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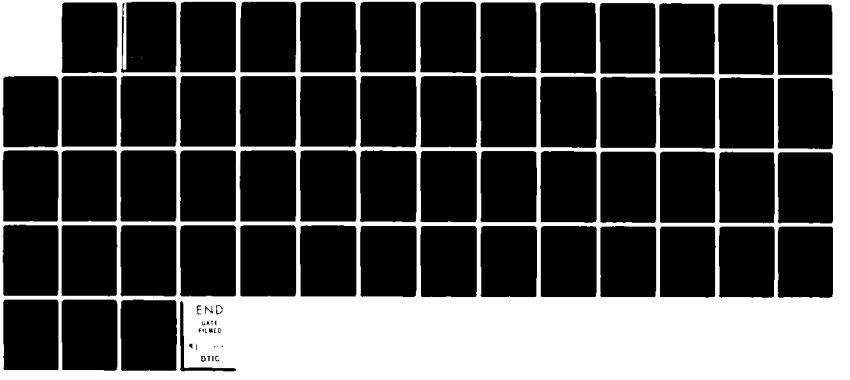

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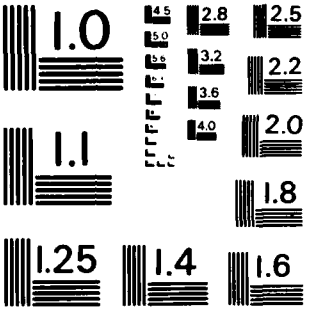

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Marilyn S. Wingersky

and

Frederic M. Lord

This research was sponsored in part by the Personnel and Training Research Programs Psychological Sciences Division Office of Naval Research, under Contract No. N00014-80-C-0402

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20. Abstract (Continued)

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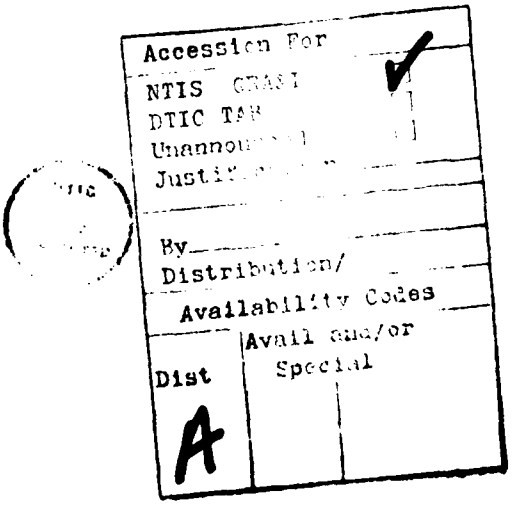

An Investigation of Methods for Reducing Sampling Error in Certain IRT Procedures*

In IRT until now, the sampling variances and covariances for maximum likelihood estimates of item parameters have usually been computed by assuming the abilities to be known; the sampling variances and covariances for ability estimates were computed by assuming the item parameters to be known. In this paper, a suggested method for computing the sampling variance-covariance matrix when all parameters are unknown (lord and Wingersky, 1983) will be used to try to answer various practical questions. Section 2 presents needed additional, though not conclusive, evidence that the new method for computing the variance-covariance matrix yields correct results. Section 3 investigates the effect of changing the number of items or the number or distribution of people on the standard errors of the item parameters and of the abilities. Section 4 presents a technique for displaying and understanding the standard errors and sampling covariances of estimates of item parameters.

Section 5 deals with the practically important situation where we have two tests that contain a set of items in common and these tests are administered to two separate groups of examinees. A problem in item

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banking or test equating is putting the parameter estimates for the two tests on a common scale. One way to do this is to estimate all of the parameters for both tests in one calibration run. When this is done, how does the number and quality of the common items affect the standard errors of the parameter estimates for the unique (noncommon) items?

1. Preliminaries

The three-parameter Birnbaum logistic model is used throughout. The probability of examinee a answering item i correctly is

$$P_{ia} = c_i + (1 - c_i)/(1. + \exp(-1.7a_i(\theta_a - b_i))) \qquad (1)$$

where a_i is the discrimination of item i; b_i is the difficulty for the item, c_i is the lower asymptote of the item response function, and θ_a is the ability for examinee a. In a typical calibration run, poorly estimatable c_i are ordinarily fixed at some common value. In this paper, however, all c_i are considered unknown and must be estimated. In treating all of the c_i as unknown we are looking at the "worst case" standard errors.

In IRT, the origin and unit of measurement of the ability scale is arbitrary. Until this scale is specified all parameters except the c_i are unidentifiable. The origin and unit of the ability scale must be specified in terms of (as a function of) the true parameters. If the

origin and unit of the ability scale were specified in terms of the parameter estimates, then the true parameters would be undefined. Since the true parameters are unknown but depend on the scale used, this means that the scale origin and the scale unit (each defined as a function of the true parameters) must be estimated from the data. The estimated origin and scale unit are obviously subject to sampling errors, which affect the accuracy of all parameter estimates. It is therefore important to define the origin and unit each by a function of parameters that can be estimated with good accuracy.

The scale recommended in Lord and Wingersky (1983) and used here requires that the mean of the difficulty parameters of certain selected items be 0 (the origin) and that the difference between two such means for two sets of selected items be 1 (the scale unit). This scale will be referred to as the "capital" scale: parameters on this scale will be denoted by the capital letters A_i , B_i , C_i , Θ_a . The "small" scale or the "LOGIST" scale, referred to by lower-case letters, is the scale used by the LOGIST program (Wingersky, Barton, and Lord (1982)), the computer program used here for estimating the parameters of (1) by maximum likelihood. LOGIST sets a truncated mean of the estimated abilities to 0 and a truncated standard deviation of the estimated abilities to 1. The following formulas convert the parameters from the LOGIST scale to the capital scale:

$$\Theta_a = (\theta_a - \bar{b}_0)/k \quad ,$$

$$k = \bar{b}_1 - \bar{b}_0 \quad ,$$

$$A_i = ka_i \quad ,$$

$$B_i = (b_i - \bar{b}_0)/k \quad ,$$

$$C_i = c_i \quad ,$$

where $\bar{b}_0$ and $\bar{b}_1$ are means of the b_i for two selected subsets of items. The capital scale is a linear transformation of the LOGIST scale. The c_i are not affected by the scale.

2. Variance of p_i, the Proportion Correct

If we could prove that the maximum likelihood parameter estimates for the Birnbaum model are consistent when all item and ability parameters are estimated simultaneously, the sampling variance-covariance matrix described in Lord and Wingersky (1983) would be the correct one to use. Since consistency has not yet been proven mathematically any results that confirm the appropriateness of this variance-covariance matrix makes one feel more comfortable about using it.

The sampling variance of p_i, the proportion of examinees in the sample who answer item i correctly, can be computed directly from familiar standard formulas; it can also be computed with some effort from the sampling variance-covariance matrix obtained by Lord and Wingersky (1983). These two methods should give the same results if the Lord-Wingersky matrix is correct.

The usual likelihood equations for $\hat{b}_i$ and for $\hat{c}_i$, obtained by setting the derivative of the likelihood function equal to zero, are (Lord, 1980, eq. 12.1 and 12.2)

$$\sum_{a=1}^{N} (u_{ia} - \hat{P}_i(\hat{\theta}_a))(\hat{P}_i(\hat{\theta}_a) - \hat{c}_i)/\hat{P}_i(\hat{\theta}_a)) = 0 \quad , \tag{2}$$

$$\sum_{a=1}^{N} (u_{ia} - \hat{P}_i(\hat{\theta}_a))/\hat{P}_i(\hat{\theta}_a) = 0 \quad , \tag{3}$$

where u_{ia} is the score (0 or 1) of examinee a on item i, N is the number of examinees, and a caret denotes substitution of parameter estimates for true parameter values. Multiplying (3) by c_i, adding to (2), and transposing gives

$$\sum_{a=1}^{N} \hat{P}_i(\hat{\theta}_a) = \sum_{a=1}^{N} u_{ia} \quad .$$

Since

$$p_i = \frac{1}{N} \sum_{a=1}^{N} u_{ia} \quad , \tag{4}$$

we have

$$p_i = \frac{1}{N} \sum_{a=1}^{N} \hat{P}_i(\hat{\theta}_a) \tag{5}$$

From (4) and (5), we can derive two separate formulas for the variance of p_i.

