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ESTIMATION OF FINITE POPULATION MEAN USING KNOWN CORRELATION COEFFICIENT BETWEEN AUXILIARY CHARACTERS

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1. INTRODUCTION

Let $U=\{U_1,U_2,...,U_N\}$ be a finite population of N units. Suppose two auxiliary variables X_1 and X_2 are observed on $U_i(i=1,2,...,N)$, where X_1 is positively and X_2 is negatively correlated with the study variable *Y*. A simple random sample without replacement (SRSWOR) of size *n* with *n* < *N*, is drawn from the population U to estimate $\bar{Y}=\sum_{i=1}^{N} y_i/N$, the population mean of Y, when the population means $\bar{X}_1=\sum_{i=1}^{N} x_{1i}/N$ and $\bar{X}_2=\sum_{i=1}^{N} x_{2i}/N$ of X_1 and X_2 are respectively, known. For estimating $\bar{Y}$, Singh (1967) suggested a ratio-cum-product estimator

$$\hat{\bar{Y}}_1=\bar{y}\left(\frac{\bar{X}_1}{\bar{x}_1}\right)\left(\frac{\bar{x}_2}{\bar{X}_2}\right) \tag{1}$$

where $\bar{y}=\sum_{i=1}^{n} y_i/n$, $\bar{x}_1=\sum_{i=1}^{n} x_{1i}/n$ and $\bar{x}_2=\sum_{i=1}^{n} x_{2i}/n$.

Wide applicability of the estimator $\hat{\bar{Y}}_1$ has led many authors to suggest unbiased versions of $\hat{\bar{Y}}_1$ with their properties, for instance, see Sahoo and Swain (1980), Biradar and Singh (1992-93) and Tracy *et al.* (1998). Sahai and Sahai (1985) and Singh (1987 b) have mentioned that the past association with experimental material might provide a close guess for the correlation coefficient ρ_{yx_1} between study variate Y and auxiliary character X_1 i.e. ρ_{yx_1} can be guessed quite accurately. Recently, Singh and Tailor (2003) have utilized the information on ρ_{yx_1} and suggested a modified ratio estimator for $\bar{Y}$ with its properties. Further Singh and Singh (1984) advocated that the correlation coefficient $\rho_{x_1x_2}$ between auxiliary variates X_1 and X_2 may be known in many practical situations and hence utilizing the known value of $\rho_{x_1x_2}$ suggested a class of estimators for population variance σ_y^2 of Y with its properties. This led authors to suggest modified ratio-cum-product estimator using $\rho_{x_1x_2}$ with its properties.

A jackknife version of the suggested estimator $\hat{\bar{Y}}_2$ is also given and its properties are studied. An empirical study is carried out in support of the proposed estimator.

2. SUGGESTED RATIO-CUM-PRODUCT ESTIMATOR

Assuming that the correlation coefficient $\rho_{x_1x_2}$ between auxiliary characters X_1 and X_2 is known, we define a ratio-cum-product estimator of $\bar{Y}$ as

$$\hat{\bar{Y}}_2 = \bar{y}\left(\frac{\bar{X}_1 + \rho_{x_1x_2}}{\bar{x}_1 + \rho_{x_1x_2}}\right)\left(\frac{\bar{x}_2 + \rho_{x_1x_2}}{\bar{X}_2 + \rho_{x_1x_2}}\right). \tag{2}$$

To the first degree of approximation, the bias and mean square error (MSE) of the proposed estimator $\hat{\bar{Y}}_2$ are respectively, given by

$$B(\hat{\bar{Y}}_2) = \theta \bar{Y}[\mu_1^* C_{x_1}^2 (\mu_1^* - K_{yx_1}) + \mu_2^* C_{x_2}^2 (K_{yx_2} - \mu_1^* K_{x_1x_2})] \tag{3}$$

and

$$MSE(\hat{\bar{Y}}_2) = \theta \bar{Y}^2[C_y^2 + \mu_1^* C_{x_1}^2 (\mu_1^* - 2K_{yx_1}) + \mu_2^* C_{x_2}^2 \{\mu_2^* + 2(K_{yx_2} - \mu_1^* K_{x_1x_2})\}], \tag{4}$$

