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ADMISSIBLE MINIMAX ESTIMATION OF A MULTIVARIATE NORMAL MEAN WITH ARBITRARY QUADRATIC LOSS

BY JAMES O. BERGER

Purdue University

The problem of estimating the mean of a p-variate ($p \geqq 3$) normal distribution is considered. It is assumed that the covariance matrix Σ is known and that the loss function is quadratic. A class of minimax estimators is given, out of which admissible minimax estimators are developed.

1. Introduction. Let $X = (X_1, \cdots, X_p)^t$ be an observation from a p-variate normal population with mean vector $\theta = (\theta_1, \cdots, \theta_p)^t$ and known positive definite covariance matrix Σ. Assume that the loss incurred in estimating θ by $\delta(X) = (\delta_1(X), \cdots, \delta_p(X))^t$ is the quadratic loss $(\delta - \theta)^t Q(\delta - \theta)$. Assume also that Q is positive definite and that $p \geqq 3$.

The special situation $Q = \Sigma^{-1}$ has been considered by several authors (James and Stein (1960), Baranchik (1970), Strawderman (1971) among them), and wide classes of minimax and admissible estimators of θ have been found. Results for arbitrary Q and Σ, however, are incomplete. Bhattacharya (1966) and Bock (1975) found some particular minimax estimators for the general situation. In this paper, a different and simpler class of minimax estimators is given, out of which admissible minimax estimators are developed.

2. A class of minimax estimators. Define $||x||^2 = x^t\Sigma^{-1}Q^{-1}\Sigma^{-1}x$, and let I denote the $p \times p$ identity matrix. Estimators of the form

$$\delta(X) = (I - r(||X||^2)Q^{-1}\Sigma^{-1}/||X||^2)X \tag{1}$$

will be considered, where r is a measurable function from $R^1 \to R^1$.

THEOREM 1. *The estimator δ, given by* (1), *is minimax if*

(i) $0 \leqq r(\cdot) \leqq 2(p - 2)$, *and*
(ii) $r(\cdot)$ *is nondecreasing.*

PROOF. This theorem has since been extended to a more general theorem, dealing with a wide class of densities, in Berger (1975). The above theorem is a special case of Theorem 3 in that paper. □

It has been brought to the author's attention that the above result was independently discovered by Malcolm Hudson (1974).

3. Admissible minimax estimators. The obvious question which arises is how

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should the function r, in (1), be chosen? In this section, choices of r which give rise to admissible minimax estimators are developed.

Let α denote the smallest characteristic root of the matrix ΣQ. The following choices of r will be considered:

$$r_c(t) = \frac{t \int_0^\alpha \lambda^{(p/2-c+1)} \exp\{-\lambda t/2\}\, d\lambda}{\int_0^\alpha \lambda^{(p/2-c)} \exp\{-\lambda t/2\}\, d\lambda}\,, \qquad c < 1 + p/2\,. \tag{2}$$

A simple integration by parts in the numerator of the above expression gives

$$r_c(t) = (p - 2c + 2) - \frac{2\alpha^{(p/2-c+1)} \exp\{-\alpha t/2\}}{\int_0^\alpha \lambda^{(p/2-c)} \exp\{-\lambda t/2\}\, d\lambda}\,. \tag{3}$$

Note that when $Q = \Sigma^{-1}$ (and hence $\alpha = 1$), r_c gives rise to the admissible minimax estimator found in Strawderman (1971).

THEOREM 2. *Assume that δ is given by* (1), *with $r = r_c$.*

(a) *If $3 - p/2 \leqq c < 1 + p/2$, then δ is minimax.*
(b) *If $3 - p/2 \leqq c < 2$, then δ is admissible.*
(c) *If $3 - p/2 \leqq c < 1$, then δ is proper Bayes.*

PROOF. To prove (a), it is only necessary to verify conditions (i) and (ii) of Theorem 1. From (2), it is clear that $r_c(\cdot) > 0$. Using (3) and the assumption that $c \geqq 3 - p/2$, it is also clear that $r_c(\cdot) < 2(p - 2)$. Hence condition (i) of Theorem 1 is satisfied. From (3) and the fact that $\exp\{(\alpha - \lambda)t/2\}$ is nondecreasing in t for $0 \leqq \lambda \leqq \alpha$, it follows that $r_c(t)$ is nondecreasing in t. Condition (ii) of Theorem 1 is thus satisfied and the conclusion follows.

To prove (b) and (c), δ must first be shown to be a generalized Bayes estimator. The notation will be considerably simplified by considering only the case $Q = I$ and $\Sigma = A$, where A is a $p \times p$ diagonal matrix with diagonal elements $a_i > 0$. Since Q and Σ are positive definite, it is easy to check that the problem can always be transformed into this diagonal case. Note that $||X||^2 = \sum X_i^2/a_i^2$ and that $\alpha = \min\{a_i\}$. For notational convenience, define $b_i(\lambda) = a_i(a_i - \lambda)/\lambda$.

For $c < 1 + p/2$, consider the generalized prior density

$$g_c(\theta) = \int_0^\alpha \left[\prod_{i=1}^p b_i(\lambda)^{-\frac{1}{2}}\right] \exp\{-\tfrac{1}{2} \sum_{i=1}^p \theta_i^2/b_i(\lambda)\}\lambda^{-c}\, d\lambda\,. \tag{4}$$

It is easy to check that $g_c(\cdot)$ is a bounded function for the given choice of c. (Clearly $b_i(\lambda)$ behaves like a_i^2/λ near $\lambda = 0$.) Note also that g_c has finite mass if $c < 1$.

