
Sarikavanij, et al. (2014) studied the location and scale estimators of a two-parameter exponential distribution using simple random sample and ranked set sample in terms of generalized variance.
Also, see Kim and Arnold (1999), Zhao and Chen (2002), Bhoj (2000), Badmus et al. (2011), Shaibu and Muttlak (2002), Al-Rawwash et al. (2014), Abu-Dayyeh et al. (2013), El-Neweihi and Sinha (2000), Al-Odat and Omari (2012), Chen et al. (2014), Chacko and Thomas (2007, 2008, 2009), Al-Saleh and Diab (2009), Al-Saleh and Ananbeh (2005, 2007), Bhoj and Ahsanullah (1996), Al-Hadhrami (2010b), Singh and Mehta (2014), Sadek and Alharbi (2014), and Modarres et al. (2006).

4. THE ISSUE OF RANKING AND CONCOMITANT VARIABLE

The efficiency of RSS is affected by many factors. One important aspect of the RSS method is the ranking steps. The ranking can be based by judgment or on concomitant (auxiliary) variable that is correlated to the variable of interest. The efficiency depends on the success in ranking.

Dell and Clutter (1972) was the first one who investigated the performance of RSS when the ranking is done with errors. They used the fundamental identity $f(x)=\frac{1}{m}\sum_{i=1}^{m} f_{[i:m]}(x)$, and showed that $\mu_{[i:m]}$ and $\sigma^2_{[i:m]}$ of the *i*th realized order statistics can be estimated from the recorded data where, $(i:m)$ represent the realized order statistics obtained by the ranking process, and $[i:m]$ denote that the ordering is based on the perceived ranks. Also, the fundamental identity is still true, the relative saving can be written as $RS=\frac{1}{m\sigma^2}\sum_{i=1}^{m}\left(\mu_{[i:m]}-\mu\right)^2$, and $\bar{X}_{RSS}$ remains unbiased estimator of the population mean with errors in ranking. Moreover, if the ranking process is completely done with errors, then $\mathrm{Var}\left(\bar{X}_{SRS}\right)=\mathrm{Var}\left(\bar{X}_{RSS}\right)$

Stokes (1977) studied the ranked set sapling with concomitant variables. She supposed that the study variable *X* has a linear relation with other variable *Y* that is easy to rank, and showed that $RS_{[X:Y]}=\rho^2 RS_Y$, where ρ is the correlation between *Y* and *X*, and

$$RS_{[X:Y]}=\frac{1}{m}\sum_{i=1}^{m}\left(\frac{\mu_{(i:m)}-\mu_x}{\sigma_x}\right)^2,\ \text{and}\ RS_Y=\frac{1}{m}\sum_{i=1}^{m}\left(\frac{E\left(Y_{(i:m)}\right)-\mu_y}{\sigma_y}\right)^2,$$

where $RS_{[X:Y]}$ is the *RS* in estimating μ_x if the ranking is implemented by the concomitant variable *Y*, while RS_Y is the *RS* in estimating μ_Y under perfect ranking on *Y*. She showed that, $RS_Y=RS_X$ if $X\overset{d}{=}Y$ up to a linear transformation since the RSs' are unaffected by linear transformations of the variable of interest. This implies that $RS_{[X:Y]}=\rho^2 RS_X$.

Kaur et al. (1996) made a comparison between RSS and stratified SRS when using a concomitant variable based on equal and optimum allocations of units for estimating the population mean.
Muttlak (1998a) conducted a study of the performance of median RSS (MRSS) for estimating the population mean of interest when the ranking is performed using a concomitant variable. Also, based on an auxiliary variable the regression estimator is proposed to estimate the population mean. According to this study, Muttlak showed that the MRSS estimator is more efficient than RSS and regression estimators.
Ridout and Cobby (1987) investigated the performance of RSS with non-random selection of sets based on the concomitants model of David and Lavine (1972) and Stokes (1977), and also they investigated the performance of RSS with errors in ranking based on cluster sampling.
Muttlak (1988c) studied the RSS with size biased probability of chosen to estimate population parameters as population total based on a concomitant variable, population mean, population size using perfect and imperfect ranking process and the population total using a two-stage sampling method in which the RSS is used at the second stage while at the first stage the sampling is done with probability proportional to the size.