where

$$K_{yx_1} = \rho_{yx_1}(C_y / C_{x_1}),\ K_{yx_2} = \rho_{yx_2}(C_y / C_{x_2}),\ K_{x_1x_2} = \rho_{x_1x_2}(C_{x_1} / C_{x_2}),$$

$$\mu_i^* = \bar{X}_i / (\bar{X}_i + \rho_{x_1x_2}),\ i = (1,2);\theta = \left(\frac{1}{n} - \frac{1}{N}\right),\ C_y = S_y / \bar{Y},$$

$$C_{x_i} = S_{x_i} / \bar{X}_i, (i = 1,2);\ \rho_{y\,x_i} = S_{yx_i} / (S_y S_{x_i}), (i = 1,2);$$

$$S_y^2 = \sum_{j=1}^{N} (y_i - \bar{Y})^2 / (N-1),\ S_{x_i}^2 = \sum_{j=1}^{N} (x_{ij} - \bar{X}_i)^2 / (N-1), (i = 1,2)$$

and

$$S_{yx_i} = \sum_{j=1}^{N} (y_i - \bar{Y})(x_{ij} - \bar{X}_i)/(N-1), (i = 1, 2).$$

When no auxiliary information is used the estimator $\hat{\bar{Y}}_2$ reduces to the conventional unbiased estimator $\bar{y}$. If the information only on auxiliary variate X_1 is used, then the estimator $\hat{\bar{Y}}_2$ tends to the usual ratio estimator $\bar{y}_R = \bar{y}(\bar{X}_1/\bar{x}_1)$. On the other hand if the information is available on auxiliary variate X_2 only, $\hat{\bar{Y}}_2$ reduces to the usual product estimator $\bar{y}_P = \bar{y}(\bar{x}_2/\bar{X}_2)$.

It is well known that sample mean $\bar{y}$ is an unbiased estimator of $\bar{Y}$ and its variance under SRSWOR sampling scheme is given by

$$V(\bar{y}) = \theta \bar{Y}^2 C_y^2. \tag{5}$$

To the first degree of approximation, the biases and MSEs of $\bar{y}_R$, $\bar{y}_P$ and $\hat{\bar{Y}}_1$ are respectively given by

$$B(\bar{y}_R) = \theta \bar{Y} C_{x_1}^2 (1 - K_{yx_1}), \tag{6}$$

$$B(\bar{y}_P) = \theta \bar{Y} C_{x_2}^2 K_{yx_2}, \tag{7}$$

$$B(\hat{\bar{Y}}_1) = \theta \bar{Y} [C_{x_1}^2 (1 - K_{yx_1}) + C_{x_2}^2 (K_{yx_2} - K_{x_1x_2})], \tag{8}$$

$$MSE(\bar{y}_R) = \theta \bar{Y}^2 [C_y^2 + C_{x_1}^2 (1 - 2K_{yx_1})], \tag{9}$$

$$MSE(\bar{y}_P) = \theta \bar{Y}^2 [C_y^2 + C_{x_2}^2 (1 + 2K_{yx_2})], \tag{10}$$

and

$$MSE(\hat{\bar{Y}}_1) = \theta \bar{Y}^2 [C_y^2 + C_{x_1}^2 (1 - 2K_{yx_1}) + C_{x_2}^2 \{1 + 2(K_{yx_2} - K_{x_1x_2})\}]. \tag{11}$$

3. EFFICIENCY COMPARISIONS

It follows from (4), (5), (9), (10) and (11) that

(i) $MSE(\bar{y}_R) < V(\bar{y})$ if

$$K_{yx_1} > \frac{1}{2} \tag{12}$$

(ii) $MSE(\bar{y}_P) < V(\bar{y})$ if

$$K_{yx_2} < -\frac{1}{2} \tag{13}$$

(iii) $MSE(\hat{\bar{Y}}_1) < V(\bar{y})$ if

$$[C_{x_1}^2(1-2K_{yx_1}) + C_{x_2}^2\{1+2(K_{yx_2} - K_{x_1x_2})\}] < 0$$

which is always true if

$$K_{yx_1} > \frac{1}{2} \quad \text{and} \quad K_{yx_2} < \left(K_{x_1x_2} - \frac{1}{2}\right) \tag{14}$$

