SELECTION OF THE BEST WITH A PRELIMINARY
TEST FOR LOCATION-SCALE PARAMETRIC MODELS *

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Dedicated to the late Professor P. R. Krishnaiah

ABSTRACT

This paper deals with the problem of selecting the best population from among \( k (\geq 2) \) populations which can be described by location-scale parametric models. New selection procedures are proposed for selecting the unique best in terms of the largest location parameter. The procedures include preliminary tests which allow the experimenter to have an option to not select if the statistical evidence is not significant. Two probabilities, the probability to make a selection and the probability of a correct selection, are controlled by these selection procedures. Applications to the normal mean models are considered. Comparisons between the proposed selection procedures and certain earlier existing procedures are also made. Finally, a two-stage procedure for the normal means problem is considered.

Key Words and Phrases: best population; correct selection; preliminary test; indifference zone; nonselection zone; preference zone; location-scale parametric model; two-stage procedure.

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1. **INTRODUCTION**

The problem of selecting the best population from among \( k(\geq 2) \) populations has been studied extensively. Many selection procedures have been derived for different selection goals by several authors. Among them, Bechhofer (1954) introduced the indifference zone approach for selecting the normal population with the largest mean. In his approach, the determination of the sample size depends entirely on the indifference zone assumption. Also, the probability of a correct selection depends on the unknown parameters and is analogous to the power of a test. However, in this formulation, a probability that is analogous to the probability of type-I error of a test was not taken into consideration. Bechhofer's procedure forces the experimenter to make a selection, and often that procedure is not used in practical applications because of the lack of a statistical test for the homogeneity of the parameters as stated in Simon (1977). It should be pointed out that in another approach called the subset selection this drawback, namely, the assumption of indifference zone, is not there. This approach due to Gupta (1956, 1965) as well as Bechhofer's approach are both discussed in detail in the monograph by Gupta and Panchapakesan (1979).

Based on the preceding reasoning, Bishop and Pirie (1979) introduced a selection procedure in which a test of homogeneity was conducted. The procedure allows the experimenter to have the option not to make a selection if the statistical evidence is not significant. Later, Chen (1985) proposed a modified selection procedure for the problem of selecting the best normal population. He considered a preliminary test based on the sampled spacing between the largest and the second largest order statistics. If the statistical evidence of the preliminary test is not significant, the experimenter decides not to make
a selection. Otherwise, he or she selects the population yielding the largest sample mean value as the best population. The sample size is determined to control both the probability of type-I error for the preliminary test and the probability of a correct selection. Analogous to Chen (1985), Chen and Mithongtae (1986) proposed selection procedures for two-parameter exponential distribution models. However, both of their procedures cannot be applied to a case where the common scale parameter is unknown. Later, Leu and Liang (1990) proposed selection procedures which improve the results of Chen and Mithongtae (1986) and discussed the case where Chen and Mithongtae's procedures cannot be applied.

In this paper, we generalize the problem proposed by Chen (1985) to location-scale parametric models. Selection procedures based on one-sample are derived according to whether the common scale parameter is known or unknown. Exact sample sizes are determined to control both the probability of type-I error and the probability of a correct selection. Applications to the normal model cases are considered. Comparisons between the proposed selection procedures and certain earlier existing procedures are also made. Finally, a two-stage procedure for the normal means problem is considered.

2. FORMULATION OF THE PROBLEM

Let $\pi_1, \ldots, \pi_k$ denote $k(\geq 2)$ independent location-scale parametric models which have absolutely continuous cumulative distribution functions (cdf) $G(\frac{x-\theta_1}{\sigma}), \ldots, G(\frac{x-\theta_k}{\sigma})$, respectively, where $\sigma > 0$, $-\infty < \theta_i < \infty$, $i = 1, \ldots, k$ and $-\infty < x < \infty$. Let $\theta = (\theta_1, \ldots, \theta_k)$ and let $\theta_{[1]} \leq \ldots \leq \theta_{[k]}$ denote the ordered values of $\theta_1, \ldots, \theta_k$. It is assumed that the exact pairing between the ordered parameters and the unordered parameters is unknown. The population associated with the largest location parameter $\theta_{[k]}$ is called the
best population. Assume that the experimenter is interested in the selection of the best population. Let

$$\Omega = \{(\theta, \sigma) | \theta = (\theta_1, \ldots, \theta_k), -\infty < \theta_i < \infty, \sigma > 0\}$$

be the parameter space. We partition the parameter space into the following three sub-spaces:

- the preference zone: $\Omega(PZ) = \{(\theta, \sigma) \in \Omega | \frac{\theta[k] - \theta[k-1]}{\sigma} \geq \delta, \delta > 0\}$,
- the nonselection zone: $\Omega(NZ) = \{(\theta, \sigma) \in \Omega | \theta[k-1] = \theta[k]\}$,
- the indifference zone: $\Omega(IZ) = \Omega - \Omega(PZ) - \Omega(NZ)$,

where $\delta$ is a known positive constant. Note that our $\Omega(IZ)$ is different from the indifference zone of Bechhofer (1954).