Muttlak and McDonald (1990a, 1999b) have suggested RSS with concomitant variable and size biased probability of chosen proportional to size using both perfect ranking and in the presence of errors in ranking.
Bohn and Wolfe (1992) generalized the Mann-Whitney-Wilcoxon two sample test to the case when a balanced ranked set sample is selected from each of the two populations under perfect ranking. Also, the authors extended the test to the case of imperfect ranking, and study the effects of judgment errors on the features of the test.
Bohn and Wolfe (1994) extended the work of Bohn and Wolfe (1992) to the situations of imperfect ranking and investigate the effects of judgment error on features of Mann-Whitney-Wilcoxon test. The proposed Mann-Whitney-Wilcoxon statistic U_{RSS} is the *U*-statistic, which is computed in as the same way as for RSS under perfect judgment ranking by Bohn and Wolfe (1992) based on the same expression of the mean and variance.
Norris et al. (1995) discussed two alternative methods. The first is a modification of an idea of Takahasi (1970). The second is to use an unbalanced data allocation using Neyman allocation for the characteristic of primary interest, treated this as a concomitant variable for the other variables under consideration.
Kaur and Taillie (2000) appointed the optimal RSS allocations for two classes of symmetric distributions.

A general form for the RSS Fisher information matrix is introduced by Barabesi and El-Sharaawi (2001). If *X* is a continuous random variable with distribution function $F(X;\theta)$, they showed that based on a sample of size *m*, the RSS Fisher information matrix can be seperated into the sum of the SRS Fisher information matrix and a semi-positive definite matrix as

$$\mathrm{I}_{RSS}(\theta) = \mathrm{I}_{SRS}(\theta) + \mathrm{I}_R(\theta),$$

where $\mathrm{I}_{SRS}(\theta)$ is the Fisher information matrix based on a SRS of size *k* which is given by

$$\mathrm{I}_{SRS}(\theta) = -kE\left\{\frac{\partial^2 \ln f(X;\theta)}{\partial\theta\,\partial\theta^T}\right\},$$

and

$$\mathrm{I}_R(\theta) = -k(k-1)E\left\{\frac{\partial \ln F(X;\theta)}{\partial\theta}\frac{\partial \ln F(X;\theta)}{\partial\theta^T}\right\}.$$

In the case of error in ranking, the authors showed that $\mathrm{I}_{\mathrm{RSS}}(\theta) \ge \mathrm{I}_{\mathrm{SRS}}(\theta)$.

Barabesi and Marcheselli (2004) proposed RSS protocol in sampling surveys when an auxiliary variable is available in addition to the target variable. Based on the newly suggested sampling method, the estimators proposed in surveys with auxiliary information such as the regression or the ratio estimators. They showed that the suggested method provides more precise estimation than the simple random sampling counterpart.
Bai and Chen (2003) studied the RSS and its ramifications involving RSS with imperfect ranking, RSS with multivariate samples and RSS by a concomitant variable ranking.
Chen and Shen (2003) suggested two layers RSS with concomitant variable. In the first layer of the procedure, sampling units are ranked with respect to one concomitant variable, and in the second layer, the sampling units are ranked with respect to another concomitant variable. The authors showed that the two-layers RSS method satisfies the fundamental equality $F(x) = \frac{1}{m}\sum_{i=1}^{m} F_{(i:m)}(x)$, which is important for the usual RSS and then more efficient than SRS method. Since the two-layers RSS falls into the framework of the general RSS, they claimed that the all features of the general RSS can be applied for the two-layers RSS without any details of the proofs. The results of the simulations illustrated the superiority of the two-layers RSS over the marginal RSS.

Nahhas et al. (2004) considered a visual judgment error model which is based on ratios of sizes of pairs of observations; that is, $Y_i = X_i\varepsilon_i$, $i = 1, 2, ..., m$, where $\varepsilon_1, \varepsilon_2, ..., \varepsilon_m$ are iid lognormal with parameters 0 and β^2, where the ε's and *X*'s are mutually independent. They derived the efficiency with balanced ranked set sampling under this multiplicative model and considered methods for estimating the variance of the error.