The generalized Bayes estimator of θ, with respect to g_c, is given componentwise by

$$\delta_i^c(X) = \frac{\int \theta_i \exp\{-\frac{1}{2} \sum_{i=1}^p (X_i - \theta_i)^2/a_i\} g_c(\theta)\, d\theta}{\int \exp\{-\frac{1}{2} \sum_{i=1}^p (X_i - \theta_i)^2/a_i\} g_c(\theta)\, d\theta}\,. \tag{5}$$

Consider first the numerator of the above expression. Using the definition of $g_c(\theta)$, interchanging orders of integration (g_c is a bounded function), and completing squares gives

$$(6) \qquad \int \theta_i \exp\{-\tfrac{1}{2} \textstyle\sum_{i=1}^p (X_i - \theta_i)^2/a_i\} g_c(\theta)\, d\theta = \int_0^\alpha \int_{R^p} \theta_i \exp\{\varphi\}[\prod_{j=1}^p b_j(\lambda)^{-\frac{1}{2}}]\lambda^{-c}\, d\theta\, d\lambda\,,$$

where

$$\varphi = -\tfrac{1}{2} \textstyle\sum_{j=1}^p ([a_j^{-1} + b_j(\lambda)^{-1}][\theta_j - X_j/\{1 + a_j b_j(\lambda)^{-1}\}]^2 + X_j^2/\{a_j + b_j(\lambda)\})\,.$$

Integrating out over θ, on the right hand side of (6), gives that the numerator of (5) equals

$$\int_0^\alpha [X_i/\{1 + a_i b_i(\lambda)^{-1}\}][\prod_{j=1}^p \{1 + a_j^{-1} b_j(\lambda)\}^{-\frac{1}{2}}] \exp\{-\tfrac{1}{2} \textstyle\sum_{j=1}^p [X_j^2/\{a_j + b_j(\lambda)\}]\}\lambda^{-c}\, d\lambda\,.$$

Using the identities

$$a_j + b_j(\lambda) = a_j + a_j(a_j - \lambda)/\lambda = a_j^2/\lambda\,,$$
$$\{1 + a_j b_j(\lambda)^{-1}\}^{-1} = 1 - \lambda/a_j\,,$$
$$\{b_j(\lambda)a_j^{-1} + 1\}^{-\frac{1}{2}} = (\lambda/a_j)^{\frac{1}{2}}\,,$$

it is thus clear that the numerator of (5) equals

$$(7) \qquad \int_0^\alpha (1 - \lambda/a_i) X_i \exp\{-\lambda \|X\|^2/2\}\lambda^{(p/2-c)}[\prod_{j=1}^p a_j^{-\frac{1}{2}}]\, d\lambda\,.$$

It can similarly be shown that

$$(8) \qquad \int \exp\{-\tfrac{1}{2} \textstyle\sum_{i=1}^p (X_i - \theta_i^2)/a_i\} g_c(\theta)\, d\theta = \int_0^\alpha \exp\{-\lambda\|X\|^2/2\}\lambda^{(p/2-c)}[\prod_{j=1}^p a_j^{-\frac{1}{2}}]\, d\lambda\,.$$

Combining (5), (7), (8), and (2) gives

$$\delta_i^c(X) = \frac{\int_0^\alpha (1 - \lambda/a_i) X_i \exp\{-\lambda\|X\|^2/2\}\lambda^{(p/2-c)}\, d\lambda}{\int_0^\alpha \exp\{-\lambda\|X\|^2/2\}\lambda^{(p/2-c)}\, d\lambda} = [1 - r_c(\|X\|^2)/(a_i\|X\|^2)]X_i\,.$$

Thus $\delta = \delta^c$ is indeed the generalized Bayes estimator with respect to g_c. Part (c) of the theorem follows immediately from this and the observation that g_c has finite mass if $c < 1$.

To prove part (b) of the theorem, a result from Brown (1971) will be used. Note first that (8) and a change of variables give

$$\begin{aligned} f^*(X) &= \int_{R^p} [\prod_{i=1}^p (2\pi a_i)^{-\frac{1}{2}}] \exp\{-\tfrac{1}{2} \textstyle\sum_{i=1}^p (X_i - \theta_i)^2/a_i\} g_c(\theta)\, d\theta \qquad \text{(definition)} \\ &= (2\pi)^{-p/2} [\prod_{i=1}^p a_i^{-1}] \int_0^\alpha \exp\{-\lambda\|X\|^2/2\}\lambda^{(p/2-c)}\, d\lambda \\ &= (2\pi)^{-p/2} [\prod_{i=1}^p a_i^{-1}] \|X\|^{-(p-2c+2)} \int_0^{\alpha\|X\|^2} \exp\{-\lambda/2\}\lambda^{(p/2-c)}\, d\lambda \\ &\leqq K|X|^{(2c-p-2)}\,. \end{aligned}$$

(Here $|X|$ denotes the usual Euclidean norm of X.) Using Corollary 4.3.4 and Theorem 5.1.1 (B) of Brown (1971), together with the assumption that $c < 2$, it can be concluded that δ is admissible. □

Note that for the estimator of Theorem 2 to be minimax, it is necessary to have $p \geqq 3$. Clearly admissible, minimax estimators of the given form do exist for $p \geqq 3$. Proper Bayes versions, however, exist only if $p \geqq 5$.

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MATHEMATICAL SCIENCES BUILDING
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