For some group of examinees whose abilities are specified by the vector $\underset{\sim}{\theta} \equiv \{\theta_1, \theta_2, \ldots, \theta_N\}$, we have from (4) that

$$\text{var}(\hat{p}_i | \underset{\sim}{\theta}) = \frac{1}{N^2} \sum_{a=1}^{N} \sum_{a'=1}^{N} \text{cov}(u_{ia}, u_{ia'} | \underset{\sim}{\theta}) \quad ,$$

$$= \frac{1}{N^2} \sum_{a=1}^{N} \text{var}(u_{ia} | \underset{\sim}{\theta}) \quad ,$$

$$= \frac{1}{N^2} \sum_{a=1}^{N} P_i(\theta_a) Q_i(\theta_a) \quad , \tag{6}$$

with

$$Q_i(\theta_a) = 1 - P_i(\theta_a) \quad ,$$

since $\text{cov}(u_{ia}, u_{ia'} | \underset{\sim}{\theta}) = 0$ when $a \neq a'$. Similarly,

$$\text{cov}(p_i, p_j | \underset{\sim}{\theta}) = 0 \quad . \tag{7}$$

By the formula for the covariance between two sums, we have from (5) for the same group of examinees that

$$\mathrm{var}(p_i|\underset{\sim}{\theta}) = \frac{1}{N^2}\sum_{a=1}^{N} \mathrm{cov}[\hat{P}_i(\hat{\theta}_a),\hat{P}_i(\hat{\theta}_a)|\underset{\sim}{\theta}] \quad , \tag{8}$$

$$\mathrm{cov}(p_i,p_j|\underset{\sim}{\theta}) = \frac{1}{N^2}\sum_{a=1}^{N}\sum_{b=1}^{N} \mathrm{cov}[\hat{P}_i(\hat{\theta}_a),\hat{P}_j(\hat{\theta}_b)|\underset{\sim}{\theta}] \quad , \tag{9}$$

The $\mathrm{cov}[\hat{P}_i(\hat{\theta}_a)\ ,\ \hat{P}_j(\hat{\theta}_b)|\underset{\sim}{\theta}]$ are evaluated by applying the delta method (Kelley, 1947, pp. 524-526; Kendall and Stuart, 1969, Section 10.6) to (1). For fixed $\underset{\sim}{\theta}$ (for simplicity, the notation " $|\underset{\sim}{\theta}$ " is omitted from the following formula)

$$\begin{aligned}
\mathrm{cov}(\hat{P}_i(\hat{\theta}_a),\hat{P}_j(\hat{\theta}_b)) = w_{ia}w_{jb}\{&t_{ia}t_{jb}[\mathrm{cov}(\hat{\theta}_a,\hat{\theta}_b) - \mathrm{cov}(\hat{b}_i,\hat{\theta}_b) \\
&- \mathrm{cov}(\hat{\theta}_a,\hat{b}_j) + \mathrm{cov}(\hat{b}_i,\hat{b}_j)] + v_{ia}t_{jb}[\mathrm{cov}(\hat{a}_i,\hat{\theta}_b) - \mathrm{cov}(\hat{a}_i,\hat{b}_j)] \\
&+ v_{jb}t_{ia}[\mathrm{cov}(\hat{\theta}_a,\hat{a}_j) - \mathrm{cov}(\hat{b}_i,\hat{a}_j)] + v_{ia}v_{jb}\,\mathrm{cov}(\hat{a}_i,\hat{a}_j) \\
&+ t_{jb}[\mathrm{cov}(\hat{c}_i,\hat{\theta}_b) - \mathrm{cov}(\hat{c}_i,\hat{b}_j)]/1.7 + [v_{jb}\,\mathrm{cov}(\hat{c}_i,\hat{a}_j) \\
&+ v_{ia}\,\mathrm{cov}(\hat{a}_i,\hat{c}_j)]/1.7 + t_{ia}[\mathrm{cov}(\hat{\theta}_a,\hat{c}_j) - \mathrm{cov}(\hat{b}_i,\hat{c}_j)]/1.7 \\
&+ \mathrm{cov}(\hat{c}_i,\hat{c}_j)/(1.7)^2\} \quad ,
\end{aligned} \tag{10}$$

where

$$w_{ia} = \frac{1.7Q_i(\theta_a)}{1 - c_i} \quad ,$$

$$t_{ia} = a_i(P_i(\theta_a) - c_i) \quad ,$$

$$v_{ia} = (\theta_a - b_i)(P_i(\theta_a) - c_i) \quad .$$

The standard errors for p_i were calculated from (5) and again from (8) and (10) for each of the 45 items in the test described in Section 3. The results from the two different approaches agree to at least three significant digits for each item. The $\text{cov}(\hat{p}_i, \hat{p}_j | \underset{\sim}{\theta})$ obtained from (9) and (10) were all of order 10^{-7} or less. This gives us increased confidence in the Lord-Wingersky sampling covariance matrix.

3. Effects of Changing Number of Items, Number of Examinees, or the Frequency Distribution of Ability

To investigate the effect of changing the number of items, the number of examinees, or the distribution of abilities on the sampling errors of parameter estimates, various sets of parameters were specified. The simplest set of parameters represents the administration of a 45-item test to 1500 examinees. The numerical values used as the true θ_a were a spaced sample of 1500 $\hat{\theta}_a$ drawn from the ability estimates obtained by LOGIST for a regular administration of the Test of English as a Foreign Language (TOEFL). A spaced sample of fifteen items were drawn from the sixty TOEFL items whose parameters were estimated in the same run as the abilities. The estimated parameters for these fifteen items were used as the true parameters. These fifteen items were then replicated twice to get a total of 45 items, where items 16-30 and items 31-45 have the same item parameters as items 1-15. Note that various parameters were specified, but no sets of artificial data were generated for this study, since sampling variances and covariances depend only on the true parameters, not on sample observations.

To investigate the effect of increasing the number of examinees, each of 1500 θ_a was repeated four times to represent the θ_a of 6000 examinees. To study the effect of increasing the number of items, another 45 items were added exactly like the first 45 to create a 90-item test. For a different distribution of abilities, a rectangular distribution of 1500 θ_a between -3 and 3 was randomly generated.

Tables 1-4 give the standard errors of the parameter estimates that would be obtained from actual data in the various situations investigated. Only the standard errors for the fifteen unique items are given in the tables of the standard errors for the item parameters. The abilities are grouped into 16 intervals between -4 and 3. Two of the intervals had no examinees. N is the number of examinees and n is the number of items. The values of both the "small" and "capital" parameters are given. The constants to convert from the small scale to the capital scale are $\bar{b}_0 = -.305$ and $k = 0.976$.

Figure 1 contains plots corresponding to these tables. Gaps in the curve for the $\hat{B}_i$ are due to some points out of the range of the plot. The standard error for $\hat{C}_i$ was not plotted against C_i , since most of the C_i were equal, but against $B_i - 2/A_i$ instead. $B_i - 2/A_i$ is an indicator of the ability level at which the item response curve becomes asymptotic. The higher $B_i - 2/A_i$, the better one should be able to estimate C .

As expected, quadrupling the number of examinees halved the standard errors of the estimated item parameters; doubling the number of items, decreased the standard errors of the estimated abilities by a factor of $\sqrt{2}$. Quadrupling the number of examinees reduces the largest standard errors for $\hat{\theta}_a$ sharply, but has little effect on the smaller standard errors; doubling the number of items has only a moderate or

Table 1

Standard Errors for $\hat{A}_i$

Item No.	a_i	A_i	Standard Errors for $\hat{A}_i$			
			Bell-shaped distribution			Rectangular
			n=45 N=1500	n=90 N=1500	n=45 N=6000	n=45 N=1500
1	0.99	0.96	0.234	0.192	0.117	0.178
2	0.35	0.34	0.134	0.131	0.067	0.072
3	1.38	1.34	0.318	0.243	0.159	0.235
4	0.78	0.76	0.147	0.126	0.073	0.099
5	0.42	0.41	0.100	0.106	0.050	0.055
6	0.92	0.90	0.178	0.145	0.089	0.120
7	0.92	0.90	0.179	0.147	0.089	0.119
8	1.06	1.04	0.209	0.168	0.104	0.141
9	1.34	1.31	0.262	0.205	0.131	0.180
10	1.50	1.46	0.317	0.259	0.158	0.231
11	0.87	0.85	0.180	0.151	0.090	0.117
12	0.62	0.60	0.142	0.128	0.071	0.086
13	1.09	1.06	0.234	0.197	0.117	0.153
14	1.39	1.36	0.311	0.265	0.156	0.204
15	1.50	1.46	0.333	0.283	0.166	0.209

Table 2

Standard Errors for $\hat{B}_i$

Item No.	b_i	B_i	Standard Errors for $\hat{B}_i$			
			Bell-shaped distribution			Rectangular
			n=45 N=1500	n=90 N=1500	n=45 N=6000	n=45 N=1500
1	-2.01	-1.75	0.516	0.466	0.258	0.339
2	-1.61	-1.33	2.544	2.344	1.272	1.470
3	-1.09	-0.80	0.353	0.259	0.177	0.242
4	-0.77	-0.48	0.257	0.240	0.128	0.177
5	-0.67	-0.38	0.965	0.929	0.483	0.591
6	-0.34	-0.04	0.191	0.161	0.095	0.141
7	-0.15	0.16	0.165	0.141	0.082	0.128
8	0.00	0.31	0.143	0.117	0.071	0.113
9	0.11	0.42	0.124	0.096	0.062	0.096
10	0.26	0.58	0.110	0.092	0.055	0.097
11	0.46	0.79	0.103	0.101	0.051	0.098
12	0.57	0.90	0.178	0.179	0.089	0.148
13	0.68	1.01	0.085	0.086	0.043	0.086
14	0.90	1.23	0.082	0.080	0.041	0.076
15	1.16	1.50	0.103	0.089	0.052	0.077