Chen et al. (2006a) investigated the empirical evaluation of the accuracy in RSS rankings. They showed that, the ranking accuracy can be assessed through an $m \times m$ matrix, $\mathrm{P} = \left(p_{ij}\right)$, where p_{ij} is the probability that the unit with real rank *i* in a given set with size *m* is considered to be the *j*th judgment order statistics (Bohn & Wolfe 1994; Stark & Wolfe 2002). In particular, $p_{ij} = P\left(X_{(i:m)} = X_{[j:m]}\right)$. Now, the value of *p* depends on the process of ranking, i.e., under perfect ranking, $p_{ii} = 1$ for $i = 1, 2, ..., m$ and $p_{ij} = 0$ for $i \neq j = 1, ..., m$, while $p_{ij} = \dfrac{1}{m}$ for any *i* and *j* if the ranking process is completely random. Since the *j*th judgment order statistics will equal to the *i*th true order statistics for some *i*, then $\sum_{i=1}^{m} p_{ij} = 1$. They illustrated that under any ranking issue the ranking errors increase progressively when the set size increases, which indeed, has a negative effect of the efficiency of a RSS estimator. On the other hand, the efficiency of RSS estimator increases with larger sample size under perfect ranking. Hatefi and Jozani (2013) investigated a Fisher information in different types of imperfect and perfect ranked set samples selected from finite mixture models. Chen (2002) considered multiple concomitant variables in adaptive ranked set sampling. Also, see Mode et al. (2002), Park and Lim (2012), Badmus and Ikegwu (2013), Scaria and Nair (1999), Presnell and Bohn (1999), Nahhas (2004), Wang et al. (2006), Vock and Balakrishnan (2011), Li and Balakrishnan (2008), Frey et al. (2007), Bouza (2001), Ozturk (2000c, 2007b, 2009, 2010), Zheng (2004), and Patil et al. (1994).

5. STATISTICAL QUALITY CONTROL CHARTS BASED ON RSS

In this section, discussions are given on the usual RSS or any of its modifications may considered to improve control charts for the mean. The RSS charts are compared to the commonly quality control charts for variables using the usual SRS. Let X_{ij}, $i = 1, 2, ..., m, j = 1, 2, ..., n$ be n independent samples each of size *m* from the normal distribution, $N\left(\mu, \sigma^2\right)$. It is well known that the distribution of the sample mean $\bar{X}_j$ is $N\left(\mu, \sigma^2 / m\right)$. Shewhart's $\bar{X}$ control charts have been considered widely in industries to detect the shift in process mean. Salazar and Sinha (1997) are fist who used the RSS in quality control charts for variables. If both μ and σ^2 are known, then the upper control limit (UCL), center limit (CL) and lower control limit (LCL) for the $\bar{X}$ charts as suggested by Salazar and Sinha (1997) are given by

$$\mathrm{UCL_{RSS}} = \mu + 3\sigma_{\bar{X}_{\mathrm{RSS}}}, \ \mathrm{CL_{RSS}} = \mu, \ \mathrm{LCL_{RSS}} = \mu - 3\sigma_{\bar{X}_{\mathrm{RSS}}},$$

where

$$\sigma_{\bar{X}_{\mathrm{RSS}}} = \sqrt{\frac{\sigma^2}{m} - \frac{1}{m^2} \sum_{i=1}^{m} \left(\mu_{i(i:m)} - \mu\right)^2} .$$

The RSS control charts are compared with the SRS control charts in terms of ARL, which is the average number of points that must be plotted before a point indicates an out-of-control condition. The ARL is defined as $\mathrm{ARL} = 1/\alpha$ for any Shewhart control chart if the process is in control and the process observations are uncorrelated, and $\mathrm{ARL} = 1/(1-\beta)$ if the process is get out of control where α and β denote the probabilities of Type-I and Type-II errors, respectively.

Based on the ARL, the process is in control with mean μ_o and standard deviation σ_o, otherwise, the process has a shift about $\delta\sigma_o / \sqrt{m}$, i.e., a shift in mean from μ_o to $\mu_o + \delta\sigma_o / \sqrt{m}$, where δ is nonnegative and selected to dominate on the shift in the mean μ. When the process is in control and follow a normal distribution with parameters μ_o and σ_o, the shift in the process mean is known as $\delta = \sqrt{m}\left|\mu_1 - \mu_o\right| / \sigma_o$, see Montgomery (2005). Table 5.1 shows some ARL values using RSS and SRS in quality control charts.

As shown in Table 5.1 The ARL values based on RSS, will decrease much faster as compared to the values under SRS. As the sample size increases, generally, the RSS ARL values decreases.