Denote the event of a correct selection by $CS$ and the event of a selection by $S$. The goal is to develop a selection procedure $R$ to select a single best population with a minimum sample size from each of the $k$ populations such that the following probability requirements are satisfied:

$$P_{(\theta, \sigma)}(S|R) \leq \alpha \text{ for all } (\theta, \sigma) \in \Omega(NZ) \quad (2.1)$$

and

$$P_{(\theta, \sigma)}(CS|R) \geq P^* \text{ for all } (\theta, \sigma) \in \Omega(PZ) \quad (2.2)$$

where $\alpha \in (0,1)$ and $P^* \in (1/k, 1)$ are preassigned probability levels.

The selection procedure $R$ depends on whether the common scale parameter $\sigma$ is known or unknown.
3. SELECTION PROCEDURE WITH KNOWN SCALE PARAMETER

Let $X_{ij}$, $j = 1, \ldots, n$ be $n$ independent observations from population $\pi_i$, where $\pi_i \sim G(\frac{x - \theta_i}{\sigma})$, $i = 1, \ldots, k$, respectively. Let $Y_i = Y(X_{i1}, \ldots, X_{in})$ be an appropriate statistic for $\theta_i$. We assume that $Y_i$ has the cdf $F_n(\frac{y - \theta_i}{\sigma})$. Also let $Y_{[1]} \leq \ldots \leq Y_{[k]}$ denote the order statistics of $Y_1, \ldots, Y_k$. When $\sigma$ is known, we propose a selection procedure as follows:

$R_1$: Select the population yielding $Y_{[k]}$ as the best population if $Y_{[k]} - Y_{[k-1]} > \lambda(n, \alpha)\sigma$; otherwise, do not make a selection, where $n$ and $\lambda(n, \alpha)$ are chosen to satisfy the probability requirements (2.1) and (2.2).

For the given rule $R_1$, we need to investigate the supremum of $P(\theta, \sigma)(S|R_1)$ for $(\theta, \sigma) \in \Omega(NZ)$ and the infimum of $P(\theta, \sigma)(CS|R_1)$ for $(\theta, \sigma) \in \Omega(PZ)$. We need to make two assumptions here:

(A) The probability density function (pdf) $f_n$ of $F_n$ is log-concave.

(B) For each $\delta > 0$ and any positive integer $m$, $\int_{-\infty}^{\infty} F_n^m(y + \delta) dF_n(y)$ is strictly increasing in $n$, and tends to 1 as $n \to \infty$.

These two assumptions are appropriate for many applications. First, we consider a lemma derived by Kim (1986).

Lemma 3.1. Assume that $\log f_n(y)$ is concave. Then for any fixed $c > 0$, $P(\theta, \sigma)\{Y_{[k]} - Y_{[k-1]} > c\}$ is non-increasing in $\theta_{[1]}$ and hence

$$P(\theta, \sigma)\{Y_{[k]} - Y_{[k-1]} > c\} \leq P(\theta^0, \sigma)\{|Y_{(k)} - Y_{(k-1)}| > c\}$$

for all $(\theta, \sigma) \in \Omega$, where $\theta^0 = (\theta^0_1, \ldots, \theta^0_k), \theta^0_{[i]} = -\infty$, $i = 1, \ldots, k - 2, \theta^0_{[i]} = \theta_{[i]}$ for $i = k - 1, k$, and $Y_{(i)}$ is the statistic associated with parameter $\theta_{[i]}$, $i = 1, \ldots, k$.

In the sequel, we let $Z_i = \frac{Y_{(i)} - \theta_{[i]}}{\sigma}$, $i = 1, \ldots, k$ and let $H_n(t)$ be the cdf of $Z_1 - Z_2$. 

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Then

\[ H_n(t) = \int_{-\infty}^{\infty} F_n(y + t) dF_n(y). \quad (3.1) \]

We note that \( H_n(t) \) is a symmetric distribution function and hence \( H_n(-t) = 1 - H_n(t) \).