Table 3

Standard Errors for $\hat{C}_i$

Item No.	c_i	C_i	Standard Errors for $\hat{C}_i$: Bell-shaped distribution n=45 N=1500	Bell-shaped distribution n=90 N=1500	Bell-shaped distribution n=45 N=6000	Rectangular n=45 N=1500
1	0.17	0.17	0.598	0.469	0.299	0.316
2	0.17	0.17	0.715	0.628	0.358	0.409
3	0.17	0.17	0.096	0.083	0.048	0.045
4	0.17	0.17	0.144	0.123	0.072	0.080
5	0.17	0.17	0.318	0.280	0.159	0.183
6	0.17	0.17	0.071	0.064	0.035	0.039
7	0.17	0.17	0.059	0.054	0.029	0.033
8	0.17	0.17	0.041	0.039	0.021	0.025
9	0.13	0.13	0.026	0.025	0.013	0.018
10	0.34	0.34	0.026	0.026	0.013	0.021
11	0.17	0.17	0.039	0.038	0.020	0.025
12	0.17	0.17	0.068	0.064	0.034	0.039
13	0.25	0.25	0.027	0.027	0.014	0.021
14	0.29	0.29	0.020	0.020	0.010	0.018
15	0.18	0.18	0.015	0.015	0.007	0.015

Table 4

Standard Errors for $\hat{\Theta}_a$

		Standard Errors for $\hat{\Theta}_a$			
		Bell-shaped distribution			Rectangular
θ_a	Θ_a	n=45 N=1500	n=90 N=1500	n=45 N=6000	n=45 N=1500
-2.75	-2.51	2.090	1.478	1.331	1.453
-2.25	-1.99	1.296	0.917	0.879	0.955
-1.75	-1.48	0.861	0.609	0.621	0.669
-1.25	-0.97	0.607	0.429	0.460	0.491
-0.75	-0.46	0.456	0.322	0.373	0.390
-0.25	0.06	0.349	0.247	0.309	0.317
0.25	0.57	0.278	0.196	0.266	0.268
0.75	1.08	0.261	0.185	0.260	0.261
1.25	1.59	0.303	0.214	0.292	0.295
1.75	2.11	0.422	0.298	0.394	0.401
2.25	2.62	0.628	0.444	0.589	0.599
2.75	3.13	0.931	0.658	0.888	0.900

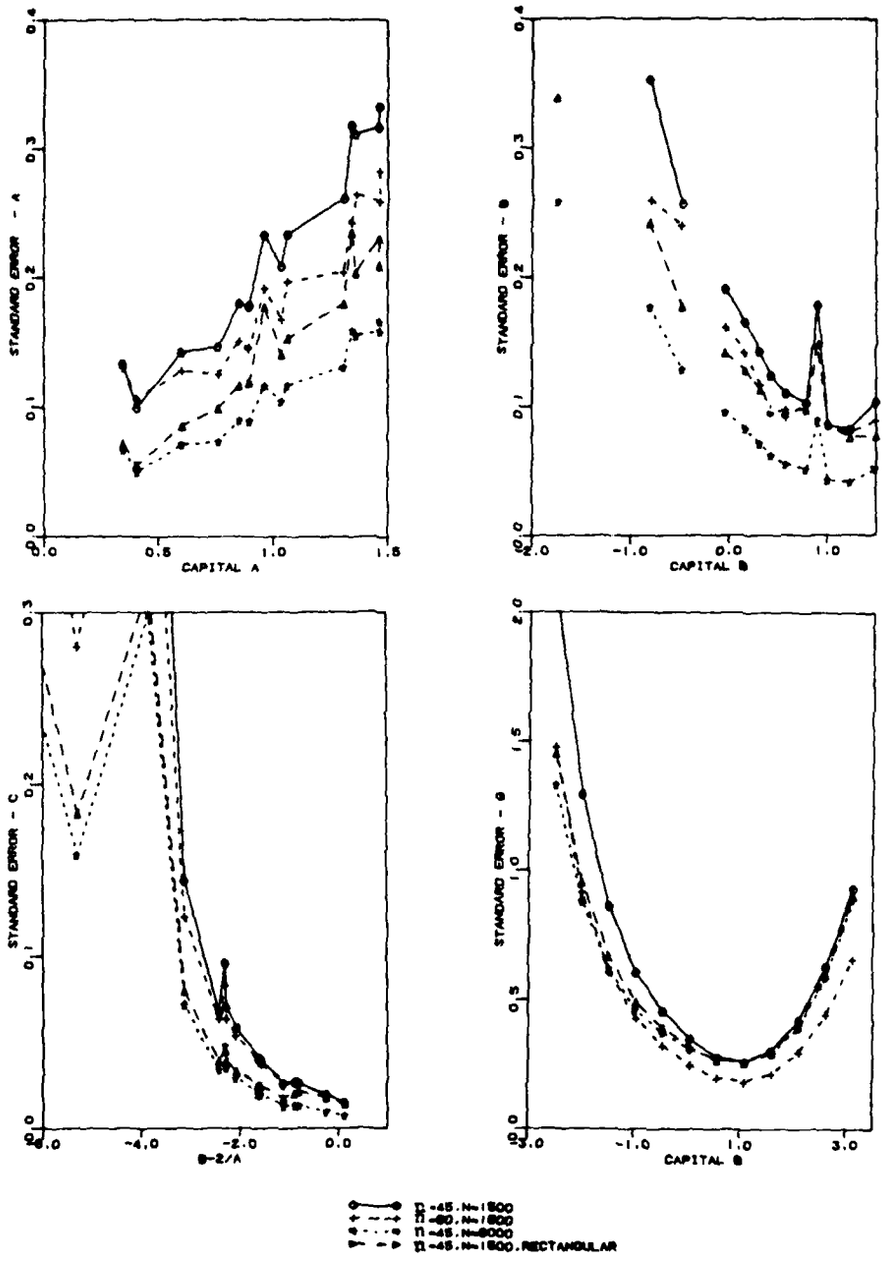


Figure 1. Comparison of the standard for $\hat{A}_i$, $\hat{B}_i$, $\hat{C}_i$, and $\hat{\theta}_a$ for different numbers of items, different numbers of examinees and for a different distribution of examinees

small effect on the standard errors of item parameter estimates. Note that the effects discussed in the previous sentence cannot be investigated at all using the usual standard error formulas, which assume either that the item parameters are known or else that the θ_a are known.

The rectangular distribution of abilities definitely gives better estimates of the item parameters than the bell-shaped distribution of abilities. For C_i where $B_i - 2/A_i$ is low, the rectangular distribution gave standard errors nearly as low as the standard errors with quadruple the number of examinees.

4. Displaying Standard Errors and Sampling Covariances

In looking at tables of standard errors it is hard to see how the standard errors for $\hat{A}_i$, $\hat{B}_i$, and $\hat{C}_i$ interrelate and how the standard errors relate to the magnitude of the parameters. A plot of the three-dimensional asymptotic joint normal distribution of $\hat{A}$, $\hat{B}$, and $\hat{C}$ would be useful but difficult to read. However, projections of the contours of this distribution onto the three two-dimensional planes will give a graphical representation not only of the magnitude of the standard errors but also of the sampling correlations between the parameter estimates. The projected contours are two-dimensional ellipses. These plots are a refinement of a suggestion by Thomas Warm (personal communication, 1982).

For convenience, the subscript i will now be dropped. To plot the projection of the three dimensional contour onto the $(\hat{A},\hat{B})$-plane, only $\text{var}(\hat{A})$, $\text{var}(\hat{B})$, and $\text{cov}(\hat{A},\hat{B})$ are needed. The exponent of

the asymptotic bivariate normal distribution of $\hat{A}$ and $\hat{B}$ is given by the right side of (11). The quadratic in brackets is asymptotically distributed as chi square with 2 degrees of freedom. The 95th percentile for a χ^2 with 2 degrees of freedom is 5.99. Thus 95 percent of the time the obtained $(\hat{A},\hat{B})$ will lie within the ellipse given by the equation

$$5.99 = \frac{1}{1-\rho^2}\left[\frac{(\hat{A}-A)^2}{\mathrm{Var}(\hat{A})} - \frac{2\rho(\hat{A}-A)(\hat{B}-B)}{\sqrt{\mathrm{Var}(\hat{A})\,\mathrm{Var}(\hat{B})}} + \frac{(\hat{B}-B)^2}{\mathrm{Var}(\hat{B})}\right] \qquad (11)$$

where

$$\rho = \frac{\mathrm{Cov}(\hat{A},\hat{B})}{\sqrt{\mathrm{Var}(\hat{A})\,\mathrm{Var}(\hat{B})}} \quad .$$

Similar equations apply for the projections onto the $(\hat{A},\hat{C})$ - and $(\hat{B},\hat{C})$ - planes. The ellipse plotted from (11) for a given N is identical to the 53-percent ellipse that would be plotted for a sample size $N/4$.

The following procedure was used to plot a representative set of ellipses. A hypothetical test of 60 items was created by selecting 60 items from an operational SAT mathematics test and treating these item parameter estimates as the true parameters. A standard normal distribution of 1000

abilities was generated. We then created 15 new items with all combinations of the parameters $a = .5, 1.0, 1.5$; $b = -2, -1, 0, 1, 2$; and $c = .15$. Using these new items, fifteen 61-item tests were created, each containing the 60 original items and one of the new items. The sampling variance-covariance matrix for each of the fifteen 61-item tests was obtained. These matrices differ only because the 61st item differs for each matrix. Only the variances and covariances for the 61st item were used in (11) to compute the ellipses.

