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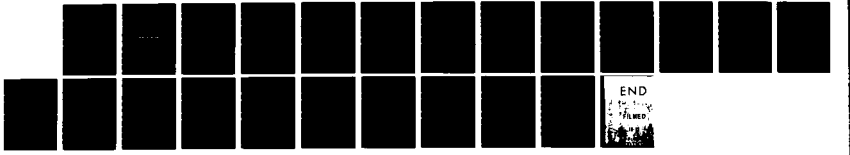

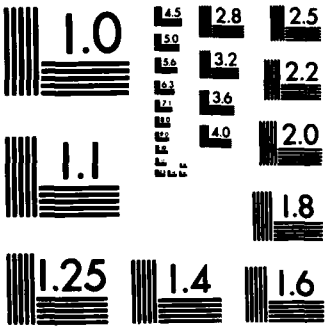

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Klaus J. Miescke

University of Illinois at Chicago

Technical Report #83-39

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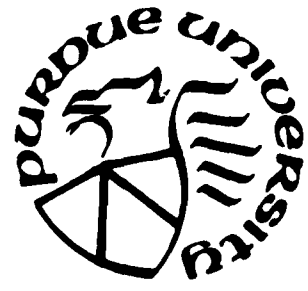




## DEPARTMENT OF STATISTICS

# TWO-STAGE SELECTION PROCEDURES BASED ON TESTS*

by

Klaus J. Miescke

University of Illinois at Chicago

Technical Report #83-39

Department of Statistics
Purdue University

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## Abstract

Suppose that k new treatments have been developed with the purpose of replacing the standard treatment with the best new one, provided that it is actually an improvement on the standard treatment. In a parametric approach, mainly under the assumption of (MLR), procedures are considered which, at a first stage, screen out inferior treatments through statistical tests at a common level of significance. If none (exactly one) is not eliminated, none (this one) will be used as a replacement. Otherwise, if more than one treatment overcomes this screening process, that one of the non-eliminated treatments will be chosen as the replacement which is judged to be the best, after additional data have been observed from the selected treatments. Topics of this paper are the questions of how to choose the terminal decision at the second stage and the tests at the first stage, respectively, and how to implement the appropriate procedures at certain pre-specified performance criteria.

Key Words: Multiple comparisons with a control; 2-stage procedures; screening procedures.

1. Introduction The following procedure, which is used in certain clinical studies, may serve as a motivation for the considerations in this paper. Suppose that k new treatments have been developed with the purpose of replacing the standard treatment with the best new one, provided that it is actually an improvement on the standard treatment. In a pilot study, each new treatment is applied several times and screened out if it is not considered to be significatly better than the standard treatment. Hereby, judgement is gained through suitable statistical tests at a fixed level of $\alpha_0$. If all k new treatments are eliminated the standard treatment will not be replaced. If exactly one new treatment is not eliminated this will be taken as a replacement. In all other cases, the remaining treatments are further examined in a follow up study through additional applications, and finally that one which appears to be the best will be used as a replacement of the standard treatment. The natural questions of how to choose the tests in the first stage and the terminal decision in the second stage are the topic of this paper.

Let $\pi_1,\ldots,\pi_k$ be k populations associated with unknown parameters $\theta_1,\ldots,\theta_k \in \Omega \subseteq \mathbb{R}$. Let $\theta_0 \in \Omega$ be a control value which may be known or unknown. In the latter case, assume that there is also a control population $\pi_0$. A population $\pi_i$ is considered to be better than $\pi_0$ if $\theta_i > \theta_0$, i=1,...,k. The goal is to determine, in two stages, whether there is any population better than the control and, in the affirmative, which one is associated with the largest parameter. Assume that samples $\underline{X}_i = \{X_{ij}\}_{j=1,\ldots,n_i}$, i=0,1,...,k, and $\underline{Y}_j = \{Y_{ij}\}_{j=1,\ldots,m_j}$, i = 1,...,k, can be drawn from $\pi_0,\pi_1,\ldots,\pi_k$ at

the first and at the second stage, respectively, which are mutually independent. Let $\{f_\theta\}_{\theta\in\Omega}$ be a given family of densities with respect to $\mu$, the Lebesgue measure on IR or the counting measure on any lattice in IR, and assume that for every $i\in\{0,1,\ldots,k\}$ all observations from $\pi_i$ have a common distribution with density $f_{\theta_i}$. Later on, after Theorem 1 has been proved, we will make the additional assumption that for every sample $\underline{Z}$ of size n from one population there exists a sufficient statistic $T_n(\underline{Z})$ such that the family of joint densities has nondecreasing likelihood ratios in $T_n$. For notational convenience let them be in the following denoted by $U_i = T_{n_i}(\underline{X}_i)$, $i = 0,1,\ldots,k$, $V_i = T_{m_i}(\underline{Y}_i)$, and $W_i = T_{n_i+m_i}(\underline{X}_i,\underline{Y}_i)$, $i=1,\ldots,k$.