(iv) $MSE(\hat{\bar{Y}}_2) < V(\bar{y})$ if

$$[C_{x_1}^2\mu_1^*(\mu_1^* - 2K_{yx_1}) + C_{x_2}^2\mu_2^*\{\mu_2^* + 2(K_{yx_2} - \mu_1^* K_{x_1x_2})\}] < 0$$

which always holds if

$$K_{yx_1} > \frac{\mu_1^*}{2} \quad \text{and} \quad K_{yx_2} < \left(\mu_1^* K_{x_1x_2} - \frac{\mu_2^*}{2}\right) \tag{15}$$

(v) $MSE(\hat{\bar{Y}}_1) < MSE(\bar{y}_R)$ if

$$K_{yx_2} < K_{x_1x_2} - \frac{1}{2} \tag{16}$$

(vi) $MSE(\hat{\bar{Y}}_1) < MSE(\bar{y}_P)$ if

$$K_{yx_1} > -K_{x_2x_1} + \frac{1}{2}, \tag{17}$$

where $K_{x_2x_1} = \rho_{x_1x_2}(C_{x_2}/C_{x_1})$.

(vii) $MSE(\hat{\bar{Y}}_2) < MSE(\bar{y}_R)$ if

$$[(1-\mu_1^*)\{2K_{yx_1} - (1+\mu_1^*)\}C_{x_1}^2 + \mu_2^*\{\mu_2^* + 2(K_{yx_2} - \mu_1^* K_{x_1x_2})\}C_{x_2}^2] < 0$$

which is always true if

$$K_{yx_1} < \frac{(1+\mu_1^*)}{2} \quad \text{and} \quad K_{yx_2} < \left(\mu_1^* K_{x_1x_2} - \frac{\mu_2^*}{2}\right) \tag{18}$$

(viii) $MSE(\hat{\bar{Y}}_2) < MSE(\bar{y}_P)$ if

$$[\mu_1^*\{\mu_1^* - 2(K_{yx_1} + \mu_2^* K_{x_2x_1})\}C_{x_1}^2 - (1-\mu_2^*)\{(1+\mu_2^*) + 2K_{yx_2}\}C_{x_2}^2] < 0$$

which always holds if

$$K_{yx_1} > -\mu_2^* K_{x_2x_1} + \frac{\mu_1^*}{2} \quad \text{and} \quad K_{yx_2} > -\frac{(1+\mu_2^*)}{2} \qquad (19)$$

and

(ix) $MSE(\hat{\bar{Y}}_2) < MSE(\hat{\bar{Y}}_1)$ if

$$[C_{x_1}^2(1-\mu_1^*)\{2K_{yx_1} - (1+\mu_1^*)\} + C_{x_2}^2\{2K_{x_1x_2}(1-\mu_1^*\mu_2^*) - -(1-\mu_2^*)(1+\mu_2^* + 2K_{yx_2}\}] < 0$$

which is always true if

$$K_{yx_1} < \frac{(1+\mu_1^*)}{2} \quad \text{and} \quad K_{yx_2} > \left[\frac{K_{x_1x_2}(1-\mu_1^*\mu_2^*)}{(1-\mu_2^*)} - \frac{(1+\mu_2^*)}{2}\right]. \qquad (20)$$