The plots were made for an N of 16,000 to avoid confusing overlap of the ellipses. These ellipses are also the 53% confidence ellipses for an N of 4000. The left and bottom axes are labeled with the "small" scale, the right and top axes are labeled with the "capital" scale. The standard errors used are for parameter estimates on the capital scale. The transformation parameters to transform from the small to the capital scale are $\bar{b}_0 = .001$, $k = 1.336$. The center of the ellipse is marked by a "+".

Figure 2 shows the ellipses on the $(\hat{A},\hat{B})$ -plane . The plot shows that the standard error of $\hat{A}$ increases with A . The standard error of $\hat{B}$ increases as B approaches the extremes. The sampling correlation between $\hat{A}$ and $\hat{B}$ is moderately or strongly positive for easy items and moderately or strongly negative for hard items.

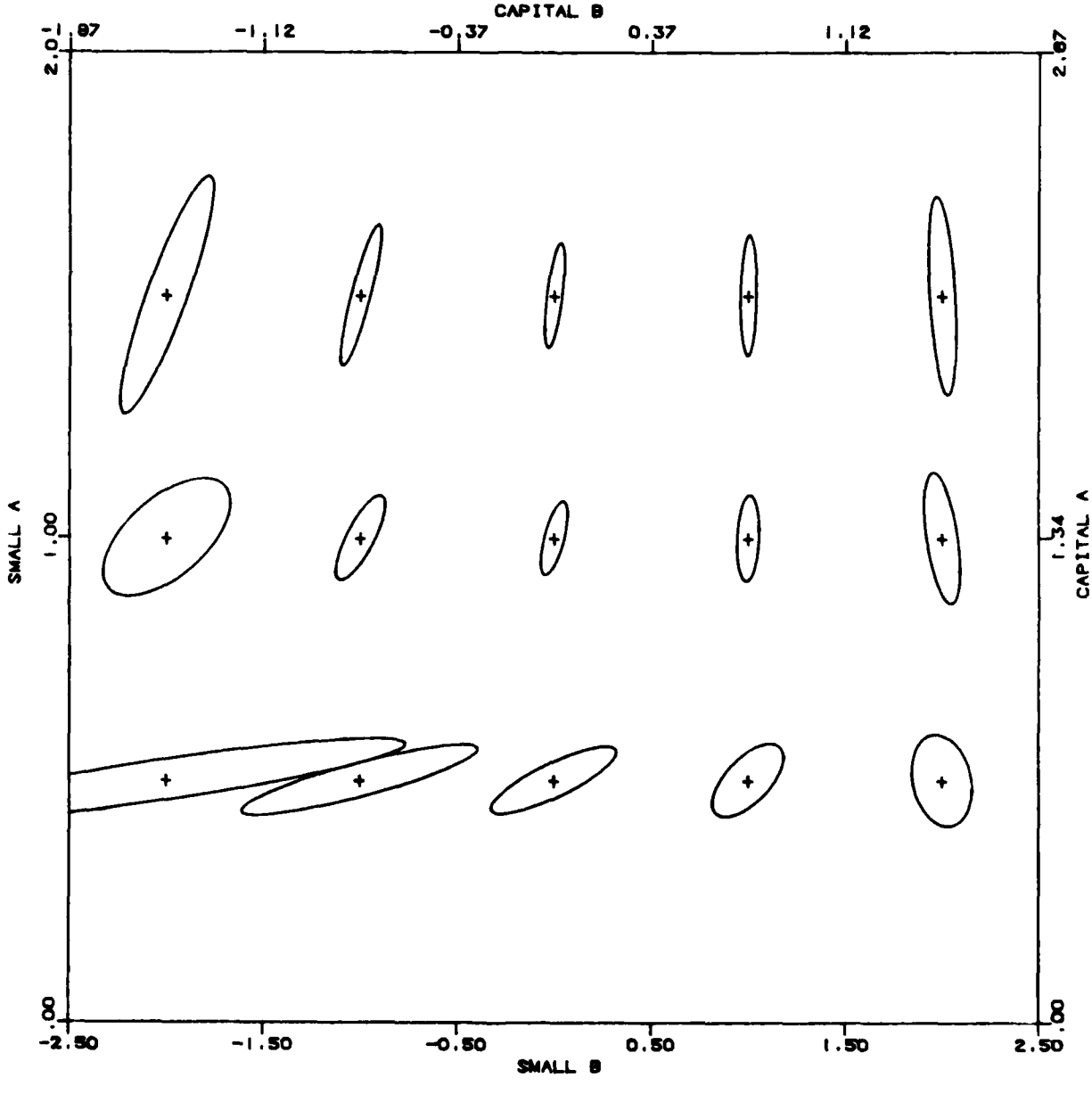


Figure 2. Projections onto the $(\hat{A},\hat{B})$ -plane of the 95% ellipses for an N of 16,000.

Figure 3 shows the projections onto the $(\hat{B},\hat{C})$ -plane. At each value of B there are three ellipses, which are concentric because $c = C = .15$ for all items. The longest ellipse along the C axis is for $a = .5$, the middle ellipse is for $a = 1.0$, and the shortest is for $a = 1.5$. The other triples of ellipses are similarly ordered on a. The standard error of $\hat{C}$ is large for easy items and moderately small for difficult items; the standard error of $\hat{C}$ decreases as a increases. As a decreases, the sampling correlation between $\hat{B}$ and $\hat{C}$ becomes strongly positive except for hard items where $\hat{C}$ is well determined.

Figure 4 shows the projections onto the $(\hat{A},\hat{C})$ -plane. There are five concentric ellipses for each value of A. The ellipse with the longest c -axis is for $b = -2.0$, the ellipse with the shortest c -axis is for $b = 2.0$. Again $\hat{C}$ has large standard errors for easy items and for items with low a's. For hard items the sampling correlation between $\hat{A}$ and $\hat{C}$ is positive and sometimes high; for easy items, the correlation is negative.

4. Standard Errors for Two Tests with Common Items

Suppose that each of two tests measuring the same ability is administered to a different group of examinees. We want to use item response theory either to put the items for both tests into a common item pool or to equate the two tests. For either purpose it is necessary that all the estimated parameters be on the same scale.

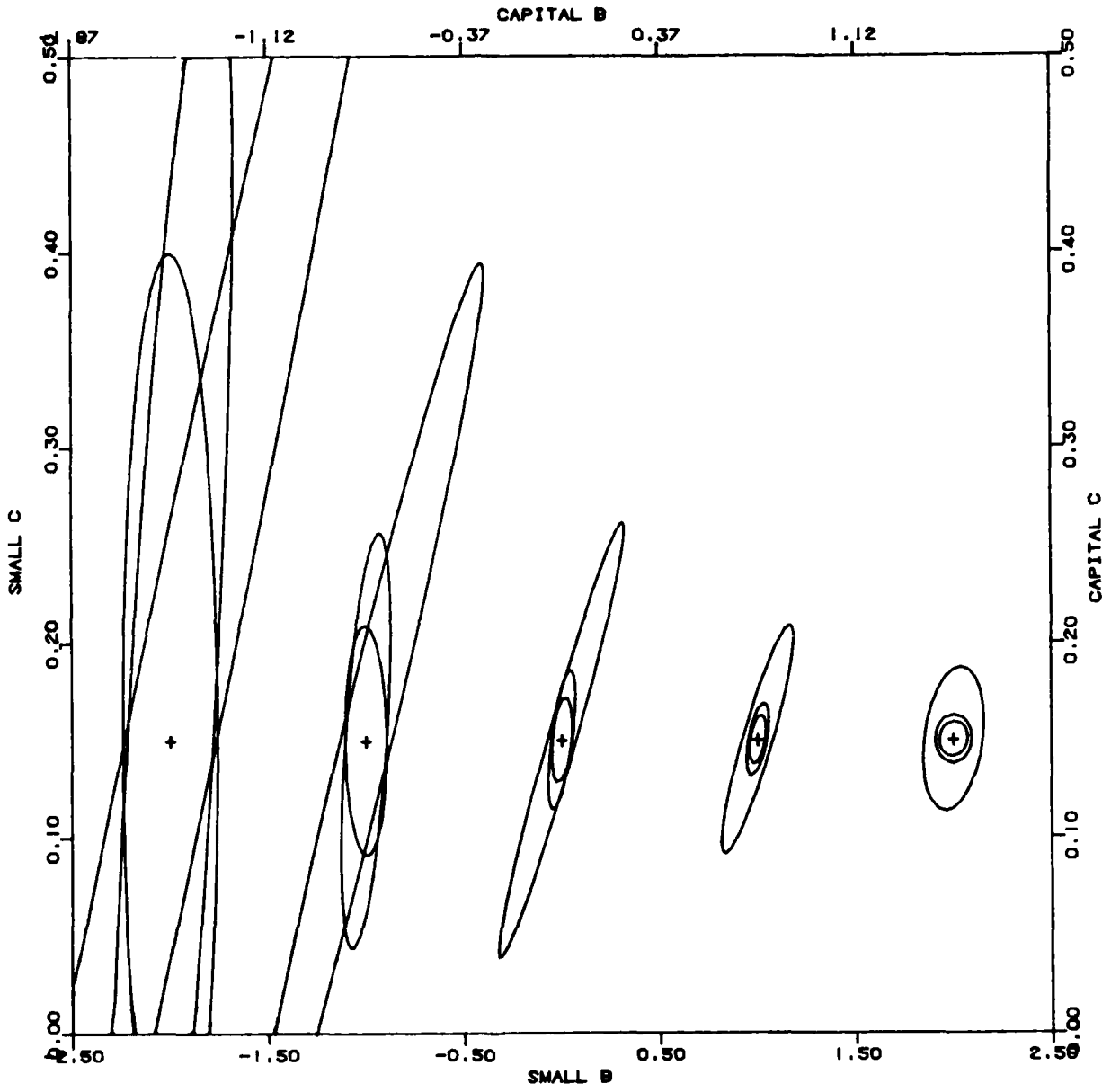


Figure 3. Projections onto the $(\hat{B},\hat{C})$ -plane of the 95% ellipses for an N of 16,000.