To simplify the presentation the case of a known control value $\theta_0$ will be considered first. Before we define a natural class of two-stage procedures in a concise way, let us briefly describe how these procedures will be typically applied. For every testing problem $H_i\colon \theta_i \le \theta_0$ versus $K_i\colon \theta_i > \theta_0$ the experimenter chooses a test based on $\underline{X}_i$ with a fixed level $\alpha_0$ and another test based on $(\underline{X}_i,\underline{Y}_i)$ with a variable level $\alpha$. At Stage 1 he discards all populations which are not significant at level $\alpha_0$ under the first set of tests. If none (exactly one) is left, he decides that none (this one) is better than the control and is the best population. If more than one population survives he proceeds to Stage 2. At Stage 2, he draws additional samples $\underline{Y}_i$ from those populations which have been selected at Stage 1 and makes a final decision in favor of that population among the selected ones which has the smallest p-value (i.e. is most significant) under the associated second test.

If these tests are upper level tests, which for simplicity may be non-randomized for a moment to fix ideas, based on some real-valued statistics $\tilde{U}_i$ and $\tilde{W}_i$, say, $i = 1,\ldots,k$, then the procedure considered above can be equivalently described as follows: At Stage 1 all $\pi_i$'s are selected with $\tilde{U}_i \geq c_i$ (where $c_i$ is the $\alpha_0$-fractile of $\tilde{U}_i$ under $\theta_i = \theta_0$), and a final decision is made in terms of the largest $\tilde{W}_i$ among the selected $\pi_j$'s. The truncated versions of such procedures (i.e. which perform Stage 1 only) have been studied by several authors, see, for example, Gupta and Sobel (1958) and Lehmann (1961). For further references see Gupta and Panchapakesan (1979) Chapter 20. Some preliminary results concerning two-stage procedures of the type described above in the case of $n_1 = \ldots = n_k$ and $m_1 = \ldots = m_k$ can be found in Gupta and Miescke (1982), which include a comparison with the one-stage analog by Bechhofer and Turnbull (1978).

To begin with, let us point out that several definitions given in Miescke (1979) will be relevant in the sequel but for brevity are not repeated here. Especially, tests may be randomized ones taking values in [0,1]. This typically occurs in discrete cases or in continuous type cases where nonparametric (rank) tests are under concern. Thus significance statements as well as p-values are understood to be based on additional randomization schemes as are used in Miescke (1979). To be more specific, let $\underline{A} = (A_1,\ldots,A_k)$ and $\underline{B} = (B_1,\ldots,B_k)$ be the randomization schemes for the first and the second stage, respectively. Note that the $\underline{X}_i$'s, $\underline{Y}_j$'s, $\underline{A}$ and $\underline{B}$ altogether are assumed to be mutually independent.

The class $\mathcal{A}$ of two-stage procedures. For $i = 1,\ldots,k$, let $\varphi_i = \{\varphi_{i,\alpha}\}_{\alpha\in[0,1]}$ be a right continuous and monotone (in $\alpha$) unbiased test for $H_i$ versus $K_i$ which

is standardized at $\theta_i = \theta_0$. Assume that $\varphi_{i,1} \equiv 1$ outside of the support of the distribution of $\underline{X}_i$ at $\theta_i = \theta_0$. Let $\underline{\varphi}_{\alpha_0} = (\varphi_{1,\alpha_0},\ldots,\varphi_{k,\alpha_0})$ where $0 < \alpha_0 < 1$ is fixed. Analogously, let $\psi_i = \{\psi_{i,\alpha}\}_{\alpha\in[0,1]}$ be such a test for $H_i$ versus $K_i$ based on $(\underline{X}_i,\underline{Y}_i)$. Let $\underline{\psi} = (\psi_1,\ldots,\psi_k)$. Let $\mathcal{S}$ be the class of all procedures of the following type $(\underline{\varphi}_{\alpha_0}, \underline{\psi})$:

Stage 1: Select $\pi_i$ if $p_{\varphi_i}(\underline{X}_i,A_i)$, the p-value of $\underline{X}_i$ under $\varphi_i$, is smaller than $\alpha_0$, $i = 1,\ldots,k$. If none (exactly one) of the populations is selected, stop and decide none (this one) is better than $\pi_0$ and is the best population. Otherwise proceed to Stage 2.