By using Lemma 3.1, we have the following theorem:

**Theorem 3.2.** \( \sup_{\Omega(NZ)} P_{(\theta, \sigma)}(S|R_1) = 2H_n(-\lambda(n, \alpha)) \). \quad (3.2)

**Proof:** By Assumption (A) and Lemma 3.1, we have

\[
P_{(\theta, \sigma)}(S|R_1) = P_{(\theta, \sigma)} \{ Y_{[k]} - Y_{[k-1]} > \lambda(n, \alpha) \sigma \} \leq P_{(\theta^0, \sigma)} \{ |Y(k) - Y_{(k-1)}| > \lambda(n, \alpha) \sigma \}.
\]

Hence

\[
\sup_{\Omega(NZ)} P_{(\theta, \sigma)}(S|R_1) = P\{ |Z_k - Z_{k-1}| > \lambda(n, \alpha) \} = 2H_n(-\lambda(n, \alpha)). \quad \square
\]

In order to satisfy the probability requirement (2.1), we may let \( 2H_n(-\lambda(n, \alpha)) = \alpha \).

That is,

\[
\lambda(n, \alpha) = -H_n^{-1}(\alpha/2). \quad (3.3)
\]

**Remark 3.1.** \( \lambda(n, \alpha) \geq 0 \), since \( H_n \) is symmetric.

**Lemma 3.3.** The \( \lambda(n, \alpha) \) defined by (3.3) has the properties: \( \lambda(n, \alpha) \) is decreasing in \( n \) and \( \lambda(n, \alpha) \to 0 \) as \( n \to \infty \), for each fixed \( \alpha, 0 < \alpha < 1 \).

**Proof:** For \( n_1 \geq n_2 \), if \( \lambda(n_1, \alpha) > \lambda(n_2, \alpha) \), then by Assumption (B) we have

\[
\alpha/2 = H_{n_1}(-\lambda(n_1, \alpha)) = 1 - H_{n_1} (\lambda(n_1, \alpha)) \leq 1 - H_{n_1}(t) < 1 - H_{n_2}(t) \leq 1 - H_{n_2}(\lambda(n_2, \alpha)) = H_{n_2}(-\lambda(n_2, \alpha)) = \alpha/2, \text{ for } \lambda(n_1, \alpha) > t > \lambda(n_2, \alpha).
\]
This is a contradiction and hence \( \lambda(n_1, \alpha) \leq \lambda(n_2, \alpha) \).

If \( \lim_{n \to \infty} \lambda(n, \alpha) = c > 0 \), then \( \lambda(n, \alpha) \geq c \) for all \( n \), since \( \lambda(n, \alpha) \) is decreasing in \( n \) for each fixed \( \alpha \). Thus, by the definition of \( \lambda(n, \alpha) \), \( 1 - \frac{\alpha}{2} = H_n(\lambda(n, \alpha)) \geq H_n(c) \).

By Assumption (B), as \( c > 0 \), \( H_n(c) \to 1 \) as \( n \to \infty \). Therefore, for \( n \) sufficiently large, \( H_n(c) \geq 1 - \frac{\alpha}{4} > H_n(\lambda(n, \alpha)) \), which is a contradiction. Hence, we have \( \lambda(n, \alpha) \to 0 \) as \( n \to \infty \).

We next evaluate the infimum of \( P_{(\theta, \sigma)}(CS|R_1) \) over \( (\theta, \sigma) \in \Omega(PZ) \).

**Theorem 3.4.** The infimum of \( P_{(\theta, \sigma)}(CS|R_1) \) over \( \Omega(PZ) \) occurs at the configuration \( \theta[1] = \ldots = \theta[k-1] = \theta[k] - \delta \sigma \) and

\[
\inf_{\Omega(PZ)} P_{(\theta, \sigma)}(CS|R_1) = \int_{-\infty}^{\infty} F_n^{k-1}(y + \delta - \lambda(n, \alpha))dF_n(y). \tag{3.4}
\]

**Proof:**

\[
P_{(\theta, \sigma)}(CS|R_1) = P_{(\theta, \sigma)}\{Y_{[k]} - Y_{[k-1]} > \lambda(n, \alpha)\sigma, Y_{(k)} = Y_{[k]}\}
= P_{(\theta, \sigma)}\{Z_j < Z_k + \frac{\theta_{[k]} - \theta_{[j]}}{\sigma} - \lambda(n, \alpha), j = 1, \ldots, k - 1\}
\geq P_{(\theta, \sigma)}\{Z_j < Z_k + \delta - \lambda(n, \alpha), j = 1, \ldots, k - 1\}. \tag{3.5}
\]