Now combining (12), (16) and (20) we get that the proposed estimator $\hat{\bar{Y}}_2$ is more efficient than $\bar{y}$, $\bar{y}_R$ and Singh's (1967) estimator $\hat{\bar{Y}}_1$ i.e. $MSE(\hat{\bar{Y}}_2) < MSE(\hat{\bar{Y}}_1) < MSE(\bar{y}_R) < V(\bar{y})$ if

$$\frac{1}{2} < K_{yx_1} < \frac{(1+\mu_1^*)}{2} \quad \text{and} \quad \left[\frac{K_{x_1x_2}(1-\mu_1^*\mu_2^*)}{(1-\mu_2^*)} - \frac{(1+\mu_2^*)}{2}\right] < K_{yx_2} < \left(K_{x_1x_2} - \frac{1}{2}\right). \qquad (21)$$

Further combining (20), (17) and (13) we obtained that the suggested estimator $\hat{\bar{Y}}_2$ is more efficient than $\bar{y}$, $\bar{y}_P$ and Singh's (1967) estimator $\hat{\bar{Y}}_1$

i.e. $MSE(\hat{\bar{Y}}_2) < MSE(\hat{\bar{Y}}_1) < MSE(\bar{y}_P) < V(\bar{y})$ if

$$\left(K_{x_2x_1} + \frac{1}{2}\right) < K_{yx_1} < \frac{(1+\mu_1^*)}{2} \quad \text{and} \quad \left[\frac{K_{x_1x_2}(1-\mu_1^*\mu_2^*)}{(1-\mu_2^*)} - \frac{(1+\mu_2^*)}{2}\right] < K_{yx_2} < -\frac{1}{2}. \qquad (22)$$

It is to be noted that the suggested estimator $\hat{\bar{Y}}_2$ is biased. In some applications, bias is a major disadvantage. Keeping this in view, we have discussed the unbiasedness of the proposed estimator $\hat{\bar{Y}}_2$, and using the technique suggested by Quenouille (1956) known as 'Jack-knife' technique, proposed a family of almost unbiased estimators with its properties.

4. FAMILY OF UNBIASED ESTIMATORS OF POPULATION MEAN $\overline{Y}$ USING JACKKNIFE TECHNIQUE

Let a simple random sample of size *n* = *gm* drawn without replacement and split at random into *g* sub-samples, each of size *m*. Then we define the Jack-knife ratio-cum-product estimator for population mean $\overline{Y}$ as

$$\hat{\overline{Y}}_{2J} = \frac{1}{g}\sum_{j=1}^{g}\overline{y}_j'\left(\frac{\overline{X}_1+\rho_{x_1x_2}}{\overline{x}_{1j}'+\rho_{x_1x_2}}\right)\left(\frac{\overline{x}_{2j}'+\rho_{x_1x_2}}{\overline{X}_2+\rho_{x_1x_2}}\right) \tag{23}$$

where $\overline{y}_j' = (n\overline{y} - m\overline{y}_j)/(n-m)$ and $\overline{x}_{ij}' = (n\overline{x}_i - m\overline{x}_{ij})/(n-m)$, $i = 1,2$; are the sample means based on a sample of (*n-m*) units obtained by omitting the j^{th} group and $\overline{y}_j$ and $\overline{x}_{ij}$ $(i = 1,2; \quad j = 1,2,...,g)$ are the sample means based on the j^{th} sub samples of size $m = n/g$.

The bias of $\hat{\overline{Y}}_{2J}$, to terms of order n^{-1}, can be easily obtained as

$$B(\hat{\overline{Y}}_{2J}) = \frac{(N-n+m)}{N(n-m)}\overline{Y}[\mu_1^* C_{x_1}^2(\mu_1^* - K_{yx_1}) + \mu_2^* C_{x_2}^2(K_{yx_2} - \mu_1^* K_{x_1x_2})]. \tag{24}$$