Stage 2: Among the selected populations decide finally in favor of that $\pi_j$ which has the smallest p-value $p_{\psi_j}(\underline{X}_j,\underline{Y}_j,B_j)$ under $\psi_j$.

The following result will prove to be useful in various aspects, except for the important question of how to optimize the component $\underline{\psi}$ in $(\underline{\varphi}_{\alpha_0}, \underline{\psi})$.

Theorem 1. Let $(\underline{\varphi}_{\alpha_0}, \underline{\psi}) \in \mathcal{S}$. For notational convenience, let $E_i = E_{\theta_i}(\varphi_{i,\alpha_0}(\underline{X}_i))$ and $F_i(\alpha) = E_{\theta_i}(\varphi_{i,\alpha_0}(\underline{X}_i)\psi_{i,\alpha}(\underline{X}_i,\underline{Y}_i))$, $\alpha\in[0,1]$, $i = 1,\ldots k$, $\underline{\theta}\in\Omega^k$. Then for every non-empty $D \subseteq \{1,\ldots,k\}$ and $\theta\in\Omega^k$,

(1) $P_{\underline{\theta}}$ {final decision falls into D}

$$= \int_0^1 \prod_{j\notin D} [1-F_j(\alpha)]\, d\Big(1- \prod_{i\in D} [1-F_i(\alpha)]\Big),$$

(2) $P_{\underline{\theta}}$ {final decision is in favor of $\pi_i$ }

$$= \int_0^1 \prod_{\substack{j=1\\ j\neq i}}^{k} [1-F_j(\alpha)]\, dF_i(\alpha), \quad i = 1,\ldots,k,$$

(3) $P_{\underline{\theta}}$ {final decision is made at Stage 1 in favor of $\pi_i$}

$$= \prod_{\substack{j=1 \\ j\neq i}}^{k} [1 - E_j]\, E_i, \qquad i = 1,\ldots,k,$$

(4) $P_{\underline{\theta}}$ {final decision is in favor of the control}

$$= \prod_{j=1}^{k} [1 - E_j].$$

Proof: It has been shown in Miescke (1979) [cf. (2.3) - (2.5) loc. cit.] that the distribution function of each p-value appearing in $(\underline{\varphi}_{\alpha_0}, \underline{\psi})$ equals to the power function of the corresponding test, which is a continuous function of $\alpha \in [0,1]$ at every fixed parameter point, and which at $\alpha = 1$ assumes the value one.

Let now D be a non-empty subset of $\{1,\ldots,k\}$. For $j = 1,\ldots,k$, let $p^*_{\psi_j}(\underline{X}_j, \underline{Y}_j, B_j)$ be equal to $p_{\psi_j}(\underline{X}_j, \underline{Y}_j, B_j)$ if $p_{\varphi_j}(\underline{X}_j, A_j) \leq \alpha_0$, and let it be equal to 1 otherwise. Then it is easy to see that the l.h.s. of (1) is equal to

$$P_{\underline{\theta}} \{ \min_{i\in D} p^*_{\psi_i}(\underline{X}_i, \underline{Y}_i, B_i)\} < \min_{j\notin D} \{p^*_{\psi_j}(\underline{X}_j, \underline{Y}_j, B_j)\}\} .$$

Since for $i \in \{1,\ldots,k\}$ and $\theta_i \in \Omega$, $P_{\theta_i}\{p^*_{\psi_i}(\underline{X}_i, \underline{Y}_i, B_i) \leq \alpha\}$ is equal to $F_i(\alpha)$ if $0 \leq \alpha < 1$, and is equal to 1 if $\alpha = 1$, (1) follows by standard arguments. (2) is a special case of (1) which was stated only because of its relevance in later applications. The verification of (3) and (4) is straightforward and therefore the proof is omitted.