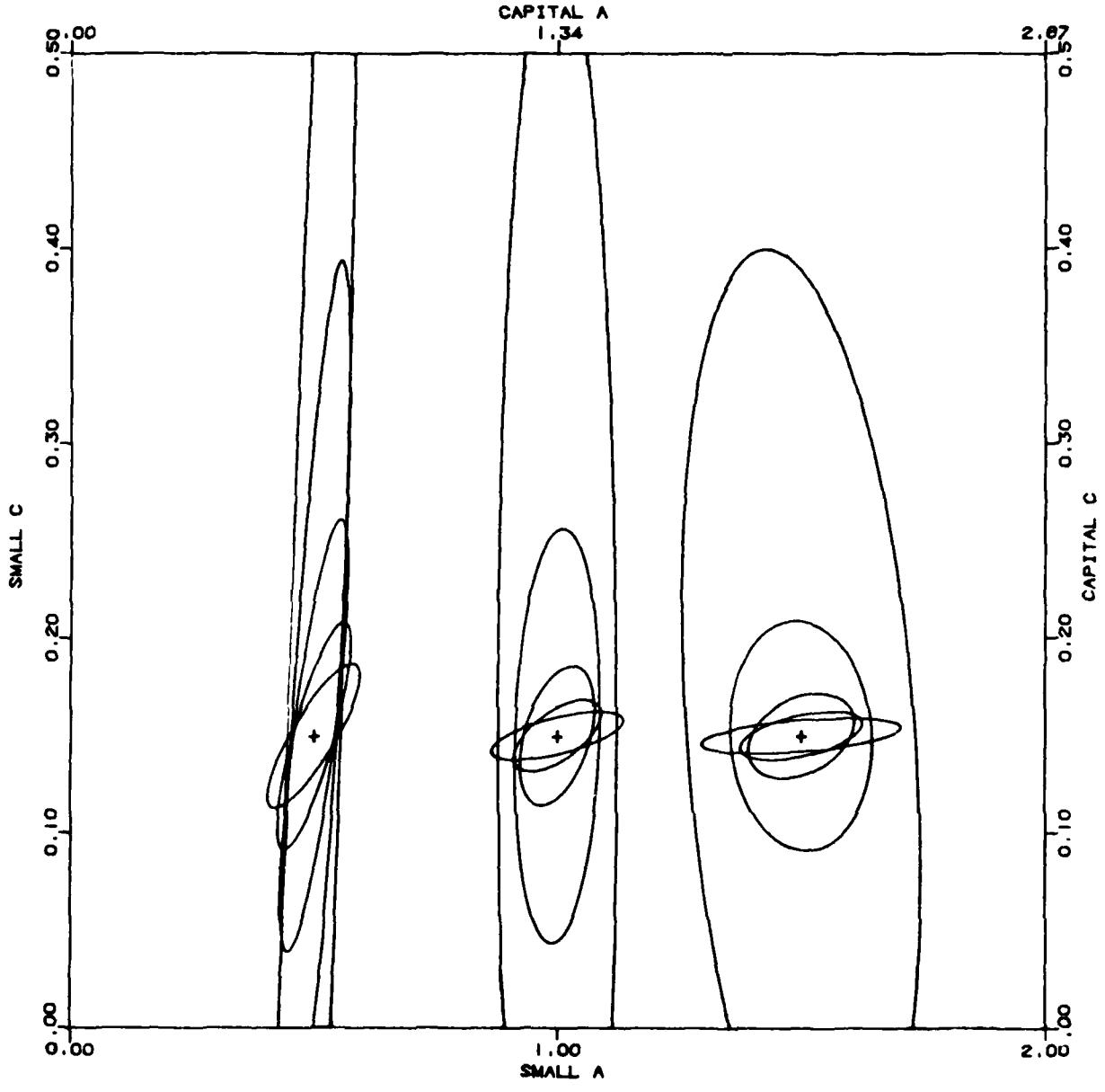


Figure 4. Projections onto the $(\hat{A},\hat{C})$ -plane of the 95% ellipses for an N of 16,000.

Unless equivalent groups of examinees are used, methods for doing this usually require a subset of items that are common to both tests. The unique items are the items in each test that are not common to the other test. The item parameters for each test can then either be estimated separately in two calibration runs or together in one calibration run. If the parameters are estimated in two separate runs, there are two different parameter estimates for each common item. These should be the same except for sampling error and the arbitrary origin and unit of measurement of the ability scale. There are several methods for determining the linear transformation necessary to transform the item parameter estimates for both tests to the same scale. These methods will not be described here (see Stocking and Lord, 1983). However, if all of the items for both tests are calibrated in one run, called a concurrent calibration, the parameters for both tests are automatically put on the same scale and no linear transformation is necessary. This concurrent procedure is most efficient; it provides smaller standard errors and involves fewer assumptions than other procedures. The concurrent procedure is the procedure studied here.

One question that arises when applying the common item method for putting the parameters for both tests on a common scale is: How many common items are necessary? Vale, Maurelli, Gialluca, Weiss, and Ree (1981) investigated this problem using simulated data with 5, 15, and 25 common items and three different shapes of the common item section test information curve: peaked, normal, and rectangular. They also investigated many other linking methods. For the common item method, they assumed that one already had good estimates of the parameters for the common items and required that one have enough common and unique items to get good estimates of the abilities. They

used two estimates of the abilities, one obtained from the common items, the other from the unique items to determine the transformation to put the unique items onto the common scale. They found that 15 to 25 items were necessary and that the common item sections with a rectangular or normal information function were better than those with a peaked information function.

Another study to determine the number of common items necessary was done by McKinley and Reckase (1981). They compared the concurrent method and several other methods for obtaining the linear transformations using the two sets of item parameter estimates for the common items. A large set of items using real data from a multidimensional achievement test covering seven subareas was calibrated in one calibration run and these parameter estimates were used as the criterion for determining how well the other linking procedures put the parameter estimates for subsets of these items on a common scale. A chain of three links was created, that is, test A was linked to test B through one set of common items, test B to test C through another set of common items, and test C to test D through a third set. Five sample sizes ranging from 100 examinee to 2000 examinees were used. All four tests were then calibrated in one run for the concurrent method for each sample. The linking was done with 5, 15 and 25 common items. Each individual test was 50 items long including the common items. McKinley and Reckase concluded that 5 items were not adequate, 25 items were better than 15, but 15 were adequate for linking with the concurrent method.

Given the sampling variance-covariance matrix for all parameter estimates in our single concurrent run when all parameters are treated as unknown, we

can investigate what effect the number of common items has on the sampling standard errors of the unique items in both tests. Note that this problem cannot be investigated at all with the limited sampling-error formulas that assume that either item or ability parameters are known.

Numerical Procedures

Suppose test 1 has a section of unique items labeled V4 , and test 2 has a section of unique items labeled Z5 . Both tests have the same set of common items labeled C0 . One group of examinees, group X , took test 1, another group of examinees, group Y , took test 2. The information matrix $\|I_{pq}\|$, which must be inverted to get the variance-covariance matrix, has the following structure (Lord and Wingersky, 1983):

$\|I_{pq}\| =$

		Items			Examinees	
		V4	C0	Z5	Group X	Group Y
Items		S_{11}	0	0	F_{11}	0
		0	S_{22}	0	F_{21}	F_{22}
		0	0	S_{33}	0	F_{32}
Examinees		F_{11}	F_{21}	0	T_{11}	0
		0	F_{22}	F_{32}	0	T_{22}

The S submatrices (S_{11} for the V4 items; S_{22} for the common items; S_{33} for the Z5 items) contain 3 x 3 Fisher information matrices for a_i , b_i , c_i on the diagonal. The T submatrices are the diagonal information matrices for the examinees: T_{11} for the examinees that took test 1; T_{22} for the examinees that took test 2. The F submatrices contain the vectors $\underset{\sim}{f}_{ia}$, each of which is the 3 x 1 Fisher information vector for item i and examinee a . Note that for Group Y , this is $\underset{\sim}{0}$ for the V4 items; for Group X , this is $\underset{\sim}{0}$ for Z5 .

The matrix $\|I_{pq}\|$ is inverted by grouping the abilities for group X into sixteen groups and by grouping the abilities for group Y into another set of sixteen groups. Then the formulas for inverting a partitioned matrix using the method described in Lord and Wingersky (1983) are successively applied.

Data and Results

To study the effect of the number of common items on the standard errors of the parameter estimates for the unique items, we selected two 60-item SAT Mathematics tests with an additional 25-item common-item section. The 60 unique items in the first test will be referred to as V4 and the 60 unique items in the second test will be referred to as Z5 . Estimates of all of the parameters were obtained in one concurrent LOGIST run. These estimates were treated as true parameter values in computing the standard errors for all 145 items.