From (3) and (24) we have

$$\frac{B(\hat{\overline{Y}}_2)}{B(\hat{\overline{Y}}_{2J})} = \frac{(N-n)(n-m)}{n(N-n+m)} \tag{25}$$

or $\quad B(\hat{\overline{Y}}_2) = \frac{(N-n)(n-m)}{n(N-n+m)}B(\hat{\overline{Y}}_{2J})$

or $\quad B(\hat{\overline{Y}}_2) - \frac{(N-n)(n-m)}{n(N-n+m)}B(\hat{\overline{Y}}_{2J}) = 0$

or $\quad \lambda^* B(\hat{\overline{Y}}_2) - \delta^* \lambda^* B(\hat{\overline{Y}}_{2J}) = 0 \qquad (26)$

for any scalar λ^*, we have

$$\delta^* = \frac{(N-n)(n-m)}{n(N-n+m)}. \tag{27}$$

From (26), we have

$$\lambda^* E(\hat{\bar{Y}}_2 - \bar{Y}) - \delta^* \lambda^* E(\hat{\bar{Y}}_{2J} - \bar{Y}) = 0$$

or $$\lambda^* E(\hat{\bar{Y}}_2 - \bar{y}) - \delta^* \lambda^* E(\hat{\bar{Y}}_{2J} - \bar{y}) = 0$$

or $$E[\lambda^* \hat{\bar{Y}}_2 - \lambda^* \delta^* \hat{\bar{Y}}_{2J} - \bar{y}\{\lambda^*(1-\delta^*) - 1\}] = \bar{Y} .$$

Hence, the general family of almost unbiased ratio-cum-product estimators of $\bar{Y}$ as

$$\hat{\bar{Y}}_{2u} = [\bar{y}\{1-\lambda^*(1-\delta^*)\} + \lambda^* \hat{\bar{Y}}_2 - \lambda^* \delta^* \hat{\bar{Y}}_{2J}] \tag{28}$$

see Singh (1987 a).

Remark 4.1. For $\lambda^* = 0$, $\hat{\bar{Y}}_{2u}$ yields the usual unbiased estimator $\bar{y}$ while $\lambda^* = (1-\delta^*)^{-1}$, gives an almost unbiased estimator for $\bar{Y}$ as

$$\begin{aligned}\hat{\bar{Y}}^*_{2u} = {} & \frac{(N-n+m)}{N} g\,\bar{y}\left(\frac{\bar{X}_1 + \rho_{x_1x_2}}{\bar{x}_1 + \rho_{x_1x_2}}\right)\left(\frac{\bar{x}_2 + \rho_{x_1x_2}}{\bar{X}_2 + \rho_{x_1x_2}}\right) \\ & - \frac{(N-n)(g-1)}{Ng} \sum_{j=1}^{g} \bar{y}'_j \left(\frac{\bar{X}_1 + \rho_{x_1x_2}}{\bar{x}'_{1j} + \rho_{x_1x_2}}\right)\left(\frac{\bar{x}'_{2j} + \rho_{x_1x_2}}{\bar{X}_2 + \rho_{x_1x_2}}\right)\end{aligned} \tag{29}$$

which is Jack-knifed version of the proposed estimator $\hat{\bar{Y}}_2$.

Many other almost unbiased estimator from (28) can be generated by putting suitable values of λ^*.

5. SEARCH OF AN OPTIMUM ESTIMATOR IN FAMILY $\hat{\bar{Y}}_{2u}$ AT (28)

The family of almost unbiased estimator $\hat{\bar{Y}}_{2u}$ at (28) can be expressed as

$$\hat{\bar{Y}}_{2u} = \bar{y} - \lambda^* \bar{y}_1 , \tag{30}$$

where $\bar{y}_1 = [(1-\delta^*)\bar{y} - \bar{y}_2]$ and $\bar{y}_2 = \hat{\bar{Y}}_2 - \delta^* \hat{\bar{Y}}_{2J}$. The variance of $\hat{\bar{Y}}_{2u}$ is given by

$$V(\hat{\bar{Y}}_{2u}) = V(\bar{y}) + \lambda^{*2} V(\bar{y}_1) - 2\lambda^* Cov(\bar{y}, \bar{y}_1) \tag{31}$$