Table 5.1: ARL values based on 3-sigma limits using SRS and RSS.

	$m=3$		$m=4$		$m=5$	
δ	**SRS**	**RSS**	**SRS**	**RSS**	**SRS**	**RSS**
0.0	370.033	343.461	368.365	344.649	370.068	345.422
0.1	347.209	316.655	357.001	308.142	354.103	307.063
0.2	312.227	247.215	312.711	241.397	312.572	222.178
0.3	253.054	185.083	255.764	166.877	252.103	151.282
0.4	202.840	132.668	198.966	115.297	200.197	98.075
0.8	72.247	34.190	71.595	26.380	71.599	21.046
1.2	27.530	11.001	27.953	8.180	27.596	6.274
1.6	12.506	4.672	12.598	3.462	12.039	2.715
2.0	6.294	2.466	6.353	1.911	6.308	1.583
2.4	3.666	1.585	3.623	1.330	3.639	1.198
3.2	1.738	1.080	1.741	1.029	1.725	1.011

Muttlak and Al-Sabah (2003a) developed many control charts for the sample mean by using pair RSS and selected RSS methods. Muttlak and Al-Sabah (2003b) used the RSS and some of its modifications, namely, median RSS and extreme RSS methods to develop different quality control charts of the mean. Abujiya and Muttlak (2004) investigated quality control charts of the sample mean using the median RSS and double RSS methods. For more about RSS in quality control charts see Al-Omari and Haq (2012), Al-Nasser and Al-Rawwash (2007), Al-Sabah (2010), Abbasi and Miller (2012), Al-Omari and Al-Nasser (2011), Haq (2012), Pongpullponsak and Sontisamran (2013), Mehmood et al. (2013, 2014), Riaz et al. (2011), Haq et al. (2014b), Lee and Riaz (2014), and Abujiya et al. (2012).

6. RSS WITH FINITE POPULATIONS

Several researchers have considered different problems belonging to the ranked set sampling scheme when the sampling is proceed on an infinite population. Recently, some authors proposed RSS procedure when the population is finite.

Kowalczyk (2004) studied the efficiency of RSS in estimating the finite population mean and total. The author showed that the new aspect of RSS includes RSS unit ranking or semi-ranking done based on grouped data. If Y_i and y_i denote the ith population element and the corresponding value of M units $(i=1,2,...,M)$, so that each element in the population has probability $1/M$. The RSS finite population mean is $\bar{y}_{RSS}=\frac{1}{m}\sum_{i=1}^{m}y_{i(i:m)}$, with variance

$$D^2\left(\bar{y}_{RSS}\right)=\frac{M-1}{mM}S^2-\frac{1}{m^2}\sum_{i=1}^{m}\left[E\left(y_{(i:m)}\right)-\bar{Y}\right]^2,$$

where $S^2=\frac{1}{M-1}\sum_{j=1}^{M}\left(Y_j-\bar{Y}\right)^2$. While the unbiased estimator of the population total is $y_{RSS}=M\,\bar{y}_{RSS}$, with variance

$$D^2\left(y_{RSS}\right)=\frac{M^2-M}{m}S^2-\left(\frac{M}{m}\right)^2\sum_{i=1}^{m}\left[E\left(y_{(i:m)}\right)-\bar{Y}\right]^2.$$

The author concluded that $\bar{y}_{RSS}$ and $y_{RSS}=M\,\bar{y}_{RSS}$ are unbiased estimators of the population mean and total, respectively and they are more efficient than their counterparts using SRS with replacement. Under SRS without replacement, the author showed that when the sample size is comparatively small as compared to the population size, then $D^2\left(\bar{y}_{SRS}\right)\approx D^2\left(\bar{y}_{SRSWOR}\right)$ and the RSS should also be superior to SRS.

Deshpande et al. (2006) suggested three protocols for drawing a RSS from a finite population. For small population size, the sampling methods lead to highly divergent sampling distribution, while when the population size increases the differences decreases. Nonparametric confidence intervals are found to be shorter than those produced by SRS based on the three sampling protocols. Gokpinar and Ozdmir (2014) considered inclusion probabilities in ranked set sat sampling in finite populations.