Remark 1. It will be shown in Section 2 that under the assumption of monotone likelihood ratios (MLR), every $(\underline{\varphi}_{\alpha_0}, \underline{\psi}) \in \mathfrak{S}$ is dominated by $(\underline{\varphi}_{\alpha_0}, \underline{\psi}^*)$ if $n_1 + m_1 = \ldots = n_k + m_k$, where $\underline{\psi}^*$ consists of the uniformly most powerful (UMP) tests. Hereby the results of Theorem 1 will not be of great help. There is, however, a particular situation where (1) and (2) can be used for a similar purpose. Suppose that the data of Stage 1 are not available but the information which populations have been significant is a hand. Then one has to use tests at Stage 2 which depend only on the $\underline{Y}_i$'s from the selected populations. In this case every $F_i(\alpha)$ factorizes into the product of the two power functions of $\varphi_{i,\alpha_0}$ and $\psi_{i,\alpha}$, respectively, and therefore (1) and (2) are completely determined through these power functions. For example, if $D_+(\underline{\theta}) = \{i \mid \theta_i > \theta_0, i \in \{1,\ldots,k\}\}$ is not empty, then (1) for $D = D_+(\underline{\theta})$ is maximized by the procedure which uses the UMP-tests at both stages. This is true even if the $n_i + m_i$'s are not assumed to be all equal. Since these and related results in such a special case, however, are considered to be of less statistical importance they will not be discussed in further detail.

In the case of an unknown control parameter $\theta_0$ some obvious changes have to be made. First of all the tests $\varphi_{i,\alpha_0}$ depend now on $(\underline{X}_i, \underline{X}_0)$, $i=1,\ldots,k$, whereas $\underline{\psi}$ remains the same as before. Let $\mathfrak{S}'$ denote the class of two-stage procedures of the type $(\underline{\varphi}_{\alpha_0}, \underline{\psi})$ in this case. The analog of Theorem 1 for $\mathfrak{S}'$ can be attained by replacing the right hand sides of (1) - (4) by their integrals with respect to the distribution of $\underline{X}_0$. If not explicitly stated otherwise, the results to be derived in the sequel for $\mathfrak{S}$ have analogous counterparts for $\mathfrak{S}'$ which will not be formulated or proved for brevity because of the close similarities.

2. Optimality Results For MLR-Families. In the following let $D_+(\underline{\theta}) = \{i|\theta_i > \theta_0,\ i\in\{1,\ldots,k\}\}$ be the "good" populations and $D_-(\underline{\theta}) = \{1,\ldots,k\} \setminus D_+(\underline{\theta})$ be the "bad" ones, $\underline{\theta}\in\Omega^k$. Also let us partition the parameter space $\Omega^k$ into $\Omega^k_- = \{\underline{\theta}\in\Omega^k \mid \theta_i \le \theta_0,\ i = 1,\ldots,k\}$ and its complement $\Omega^k_+$, say. A procedure is said to make a correct selection (CS) at $\underline{\theta}\in\Omega^k_-$ if all populations are eliminated at Stage 1, and it is said to make a correct selection at $\underline{\theta}\in\Omega^k_+$ if a final decision is made in favor of a population with the largest $\theta$-value. Let the goal be now to find a procedure in $\mathcal{S}$ which has a large probability of a correct selection (PCS) on $\Omega^k$. From now on we assume that the family $\{f_\theta\}_{\theta\in\Omega}$ has the MLR-property as specified in Section 1. Then the following partial solution to our problem can be given.

Theorem 2. Let $(\varphi_{\alpha_0}, \underline{\psi}) \in \mathcal{S}$. If $n_1 + m_1 = \ldots = n_k + m_k$ and $\psi_1 = \ldots = \psi_k$ then for all $\underline{\theta}\in\Omega^k$,

$$(5)\qquad \sum_\sigma P_{\sigma(\underline{\theta})}\{\text{CS under } (\varphi_{\alpha_0}, \underline{\psi})\} \le \sum_\sigma P_{\sigma(\underline{\theta})}\{\text{CS under } (\varphi_{\alpha_0}, \underline{\psi}^*)\},$$

where $\underline{\psi}^*$ consists of the UMP-tests for $H_i$ versus $K_i$, $i = 1,\ldots,k$, which in this case are all identical, and where the summation is with respect to all $k!$ permutations of $(1,\ldots,k)$. $\sigma(\underline{\theta}) = (\theta_{\sigma(1)},\ldots,\theta_{\sigma(k)})$.

Proof: Only an outline of the proof will be given since it follows by similar decision theoretic arguments as have been used previously in Gupta and Miescke (1983).

Under the assumptions stated above, the associated decision function of $(\varphi_{\alpha_0}, \underline{\psi})$ which determines final selections at Stage 2 is permutation invariant. The loss function which is implicitly employed is zero if a correct selection

is made, and is one, otherwise. Its component which is associated with final selections at Stage 2 is permutation invariant and favors selections of populations with large parameters.