We then doubled the length of the common item section by simply replicating the parameters for the 25 common items. Surprisingly, the standard errors for the 120 unique items in V4 and Z5 computed with 50 common items agreed with the standard errors computed with only 25 common items to two decimal places. If doubling the number of common items makes so little difference, what is the effect of halving the number of common items? Or at the extreme, reducing the number of common items to 2?

This is really not as absurd as it sounds. Providing the common items are not part of the test score, other than improving the estimates of the abilities, the function of the common items is to put the parameters for the two sets of unique items on the same metric. If the model holds, only a linear transformation is required to convert the parameters from one scale to another. Only 2 parameters are necessary to determine this linear transformation. With 2 common items we are estimating four parameters that affect the scale, the two a's influence the scale unit and the two b's influence both the scale unit and origin. The two c's are not affected by the scale. Consequently with 2 items we actually have two more parmeters than absolutely necessary. However, if the 2 common items have parameter estimates with large standard errors, the scale will be less well determined than if the estimates have small standard errors.

To study the effect of two common items on the standard errors of the unique items, we selected 2 "good" items and 2 "bad" items from the 25 common items. The item parameters and their standard errors for the 2 "good" items were

a	SE($\hat{A}$)	b	SE($\hat{B}$)	c	SE($\hat{C}$)
.98	.09	-.10	.02	.06	.02
.96	.10	.21	.02	.15	.02

The item parameters and their standard errors for the 2 "bad" common items were

a	SE($\hat{A}$)	b	SE($\hat{B}$)	c	SE($\hat{C}$)
.32	.10	-1.51	.47	.07	.24
.53	.07	-1.19	.12	.07	.10

These standard errors were computed for the situation where all 25 common items are included in the parameter estimation run.

We then obtained the variance-covariance matrix for the V4 and Z5 items when only the 2 good common items are included in the estimation run and also the variance-covariance matrix when only the 2 bad common items are used. The constants to transform from the small scale to the capital scale are $\bar{b}_0 = -.261$ and $k = 1.914$. Only V4 and Z5 items were used to compute $\bar{b}_0$ and k so that the same transformation would apply to all four variance-covariance matrices.

Table 5 gives the medians, and the bottom and top quartiles of the standard errors for $\hat{A}$, $\hat{B}$, and $\hat{C}$, for the Z4 and V5 unique items computed for four different situations: using 50 common items, using 25 common items, using 2 good common items, and using 2 bad common items. Using 2 good common items gives smaller standard errors for the unique items than using 2 bad common items. The standard errors using the 2 good items

Table 5

Comparison of the Standard Errors of Estimated Item Parameters across the Four Sets of Common Items

	50 Common Items	25 Common Items	2 Good Common Items	2 Bad Common Items
Standard Errors for A				
First Quartile	0.114	0.115	0.123	0.131
Median	0.140	0.141	0.151	0.163
Third Quartile	0.224	0.226	0.236	0.243
Standard Errors for B				
First Quartile	0.029	0.030	0.034	0.041
Median	0.042	0.042	0.048	0.056
Third Quartile	0.066	0.067	0.072	0.076
Standard Errors for C				
First Quartile	0.013	0.013	0.013	0.013
Median	0.027	0.027	0.028	0.027
Third Quartile	0.055	0.055	0.058	0.056

are not much larger than the standard errors using 25 common items. Even reliance on just 2 bad common items gives surprisingly good results. Since the purpose of the common items is to determine the scale, it is not surprising that the number of common items has a negligible effect on the standard error of $\hat{C}$, since c is independent of the ability scale.

Table 6 gives the standard errors for the abilities computed with the four different sets of common items. Not surprisingly, if we increase the number of common items to 50 we reduce the standard error of the abilities, although not uniformly as shown by the ratio column. The standard error for the abilities at -2 were lower when computed using the two bad common items, which were easy items, than when computed using the two good common items.

Even though there is little difference between the standard errors when there are 2 common items and when there are 25 common items, the parameter estimates for the V4 and Z5 items will not have been adequately put on the same scale if all of the parameter estimates for V4 items err in one direction and all of the parameter estimates for Z5 items err in the opposite direction. Is this what will happen in practice? To determine how well an anchor test of only 2 common items puts tests V4 and Z5 on the same scale, we reestimated the parameters twice, once in a LOGIST run with the items for Z5 and V4 and the two "good" common items, the other in a LOGIST run with the items for Z5 and V4 and the two "bad" common items.

The estimated parameters for Z5 and V4 computed with the 25 common items will be used as the criterion for evaluating the calibrations

Table 6

Comparison of the Standard Errors of Estimated Abilities across the Four Sets of Common Items

θ_a	Θ_a	50 Common Items S.E	25 Common Items S.E.	Ratio	2 Good Common Items S.E.	2 Bad Common Items S.E.
2.00	1.18	0.097	0.109	0.894	0.127	0.132
1.00	0.66	0.089	0.102	0.870	0.122	0.126
0.0	0.14	0.100	0.115	0.874	0.134	0.138
-1.00	-0.39	0.129	0.145	0.892	0.165	0.167
-2.00	-0.91	0.221	0.248	0.891	0.288	0.281

with 2 common items. The 2 good common items did fairly well at putting the parameters on this scale. The 2 bad items did not do so well. The top plot in Figure 5 compares the b's for the 60 unique V4 items estimated with 2 good items with the b's estimated with 25 common items. Similarly, the bottom plot compares the $\hat{b}$'s for the unique Z5 items. If the parameters were on the same metric the $\hat{b}$'s in both plots should fall on a 45° line. The difference from the 45° line is hard to distinguish. The two points for Z5 that are far away from the 45° line had the c's fixed by LOGIST at the common $\hat{c}$ value in one calibration but not in the other.

Figure 6 shows the plots for the $\hat{a}$'s for V4 and Z5 respectively. Here it definitely looks as if the $\hat{a}$'s are not on the same scale. The $\hat{a}$'s for the V4 items have a slope greater than 45°.

Figure 7 compares the b's estimated with the 2 bad common items with the b's estimated with 25 common items. Here the points for the V4 items are above the 45° line, and points for the Z5 items are below the line. The plots comparing the $\hat{a}$'s in Figure 8 confirm that the 2 bad common items do not put the parameters for Z5 and V4 on the same metric. As suspected, with the 2 bad items the parameters for one set of the unique items err in one direction and for the other set, in the opposite direction.

The reason for putting Z5 and V4 on the same scale was to equate Z5 to V4 using true-score equating. What effect does using only 2 common items to put the two forms on the same scale have on the true-score equating