The equality in (3.5) holds when \( \theta_{[1]} = \ldots = \theta_{[k-1]} = \theta_{[k]} - \delta \sigma \). Hence

\[
\inf_{\Omega(PZ)} P_{(\theta, \sigma)}(CS|R_1) = P_{(\theta, \sigma)}\{Z_j < Z_k + \delta - \lambda(n, \alpha), j = 1, \ldots, k - 1\}
= \int_{-\infty}^{\infty} F_n^{k-1}(y + \delta - \lambda(n, \alpha))dF_n(y). \quad \square
\]

**Remark 3.2.**

(1) Since \( \lambda(n, \alpha) \to 0 \) as \( n \to \infty \) for each \( 0 < \alpha < 1 \), hence, by Assumption (B), we have, for each fixed \( \delta > 0 \), \( \int_{-\infty}^{\infty} F_n^{k-1}(y + \delta - \lambda(n, \alpha))dF_n(y) \to 1 \) as \( n \to \infty \).
(2) In order to satisfy the probability requirement (2.2), we may let the right hand side of (3.4) equal \( P^* \). In practice, \( n \) is chosen to be the smallest integer such that (2.1) and (2.2) are satisfied.

4. SELECTION PROCEDURE WITH UNKNOWN SCALE PARAMETER

When the scale parameter \( \sigma \) is unknown, it is assumed that \( Y_i = Y(X_{i1}, \ldots, X_{in}) \) is a complete sufficient statistic for the parameter \( \theta_i \) for each fixed \( \sigma > 0 \). Let \( T_i = T(X_{i1}, \ldots, X_{in}) \) be a nonnegative function of \( X_{i1}, \ldots, X_{in} \), such that \( T(X_{i1} - a, \ldots, X_{in} - a) = T(X_{i1}, \ldots, X_{in}) \) and \( T(cX_1, \ldots, cX_n) = cT(X_1, \ldots, X_n) \) for any \( c > 0 \). Also, let \( S = S(T_1, \ldots, T_k) \) be a nonnegative function of \( T_1, \ldots, T_k \) such that \( S(cT_1, \ldots, cT_k) = cS(T_1, \ldots, T_k) \) for any \( c > 0 \). Since \( Y_i \) is a complete sufficient statistic for \( \theta_i \) and the distribution of \( T_i \) is independent of \( \theta_i \), hence \( T_i \) is independent of \( Y_i \). Therefore, \( S \) is independent of \( (Y_1, \ldots, Y_k) \). Also, by the preceding assumption, the distribution of \( W = S/\sigma \) is independent of the parameters \( \theta_1, \ldots, \theta_k \) and \( \sigma \). Hence, we propose a selection procedure \( R_2 \) as follows:

\( R_2 \): Select the population yielding \( Y_{[k]} \) as the best population if \( Y_{[k]} - Y_{[k-1]} \) \( \gtrsim \lambda(n, \alpha)S \); otherwise, do not make a selection, where \( n \) and \( \lambda(n, \alpha) \) are chosen to satisfy the probability requirements (2.1) and (2.2).

First, we evaluate the supremum of \( P(\theta, \sigma) (S| R_2) \) over \( (\theta, \sigma) \in \Omega(NZ) \). In the following, let \( G(w) \) denote the cdf of \( W = S/\sigma \).

**Theorem 4.1.**

\[
\sup_{\Omega(NZ)} P(\theta, \sigma) (S| R_2) = 2 \int_{0}^{\infty} \int_{-\infty}^{\infty} F_n(y - \lambda(n, \alpha)w) dF_n(y) dG(w). \quad (4.1)
\]
Proof:

\[ P_{(\hat{\theta}, \sigma)}(S|R_2) = P_{(\hat{\theta}, \sigma)}\{Y[k] - Y_{[k-1]} > \bar{\lambda}(n, \alpha)S\} \]

\[ = E\{P_{(\hat{\theta}, \sigma)}\{Y[k] - Y_{[k-1]} > \bar{\lambda}(n, \alpha)\sigma W|W\}\} \]

\[ = \int_0^\infty P_{(\hat{\theta}, \sigma)}\{Y[k] - Y_{[k-1]} > \bar{\lambda}(n, \alpha)\sigma w\}dG(w). \]

By Theorem 3.2, we have

\[ \sup_{\Omega(NZ)} P_{(\hat{\theta}, \sigma)}\{Y[k] - Y_{[k-1]} > \bar{\lambda}(n, \alpha)\sigma w\} = 2H_n(-\bar{\lambda}(n, \alpha)w). \]