Let now $\underline{\theta} \in \Omega_+^k$ be fixed and, in a Bayes approach, assume that the unknown parameter vector is random and has a prior distribution which gives equal mass 1/k! to all permutations $\sigma(\underline{\theta})$ of $\underline{\theta}$. Then the posterior distribution of the parameter vector, given $W_i = w_i$, $i = 1,\ldots,k$, has the decreasing in transposition (DT) property. From this fact and the properties of the loss function stated above it follows that the optimal final decision at Stage 2 is the natural one which is made with respect to the non-eliminated population with the largest $W_i$, where ties are broken at random. Clearly this is equivalent to selecting the non-eliminated population with the smallest p-value under test $\psi^*$. The proof is now completed by noting that (5) gives a comparison of the corresponding Bayes risks, where of course at all $\underline{\theta} \in \Omega_-^k$ the probabilities of a correct selection are the same for both procedures.

Corollary 1. If, under the assumptions of Theorem 2, additionally $n_1 = \ldots = n_k$ and $\varphi_{1,\alpha_0} = \ldots = \varphi_{k,\alpha_0}$ is given then for all $\underline{\theta} \in \Omega^k$

$$(6) \qquad P_{\underline{\theta}}\{\text{CS under } (\varphi_{\alpha_0}, \underline{\psi})\} \leq P_{\underline{\theta}}\{\text{CS under } (\varphi_{\alpha_0}, \underline{\psi}^*)\} .$$

Proof: Since both procedures considered here are completely permutation invariant, and since also the 0-1 loss function employed is permutation invariant, their risk functions are symmetric functions of $\underline{\theta} \in \Omega^k$. Therefore

all summands on the l.h.s. of (5) coincide and the same holds for the summands on the r.h.s. of (5).

Remark 2. The proof of Theorem 2 actually applies more generally to the following situation. No matter of how the populations are eliminated at Stage 1, if $n_1 + m_1 = \ldots = n_k + m_k$, and if only permutation invariant final decision functions are admitted at Stage 2, then every Bayes procedure w.r.t. any symmetric prior employs the natural rule at State 2. Of course, also from a non-Bayesian point of view, (5) is an intuitively appealing criterion. It simply reflects the lack of knowledge of how the sample sizes are associated with the k ordered populations parameters. Since there is not even an approximately similar result available in the case of unequal $n_i + m_i$'s, it is strongly recommended to "repair" the design of every experiment with unequal $n_i$'s by choosing the $m_i$'s appropriately to get equal overall sample sizes. Let us assume from now on that $n_1 + m_1 = \ldots = n_k + m_k = N$, say, holds.

Actually, in various selection problems authors have chosen their designs such that the statistics on which the natural final decision rule is based have joint distributions with the DT property. To mention a few relevant examples, Bechhofer (1954), Bechhofer, Dunnett and Sobel (1954), and Dudewicz and Dalal (1975) have done so to be able to implement their procedures at certain specified performance requirements. It may now be added that exactly in these designs the employed natural final decisions are optimal in terms of the risk or the PCS, respectively, uniformly on all parameter configurations.

Lack of the DT property in distributions of statistics used for final

decisions, in cases where this property could be attained in principle, should be considered as pathological designs. Even in the simplest case of a one-stage selection procedure serious difficulties arise, as may be illustrated by the following problem which was emphasized by Bechhofer (1982).

Example. Let $\bar{X}_1,\ldots,\bar{X}_k$ be independent sample means with unknown expectations $\theta_1,\ldots,\theta_k$ and with known but different variance $q_1,\ldots,q_k$ generated from k normal populations. For the problem of selecting a population with the largest $\theta$-value no procedure exists which has a largest PCS, uniformly in $\underline{\theta} \in \mathbb{R}^k$. If k = 2 the natural rule is Bayes w.r.t. independent priors $N(0,q_1)$ and $N(0,q_2)$, respectively, and it can be seen to be admissible under the 0-1 loss function. However, if $k \geq 3$ the natural rule cannot be Bayes with respect to any (multivariate) normal prior. The question of whether it is admissible is still open. On the other hand, for every normal prior with expectation $(0,\ldots,0)$, for which the posterior distribution has the DT -property (which leads to a simple solution) the Bayes rule selects in terms of the largest $\bar{X}_i/q_i$, $i = 1,\ldots,k$.