which is minimized for

$$\lambda^* = Cov(\bar{y}, \bar{y}_1)/V(\bar{y}_1)\,. \tag{32}$$

Substitution of (32) in (31) yields minimum variance of $\hat{\bar{Y}}_{2u}$ as

$$\min.V(\hat{\bar{Y}}_{2u}) = V(\bar{y}) - \frac{\{Cov(\bar{y}, \bar{y}_1)\}^2}{V(\bar{y}_1)}$$

$$= V(\bar{y})(1-\rho_{01}^2)\,, \tag{33}$$

where ρ_{01} is the correlation coefficient between $\bar{y}$ and $\bar{y}_1$.
From (33) it is immediate that

$$\min.V(\hat{\bar{Y}}_{2u}) < V(\bar{y})\,.$$

To obtain the explicit expression of the variance of $\hat{\bar{Y}}_{2u}$, we write the following results to terms of order n^{-1}, as

$$MSE(\hat{\bar{Y}}_{2J}) = Cov(\hat{\bar{Y}}_2, \hat{\bar{Y}}_{2J}) = MSE(\hat{\bar{Y}}_2) \tag{34}$$

and

$$Cov(\bar{y}, \hat{\bar{Y}}_2) = Cov(\bar{y}, \hat{\bar{Y}}_{2J}) = \theta \bar{Y}^2 [C_y^2 - \mu_1^* \rho_{yx_1} C_y C_{x_1} + \mu_2^* \rho_{yx_2} C_y C_{x_2}] \tag{35}$$

where $MSE(\hat{\bar{Y}}_2)$ is given by (4).

Now using the results from (4), (5) and (35) into (31) we get the variance of $\hat{\bar{Y}}_{2u}$ to the terms of order n^{-1} as

$$V(\hat{\bar{Y}}_{2u}) = \theta \bar{Y}^2 [C_y^2 + \lambda^{*2}(1-\delta^*)^2(\mu_1^{*2} C_{x_1}^2 + \mu_2^{*2} C_{x_2}^2 - 2\rho_{x_1x_2} C_{x_1} C_{x_2} \mu_1^* \mu_2^*)$$

$$-2\lambda^*(1-\delta^*)(\mu_1^* \rho_{yx_1} C_y C_{x_1} - \mu_2^* \rho_{yx_2} C_y C_{x_2})] \tag{36}$$

which is minimized for

$$\lambda^* = \frac{(\mu_1^* \rho_{yx_1} C_y C_{x_1} - \mu_2^* \rho_{yx_2} C_y C_{x_2})}{(1-\delta^*)(\mu_1^{*2} C_{x_1}^2 + \mu_2^{*2} C_{x_2}^2 - 2\mu_1^* \mu_2^* \rho_{x_1x_2} C_{x_1} C_{x_2})} = \lambda_{opt}^*\,. \tag{37}$$

Substitution of λ_{opt}^* in $\hat{\bar{Y}}_{2u}$ yields the optimum estimator $\hat{\bar{Y}}_{2u(opt)}$ (say). Thus the resulting minimum variance of $\hat{\bar{Y}}_{2u}$ is given by

$$\min.V(\hat{\bar{Y}}_{2u})=\theta\bar{Y}^2C_y^2\left[1-\frac{(\mu_1^*\rho_{yx_1}C_{x_1}-\mu_2^*\rho_{yx_2}C_{x_2})^2}{(\mu_1^{*2}C_{x_1}^2+\mu_2^{*2}C_{x_2}^2-2\mu_1^*\mu_2^*\rho_{x_1x_2}C_{x_1}C_{x_2})}\right]=V(\hat{\bar{Y}}_{2u(opt)})\,. \tag{38}$$