Hence

\[ \sup_{\Omega(NZ)} P_{(\hat{\theta}, \sigma)}(S|R_2) = \int_0^\infty 2H_n(-\bar{\lambda}(n, \alpha)w)dG(w) \]

\[ = 2\int_0^\infty \int_{-\infty}^\infty F_n(y - \bar{\lambda}(n, \alpha)w) dF_n(y) dG(w). \]

For each 0 < \alpha < 1, let \bar{\lambda}(n, \alpha) be the solution of the equation

\[ \int_0^\infty H_n(-\bar{\lambda}(n, \alpha)w)dG(w) = \alpha/2. \quad (4.2) \]

Then the probability requirement (2.1) is satisfied.

Remark 4.1. \( \bar{\lambda}(n, \alpha) \) is nonnegative.

Lemma 4.2. The \( \bar{\lambda}(n, \alpha) \) defined by (4.2) has the properties: \( \bar{\lambda}(n, \alpha) \) is decreasing in \( n \) and \( \bar{\lambda}(n, \alpha) \to 0 \) as \( n \to \infty \).

Proof: If \( n_1 \geq n_2 \) and \( \bar{\lambda}(n_1, \alpha) > \bar{\lambda}(n_2, \alpha) \), then

\[ \alpha/2 = \int_0^\infty H_{n_1}(-\bar{\lambda}(n_1, \alpha)w)dG(w) = 1 - \int_0^\infty H_{n_1}(\bar{\lambda}(n_1, \alpha)w)dG(w) \]

\[ < 1 - \int_0^\infty H_{n_2}(\bar{\lambda}(n_2, \alpha)w)dG(w) = \int_0^\infty H_{n_2}(-\bar{\lambda}(n_2, \alpha)w)dG(w) = \alpha/2, \]

which is a contradiction. Hence \( \bar{\lambda}(n, \alpha) \) is decreasing in \( n \) for each fixed \( \alpha, 0 < \alpha < 1 \).

Moreover, if \( \lim_{n \to \infty} \bar{\lambda}(n, \alpha) = c > 0 \), then

\[ 1 - \alpha/2 = \int_0^\infty H_n(\bar{\lambda}(n, \alpha)w)dG(w) \geq \int_0^\infty H_n(c)dG(w). \]
However, as $c > 0$, by Assumption (B), $\int_0^\infty H_n(c) dG(w) \to 1$ as $n \to \infty$, which leads to a contradiction. Hence, $\lambda(n, \alpha) \to 0$ as $n \to \infty$. 

Now we evaluate the infimum of $P_{(\theta, \sigma)}(CS|R_2)$ over $(\theta, \sigma) \in \Omega(PZ)$.

**Theorem 4.3.** The infimum of $P_{(\theta, \sigma)}(CS|R_2)$ over $\Omega(PZ)$ occurs at the configuration $\theta_{[1]} = \ldots = \theta_{[k-1]} = \theta_{[k]} - \delta \sigma$ and

$$
\inf_{\Omega(PZ)} P_{(\theta, \sigma)}(CS|R_2) = \int_0^\infty \int_{-\infty}^\infty F_n^{k-1}(y + \delta - \lambda(n, \alpha)w) dF_n(y) dG(w). \quad (4.3)
$$

**Proof:**

$$
P_{(\theta, \sigma)}(CS|R_2) = P_{(\theta, \sigma)}\{Y_{[k]} - Y_{[k-1]} > \lambda(n, \alpha)S, Y_{(k)} = Y_{[k]}\}
$$

$$
= P_{(\theta, \sigma)}\{Z_j < Z_k + \frac{\theta_{[k]} - \theta_{[j]}}{\sigma} - \lambda(n, \alpha)W, j = 1, \ldots, k - 1\}
$$

$$
\geq P\{Z_j < Z_k + \delta - \lambda(n, \alpha)W, j = 1, \ldots, k - 1\} \quad (4.4)
$$

$$
= \int_0^\infty \int_{-\infty}^\infty F_n^{k-1}(y + \delta - \lambda(n, \alpha)w) dF_n(y) dG(w).
$$

The equality in (4.4) holds when $\theta_{[1]} = \ldots = \theta_{[k-1]} = \theta_{[k]} - \delta \sigma$. 