In the remainder of this paper only procedures of the type $(\varphi_{\alpha_0}, \underline{\psi}^*) \in \mathcal{S}$ will be considered. This is justified in view of the assumption $n_1 + m_1 = \ldots = n_k + m_k = N$ and of Theorem 2. The following result, which generalizes Theorem 3 of Gupta and Miescke (1982), can be used to find least favorable parameter configurations (LFC) of such procedures on suitable subspaces of $\Omega^k$.

Theorem 3. Let $(\varphi_{\alpha_0}, \underline{\psi}^*) \in \mathcal{S}$ where for every $i \in \{1,\ldots,k\}$ the power function of $\varphi_{i,\alpha_0}$ is nondecreasing in $\theta_i$. Then the performance characteristics con-

sidered in Theorem 1 have the following monotonicity properties. (1) is nondecreasing in $\theta_i$, $i \in D$ and nonincreasing in $\theta_j$, $j \notin D$. (2) and (3) are nondecreasing in $\theta_i$ and nonincreasing in $\theta_j$, $j \neq i$. (4) is nonincreasing in $\theta_1,\ldots,\theta_k$.

Proof: The assertions concerning (3) and (4) are obviously true. To prove those concerning (1) and (2) note that for $i \in \{1,\ldots,k\}$, $\underline{\theta} \in \Omega^k$, and $\alpha \in [0,1]$,

$$(7) \qquad F_i(\alpha) = E_{\theta_i}(\varphi_{i,\alpha_0}(\underline{X}_i))\, E_{\theta_i}\{\psi^*_{i,\alpha}(\underline{X}_i,\underline{Y}_i) \mid P_{\varphi_i}(\underline{X}_i,A_i) \leq \alpha_0\}.$$

The first factor on the r.h.s. of (7) is nondecreasing in $\theta_i$ according to the assumptions made above. The second factor can be seen to have the same property by applying Theorem 1 of Simons (1980) which guarantees that for every sample from a MLR-family, conditionally on any proportion of the information in the sample which one might choose to extract, likelihood ratios are still stochastically nondecreasing. Since $\psi^*_{i,\alpha}(\underline{X}_i,\underline{Y}_i)$ is a nondecreasing function of $W_i$, the proof is completed by noting that $A_i$ could be ignored since the arguments apply to every situation $A_i = a_i$, where $a_i \in [0,1]$ is held fixed.

Corollary 2. Under the assumptions of Theorem 3, let $n_1 = \ldots = n_k$ and $\varphi_{1,\alpha_0} = \ldots = \varphi_{k,\alpha_0}$. Then for every $\underline{\theta} \in \Omega^k$ with $\theta_1 \leq \ldots \leq \theta_k$, the probability of a final decision in favor of population $\pi_i$ is nondecreasing in $i \in \{1,\ldots,k\}$.

Proof: Given the assumptions above it can be seen from the proof of Theorem 3 that for $\theta_1 \leq \ldots \leq \theta_k$,

(8) $$F_1(\alpha) \leq F_2(\alpha) \leq \ldots \leq F_k(\alpha),\ \alpha \in [0,1].$$

Therefore the assertion follows from (2) by the same technique which was used in (IV) of Miescke (1979).

Theorem 4. Let $(\varphi_{\alpha_0}, \underline{\psi}^*) \in \mathcal{D}$ where $\varphi_{\alpha_0}$ consists of consistent tests. Then for increasing sample sizes $n_i$ and $m_i = N - n_i$, $i = 1,\ldots,k$, the probability of a correct selection tends to one at all $\underline{\theta} \in \Omega_+^k$ and at all $\underline{\theta} \in \Omega_-^k$ with $\theta_i < \theta_0$, $i = 1,\ldots,k$.

Proof: If $\underline{\theta} \in \Omega_+^k$ with exactly one coordinate greater than $\theta_0$, or if $\underline{\theta} \in \Omega_-^k$ with $\theta_i < \theta_0$, $i = 1,\ldots,k$, the assertion follows immediately from (3) and (4), respectively. For all other $\underline{\theta} \in \Omega_+^k$, the probability that all populations $\pi_i$ with $\theta_i > \theta_0$ will not be eliminated at Stage 1 tends to one. Moreover, $P_{\underline{\theta}}\{W_i > W_j \text{ for all } j \neq i\}$ also tends to one if $\theta_i$ is the unique maximum of $\theta_1,\ldots,\theta_k$. This can be seen as follows. Selecting in terms of the largest $W_j$ is equivalent to selecting in terms of the smallest p-value under tests $\tilde{\psi}_j$ which are essentially the same tests as $\psi_j^*$ but now standardized at $\theta_i$. By Theorem 2 of Miescke (1979) it follows that

(9) $$P_{\underline{\theta}}\{W_i > W_j \text{ for all } j \neq i\} = \int_0^1 \prod_{j \neq i} [1 - E_{\theta_j}(\tilde{\psi}_{j,\alpha}(\underline{X}_j, \underline{Y}_j))]d\alpha,$$

which now can be seen to tend to one if N tends to infinity. If, however, all the good populations are not eliminated at Stage 1 and $W_i > W_j$, for all $j \neq i$, then a correct selection is made. This completes the proof in the given parameter configuration. The case of more than one best population can be treated similarly.