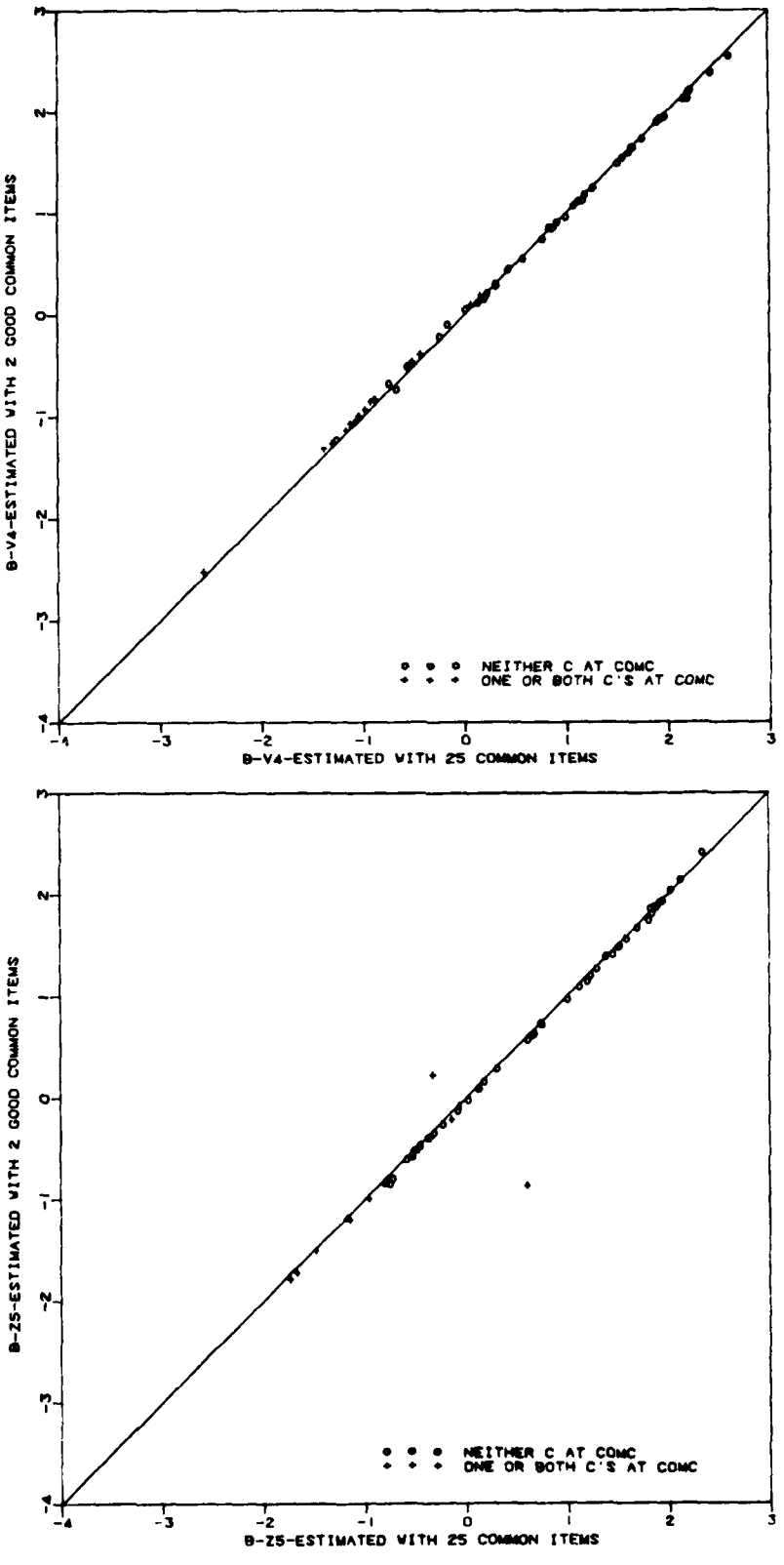
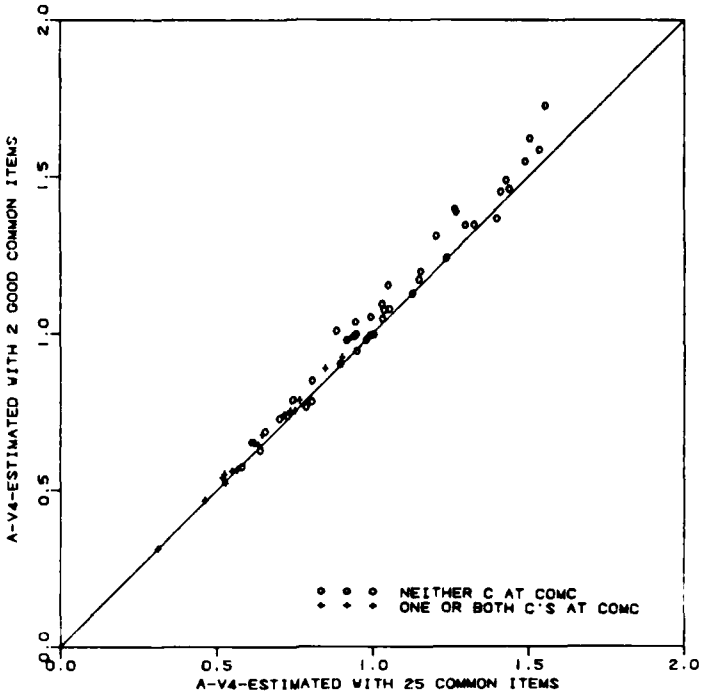


Figure 5. Comparison of the b 's estimated with 2 good common items and the b 's estimated with 25 common items, separately for V4 and Z5.



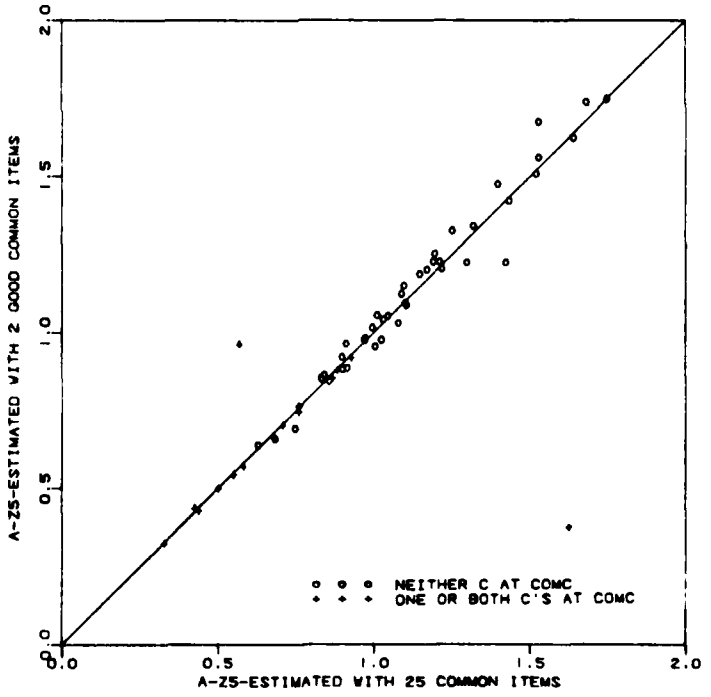


Figure 6. Comparison of the a 's estimated with 2 good common items and the a 's estimated with 25 common items, separately for V4 and Z5.

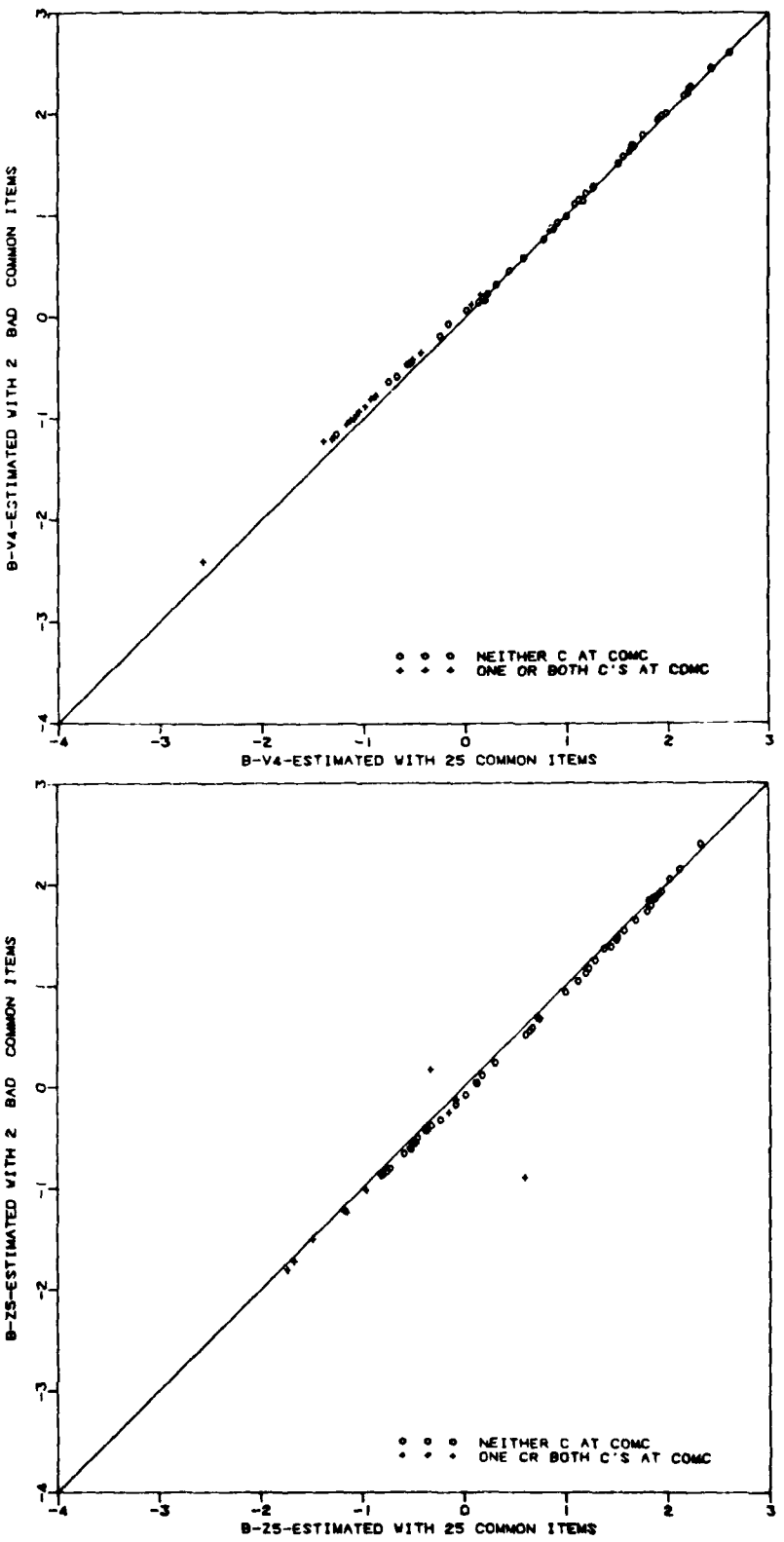


Figure 7. Comparison of the b 's estimated with 2 bad common items and the b 's estimated with 25 common items, separately for V4 and Z5.