**Remark 4.2.**

1. For fixed $w > 0$ and $0 < \alpha < 1$, we have $\lambda(n, \alpha)w \to 0$ as $n \to \infty$. Now by Assumption (B), we have

$$
\int_0^\infty \int_{-\infty}^\infty F_n^{k-1}(y + \delta - \lambda(n, \alpha)w) dF_n(y) dG(w) \to 1 \text{ as } n \to \infty.
$$

2. In order to satisfy the probability requirement (2.2), we may choose the smallest $n$ such that

$$
\int_0^\infty \int_{-\infty}^\infty F_n^{k-1}(y + \delta - \lambda(n, \alpha)w) dF_n(y) dG(w) \geq P^*. \quad (4.5)
$$
5. **NORMAL MODEL CASE**

In this section, it is assumed that the populations \( \pi_1, \ldots, \pi_k \) have normal distributions with means \( \theta_1, \ldots, \theta_k \), respectively, and a common variance \( \sigma^2 \). Let \( X_{ij}, j = 1, \ldots, n, \) be independent samples from \( \pi_i, i = 1, \ldots, k \) and \( Y_i = \bar{X}_i = \frac{1}{n} \sum_{j=1}^{n} X_{ij}, i = 1, \ldots, k \). Then \( Y_i \sim N(\theta_i, \frac{\sigma^2}{n}) \) and \( F_n(x) = \Phi(\sqrt{n}x) \) where \( \Phi \) is the cdf of the standard normal. One can easily check that Assumptions (A) and (B) are satisfied. Also, we have \( H_n(t) = \Phi(\sqrt{\frac{n}{2}}t) \).

We discuss both cases when the scale parameter \( \sigma \) is known or unknown.

(a) \( \sigma \) known case:

It follows from (3.3) that

\[
\lambda(n, \alpha) = \sqrt{\frac{2}{n}} z_{\alpha/2}
\]  

(5.1)

where \( z_{\alpha/2} \) denotes the upper \( \alpha/2 \) quantile of the standard normal distribution. Thus, the selection procedure is:

**R_1:** Select the population yielding \( Y_{[k]} \) as the best population if \( Y_{[k]} - Y_{[k-1]} > \sqrt{\frac{2}{n}} z_{\alpha/2} \sigma \); otherwise, do not make a selection.

This is the same procedure as the one proposed by Chen (1985).

From equation (3.4), we have

\[
\inf_{\Omega(P, \pi)} P_{(\theta, \sigma)}(CS|R_1) = \int_{-\infty}^{\infty} \Phi^{k-1}(y + \sqrt{n}\delta - \sqrt{2z_{\alpha/2}}) d\Phi(y).
\]

For given \( k, P^* \) and \( \alpha \), let \( d \) be the solution to the equation

\[
\int_{-\infty}^{\infty} \Phi^{k-1}(y + d) d\Phi(y) = P^*.
\]  

(5.2)

Then the sample size required to satisfy the probability requirements (2.1) and (2.2) is given by

\[
n = \langle \frac{d + \sqrt{2z_{\alpha/2}}}{\delta} \rangle,
\]
where \( \lfloor x \rfloor \) is the smallest integer not less than \( x \). The solutions of \( d \)-values in equation (5.2) can be found in Bechhofer (1954), Gupta (1963) and Gupta, Nagel and Panchapakesan (1973).

(b) \( \sigma \) unknown case:

Let \( T_i^2 = \sum_{j=1}^{n} (X_{ij} - Y_i)^2 \) and \( S^2 = \frac{1}{v} \sum_{i=1}^{k} T_i^2 \), where \( v = k(n - 1) \). Then \( vS^2/\sigma^2 = vW^2 \) follows a \( \chi^2 \)-distribution with \( v \) degrees of freedom. We denote the distribution of the nonnegative random variable \( W \) by \( G \).

For \( \alpha \in (0, 1) \), we determine \( \tilde{\lambda}(n, \alpha) \) by solving the equation

\[
\int_0^\infty H_n(-\tilde{\lambda}(n, \alpha)w)dG(w) = \alpha/2.
\]

However, we can solve an easy equation in this case. Because

\[
\int_0^\infty H_n(-\tilde{\lambda}(n, \alpha)w)dG(w) = P\{Z \leq -\tilde{\lambda}(n, \alpha)W\}
\]

where \( Z \sim N(0, \frac{2}{n}) \) and \( VW^2 \sim \chi^2_v \) are independent, thus

\[
\int_0^\infty H_n(-\tilde{\lambda}(n, \alpha)w)dG(w) = P\left\{\sqrt{\frac{n}{2}} Z \leq -\sqrt{\frac{n}{2}} \tilde{\lambda}(n, \alpha)\right\}
\]

where \( \sqrt{\frac{n}{2}} Z \sim t_v \), the \( t \)-distribution with degrees of freedom \( v \). Hence

\[
\tilde{\lambda}(n, \alpha) = \sqrt{\frac{n}{2}} t_v, \alpha/2,
\]

where \( t_v, \alpha/2 \) is the upper \( \alpha/2 \) quantile of the \( t_v \) distribution.