From (4), (11) and (38) we have

$$V(\bar{y})-\min.V(\hat{\bar{Y}}_{2u})=\theta\bar{Y}^2C_y^2\left[\frac{(\mu_1^*\rho_{yx_1}C_{x_1}-\mu_2^*\rho_{yx_2}C_{x_2})^2}{(\mu_1^{*2}C_{x_1}^2+\mu_2^{*2}C_{x_2}^2-2\mu_1^*\mu_2^*\rho_{x_1x_2}C_{x_1}C_{x_2})}\right]\geq 0 \tag{39}$$

and

$$MSE(\hat{\bar{Y}}_2)-\min.V(\hat{\bar{Y}}_{2u})=$$
$$=\theta\bar{Y}^2\left[\frac{(\mu_1^{*2}C_{x_1}^2+\mu_2^{*2}C_{x_2}^2-2\mu_1^*\mu_2^*\rho_{x_1x_2}C_{x_1}C_{x_2}-\rho_{yx_1}C_yC_{x_1}\mu_1^*+\rho_{yx_2}C_yC_{x_2}\mu_2^*)^2}{(\mu_1^{*2}C_{x_1}^2+\mu_2^{*2}C_{x_2}^2-2\mu_1^*\mu_2^*\rho_{x_1x_2}C_{x_1}C_{x_2})}\right]\geq 0. \tag{40}$$

Thus from (39) and (40) we have the following inequalities:

$$\min.V(\hat{\bar{Y}}_{2u})\leq V(\bar{y}) \tag{41}$$

and

$$\min.V(\hat{\bar{Y}}_{2u})\leq MSE(\hat{\bar{Y}}_2) \tag{42}$$

which follows that $\hat{\bar{Y}}_{2u}$ with $\lambda^*=\lambda_{opt}^*$ is more efficient than $\bar{y}$ and $\hat{\bar{Y}}_2$.

When λ^* does not coincide with λ_{opt}^* then from (5) and (36) we note that $V(\hat{\bar{Y}}_{2u})\leq V(\bar{y})$ if

$$\left.\begin{array}{ll} \textit{either} & 0<\lambda^*<2\lambda_{opt}^* \\ \textit{or} & 2\lambda_{opt}^*<\lambda^*<0 \end{array}\right\} \tag{43}$$

It is observed from (11) and (36) that $MSE(\hat{\bar{Y}}_{2u})<MSE(\hat{\bar{Y}}_1)$ if

$$\frac{B-\sqrt{(B^2-AC)}}{(1-\delta^*)A}<\lambda^*<\frac{B+\sqrt{(B^2-AC)}}{(1-\delta^*)A}\ , \tag{44}$$

$$A=(\mu_1^{*2}C_{x_1}^2+\mu_2^{*2}C_{x_2}^2-2\mu_1^*\mu_2^*\rho_{x_1x_2}C_{x_1}C_{x_2}),$$

$$B = (\mu_1^* \rho_{yx_1} C_y C_{x_1} - \mu_2^* \rho_{yx_2} C_y C_{x_2}),$$

$$C = [C_{x_1}^2 (1 - 2K_{yx_1}) + C_{x_2}^2 \{1 + 2(K_{yx_2} - K_{x_1x_1})\}].$$

We also note from (4) and (36) that the estimator $\hat{\bar{Y}}_{2u}$ is better than $\hat{\bar{Y}}_2(or \hat{\bar{Y}}_{2u}^*)$ if

$$\left.\begin{array}{ll} either & \frac{1}{(1-\delta^*)} < \lambda^* < \left[2\lambda_{opt}^* - \frac{1}{(1-\delta^*)}\right] \\ or & \left[2\lambda_{opt}^* - \frac{1}{(1-\delta^*)}\right] < \lambda^* < \frac{1}{(1-\delta^*)} \end{array}\right\} \quad (45)$$

The optimum value λ_{opt}^* of λ^* can be obtained quite accurately through past data or experience.

6. EMPIRICAL STUDY

To observe the relative performance of different estimators of $\bar{Y}$, we consider a natural population data set given in Steel and Torrie (1960, p.282). The population description is given below:

y : Log of leaf burn in sec.
x_1 : Potassiam percentage
x_2 : Clorine percentage.

The required population values are:

$\bar{Y} = 0.6860$, $C_y = 0.4803$, $\rho_{yx_1} = 0.1794$, N=30 ,
$\bar{X}_1 = 4.6537$, $C_{x_1} = 0.2295$, $\rho_{yx_2} = -0.4996$, n=6,
$\bar{X}_1 = 0.8077$, $C_{x_2} = 0.7493$, $\rho_{x_1x_2} = 0.4074$, g=2.