Focussing now on the first component $\underline{\varphi}_{\alpha_0}$ in the procedures $(\underline{\varphi}_{\alpha_0}, \underline{\psi}^*) \in \mathcal{D}$, the natural choice is of course $\underline{\varphi}^*_{\alpha_0}$ which consists of the corresponding UMP-tests for $H_i$ versus $K_i$ based on $\underline{X}_i$, or more precisely, based on $U_i, i = 1,\ldots,k$. Even though such a choice cannot be justified (not even in the case of $n_1 = \ldots = n_k$) by an overall improvement on the PCS, several strong reasons can be quoted in support of choosing $(\underline{\varphi}^*_{\alpha_0}, \underline{\psi}^*)$. First, of course, all results derived hitherto hold for this procedure. Second, the following can be stated.

Theorem 5. Among all procedures in $\mathcal{D}$, $(\underline{\varphi}^*_{\alpha_0}, \underline{\psi}^*)$ maximizes

(10) $P_{\underline{\theta}}\{CS\}$ at every $\underline{\theta} \in \Omega^k$,

(11) $P_{\underline{\theta}}\{CS \text{ at Stage 1}\}$ at every $\underline{\theta} \in \Omega^k_+$ with exactly one $\theta_i > \theta_0$,

(12) $E_{\underline{\theta}}$ (number of good populations selected at Stage 1) at every $\underline{\theta} \in \Omega^k_+$,

(13) $E_{\underline{\theta}}$ (number of bad populations eliminated at Stage 1) at every $\underline{\theta} \in \Omega^k$.

Proof: (10) follows from (4) and (11) follows from (3). All arguments are standard and are based on the well known properties of the power functions of the UMP-tests $\varphi^*_{i,\alpha_0}, i = 1,\ldots,k$. Therefore, no further details will be given.

Third, all permutation invariant procedures $(\underline{\varphi}_{\alpha_0}, \underline{\psi}^*)$, except $(\underline{\varphi}^*_{\alpha_0}, \underline{\psi}^*)$ ifself, can be modified, without changing the sizes of the selected subsets of populations at Stage 1, in a specific way which leads to an improvement on the PCS, provided that the family $\{f_\theta\}_{\theta \in \Omega}$ is a strongly unimodal exponential family. It should be noted that the modified procedure is also based on tests but is no longer a member of the class $\mathcal{D}$. More precisely, from Corollary 2 in Gupta and Miescke (1983) the following result can be derived.

Theorem 6. Let $\{f_\theta\}_{\theta \in \Omega}$ be a strongly unimodal (i.e. log-concave) exponential family. Then every $(\underline{\varphi}_{\alpha_0}, \underline{\psi}^*) \in \mathcal{S}$ with $n_1 = \ldots = n_k$ and $\varphi_{1,\alpha_0} = \ldots = \varphi_{k,\alpha_0}$ can be improved by simply replacing the selected populations at Stage 1 by the same number of populations, but now by those which are associated with the largest $U_i$'s, where ties are broken at random. Then for all $\underline{\theta} \in \Omega^k$,

$$(14) \qquad P_{\underline{\theta}} \{\text{CS under } (\underline{\varphi}_{\alpha_0}, \underline{\psi}^*)\} \leq P_{\underline{\theta}} \{\text{CS under the modified procedure}\}.$$