Furthermore, from (4.5), \( n \) is the smallest integer such that

\[
\int_0^\infty \int_{-\infty}^{\infty} \Phi^{k-1}(y + \sqrt{n}\delta - \sqrt{2}t_v, \alpha/2w)d\Phi(y)dG(w) \geq P^*.
\]

Remark: Although this result is the same as that of Chen (1985), the concept is different.

Chen's method should specify the ratio \( \delta^*/\sigma \) in advance. However, we need not make this assumption in our approach.
6. A TWO-STAGE PROCEDURE WHEN SCALE PARAMETER IS UNKNOWN

If we partition the parameter space $\Omega$ into three parts as follows.

$$\tilde{\Omega}(PZ) = \{(\theta, \sigma) \in \Omega | \theta_{[k]} - \theta_{[k-1]} \geq \delta^*, \delta^* > 0\}$$

$$\tilde{\Omega}(NZ) = \{(\theta, \sigma) \in \Omega | \theta_{[k-1]} = \theta_{[k]}\}$$

and

$$\tilde{\Omega}(IZ) = \Omega - \tilde{\Omega}(PZ) - \tilde{\Omega}(NZ).$$

If $\sigma$ were known, we can take $\delta = \delta^*/\sigma$, then the result is the same as that of Section 3.

When $\sigma$ is unknown, a single-stage procedure does not exist for this problem. In the following, we consider the normal case only. Analogous to that of Bechhofer, Dunnett and Sobel (1954), a two-stage selection procedure $R_3$ is proposed as follows:

(i) Take an initial sample of size $n_0$ from each of the $k$ populations, say $X_{i1}, \ldots, X_{in_0}, i = 1, \ldots, k$.

Let $Y_i(n_0) = \frac{1}{n_0} \sum_{j=1}^{n_0} X_{ij}$ and $S_0^2 = \sum_{i=1}^{k} \sum_{j=1}^{n_0} (X_{ij} - Y_i(n_0))^2 / v_0, v_0 = k(n_0 - 1)$, and $W_0 = S_0/\sigma$.

(ii) Define $N = \max\{n_0, < \frac{h^2 S_0^2}{\delta^2}, >\}$, where $h > \sqrt{2t_{v_0, \alpha/2}}$ is determined by

$$\int_{0}^{\infty} \int_{-\infty}^{\infty} \Phi^{k-1}(y + (h - \sqrt{2t_{v_0, \alpha/2}})w) d\Phi(y) dG_0(w), \quad (6.1)$$

where $G_0(w)$ is the cdf of $W_0$.

(iii) If necessary, take $N - n_0$ additional observations from each of the $k$ populations and compute

$$Y_i(N) = \frac{1}{N} \sum_{j=1}^{N} X_{ij}, \ i = 1, \ldots, k.$$ 

(iv) The selection rule $R_3$ is defined by:
$R_3$: Select the population yielding $Y_{[k]}(N)$ as the best population if $Y_{[k]}(N) - Y_{[k-1]}(N) > \frac{\lambda}{\sqrt{N}}S_0$; otherwise do not make a selection; here $\lambda$ is chosen to satisfy the probability requirement (2.1).

For the procedure $R_3$ defined above, we have the following result:

**Theorem 6.1.**

\[
\sup_{\mathcal{H}(NZ)} P_{(\mathcal{E}, \sigma)}(S|R_3) = P\{|T| > \frac{\lambda}{\sqrt{2}}\}, \tag{6.2}
\]

where $T$ follows a Student’s $t$-distribution with $v_0$ degrees of freedom.

Proof: Let $A_n = \{N = n\}$, then $A_{n_0} = \{W_0^2 \leq \frac{n_0 \delta^2}{h^2 \sigma^2}\}$ and

\[
A_n = \{(n - 1)\delta^2 < W_0^2 \leq \frac{n \delta^2}{h^2 \sigma^2}\} \text{ if } n > n_0.
\]