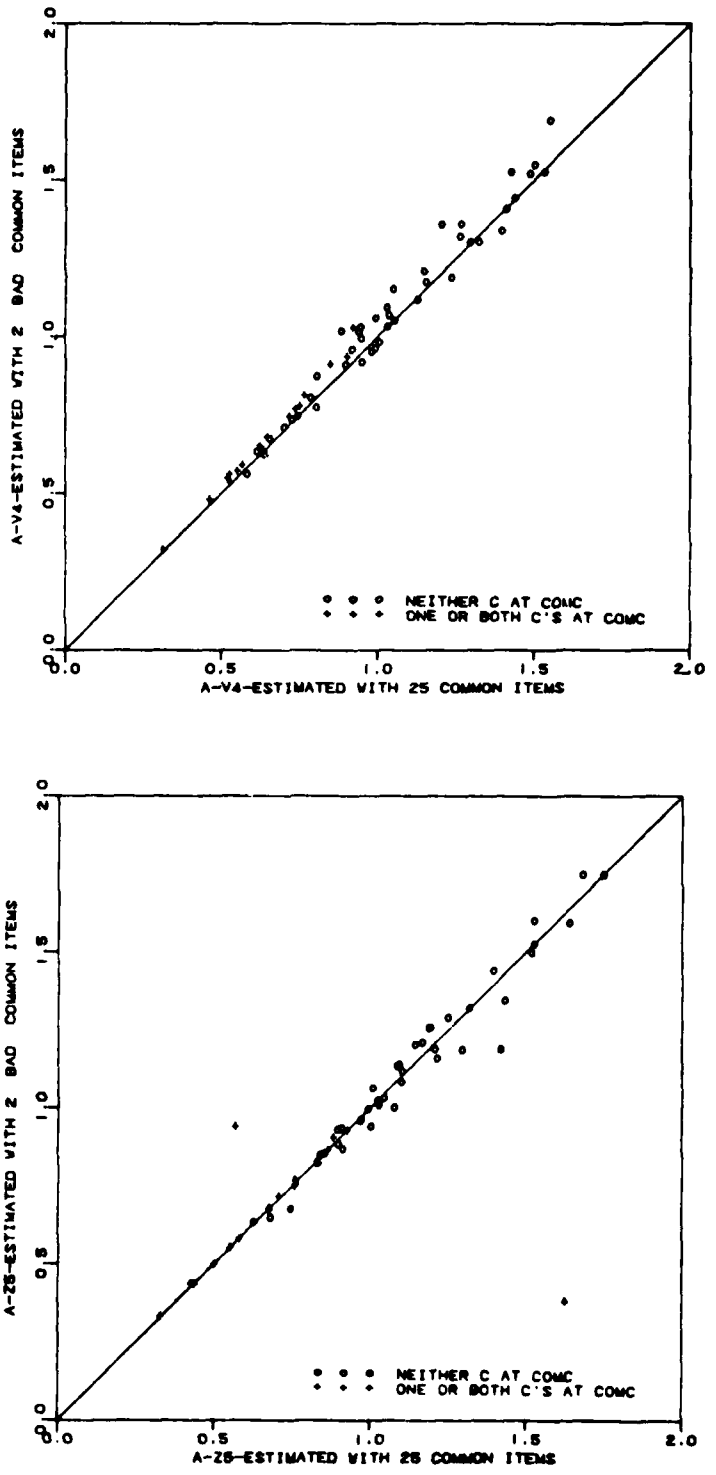


Figure 8. Comparison of the a's estimated with 2 bad common items and the a's estimated with 25 common items, separately for V4 and Z5.

between the two forms? Figure 9 shows three true-score equating lines: the solid line is the equating line found when the parameters are estimated using 25 common items, the dotted line is the equating line found when the parameters are estimated using just the 2 good common items, the dashed line is found when the parameters are estimated using just the 2 bad common items. For this equating, true scores on form Z5 are first equated to true scores on V4 . Then the true scores on V4 are converted to scaled scores between 100 and 800 by a linear transformation. Using the equating line with the 25 items as a criterion, the equating using 2 bad common items is worse than the equating using 2 good common items. The equating using the 2 good common items is close to the equating with 25 common items; the maximum scaled score difference is 8 points.

All of these results assume that the item parameters estimated using 25 common items are on the same scale. This analysis should be repeated in a situation where one knows that all of the parameters used as a criterion are on a common scale. From the results so far, it appears that good linking may be obtained with as few as five common items or less. However, these results only apply when the item parameters for the two forms are put on a common scale by estimating all of them in one calibration run. These results do not apply when the two tests are calibrated in two separate runs and the parameters are put on a common scale using some linear transformation determined from the common items.

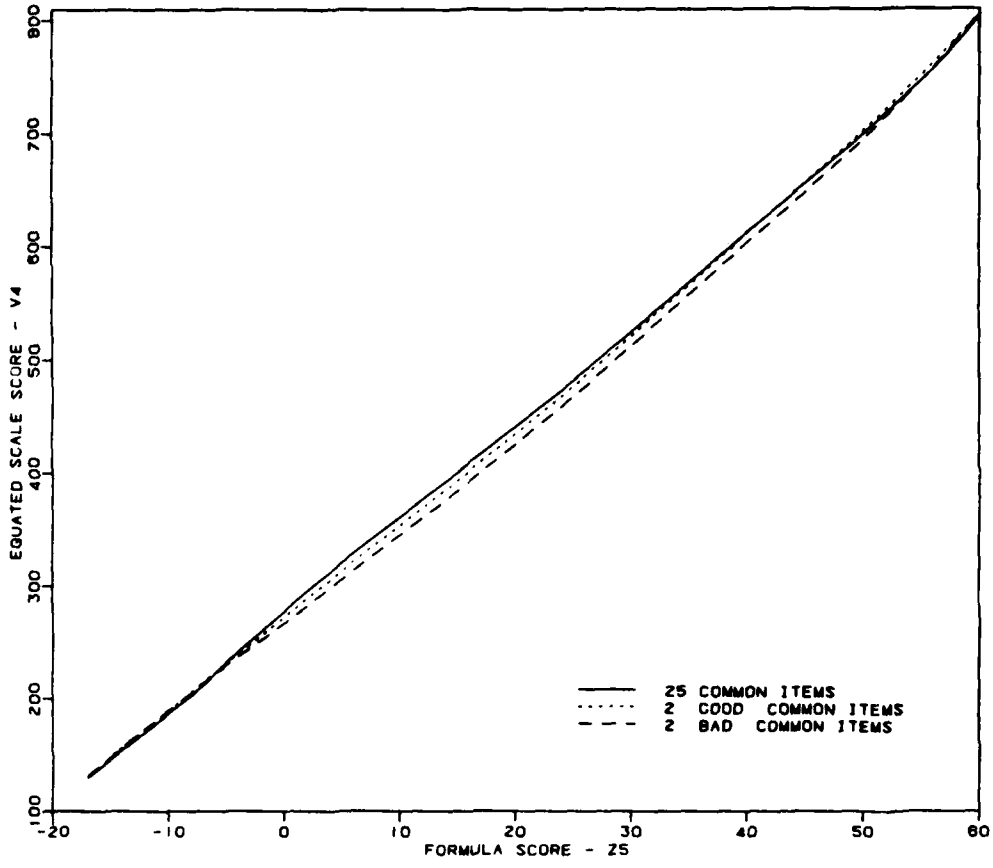


Figure 9. Comparisons of the three true-score equatings of test Z5 to test V4 : using 25 common items, using 2 good common items, and using 2 bad common items.

The conclusion that good linking may be obtained with as few as five common items is more optimistic than the conclusions reached by Vale et al. (1981) and by McKinley and Reckase (1981). Our differences with Vale et al. may be due to the facts that 1) their scaling was based on estimated θ 's, and 2) they used three estimation runs instead of one concurrent run. Our differences with McKinley and Reckase are probably due to the facts that in their study 1) the responses of some examinees to some items (as we understand it) often appeared twice in the same concurrent LOGIST run, violating the assumption of local independence; and, more importantly, 2) they pooled the Iowa Tests of Educational Development covering seven different achievement areas, and analyzed the resulting multidimensional pool of items as if it were unidimensional.

Summary

The asymptotic sampling variance-covariance matrix of maximum likelihood estimators when both abilities and item parameters are unknown was used to study several problems in item response theory, such as the extent to which more items, more examinees, or a different distribution of abilities will provide better estimates of parameters. It was found for the values of n and N studied that that the standard error of $\hat{\theta}$ varies inversely as $\sqrt{n}$, but is only moderately affected by changes in N ; the standard error of the estimated item parameters varies inversely as $\sqrt{N}$, but is only slightly affected by changes in n .

A rectangular distribution of abilites gives smaller standard errrors for the item parameters than doubling the number of items. In fact, for low A 's, also for C 's for items with B - 2/A less than -1, the standard errors computed with a rectangular distribution of ability were nearly as low as the standard errors computed with a bell-shaped distribution and quadruple the number of people.

With the variance-covariance matrix computed when all parameters are treated as unknown, one can study the effect of the number of common items on the standard errors of the unique items when each of two tests containing common items is administered to a different group of examinees and the parameters for both tests are calibrated in one LOGIST run. This problem cannot be dealt with at all by previously available sampling error formulas. The number of common items has little effect on the standard errors of the parameters for the unique items. The standard errors indicate that as few as 2 items may be sufficient providing the parameter estimates for these two items are well determined. However when two tests were actually calibrated in one LOGIST run using 2 common items that had parameter estimates with low standard errors, the parameters were not quite on the same scale as the parameters estimated with 25 common items. The $\hat{b}$'s were very close to the same scale but the $\hat{a}$'s for one of the tests were on a slightly different scale. Although 2 items are not quite enough, adequate linking may be possible with as few as five items.

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