3. Applications With Illustrations In The Normal Case. In applications the procedure $(\underline{\varphi}^*_{\alpha_0}, \underline{\psi}^*)$ usually will be implemented as to meet certain performance requirements. This will be described in this section and will be illustrated by the example of k normal populations $N(\theta_i, \sigma_i^2)$, $i = 1,\ldots,k$, where $U_i$ and $W_i$ are the corresponding sample means, $i = 1,\ldots,k$. Here the procedure can be considered to be the two-stage analog of the one-stage procedure by Bechhofer and Turnbull (1978). At first consider the basic requirement

$$(15) \qquad \inf \{ P_{\underline{\theta}} \{\text{CS under } (\underline{\varphi}^*_{\alpha_0}, \underline{\psi}^*) \mid \underline{\theta} \in \Omega^k_- \} = P_0^*,$$

where $P_0^*$ is a predetermined constant. In view of (4) this can be accomplished by choosing $\alpha_0$ to satisfy

$$(16) \qquad (1 - \alpha_0)^k = P_0^*.$$

Then in the normal case the procedure is of the following form. At Stage 1, population $\pi_i$ is selected if $n_i^{1/2} (U_i - \theta_0) / \sigma_i \geq \Phi^{-1}(1 - \alpha_0)$, $i = 1,\ldots,k$, where $\Phi$ denotes the c.d.f. of N(0,1). And at Stage 2, a final decision is made in terms of the largest $W_j$ from the selected populations.

Since (15) actually involves only the properties of the procedure at the

first stage, the $P_0^*$ - condition can be attained by employing techniques used for one-stage procedures. Thus (15) can be solved by taking recourse to relevant papers in this area. For further details see Gupta and Panchapakesan (1979).

A second requirement will typically employ the indifference zone approach which is due to Bechhofer (1954). Let $\Delta > 0$ be fixed and let $\Omega_\Delta^k = \{ \underline{\theta} \in \Omega^k \mid \underline{\theta}_i + \Delta \geq \theta_0, \theta_1, \ldots, \theta_{i-1}, \theta_{i+1}, \ldots \theta_k \text{ for some } i\}$. Now consider the requirement

$$\text{(16)} \qquad \inf \{P_{\underline{\theta}} \{\text{CS under } (\underline{\varphi}^*_{\alpha_0}, \underline{\psi}^*)\} \mid \underline{\theta} \in \Omega_\Delta^k\} = P_1^* ,$$

where $P_1^*$ is a second predetermined constant. Even though Theorem 3 can be used to find the LFC, it is technically too difficult to attain (16) exactly. Therefore, the following conservative approach, which over-protects the experimenter with respect to (16), is recommended and is easy to perform. Let $\underline{\theta} \in \Omega_\Delta^k$ with $\theta_i = \max \{\theta_1, \ldots, \theta_k\}$, say. If population $\pi_i$ is selected at Stage 1 and $W_i$ is the unique maximum of $W_1, \ldots, W_k$ then a correct selection is made. Therefore if the following two conditions are fulfilled, and $\beta_1$ and $\beta_2$ are chosen to meet $\beta_1 + \beta_2 - 1 = P_1^*$ then by Bonferroni's inequality it follows that the l.h.s. of (16) is not smaller than $P_1^*$. The conditions are

$$\text{(17)} \qquad E_{\theta_0 + \Delta}(\varphi^*_{j,\alpha_0}(\underline{X}_j)) \geq \beta_1, \; j = 1, \ldots, k, \text{ and}$$

$$\text{(18)} \qquad \inf \{ P_{(\theta+\Delta, \theta, \ldots, \theta)} \{W_1 > W_2, \ldots, W_k\} \mid \theta \in \Omega \} \geq \beta_2 .$$

In the normal case it is well known (cf. Tamhane and Bechhofer (1979)) that Slepian's inequality leads to better results than Bonferroni's inequality.

Use of the former allows to choose $\beta_1$ and $\beta_2$ according to the condition $\beta_1\beta_2 = P_1^*$ which is preferable since $\beta_1 + \beta_2 - 1 < \beta_1\beta_2$ for $0 < \beta_1, \beta_2 < 1$.

To meet (17), standard techniques from the theory of testing hypotheses can be used. And (18) can be attained by using results of single-stage selection procedures in the indifference zone approach due to Bechhofer (1954).

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Suppose that k new treatments have been developed with the purpose of replacing the standard treatment with the best new one, provided that it is actually an improvement on the standard treatment. In a parametric approach, mainly under the assumption of MLR, procedures are considered which, at a common level of significance $\alpha_0$. If none (exactly one) is not eliminated, none (this one) will be used as a replacement. Otherwise, if more than one treatment overcomes this screening process, that one of the non-eliminated

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treatments will be chosen as the replacement which is judged to be the best, after additional data have been observed from the selected treatments. Topics of this paper are the questions of how to choose the terminal decision at the second stage and the tests at the first stage, respectively, and how to implement the appropriate procedures at certain pre-specified performance criteria.

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