Thus,

\[
P_{(\mathcal{E}, \sigma)}(S|R_3) = P_{(\mathcal{E}, \sigma)}\{Y_{[k]}(N) - Y_{[k-1]}(N) > \frac{\lambda}{\sqrt{N}}S_0\}
\]

\[
= \sum_{n=n_0}^{\infty} \int_{A_n} P_{(\mathcal{E}, \sigma)}\{\sqrt{N} \frac{Y_{[k]}(N)}{\sigma} - \sqrt{N} \frac{Y_{[k-1]}(N)}{\sigma} > \lambda w\}dG_0(w)
\]

\[
\leq \sum_{n=n_0}^{\infty} \int_{A_n} P_{(\mathcal{E}, \sigma)}\{|\sqrt{N} \frac{Y_{(k)}(N)}{\sigma} - \sqrt{N} \frac{Y_{(k-1)}(N)}{\sigma}| > \lambda w\}dG_0(w).
\]

Therefore, we have

\[
\sup_{\mathcal{H}(NZ)} P_{(\mathcal{E}, \sigma)}(S|R_3) = \sum_{n=n_0}^{\infty} \int_{A_n} 2\Phi(-\frac{\lambda w}{\sqrt{2}})dG_0(w)
\]

\[
= \int_{0}^{\infty} 2\Phi(-\frac{\lambda w}{\sqrt{2}})dG_0(w)
\]

\[
= P\{|T| > \frac{\lambda}{\sqrt{2}}\}.
\]

In order to satisfy the probability requirement (2.1), we set $P\{|T| > \frac{\lambda}{\sqrt{2}}\} = \alpha$. Then

$\lambda = \sqrt{2t_{v_0, \alpha/2}}$.

Now, we evaluate the infimum of $P_{(\mathcal{E}, \sigma)}(CS|R_3)$ over $(\mathcal{E}, \sigma) \in \tilde{\mathcal{H}}(PZ)$. 

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Theorem 6.2. The infimum of $P(\theta, \sigma)(CS|R_3)$ over $\tilde{\Omega}(PZ)$ occurs at the configuration
\[ \theta_{[1]} = \ldots = \theta_{[k-1]} = \theta_{[k]} - \delta^* \] and
\[ \inf_{\tilde{\Omega}(PZ)} P(\theta, \sigma)(CS|R_3) \geq \int_0^\infty \int_{-\infty}^\infty \Phi^{k-1}(y + (h - \sqrt{2}t_{v_0, \alpha/2})w)d\Phi(y)dG_0(w). \quad (6.3) \]

Proof:
\[ P(\theta, \sigma)(CS|R_3) = P(\theta, \sigma)\{Y(j)(N) < Y(h)(N) - \frac{\lambda}{\sqrt{N}}S_0, \ j = 1, \ldots, k - 1\} \]
\[ = P(\theta, \sigma)\{Z_j < Z_k + \frac{\sqrt{N}(\theta_{[k]} - \theta_{[j]})}{\sigma} - \lambda W_0, \ j = 1, \ldots, k - 1\} \]
\[ \geq P(\theta, \sigma)\{Z_j < Z_k + \frac{\sqrt{N} \delta^*}{\sigma} - \lambda W_0, \ j = 1, \ldots, k - 1\} \quad (6.4) \]
\[ \geq P(\theta, \sigma)\{Z_j < Z_k + (h - \lambda)W_0, \ j = 1, \ldots, k - 1\}, \text{ since } \sqrt{N} \geq \frac{hS_0}{\delta^*} \]
\[ = \int_0^\infty \int_{-\infty}^\infty \Phi^{k-1}(y + (h - \lambda)w)d\Phi(y)dG_0(w). \quad (6.5) \]
The equality in (6.4) holds when \( \theta_{[1]} = \ldots = \theta_{[k-1]} = \theta_{[k]} - \delta^* \). The \( \lambda \) in (6.5) is \( \sqrt{2}t_{v_0, \alpha/2} \).
\[ \Box \]

In order to satisfy the probability requirement (2.2), let \( d \) be the solution of
\[ \int_0^\infty \int_{-\infty}^\infty \Phi^{k-1}(y + dw)d\Phi(y)dG_0(w) = P^*. \quad (6.6) \]
Then \( h = \sqrt{2}t_{v_0, \alpha/2} + d \).
BIBLIOGRAPHY


This paper deals with the problem of selecting the best population from among \( k \geq 2 \) populations which are location-scale models. New selection procedures are proposed for selecting the unique best in terms of the largest location parameter. The procedures include preliminary tests which allow the experimenter to have an option to not select if the statistical evidence is not significant. Two probabilities, the probability to make a selection and the probability of a correct selection, are controlled by these selection procedures. Applications to the normal mean models are considered. Comparisons between the proposed selection procedures and certain earlier existing procedures are also made. Finally, a two-stage procedure for the normal means problem is considered.