and where,

$$\bar{x} = \sum_{i=1}^{n} x_i/n,$$

$$s^2 = \sum_{i=1}^{n} (x_i - \bar{x})^2/(n - 1),$$

and

$$\sqrt{n}k = t(n - 1, u_\alpha\sqrt{n}, 1 - P).$$

Here $t(f, \delta, \epsilon)$ [1] is the $100\epsilon$ percentage point of the non-central $t$ distribution with $f$ degrees of freedom, $\delta$ is the measure of non-centrality in the definition of $t$, and $u_\alpha$ is the $100(1 - \alpha)$ percentage point of the unit normal distribution with zero mean.

By use of the tables (especially Table IV) and iteration of the approximations given by Johnson and Welch in [1] the authors obtain values of the coefficient $\sqrt{n}k$ to 4S, for $n = 5(1)20(5)50(10)\ 100(100)\ 300$, for $P$ and $\alpha = .90, .95, .99$. A method for determination of these coefficients is given in [1], but the calculations are, of course, quite tedious, so that the present tables render a valuable service for practical applications to one-sided tolerance limits.

S. B. LITTAUER

Columbia University
New York, N. Y.

1. N. S. JOHNSON & B. L. WELCH, "Applications of the non-central $t$-distribution," *Biometrika*, v. 31, 1939, p. 362–389.

**75[K].**—J. PACHARES, "Tables of the upper 10% points of the Studentized range," *Biometrika*, v. 46, 1959, p. 461–466.

Let $q = w/s$, where $w$ is a sample range based on $n$ values, and $s$ is an independent estimate of standard deviation based on $m$ values. Then tables of $q'$ have been prepared for $Pr(q \geqq q') = \alpha$, where $\alpha = .01, .05, .10$, $n = 2\ (1)\ 20$, and $m = 1\ (1)\ 20, 24, 30, 40, 60, 120, \infty$. Three significant figures are given throughout. The work of Harter [1] has been used in improving the accuracy throughout, particularly for $\alpha = .01$. For $\alpha = .01$ and $.05$, these tables correct errors in [2].

I. R. SAVAGE

University of Minnesota
Minneapolis, Minn.

1. H. L. HARTER, *The Probability Integrals of the Range and of the Studentized Range*, WADC Technical Report 58-484, vols. I & II, 1959. Office of Technical Services, U. S. Dept. of Commerce, Washington 25, D.C.
2. J. M. MAY, "Extended and corrected tables of the upper percentage points of the 'Studentized' range," *Biometrika*, v. 39, 1952, p. 192–193. [**RMT** 1080, *MTAC*, v. 7, 1953, p. 94]

**76[K].**—K. C. S. PILLAI & PABLO SAMSON, JR., "On Hotelling's generalization of $T^2$," *Biometrika*, v. 46, 1959, p. 160–168.

Let $S_1/n_1$, $S_2/n_2$ denote independent covariance matrices arising from samples of sizes $n_1$ and $n_2$ from two $p$-variate normal populations, and $U^{(s)} = \text{trace } S_2^{-1}S_1$, where $s$ is the number of non-zero roots. Two approximations are compared with the exact values for the upper 5 and 1 percentage points of $U^{(2)}$ for several values of $m = (n_1 - s - 2)/2$ and $n = (n_2 - s - 2)/2$. The approximations for the upper 5 and 1 percentage points of $U^{(3)}$ and $U^{(4)}$ are given to 3 or 4D for $m = 0, 5$, $n = 15(5)50, 60(20)100$.

I. Olkin

University of Minnesota
Minneapolis, Minn.

**77[K].**—J. G. Saw, "Estimation of the normal population parameters given a singly censored sample," *Biometrika*, v. 46, 1959, p. 150–159.

As estimators of the mean and variance of a normal distribution, given an ordered sample $x_1 < x_2 \cdots < x_n$ censored above $x_r$, the author proposes

$$\mu^* = \bar{x}_{r-1} + (1 - \epsilon)x_r, \quad \text{where} \quad \bar{x}_{r-1} = \sum_{i=1}^{r-1} x_i/(r - 1),$$

$$\eta^* = \alpha \sum_{i=1}^{r-1} (x_i - x_r)^2 + \beta \sum_{i=1}^{r-1} (x_i - x_r)^2,$$

respectively, where $\epsilon$ is chosen to make $\mu^*$ unbiased, and $\alpha$ and $\beta$ are chosen to make $\eta^*$ unbiased and of minimum variance. To facilitate use of these estimators, three tables are appended. Table 1 consists of entries of the weight factor $\epsilon$ and Var $(\mu^*/\sigma)$ to 10D for $1 < r < n \leqq 20$. Table 2 contains coefficients of $(n + 1)^{-i}$ in series approximations to $\epsilon$ and to Var $(\mu^*/\sigma)$. Weight factors $\alpha$ and $\beta$ are not tabulated directly, and consequently routine application of the author's estimates may be hampered. However, in order to permit calculation of these factors, Table 3, containing coefficients of $(n + 1)^{-i}$ in series approximations to them, has been included. These entries are given to 6D for $p_r = .50(.05).80$, where $p_r = r/(n + 1)$.

A. C. Cohen, Jr.

University of Georgia
Athens, Georgia

**78[K].**—Minoru Siotani, "The extreme value of the generalized distances of the individual points in the multivariate normal sample," *Ann. Inst. Statist. Math. Tokyo*, v. 10, 1959, p. 183–208.

Let $x_\alpha' = (x_{1\alpha}, \cdots, x_{p\alpha})$, $\alpha = 1, \cdots, n$, be $n$ independent observations from a $p$-variate normal population with mean vector $m' = (m_1, \cdots, m_p)$ and covariance matrix $\Lambda$, and let $\bar{x}' = (\bar{x}_1, \cdots, \bar{x}_p)$. The upper 5, $2\frac{1}{2}$, and 1 percentage points of the extreme deviate $\hat{\chi}^2_{\max D} = \max_i[(x_i - \bar{x})'\Lambda^{-1}(x_i - \bar{x})]$ is given to 2D for $n = 3(1)10(2)20(5)30$, $p = 2, 3, 4$. When $\Lambda$ is unknown, let $L$ be a $p \times p$ matrix whose elements, $l_{ij}$, are unbiased estimates of $\lambda_{ij}$, and have a Wishart distribution with $\nu$ degrees of freedom. The upper 5, $2\frac{1}{2}$, and 1 percentage points of the Studentized extreme deviate $\hat{T}^2_{\max D} = \max^i [(x_i - \bar{x})'L^{-1}(x_i - \bar{x})]$ is given to 2D for $n = 3(1)12, 14$, $\nu = 20(2)40(5)\ 60, 100, 150, 200$.

I. Olkin