7. CONFIDENCE INTERVALS AND HYPOTHESIS TESTING IN RSS

Hartlaub and Wolfe (1999) extended the concept of ranked set samples to the *m*-sample location setting when the treatment effect parameters follow a restricted umbrella pattern. They developed the distribution-free testing for both the case where the peak of the umbrella is known and for the case where it is unknown.
Bohn (1994) considered technique based on ranked-set samples, analogues of the standard one-sample sign statistics.
Hettmansperger (1995) proposed procedures based on RSS, analogues of the standard one-sample signed-rank test.
Kaur et al. (1996) examined the performance of the RSS sign test under unequal allocations and found that the optimal allocation for testing the median assigns all observations to the middle rank order regardless of the parent distribution. Kaur et al. (1997) considered the unequal allocations using RSS for skew distributions.
Ozturk and Wolfe (2000) suggested alternative ranked set sampling methods for the sign test statistic.
Al-Saleh and Zheng (2003) suggested a method based on ranked set sampling to get an approximate preferable sample from a population of interest. This method can be considered to get sample points that goes to be from a specific portion of a distribution such as the upper or lower quartiles.
Kim and Kim (2003) suggested ranked ordering-set sampling (ROSS) and compared it to the usual RSS. They proposed the test statistic using sign test on ROSS and found that the ROSS has more information than RSS, where the asymptotic efficiency of ROSS relative to RSS is always greater than 1 except when sample size is 2.
Ozturk et al. (2004) developed a nonparametric multi-sample inference for simple-tree samples alternatives is distribution-free. They have constructed the multi-sample inference based on compare the simultaneous one-sample sign confidence intervals for the medians.
Hui et al. (2005) investigated the resampling methods to obtain confidence intervals for the regression estimator of the population mean.
Ozturk and Deshpande (2006) studied the properties of the quantile intervals using RSS and improved distributional properties of the RSS order statistics. The interpolation of adjacent order statistics suggested by Hettmansperger and Sheather (1996) was considered by the authors to have confidence interval by extending the interpolated confidence intervals to the RSS data for small sample sizes. The authors concluded that the quantile intervals using RSS have shorter expected lengths or higher coverage probabilities than their simple random sample competitors.

Terpstra and Miller (2006) studied the hypothesis testing and confidence interval for a population proportion based on a RSS. For a RSS data, let $X_{(i:m)h}$ where $i=1,2,...,m$ and $h=1,2,...,n$, consists of a $m\times n$ matrix of independent binary order statistics. The distribution of $X_{(i:m)h}$ is assumed to be $B\left(1,\pi_{mi}(p)\right)$, where *p* denotes the success probability and

$$\pi_{mi}(p)=P\left(B(m,p)\geq m-i+1\right),\ i=1,2,...,m.$$

The RSS estimator of the population proportion is given by

$$\hat{p}_{RSS}=\frac{X_{(\cdot).}}{nm}=\frac{1}{nm}\sum_{i=1}^{m}X_{(i:m).},$$

where $X_{(\cdot)}$ is the number of success which is a $B\left(n,\pi_{mi}(p)\right)$ random variable. They showed that, the RSS inference procedures are generally more efficient than their competitors using SRS, referring to the shorter confidence intervals and more powerful test.
Ozturk (2007) improved an exact nonparametric test for the equality of medians in a two-sample or design. Based on general ranking method, he showed that the asymptotic null distribution is normal with imperfect ranking.
Albatineh et al. (2014) studied a confidence interval of the signal-to-noise ratio using RSS. Zhang et al. (2014) considered sign tests using ranked set sampling with unequal set sizes. Gaur et al. (2013) investigated a nonparametric test for a multi-sample scale problem using RSS data. Frey and Wang (2013) considered most powerful rank tests for perfect rankings. Omar et al. (2012) studied cconfidence interval estimation of the shape parameter of Pareto distribution using extreme order statistics. Also, see Samawi et al. (2006), Vock and Balakrishnan (2013), Ozturk and Wolfe (2001, 2000a, 2000b, 2000d), Kim et al. (1999), Samawi and Abu-Dayyeh (2003), Liangyong and Xiaofang (2010), Koti and Babu (1996), Al-Rawwash et al. (2010), Ozturk (1999), Barabesi (1998), Ozturk and Balakrishnan (2009),

Balakrishnan and Li (2006), Dong and Cui (2010), Ozturk (2001), Wang and Zhu (2005), Kim and Kim (1998), and Ozturk (2007a).

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