The percentage relative efficiencies (PREs) of various estimators of $\bar{Y}$ with respect to $\bar{y}$ have been computed and presented in Table 1.

TABLE 1

Percent relative efficiencies of different estimators of $\bar{Y}$ with respect to $\bar{y}$

Estimator	$\bar{y}$	$\bar{y}_R$	$\bar{y}_P$	$\hat{\bar{Y}}_1$	$\hat{\bar{Y}}_2(\hat{\bar{Y}}_{2u}^*)$	$\hat{\bar{Y}}_{2u}$ with $\lambda_{opt}^* = 1.19751$
PRE $(\bullet, \bar{y})$	100.00	94.62	53.33	75.50	142.18	165.88

Table 1 clearly indicates that the suggested estimators $\hat{\bar{Y}}_2 (or\, \hat{\bar{Y}}_{2u}^*)$ and $\hat{\bar{Y}}_{2u}$ with $\lambda^* = \lambda_{opt}^*$, are more efficient than usual unbiased estimator $\bar{y}$, ratio estimator $\bar{y}_R$, product estimator $\bar{y}_P$, and Singh's (1967) ratio-cum-product estimator $\hat{\bar{Y}}_1$ with considerable gain in efficiency.

7. CONCLUDING REMARKS

Usually information regarding correlation coefficient $\rho_{x_1 x_2}$ between the two auxiliary variates X_1 and X_2 is known or can made known to the experimenter through past studies or with the familiarity of experimental material. When $\rho_{x_1 x_2}$ is known an improved version $\hat{\bar{Y}}_{2u}$ of Singh's (1967) estimator $\hat{\bar{Y}}_1$ is suggested with its properties. Using 'Jack-knife' technique envisaged by Quenouille (1956), a family of unbiased estimators $\hat{\bar{Y}}_{2u}$ is also proposed. A large number of unbiased estimators can be generated from $\hat{\bar{Y}}_{2u}$. Asymptotically optimum estimator (AOE) in the family of estimators $\hat{\bar{Y}}_{2u}$ is identified with its variance formula. It is shown that the suggested family of estimators $\hat{\bar{Y}}_{2u}$ is more efficient than $\bar{y}$ and $\hat{\bar{Y}}_2$ at optimum conditions. Empirical study also suggests that the suggested estimators $\hat{\bar{Y}}_2 (or\, \hat{\bar{Y}}_{2u}^*)$ and $\hat{\bar{Y}}_{2u}$ with $\lambda^* = \lambda_{opt}^*$ are better than $\bar{y}$, $\bar{y}_R$, $\bar{y}_P$ and Singh's (1967) estimator $\hat{\bar{Y}}_1$. Thus we conclude that the proposed estimators $\hat{\bar{Y}}_2 (or\, \hat{\bar{Y}}_{2u}^*)$ and $\hat{\bar{Y}}_{2u}$ are to be preferred in practice.

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RIASSUNTO

Stima della media di un popolazione finita con coefficiente di correlazione tra caratteri ausiliari noto

Il contributo propone uno stimatore *ratio-cum-product* modificato della media di una popolazione finita di una variabile oggetto di studio Y sfruttando il coefficiente di correlazione noto tra due caratteri ausiliari X_1 e X_2. Si ottiene uno stimatore *ratio-cum-product* quasi corretto attraverso la tecnica Jacknife del tipo previsto da Quenille (1956). In seguito vengono esaminati con un esempio numerico i meriti dello stimatore proposto.

SUMMARY

Estimation of finite population mean using known correlation coefficient between auxiliary characters

This paper proposes a modified ratio-cum-product estimator of finite population mean of the study variate Y using known correlation coefficient between two auxiliary characters X_1 and X_2. An almost unbiased ratio-cum-product estimator has also been obtained by using Jackknife technique envisaged by Quenouille (1956). The merits of the proposed estimator are examined through a numerical illustration.